International Differences in the Cost of Equity Capital: Do Legal Institutions and Securities Regulation Matter?*

Luzi Hail

The Wharton School
University of Pennsylvania

and

Christian Leuz

The Wharton School
University of Pennsylvania

December 2005

Forthcoming in Journal of Accounting Research June 2006

ABSTRACT

This paper examines international differences in firms' cost of equity capital across 40 countries. We analyze whether the effectiveness of a country's legal institutions and securities regulation is systematically related to cross-country differences in the cost of equity capital. We employ several models to estimate firms' implied or ex ante cost of capital. Our results support the conclusion that firms from countries with more extensive disclosure requirements, stronger securities regulation and stricter enforcement mechanisms have a significantly lower cost of capital. We perform extensive sensitivity analyses to assess the potentially confounding influence of countries' long-run growth differences on our results. We also show that, consistent with theory, the cost of capital effects of strong legal institutions become substantially smaller and, in many cases, statistically insignificant as capital markets become globally more integrated.

JEL classification: G14, G15, G38, G30, K22, M41

Key Words: *Cost of equity, Disclosure regulation, Law and finance, International finance, Country risk, Legal system*

* We thank Yakov Amihud, Ray Ball, Phil Berger (the editor), Gauri Bhat, Mariassunta Giannetti, Wayne Guay, Cam Harvey, Ole-Kristian Hope, Bob Holthausen, Leora Klapper, Rafael La Porta, David Larcker, DJ Nanda, Felix Oberholzer, Shiva Rajgopal, Tjomme Rusticus, Terry Shevlin, René Stulz, Surjit Tinaikar, Joe Weber, Peter Wysocki, Stephen Young, two anonymous reviewers and workshop participants at Duke University, Harvard Business School, Rutgers University, Singapore Management University, University of Chicago, University of Washington, the UC Davis Corporate Governance conference, the 2004 European Accounting Association meeting, and the 2004 European Finance Association meeting for helpful comments on earlier drafts. We also thank Bryan Chao, Nick Vedder and Ian Weiliang for their research assistance. Luzi Hail acknowledges the financial support by the Research Commission of the University of Zurich. Analyst forecast data has been generously provided by I/B/E/S (Thomson Financial). Campbell Harvey, Ole-Kristian Hope and Leora Klapper have graciously provided supplemental data. The study is one of the recipients of the 2004 Geewax and Terker Company Prizes.

1. Introduction

In this paper, we examine international cost of equity capital differences across 40 countries. Specifically, we investigate whether the effectiveness of securities regulation and supporting legal institutions are systematically related to firms' cost of equity capital, over and above traditional proxies for firm and country risk. Our analysis exploits cross-sectional variation in securities regulation around the world using the La Porta, Lopez-de-Silanes, and Shleifer [2005] dataset on regulation mandating and enforcing disclosures in security offerings. Disclosure regulation is often justified by regulators with the argument that it reduces firms' cost of capital (e.g., Levitt [1998]). However, as noted in a survey by Healy and Palepu [2001], there is little empirical evidence on this link and, more generally, the costs and benefits of disclosure regulation.

Prior research demonstrates that legal institutions and securities regulation are associated with the development of equity markets (e.g., La Porta et al. [1997, 2005]). The basic idea is that well-functioning legal systems protect outside investors which in turn should improve firms' ability to raise external finance and to exploit growth opportunities. In particular, strong investor protection limits expropriation by insiders which should lead to less price protection on the part of outside investors. Consistent with these arguments, La Porta et al. [2002] provide evidence that firms in countries with stronger investor protection and more effective legal systems enjoy higher equity valuations. However, the mechanisms by which legal institutions affect firms' equity valuations are still unclear. It is possible that the valuation effects primarily reflect differences in the level of expropriation and firms' growth opportunities. But effective legal institutions may also reduce the risk premium demanded by investors and hence firms' cost of capital.

To explore one way in which legal institutions could affect equity valuations, we examine whether differences in countries' securities regulation explain international differences in cost of equity capital. Our analysis focuses on regulation mandating and enforcing disclosures because its relation with the cost of capital is probably better supported by extant theory than the link between the cost of capital and other institutions.[1] That said, even the relation between disclosure and the cost of capital is far from obvious.

It is straightforward to show that a credible commitment to disclosure reduces uncertainty and information asymmetries between the firm and its investors or among investors (e.g., Verrecchia [2001]). However, it is less clear whether differences in disclosure lead to differences in non-diversifiable risk and hence are reflected in the cost of capital. One possible explanation is that more disclosure reduces parameter uncertainty and estimation risk, parts of which can be non-diversifiable (Barry and Brown [1985], Coles, Loewenstein, and Suay [1995], Lambert, Leuz, and Verrecchia [2005]). In capital markets with incomplete information, disclosure can enhance investor recognition, thereby enlarge the investor base and improve risk sharing (Merton [1987]).[2] Lombardo and Pagano [2002a] argue that better disclosure reduces out-of-pocket monitoring costs borne by investors and hence the compensation they demand for holding equity. More recently, Lambert et al. [2005] show that, even in a CAPM world, improved disclosure regulation has the effect of decreasing firms' cost of capital by generally lowering the covariance between a firm's future cash flows with the future cash flows of the other firms in the economy.

While all these explanations predict that more disclosure is associated with a lower cost of capital, the empirical magnitude of these effects is still an open issue. It is possible that the hypothesized effects lead to minor differences in non-diversifiable risk or that they are largely

---

1 A rare example for a different link is Himmelberg, Hubbard, and Love [2004] providing an explanation for why weak investor protection may lead to a higher (marginal) cost of capital.

2 A similar risk-sharing argument is used to explain why market integration reduces firms' cost of capital (e.g., Lombardo and Pagano [2002a]).

captured by traditional proxies for risk. Moreover, the effects of country-specific factors, including disclosure regulation, likely decrease as capital markets become integrated and the opportunities for risk sharing and diversification expand (e.g., Harvey [1991]). Thus, the effect of disclosure regulation on the cost of capital is ultimately an empirical question.

We compute estimates for firms' cost of equity capital from 1992 to 2001 and across 40 countries. Our primary analysis is based on four models suggested in the literature to obtain estimates for the cost of capital implied in share prices and analyst forecasts.[3] Based on these estimates, we document statistically and economically significant differences in the cost of equity capital across countries. We show that a substantial portion of this cross-sectional variation is explained by traditional proxies for firm risk, i.e., size, volatility and the book-to-market ratio, as well as country factors capturing differences in inflation and macroeconomic variability. Together, these variables explain about 60% of the country-level variation and over 35% of the firm-level variation in the cost of equity capital around the world.

Going beyond a descriptive analysis, we investigate whether and how effective securities regulation is related to firms' cost of capital. We introduce proxies for the level of disclosure and securities regulation as well as the overall quality of the legal system into our country-level regressions. The results indicate that these institutional proxies are significantly related to international differences in the cost of equity capital over and above traditional proxies for firm and country risk. Firms in countries with more extensive disclosure requirements, stronger securities regulation and more effective legal systems have a significantly lower cost of capital.

---

[3] See Claus and Thomas [2001], Gebhardt, Lee, and Swaminathan [2001], Easton [2004], and Ohlson and Juettner-Nauroth [2005]. In sensitivity analyses, we also use the regression-based approaches in Easton et al. [2002] and Easton [2004].

We perform extensive analyses to assess the potentially confounding influence of growth differences across countries on our findings. Such differences can arise from differential real economic growth or from different accounting rules and their impact on the long-run growth rate in accounting-based valuation models. We address this issue by gauging the sensitivity of our findings to different assumptions about long-run growth, by simultaneously estimating the implied cost of capital and implied growth rate based on Easton et al. [2002] and Easton [2004], by introducing growth proxies as control variables and by providing results using non-accounting-based cost of capital estimates, i.e., dividend yields and expected stock returns derived from country credit-risk ratings as in Erb, Harvey, and Viskanta [1996a]. These additional analyses strongly corroborate our findings, with the exception of the results using estimates based on Easton et al. [2002] and Easton [2004] where we obtain consistent and significant findings for our base model of controls but lose significance for our most conservative specification. But even then the coefficients on disclosure regulation are directionally consistent suggesting that statistical power is the main issue. In addition, we perform numerous sensitivity analyses to mitigate concerns about correlated omitted variables, including differences in equity market size, tax rates, or analyst forecast properties. We also extract country-fixed effects from firm-level regressions controlling for traditional risk factors and show that the institutional proxies are related to the fixed effects in the predicted fashion.

Finally, we investigate to what extent the effect of securities regulation differs by capital market integration. Using several integration proxies, we find that, as expected, the effect of regulation is decreasing in capital market integration. In segmented markets, the estimated effect on the cost of capital is about 200 basis points going from the 25^{th} to the 75^{th} percentile of the securities regulation variable that combines disclosure rules and associated enforcement. Among countries with integrated capital markets, extensive disclosure rules continue to be negatively associated with the cost of capital, but the effect becomes much smaller (about 60 basis points over the inter-quartile range). Once we combine disclosure requirements and related enforcement, the effect of securities regulation on the cost of capital is significant only in segmented capital markets. Overall, these findings are consistent with the general notion that there are local factors under segmentation and that these factors become less important with integration (e.g., Bekaert and Harvey [1995], Karolyi and Stulz [2003]).

Our study builds on recent advances in the finance and accounting literature on the role of legal institutions (La Porta et al. [1997, 2000a], Leuz, Nanda, and Wysocki [2003], Bushman, Piotroski, and Smith [2004]). We extend this literature by presenting evidence that securities regulation and, in particular, disclosure requirements are associated with international differences in the cost of equity capital. Specifically, our paper complements the studies by La Porta et al. [2005] showing that securities regulation explains cross-sectional variation in equity market development, by La Porta et al. [2002] demonstrating that strong shareholder rights and legal systems are associated with higher equity valuations using Tobin's Q, and by Lee and Ng [2002] suggesting that firms in corrupt countries trade at lower multiples. Our study also relates to Bhattacharya and Daouk [2002] demonstrating that enforcement of insider trading regulations lowers firms' cost of capital.

Our paper is most closely related to a study by Lombardo and Pagano [2002b]. They document that proxies for the quality of the legal system (e.g., judicial efficiency) are *positively* associated with returns on equity using realized stock returns, dividend yields, and earnings-to-price ratios. These proxies, however, capture not only differences in firms' cost of capital but may also reflect shocks to firms' growth opportunities and differences in expected growth rates (e.g., Bekaert and Harvey [2000], Hail and Leuz [2005]). This may explain the contrast to our findings, which are based on proxies that make an explicit attempt to control for growth differences in estimating the cost of capital.

We also contribute to a strand of literature examining the link between disclosure and the cost of capital. Prior evidence suggests that firms providing more disclosures have a lower cost of capital (e.g., Botosan [1997], Hail [2002], Francis et al. [2004]). However, as these studies are based on firm-level data within one country, they have to rely on voluntary disclosures, which may reflect self-serving choices, rather than a commitment to disclosure (Leuz and Verrecchia [2000], Verrecchia [2001]).[4] In contrast, securities regulation mandates certain disclosures at the country (or exchange) level and hence constitutes a commitment if the rules are properly enforced. Moreover, as Lambert et al. [2005] argue, the effects of firm-specific disclosures on the cost of capital are more ambiguous than the economy-wide effects of mandated disclosures. Thus, cross-country settings provide a promising way to explore the link between disclosure and the cost of capital.[5]

Our work further contributes to the international finance literature. The study is the first to analyze international cost of capital differences using analyst forecasts for a large set of countries around the world.[6] Prior studies are generally based on realized stock returns and find that the explanatory power of the international CAPM is fairly low. This result is often attributed to market segmentation (e.g., Harvey, [1995]). The issue can be addressed by jointly estimating the

4 Consistent with this concern, Miller [2002] shows that time-series variation in discretionary disclosures reflects earnings performance. That is, firms provide more disclosures following good performance and reverse them after performance declines.

5 An alternative approach is to study major changes in the securities regulation of a country. However, such changes are rare. Benston [1973], Greenstone, Oyer, and Vissing-Jorgensen [2005], and Bushee and Leuz [2005] study changes in U.S. securities regulation in 1934, 1964 and 1999, respectively, but none of them examines cost of capital effects.

6 In concurrent studies, Chen, Chen, and Wei [2003], Chen, Jorgensen, and Yoo [2004], and Lee, Ng, and Swaminathan [2004] analyze differences in the implied cost of capital for firms in 9 Asian countries and the G7 countries, respectively.

ex ante cost of capital and the degree of market integration (Bekaert and Harvey [1995]). However, standard techniques to obtain unbiased estimates of expected returns from realized stock returns require fairly long time-series to wash out the effects of shocks to firms' growth opportunities (e.g., Elton [1999], Stulz [1999]). Bekaert and Harvey [2000] show in Monte Carlo simulations that average *realized* returns often increase when a shock decreases the cost of capital in a population, and that returns perform worse than dividend yields. Our study builds on models that estimate an *ex ante* return required by investors using market prices and analyst forecasts. This alternative approach allows us to present evidence from 40 countries. Aside from being novel at the descriptive level, our findings complement prior return-based studies in providing further evidence on the determinants of international cost of capital differences and on the mitigating role of capital market integration.

