Corporate Diversification and the Cost of Capital

by

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September 2009

Abstract

We examine whether organizational form matters for a firm's cost of capital. Contrary to the conventional view, our model shows that coinsurance among a firm's business units can reduce systematic risk through the alleviation of countercyclical deadweight costs. 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, diversified firms with less correlated segment cash flows have a lower cost of capital, consistent with a coinsurance effect. Holding the magnitude of cash flows constant, our estimates imply an average value gain of approximately 6% when moving from the highest to the lowest cash flow correlation quintile.

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1. Introduction

The conventional view among practitioners and researchers is that organizational form does not matter for a firm's cost of capital because, while the imperfect correlation of business unit cash flows may help reduce idiosyncratic risk, this should have no effect on systematic risk.[1] In this paper, we present evidence that is contrary to the conventional view. We show that diversified firms have a lower cost of capital than portfolios of comparable stand-alone firms and that the reduction is strongly related to the correlation of business unit cash flows, consistent with a coinsurance effect.

A large body of research inspired by Coase's (1937) fundamental question about the boundaries of the firm points to various costs (Rajan, Servaes, and Zingales (2000), Scharfstein and Stein (2000)) and benefits (Matsusaka and Nanda (2002), Stein (1997)) of integration. In theory, if the cash flows due to integration costs and benefits carry precisely the same systematic risk as the cash flows of the underlying businesses, 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 differ from that of its business units as 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.

We argue that organizational form can matter, and in particular, coinsurance among a firm's business units can reduce the firm's cost of capital. The economic intuition underlying our argument is easily illustrated: (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

[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 (e.g. the stand-alone principle)..

has a systematic component (Elton, Gruber, Agrawal, and Mann (2001), Almeida and Philippon (2007)). Intuitively, if coinsurance enables a diversified firm to avoid systematic financial distress costs that its business units would otherwise incur if they were stand-alone firms, then coinsurance should lead to a reduction in the diversified firm's systematic risk and hence its cost of capital. In this paper, we show in a parsimonious model that the coinsurance idea outlined above is more general and that it extends to an all-equity firm if one replaces costs of financial distress with other kinds of systematic 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 reduce systematic risk when deadweight costs are partly systematic.[2] In addition, we show that the coinsurance effect is stronger when the firm's units have less correlated cash flows.[3]

We examine the connection between organizational form and a firm's cost of capital using a 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 expected returns for both components of financing: implied cost of equity constructed from analyst forecasts to proxy for expected equity returns and yields from the Barclays Capital Aggregate Bond Index to proxy for expected debt returns.[4] Thus, our study avoids the many pitfalls of using ex post measures such as stock or bond returns as proxies for expected returns

[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 could 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.

[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 understate the effect of coinsurance on total cost of capital.

(Elton (1999)). 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 a significantly lower cost of capital compared to portfolios of stand-alone firms, rejecting the conventional view that organizational form does not matter for a firm's cost of capital. To explore whether the difference is due to coinsurance, we consider the correlation of cash flows among a firm's segments as an inverse measure of coinsurance. Consistent with a coinsurance effect, we find a significant and positive association between excess cost of capital and cross-segment cash flow correlations. These findings are robust to using alternative measures of implied cost of equity capital (Claus and Thomas (2001), Easton (2004)) and coinsurance (Duchin (2008)). These findings are also economically significant. 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.[5]

Our estimates of implied cost of equity are subject to potential measurement errors arising from analyst forecast bias.[6] We therefore perform various sensitivity tests to address this issue. First, we control for analyst forecast errors in our main multivariate regression analysis and find similar results. Second, we perform analysis based on Easton and Monahan's (2006)

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[5] It is possible that these estimates represent a lower bound on the effect of coinsurance because our proxies are limited to segment data and do not capture coinsurance among different product lines or geographic areas.

[6] Our analysis is also subject to a potential self-selection bias, an issue that has been addressed extensively in the diversification discount literature (Campa and Kedia (2002), Graham, Lemmon, and Wolf (2002), Villalonga (2004)). However, it is unclear how a strong monotonic relation between our continuous coinsurance measures and excess cost of capital would be driven by a dichotomous selection mechanism that pushes some business units to conglomerate. Nevertheless, we perform a sensitivity test using Heckman's two-stage method to correct for a potential selection bias and we find similar results.

finding that implied cost of equity estimates are generally reliable when analysts' forecast accuracy is high. We partition our sample based on absolute forecast errors and find that our results are strongest in the subsample with the lowest errors, which suggests that our findings are weakened by forecast errors, rather than induced by them. Finally, we use Fama-French factor loadings to estimate cost of equity capital and find remarkably similar results.

