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Testing International Asset Pricing Models Using Implied Costs of Capital

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Abstract

This paper tests international asset pricing models using firm-level expected returns estimated from an implied cost of capital approach. We show that the implied approach provides clear evidence of economic relations that would otherwise be obscured by the noise in realized returns. Among G-7 countries, expected returns based on implied costs of capital have less than one-tenth the volatility of those based on realized returns. Our tests show that firm-level expected returns increase with world market beta, idiosyncratic volatility, financial leverage, and book-to-market ratios, and decrease with currency beta and firm size.

I. Introduction

Empirical asset pricing is plagued by a major problem: the fact that realized returns are extremely noisy proxies of expected returns.[1] This problem is exacerbated in international settings, where data availability often limits the time period of examination. As a result, financial researchers testing international asset pricing models face the risk that, during the time period of investigation, economically significant relations can be rendered statistically insignificant by the noise in realized returns.

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[1] See Elton (1999), Gebhardt, Lee, and Swaminathan (2001), and Pastor, Sinha, and Swaminathan (2008) for a more detailed discussion of the problem.

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The international asset pricing literature has traditionally dealt with this problem by focusing on the right-hand side of the asset pricing equation (e.g., by introducing conditional asset pricing models or more sophisticated econometric specifications of factor risk premiums and factor loadings).[2] In this paper, we reexamine the left-hand side of the asset pricing equation and use an approach based on the implied cost of capital to estimate expected returns for the firms in G-7 countries (Canada, France, Germany, Italy, Japan, the U.K., and the U.S.).[3]

The implied cost of capital approach estimates firm-level expected return as the internal rate of return (IRR) that equates the current stock price to the present value of forecast free cash flows to equity holders. We use these expected return measures to test international asset pricing models and to contrast the results with those based on realized returns. The key objectives of our paper are i) to introduce the implied cost of capital approach to international asset pricing, and ii) to examine whether the implied approach leads to sharper inferences, particularly in economic relations that may be difficult to detect with realized returns.

To our knowledge, this is the first study to conduct formal international asset pricing tests using implied expected return estimates.[4] We show that expected returns based on implied costs of capital are much less noisy than those based on realized returns. In our sample, expected returns based on the implied approach have less than one-tenth of the volatility of the estimates computed using realized returns (see Table 1). Our main analyses exploit this feature of the implied expected return measures to test a number of international asset pricing models.

Opinions vary widely as to the "right" model for international asset pricing. One central issue is whether risks are priced locally or globally—i.e., whether local market returns should be included as a risk factor (see Karolyi and Stulz (2003) for a summary). Another point of contention is the existence of a currency risk factor. While many theoretical models (e.g., Solnik (1974), Sercu (1980), Stulz (1981), and Adler and Dumas (1983)) suggest that a currency factor should be included, empirical support for this factor is mixed.[5]

In this study, we estimate factor loadings for a three-factor asset pricing model consisting of a world market factor, a country-specific local market factor,

[2]For example, Merton (1980), Harvey (1991), Chan, Karolyi, and Stulz (1992), Dumas and Solnik (1995), De Santis and Gerard (1998), Fama and French (1998), Ferson and Harvey (1993), Griffin (2002), and Hou, Karolyi, and Kho (2006). See the next section and Karolyi and Stulz (2003) for a more detailed discussion.

[3]Gebhardt et al. (2001) use a similar approach to examine asset pricing models using U.S. data. Pastor et al. (2008) use this approach to examine the intertemporal asset pricing relationship between market-wide conditional expected returns and conditional volatility. Other papers that have used this approach include Claus and Thomas (2001), Fama and French (2002), Gode and Mohanram (2003), Easton (2004), Hail and Leuz (2006), Botosan and Plumlee (2005), and Ohlson and Juettner-Nauroth (2005). We discuss these studies in detail later in this paper.

[4]Hail and Leuz (2006) use implied cost of capital in an international setting, but they emphasize country-level factors. Specifically, they explore legal institutions and regulations that may affect country-level cost of capital, and they do not address international asset pricing issues.

[5]Some empirical studies (e.g., Dumas and Solnik (1995), De Santis and Gerard (1998)) find that currency risk is priced, but others do not (e.g., Jorion (1991)).

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and a currency factor.[6] In addition, we include several firm characteristics that have been shown in prior literature to be significantly related to the cross section of realized returns (i.e., idiosyncratic volatility, size, book-to-market (B/M) ratio, and financial leverage).[7] For completeness, we also evaluate a model that features Fama-French three-factor betas (see Table 8).

We find that implied expected returns are significantly related to a number of these factor loadings and characteristics. First, we show that world market betas and currency betas are reliably related to firm-level expected returns (local betas do not exhibit the same statistical reliability—which may not be surprising, given our sample of G-7 firms). Second, we document a clear positive relation between idiosyncratic volatility and expected returns (contrary to the mixed results in prior studies).[8] Third, we find a reliably positive relation between leverage and expected returns (again, contrary to the weak results in prior studies). Finally, we show that both size and B/M ratio are strongly related to expected returns among G-7 firms (i.e., smaller and higher B/M firms are associated with higher expected returns).

These findings are robust to a variety of modifications to the primary research approach used in this paper. We show that these relations hold in both univariate sorts and multivariate regressions. They are insensitive to adjustments for country-level analyst forecast biases, alternative exchange rate forecasts, as well as for a range of reasonable perturbations in growth rate assumptions, fade rates, and forecast horizons. They are also robust across four different measures of implied costs of capital.[9] An empirical asset pricing model that includes three betas (world, local, and currency) and four characteristics (volatility, leverage, B/M ratio, and size) consistently explains about 20% of the cross-sectional variation in firm-level risk premiums across G-7 countries.

For comparison purposes, we show that results from identical tests are much weaker when realized returns are used as proxies for expected returns. We find that both realized and implied risk premiums are positively related to world beta, local beta, idiosyncratic volatility, leverage, and B/M ratio; and that both are negatively related to currency beta and size. However, in tests using realized returns, only the currency beta is uniformly significant. In contrast, results based on implied expected returns are significant for all seven variables.

In sum, the main conclusion of our paper is that the implied cost of capital approach can provide important new insights into the cross-sectional determinants

[6]We follow Chan et al. (1992) and Bekaert and Harvey (1995) in including a local market factor. See also Errunza and Losq (1985). Additionally, we include a currency risk factor as suggested in Solnik (1974), Sercu (1980), and Adler and Dumas (1983).

[7]Stocks with high B/M ratios are found to earn higher realized returns (Fama and French (1998), Griffin (2002)); small stocks outperform large stocks (Banz (1981), Heston, Rouwenhorst, and Wessels (1995)); highly leveraged firms tend to earn higher returns (Bhandari (1988)), although Fama and French (1992) find that leverage is dominated by B/M and size; finally, Ang, Hodrick, Xing, and Zhang (2006), (2009), Spiegel and Wang (2006), and Malkiel and Xu (2002) examine the relation between volatility and realized returns, with mixed findings.

[8]Empirical evidence with realized returns is mixed. For example, Spiegel and Wang (2006) find that at the monthly horizon, stocks with higher idiosyncratic risk have lower liquidity and higher returns, while Ang et al. (2006), (2009) find that stocks with higher idiosyncratic volatility have lower subsequent returns.

[9]We conduct our tests using valuation models featured in four prior studies: Gebhardt et al. (2001), Claus and Thomas (2001), Gode and Mohanram (2003), and Easton (2004).

of firm-level expected returns in the international context. We recognize that our study is limited in scope, particularly with respect to country membership. However, we view these findings as a significant first step toward a broader investigation of the role of implied risk premiums (IRPs) across international borders.

The rest of the paper proceeds as follows: Section II reviews the international asset pricing literature and the existing evidence. Section III outlines the methodology for computing cost of capital. Section IV describes the data and the firm characteristics and factor model used to estimate betas. Section V discusses the summary statistics. Section VI discusses the cross-sectional findings, and Section VII concludes.

II. International Asset Pricing Models

Asset pricing models call for the use of expected returns. The use of realized returns rests on the assumption that, over a long period, realized returns equal expected returns. However, in international markets, the time period of examination is substantially shorter than in the U.S. Thus, researchers run the risk that during the period under consideration, tests based on realized returns may obscure the true relation between time-varying risk and expected returns.

To circumvent these problems with realized returns, the international asset pricing literature has followed the domestic asset pricing literature in using conditional asset pricing models (e.g., Harvey (1991), Chan et al. (1992), and Ferson and Harvey (1993)). Many of these tests follow Harvey (1991) in modeling time-varying expected returns as a function of instrumental variables. These conditional pricing models have been a source of much debate and controversy in the literature. A major problem with this approach is that forecasts of expected returns are only as good as the predictive power of the instruments used. For instance, Bossaerts and Hillion (1999) fail to find out-of-sample predictability in the commonly used instrumental variables in 14 countries.[10]

Other main points of contention in international asset pricing revolve around i) whether the risks are priced locally or globally, and ii) whether currency risk is priced. Using unconditional asset pricing models, Griffin (2002) finds that risks are local, while papers using conditional asset pricing models (see Harvey (1991), Ferson and Harvey (1993)) largely support global models. Additionally, while unconditional models (Jorion (1991)) find little evidence for priced currency risk, conditional models (Dumas and Solnik (1995), De Santis and Gerard (1998)) find the opposite. Thus, just as in the domestic asset pricing literature, the conclusions reached in the international asset pricing literature appear specific to the asset pricing methodology adopted.

As discussed in the introduction, the purpose of this paper is not to critique the various asset pricing methodologies but rather to present an alternate approach to estimating expected returns. Our focus is on the left-hand side of the asset pricing equation as opposed to the right-hand side, and we hope to present new

[10]Goyal and Welch (2003) fail to find evidence of predictability in the U.S., but their findings are contested by Ang and Bekaert (2007).

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evidence on some of the key issues discussed above. In Section III, we describe the methodology for computing the implied cost of capital.

III. The Methodology for Computing Implied Cost of Capital

We compute our main measure of cost of equity capital for each firm as the IRR that equates the present value of future free cash flows to equity (FCFE) to current stock price where the numeraire currency is the U.S. dollar. The free cash flow model, in infinite-horizon form, can be written as

$$P_t = \sum_{k=1}^{\infty} \frac{E_t\left(\mathrm{FCFE}_{t+k}\right)}{\left(1+r_e\right)^k}, \tag{1}$$

where P_t is the current stock price in U.S. dollars, $E_t(\mathrm{FCFE}_{t+k})$ is the expected future free cash flow to equity for period $t+k$ conditional on information available at time t in U.S. dollars, and r_e is the cost of equity capital, also denominated in U.S. dollars, based on the information set at time t. All units are expressed in nominal terms. This definition assumes a flat term structure of discount rates.