Finally, we complement recent work on implied cost of capital models. Our finding that implied cost of capital models produce estimates that are highly associated with traditional proxies for firm and country risks across 40 countries extends recent studies on the associations of implied cost of capital estimates based on U.S. firms (Gode and Mohanram [2003], Botosan and Plumlee [2005]) and for the G7 countries (Chen et al. [2004], Lee et al. [2004]). As several prior studies indicate that implied cost of capital estimates can be unreliable when based on low quality analyst forecasts (e.g., Easton and Monahan [2005], Guay, Kothari, and Shu [2005]), we show that we obtain similar findings using cost of capital proxies that do not rely on analyst forecasts. While these results are comforting, we hasten to add that the key constructs in our analysis are notoriously difficult to measure and that our findings should be interpreted carefully.

The paper is organized as follows. Section 2 describes the sample and the construction of the implied cost of capital estimates. In Section 3, we present our base results relating traditional proxies for firm and country risks to firms' cost of capital. In Sections 4 and 5, we present the empirical evidence on the role of institutional factors and report extensive sensitivity analyses and the effects of capital market integration. Section 6 concludes the study.

2. *Data and Research Design*

2.1 SAMPLE SELECTION AND COST OF CAPITAL ESTIMATION

To compute the cost of capital proxies, we obtain financial data from *Worldscope* and analyst forecasts and share price information from *I/B/E/S*. We download all firms contained in *Worldscope* from 1992 to 2001 and match them to firms covered in *I/B/E/S*. As described in more detail in the Appendix, we require each observation to have one-year-ahead and two-year-ahead, non-negative earnings forecasts, either a long-term growth forecast or a three-year-ahead earnings forecast, and a contemporaneous share price in order to be included in the cost of capital computation. All items are denominated in local currency. Financial data are measured as of the fiscal-year end and analyst forecasts and stock prices are measured as of month +10. We deliberately choose to compute our estimates ten months after the fiscal-year end to assure that financial data are publicly available and priced at the time of our computations.[7] These data requirements result in a sample of 36,202 firm-years. We eliminate firm-years (a) if there are less than five observations for the country in that year, (b) if the inflation rate for the country in that year is above 25% and (c) if we do not have institutional data for the country.[8] The final sample consists of 35,118 firm-year observations from 40 countries between 1992 and 2001.

---

7 To account for the fact that some of the input data are taken as of the fiscal-year end (e.g., book values) whereas prices and forecasts stem from month +10, we discount the price at month +10 back to the beginning of the fiscal year using the imputed cost of capital. See Appendix A.2 for further details. We repeat our analyses using data as of month +7. In this case, we have a slightly smaller sample, i.e., 339 country-years, but the results are very similar and the inferences and conclusions are the same.

8 The last criterion is obviously necessary for our institutional analysis. The two other criteria are imposed to prevent extreme observations from unduly affecting our country-year regressions. However, the results do not hinge on these criteria and are very similar without them. Moreover, we repeat our analyses requiring a

For each firm-year observation, we compute the ex ante cost of capital implied in contemporaneous stock price and analyst forecast data. In our primary analyses, we use four different models suggested in Claus and Thomas [2001], Gebhardt, Lee, and Swaminathan [2001], and Ohlson and Juettner-Nauroth [2005] as implemented by Gode and Mohanram [2003], and in Easton [2004]. The first two are special cases of the residual income valuation model described by Ohlson [1995], while the latter two are based on the abnormal earnings growth valuation model developed by Ohlson and Juettner-Nauroth [2005]. The basic idea of all four models is to substitute price and analyst forecasts into a valuation equation and to back out the cost of capital as the internal rate of return that equates current stock price and the expected future sequence of residual incomes or abnormal earnings. The individual models differ with respect to the use of analyst forecast data, the assumptions regarding short-term and long-term growth, the explicit forecasting horizon, and whether and how inflation is incorporated into the steady-state terminal value. We describe the models' key assumptions and provide more details on data requirements and implementation choices in the Appendix.

A primary concern about cost of capital comparisons across countries is that the estimates reflect countries' growth differences. In using analyst forecasts, implied cost of capital models make an explicit attempt to capture (short-term) growth differences, which is why we use them for our analysis. However, the underlying valuation models have to make assumptions about firm growth beyond the explicit forecast horizon and, as a consequence, the estimates could be quite sensitive to these assumptions about long-run growth (e.g., Easton et al. [2002]). Furthermore, implied cost of capital models are based on earnings (rather than cash flow) forecasts creating the

---

minimum of 10 (20) observations per country and year. Imposing these restrictions eliminates several smaller countries and reduces the sample to 333 (288) country-year observations, but provides very similar results and does not change our inferences.

possibility that accounting differences bias our cost of capital estimates via their influence on growth beyond the forecast horizon. For instance, a more conservative accounting system implies that, generally, a smaller fraction of firm value is captured during the explicit forecast horizon. Thus, firms from countries with more conservative accounting systems should exhibit higher growth rates in residual income or abnormal earnings beyond the explicit forecast horizon as accounting earnings have to "catch up" with economic earnings. We address these concerns in Section 4.2 by presenting analyses that assess the influence of growth differences on our results.

In addition, there are serious concerns about measurement error in our implied cost of capital proxies, in particular when there are systematic deficiencies in analysts' forecasting behavior. Easton and Monahan [2005] use a variance decomposition approach to evaluate several implied cost of capital measures. They conclude that the absolute magnitude of noise is generally large and that even aggregating across firms or an instrumental variables approach may not help much.[9] Guay et al. [2005] show that implied cost of capital models may perform poorly in cross-sectional return regressions and that these problems are partly attributable to inaccurate and sluggish analyst forecasts. We attempt to mitigate these measurement concerns in three ways. First, following Gode and Mohanram [2003], and Botosan and Plumlee [2005], we show that our cost of capital estimates are systematically related to various traditional risk and country factors. Second, we subject our findings to multiple robustness checks to address concerns about systematic deficiencies in analyst forecasting behavior. Specifically, we lag stock price by three months relative to the forecast measurement date, as suggested by Guay et al. [2005]. We also

[9] Easton and Monahan [2005] show that a naïve proxy such as the price-to-forward-earnings ratio may perform similar to, if not better than, estimates from more complex valuation models. However, this finding is unlikely to extend to international settings where accounting for the vast differences in firms' growth opportunities is probably beneficial. When we use the suggested metric in our analyses, the results are similar, although at lower levels of statistical significance. As expected, the results become stronger once we control for differences in forecasted growth.

compute accuracy-weighted country-year means of our cost of capital estimates, giving more weight to observations with higher forecast accuracy and reducing the influence of estimates with noisy inputs. As a final check, we conduct sensitivity analyses using proxies for the cost of capital that do not rely on analyst forecast data, namely dividend yields and expected returns derived from country credit-risk ratings and country-index returns.

Our analyses are based on country-year medians of the cost of capital estimates.[10] We choose to do so for several reasons. First, the basic idea of this study is to exploit variation in securities regulation, which varies at the country and not the firm level. Second, in using multiple observations per country, we can use time-series variation in traditional risk factors to tease out the effects of regulation on the cost of capital. Finally, a country-year analysis does not give undue weight to large countries with many firm-year observations. An alternative approach that also avoids these issues, but nevertheless exploits firm-level information is to analyze country-fixed effects extracted from firm-level regressions. We perform this analysis in Section 4.3.

Table 1 reports descriptive statistics for the implied cost of capital estimates. Panel A provides descriptive information on the distribution of the four models' estimates. It shows that the Gebhardt et al. model generally yields the lowest estimates. The three other models are fairly close and on average around 13.5 percent. All four models provide estimates that are within reasonable ranges. Panel B reports the correlation coefficients and shows that all four estimates are highly correlated. The Gebhardt et al. estimates exhibit the lowest correlations, presumably because they have the longest explicit forecasting horizon (12 years) and incorporate industry information, both of which are distinct features of the Gebhardt et al. model.

---

10 We repeat our analyses using country-year *means* for all variables after deleting extreme observations in the 1st and the 99th percentile. This procedure yields a sample of 356 country-year observations, for which we obtain very similar results and the same inferences as those reported.

In subsequent analyses we report results using the average over the four proxies. Panels A and B provide descriptive statistics and correlation coefficients for the mean cost of capital, r_{AVG}. We also perform (but do not tabulate) analyses using estimates from the four individual models or, alternatively, the first principal component of the four individual estimates. The results of our main analyses are consistent across all four individual models and similar to those reported, albeit at slightly lower significance levels for some models. The inferences, however, remain virtually unchanged. If we use the first principal component, the results are even stronger than those reported. These findings are consistent with the notion that aggregating across the four models reduces some measurement error.

Panel C of Table 1 reports for each sample country the number of firm-year observations used to compute the cost of capital estimates, the number of country-year observations available for our analyses as well as descriptive information on the institutional variables used in subsequent analyses. We also provide the time-series average cost of capital by country. The panel shows considerable variation in the cost of capital across countries. However, simple comparisons across countries can be misleading because they do not control for various factors known to affect firms' cost of capital. We therefore control for a number of risk and country factors before introducing the institutional variables of interest.

2.2 RISK AND COUNTRY CONTROL VARIABLES

An important factor is the inflation rate. Analyst forecasts are expressed in nominal terms and local currency, which implies that the resulting estimates for the cost of capital reflect countries' expected inflation rates. This effect explains, for instance, the relatively low cost of capital estimate for Japan, which experienced deflation over parts of the sample period. Thus, it is important to control for international differences in expected inflation rates.

One approach is to subtract the expected future inflation rates from the cost of capital estimates and to conduct a regression analysis on the resulting inflation-adjusted estimates.[11] However, this approach essentially forces a coefficient of minus one on the inflation proxy. As the market's expectation for future inflation is only imperfectly observable, we prefer to introduce a separate control variable for cross-sectional differences in inflation. This approach lets the data determine the relation between the inflation proxy and the cost of capital estimate. We expect the coefficient to be positive but smaller than one, as measurement error likely biases the coefficient towards zero. We compute monthly inflation rates for each country using the consumer price indices provided in the *Datastream* and *Worldbank* databases and use the median of next year's (annualized) monthly inflation rates as a proxy for the expected future inflation.

Another factor is time-series variation in the risk-free interest rates. It is common in international studies to convert local returns into U.S. dollar returns and use the U.S. Treasury bill as a proxy for the risk-free rate in all countries (e.g., Harvey [1995]). This approach essentially assumes that exchange rates reflect inflation differences and that time preferences and real rates are similar across countries. Thus, expressing returns in excess of the T-bill rate controls for time-series variation in the risk-free rate. In our country-year analysis, the T-bill rate is a yearly constant and hence year-fixed effects control for time-series variation in the risk-free rate.[12]

Next, we introduce a number of controls for risk. Based on prior empirical studies on the cross-sectional determinants of returns, we expect the cost of capital to be negatively associated with firm size and to be positively associated with stock return volatility (or beta) and the book-to-market ratio (e.g., Fama and French [1992, 1993]). We measure size as the firm's market

---

[11] We report this approach as a sensitivity check in Section 4.3.

[12] It is of course possible that there are differences in the real (risk-free) rates across countries. We address this concern in Section 4.3 and show that such differences do not unduly affect our results.

capitalization as of the fiscal year end, return variability as the standard deviation of monthly stock returns over the last twelve months and the book-to-market ratio as the ratio of book value to market value of equity at the end of the fiscal year.

The Fama and French three-factor model loosely motivates these three variables. We use return variability rather than beta factors for two reasons. First, the estimation of beta presupposes a stance on the degree of capital market integration. If capital markets are integrated, it is appropriate to use the world market portfolio (e.g., Solnik [1974], Stulz [1981]). But we do not know to what extent our sample markets are integrated. In fact, one reason for using the implied cost of capital approach is that it does not require a choice of a market portfolio and avoids one of the difficulties return-based studies face in an international context. Second, prior studies find that future returns in emerging markets exhibit no or even a negative relation with beta factors computed with respect to the world market portfolio (e.g., Harvey [1995], Erb et al. [1996a]).