We believe our study is the first to establish a link between 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 (1971), a stream of research studies coinsurance in the context of 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)), a hypothesis supported by the findings of Mansi and Reeb (2002) based on segment disclosures. More recently, Duchin (2008) studies the relation between coinsurance and firms' cash retention policies. Our paper combines with Duchin's paper to form a nascent literature examining the implications of coinsurance for corporate finance in general.

Our study also complements the literature on corporate diversification and firm value (Lang and Stulz (1994) and Berger and Ofek (1995)) by exploring an important dimension that thus far has received little attention, namely, cost of capital. 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. 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. By exploring whether the cross-sectional variation in cost of capital is due to coinsurance, our work deepens the foundations of this literature.

Our work is also related to Ortiz-Molina and Phillips (2009), who find that firms with more liquid real assets have a lower cost of capital using the implied cost of equity developed by Gebhardt et al. (2001). To the extent their measure of real asset liquidity is inversely related to deadweight costs that firms incur when selling assets, their findings confirm our model assumption that deadweight costs have a systematic component. Benmelech and Bergman's (2009) recent work showing that debt tranches of airlines secured with more redeployable collateral have higher credit ratings and lower credit spreads also supports this notion.

Our evidence also 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)).[7]

[7] 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

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.

Our basic model assumes all-equity financing and an efficient internal capital market to illustrate the coinsurance effect. In the Appendix, we relax these assumptions and extend the basic model to incorporate debt financing and the possibility of rent-seeking activities and inefficient transfers in internal capital markets.

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.

adjustments in capital budgeting. Schall's analysis rules out by assumption the possibility that synergistic cash flows may be a function of covariance.

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 equilibrium, 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

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

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

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

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* $t=0$*:*

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

Let $\beta^i \equiv (C_g^i / C_b^i - 1)$. Note that β^i is monotone in the conventional measure of systematic risk $-\operatorname{cov}(C^i, m)$ because $m_g < m_b$. This means that we can use β^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 relation.

Lemma 1 *Given any equilibrium summarized by* (p_g, m_g, m_b), $E[r^i]$ *depends only on* β^i *and it 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. 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 the first-best amount of financial slack to avoid any future deadweight cost.

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 costly defections by important stakeholders such as suppliers and customers, in which case the firm incurs a deadweight loss L and investors receive 0.[8]

Suppose further 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 $\beta^S \equiv (C_g^S / C_b^S - 1)$, where $C_g^S = \theta_g H$ and $C_b^S = \theta_b H$. A stand-alone firm whose $\theta_g > \theta_b$ carries positive systematic risk, whereas a stand-alone firm whose $\theta_g < \theta_b$ carries negative systematic risk. Accordingly, 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.

[8] The assumption that investors receive 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.

Combining Two Stand-alone Firms into One Diversified Firm

Suppose that two identical stand-alone firms can be combined under one roof.[9] 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 the low-outcome project, or alternatively, the firm can 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.[10] 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

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

Without the terms involving L, the expected cash flow of a diversified firm C_e^D equals twice the expected cash flow of a stand-alone firm, $2C_e^S$, 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.

[9] 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*.

[10] 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.

As the next proposition shows, one implication of coinsurance may be to reduce systematic risk in addition to increasing cash flows $(C_g^D > 2C_g^S, C_b^D > 2C_b^S)$.[11]

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,

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

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

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

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

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 (i.e., $(1-\theta_g)$ in the good state and $(1-\theta_b)$ in the bad state). In other words, deadweight costs are partly systematic. Since β^D is a weighted average 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 zero.

[11] Coinsurance effects arise even if we introduce integration costs to make the analysis cash-neutral. We incorporate integration costs in the model in the Appendix.

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 $w \in \{1,..,N\}$ 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 (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 case of correlated idiosyncratic risks 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:

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

The case in which ρ equals 0 corresponds to the case of independence in Proposition 1. 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 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* ρ*, and reach those of a stand-alone firm in the limit when* ρ *equals* 1.

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

Using the new probability structure,

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

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 2 demonstrates that 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 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. 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.

2.3. Testable Predictions

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

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 of 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 motivation for the latter test is that coinsurance may lower systematic risk through internal capital markets that help firms avoid the deadweight costs of external financing by channeling resources to business units with superior investment opportunities (Matsusaka and Nanda (2002)).

3. Research Design

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 (Elton (1999)). To address this concern, recent literature in accounting and finance has developed an ex ante approach to measuring expected 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.[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

[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)).

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

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

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 linearly to the industry median from year 3 until year T, and all other variables are as defined previously.

[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.

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 (where we use Fama and French (1997) industry definitions). The cost of equity is calculated numerically by 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

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[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).