Clean surplus accounting requires that all gains and losses affecting book value are also included in earnings; that is, the change in book value from period to period is equal to earnings minus free cash flows to equity or net dividends ($b_t = b_{t-1} + \mathrm{NI}_t - \mathrm{FCFE}_t$). The model in equation (1) can be written in terms of economic profits or residual income earned by a firm, if the clean surplus relationship holds:

$$P_t = B_t + \sum_{k=1}^{\infty} \frac{E_t\left[\mathrm{NI}_{t+k} - r_e B_{t+k-1}\right]}{\left(1+r_e\right)^k}, \tag{2}$$

where B_t is the book value of equity at time t, NI_{t+k} equals net income for period $t+k$, and the term inside the expectation operator is the economic profit/residual income for period $t+k$. Equation (2) is commonly referred to as the residual income model.

Gebhardt et al. (2001) compute the implied cost of equity using the residual income model. They forecast earnings by assuming that the return on equity (ROE) of each firm reverts to an industry median ROE by the terminal period. This paper adopts a modified version of their approach, in which individual firms' earnings growth rates are assumed to revert to the long-run nominal world gross domestic product (GDP) growth rate. This alternate approach imposes less stringent data requirements and hence is easier to implement in the international context. We describe this approach in detail below.

A. Empirical Implementation of the Free Cash Flow Model

Equation (2) expresses stock price in terms of an infinite series, but in actual implementation, free cash flows are only explicitly forecast for a finite horizon. The present value of free cash flows beyond the last explicit forecast period is captured in the "terminal value" estimation. In other words, free cash flow models estimate firm value in two parts: i) the present value of free cash flows up to

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a terminal period $t + T$, and ii) a continuing value that captures free cash flows beyond the terminal period. We derive future free cash flows up to year $t + T$ as the product of annual earnings forecasts and one minus the plowback rate:

$$E_t\left(\mathrm{FCFE}_{t+k}\right) = \mathrm{FE}_{t+k} \times (1 - b_{t+k}), \tag{3}$$

where FE_{t+k} and b_{t+k} are the earnings forecasts in U.S. dollars and the plowback rate forecasts for year $t + k$, respectively. To obtain equation (3), define net new equity investment as capital expenditures plus change in working capital minus depreciation and amortization minus net new issues of debt and preferred stock. Thus, FCFE is net income minus net new equity investment. Defining the plowback rate as net new equity investment divided by net income, we obtain equation (3).

We forecast earnings up to year $t+T$ in three stages: i) we explicitly forecast earnings (in dollars) for years $t+1$ and $t+2$; ii) we then use the growth rate implicit in the forecasts in years $t+1$ and $t+2$ to forecast earnings in year $t+3$; and, finally, iii) we forecast earnings from year $t + 4$ to year $t + T + 1$ implicitly by assuming that the year $t + 3$ earnings growth rate (g_3) reverts to steady-state values by year $t + T + 2$.

We use the world average real GDP growth rate (g) over the last 10 years plus the average U.S. inflation rate over the same period as the nominal steady-state growth rate starting in year $t+T+2$.[11] We further impose an exponential rate of decline by allowing the year $t+3$ growth to revert to the steady-state growth.[12] Specifically, earnings growth rates and earnings forecasts using the exponential decline are computed as follows for years $t + 4$ to $t + T + 1$ $(k = 4, \ldots, T + 1)$:

$$\begin{aligned} g_{t+k} &= g_{t+k-1} \times \exp\left[\log\left(g/g_3\right)/(T-1)\right] \quad \text{and} \\ \mathrm{FE}_{t+k} &= \mathrm{FE}_{t+k-1} \times (1 + g_{t+k}). \end{aligned} \tag{4}$$

We forecast plowback rates using a two-stage approach: i) we explicitly forecast plowback rates for years $t + 1$ and $t + 2$ (see Section III.D), and ii) we assume that the plowback rate in year $t + 2$, b_2 reverts linearly to a steady-state value by year $t + T + 1$ computed from the sustainable growth rate formula.[13] This formula assumes that, in the steady state, the product of the return on new investments and the plowback rate $(\mathrm{ROI} * b)$ is equal to the growth rate in earnings (g).[14] We further impose the condition that, in the steady state, ROI equals r_e for new investments, because competition will drive returns on these investments down to the cost of equity. Thus, there are two major assumptions underlying our

[11] We use the average real GDP growth rates across the G-7 countries as our proxies for world real GDP growth. By reverting to world average GDP growth rates in U.S. dollar terms, we implicitly assume that, in the long run, purchasing power parity (PPP) holds.

[12] The exponential rate of decline reverses high growth rates faster than a simple linear rate of decline. We choose this rate of decline because prior studies such as Nissim and Penman (2001) and Chan, Karceski, and Lakonishok (2003) show that high earnings growth rates decay rapidly.

[13] We assume that year $t + k$ plowback affects year $t + k + 1$ earnings growth. We assume a linear decline in plowback rate and an exponential decline in growth rate because, empirically, growth rates revert much faster than plowback rates and ROI (e.g., see Nissim and Penman (2001)).

[14] See Brealey and Myers (2000).

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model: i) the growth rate is assumed to revert to the long-run GDP growth rate, and ii) the ROI is assumed to revert to the cost of equity (r_e).

Substituting ROI with cost of equity (r_e) in the sustainable growth rate formula and solving for plowback rate (b) provides the steady-state value for the plowback rate, which equals the steady-state growth rate divided by cost of equity (g/r_e). The intermediate plowback rates from $t+3$ to $t+T$ ($k=3,\ldots,T$) are computed as follows:

$$b_{t+k} = b_{t+k-1} - \frac{b_2 - b}{T-1}. \tag{5}$$

The terminal value (TV) is computed as the present value of a perpetuity equal to the ratio of the year $t+T+1$ earnings forecast divided by the cost of equity:

$$\mathrm{TV}_{t+T+1} = \frac{\mathrm{FE}_{t+T+1}}{r_e}, \tag{6}$$

where FE_{t+T+1} is the earnings forecast for year $t+T+1$. Note that the use of the no-growth perpetuity formula does not imply that earnings or cash flows do not grow after period $t+T$. Rather, it simply means that any new investments after year $t+T$ earn zero economic profits. In other words, any growth in earnings or cash flows after year T is value-irrelevant. It is easy to show that the Gordon growth model for TV will simplify to equation (6) when ROI equals r_e.

Substituting equations (3) to (6) into the infinite-horizon free cash flow valuation model in equation (2) provides the following empirically tractable finite-horizon model:

$$P_t = \sum_{k=1}^{T} \frac{\mathrm{FE}_{t+k} \times (1 - b_{t+k})}{(1+r_e)^k} + \frac{\mathrm{FE}_{t+T+1}}{r_e\,(1+r_e)^T}. \tag{7}$$

In this paper, we implement the model in equation (7) using a 15-year horizon ($T=15$) and compute r_e as the rate of return that equates the present value of free cash flows to the current stock price in U.S. dollars.[15] We subtract the yield on a 10-year U.S. government bond from r_e to compute the IRPs. Our empirical tests are based on the IRPs.

B. Explicit Earnings Forecasts over the First Two Years

We obtain explicit earnings forecasts in the respective currencies for years $t+1$ and $t+2$ for each firm from the Institutional Brokers' Estimate System (IBES) domestic and international databases. IBES analysts supply a one-year-ahead (FY1) and a two-year-ahead (FY2) earnings per share (EPS) forecast in the

[15] For simplicity, we assume the same number of years on the forecast horizon before the terminal period for all firms. In Section VI, we conduct robustness checks which show that the forecast horizon assumption has little effect on our results. It is possible that results can be further improved by allowing the forecast horizon to vary for firms in different industries or growth stages. This is left for future research.

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local currency for each firm in the IBES database. These forecasts are aggregated across analysts monthly, and a consensus (or mean) forecast is computed. We use the consensus one- and two-year-ahead EPS forecasts. We translate these EPS forecasts into dollar terms using one- and two-year-ahead forecast exchange rates and denote these dollar forecasts as FE_{t+1} and FE_{t+2}.

$$\begin{aligned} FE_{t+1} &= FY1 \times FS_{t+1} \quad \text{and} \\ FE_{t+2} &= FY2 \times FS_{t+2}, \end{aligned} \tag{8}$$

where FY1 and FY2 are local currency consensus EPS forecasts, and FS_{t+1} and FS_{t+2} are one- and two-year-ahead exchange rate forecasts expressed in dollars per foreign currency. In using dollar earnings forecasts, we are essentially invoking long-run PPP. Exchange rate forecasts are described in the next subsection.

We then estimate the growth rate for year 3 (g_3) from the consensus forecasts in years 1 and 2 by assuming that $g_3 = FE_{t+2}/FE_{t+1} - 1$, and we use this growth rate to compute a three-year-ahead earnings forecast, as given by $FE_{t+2}(1 + g_3)$. Firms with growth rates above 100% (below zero) are given values of 100% (zero).[16] These earnings forecasts, combined with the plowback rates, allow us to generate explicit forecasts of FCFE.

C. Exchange Rate Forecasts

To compute implied cost of capital in U.S. dollars for firms in different countries, we use exchange rate forecasts to convert local currency earning forecasts into U.S. dollars. The exchange rate forecasts are obtained from the Economist Intelligence Unit (EIU). All of our reported results are based on risk premiums computed in U.S. dollars using the EIU forecasts.[17]

D. Plowback Rates

We compute the plowback rate for each firm as one minus that firm's dividend payout ratio. We estimate the dividend payout ratio by dividing actual dividends from the most recent fiscal year by earnings over the same time period. We exclude share repurchases due to the practical problems associated with determining the likelihood of their recurrence in future periods and the limited availability of this data for international firms. For firms with negative earnings, we estimate the normalized earnings as the product of the firm's total assets and the median return on total assets ratio (ROA) of the industry-size portfolio (based on three market cap portfolios in each industry group) to which the firm belongs.[18] The

[16] We have relaxed this assumption to allow growth rates above 100% and repeated our cross-section tests. All of the major results remain unchanged.