We also note that there is some debate about the inclusion of the book-to-market ratio in implied cost of capital regressions (e.g., Gode and Mohanram [2003]). But even setting its empirical relevance in asset pricing models aside, we think it is prudent to control for book-to-market differences in our study. As explained in more detail in Section 4.2, book-to-market ratios capture differences in firms' growth opportunities (e.g., La Porta et al. [2002]) as well as differences in the accounting rules (e.g., Joos and Lang [1994]). Thus, we include the book-to-market ratio in our models, as a way to control for these differences. This approach is conservative because market-based controls may absorb the effects of the institutional variables, if better institutions manifest in a lower cost of capital and hence higher valuations.[13]

In addition to the three proxies for firm risk, we include industry controls in all our regressions. Fama and French [1997] find that there is substantial variation in factor loadings across industries. To construct our industry controls we use the industry classification in Campbell [1996] and compute the percentage of firms in each of the 12 industry classes by country and year.[14] Furthermore, it makes sense to control for differences in macroeconomic variability and a country's exposure to global economic risks and shocks (Ferson and Harvey [1998]). Similarly, differences in risk-sharing opportunities across countries can affect corporate investment, e.g., lead to more or less specialization and risk taking, which in turn can affect the variability of aggregate output or performance (Obstfeld [1994]).

To capture cross-country differences in macroeconomic variability, we consider four variables: (1) the standard deviation of annual earnings per share over the last five years scaled by total assets per share, (2) the standard deviation of accounting returns on equity over the last five years, (3) the standard deviation of the residuals from a regression of annual GDP growth rates on a time index over the sampling period, and (4) the coefficient of variation of yearly average exchange rates (US$ to local currency) over the sampling period.[15]

---

[13] Consistent with this conjecture, the coefficients for the institutional variables tend to be larger if we do not include the book-to-market ratio. See Section 4.3 for results using accounting-based controls only.

[14] Our results remain qualitatively unchanged (and in some cases become even stronger) when we use one-digit SIC codes to calculate the industry controls. Note further that dropping the industry controls does not materially alter our results.

[15] As noted for the market-based controls, macroeconomic proxies may reflect the effects of legal institutions. For instance, good legal institutions may manifest in lower GDP volatility. In general, such effects should make it harder for us to find significant relations for the institutional variables in the presence of macroeconomic controls. We therefore also present results without macroeconomic variability in Table 4.

The first two variables are based on firm-level data, but like all other variables they are measured as country-year medians. Hence, they are likely to capture the country-level or macroeconomic variability. Consistent with this claim, we find that all four variables are highly correlated with each other. Factor analysis shows that there is only one factor in the data with an eigenvalue above one. We therefore summarize the four variables into a single control variable for macroeconomic variability using the first principal component.

The final control variable captures differences in forecast bias, which is not a risk factor. But as the cost of capital estimates rely on analyst forecasts, we are concerned that international differences in the forecasting behavior could "mechanically" affect our results. For instance, if forecasts in a particular country tend to be optimistic but market participants understand this bias and properly adjust prices, implied cost of capital models yield upwardly biased estimates (see also Botosan and Plumlee [2005]). Hope [2003] shows that forecast accuracy differs significantly across countries and that forecast accuracy is related to firms' disclosure policies. We compute forecast bias at the country-year level using firm-level forecast errors. We define the forecast error as the mean one-year-ahead consensus forecast minus the actual earnings reported in I/B/E/S. Thus, if forecasts tend to be optimistic, the forecast bias variable assumes positive values. We expect a positive coefficient if markets back out the bias.

Table 2 presents summary statistics and correlation coefficients for the control variables described in this section. There is considerable cross-sectional variation in all variables and most variables are significantly correlated. The bottom row of Panel B also reports correlation coefficients for the average cost of capital and the control variables. The average cost of capital displays the predicted correlations with all control variables.

3. *Implied Cost of Capital Estimates and Controls for Firm and Country Risk*

In this section, we establish that cross-sectional differences in the implied cost of capital are systematically related to firm and country proxies for risk in the predicted fashion. Similarly, Gode and Mohanram [2003], and Botosan and Plumlee [2005] validate implied cost of capital estimates by showing that they are related to proxies capturing various sources of risk. The regressions in Table 3 provide a benchmark for our institutional analysis.

The first model controls for inflation, firm size, beta, the book-to-market ratio, industry and year effects. As explained in Section 2, we introduce year-fixed effects to capture time-series variation in the risk-free rate. Table 3 shows that this model explains almost 58% of the cross-sectional variation in the implied cost of capital across 40 countries. All control variables are highly significant and have the predicted sign. As expected, the coefficient on inflation is smaller than one. This finding probably reflects measurement error in the proxy for future inflation.[16]

In the next model, we replace the beta factor with return variability. Although beta is positively associated with the implied cost of capital, its association is weaker than the association of return variability. Moreover, we find that beta is insignificant once we include all other control variables, whereas return variability is not. Thus, to be conservative, we use return variability in our regressions. The results for the institutional variables are even stronger if we use the beta factor, and very similar to those reported in Sections 4 and 5 if we include both variables together.

As capital structures differ across countries and have a predictable effect on the cost of equity capital, we also check whether adding financial leverage as control alters our findings. It does

[16] Illustrating the measurement issue, the contemporaneous inflation rate yields a smaller coefficient and lower R^2 than the one-year-ahead realized inflation rate, which we use to proxy for the expected inflation. Note, however, that our results are not affected by this choice.

not. We find (but do not tabulate) that leverage is insignificant if either the beta factor or return variability are used in the model, suggesting that these variables sufficiently control for capital structure differences across countries.

In the third (or full) model, we add controls for macroeconomic variability and differences in forecast bias across countries. As described in Section 2, MACVAR represents the first principal component of four proxies: earnings per share variability, ROE variability, volatility in GDP growth and exchange rate variability. At the country-year level, the four proxies are highly correlated, which is why we summarize them. However, the results are not materially affected if we use any one of them as our proxy instead of the principal component of all four variables. In all specifications, the coefficient on the proxy for macroeconomic variability is positive and significant. While Table 2, Panel B, shows that forecast bias is positively correlated with our estimates for the cost of capital, the coefficient on forecast bias is not significant in the regression. Thus, our country-year estimates do not appear to be significantly biased due to differences in forecasting behavior across countries.[17] Overall, the full model explains 60% of the country-level variation in the implied cost of equity capital around the world.

The last column in Table 3 (Model 4) presents the full model using firm-level instead of country-level controls. We therefore replace the macroeconomic variability with earnings variability at the firm-level. All variables exhibit the expected associations and are highly significant. The results are very similar to those in the country-level regressions. The only notable difference is that the coefficient on forecast bias is now significant. Using firm-level controls, our model still explains 36% of the variation in the implied cost of capital.

[17] Our results in Table 3 are very similar if we use forecast accuracy, i.e., the absolute value of the one-year head or two-year ahead forecast error, as control variable. Similarly, lagging price by three months to account for sluggish analyst forecasts, as suggested in Guay et al. [2005], does not materially alter our findings.

4. *The Role of Securities Regulation and Legal Institutions*

4.1 CROSS-SECTIONAL EFFECTS ON THE COST OF EQUITY CAPITAL

In this section, we examine whether differences in countries' securities regulation mandating and enforcing disclosures can explain international differences in cost of equity capital, over and above those factors previously introduced into the model.

One potential role of securities regulation is to serve as a commitment device. Disclosures reduce information asymmetries between the firm and its investors as well as among investors, but only if they are credible and not self-serving (e.g., Verrecchia [2001]). Without commitment, firms may have incentives to withhold or manipulate information in certain situations, e.g., when performance is poor. Effective securities regulation binds firms to provide disclosures in good and bad times, which reduces information asymmetries and increases liquidity in secondary markets (e.g., Bushee and Leuz [2005]). However, it is less clear whether differences in securities regulation also manifest in differences in non-diversifiable risk and hence the cost of capital.

To support such a relation, we can draw on several theoretical models suggesting a link between disclosure and the cost of capital. As already discussed, these explanations are based on the idea of estimation risk, investor recognition, and out-of-pocket monitoring costs (see Barry and Brown [1985], Merton [1987], Lombardo and Pagano [2002a], respectively). More recently, Hughes, Liu, and Liu [2005], and Lambert et al. [2005] show that information quality manifests itself in the market risk premium and firms' cost of capital, respectively, even if the economy becomes large and investors hold diversified portfolios.

However, the empirical relevance of these explanations, especially for firm-level disclosures, is not obvious: estimation risk is in part diversifiable and may be captured by traditional risk proxies (Clarkson, Guedes, and Thompson [1996], Lambert et al. [2005]), investor-base effects are susceptible to arbitrage (Merton [1987], Easley and O'Hara [2004]), and monitoring costs can be reduced by information intermediation. Thus, while a negative relation between disclosure and the cost of capital can be supported by various theories, the link is far from obvious and still debated.

To explore these issues, we analyze variables that capture cross-country differences in securities regulation, i.e., disclosure rules and supporting enforcement institutions. We build on a recent study by La Porta et al. [2005] analyzing the role of securities regulation for financial market development. Based on answers to an extensive questionnaire distributed to security-law attorneys in 49 countries, La Porta et al. [2005] construct a series of quantitative metrics capturing the current status of rules and regulations governing security issuance. Each score ranges from zero to one with higher values indicating more extensive requirements or stricter enforcement. The database is constructed as of December 2000 and provides three main indices: (1) the disclosure requirements index capturing several aspects of prospectus disclosure in security offerings, (2) the liability standard index capturing the procedural difficulties in recovering losses from the issuer and its directors in a civil liability case, and (3) the public enforcement index capturing market supervision by a regulator and its investigative powers and sanctions.

We create two constructs from these indices to measure international differences in disclosure and securities regulation. Given the motivation of our study, we focus on the disclosure requirements index, DISREQ, as our primary proxy for disclosure regulation. It consists of the arithmetic mean of several sub-indices scoring disclosure requirements at the country's largest stock exchange in the areas of prospectus requirements, directors' compensation, ownership structure and inside ownership, related-party transactions and contracts. We refrain from using variables that capture disclosure *practice*, such as the CIFAR index, because our study focuses on the role of *legal* institutions and disclosure *regulation*. Variables such as the CIFAR index also capture voluntary disclosures by firms, which makes it harder to attribute the estimated effects to securities regulation. We revisit this issue in our sensitivity analyses in Section 4.3.

Prior research suggests that rules alone are unlikely to be effective without proper enforcement (e.g., Bhattacharya and Daouk [2002], Berkowitz, Pistor, and Richard [2003]). As our first variable is rules-based, we also create a second institutional variable, SECREG, which captures the effectiveness of a country's securities regulation by combining the disclosure requirements index with the liability standard and the public enforcement indices. We construct this variable by computing the arithmetic mean of the three La Porta et al. [2005] indices. We believe that combining all indices in this fashion to construct a comprehensive measure of securities regulation is appropriate for two reasons. First, countries with effective securities regulation are likely to have both proper rules and supporting enforcement mechanisms in place. This logic implies that the three securities regulation indices in La Porta et al. [2005] should exhibit a relatively high correlation and that separately adding the indices to the model would not properly capture the complementary nature of the underlying constructs. Second, the various enforcement mechanisms considered by La Porta et al. [2005] could be substitutes, at least to some degree. In principle, countries may be able to choose different combinations of institutions to enforce their securities laws with similar outcomes. By aggregating the enforcement indices into a comprehensive measure we allow for substitution among the various enforcement mechanisms.

We also use the rule of law index from La Porta et al. [1997], LAW, as a proxy for the overall quality of a country's legal system. It has been extensively used in the literature as a variable indicating how well a country's legal system works.[18] In most of our analyses, we simultaneously include disclosure regulation (DISREQ) or securities regulation (SECREG) together with the rule of law variable (LAW). Although the latter two variables overlap with respect to enforcement aspects, we believe that the two constructs are sufficiently different in nature to warrant a separate treatment and that our specifications allow us to estimate the effect of securities regulation over and above the effect of the quality of a country's legal system.

To check whether these research design choices are reasonable, we conduct a factor analysis using the disclosure requirements, the liability standard, the public enforcement and the rule of law indices. We find that there are two principal factors in the institutional data. The first factor exhibits high loadings (0.52-0.60) for the three securities regulation indices and a low and negative loading with LAW (-0.11). These factor loadings indicate that the three variables capture one construct, presumably the overall effectiveness of a country's securities regulation, and hence support our construction of SECREG. The second factor displays high loadings (0.96) with rule of law and low loadings with all three securities regulation variables. We view this factor as capturing the overall quality of the legal system.[19] Thus, the factor structure of the institutional data supports our research design choices.

The first five models in Table 4 present results using the full set of controls from Model 3 in Table 3. The results are very similar and typically stronger than those reported in the table, if we

---

18 La Porta et al. [1997] provide several other variables capturing the effectiveness of the legal system. They are all highly correlated with the rule of law variable. Aggregating them into a legal quality variable yields similar results. See also Berkowitz et al. [2003].