(2001), who define total cost of capital as the weighted average of a firm's realized equity return and the return on an aggregate bond index. Instead of using realized equity and bond returns, however, we use ex ante measures of the implied cost of equity and bond yields to proxy for expected equity and debt returns, respectively. More specifically, the COC for each firm *i* and year *t* is computed as follows:

$$COC_{i,t} = D_{i,t-1} Y_t^{BC} + (1 - D_{i,t-1})\, COEC_{i,t},$$

where Y_t^{BC} is the aggregate bond yield from the Barclays Capital Aggregate Bond Index (formerly, the Lehman Brothers Aggregate Bond Index), $COEC_{i,t}$ is the implied cost of equity (GLS), and $D_{i,t-1}$ is the firm's book value of debt divided by total value (book value of debt plus market value of common equity).[15]

This cost of capital measure has the limitation that our proxy for the cost of debt does not capture any firm-specific variation in expected debt returns.[16] To the extent that coinsurance reduces the cost of debt (which we show in the Appendix), our results understate the coinsurance effect on cost of capital. Despite this limitation, 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

To compare the cost of capital of a diversified firm to the cost of capital that its segments would have if they were stand-alone businesses, we compute excess COC as the natural logarithm of the ratio of the firm's COC to its imputed COC (defined below). An excess COC

[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 because bond yields reflect both systematic and idiosyncratic risk.

below (above) zero is consistent with diversification reducing (increasing) the firm's cost of capital.

We calculate a firm's *imputed* COC as a value-weighted average of the imputed COC of its segments:

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

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

The procedure for estimating segments' imputed market values is described in detail in Berger and Ofek (1995). In short, the estimation consists of: (1) estimating the median ratio of enterprise value to sales for all single-segment firms in the industry to which the segment belongs, and (2) multiplying the segment's sales by the median industry ratio. Industry definitions are 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 among a firm's segments; specifically the correlation among the idiosyncratic part of segments' future free cash flows. A precise measure of coinsurance, however, is difficult to obtain because the distribution of segments' future cash flows is not observable. Moreover, using the distribution of historical segment-level cash flow to estimate coinsurance is problematic because firm composition usually changes over time. Accordingly, we construct empirical proxies of coinsurance using industry-level cash flow series based on single-segment firms. To ensure that estimated correlations are not contaminated with systematic risk, we perform the computation in two stages.

First, for each 2-digit SIC code industry in a given year in our sample, we compute idiosyncratic industry-level cash flows for the prior ten years as residuals from a regression of average industry cash flow on average market-wide cash flow over the same period.[17]

Next, for each year in our sample, we estimate correlations between every industry pair based on the prior ten-year idiosyncratic industry-level cash flows. We then use these estimated correlations to construct our cash flow coinsurance measure. In constructing our investment coinsurance measure, we use capital expenditures but otherwise follow the same procedure.

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 estimated correlation of idiosyncratic industry cash flows or investments between industries j and k over the ten-year period before year y. We obtain similar results using an alternative coinsurance measure, which also includes the standard deviation of 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.

[17] We measure cash flow as operating income before depreciation scaled by total assets.

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 sample restrictions. 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 be within 1% of the total sales of the firm to ensure the integrity of segment data; (3) all of the firm's segments for a given year have at least five firms in the same 2-digit SIC code industry with non-missing firm value to 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) be 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.

[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.

4.2. Control Variables for Cross-sectional Analysis

Return Patterns

To ensure that our results on the relation between coinsurance and cost of capital are distinct from the well-documented return patterns (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. 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, which could induce a negative correlation between past returns and the cost of capital measures. [19]

In addition, we include I/B/E/S's long-term growth forecast to control for LaPorta's (1996) finding that forecasted long-term growth in earnings is negatively associated with returns.

Analyst Forecast Dispersion

Gebhardt et al. (2001) show that the GLS COC measure is positively correlated with dispersion in analysts' forecasts. Accordingly we control for dispersion in analysts' forecasts, as measured by the log of standard deviation of analyst forecasts.

Leverage

We also control for leverage to account for tax-shield benefits of debt in the weighted average cost of capital. In robustness specifications with cost of equity as the dependent variable, we expect a positive relation due to increased financial risk.

[19] It is possible that we are overcontrolling by including size and the book-to-market ratio in our regressions. First, book-to-market may be associated with coinsurance related forward-looking betas in a conditional asset-pricing model (e.g., Petkova and Zhang (2005)). Second, size may serve as an alternative proxy for coinsurance. Larger firms are likely to have a larger number of unrelated projects, which can lead to greater coinsurance.