[17] We also conduct our analysis using our own exchange rate forecasts based on a transfer function (i.e., a multivariate autoregressive moving average (ARMA) model of past exchange rate changes with uncovered interest parity variables). As a further robustness check, we also redo all of the analysis based on risk premiums computed completely in local currency terms, thus eliminating the need to forecast exchange rates. All the main results remain the same.

[18] As a robustness check, we also use 6% of total assets as a proxy for normal earnings levels when current earnings are negative (6% is the average long-run ROA in the U.S.). None of our results change significantly.

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industry-size portfolio is constructed by sorting firms into three size portfolios within each industry. Payout ratios of less than zero (greater than one) are assigned a value of zero (one).

E. Measurement Error

One source of measurement error in calculating the implied cost of capital is associated with slow updating of analyst forecasts in the face of big price moves. As Guay, Kothari, and Shu (2005) show, when there is a large upward movement in stock price prior to estimating the cost of capital, the change in stock price reflects the market's revision of future firm earnings. Because analysts do not update their earning forecasts instantaneously, the valuation model will be based on downwardly biased cash flow forecasts, yielding an artificially low cost of capital. Similarly, if there is a large drop in stock price, the cost of capital will be artificially high. Guay et al. (2005) recommend a simple correction for this effect, which we incorporate in all of our implied cost of capital measures.[19]

F. Three Alternate Implied Costs of Capital

While the main results in our paper use the implied cost of capital that we feature, we check the robustness of our results using three other implied costs of capital: Ohlson and Juettner-Nauroth (2005), the Modified-PEG model of Easton (2004), and Claus and Thomas (2001). Our computation of these three implied costs of capital (r_{MPEG}, r_{OJ}, and r_{CT}) follows closely the implementation in Hail and Leuz (2006). The alternate implied risk premiums (AIRPs) for these three models are denoted by $\text{AIRP}_{\text{MPEG}}$, AIRP_{OJ}, and AIRP_{CT} and are computed as the respective implied costs of capital minus the U.S. 10-year interest rate. The term r_{MPEG} is based on the modified PEG ratio model by Easton (2004),

$$P_t = \left(\frac{\text{FE}_{t+2} + r_{\text{MPEG}} d_{t+1} - \text{FE}_{t+1}}{r_{\text{MPEG}}^2} \right). \tag{9}$$

The term r_{OJ} is based on Ohlson and Juettner-Nauroth (2005) and Gode and Mohanram (2003),

$$P_t = \left(\frac{\text{FE}_{t+1}}{r_{\text{OJ}}} \right) \cdot \left(\frac{g_{\text{ST}} + r_{\text{OJ}} \frac{d_{t+1}}{\text{FE}_{t+1}} - g_{\text{LT}}}{r_{\text{OJ}} - g_{\text{LT}}} \right), \tag{10}$$

where d_{t+1} is the dividend that is set as a constant fraction of forecast earnings following the current payout ratio of the firm, the short-term growth rate g_{ST} is estimated as the IBES forecast percentage change in earnings, and the long-term

[19] To correct for the sluggish earnings revisions, Guay et al. (2005) explicitly adjust analyst earnings forecasts by the expected forecast error. We apply this technique in our paper. Specifically, we sort firms into 12 portfolios based on stock returns each year. In each portfolio, we calculate the time-series median of portfolio median forecast errors and subtract these portfolio median forecast errors from each firm's FE1 and FE2 to adjust for near-term analyst sluggishness.

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earnings growth rate g_{LT} is the historical mean GDP growth rate in the G-7 countries plus the historical mean U.S. inflation rate.[20] Finally, r_{CT} is based on Claus and Thomas (2001):

$$P_t = B_t + \sum_{k=1}^{T} \frac{(\mathrm{FE}_{t+k} - r_{CT}B_{t+k-1})}{(1+r_{CT})^k} + \frac{(\mathrm{FE}_{t+k} - r_{CT}B_{t+k-1})(1+g)}{(r_{CT}-g)(1+r_{CT})^T}, \tag{11}$$

where actual book value B_t and IBES forecast earnings up to three years ahead are used to derive the expected future residual income sequence.[21] Future book value is derived through clean surplus accounting. Each residual income in the future equals the forecast earnings minus the book value times the implied cost of capital.

IV. Data Description

Our sample of international firms is derived from the Worldscope database. To ensure that we have a reasonable number of firm-level observations in each country, we focus our analysis on the 1990 to 2000 time period, which has the widest coverage in the Worldscope database. This database requires each firm's home country (both country of origin and country of domicile) to be clearly identified. Due to data limitations, we confine our analysis to firms in the G-7 countries (Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States). Data for U.S. firms are obtained from Compustat. Together, there are 118,244 firm-year observations.

We obtain international stock returns data from Datastream and the U.S. returns data from the Center for Research in Security Prices (CRSP). We require our firms to have enough return observations available from Datastream and CRSP that we can estimate their betas and compare their one-year-ahead realized returns. The data sample includes 83,678 firm-year observations.

We obtain total market capitalization for each firm based on the closing market price as of June 30 of each year. We require the availability of the following data items, measured as of the most recent fiscal year ending at least six months prior to June 30: common dividend, net income, book value of common equity, fiscal year-end date, and currency denomination. In addition, each firm needs to have a one- and two-year-ahead consensus earnings forecast (in local currencies) and one- and two-year-ahead actual earnings in the IBES International database in June of each year. The stock price P_t as of June 30 is also obtained from IBES. We make adjustments for measurement errors in the forecast earnings based on the methodology of Guay et al. (2005). After these screens, there are 30,486 firm-year observations.

[20] As in Hail and Leuz (2006), this model requires a positive change in earnings forecasts in order to obtain a solution.

[21] Beyond the third year, the nominal residual income is assumed to grow at the historical mean GDP growth rate in the G-7 countries plus the historical mean U.S. inflation rate. As in our main implied cost of capital measure, we estimate the growth rate for year 3 (g_3) from the consensus forecasts in years 1 and 2: $g_3 = \mathrm{FE}_{t+2}/\mathrm{FE}_{t+1} - 1$. We use this growth rate to compute a three-year-ahead earnings forecast: $\mathrm{FE}_{t+3} = \mathrm{FE}_{t+2}(1+g_3)$.

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Only firms from the largest exchange in each country are included, except in Japan (where stocks from both the Osaka and Tokyo exchanges are included) and in the United States (where NYSE, AMEX, and NASDAQ stocks are all included).[22] To reduce errors in Datastream, we follow Hou et al. (2006) and Ince and Porter (2006) in applying a similar screening procedure for international stock returns.[23] Firms with negative common equity and negative one- and two-year-ahead earnings forecasts are excluded. After these screens, there are 28,124 firm-year observations.

We obtain country-level data from the International Country Risk Guide (ICRG) and idiosyncratic volatility data from Ang et al. (2006), (2009). The data include inflation and real GDP growth. The 10-year interest rates we use are the yields to maturity on 10-year U.S. Treasury bonds as of the end of June, obtained from Datastream. The inflation rates are calculated from the Consumer Price Index (CPI) series of different countries from the Conference Board. To eliminate the influence of outliers, we rank firms in each country annually by their estimated risk premiums and exclude observations in the top and bottom 1%. We also exclude observations at the top and bottom 1% of B/M ratio and leverage.

After these filters, there are 24,913 firm-year observations across the G-7 countries. The number of firms per year is listed in Table 1. For each firm, we estimate the cost of equity, r_e, as the IRR from equation (2) at the end of June for each year from 1990 to 2000. We then subtract the end-of-June yield on the long-term (10-year) U.S. government bonds from the IRR to obtain an (annualized) IRP for each firm in the sample.

Next, we describe our systematic risk measures (betas) as well as various firm-specific characteristics used in our analysis. As discussed earlier, we take a partial segmentation approach and adopt the following three-factor model as our primary specification for measuring systematic risk:[24]

$$r_t - r_{ft} = a + b_W(r_{Wt} - r_{ft}) + b_L\,(r_{Lt} - r_{ft}) + b_E\, e_t + u_t, \tag{12}$$

where b_W is the world market beta, b_L is the local market beta, and b_E is the currency beta; r_{Wt} represents the monthly dollar return on the world market portfolio, r_{Lt} represents the monthly dollar return on the local market portfolio; e_t is the monthly trade-weighted exchange rate return index; and r_{ft} is the monthly one-month U.S. Treasury bill return. Note that all returns are measured in U.S. dollars. The exchange risk is approximated by a single currency risk factor using the approach of Ferson and Harvey (1993).[25] Specifically, our currency risk factor

[22] Because they are not domestic firms, depositary receipts are excluded. There are two basic means by which we identify depositary receipts. First, Worldscope marks some firms with a depositary receipt indicator. Second, the names of some firms are clearly labeled as depositary receipts.

[23] We require a minimum price of $5 in the previous month for a valid monthly return. Any return above 300% that is reversed within one month is set to missing. That is, if R_t or R_{t-1} is greater than 300% and $(1 + R_t)\,(1 + R_{t-1}) - 1 < 50\%$, then both R_t and R_{t-1} are considered missing. In order to exclude other outliers, we treat as missing the monthly returns that fall outside of the 0.5% and 99.5% percentile ranges in each country.

[24] We have also estimated a two-factor model excluding the local factors, and our cross-sectional findings are similar.

[25] In a model featuring currency risk, the number of exchange risk factors can be as high as the number of currencies other than the numeraire currency because, theoretically, investors bear exchange

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TABLE 1

Risk Premiums by Country

Table 1 reports the equal-weighted means of the implied risk premiums (IRPs) (in percent per annum) and realized risk premiums (RRPs) in each country in the G-7 as of June-end every year from 1990 to 2000. The IRPs are computed as the differences between implied costs of capital and the yields on the U.S. 10-year government bond. The implied costs of capital are computed based on the 15-year discounted cash flow model, described in the main text, that uses analyst forecasts of earnings and exchange rates and the mean reversion to long-run average GDP growth rates across G-7 countries. RRPs are defined as the realized returns for firms in each country in the subsequent year, minus the realized returns on the U.S. 10-year government bond. The time-series averages and standard deviations across the years are reported. The number of firms *N* is reported for each country-year. SD ratio refers to the ratio of the time-series standard deviation of RRPs to the time-series standard deviation of IRPs. We calculate the SD ratio for each firm and report (in bold) the average of the SD ratios across all firms in each country.