19 We obtain very similar results to those reported in Table 4 when we use the first two principal components of the four institutional variables instead of the raw institutional scores.

use only the controls from Models 1 or 2 in Table 3, which is consistent with the notion that our full model is a conservative specification. As the dependent variable likely exhibits serial correlation, i.e., a country with a below-average cost of capital in one year is likely to have a lower-than-average cost of capital in other years, we report t-statistics based on Newey and West [1987] corrected standard errors, which are also robust to heteroscedasticity. When investigating the institutional factors separately, we find that more extensive disclosure requirements are negatively and significantly associated with the level of firms' cost of equity capital. The effect is roughly 90 basis points, going from the 25th to the 75th percentile of the disclosure requirements index. Similarly, stricter securities regulation exhibits a negative and significant association with firms' cost of capital. The effect is also economically significant and approximately 60 basis points, going from the 25th and the 75th percentile of the securities regulation index.[20] The rule of law variable also exhibits a negative coefficient. However, in the full model, the effect is not significant at conventional levels. Thus, the effect is strongest for DISREQ, which also has the strongest theoretical basis.

As mentioned above, we extend our analysis by including DISREQ or SECREG in conjunction with the LAW variable. In this model, both securities regulation variables continue to be negatively and significantly associated with the cost of capital. In fact, the coefficients slightly increase once we control for LAW. The quality of the legal system also exhibits a negative coefficient, which is significant at the 5% level in combination with DISREQ and is marginally significant together with SECREG (two-sided p-value = 0.16). These findings suggest that both securities regulation as well as the quality of the legal system have an effect on firms'

---

[20] The marginal effects (in %) are computed as (estimated coefficient*interquartile range*100). The interquartile ranges for DISREQ, SECREG and LAW are 0.333, 0.197 and 0.382, respectively.

cost of capital. The slightly weaker result for LAW in the regression using SECREG could reflect the fact that both SECREG and LAW capture enforcement aspects. Another explanation is that, as noted in Section 2, the full model captures some of the effects of strong legal institutions through the control variables, such as historical stock price volatility or macroeconomic variability. Consistent with this notion, columns 6 and 7 in Table 4 report that we obtain stronger and significant results for all three institutional variables using only the controls from our base model (Model 1 in Table 3).

Finally, in an attempt to address concerns about measurement error in our implied cost of capital estimates, we use the accuracy-weighted, country-year average cost of capital as dependent variable in the regressions. The idea is to give more weight to observations with higher forecast accuracy and reduce the influence of estimates that are likely to be noisy because of noisy inputs. Studies evaluating the quality of implied cost of capital estimates show that inaccurate analyst forecasts can lead to poor implied cost of capital estimates (e.g., Easton and Monahan [2005], Guay et al. [2005]). We compute the weighted country-year average cost of capital, r_{WEIGHT}, weighing each firm-year estimate with analyst forecast accuracy, which is measured by the absolute one-year-ahead forecast error scaled by forecast-period stock price. We report the regressions using accuracy-weighted estimates in the last two columns of Table 4. The findings show that both the coefficient magnitudes and the significance levels for the institutional variables increase, suggesting that accuracy weighing helps in reducing measurement error and that the negative associations for the institutional variables are not driven by inaccurate analyst forecasts. Taken together, the results in Table 4 provide strong support for the conclusion that both the effectiveness of securities regulation and the overall quality of a country's legal system are negatively associated with firms' cost of equity capital.

4.2 THE INFLUENCE OF GROWTH DIFFERENCES ACROSS COUNTRIES

An important concern about international cost of capital comparisons and our analyses in Table 4 is that growth differences across countries unduly influence the results. In using analyst forecasts, implied cost of capital models make an explicit attempt to capture market expectations about (short-term) growth, which is why we use these models for our analysis. However, the underlying valuation models have to make assumptions about growth in residual income or abnormal earnings beyond the explicit forecast horizon and, as a consequence, the estimates could be quite sensitive to these assumptions about long-run growth (e.g., Easton et al. [2002]). In this section, we examine the potentially confounding influence of growth differences on our findings in three ways: (1) we gauge the sensitivity of our results to different assumptions about long-run growth, (2) assess whether accounting differences bias our estimates via their influence on growth beyond the forecast horizon, and (3) provide results for cost of capital estimates that do not rely on accounting-based valuation models.

The first set of analyses assesses whether the results are sensitive to the growth assumptions in the terminal value computations. As explained in more detail in the Appendix, we use a proxy for countries' expected inflation rates to capture differences in nominal growth beyond the explicit forecast horizon. This assumption is similar in spirit to the one that is often used for U.S. firms (e.g., Gode and Mohanram [2003]). We repeat the analyses using two alternate terminal value computations.[21] First, we assume a constant 3% inflation rate for all countries. Second, we attempt to incorporate international differences in real growth into the terminal value computation. For instance, developing countries may exhibit abnormal earnings or positive

---

[21] Note that only estimates from two of the four models, r_{CT} and r_{OJ}, are affected by changing growth assumptions beyond the explicit forecasting horizon.

residual incomes for longer time periods while they "catch up" with the rest of the world.[22] To accommodate such differences, we set the perpetual growth rate equal to a country's annual change in real GDP plus its long-run inflation rate, where the latter is measured as the median inflation rate over our sample period. The second specification assumes that growth differences across countries persist in perpetuity whereas the first specification assumes that countries' growth rates converge to a competitive equilibrium with zero real growth (or 3% nominal growth) in perpetuity. In this sense, the two alternative terminal value computations represent opposite ends of the spectrum of perpetual growth assumptions. Our primary implementation (reported in the tables) uses country-specific inflation rates and falls somewhere in the middle.[23]

The cost of capital estimates from the first alternative specification, assuming uniform long-run growth, have similar mean and median values as the primary estimates but a slightly lower variance, as expected. The second alternative specification, assuming persistent differences in real growth, leads to an increase in the mean, median and variance of r_{AVG}, which is also not surprising. The new estimates from either specification are highly correlated ($\rho > 0.95$) with the primary estimates using country-specific inflation rates. This finding illustrates that our cost of capital estimates are not very sensitive to the terminal value assumptions about long-run growth.[24] In the first two columns of Table 5, we report our main regressions using the two alternative terminal value assumptions, respectively. We find that, regardless of the long-run growth assumption, the results are very similar to those presented in our main analysis (Table 4). In

22 In principle, ignoring differences in real (economic) growth biases the implied cost of capital estimates of high-growth countries downwards, which likely biases against our findings, considering that high-growth countries tend to be developing countries with weaker securities regulation.

23 Using country-specific inflation rates implies either that forecasted inflation differences persist or that countries with high inflation tame price increases in the future but in the process experience higher real growth rates than countries with low inflation. The latter assumption seems quite plausible.

24 This conclusion also holds if we restrict attention to estimates from the two models that are affected by the different growth assumptions, r_{CT} and r_{OJ}, rather than our average cost of capital estimate.

untabulated tests, we also confirm that explicitly introducing control variables for differences in economic growth into the full model does not materially affect our results (using countries' GDP growth rates, realized EPS growth rates and forecasted earnings growth rates as alternative proxies).

Our second set of analyses addresses the (related) concern that implied cost of capital models are based on earnings (rather than cash flow) forecasts and that, as a consequence, international accounting differences could affect our results. The issue is that most models are based on rather short forecasting horizons, so that accounting differences can give rise to growth differences beyond the forecast horizon (Easton et al. [2002]). For instance, a more conservative accounting system implies that, generally, a smaller fraction of firm value is captured during the explicit forecast horizon. Thus, firms from countries with more conservative accounting systems should exhibit higher growth rates in residual income or abnormal earnings after the explicit forecast horizon because accounting earnings have to "catch up" with economic earnings (Easton [2004]). Such differences likely bias our cost of capital estimates and could affect our findings, if international differences in conservatism are systematically associated with differences in disclosure regulation.[25]

We begin addressing this concern by pointing out that our regression models already include the book-to-market ratio, which among other things reflects differences in accounting conservatism.[26] That is, firms from countries with more (balance-sheet) conservatism typically

---

25 Failing to capture the effect of accounting conservatism on growth beyond the forecast horizon should downward (upward) bias the cost of capital estimates for countries with high (low) conservatism. We believe that this effect is likely to work against our findings considering that countries that are traditionally viewed as having a conservative accounting system, such as Germany, are not known for a high quality disclosure regime.

26 We further point out that estimates from abnormal earnings growth valuation models, such as r_{PEG} and r_{OJ}, are less likely to be affected by accounting differences than those from models using book values, such as r_{CT} and

exhibit lower book-to-market ratios (Joos and Lang [1994]). However, the book-to-market ratio also reflects a firm's investment opportunity set and hence may be a noisy proxy for accounting conservatism. We therefore use the accounting return on assets as an additional control variable because this ratio also reflects accounting differences across countries (Joos and Lang [1994]). As reported in Table 5, adding this control variable does not materially alter our findings.[27]

While these results mitigate concerns about the influence of accounting conservatism, they do not directly get at the issue of how accounting differences manifest in different growth rates *beyond the forecast horizon*. Easton et al. [2002] and Easton [2004] have developed implied cost of capital models that address this issue. These models use regressions to simultaneously estimate the cost of capital and the long-run growth rate in residual income or change in abnormal earnings growth for a portfolio of firms. We implement these models at the country-year level, requiring at least 10 observations per country and year for each regression. We find that the resulting portfolio-based cost of capital estimates (r_{ETSS} and r_{Easton}) are highly correlated with our cost of capital proxy, r_{AVG}, that is aggregated up from firm-level estimates ($\rho = 0.69$ for r_{ETSS} and $\rho = 0.53$ for r_{Easton}). Despite these correlations, the institutional variables are generally insignificant when we replace our dependent variable with the Easton et al. [2002] and Easton [2004] estimates. However, we find that these results are very sensitive to the inclusion of the simultaneously estimated implied growth rate. Including the respective implied growth rate as control (see Models 4 and 5 in Table 5), the coefficients on disclosure regulation are still

---

r_{GLS}. It is therefore reassuring that all four models individually produce similar results to those reported in Table 4. Chen et al. [2004] make a related point regarding violations of the clean surplus assumption.

27 We also refer to a recent study by Bushman and Piotroski [2005] on international accounting conservatism and use the difference between bad news and good news coefficients from reverse Basu [1997]-style regressions as a measure of (earnings) conservatism. Using this control, our results for DISREQ and LAW become stronger whereas the results for SECREG and LAW are slightly attenuated. The inferences, however, remain the same.

insignificant but fairly close to conventional levels (two-sided p-values = 0.13 and 0.14, respectively). Moreover, comparing the coefficients for both institutional variables of interest to those in Table 4, we note that the magnitudes are quite similar, especially when using the Easton [2004] estimates. We therefore explore whether our conservative set of full controls, which – as noted before – may capture some institutional effects, is responsible for these findings and, in particular, for the insignificance of LAW. Using less conservative specifications from Table 3, we find that generally disclosure regulation *and* legal quality exhibit *significantly* negative coefficients. Similar to Table 4, we report one of these specifications using the base model from Table 3. The coefficients on DISREQ and LAW are both significant and only slightly attenuate relative to the coefficients in Table 4 (compare Model 6 in Table 5 to column 6 in Table 4). Thus, we believe that the attenuated and in some cases insignificant findings for the Easton et al. [2002] or Easton [2004] approach are primarily a result of noise in the estimates, possibly caused by smaller country-year samples or the heterogeneous nature of our country-year portfolios.

An alternative way to exploit the portfolio-based approaches, but to avoid small samples is to estimate the (implied) growth rate beyond the forecast horizon at the country level and then to use this variable as an additional control in our Table 4 regressions. If our findings primarily reflect cross-country differences in economic growth beyond the forecast horizon and if the portfolio-based approaches capture these growth differences, then controlling for them should render the institutional coefficients insignificant, or at least substantially attenuate the results.[28] Based on this logic, we introduce country-specific controls for (implied) growth beyond the forecast

[28] We note that this procedure does not work, if the primary concern is the *model-specific* impact of accounting differences on growth beyond the forecast horizon. In this case, the proper growth rate is different (and potentially uncorrelated) across models. We thank the reviewer for pointing this out to us. One way to address this concern is to use cost of capital proxies that are not influenced by accounting differences, as we do below.

horizon into our regressions using estimates from Easton et al. [2002] or Easton [2004], respectively. We report one of these specifications in Table 5 (Model 7) and conclude that the results and inferences from these regressions are very similar to those presented in Section 4.1.

An entirely different way to address concerns about the effect of accounting differences is to use alternative proxies that do not rely on inverting an accounting-based valuation model. Such proxies can also provide evidence on the validity of our implied cost of capital estimates. We follow this approach in our third set of analyses. We focus on alternative cost of capital proxies that have been used in the finance literature and that do not rely on analyst forecasts, which also addresses concerns that our results are affected by noisy or biased analyst forecast behavior.

Dividend yields have been used extensively in prior studies in finance (e.g., Bekaert and Harvey [2000], Errunza and Miller [2000], Lombardo and Pagano [2002b]). Bekaert and Harvey [2000] argue and provide evidence from simulations that dividend yields are less affected by shocks to firms' growth opportunities than ex post stock returns. But they also note that dividend yields are far from perfect and are still likely to reflect differences in growth expectations. We find that the country-year median dividend yield (r_{DIV}), calculated as last year's actual dividends divided by stock price as of month +10, is highly and significantly correlated (ρ = 0.43) with our implied cost of capital estimates. The (partial) correlation coefficient increases if we control for differences in forecasted growth (ρ_{part} = 0.48), illustrating the fact that dividend yields reflect differences in growth expectations.