We summarize the definitions of the control variables below (numbered items refer to the Compustat annual database):

Log(market capitalization)	=	Natural logarithm of fiscal year-end stock price times shares outstanding from Compustat (#199 * #25);
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*) from CRSP.

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

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, our model predicts that diversified firms 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, though 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. The difference in means between the single- and multi-segment subsamples is negative and significant (0.010 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.[20] The modest result is due to the pooling of all multi-segment firms, many of which operate within a single industry and thus enjoy little cross-segment coinsurance as captured by our measure. Indeed, as presented in the next section, we find economically important cross-sectional differences when we sort multi-segment firms based on cash flow and investment correlations.

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

5.2.1. Nonparametric Univariate Sorts

In Table 2, we sort our sample of multi-segment firms into quintiles based on the two coinsurance measures and report the average excess COC for each quintile. The results from cash flow and investment correlation sorts are reported in the left and right panels, respectively. We also present results for the single-segment firms. Note that single-segment firms can be

[20] 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.

viewed as limit observations with respect to the degree of coinsurance – for these firms, cash flow and investment correlations are equal to one by definition.

Because the results are similar, we focus our discussion on the cash flow correlation sort. Consistent with the coinsurance hypothesis, we observe a monotonic increase in excess COC from the highest coinsurance quintile (Q1) to the lowest coinsurance quintile (Q5) (recall that a higher cash flow correlation means lower coinsurance). The mean difference between Q5 and Q1 is a statistically significant 3.2%, i.e., a cost of capital reduction relative to single-segment firms 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%, consistent with a significant coinsurance effect. Overall, these nonparametric results support Prediction 2 – diversified firms that consist of businesses with less correlated cash flows have a lower total cost of capital.

5.2.2. Multivariate Analysis

Next, we investigate whether the negative relation between excess COC and coinsurance is robust to controlling for the set of firm characteristics discussed in Section 4.2.

In the first set of regressions, Models 1 and 2, we regress excess COC on cross-segment cash flow and investment correlations, respectively, and control for all variables except for the number of segments and the natural logarithm of market capitalization. We exclude these two measures 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 them in the regressions could overcontrol for the coinsurance effect.

In the second set of regressions, Models 3 and 4, we use the number of segments and the natural logarithm of market capitalization, respectively, as alternative measures of coinsurance. As discussed earlier, both are measures of firm size, and hence they are likely to capture some degree of coinsurance that is not captured by the correlation-based measures.

In the last set of specifications, Models 5 and 6, we include all control variables, *including* the number of segments and the natural logarithm of market capitalization, to disentangle other possible "size effects" from the coinsurance effect that is captured by the cash flow and investment correlation measures. We therefore view this last set of specifications as the most demanding test 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 results based on univariate sorts in the previous section, the coefficient on the cash flow correlation measure is positive and significant in both Models 1 and 5 (with $p<0.01$). In Models 3 and 4, we find a negative and significant coefficient on the number of segments and the natural logarithm of market capitalization, respectively, at conventional levels. This result suggests that larger firms and firms with more segments, which may have more product lines with coinsurance potential, 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 support Prediction 2: firms with lower cross-segment cash flow correlations have a lower cost of capital, i.e., the coinsurance effect increases as cross-segment cash flow correlation decreases.

5.2.3. Robustness Tests

Excluding Single-segment Firms

Our main regression analysis in the previous subsection includes both single- and multi-segment firms. To investigate the possibility that our results may be spuriously driven by differences between stand-alone and diversified firms, we perform our main analysis using multi-segment firms only. 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 $p<0.01$). These results suggest that our main finding on coinsurance and cost of capital are unlikely driven by differences between 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 perform the following sensitivity tests.

First, we control for one- 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 ratios, (4) long-term growth forecasts, and (5) investments in property, plant, and equipment. Estimation of the predicted forecast error 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- 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 Panel A of Table 5, continue to show a positive and significant coefficient on the cash flow and investment correlation measures, suggesting that our main findings are unlikely driven by systematic differences in analyst forecast biases between single- and multi-segment firms.

Second, Easton and Monahan (2006) find that the reliability of implied cost of equity estimates increases as analyst forecast accuracy improves. Accordingly, we partition our sample into terciles using absolute forecast errors in one-year-ahead earnings and estimate cost of capital regressions within each sub-sample. The results are reported in Panel B of Table 5. The cost of capital effect is strongest in the subsample with low absolute forecast error. These results suggest that our findings are unlikely driven by measurement errors in the implied cost of equity estimates that are induced by biased forecasts. Rather, our results are weakened by them.

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 introduce two alternative measures of implied COE.