		1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	Avg	Std Dev	SD Ratio
Canada														
IRP	Mean	10.2	11.5	12.3	8.9	8.6	8.1	6.2	8.1	10.7	7.1	9.2	1.9	**12.13**
RRP	Mean	−11.8	2.5	−8.3	11.9	2.7	48.2	17.2	−2.2	18.7	5.3	8.4	17.2	
	N	26	23	29	31	33	40	45	46	60	60			
France														
IRP	Mean	4.4	5.8	6.4	7.4	7.0	6.8	5.6	5.4	8.2	7.4	6.4	1.1	**16.20**
RRP	Mean	31.1	−24.4	26.8	5.9	11.1	0.7	31.7	−20.7	23.3	−18.8	6.7	21.9	
	N	26	52	54	59	81	86	95	113	129	157			
Germany														
IRP	Mean	1.0	0.5	2.9	3.2	4.3	4.4	3.8	5.0	6.7	8.7	4.1	2.5	**13.70**
RRP	Mean	6.8	−28.7	19.7	1.8	−8.9	17.3	6.1	−26.6	−5.8	−38.4	−5.7	19.9	
	N	19	40	50	69	73	75	86	84	94	131			
Italy														
IRP	Mean	2.5	8.5	6.2	8.2	10.1	6.3	5.9	5.4	7.2	7.0	6.7	2.1	**14.53**
RRP	Mean	−11.5	−9.1	24.1	−4.8	−8.3	−3.4	67.3	−10.8	12.5	−26.6	2.9	26.5	
	N	11	12	20	21	25	31	32	40	38	48			
Japan														
IRP	Mean	−1.4	1.7	2.3	2.8	3.7	1.3	1.7	3.7	5.5	5.2	2.6	2.0	**17.30**
RRP	Mean	−29.6	32.4	17.7	−21.8	10.4	−17.5	−41.4	42.3	26.8	−26.8	−0.8	29.9	
	N	115	80	98	110	215	298	477	274	378	500			
U.K.														
IRP	Mean	5.8	3.5	5.7	6.2	5.5	3.9	3.5	4.8	7.4	7.3	5.4	1.4	**12.61**
RRP	Mean	34.7	0.6	15.2	7.3	16.7	8.2	24.4	−11.3	11.3	−0.3	10.7	13.1	
	N	112	135	171	188	199	206	200	227	215	195			
U.S.														
IRP	Mean	6.2	7.0	7.4	6.2	7.0	6.1	5.3	5.8	6.5	5.6	6.3	0.7	**18.33**
RRP	Mean	8.1	19.2	−0.8	17.7	23.4	17.4	15.7	−0.9	13.0	18.5	13.1	8.4	
	N	1,262	1,438	1,636	1,837	2,047	2,061	2,105	1,832	1,718	1,474			

is the return on a trade-weighted exchange rate index against major trading partners of the U.S., as provided by the Federal Reserve Board.[26] The exchange rate index represents the dollar value of the basket of currencies. An increase in the value of the index represents a depreciation of the value of the dollar.

To make the results easier to interpret, we orthogonalize local market returns relative to the world market returns, following Jorion and Schwartz (1986). Using ordinary least squares regression, the local market returns are regressed upon the world market returns. The error term of the regression is the new local market return. This regression is conducted every two to five years to allow for time-varying factor loadings.

rate risks from all currencies. This, of course, is difficult to implement empirically, as the number of factors could easily exceed the number of time-series observations used in the estimation of the factor model.

[26]We use a nominal exchange rate in most specifications, but we test the robustness of our findings using real exchange rates as well. The nominal exchange rate index we adopt is a trade-weighted exchange index vis-à-vis currencies of major trading partners including countries using the Euro and Canada, Japan, the United Kingdom, Switzerland, Australia, and Sweden. The index is obtained from the Federal Reserve Board's Web site (http://research.stlouisfed.org/fred2/series/TWEXBMTH/15).

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For each firm, in June of each year, we estimate the factor model in equation (12) using at least 24 months of data over the prior 60 months. The three estimated betas are then used in cross-sectional regression tests, which are discussed later in this paper. For U.S. firms, individual stock returns are obtained from CRSP. For firms in other countries, returns data are obtained from Datastream. Returns on value-weighted local and world stock market indices are from Morgan Stanley Capital International (MSCI).

In addition to the betas, we include size, B/M ratio, leverage, and idiosyncratic volatility as explanatory variables in some of our regression specifications. Based on average realized returns, Fama and French (1992) suggest that smaller and higher B/M ratio firms might be riskier.[27] To reduce the impact of outliers, we use the natural log of market capitalization (LnSZE) and the natural log of B/M ratios (LnBM) in our empirical analysis.

We also investigate whether financial leverage (LEV) affects the cost of equity. Modigliani and Miller (1958) predict that a firm's cost of equity should be increasing in its financial leverage. Since we already include the B/M ratio in the regression, we use a measure of book leverage in our cross-sectional regression tests. The book leverage ratio is defined as the log of the ratio of total debt to total book value of equity from the most recent fiscal year end. Firms with no reported total debt are assigned a value of zero. To avoid any spurious relationship, we use lagged firm characteristics from the previous year in the regression.

Finally, we include the idiosyncratic volatility (VOL; annualized standard deviation based on the last 12 months' daily stock returns) as a measure of total risk.[28] Firms with higher volatilities are found to have lower subsequent returns in Ang et al. (2006), (2009) but higher subsequent returns in other studies (Spiegel and Wang (2006), Malkiel and Xu (2002)). Given these mixed results, we are interested in understanding the relationship between IRPs and idiosyncratic volatility.

V. Summary Statistics

In Section V, we present risk premiums computed for country portfolios. Our main goal in this section is to provide a benchmark comparison of the IRP observed in each country to its historical risk premium, computed using realized returns. In Section VI, we present the results of cross-sectional asset pricing tests.

Table 1 reports the equal-weighted average risk premiums in percentage for each G-7 country in each year from 1991 to 2000. As noted earlier, the IRPs are computed as the difference between the dollar cost of equity obtained from the discounted cash flow model in equation (7) as of the end of June each year and the yield on the 10-year U.S. Treasury bond at the end of June. The realized risk

[27] An alternative interpretation is that B/M ratios are proxies of mispricing, whereby high B/M ratios reflect undervaluation and low B/M ratios reflect overvaluation (see Lakonishok, Shleifer, and Vishny (1994)).

[28] Xiaoyan Zhang has kindly provided us with the monthly volatility data from Ang et al. (2006), (2009). We also test the robustness of our results using the standard deviation of monthly returns computed using at least 24 months of data over the prior 60 months as a measure of total risk. The results are very similar.

premiums (RRPs) are computed as the one-year-ahead ex post excess returns. The third-to-last and second-to-last columns of the table report the time-series averages and standard deviations across the IRPs and RRPs. The average IRP ranges from 2.6% in Japan to 9.2% in Canada. In comparison, the average equal-weighted RRP ranges from −0.8% in Japan to 13.1% in the U.S.

The most striking finding in Table 1 is the difference in volatility between the RRP and the IRP. The time-series standard deviation of the RRP for each firm is substantially larger than that of the IRP. To facilitate the comparison, we compute a standard deviation ratio (SD ratio), which is the ratio of the standard deviation of the RRP to the standard deviation of the IRP. We calculate the SD ratio for each firm and report the average of the SD ratios across all firms in each country.[29]

Table 1 shows that the SD ratio (in bold) ranges from a low of 12.13 (Canada) to a high of 18.33 (U.S.). To put these numbers in context, an SD ratio of 12.13 for Canada means that the standard deviation of RRP is 12.13 times that of the IRP for Canadian firms. For every country in our sample, the standard deviation of the RRP is more than 10 times the standard deviation of the IRP. The cross-country average of the SD ratios is 15.

Table 2 reports the summary statistics for risk premiums based on alternate implied cost of capital measures (see Section III) and various firm characteristics. The mean values across G-7 countries for our implied risk premium (IRP) measure and the alternate measures, $AIRP_{MPEG}$, $AIRP_{OJ}$, and $AIRP_{CT}$, are 5.80%, 4.83%, 7.57%, and 5.08%, respectively. While the means are comparable across IRP, $AIRP_{MPEG}$, and $AIRP_{CT}$, the Ohlson-Juettner-Nauroth measure ($AIRP_{OJ}$) is substantially higher, suggesting a possible overestimation of long-term growth in estimating the implied cost of capital. Perhaps more importantly, the standard deviations of all four measures are narrowly distributed, ranging from 1.28% to 1.69%. In contrast, the RRP is more than 10 times as volatile as the IRP, irrespective of the valuation model. We also report summary statistics for the error spread, which is the difference between RRP and IRP. The volatility of the error spread is comparable to that of the RRP.

In additional (untabulated) analyses, we also compute the pairwise correlation among the IRP, the three AIRPs, and the RRP. Although the IRP and AIRPs are derived under different assumptions, the correlations among them are quite high, ranging from 0.45 to 0.94. The measure that is least correlated with the other three measures is $AIRP_{CT}$, with correlations ranging from 0.44 to 0.49. None of these four implied measures, however, is highly correlated with the RRPs: Realized returns exhibit correlations of only 0.04 to 0.06 with the implied measures. This is perhaps not surprising, as we know that realized returns are quite noisy.

To further examine this issue, Table 3 provides evidence of the positive relation between the IRP and ex post realized returns using a portfolio approach. Specifically, Table 3 reports the average IRP and RRP for firms sorted into quintiles based upon their IRPs as of June 30 of each year. The results show that firms with higher IRPs (Q5) earn higher average excess returns than firms with lower IRPs (Q1) in each of the next two years. This relation is monotonic across

[29]To limit the effect of outliers, we windsorize these SD ratios at the top and bottom 1% level.