Erb, Harvey, and Viskanta [1996a, 1996b] suggest an alternative method to estimate expected returns using the *Institutional Investor's* country credit-risk ratings. The basic idea is that country credit-risk ratings reflect fundamental country risks, but are less likely than returns to reflect past shocks to countries' growth opportunities. Thus, credit-risk ratings can be used to purge ex post stock returns and to derive future expected returns by regressing ex post country index returns on the country credit-risk ratings. Erb et al. [1996b] demonstrate that this method produces reasonable cost of capital estimates at the country level. In a recent study, Harvey [2004] shows that country-credit risk ratings in fact contain information about future stock returns using trading simulations. Following Erb et al. [1996a], we estimate a regression of semi-annual stock returns using *Datastream* country indices for our sample countries on the natural log of semi-annual country credit-risk ratings (from *Institutional Investor*).[29] The annualized fitted values from this regression serve as proxy for future expected returns, r_{CRED}. We find that this measure is highly and significantly correlated with our implied cost of capital estimates (ρ = 0.64), despite the fact that they are derived from entirely different methods and input variables.

Next, we replicate our regressions from Table 4 using the two alternative non-accounting-based proxies as dependent variables. That is, we use the same control variables as before, except for the forecast bias variable which serves no purpose in these regressions. Following prior studies, we also include a control for earnings growth for the dividend yield regressions (e.g., Bekaert and Harvey [2000], Lombardo and Pagano [2002b]).[30] Consistent with our earlier results, we find that dividend yields are negatively associated with the strength of countries' disclosure requirements and securities regulation (Model 8 in Table 5). The coefficient on legal quality is positive and, using securities regulation (not tabulated), significant at the 5% level. This finding is opposite to our earlier results, but confirms Lombardo and Pagano [2002b], documenting a significantly positive coefficient for legal quality using dividend yields.[31]

---

[29] We thank Campbell Harvey for providing us with the necessary semi-annual country credit-risk ratings.

[30] We include earnings growth in the regression to be consistent with prior literature. However, the results do not depend on either the inclusion or the specific proxy of the earnings growth variable.

[31] In addition, we notice that dividend yields exhibit only weak associations with our controls for risk. We interpret this evidence as suggesting that dividend yields are a relatively poor country-level proxy for firms' cost of capital. Aside from the limitations mentioned above, a possible explanation is that dividend *payout* ratios differ across countries, reflecting among other things the level of investor protection (La Porta et al.

Finally, we use expected returns based on country credit-risk ratings (r_{CRED}) as dependent variable. We find that all three institutional variables of interest exhibit negative and highly significant coefficients, confirming our results in Table 4. In fact, the estimated effects of disclosure requirements and securities regulation are of similar magnitude as in Table 4, increasing our confidence in the implied cost of capital estimates. The coefficient on LAW increases substantially relative to Table 4 and exhibits very high t-values (see Model 9, Table 5). This finding is not surprising considering that, from a creditor's perspective, the ability to enforce claims via a country's legal system is of paramount importance and probably reflected in the country credit-risk ratings.

In sum, while growth differences are a major issue in international cost of capital comparisons, our findings are generally robust and the influence of growth differences, while present, does not appear to be the main driver of our institutional findings.

4.3 ADDITIONAL SENSITIVITY ANALYSES

In this section, we conduct a battery of additional sensitivity analyses. First, we address concerns about correlated omitted variables. Next, we check whether our results are robust to alternative specifications, susceptible to certain econometric problems, and sensitive to our sample composition. Finally, we estimate country-fixed effects and relate them to the institutional variables. Table 6 summarizes these sensitivity analyses. For brevity, we tabulate only the coefficients and t-statistics of the institutional variables.

[2000b]). Thus, in international settings, payout ratio effects could lead to differences in firms' dividend yields that do not reflect differences in the cost of capital.

Potential Correlated Omitted Variables

Prior work shows that per-capita GDP partially explains differences in financing, ownership, and payout policies across countries. Consequently, we re-estimate our primary regressions using per-capita GDP as an additional control variable. We find that the coefficient on GDP is generally negative, but never significant. Panel A in Table 6 shows that the previously documented relations between the institutional variables and the cost of capital do not materially change if we introduce per-capita GDP into the model.

Next, we investigate whether our results are robust to controls for the size of a country's equity market. La Porta et al. [1997, 2005] provide evidence that equity market development is an outcome of strong legal institutions. For this reason, equity market development may not be an appropriate right-hand side variable in our models due to its endogenous relation with the legal variables. However, we are not interested in the effect of equity market size per se; we solely test whether our results are sensitive to its inclusion in the model. Using the yearly ratio of equity market capitalization to GDP as a control, we find that the results and inferences are not materially affected and, if anything, become slightly stronger (Panel A, Table 6).[32]

We further check whether differences in voluntary disclosures influence our findings. The cost of capital estimates may also reflect firms' voluntary disclosure practices. Thus, to the extent that firms can privately commit to certain disclosures, the regressions in Table 4 may overstate the effect of securities regulation. To address this concern and capture variation in disclosure practice, we introduce the CIFAR disclosure index as an additional control variable (Hope [2003], Bushman et al. [2004]). We calculate a disclosure practice score by averaging the

[32] We also check whether the results are robust to controls for differences in market liquidity (e.g., Amihud and Mendelson [1986]). The results in Table 4 become even stronger when we add yearly share turnover as control.

1995 CIFAR firm-level disclosure ratings by country. We find that the coefficient on the CIFAR index has a negative sign, but is not significant. Our institutional variables continue to be significant, although at lower levels, and the coefficients are slightly attenuated. Thus, our results do not appear to be driven primarily by differences in firms' voluntary disclosures.[33]

Implied cost of capital models are based on earnings forecasts after corporate but before personal taxes. Thus, international differences in personal and corporate taxation are likely to be reflected in our cost of capital estimates and, at the same time, are likely to be associated with various institutional variables. As it is impossible to account for these differences directly in the cost of capital estimation, we make an attempt to control for them by including separate tax variables in our regressions. We use the average personal top income tax rate and the average top corporate tax rate, both taken from the *World Tax Database* of the University of Michigan. The introduction of tax rates slightly increases the coefficients and significance levels of DISREQ and SECREG, but it does not have any major impact on our findings.

Alternative Specifications and Estimation Methods

In this sub-section, we address concerns regarding our specifications and estimation methods. First, our design assumes that differences in the nominal risk-free rate stem only from differences in expected inflation rates. Although this assumption is common in the international finance literature, it is likely that real interest rates differ across countries, reflecting among other things saving rates or interest rate regimes.[34] Thus, it would be desirable to control for the real risk-free rate in each country. The difficulty, however, is that short-term interest rates on government

---

[33] Alternatively, we include a country's mean analyst following as an additional control variable because Lang and Lundholm [1996] show that analyst following and firms' voluntary disclosures are related. The results are very similar to those reported in Table 4.

[34] Return-based studies typically convert local currency returns to US$ returns and use the U.S. T-bill yield as risk-free rate (e.g., Harvey [1991]), assuming that inflation differences are properly reflected in the exchange rates, and that time preferences are similar across countries.

securities, which are usually taken as a proxy for the risk-free rate, are likely to reflect the quality of the institutional structure of some countries. For instance, the fact that governments have to pay extremely high interest rates or cannot even borrow in local currency often reflects institutional weaknesses. Thus, using local rates on government securities, we may partially take out the effects that are at the heart of our analysis.

For this reason, our main regressions follow the standard approach controlling for inflation differences only. To check the sensitivity of our results, we replace the proxy for the expected inflation with the nominal local risk-free rate using yields of local treasury bills, central bank papers or inter-bank loans provided by *Datastream*. The first row of Panel B in Table 6 shows that the institutional variables continue to be significant, although the coefficient magnitudes are slightly attenuated compared to Table 4.

A related issue is that we have chosen to introduce the inflation rate (or local risk-free rate) as a separate control variable, rather than to subtract it from the raw cost of capital estimates and to conduct the analysis based on risk premia. As discussed in Section 2, this approach has the advantage of not forcing a coefficient of minus one on the inflation proxy. In this section, we repeat our analyses using risk premia. The results for the two securities regulation variables are very similar (Panel B, Table 6), but the effect of LAW becomes insignificant. We also notice that the fit of these regressions is generally lower ($R^2 \approx 40\%$) than those reported in Table 4, supporting our approach of not forcing a specific coefficient on the inflation proxy.

Next, we address concerns about the endogeneity of our right-hand side variables. One issue is that market-based control variables may also reflect the effects of legal institutions. For instance, La Porta et al. [2002] show that well-functioning legal institutions lead to higher equity valuations, i.e., to differences in book-to-market ratios and market capitalizations. Similarly, good legal institutions may manifest in lower return variability. In principle, such effects should make it harder for us to find significant relations for the institutional variables once market-based controls are included.

To check whether our data actually supports this reasoning, we re-estimate our models using accounting-based controls only, which are less susceptible to the above effects. That is, we replace market capitalization, return variability and the book-to-market ratio with total assets and financial leverage, and scale the forecast bias variable by assets (rather than price) per share. Panel C of Table 6 reports the results from these regressions. The coefficients for all four institutional variables increase in the predicted direction and are highly significant.[35]

A remaining issue is that the institutional variables themselves may be endogenous. The concern is less reverse causality than selection bias, as countries choose their securities regulation. To address this issue, we need instrumental variables, which are generally hard to find. Several prior studies use a country's legal origin as instruments because they have been determined a long time ago, but at the same time are correlated with many institutional variables (e.g., Levine [1999], Leuz et al. [2003], La Porta et al. [2005]). We follow these studies and find that a binary variable indicating an English legal origin is highly correlated with the two securities regulation variables, but not with the LAW variable. Thus, we can only instrument DISREQ and SECREG. We estimate 2SLS regressions for the accounting-based model, which is more likely to have exogenous controls.[36] Panel C of Table 6 shows that both variables continue to be negatively associated with the cost of capital, mitigating endogeneity concerns for the securities regulation variables, provided the instrument is valid. Subject to the same caveat, these

[35] A similar argument can be made for the variables entering the construction of our proxy for macroeconomic variability. As expected, when we drop this variable, the findings are very similar and, if anything, stronger.

[36] The first-stage R^2 from a regression of DISREQ (SECREG) on the legal origin is 47.2% (37.9%), suggesting that a problem with weak instruments is unlikely. Note further that the IV results are similar if the full set of (market-based) controls is used.

findings also alleviate concerns that our main results are driven by correlated omitted variables because, by construction, an IV regression assures that the second-stage residuals are uncorrelated with the instrumented variable, i.e., the securities regulation measures.

Finally, we estimate Prais-Winsten regressions with panel-corrected standard errors. These standard errors go beyond the Newey-West corrected standard errors in accounting for the panel structure of our data. Panel-corrected standard errors allow for heteroscedasticity, within-panel serial correlation and cross-sectional dependence, and are more conservative than feasible GLS estimation (e.g., Beck and Katz [1995], Greene [2000]).[37] Panel C of Table 6 shows that, in both models, securities regulation and legal quality continue to be significant and, if anything, the marginal effects increase. The conclusions are similar to those from Table 4.

Sample Composition

In this sub-section, we conduct several robustness checks related to the sample period and composition. First, the securities regulation variables from La Porta et al. [2005] are measured as of December 2000. However, our analysis exploits data from 1992 to 2001 and hence assumes that securities regulation stays constant over this time frame. Presumably, securities regulation and legal institutions change only slowly. We nevertheless perform regressions using only data from the years 2000 and 2001, i.e., the time period surrounding the measurement of the securities regulation variables. Despite the fact that this restriction reduces our sample considerably, the results and inferences (Panel C, Table 6) are essentially the same as those in Table 4. The coefficient magnitudes increase noticeably, suggesting that measurement error in the institutional variables for the earlier years attenuates the coefficients in Table 4.

---

[37] We also estimate regressions using feasible generalized least squares. The results are similar and the inferences are stronger than in the Prais-Winsten regressions.

Second, as more than 30% of the firm-year observations stem from four individual countries, we re-run our analyses dropping observations from France, Japan, the U.K. and the U.S. The results (not tabulated) are very similar and, if anything, stronger than those reported in the tables. Next, we investigate whether the country-year analysis gives undue weight to small countries and ignores that the medians are based on varying numbers of firm-year observations. We address this issue with a weighted least squares regression giving higher weights to country-year observations that are based on more firm-year observations and presumably more precise. The results (not tabulated) are similar to those reported in Table 4.[38]

Country-Fixed Effects Regressions

As a final set of sensitivity analyses, we use an alternative approach to address several of the issues discussed above. We extract country-fixed effects based on firm-level regressions and then regress the country-fixed effects on the institutional variables. While this approach collapses the sample to 40 observations, it has several appealing features. First, each country enters the regression only once and receives equal weight. Second, it exploits firm-level information and controls for differences in within-country economic heterogeneity.