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,[21] g is the terminal growth rate of residual earnings, which is 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 using the 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. Easton makes two simplifying assumptions, namely, 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:

[21] 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 value by applying the long-term growth rate to the two-year-ahead forecast.

$$r_e = \sqrt{g * \frac{FEPS_{t+1}}{P_t}}, \text{ where } g = \frac{(FEPS_{t+2} - FEPS_{t+1})}{FEPS_{t+1}}.$$

The PEG COE can be estimated only for firms where two-year-ahead EPS forecasts exceed one-year-ahead EPS forecasts. In addition, the estimation is restricted 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 the 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 cost of equity are reported in Table 6. Consistent with our earlier findings, the coefficients on the cash flow and investment correlation measures are positive and significant. Overall, our main findings are robust to using CT or PEG COE 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. As such, we employ an empirical proxy that measures the weighted average of the cost of equity and debt capital. 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 using 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 the cash flow and investment correlation measures, suggesting that our main findings are at least partially driven by the cost of equity component. An interesting extension would be to examine whether our results also hold for the cost of debt.[22]

Factor-Model-Based Cost of Equity Estimates

As a further robustness test, we estimate expected returns using the Fama-French three-factor model. To obtain ex-ante estimates of cost of equity at a given point in time, we estimate factor loadings using 24 months of prior excess returns, multiply the estimated factor loadings with corresponding historic risk premiums, and add the yield on the 10-year Treasury note. To deal with low (and sometimes negative) cost of equity estimates, we set cost of equity estimates that are lower than the risk-free rate equal to the risk-free rate.

The results based on Fama-French excess cost of equity are reported in the last two columns of Table 7. The coefficients on the cash flow and investment correlation measures are positive and significant, and remarkably similar to our main findings. The standard errors are higher, reflecting a greater amount of noise in estimating factor loadings.

5.2.4. Economic Significance

To evaluate the economic significance of our findings, we estimate the effect of coinsurance-related reduction in cost of capital on 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

[22] In unreported analyses, we explore the relation between excess debt ratings and 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.

firm value is in general non-linear and depends on other inputs in the valuation formula – expected earnings and earnings growth.

To estimate the 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):

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

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^{iCOC}), where actual value is the sum of the market value of equity at the time of the cost of capital estimation and the book value of debt for the latest fiscal year. This measure of excess value captures the percentage gain or loss in market value resulting from the coinsurance effect on cost of capital.

Using this approach, we find 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 the coinsurance effect of diversification having an economically significant effect on firm value.

6. Conclusion

In this paper, we study the connection between organizational form and cost of capital. We show in a model with systematic deadweight costs that combining business units with imperfectly correlated cash flows can lead to a reduction in systematic risk and hence the combined firm's cost of capital. This coinsurance effect 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 diversified firms have on average a 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. Holding cash flows constant, these findings imply a 6% value gain when moving from the lowest to the highest cash flow correlation quintile.

The core of our findings represents a major challenge to the conventional view that corporate diversification reduces only idiosyncratic risk. In addition, our evidence suggesting that coinsurance affects firms' cost of capital has novel implications for valuation and capital budgeting as ignoring coinsurance effects may yield incorrect (i.e., understated) firm value and NPV estimates, particularly in the context of diversifying mergers and acquisitions. Moreover, because the effects that we find are economically significant, coinsurance is likely to affect optimal financial policies such as hedging and payout policy. The role of coinsurance in relation to these central corporate finance questions represents an exciting and unexplored avenue for future research.

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Appendix

In this Appendix, we develop three extensions of our basic model by (1) relaxing an assumption that the merged firm operates with an efficient internal capital market that is free of agency problems, (2) showing that the coinsurance effect can also apply to debt financing, (3) allowing deadweight losses to vary with the state of the economy.

A1. Agency Problems and Inefficient Transfers as Costs of Integration

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.[23] 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.

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

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

[23] 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)).

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,

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

and add to the coinsurance effect. If $A_g^D < A_b^D$, then integration costs work against coinsurance. Beyond a potential level effect (Prediction 1), however, integration costs do not generate a monotonic relation between a firm's cost of capital and the correlation of its segment cash flows (Prediction 2).

A2. 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, that is, $0<\Delta<\left(H-L\right)/2$ 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

$$\begin{aligned} B_g^D &= \theta_g^2(2(H-\Delta))+2\theta_g(1-\theta_g)(H+L) \\ B_b^D &= \theta_b^2(2(H-\Delta))+2\theta_b(1-\theta_b)(H+L). \end{aligned}$$

Using the expected cash flows above to compute bond betas, we have

$$\begin{aligned} \beta_B^S &= \frac{\theta_g\left(H-\Delta\right)}{\theta_b\left(H-\Delta\right)} \\ \beta_B^D &= \frac{2\theta_g(H-\Delta)+2\theta_g(L+\Delta-\theta_g(L+\Delta))}{2\theta_b(H-\Delta)+2\theta_b(L+\Delta-\theta_b(L+\Delta))}. \end{aligned}$$

$$\underbrace{\qquad\qquad}_{2B_e^S} \qquad \underbrace{\qquad\qquad\qquad\qquad}_{\text{Coinsurance}}$$

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 their cost of debt as well.