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TABLE 2

Summary Statistics of the Risk Premiums, Betas, and Firm Characteristics by Country

Table 2 reports the summary statistics of the implied risk premium (IRP), three alternative implied risk premiums (AIRPs), realized risk premium (RRP), error spread (ES), and firm characteristics by country. The IRP and AIRPs are computed as the difference between implied costs of capital and the yields on the U.S. 10-year government bond. The IRP and AIRPs are computed based on four different models of implied costs of capital: the 15-year discounted cash flow model described in the text (IRP), the modified PEG ratio model ($AIRP_{MPEG}$), the Ohlson-Juettner-Nauroth model ($AIRP_{OJ}$), and the Claus and Thomas model ($AIRP_{CT}$). We also report the RRP and the ES, defined as the difference between RRP and IRP. Details on each model are explained in the text. These characteristics include three international CAPM betas and four firm characteristics. Specifically, they are: individual firm world market beta (b_W); local market beta (b_L); exchange rate beta (b_E); individual firm idiosyncratic volatility as measured by annualized standard deviation over the last 12 months' daily stock returns (VOL); log of individual book leverage ratio (LEV), where book leverage ratio is defined as the ratio of total debt to total equity expressed in percentage; log of individual firm B/M ratio in percentage (LnBM); and log of individual firm market capitalization in millions of dollars (LnSZE).

	Risk Premiums						Betas			Firm Characteristics			
	IRP	$AIRP_{MPEG}$	$AIRP_{OJ}$	$AIRP_{CT}$	RRP	ES	b_W	b_L	b_E	VOL	LEV	LnBM	LnSZE
Canada													
Mean	9.17	7.48	9.75	6.88	8.43	−0.72	6.04	0.92	0.92	0.30	−0.72	0.33	4.02
Std Dev	1.95	1.33	1.55	1.05	17.18	18.04	0.21	0.10	0.09	0.09	18.04	0.10	0.23
France													
Mean	6.43	5.67	8.41	4.93	6.67	0.24	6.28	0.85	0.75	0.30	0.24	0.32	3.81
Std Dev	1.14	1.13	1.04	0.83	21.92	22.06	0.15	0.17	0.12	0.06	22.06	0.07	0.24
Germany													
Mean	4.06	3.48	6.78	4.08	−5.68	−9.75	6.68	0.85	0.78	0.28	−9.75	0.32	3.69
Std Dev	2.45	2.15	1.64	1.69	19.92	20.96	0.19	0.07	0.09	0.07	20.96	0.10	0.22
Italy													
Mean	6.73	5.42	8.02	4.28	2.93	−3.74	6.58	0.91	0.70	0.29	−3.74	0.33	4.04
Std Dev	2.05	1.83	1.61	1.10	26.50	26.83	0.45	0.10	0.11	0.05	26.83	0.05	0.39
Japan													
Mean	2.64	1.58	4.66	3.83	−0.75	−3.39	7.68	0.99	0.99	0.34	−3.39	0.36	3.89
Std Dev	2.03	1.89	1.32	1.70	29.91	29.23	0.39	0.12	0.13	0.08	29.23	0.09	0.16
U.K.													
Mean	5.35	4.89	7.85	5.53	10.69	5.35	5.70	0.92	0.76	0.21	5.35	0.25	3.53
Std Dev	1.43	1.52	1.20	1.16	13.11	13.25	0.37	0.13	0.07	0.05	13.25	0.06	0.16
U.S.													
Mean	6.19	5.30	7.55	6.03	11.63	5.44	6.19	0.87	0.82	0.40	5.44	0.44	3.66
Std Dev	0.76	0.80	0.58	1.42	9.38	9.24	0.35	0.05	0.26	0.06	9.24	0.08	0.14
Avg Mean	5.80	4.83	7.57	5.08	4.85	−0.94	6.45	0.90	0.82	0.30	−0.94	0.34	3.81
Avg Std Dev	1.69	1.52	1.28	1.28	19.70	19.94	0.30	0.11	0.12	0.07	19.94	0.08	0.22

TABLE 3

Relation between Realized and Implied Risk Premiums

Firms are sorted into quintiles on June 30 of each year from 1990 to 2000 based on their implied risk premiums (IRPs). The IRPs are computed as the differences between the dollar costs of equity computed from a 15-year discounted cash flow model that uses analyst forecasts of earnings and exchange rates, the mean reversion to long-run average GDP growth rates across G-7 countries, and the yields on the U.S. 10-year government bond. For each quintile, realized risk premiums (RRPs) are calculated as the time-series equal-weighted return in years 1 and 2 after the ranking period minus the realized returns on the 10-year government bond. Risk premium differences for quintile 5 minus quintile 1 are reported (Q5−Q1), along with Newey-West autocorrelation corrected *t*-statistics. ***, **, * indicate 1%, 5%, and 10% significance in the direction predicted, respectively.

	Quintile Sort						
	Smallest				Largest		
Ranked by	Q1	Q2	Q3	Q4	Q5	Q5−Q1	*t*-Stat
IRP	−0.52	1.95	3.93	7.07	16.98	17.49	18.19***
RRP (year 1)	3.42	7.07	8.49	9.02	9.58	6.16	5.05***
RRP (year 2)	3.06	6.68	6.86	7.56	9.49	6.44	2.30***

the quintiles, with the difference between Q5 and Q1 statistically significant at the 5% level.[30] Evidently, the IRP is positively related to the RRP, although the relationship is fraught with noise.

VI. Cross-Sectional Tests

A. Univariate Portfolio Tests

Table 4 provides evidence of the univariate relationship between factor betas, firm characteristics, and the IRP. To construct this table, we sort firms into quintiles based upon each factor beta or firm characteristic as of June 30 in each year and compute the mean IRP as of June in the next year for the firms in each

TABLE 4

Implied Risk Premiums and Firm Betas/Characteristics

On June 30 of each year, stocks are sorted into quintiles based on each of the 10 firm betas or characteristics. Table 4 reports the time-series average implied risk premium (IRP) and realized risk premium (RRP) for each quintile. Panel A shows the results for betas, while Panel B shows the results for firm characteristics. The betas include the individual firm world market beta (b_W), local market beta (b_L), and exchange rate beta (b_E). Firm characteristics include individual firm idiosyncratic volatility as measured by annualized standard deviation based on last 12 months' daily stock returns (VOL); log of individual book leverage ratio (LEV), where book leverage ratio is defined as the ratio of total debt to total equity; log of individual firm market capitalization (LnSZE); and log of individual firm book-to-market ratio (LnBM). The IRPs are computed as the differences between the dollar costs of equity computed from a 15-year discounted cash flow model that uses analyst forecasts of earnings and exchange rates, the mean reversion to long-run average GDP growth rates across G-7 countries, and the yields on the U.S. 10-year government bond. RRPs are defined as the annual equal-weighted returns for each portfolio in the year after the ranking, minus the realized returns on the 10-year government bond, and averaged across all years. Risk premium differences for quintile 5 minus quintile 1 are reported (Q5−Q1). Also reported are *t*-statistics computed using Newey-West correction with two lags of autocorrelations. ***, **, * indicate 1%, 5%, and 10% significance in the direction predicted, respectively.

	Ranked by						
	Smallest				Largest		
	Q1	Q2	Q3	Q4	Q5	Q5−Q1	*t*-Stat
Panel A. Firm-Level Betas							
b_W	0.25	0.64	0.86	1.08	1.59		
IRP	5.68	5.13	5.28	5.54	7.06	1.38	2.97***
RRP	7.32	8.25	7.70	6.58	8.55	1.22	0.28
b_L	−0.20	0.52	0.85	1.17	1.91		
IRP	6.28	5.24	4.97	5.30	6.88	0.60	1.90*
RRP	7.46	7.25	6.73	6.09	10.91	3.46	0.96
b_E	−1.81	−0.67	−0.20	0.24	1.10		
IRP	7.57	5.78	4.97	4.66	5.70	−1.87	−2.86***
RRP	9.14	7.88	7.67	5.84	7.84	−1.30	−0.67
Panel B. Firm Characteristics							
VOL	0.22	0.30	0.37	0.47	0.70		
IRP	3.51	4.34	5.32	6.72	9.35	5.83	8.91***
RRP	6.14	6.14	6.23	7.66	12.39	6.25	0.94
LEV	4.56	21.49	45.11	81.88	247.84		
IRP	5.36	5.37	5.85	6.20	6.77	1.41	2.82***
RRP	5.79	6.08	6.78	7.50	10.59	4.79	0.909
LnSZE	4.73	6.21	7.00	7.87	9.29		
IRP	7.95	5.39	4.49	3.80	3.18	−4.77	−15.68***
RRP	10.36	7.22	4.37	5.85	4.67	−5.69	−1.48
LnBM	2.58	3.38	3.78	4.11	4.62		
IRP	4.35	4.41	5.26	6.48	8.88	4.53	5.87***
RRP	3.46	5.57	6.23	8.72	13.70	10.24	1.92*

[30]We have also examined deciles instead of quintiles. The results are similar.

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portfolio. The one-year gap between firm characteristics and the IRP serves to minimize spurious correlations between these characteristics and the dependent variable. We compute time-series averages of cross-sectional means and report the Newey-West autocorrelation-adjusted *t*-statistics for the difference in risk premiums across the extreme quintiles (Q1 and Q5).

Panel A of Table 4 presents results for the factor betas, and Panel B presents results for the firm characteristics. These results show that firms with high world market betas have higher IRPs, although the positive relation is not monotonic across quintiles. The difference between the Q5 and Q1 mean IRPs is 1.38% per annum. This difference is statistically significant with a *t*-statistic of 2.97. The difference between the Q5 and Q1 mean RRPs is also positive and is comparable in magnitude (1.22%). However, this difference is not statistically significant (*t*-statistic of only 0.28), illustrating the difference in the power of the two tests.

Local beta (based on the local market return orthogonalized to the world market return) is also positively related to the IRP. Panel A of Table 4 shows that the spread in IRPs between the extreme quintiles is 0.60%, which is significant at the 10% level. However, as with the world market beta, the correlation between the local beta and the RRP is not statistically significant.

Panel A of Table 4 also shows that firms with high currency betas have reliably lower IRPs. The risk premium spread between extreme quintiles (Q5−Q1) is −1.87%, which is statistically significant at the 1% level. Economically, a negative risk premium implies that investors view stocks with positive currency betas (stocks that perform well when the dollar depreciates) as a hedge against currency risk (the risk of dollar depreciation). The results in Panel A suggest that investors are willing to accept a lower risk premium for such stocks.

Panel B of Table 4 reports the results for firm characteristics. Interestingly, stocks with higher idiosyncratic volatility have significantly higher IRPs. This supports the results in Spiegel and Wang (2006), that at the monthly horizon, stocks with higher idiosyncratic risk have lower liquidity and higher returns. Our findings differ from Ang et al. (2006), (2009), who report that stocks with higher idiosyncratic volatility have lower subsequent returns.