We estimate country-fixed effects using the firm-level version of the full model (Table 3, last column). The results are very similar if we use the firm-level versions of models 1 and 2 to extract the fixed effects. Panel D of Table 6 reports the results from regressions of the country-fixed effects on the institutional variables. The first set of regressions includes an intercept and either DISREQ or SECREG together with LAW. The disclosure requirements index, the composite securities regulation index and the legal quality index exhibit significantly negative

[38] We note that the estimated coefficients on the institutional variables become smaller in these regressions. Section 5 offers an explanation for this phenomenon. The weighted regressions implicitly give more weight to countries with integrated capital markets, for which the institutional effects are smaller.

associations with countries' cost of capital. We note that the coefficient magnitudes on the securities regulation variables are similar to those in Table 4, which is reassuring. Interestingly, the rule of law variable exhibits even larger coefficients than before. In the second set of regressions, we include growth in per-capita GDP as an additional control variable to make an attempt to capture international differences in real growth in these analyses as well.

As a sensitivity check, we repeat these analyses using accounting-based control variables to extract the country-fixed effects, and find that these fixed effects yield similar results. As the country-fixed effects capture observed and unobserved heterogeneity in the cost of capital, it is interesting to see how much variation in the fixed effects is explained by the institutional variables. We find that differences in securities regulation and legal quality explain a substantial portion (i.e., 35%) of the variation in countries' cost of capital.

In summary, our extensive sensitivity analyses show that the main results of this study are quite robust and generally support the conclusion that firms in countries with more extensive disclosure requirements and strong securities regulation have a lower cost of equity capital. The negative association of the quality of the legal system is not quite as robust, which may reflect the fact that the theoretical link with the cost of equity capital is weaker for this variable.

5. *Do the Effects Differ in Integrated and Segmented Markets?*

In this section, we analyze whether the previously documented cost of capital effects differ by capital market integration. This analysis serves two purposes. First, if we can show that the effects of securities regulation and supporting legal institutions differ cross-sectionally in a predicted way, such a finding further reduces concerns that our results are driven by correlated omitted variables or measurement error. Second, this analysis is an attempt to shed some light on why (and when) securities regulation appears to be priced in firms' cost of capital.

Both theory and prior empirical evidence suggest that country-specific factors become less important in asset pricing as markets become more integrated (e.g., Bekaert and Harvey [1995], Stulz [1999]). Similar evidence is available from ADR programs, which result in market integration at the firm level (e.g., Errunza and Miller [2000]). The basic idea is that, if investors can invest freely in stocks around the world, it is easier to find close substitutes in other countries, which is likely to reduce the effects of securities regulation and other legal institutions. Moreover, in integrated markets, risk is shared and diversified globally, which should reduce firms' cost of capital (Karolyi and Stulz [2003]). Lombardo and Pagano [2002a] also provide an analytical model showing that the effects of legal institutions may depend on whether countries' capital markets are integrated or segmented.

To explore these issues, we use two binary variables to capture capital market integration. The first variable is equal to one if the country's equity market is in the *MSCI Developed Markets Index* and zero otherwise (DEV). The second variable is based on portfolio flows across countries and captures more directly to what extent capital flows freely across borders. We create a binary variable and assign the value of one (zero) if the sum of a country's portfolio inflows and outflows divided by its GDP is above (below) the median (FLOW).[39] Panel C in Table 1 provides descriptive information on the two classifications.

We multiply the integration dummies with the securities regulation variables from Table 4 and introduce both the main and the interaction effects into the model. In essence, this

---

[39] In additional analyses (not tabulated), we use two alternative integration proxies: (1) the correlation of a country's equity market index with the *MSCI World Market Index* estimated from 1990 to 1995 by Erb et al. [1996a], and (2) an index of rules and restrictions on international transactions and capital flows by Harrison, Love, and McMillan [2002]. For each variable, we assign countries a value of one if they belong to the group of integrated economies and zero otherwise, using the median as cut-off (i.e., countries with above median correlation with the *MSCI World Market Index* and countries with below median capital controls). The results using these alternate integration variables are similar to those reported in Table 7.

specification estimates separate slope coefficients for the effect of disclosure rules and securities regulation in integrated and segmented markets.[40] Panel A of Table 7 presents the country-year regression results using the full set of controls. For brevity, we tabulate only the main and interaction effects of market integration (DEV or FLOW) and securities regulation (DISREQ or SECREG) as well as the main effect of the rule of law variable (LAW). We find that the main effect of capital market integration is always negative, highly significant in two out of four specifications, and marginally significant in a third. Thus, consistent with prior findings in the international finance literature, capital market integration is generally associated with a lower cost of equity capital (see Karolyi and Stulz [2003] for a survey).

The interaction effects are significantly positive in the SECREG specification and weakly positive in the DISREQ specification. These results suggest that, consistent with our expectations, the effect of securities regulation becomes smaller in integrated markets.[41] Based on the Wald tests for the sum of the coefficients α_1 and α_2, we find that extensive disclosure requirements continue to exhibit a small negative effect on the cost of equity capital in integrated markets. Using the MSCI integration variable, the effect is significant at the 0.06 level and around 70 basis points going from the 25th to the 75th percentile of the disclosure regulation variable. The composite securities regulation index exhibits an interaction effect of a similar magnitude as the main effect, resulting in an insignificant Wald test and suggesting that, in integrated markets, the effect of securities regulation on the cost of equity capital is essentially

---

[40] Given the focus of our paper, we do so for DISREQ and SECREG only. But the results are very similar to those reported if we also introduce main and interaction effects for the LAW variable. Moreover, the LAW variable exhibits a similar pattern as the other two variables with respect to capital market integration.

[41] This finding also provides an explanation for why Lee et al. [2004] find insignificant coefficients for three institutional variables from La Porta et al. [1998] included in some of their models. Their sample comprises firms from the G7 countries only, which probably have the most integrated capital markets.

zero. Together, the findings suggest that institutional effects are much smaller in integrated markets.

Correspondingly, the estimated effects of disclosure regulation and related enforcement mechanisms are much larger in segmented countries. For instance, using either integration variable, the estimated effect of securities regulation on the cost of equity capital is about 200 basis points going from the 25th to the 75th percentile of the SECREG variable. This decline in the cost of capital translates roughly into a 19-percent market value premium, holding expected cash flows constant.[42] Thus, in segmented markets, cost of capital differences alone can explain substantial valuation effects related to countries' legal institutions, as they for instance are shown in La Porta et al. [2002].

We again conduct a number of robustness checks. First, as in Section 4.3, we check whether our results are sensitive to the inclusion of additional variables. Adding any of the variables from Panel A of Table 6 or using local short-term rates instead of the expected inflation proxy does not materially alter the results presented in Table 7. The results are also similar if we use Prais-Winsten regressions with panel-corrected standard errors instead of OLS estimates.

As a last sensitivity check, we conduct the integration analysis based on country-fixed effects. The results for the two integration variables are presented in Panel B of Table 7 and are similar to those based on the country-year regressions. Although the significance levels for the main and interaction effects are generally lower, the effects of securities regulation are again much stronger in segmented economies and substantially attenuated in integrated markets.

---

[42] Evaluating at the sample mean, we can roughly gauge the effect as follows: (1/(0.1249-0.020))/(1/0.1249).

6. *Conclusions*

In this paper, we examine cross-country differences in the cost of equity capital. We analyze whether differences in countries' disclosure and securities regulation are systematically related to international cost of capital differences. The existence of this link is not obvious and depends on the extent to which differences in securities regulation lead to measurable differences in non-diversifiable risk across countries and on the degree to which capital market integration attenuates these differences.

We estimate the cost of equity capital for 40 developed and developing countries over the period from 1992 to 2001. Our primary analysis is based on four different models estimating the ex ante cost of capital implied in share prices and analyst forecasts. We document substantial variation in the cost of equity capital across countries, much of which is explained by traditional proxies for risk, i.e., size, volatility and the book-to-market ratio, as well as country controls capturing international differences in inflation and macroeconomic variability. Together, these variables explain about 60% of the country-level and close to 40% of the firm-level variation in the implied cost of equity capital around the world.

Using proxies for the extensiveness of a country's disclosure requirements, the effectiveness of its securities regulation and the overall quality of its legal system, we find that countries' legal institutions are significantly related to international differences in the cost of equity capital. Specifically, firms in countries with more extensive disclosure requirements, stronger securities regulation, and to a lesser extent, firms in countries with higher quality legal systems display a lower cost of capital, even after traditional controls for firm and country risk.

We further investigate to what extent the effect of securities regulation differs by market integration and economic development. Using several proxies, we find that the effects are strongest for markets that are least integrated. In these markets, we gauge the estimated effect of stringent disclosure requirements to be on the order of 200 basis points going from the 25th to the 75th percentile of the disclosure variable. Thus, in segmented markets, the effect of securities regulation on the cost of equity capital is economically significant and can explain substantial differences in international valuation. In contrast, the effects are substantially smaller and in several cases insignificant among countries with integrated capital markets. These findings are consistent with the notion that, in integrated economies, risk is shared and priced globally.

Finally, several caveats are in order. First, our study focuses exclusively on the link between legal institutions and firms' cost of *equity* capital. It is conceivable that countries with weak securities regulation and relatively high cost of equity capital have created institutions that lower the cost of debt. Consequently, our evidence does not provide a ranking of firms' overall (or weighted) cost of capital. Second, as many legal institutions are complementary, variables characterizing countries' institutional characteristics tend to be highly correlated. For this reason, it is difficult to disentangle the marginal effects of various legal institutions. Thus, the labels attached to our institutional variables should not be interpreted too narrowly. Third, and perhaps most importantly, the cost of capital is notoriously difficult to measure. In particular, it is difficult to ensure that the estimates do not reflect differences in growth expectations, rather than firms' cost of capital. We provide extensive sensitivity analyses to see whether growth differences across countries unduly influence our results. While these analyses suggest that our findings are quite robust and not spurious, we acknowledge that some concerns about measurement error remain. Our evidence should therefore be interpreted cautiously keeping these difficulties in mind. Lastly, our results may be subject to sample truncation bias. The data requirements of implied cost of capital models screen out lesser known firms with no or weak analyst coverage. Moreover, we do not observe firms for which the cost of publicly traded equity capital is prohibitive and which are not listed. Both effects are more likely to occur in countries with weak legal institutions, suggesting that, in these countries, the better known and more international firms enter our sample. As these firms are likely to have a lower cost of equity capital, the benefits of securities regulation and supporting legal institutions may be *under*stated in our analyses.

APPENDIX: Implied Cost of Equity Capital Models

A.1 OVERVIEW AND MODEL-SPECIFIC ASSUMPTIONS

Claus and Thomas [2001]:

$$P_t = bv_t + \sum_{\tau=1}^{T} \frac{\left(\hat{x}_{t+\tau} - r_{CT} \cdot bv_{t+\tau-1}\right)}{\left(1 + r_{CT}\right)^{\tau}} + \frac{\left(\hat{x}_{t+T} - r_{CT} \cdot bv_{t+T-1}\right)\left(1+g\right)}{\left(r_{CT} - g\right)\left(1 + r_{CT}\right)^{T}}$$

Model-specific assumptions:
This is a special case of the residual income valuation model. It uses actual book values per share and forecasted earnings per share up to five years ahead to derive the expected future residual income series. We define residual income as forecasted earnings per share less a cost of capital charge for beginning of fiscal year book value of equity per share. We assume clean surplus, i.e., future book values are imputed from current book values, forecasted earnings and dividends. Dividends are set equal to a constant fraction of forecasted earnings. At time $T = 5$, it is assumed that (nominal) residual income grows at rate g equal to the expected inflation. As a proxy for g, we use the (annualized) median of country-specific, one-year-ahead realized monthly inflation rates. Note that g sets a lower bound to the cost of capital estimates.

Gebhardt, Lee, and Swaminathan [2001]:

$$P_t = bv_t + \sum_{\tau=1}^{T} \frac{\left(\hat{x}_{t+\tau} - r_{GLS} \cdot bv_{t+\tau-1}\right)}{\left(1 + r_{GLS}\right)^{\tau}} + \frac{\left(\hat{x}_{t+T+1} - r_{GLS} \cdot bv_{t+T}\right)}{r_{GLS}\left(1 + r_{GLS}\right)^{T}}$$

Model-specific assumptions:
This is a special case of the residual income valuation model. It uses actual book values per share and forecasted earnings per share up to three years ahead to impute future expected residual income for an initial three-year period. We assume clean surplus, i.e., future book values are imputed from current book values, forecasted earnings and dividends. Dividends are set equal to a constant fraction of forecasted earnings. After the explicit forecast period of three years, the residual income series is derived by linearly fading the forecasted accounting return on equity to the sector-specific median return. We compute the historic three-year average return on equity in a given country and year for the industrial, service and financial sector. Negative sector-specific target returns are replaced by country-year medians. From $T = 12$ on residual income is assumed to remain constant.