A3. State-contingent Deadweight Loss

The basic model assumes 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 β^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 state-contingent deadweight losses would strengthen the coinsurance effect. 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 relationship-specific investments due to the larger wedge between internal and external finance during bad times.

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

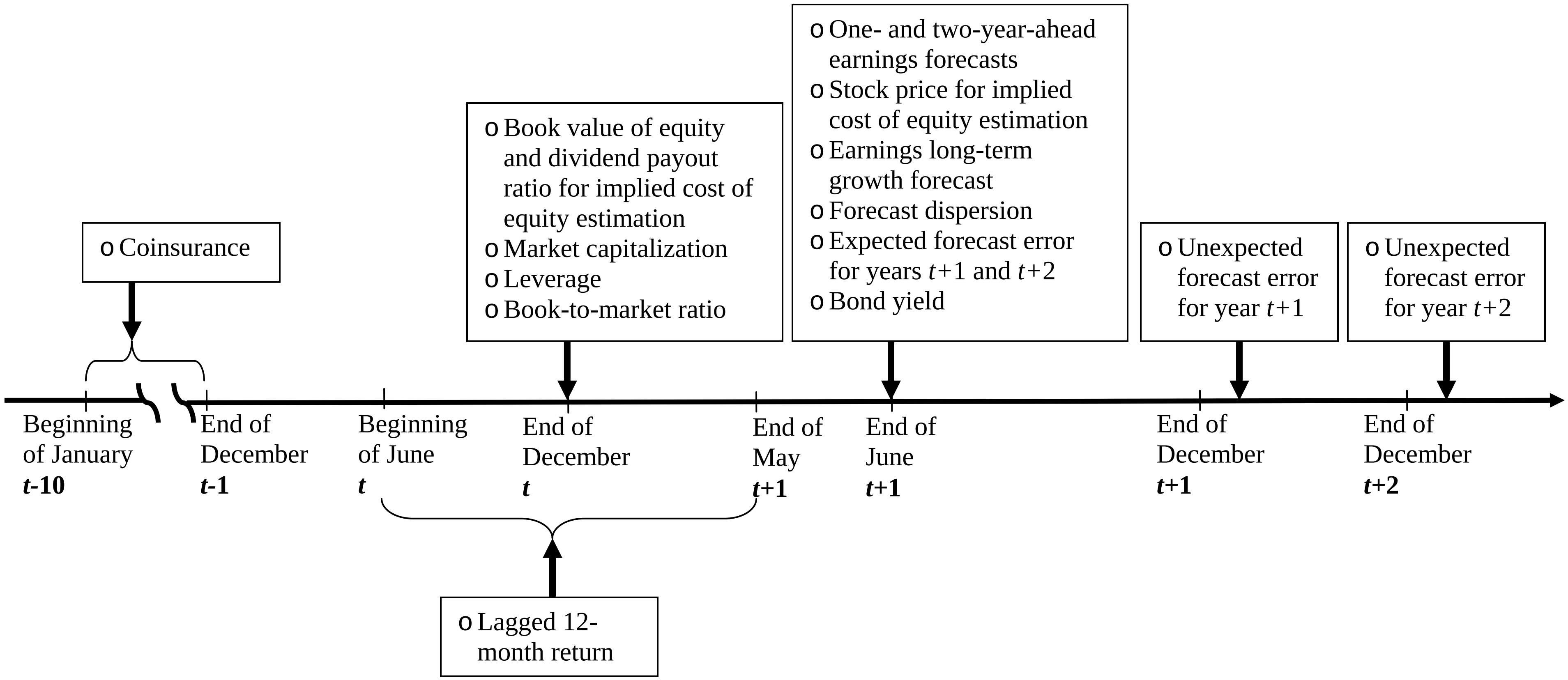

TABLE 1
Summary Statistics: Excess Cost of Capital

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 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.