Consistent with Modigliani-Miller (1958), we find that firms with higher financial leverage face higher costs of capital. This is also consistent with Bhandari's (1988) findings that leverage increases expected returns. Finally, we find that firms with smaller size and higher B/M ratios tend to have higher IRPs, which is consistent with previous empirical work using U.S. data.

B. Multivariate Cross-Sectional Regressions

In this section, we estimate the cross-sectional relationship among factor loadings, firm characteristics, and the IRPs in a multivariate context.[31] This approach allows us to evaluate the incremental explanatory power of each beta or

[31] We estimated cross-sectional regressions with value-weighted average industry-country portfolio risk premiums as the dependent variable (not reported), because portfolio risk premiums might be estimated more precisely than individual firm risk premiums. The result is similar to the firm-level result. We also estimated regressions involving only the country or only the industry risk premiums.

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firm characteristic while controlling for the other variables. Specifically, we estimate the following Fama-MacBeth regression for year t with K factor loadings and M characteristics:

$$r_{it} - r_{ft} = a + \sum_{k=1}^{K} \gamma_{kt}\beta_{ikt} + \sum_{m=1}^{M} \gamma_{mt}C_{imt} + u_{it}, \tag{13}$$

where $r_{it} - r_{ft}$ is the dollar risk premium estimate (relative to long-term U.S. Treasury yields) for firm i in year t; β_{ikt} is the factor loading, or beta, corresponding to risk factor F_k; C_{imt} is the firm characteristic m for firm i in year t; and γ_{kt} and γ_{mt} are the slope coefficients corresponding to betas and characteristics, respectively, from the cross-sectional regressions in year t.

In our empirical implementation, we make sure that all explanatory variables, except betas and idiosyncratic volatility, are from the year prior to that of the risk premium estimates. The cross-sectional regression in equation (13) is estimated each year from 1990 to 2000, with betas and characteristics as independent variables. The time-series averages and t-statistics of the estimated slope coefficients are then used to examine the cross-sectional asset pricing implications. Because the risk premiums are autocorrelated over time, we use a Newey-West autocorrelation correction with two moving average lags to compute the t-statistics. We also report the time-series average of adjusted R^2 of each regression specification over the sample period.

In Table 5, we report the results of six different regression specifications. The first specification (S1) estimates a pure beta model in which the only independent variables are betas from the factor model in equation (12). The second specification (S2) adds idiosyncratic volatility to the betas. The third specification (S3) adds leverage to the betas. The fourth specification (S4) includes size. The fifth specification (S5) adds the B/M ratio. The sixth specification (S6) includes all variables. In all six models, the dependent variable is the individual firm risk premium.

In the pure beta specification S1, the coefficient with respect to the world beta, b_W, is marginally significant, with a t-statistic of 1.72. The local market beta is not significant. The currency beta, b_E, has a negative sign and is statistically significant. These results suggest that currency beta is an important determinant of international risk premiums. In particular, the regression confirms the univariate finding that firms with high currency beta tend to have lower risk premiums.

Regression specification S2 examines the effect of the total risk variable, idiosyncratic volatility, on the cost of capital. We find that idiosyncratic volatility is highly significant, and that its introduction reduces the importance of both the currency beta and the world market beta (which now becomes insignificant). When idiosyncratic volatility is added to the model, the adjusted R^2 increases from 3% to 11%. In fact, idiosyncratic volatility continues to be one of the most significant determinants of the cross-sectional variation in international risk premiums, even with the inclusion of all of the other firm characteristics.

These regressions suffer from a lack of power because there are few cross-sectional observations. As expected, these regressions have very low R^2 and insignificant t-statistics.

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TABLE 5

Cross-Sectional Regressions of Implied Risk Premiums on Various Explanatory Variables

Table 5 presents time-series averages of slope coefficients from cross-sectional Fama-MacBeth regressions estimated as of June 30 each year from 1990 to 2000. The implied risk premiums are computed as the differences between the dollar costs of equity and the yields on the U.S. 10-year government bond. The dollar costs of equity are computed from a 15-year discounted cash flow model that uses forecast exchange rates to convert analyst forecasts of earnings into dollar denominations. Individual firm risk premiums are regressed on: individual firm world market beta (b_W); local market beta (b_L); exchange rate beta using nominal trade-weighted index (b_E); individual firm idiosyncratic volatility as measured by annualized standard deviation based on last 12 months' daily stock returns (VOL); log of individual book leverage ratio (LEV), where book leverage ratio is defined as the ratio of total debt to total equity; log of individual firm market capitalization (LnSZE); and log of individual firm book-to-market ratio (LnBM). We estimate various regression specifications (S1 to S6) using different combinations of these independent variables, with each column representing a different specification. The numbers in parentheses are *t*-statistics computed using Newey-West correction with two lags of autocorrelations. Coefficients with 5% statistical significance are indicated in bold. The average adjusted R^2 for each model over the sample period is shown. *N* represents average number of firms each year.

	Regression Specifications					
	S1	S2	S3	S4	S5	S6
Intercept	4.45 (8.03)	1.02 (1.45)	3.52 (4.46)	10.41 (14.00)	−3.31 (−1.78)	−5.24 (−1.62)
b_W	0.66 (1.72)	0.03 (0.13)	**0.87** **(2.06)**	**1.00** **(2.22)**	**1.11** **(2.16)**	**0.67** **(2.37)**
b_L	0.44 (1.23)	0.10 (0.40)	0.44 (1.33)	**0.63** **(2.12)**	0.56 (1.44)	0.40 (1.75)
b_E	**−0.72** **(−2.16)**	**−0.34** **(−1.86)**	**−0.78** **(−2.33)**	**−0.55** **(−1.83)**	**−0.75** **(−2.30)**	−0.29 (−1.77)
VOL		**11.64** **(10.12)**				**11.58** **(8.06)**
LEV			**0.22** **(2.15)**			**0.28** **(4.67)**
LnSZE				**−0.99** **(−16.66)**		**−0.42** **(−2.62)**
LnBM					**1.95** **(5.18)**	**1.96** **(4.70)**
Adj. R^2	0.03	0.11	0.03	0.10	0.08	0.195
N	2,129	2,062	1,879	2,048	2,040	1,814

Specification S3 evaluates the relation between leverage and cost of capital, controlling for other variables. Just as in the univariate case, higher leverage corresponds to higher implied cost of capital. Specifications S4 and S5 show that, as in the univariate tests, size and B/M ratio play a significant role in explaining cross-sectional IRPs. Smaller firms face a higher cost of capital, as do firms that have higher B/M ratios (perhaps due to greater distress risk).

In specification S6, all seven explanatory variables are included. This specification shows that the most significant cross-sectional determinants of firm-level risk premiums are world market beta, currency beta, idiosyncratic volatility, leverage, B/M ratio, and size. Interestingly, once firm characteristics are introduced, the world beta becomes more significant. The currency beta is significant in most specifications, although it becomes marginal in the specification that includes all variables.[32] The firm characteristics (B/M ratio, size, leverage, and idiosyncratic volatility) are generally more significant than the betas. Overall, a model that includes the betas, idiosyncratic volatility, leverage, B/M ratio, and size can explain

[32] See Griffin (2002) for similar conclusions on the noisy nature of beta estimates and the importance of local factors in an international context.

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approximately 20% (19.5%) of the cross-sectional variation in individual firm risk premiums.

C. Regressions Based on Alternate Measures of Implied Cost of Capital

In this section, we examine the robustness of our cross-sectional findings to the risk premiums based on the three alternate IRPs: $AIRP_{MPEG}$, $AIRP_{OJ}$, and $AIRP_{CT}$. Table 6 presents the results from cross-sectional regressions involving these three measures. To conserve space, we only report results for the pure beta specification and the complete specification (models S1 and S6 from Table 5). The main finding in Table 6 is that the key results in Table 5 are not sensitive to the use of alternate measures of implied costs of capital. The currency beta, idiosyncratic volatility, leverage, and B/M ratio are all important in determining the IRPs. The only difference relates to the world market beta, which is in the right direction for all six specifications but is significant only for the $AIRP_{OJ}$ measure.

TABLE 6

Cross-Sectional Regressions Based on Three Alternate Implied Risk Premiums

Table 6 presents time-series averages of slope coefficients from cross-sectional Fama-MacBeth regressions estimated as of June 30 each year from 1990 to 2000. The alternate implied risk premiums (AIRPs) are computed as the differences between the implied costs of capital and the yield on the U.S. 10-year government bond. The alternate implied costs of capital are computed using three different methods: modified PEG ratio (MPEG) in Panel A, Ohlson and Juettner-Nauroth (OJ) in Panel B, and Claus and Thomas (CT) in Panel C. Details of each method are explained in the text. Individual firm risk premiums are regressed on individual firm world market beta (b_W); local market beta (b_L); exchange rate beta (b_E); individual firm idiosyncratic volatility as measured by annualized standard deviation based on last 12 months' daily stock returns (VOL); log of individual book leverage ratio (LEV), where the leverage ratio is defined as the ratio of total debt to total equity; log of individual firm market capitalization (LnSZE); and log of individual firm book-to-market ratio (LnBM). The numbers in parentheses are *t*-statistics computed using Newey-West correction with two lags of autocorrelation. Coefficients with 5% statistical significance are indicated in bold. The average adjusted R^2 for each model over the sample period is shown. *N* represents average number of firms each year.