Ohlson and Juettner-Nauroth [2005]:

$$P_t = \left(\hat{x}_{t+1}/r_{OJ}\right) \cdot \left(g_{st} + r_{OJ} \cdot \hat{d}_{t+1}/\hat{x}_{t+1} - g_{lt}\right) \Big/ \left(r_{OJ} - g_{lt}\right)$$

Model-specific assumptions:
This is a special case of the abnormal earnings growth valuation model developed by Ohlson and Juettner-Nauroth [2005]. It uses one-year ahead forecasted earnings and dividends per share as well as forecasts of short-term and long-term abnormal earnings growth. Dividends are set equal to a constant fraction of forecasted earnings. Following Gode and Mohanram [2003], the short-term growth rate g_{st} is estimated as the average between the forecasted percentage change in earnings from year t+1 to t+2 and the five-year growth forecast provided by financial analysts on *I/B/E/S*. The model requires a positive change in forecasted earnings to yield a numerical solution. The long-term earnings growth rate g_{lt} incorporates the assumption that growth in abnormal earnings per share beyond year t+1 equals the expected rate of inflation. We use the (annualized) country-specific median of one-year-ahead realized monthly inflation rates. Note that g_{lt} sets a lower bound to the cost of capital estimates.

Modified PEG ratio model by Easton [2004]:

$$P_t = \left(\hat{x}_{t+2} + r_{PEG} \cdot \hat{d}_{t+1} - \hat{x}_{t+1}\right) \Big/ r_{PEG}^2$$

Model-specific assumptions:
This is a special case of the abnormal earnings growth valuation model developed by Ohlson and Juettner-Nauroth [2005]. It uses one-year ahead and two-year ahead earnings per share forecasts as well as expected dividends per share in period t+1 to derive a measure of abnormal earnings growth. Dividends are set equal to a constant fraction of forecasted earnings. The model embeds the assumption that growth in abnormal earnings persists in perpetuity after the initial period. Note that it requires positive changes in forecasted earnings (including re-invested dividends) to yield a numerical solution.

Notes:

P_t	=	Market price of a firm's stock at date t
bv_t	=	Book value per share at the beginning of the fiscal year
$bv_{t+\tau}$	=	Expected future book value per share at date $t+\tau$, where $bv_{t+\tau} = bv_{t+\tau-1} + \hat{x}_{t+\tau} - \hat{d}_{t+\tau}$
$\hat{x}_{t+\tau}$	=	Expected future earnings per share for period ($t+\tau$–1, $t+\tau$) using either explicit analyst forecasts or future earnings derived from growth forecasts g, g_{st}, and g_{lt}, respectively
$\hat{d}_{t+\tau}$	=	Expected future net dividends per share for period ($t+\tau$–1, $t+\tau$), derived from the dividend payout ratio times the earnings per share forecast $\hat{x}_{t+\tau}$
g, g_{st}, g_{lt}	=	Expected (perpetual, short-term or long-term) future growth rate
$r_{CT}, r_{GLS}, r_{OJ}, r_{PEG}$	=	Implied cost of capital estimates calculated as the internal rate of return solving the above valuation equations, respectively

A.2 GENERAL ASSUMPTIONS AND DATA REQUIREMENTS

For an observation to be included in our sample we require current stock price data (P_t), analyst earnings per share forecasts for two periods ahead ($\hat{x}_{t+1}$ and $\hat{x}_{t+2}$), and either forecasted earnings per share for period t+3 ($\hat{x}_{t+3}$) or an estimate of long-term earnings growth (*ltg*). We obtain this information from the *I/B/E/S* database. If explicit earnings per share forecasts for the periods t+3 through t+5 are missing, we apply the following relation: $\hat{x}_{t+\tau} = \hat{x}_{t+\tau-1} \cdot (1 + ltg)$. Alternatively, if long-term growth projections are missing, we impute *ltg* from the percentage change in forecasted earnings per share between periods t+2 and t+3. We use only positive earnings forecasts and growth rates. All estimates are mean analyst consensus forecasts.

Stock prices and analyst forecasts are measured as of month +10 after the fiscal year end (*I/B/E/S* provides updates as of the third Thursday of each month).[43] This time lag is chosen to ensure that financial data, especially earnings and book values of equity, are publicly available and impounded in prices at the time we compute the cost of capital estimate. However, this implies that the one-year ahead forecast ($\hat{x}_{t+1}$) is for a fiscal year that ends just two months later.

[43] We repeat all analyses using price and earnings per share forecast data from month +7 (instead of month +10) after the fiscal year end and obtain very similar results.

Furthermore, the pricing date ($P_{t'}$, where t' refers to month +10) diverges from the equity valuation date in the above formulas (P_t, where t refers to the end of the previous fiscal year).

This misalignment of t' and t has no effect on the earnings forecasts per se. In the absence of any new information, a US$ 2 earnings per share forecast at the beginning of the fiscal year (t) yields the same number just 10 months later (t'). Prices on the other hand increase as they move closer to future expected cash flows, even without new information. To account for this appreciation in price, we discount the month +10 price ($P_{t'}$) to the beginning of the fiscal year, using the imputed cost of capital, i.e., we use $[1+r]^{-10/12}$ as discount factor, where r equals r_{CT}, r_{GLS}, r_{OJ} and r_{PEG}, respectively.[44] This adjustment directly yields cost of equity capital estimates on an annualized basis, which at the same time reflect the information set available at month +10 after the fiscal-year end.[45] As expected, the cost of capital estimates increase due to this adjustment. But our cross-sectional results reported in Tables 3 through 7 are virtually unchanged if we ignore the misalignment, i.e., if we use price and forecast data as of month +10 (t') together with financial data as of the fiscal-year end (t).

Net dividends ($\hat{d}_{t+\tau}$) are forecasted up to the finite forecast horizon as a constant fraction of expected future earnings per share. We define the dividend payout ratio (k_t) as the historic three-year average for each firm. If k_t is missing or outside the range of zero and one, we replace it by the country-year median payout ratio. We use the (annualized) country-specific median of one-year-ahead realized monthly inflation rates as proxy for expected inflation. Negative values are

---

44 For an alternative approach see Daske et al. (2006).

45 The described procedure is essentially equivalent to using month +10 prices ($P_{t'}$) and discounting the forecasted valuation attributes to date t' (see e.g., Francis, Olsson, and Oswald [2000], Botosan and Plumlee [2005]). In that case, we add a part-year factor of $^2/_{12}$ to the discount factor, i.e., $[1+r]^{\tau-1+2/12}$, where τ indicates the forecast year and r equals r_{CT}, r_{GLS}, r_{OJ} and r_{PEG}, respectively. In unreported analyses we confirm that this alternative adjustment produces very similar results to those reported in the tables and the text.

replaced by the country's historical inflation rate, estimated as the median of the monthly inflation rates over the entire sample period, because deflation cannot persist forever. We obtain all financial data (bv_t and k_t) from the *Worldscope* database. Inflation data are gathered from the *Datastream* and *Worldbank* databases.

Since most of the valuation models do not have a closed form solution, we use an iterative procedure to determine the internal rate of return. This numerical approximation identifies the annual firm-specific discount rate that equates P_t to the right-hand side of the respective equity valuation model. We stop iterating if the imputed price falls within a 0.001 difference of its actual value. Implied cost of equity capital estimates are restricted to be positive and set to missing otherwise.

REFERENCES

AMIHUD, Y., AND H. MENDELSON. "Asset Pricing and the Bid-Ask Spread." *Journal of Financial Economics* 32 (1986): 223–49.

BARRY, C., AND S. BROWN. "Differential Information and Security Market Equilibrium." *Journal of Financial and Quantitative Analysis* 20 (1985): 407–22.

BASU, S. "The Conservatism Principle and the Asymmetric Timeliness of Earnings." *Journal of Accounting and Economics* 24 (1997): 3–37.

BECK, N., AND J. KATZ. "What To Do (and Not To Do) with Time-Series Cross-Section Data." *American Political Science Review* 89 (1995): 634–47.

BEKAERT, G., AND C. HARVEY. "Time-Varying World Market Integration." *The Journal of Finance* 50 (1995): 403–44.

BEKAERT, G., AND C. HARVEY. "Foreign Speculators and Emerging Equity Markets." *The Journal of Finance* 55 (2000): 565–613.

BENSTON, G. "Required Disclosure and the Stock Market: An Evaluation of the Securities Exchange Act of 1934." *The American Economic Review* 63 (1973): 132–55.

BERKOWITZ, D.; K. PISTOR; AND J. RICHARD. "Economic Development, Legality and the Transplant Effect." *European Economic Review* 47 (2003): 165–95.

BHATTACHARYA, U., AND H. DAOUK. "The World Price of Insider Trading." *Journal of Finance* 57 (2002): 75–108.

BOTOSAN, C. "Disclosure Level and the Cost of Equity Capital." *The Accounting Review* 72 (1997): 323–49.

BOTOSAN, C., AND M. PLUMLEE. "Assessing Alternative Proxies for the Expected Risk Premium." *The Accounting Review* 80 (2005): 21–53.

BUSHEE, B., AND C. LEUZ. "Economic Consequences of SEC Disclosure Regulation: Evidence from the OTC Bulletin Board." *Journal of Accounting and Economics* 39 (2005): 233–64.

BUSHMAN, R.; J. PIOTROSKI; AND A. SMITH. "What Determines Corporate Transparency?" *Journal of Accounting Research* 42 (2004): 207–52.

BUSHMAN, R., AND J. PIOTROSKI. "Financial Reporting Incentives for Conservative Accounting: The Influence of Legal and Political Institutions." Working paper, University of North Carolina and University of Chicago, 2005.

CAMPBELL, J. "Understanding Risk and Return." *Journal of Political Economy* 104 (1996): 298–345.

CHEN, K.; Z. CHEN; AND K.C. WEI. "Disclosure, Corporate Governance, and the Cost of Equity Capital: Evidence from Asia's Emerging Markets." Working paper, Hong Kong University of Science & Technology, 2003.

CHEN, F.; B. JORGENSEN; AND Y. YOO. "Implied Cost of Equity Capital in Earnings-Based Valuation: International Evidence." *Accounting and Business Research* 34 (2004): 323–44.

CLARKSON, P.; J. GUEDES; AND R. THOMPSON. "On the Diversification, Observability, and Measurement of Estimation Risk." *Journal of Financial and Quantitative Analysis* 31 (1996): 69–84.

CLAUS, J., AND J. THOMAS. "Equity Premia as Low as Three Percent? Evidence from Analysts' Earnings Forecasts for Domestic and International Stock Markets." *Journal of Finance* 56 (2001): 1629–66.

COLES, J.; U. LOEWENSTEIN; AND J. SUAY. "On Equilibrium Pricing Under Parameter Uncertainty." *Journal of Financial and Quantitative Analysis* 30 (1995): 347–64.

DASKE, H.; G. GEBHARDT; AND S. KLEIN. "Estimating a Firm's Expected Cost of Equity Capital Using Consensus Forecasts." *Schmalenbach Business Review*, Forthcoming (2006).

EASLEY, D., AND M. O'HARA. "Information and the Cost of Capital." *Journal of Finance* 59 (2004): 1553–83.

EASTON, P. "PE Ratios, PEG Ratios, and Estimating the Implied Expected Rate of Return on Equity Capital." *The Accounting Review* 79 (2004): 73–95.

EASTON, P., AND S. MONAHAN. "An Evaluation of Accounting-Based Measures of Expected Returns." *The Accounting Review* 80 (2005): 501–38.

EASTON, P.; G. TAYLOR; P. SHROFF; AND T. SOUGIANNIS. "Using Forecasts of Earnings to Simultaneously Estimate Growth and the Rate of Return on Equity Investment." *Journal of Accounting Research* 40 (2002): 657–76.

ELTON, E. "Expected Return, Realized Return, and Asset Pricing Tests." *Journal of Finance* 54 (1999): 1199–1220.

ERB, C.; C. HARVEY; AND T. VISKANTA. "Expected Returns and Volatility in 135 Countries." *The Journal of Portfolio Management* (1996a): 46–58.

ERB, C.; C. HARVEY; AND T. VISKANTA. "Political Risk, Economic Risk, and Financial Risk." *Financial Analysts Journal* 52 (1996b): 29–46.

ERRUNZA, V., AND D. MILLER. "Market Segmentation and the Cost of Capital in International Equity Markets." *Journal of Financial and Quantitative Analysis* 35 (2000): 577–600.

FAMA, E., AND K. FRENCH. "The Cross-Section of Expected Stock Returns." *The Journal of Finance* 47 (1992): 427–65.