	Obs.	Mean	Std. Dev.	Lower Quartile	Median	Upper Quartile
Single-Segment (SS)	20,046	-0.030***	0.219	-0.125	-0.001***	0.093
Multi-Segment (MS)	9,107	-0.040***	0.225	-0.150	-0.025***	0.093
MS-SS		-0.010***			-0.024***	

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 for a firm are measured as the portfolio weighted sum of pair-wise segment correlations estimated using historical average industry cash flow and investment based on single-segment firms over a prior ten-year period. *** 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	***			0.034	***
"Single-Segment" - "Q1"			0.029	***			0.027	***

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. Market capitalization is fiscal year-end stock price (#199) multiplied by shares outstanding (#25). Leverage is long-term debt (#9) divided by the sum of long-term debt and market capitalization. Book-to-market is book value of equity (#60) divided by market capitalization. Forecast dispersion is the standard deviation of analysts' one-year-ahead earnings forecasts from I/B/E/S. Long-term growth forecast is the median long-term growth forecast from I/B/E/S. Lagged 12-month return is buy-and-hold return from beginning of June (*t*-1) to end of May (*t*). 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	0.057***				0.055***	
	[0.014]				[0.015]	
Investment correlations		0.056***				0.052***
		[0.015]				[0.014]
Number of segments			-0.005*		0.007**	0.007**
			[0.003]		[0.003]	[0.003]
Logarithm of market capitalization				-0.026***	-0.027***	-0.027***
				[0.005]	[0.005]	[0.005]
Leverage	-0.177***	-0.177***	-0.178***	-0.178***	-0.178***	-0.178***
	[0.027]	[0.027]	[0.027]	[0.026]	[0.027]	[0.027]
Book-to-market	0.192***	0.192***	0.192***	0.141***	0.139***	0.139***
	[0.019]	[0.019]	[0.019]	[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	-0.175*	-0.176*	-0.174*	-0.272***	-0.273***	-0.273***
	[0.105]	[0.104]	[0.102]	[0.103]	[0.100]	[0.099]
Lagged 12-month return	-0.091***	-0.091***	-0.091***	-0.090***	-0.090***	-0.089***
	[0.009]	[0.009]	[0.009]	[0.007]	[0.007]	[0.007]
Constant	-0.092***	-0.091**	-0.031	0.185***	0.130**	0.132**
	[0.035]	[0.036]	[0.026]	[0.058]	[0.057]	[0.060]
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

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 in Table 1. Cash flow and investment correlations are defined in Table 2. The control variables are defined in Table 3. 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	0.043***	
	[0.015]	
Investment correlations		0.041***
		[0.014]
Number of segments	0.012***	0.012***
	[0.003]	[0.003]
Logarithm of market capitalization	-0.028***	-0.028***
	[0.006]	[0.006]
Leverage	-0.234***	-0.234***
	[0.042]	[0.041]
Book-to-market	0.175***	0.175***
	[0.028]	[0.028]
Logarithm of forecast dispersion	0.005	0.005
	[0.004]	[0.004]
Long-term growth forecast	-0.206**	-0.207**
	[0.101]	[0.100]
Lagged 12-month return	-0.081***	-0.081***
	[0.010]	[0.010]
Constant	0.097	0.098
	[0.069]	[0.071]
Observations	9,107	9,107
R-squared	0.134	0.134

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 effects of analyst forecast biases. Panel A reports regressions with expected and unexpected forecast errors added as controls. Panel B reports regressions for sub-samples partitioned on the magnitude of absolute forecast error. 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 Table 3. 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).

Panel A: Full Sample

Cash flow correlations	0.055***	
	[0.017]	
Investment correlations		0.049***
		[0.015]
Number of segments	0.008*	0.008**
	[0.004]	[0.004]
Logarithm of market capitalization	-0.023***	-0.023***
	[0.005]	[0.005]
Leverage	-0.186***	-0.186***
	[0.031]	[0.031]
Book-to-market	0.149***	0.149***
	[0.019]	[0.019]
Logarithm of forecast dispersion	0.005**	0.005**
	[0.002]	[0.002]
Long-term growth forecast	-0.331***	-0.331***
	[0.104]	[0.104]
Lagged 12-month return	-0.059***	-0.059***
	[0.008]	[0.008]
Unexpected analyst forecast error in year +1	-0.211**	-0.210**
	[0.094]	[0.094]
Unexpected analyst forecast error in year +2	-0.479***	-0.479***
	[0.089]	[0.089]
Average predicted analyst forecast error in years +1 and +2	-1.470***	-1.471***
	[0.267]	[0.267]
Constant	0.084*	0.090*
	[0.047]	[0.051]
Observations	23,270	23,270
R-squared	0.189	0.189