	Panel A $AIRP_{MPEG}$		*Panel B* $AIRP_{OJ}$		*Panel C* $AIRP_{CT}$	
	S1	S2	S3	S4	S5	S6
Intercept	4.23 (7.58)	−2.36 (−0.82)	6.53 (16.20)	−0.25 (−0.09)	4.92 (10.87)	1.41 (0.89)
b_W	0.26 (0.84)	0.39 (1.57)	0.50 (1.68)	**0.50** **(2.68)**	0.24 (1.23)	0.28 (1.36)
b_L	0.33 (1.05)	0.31 (1.54)	0.23 (0.88)	0.20 (1.24)	0.21 (1.18)	0.17 (1.47)
b_E	**−0.65** **(−2.26)**	**−0.31** **(−2.19)**	**−0.55** **(−2.11)**	−0.21 (−1.64)	**−0.39** **(−2.35)**	**−0.21** **(−2.10)**
VOL		**8.24** **(6.85)**		**9.09** **(7.23)**		**4.79** **(6.87)**
LEV		**0.28** **(4.33)**		**0.22** **(4.16)**		**0.17** **(4.38)**
LnSZE		**−0.44** **(−3.31)**		**−0.40** **(−2.42)**		**−0.24** **(−2.48)**
LnBM		**1.45** **(4.40)**		**1.43** **(3.75)**		**0.72** **(3.67)**
Adj. R^2	0.03	0.23	0.03	0.20	0.01	0.08
N	2,171	1,848	2,056	1,749	1,749	1,478

The fact that our results are invariant to different measures of implied cost of capital is encouraging because it shows that the results are not driven by spurious effects such as measurement errors specific to any particular implementation. However, we hasten to point out that all four models are based on similar

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theoretical constructs and could be vulnerable to similar problems.[33] Moreover, it is possible that commonality in the way that we handled model implementation or noise filtering could have induced a greater degree of commonality in results across these models. Despite these caveats, Table 6 provides an additional measure of assurance that the earlier results are not due to particular quirks of any given valuation model.

D. Regressions Using Realized Returns

Table 7 presents results from cross-sectional regressions that use realized excess returns in place of IRPs. Our goal in constructing this table is to compare and contrast these findings with those obtained earlier using the implied cost of

TABLE 7

Cross-Sectional Regressions of Realized Risk Premiums and Error Spreads on Various Explanatory Variables

Table 7 presents time-series averages of slope coefficients from cross-sectional Fama-MacBeth regressions estimated as of June 30 each year from 1990 to 2000. Panel A shows the results for the full sample, and Panel B shows the results for the subsample where IBES analysts' earning forecasts are available. Panel C shows the results using error spread (ES), which is the difference between the realized risk premium (RRP) and the implied risk premium (IRP). The risk premium is computed as the difference between one-year-ahead realized return for the stock and the yield on the U.S. 10-year government bond. Individual firm risk premiums and error spreads are regressed on individual firm world market beta (b_W); local market beta (b_L); exchange rate beta (b_E); individual firm idiosyncratic volatility as measured by annualized standard deviation based on last 12 months' daily stock returns (VOL); log of individual book leverage ratio (LEV), where book leverage ratio is defined as the ratio of total debt to total equity; log of individual firm market capitalization (LnSZE); and log of individual firm book-to-market ratio (LnBM). We estimate various regression specifications using different combinations of these independent variables, with each column representing a different specification. The numbers in parentheses are *t*-statistics computed using Newey-West correction with two lags of autocorrelation. Coefficients with 5% statistical significance are indicated in bold. The average adjusted R^2 for each model over the sample period is shown. *N* represents average number of firms each year.

	Panel A RRP (full sample)		*Panel B* RRP (IBES sample)		*Panel C* ES (RRP − IRP)	
	S1	S2	S3	S4	S5	S6
Intercept	6.04 (1.96)	−2.74 (−0.29)	6.10 (1.27)	17.15 (1.40)	1.65 (0.34)	22.41 (1.88)
b_W	−0.61 (−0.26)	−0.22 (−0.15)	1.22 (0.34)	0.15 (0.10)	0.56 (0.16)	−0.51 (−0.35)
b_L	1.84 (1.81)	**1.99** **(2.68)**	2.86 (1.85)	**3.02** **(2.71)**	2.42 (1.45)	**2.62** **(2.17)**
b_E	**−1.29** **(−2.68)**	**−1.50** **(−4.48)**	**−1.50** **(−2.95)**	**−1.31** **(−3.04)**	**−0.78** **(−1.98)**	**−1.02** **(−2.38)**
VOL		0.79 (0.09)		−1.83 (−0.11)		−13.45 (−0.76)
LEV		0.28 (0.30)		0.60 (0.81)		0.31 (0.44)
LnSZE		−1.28 (−1.52)		**−2.78** **(−6.94)**		**−2.36** **(−5.27)**
LnBM		**3.34** **(2.70)**		1.02 (0.61)		−0.94 (−0.74)
Adj. R^2	0.01	0.04	0.02	0.05	0.01	0.05
N	6,999	5,328	2,125	1,812	2,125	1,812

[33] For example, Pastor et al. (2008) use simulation analysis to show that even crude measures of implied cost of capital with no information about future cash flow growth should be correlated with conditional expected returns as long as some of the stock price changes are driven by changes in expected returns.

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capital. If the noise in realized returns is the primary culprit in traditional asset pricing tests, then we would hope to see similar regression coefficients in the realized return regressions, but with lower statistical significance. This is, in general, what we find in Table 7.

Panel A of Table 7 reports the results for the full sample, and Panel B reports the results for the subsample for which IBES analysts' earnings forecasts are available. The results in both panels show that only the currency beta is uniformly significant. While idiosyncratic volatility has a positive sign in Panel A, it has a negative sign in Panel B. The local beta loads more significantly than the world market beta, which is consistent with Griffin (2002), who finds that the local beta is more important than the world beta. B/M ratio and size are significant in some specifications but not in others. Leverage has the right sign but is not statistically significant. Even under the best model (S6), only 5% of the cross-sectional variation in realized returns is captured by the betas and the characteristics.

These findings illustrate the difficulty underlying the use of realized returns as proxies for expected returns. Notice that the sample of firms with analyst coverage (in Panel B of Table 7) is much smaller than the sample of firms in broad cross-sectional studies such as Hou et al. (2006) and De Moor and Sercu (2006). This is because IBES analysts tend to cover only the larger, more liquid firms in each market. The fact that both samples provide similar results with respect to realized returns alleviates concerns that our results might be specific to our more restricted sample.[34]

Panel C of Table 7 examines the relation between the difference in realized returns and the implied cost of capital (i.e., the "error spread") and betas and firm characteristics. The error spread can be interpreted as the news or the expectation error in realized returns. The error spread, however, is also dominated by the noise in realized returns because the implied expected returns vary little, while the realized returns vary to a considerable extent. As a result, the standard deviations of the error spreads are indistinguishable from the standard deviations of realized returns. For the overall sample, the standard deviation of annual realized excess returns is 19.70%, while the standard deviation of the error spread is 19.94%.

Empirically, we expect the error spread results to be similar to those based on realized returns. This is what we observe in Panel C of Table 7. The error spread results are quite similar to the IBES sample results in Panel B with the local beta, the exchange beta, and firm size registering statistical significance and the other variables registering insignificance. The slope coefficients from the error spread regression are almost exactly equal to the difference between the slope coefficients corresponding to the IRP in Table 5 and those corresponding to realized returns in Panel B. The differences are not statistically significant even if they might be economically significant. The adjusted R^2 of the error spread regression in Panel C is 5%, which is much lower than the adjusted R^2 of the IRP regressions

[34]De Moor and Sercu (2006) show that a data sample including ultra-small firms can have a large impact on the Fama and French (1993) results. Unfortunately, in any measure of implied cost of capital, IBES earning forecasts are a crucial input to construct expected cash flows. Hence, we are limited to the IBES sample.

in Tables 5 and 6, but it is comparable to the adjusted R^2 of the realized excess returns in Panels A and B. Collectively, these findings show that few systematic components of the error spreads can be extracted by the model.

We also conducted robustness checks of our results of error spread regressions by reestimating regressions for our three alternate measures of implied cost of capital and for each country, based on our risk premium measure. The results, unreported, are quite similar to those in Panel C of Table 7. For each alternate measure of implied cost of capital, only the local beta, currency beta, and firm size are statistically significant. The results for individual countries are even noisier because of reduced sample size. Finally, we have also estimated error-spread regressions as of December year-end, i.e., measuring annual returns from January to December as opposed to July to June as in Panel A, and find that under this formulation, none of the slope coefficients are statistically significant. Overall, these results indicate that the divergence in the results between implied cost of capital and realized excess returns is not specific to one country, one measure of implied cost of capital, or seasonal effects. As we discussed earlier, the reason for the divergence is the noise in realized returns.

E. Robustness Tests on Forecast Horizon, Sample, and Model

We further conducted a number of robustness checks on forecast horizon, data sample, and model. First, we have reestimated our main implied cost of capital model for a 10-year forecast horizon and a 20-year forecast horizon, instead of a 15-year horizon. The cross-sectional results are essentially unchanged.

Second, we are concerned that our sample of firms is dominated by firms in the U.S., the U.K., and Japan. To address this, we reestimate our cross-sectional regressions using only the 100 largest companies from each of the G-7 countries. The results are virtually unchanged.

Third, we reestimate our results using alternate year-end dates. One concern is that our findings may be dependent upon our arbitrary choice of June 30 as the date for the estimation of implied cost of capital. As Chan, Karceski, and Lakonishok (1998) show, there may be seasonal patterns in the returns for different B/M ratios, momentums, sizes, and other portfolios. To ensure that such seasonalities do not affect our cross-sectional findings, we estimate the implied cost of capital using market closing prices as of three different calendar months (January, March, and December). The results show that our key findings are robust to the use of alternate year-end dates.

Fourth, since asset pricing tests typically use monthly data, we reestimate the implied cost of capital on a monthly basis and reestimate Fama-MacBeth cross-sectional regressions using the monthly data. The results show that our key findings are robust to the use of monthly estimation.

Finally, we examine the cross-sectional relationship between betas estimated from an international version of the Fama and French (1993) three-factor model used by Zhang (2006). Zhang extends Fama and French (1998) by adding a size factor:

$$(14) \quad r_t - r_{ft} = a + b_{\mathrm{WMKT}}\mathrm{WMKT}_t + b_{\mathrm{WSMB}}\mathrm{WSMB}_t + b_{\mathrm{WHML}}\mathrm{WHML}_t + u_t,$$

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where WMKT is the excess return on the world market portfolio, defined as the return on the Datastream Value-Weighted Global Index less the one-month Eurodollar deposit rate; WSMB is the return on a zero-investment portfolio designed to capture small firm risks; and WHML is the return on a zero-investment portfolio designed to capture B/M risks.[35] We use the model in equation (14) to estimate betas for each firm.

The results in Table 8 indicate that size and B/M betas are significantly positively related to the cross section of IRPs even after controlling for different characteristics. The world market beta turns insignificant in the presence of size and B/M betas, while leverage, idiosyncratic volatility, and B/M ratio continue to be

TABLE 8

Cross-Sectional Regressions with Fama-French Betas

Table 8 presents time-series averages of slope coefficients from cross-sectional Fama-MacBeth regressions estimated as of June 30 each year from 1990 to 2000. The implied risk premiums are computed as the differences between the dollar costs of equity and the yields on the U.S. 10-year government bond. The dollar costs of equity are computed from a 15-year discounted cash flow model that uses forecast exchange rates to convert analyst forecasts of earnings into dollar denominations. Individual firm risk premiums are regressed on: Fama-French world market beta (b_{WMKT}); beta for a high-minus-low book-to-market factor (b_{WHML}); beta for a small-minus-big factor (b_{WSMB}); individual firm idiosyncratic volatility as measured by annualized standard deviation based on last 12 months' daily stock returns (VOL); log of individual book leverage ratio (LEV), where book leverage ratio is defined as the ratio of total debt to total equity; log of individual firm market capitalization (LnSZE); and log of individual firm book-to-market ratio (LnBM). We estimate various regression specifications (S1 to S6) using different combinations of these independent variables with each column representing a different specification. The numbers in parentheses are *t*-statistics computed using Newey-West correction with two lags of autocorrelations. Coefficients with 5% statistical significance are indicated in bold. The average adjusted R^2 for each model over the sample period is shown. *N* represents average number of firms each year.

	Regression Specifications					
	S1	S2	S3	S4	S5	S6
Intercept	4.76 (9.35)	1.36 (2.29)	3.61 (5.45)	9.77 (15.57)	−3.11 (−1.98)	−5.89 (−1.81)
b_{WMKT}	−0.18 (−0.60)	**−0.58** **(−3.53)**	0.16 (0.25)	0.64 (1.19)	0.42 (0.60)	0.17 (0.38)
b_{WHML}	0.28 (1.67)	**0.62** **(4.04)**	0.29 (1.69)	0.18 (1.11)	0.06 (0.50)	**0.34** **(3.06)**
b_{WSMB}	**1.71** **(7.43)**	**1.06** **(4.70)**	**1.77** **(7.86)**	**1.05** **(6.03)**	**1.61** **(8.54)**	**0.73** **(4.66)**
VOL		**11.28** **(7.86)**				**11.91** **(7.21)**
LEV			**0.27** **(3.33)**			**0.25** **(4.77)**
LnSZE				**−0.82** **(−15.90)**		−0.26 (−1.50)
LnBM					**1.99** **(6.23)**	**1.99** **(4.94)**
Adj. R^2	0.06	0.13	0.07	0.11	0.11	0.20
N	2,113	2,048	1,867	2,034	2,026	1,803

[35]Zhang (2006) follows Fama and French (1993), (1998) closely in her portfolio construction. Zhang (2006) constructs the WHML and WSMB factors using firm-level returns, B/M ratios, and market capitalizations from the U.S., the U.K., and Japan, as obtained from Datastream. Within each country, the high and low B/M portfolios and small and large size portfolios are first formed. These country value and size portfolios are formed using value weights. The world value and size portfolios are formed by aggregating across countries' value and size portfolios using the countries' market capitalizations. The WHML factor represents the difference in returns of the high and low world value portfolios, while the WSMB factor represents the difference in returns of world small and big size portfolios.

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significant.[36] These results suggest a possible role for size and B/M factors in international asset pricing. The challenge, however, is to understand why these factors might be important in international asset pricing, since the theoretical basis for these factors in international asset pricing is unclear.

F. Robustness Test on Analyst Forecast Errors

One potential concern with the use of analyst forecasts is the degree of bias in them. It is well known that analyst forecasts are biased (i.e., tend to be optimistic). If the bias in these forecasts differs significantly across countries, our IRP estimates could be affected. We have taken several steps to handle this potential problem.

First, prior studies show that while analyst forecasts tend to be over-optimistic, the severity of the bias is quite similar across countries. For example, Frankel and Lee (1999) report that the average bias (forecast minus actual, scaled by the forecast) in one-year-ahead forecasts is between 1% and 13% across the G-7 countries, with the largest bias (13%) observed for U.S. analysts, while the bias for everyone else is less than 10%. Similarly, Ang and Ciccone (2001) report standard deviation of analyst forecast error (but not the bias) across the G-7 countries. Their estimates range between 6% and 19%, suggesting once again that the magnitude of the forecast errors is comparable across our sample countries.

To investigate this issue further, we explicitly adjust for ex post country-level forecast biases. The ex post forecast biases are computed using subsequent realized earnings of firms in each country. We then subtract these ex post country-mean forecast biases from each firm's earnings forecasts before the implied cost of capital is computed. We find that our cross-sectional results (available from the authors) remain similar to the main results.

Second, in order to minimize the effect of these biases in long-run cash flows, we use an exponential rate of mean reversion to forecast long-run growth rates, thus ensuring that the biases in short-run forecasts have minimal effect on the long-run growth rates.

Third, we conduct a Monte Carlo simulation that examines the potential effect of analyst forecast errors on our cost of capital estimates. The results (available from the authors) show that the risk premium estimates vary from −1.0% to +1.0%, on average, based on the most extreme 5% and 95% forecasts (drawn from a simulated bivariate normal distribution of FY1 and FY2 estimates). These results suggest that analyst forecast errors are not likely to be a major factor in our results.

VII. Conclusions

This paper uses a new approach based on discounted cash flow models to estimate cost of equity capital for firms in the G-7 countries. The results show that

[36] We have also estimated cross-sectional regressions involving realized returns and the error spread based on the three-factor model in (14), and we find that these regressions are noisy compared to those based on the implied cost of capital presented in Table 8.

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firm characteristics such as idiosyncratic volatility, leverage, B/M ratio, and size are important in explaining the risk premiums. Currency betas and, to a smaller extent, world market betas, as suggested by international asset pricing theory, also have power to explain the estimated risk premiums. Together, the betas and the characteristics explain approximately 20% of the cross-sectional variation in risk premiums in G-7 countries.

Our results are subject to a few caveats. First, we acknowledge that this study is limited in scope, particularly with respect to country membership. Second, we acknowledge that any large-scale estimation of valuation models across many firms involves simplifying assumptions. Although we have done extensive robustness checks, including the use of four different valuation models, it is still possible that commonality in the way we handled model implementation or noise filtering could have induced a greater degree of commonality in results across these models. However, we see no reason that these design issues should result in the systematic relationships with betas and firm characteristics that we document.

Collectively, our findings support a broader use of the implied cost of capital in the finance literature, particularly in international asset pricing. Our findings should also be of interest to financial practitioners who wish to estimate the cost of equity capital for their international investments. The consistency of our regression results suggests that the relations we document can potentially be used to estimate cross-sectional costs of capital in an international context. We view this as a promising topic for future research.

The fact that the firm characteristics are priced also highlights the need to develop alternate asset pricing models that might accommodate these empirical findings. For example, Daniel, Hirshleifer, and Subrahmanyam (2001) entertain a behavioral asset pricing model in which expected returns are linearly related to both systematic risk and mispricing measures. Their model implies that B/M ratios should be priced. Daniel and Titman (1997) and Daniel, Titman, and Wei (2001) find that characteristics rather than covariances drive the cross-sectional expected returns in the U.S. and Japan.[37] Our results involving the role of B/M ratio and size are consistent with these findings

In summary, our main message is that the implied cost of capital approach can provide important new insights into the cross-sectional determinants of firm-level expected returns in the international context. We view these results as encouraging, and we hope that they will serve as catalysts toward a broader acceptance of implied expected returns in the asset pricing literature.

Appendix. Detailed Description of Alternate Implied Costs of Capital

Botosan and Plumlee (2005) provide a description of the assumptions and data requirements among different implied costs of capital methodologies. Our summary is a modified version of their description. Our implementations of the models follow Hail and Leuz (2006) (see Table A1).

[37] Davis, Fama, and French (2000) object to this interpretation.

TABLE A1

Detailed Description of Alternate Implied Costs of Capital

Table A1 describes the four different models of implied costs of capital: the 15-year discounted cash flow model described in the text (r_e), the Ohlson-Juettner-Nauroth model (r_{OJ}), the modified PEG ratio model (r_{MPEG}), and the Claus and Thomas model (r_{CT}). The second column labeled CS refers to whether clean surplus accounting is assumed. YES means that clean surplus accounting is assumed, while NO means that clean surplus accounting is not required. The third column refers to the assumption in the short horizon, while the fourth column refers to the assumption for the terminal value. The last column describes the IBES forecasts used.

	CS	Short-Horizon	Terminal Value	IBES Forecasts
r_e	YES	During the explicit forecast period, analysts' forecasts of earnings and book value are assumed to be equal to market expectation. During the explicit forecast period, firm ROE fades linearly to industry ROE, while firms' growth rates fade exponentially to long-term growth rates.	The explicit forecast horizon we implement is 15 years. Beyond the forecast horizon, there are two major assumptions underlying our model: i) growth rate is assumed to revert to long-run GDP growth rate, and ii) return on new investment, ROI, is assumed to revert to cost of equity, r_e.	Current year, one- and two-year-ahead earnings per share. Dividends per share. Book value per share.
r_{OJ}	NO	Analysts' forecasts of earnings in years 1 and 2 equal the market expectation. Year 1 earnings and year 2 abnormal earnings are positive. Abnormal earning is defined as: $r^{-1}(FE_2 - rD_1 - (1 + r)FE_1)$	The explicit forecast horizon is two years. Growth in abnormal earnings occurs at a constant, economy-wide growth rate. Estimated constant rate of growth in abnormal earnings is the same as the market's expectation. Constant rate of growth is below the cost of equity and above zero.	Current year dividends per share. One- and two-year-ahead earnings per share.
r_{MPEG}	NO	Analysts' forecasts of earnings in years 1 and 2 equal the market expectation. Zero dividends in year 1. Year 1 earnings and year 2 abnormal earnings are positive. Year 2 abnormal earnings are defined as: $r^{-1}(FE_2 - (1 + r)FE_1)$	The explicit forecast horizon is two years. Beyond the forecast horizon, growth in abnormal earnings are assumed to be zero.	One-year-ahead earnings per share.
r_{CT}	YES	During the explicit forecast period, analysts' forecasts of earnings up to three years ahead are assumed to equal market expectation. Each residual income in the future equals the forecast earnings minus the book value times the implied cost of capital.	The explicit forecast horizon is three years. Beyond the forecast horizon, the nominal residual income is assumed to grow at GDP growth rate.	Current book value per share. One- and two-year-ahead earnings per share.

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