Panel B. Partitions Based on Absolute Forecast Error

	Absolute forecast error					
	Low		Medium		High	
Cash flow correlations	0.065***		0.043**		0.035	
	[0.016]		[0.020]		[0.025]	
Investment correlations		0.057***		0.052***		0.018
		[0.016]		[0.016]		[0.024]
Number of segments	0.007**	0.006**	0.005	0.006*	0.008	0.007
	[0.003]	[0.003]	[0.003]	[0.004]	[0.005]	[0.005]
Logarithm of market capitalization	-0.022***	-0.022***	-0.023***	-0.023***	-0.032***	-0.032***
	[0.005]	[0.005]	[0.005]	[0.005]	[0.006]	[0.006]
Leverage	-0.144***	-0.144***	-0.183***	-0.183***	-0.217***	-0.218***
	[0.034]	[0.034]	[0.035]	[0.034]	[0.024]	[0.024]
Book-to-market	0.199***	0.200***	0.166***	0.166***	0.084***	0.085***
	[0.018]	[0.018]	[0.020]	[0.020]	[0.022]	[0.022]
Logarithm of forecast dispersion	0.006*	0.006*	0.008**	0.008**	0.010**	0.010**
	[0.004]	[0.004]	[0.004]	[0.004]	[0.004]	[0.004]
Long-term growth forecast	-0.176***	-0.176***	-0.294***	-0.296***	-0.377***	-0.376***
	[0.063]	[0.063]	[0.094]	[0.093]	[0.138]	[0.137]
Lagged 12-month return	-0.092***	-0.092***	-0.090***	-0.090***	-0.085***	-0.085***
	[0.009]	[0.009]	[0.006]	[0.006]	[0.008]	[0.008]
Constant	0.022	0.029	0.101*	0.092	0.238***	0.255***
	[0.050]	[0.050]	[0.058]	[0.061]	[0.066]	[0.075]
Observations	9,252	9,252	9,267	9,267	9,261	9,261
R-squared	0.182	0.181	0.173	0.174	0.112	0.112

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, 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 Table 3. 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	0.026***		0.051***	
	[0.010]		[0.011]	
Investment correlations		0.029***		0.056***
		[0.009]		[0.012]
Number of segments	0.010***	0.011***	0.002	0.003
	[0.002]	[0.002]	[0.002]	[0.002]
Logarithm of market capitalization	-0.025***	-0.025***	-0.029***	-0.029***
	[0.004]	[0.004]	[0.003]	[0.003]
Leverage	-0.109***	-0.109***	-0.124***	-0.125***
	[0.013]	[0.013]	[0.021]	[0.021]
Book-to-market	-0.083***	-0.083***	0.010	0.009
	[0.010]	[0.010]	[0.006]	[0.006]
Logarithm of forecast dispersion	0.018***	0.018***	0.030***	0.030***
	[0.002]	[0.002]	[0.002]	[0.002]
Long-term growth forecast	0.187***	0.186***	0.517***	0.516***
	[0.041]	[0.041]	[0.046]	[0.046]
Lagged 12-month return	-0.060***	-0.060***	-0.072***	-0.072***
	[0.008]	[0.008]	[0.007]	[0.007]
Constant	0.203***	0.200***	0.153***	0.148***
	[0.033]	[0.036]	[0.035]	[0.037]
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 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. Cost of equity is based on the approach of Gebhardt, Lee, and Swaminathan (2001) (GLS) and Fama and French (1997) (FF). The control variables are defined in Table 3. 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).

	GLS	GLS	FF	FF
Cash flow correlations	0.084***		0.099**	
	[0.016]		[0.042]	
Investment correlations		0.074***		0.085**
		[0.015]		[0.040]
Number of segments	0.011***	0.010***	0.023***	0.022***
	[0.004]	[0.004]	[0.006]	[0.005]
Logarithm of market capitalization	-0.029***	-0.029***	-0.013**	-0.013**
	[0.006]	[0.006]	[0.006]	[0.006]
Leverage	0.016	0.016	0.237***	0.237***
	[0.023]	[0.023]	[0.042]	[0.042]
Book-to-market	0.193***	0.193***	0.056***	0.056***
	[0.020]	[0.020]	[0.017]	[0.018]
Logarithm of forecast dispersion	0.003	0.002	0.017***	0.017***
	[0.003]	[0.003]	[0.006]	[0.006]
Long-term growth forecast	-0.142	-0.141	0.341***	0.342***
	[0.111]	[0.110]	[0.079]	[0.079]
Lagged 12-month return	-0.097***	-0.097***	0.001	0.001
	[0.007]	[0.007]	[0.039]	[0.039]
Constant	-0.001	0.008	-0.124*	-0.111*
	[0.063]	[0.068]	[0.072]	[0.063]
Observations	29,150	29,150	26,364	26,364
R-squared	0.158	0.158	0.014	0.014