FAMA, E., AND K. FRENCH. "Common Risk Factors on the Returns of Stocks and Bonds." *Journal of Financial Economics* 33 (1993): 3–57.

FAMA, E., AND K. FRENCH. "Industry Costs of Equity." *Journal of Financial Economics* 43 (1997): 153–94.

FERSON, W., AND C. HARVEY. "Fundamental Determinants of National Equity Market Returns: A Perspective on Conditional Asset Pricing." *Journal of Banking & Finance* 21 (1998): 1625–65.

FRANCIS, J.; R. LAFOND; P. OLSSON; AND K. SCHIPPER. "Costs of Equity and Earnings Attributes." *The Accounting Review* 79 (2004): 967–1010.

FRANCIS, J.; P. OLSSON; AND D. OSWALD. "Comparing the Accuracy and Explainability of Dividend, Free Cash Flow, and Abnormal Earnings Equity Value Estimates." *Journal of Accounting Research* 38 (2000): 45–70.

GEBHARDT, W.; C. LEE; AND B. SWAMINATHAN. "Toward an Implied Cost of Capital." *Journal of Accounting Research* 39 (2001): 135–76.

GODE, D., AND P. MOHANRAM. "Inferring the Cost of Capital Using the Ohlson-Juettner Model." *Review of Accounting Studies* 8 (2003): 399–431.

GREENE, W. *Econometric Analysis, 4th Edition*. Upper Saddle River, NJ: Prentice-Hall, 2000.

GREENSTONE, M.; P. OYER; AND A. VISSING-JORGENSEN. "Mandated Disclosure, Stock Returns, and the 1964 Securities Acts Amendments." *Quarterly Journal of Economics*, Forthcoming (2005).

GUAY, W.; S.P. KOTHARI; AND S. SHU. "Properties of Implied Cost of Capital Using Analysts' Forecasts." Working paper, The Wharton School, MIT, and Boston College, 2005.

HAIL, L. "The Impact of Voluntary Corporate Disclosures on the Ex-ante Cost of Capital for Swiss Firms." *European Accounting Review* 11 (2002): 741–73.

HAIL, L., AND C. LEUZ. "Cost of Capital and Cash Flow Effects of U.S. Cross-Listings." Working paper, University of Pennsylvania, 2005.

HARRISON, A.; I. LOVE; AND M. MCMILLAN. "Global Capital Flows and Financing Constraints." Working paper, National Bureau of Economic Research, 2002.

HARVEY, C. "The World Price of Covariance Risk." *The Journal of Finance* 46 (1991): 111–58.

HARVEY, C. "Predictable Risk and Returns in Emerging Markets." *The Review of Financial Studies* 8 (1995): 773–816.

HARVEY, C. "Country Risk Components, the Cost of Capital, and Returns in Emerging Markets." Working paper, Duke University, 2004.

HEALY, P., AND K. PALEPU. "Information Asymmetry, Corporate Disclosure, and the Capital Markets: A Review of the Empirical Disclosure Literature." *Journal of Accounting and Economics* 31 (2001): 405–40.

HIMMELBERG, C.; G. HUBBARD; AND I. LOVE. "Investor Protection, Ownership, and the Cost of Capital." Working paper, Columbia University and World Bank, 2004.

HOPE, O. "Disclosure Practices, Enforcement of Accounting Standards and Analysts' Forecast Accuracy: An International Study." *Journal of Accounting Research* 41 (2003): 235–73.

HUGHES, J.; J. LIU; AND J. LIU. "Information, Diversification and the Cost of Capital." Working paper, University of California at Los Angeles, 2005.

JOOS, P., AND M. LANG. "The Effects of Accounting Diversity: Evidence from the European Union." *Journal of Accounting Research* 32 (1994): 141–68.

KAROLYI, A., AND R. STULZ. "Are Financial Assets Priced Locally or Globally?" in *The Handbook of the Economics of Finance*, edited by G. Constantinides, M. Harris and R. Stulz. North-Holland, 2003.

LAMBERT, R.; C. LEUZ; AND R. VERRECCHIA. "Accounting Information, Disclosure and the Cost of Capital." Working paper, University of Pennsylvania, 2005.

LANG, M., AND R. LUNDHOLM. "Corporate Disclosure Policy and Analyst Behavior." *The Accounting Review* 71 (1996): 467–92.

LA PORTA, R.; F. LOPEZ-DE-SILANES; AND A. SHLEIFER. "What Works in Securities Laws?" *The Journal of Finance*, Forthcoming (2005).

LA PORTA, R.; F. LOPEZ-DE-SILANES; A. SHLEIFER; AND R. VISHNY. "Legal Determinants of External Finance." *Journal of Finance* 52 (1997): 1131–50.

LA PORTA, R.; F. LOPEZ-DE-SILANES; A. SHLEIFER; AND R. VISHNY. "Law and Finance." *Journal of Political Economy* 106 (1998): 1113–55.

LA PORTA, R.; F. LOPEZ-DE-SILANES; A. SHLEIFER; AND R. VISHNY. "Investor Protection and Corporate Governance." *Journal of Financial Economics* 58 (2000a): 3–27.

LA PORTA, R.; F. LOPEZ-DE-SILANES; A. SHLEIFER; AND R. VISHNY. "Agency Problems and Dividend Policies Around the World." *The Journal of Finance* 55 (2000b): 1–33.

LA PORTA, R.; F. LOPEZ-DE-SILANES; A. SHLEIFER; AND R. VISHNY. "Investor Protection and Corporate Valuation." *The Journal of Finance* 57 (2002): 1147–70.

LEE, C., AND D. NG. "Corruption and International Valuation: Does Virtue Pay?" Working paper, Cornell University, 2002.

LEE, C.; D. NG; AND B. SWAMINATHAN. "International Asset Pricing: Evidence from the Cross Section of Implied Cost of Capital." Working paper, Cornell University and University of Chicago, 2004.

LEUZ, C., AND R. VERRECCHIA. "The Economic Consequences of Increased Disclosure." *Journal of Accounting Research* 38 (2000): 91–124.

LEUZ, C.; D. NANDA; AND P. WYSOCKI. "Earnings Management and Investor Protection: An International Comparison." *Journal of Financial Economics* 69 (2003): 505–27.

LEVINE, R. "Law, Finance and Economic Growth." *Journal of Financial Intermediation* 8 (1999): 8–35.

LEVITT, A. "The Importance of High Quality Accounting Standards." *Accounting Horizons* 12 (1998): 79–82.

LOMBARDO, D., AND M. PAGANO. "Law and Equity Markets: A Simple Model," in *Convergence and Diversity of Corporate Governance Regimes and Capital Markets*, edited by J. McCahery, P. Moerland, T. Raaijmakers and L. Renneboog. Oxford University Press, 2002a.

LOMBARDO, D., AND M. PAGANO. "Legal Determinants of the Return on Equity." Working paper, Stanford Graduate School of Business and University of Salerno, 2002b.

MERTON, R. "A Simple Model of Capital Market Equilibrium with Incomplete Information." *Journal of Finance* 42 (1987): 483–510.

MILLER, G. "Earnings Performance and Discretionary Disclosure." *Journal of Accounting Research* 40 (2002): 173–204.

NEWEY, W., AND K. WEST. "A Simple, Positive Semi-definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix." *Econometrica* 55 (1987): 703–8.

OBSTFELD, M. "Risk-taking, Global Diversification, and Growth." *American Economic Review* 84 (1994): 1310–29.

OHLSON, J. "Earnings, Book Values, and Dividends in Equity Valuation." *Contemporary Accounting Research* 11 (1995): 661–87.

OHLSON, J., AND B. JUETTNER-NAUROTH. "Expected EPS and EPS Growth as Determinants of Value." *Review of Accounting Studies* 10 (2005): 349–65.

SOLNIK, B. "An Equilibrium Model of the International Capital Market." *Journal of Economic Theory* 8 (1974): 500–24.

STULZ, R. "A Model of International Asset Pricing." *Journal of Financial Economics* 9 (1981): 383–406.

STULZ, R. "Globalization, Corporate Finance, and the Cost of Capital." *Journal of Applied Corporate Finance* 12 (1999): 8–25.

VERRECCHIA, R. "Essays on Disclosure." *Journal of Accounting and Economics* 32 (2001): 91–180.

TABLE 1

Descriptive Statistics for the Implied Cost of Capital Estimates, Legal and Integration Variables

Panel A: Distributional Statistics for the Cost of Capital Estimates

Variable	*N*	*Mean*	*Min.*	*Percentile* *Q1*	*Q2*	*Q3*	*Max.*	*Standard Deviation*
r_{CT}	358	12.17%	3.98%	9.72%	11.36%	13.96%	28.90%	3.70%
r_{GLS}	358	9.25%	1.34%	6.79%	8.87%	11.05%	27.22%	3.83%
r_{OJ}	358	14.59%	6.19%	12.00%	13.65%	16.56%	35.72%	4.01%
r_{PEG}	358	13.96%	5.44%	11.57%	13.13%	15.37%	34.91%	3.82%
r_{AVG}	358	12.49%	4.67%	10.32%	11.63%	14.06%	27.81%	3.45%

Panel B: Pearson Correlation Coefficients among Cost of Capital Estimates

Variable	r_{CT}	r_{GLS}	r_{OJ}	r_{PEG}
r_{GLS}	0.617 **			
r_{OJ}	0.945 **	0.562 **		
r_{PEG}	0.875 **	0.506 **	0.933 **	
r_{AVG}	0.958 **	0.747 **	0.959 **	0.924 **

(continued)

The cost of capital estimates are based on 35,118 firm-year observations from 40 countries between 1992 and 2001, for which sufficient Worldscope financial data, I/B/E/S forecast and pricing data, and legal institutional data exist. The computations use country-year medians. We have 358 country-years for our analysis, after excluding country-years with less than five individual firm observations or with inflation rates above 25%. The implied cost of capital estimates, r_{CT}, r_{GLS}, r_{OJ} and $r_{PEG,}$ are derived as the internal rate of return in the Claus and Thomas [2001], the Gebhardt et al. [2001], the Ohlson and Juettner-Nauroth [2005], and the Easton [2004] models, respectively. These models are described in more detail in the Appendix. r_{AVG} is the mean of the four estimates for the implied cost of equity capital. All estimates use mean analyst consensus forecasts and realized price data as of month +10 after the fiscal-year end. ** indicates statistical significance at the 1% level (two-tailed).

TABLE 1 (continued)

Panel C: Sample Information, Cost of Capital Estimates, Legal and Integration Variables by Country

Country	*Firm-Years*	*Country-Years*	r_{AVG}	*DISREQ*	*SECREG*	*LAW*	*DEV*	*FLOW*
Argentina	179	8	12.81%	0.50	0.43	0.54	0	0
Australia	1,596	10	10.72%	0.75	0.77	1.00	1	0
Austria	228	8	11.21%	0.25	0.18	1.00	1	1
Belgium	478	9	11.00%	0.42	0.34	1.00	1	1
Brazil	323	7	20.85%	0.25	0.39	0.63	0	0
Canada	1,560	10	10.53%	0.92	0.91	1.00	1	1
Chile	147	10	12.55%	0.58	0.50	0.70	0	0
Denmark	575	10	10.78%	0.58	0.50	1.00	1	1
Egypt	15	2	25.27%	0.50	0.34	0.42	0	0
Finland	491	10	13.40%	0.50	0.49	1.00	1	1
France	2,116	10	10.37%	0.75	0.58	0.90	1	1
Germany	1,598	10	10.05%	0.42	0.21	0.92	1	1
Greece	316	8	12.16%	0.33	0.38	0.62	0	0
Hong Kong	1,193	10	14.58%	0.92	0.81	0.82	1	1
India	728	9	14.39%	0.92	0.75	0.42	0	0
Indonesia	404	9	16.60%	0.50	0.59	0.40	0	0
Ireland	182	8	12.74%	0.67	0.49	0.78	1	1
Israel	48	5	11.41%	0.67	0.65	0.48	0	0
Italy	759	10	10.61%	0.67	0.46	0.83	1	1
Japan	3,391	10	6.16%	0.75	0.47	0.90	1	1
Korea (South)	970	10	14.10%	0.75	0.55	0.54	0	0
Malaysia	1,248	10	10.65%	0.92	0.78	0.68	0	0
Mexico	260	8	15.59%	0.58	0.35	0.54	0	0
Netherlands	1,063	10	12.75%	0.50	0.62	1.00	1	1
New Zealand	327	10	11.14%	0.67	0.48	1.00	1	0
Norway	478	10	13.08%	0.58	0.43	1.00	1	1
Pakistan	75	7	19.51%	0.58	0.52	0.30	0	0
Peru	58	6	17.37%	0.33	0.59	0.25	0	0
Philippines	324	10	13.48%	0.83	0.89	0.27	0	0
Portugal	179	9	11.46%	0.42	0.55	0.87	0	1
Singapore	855	10	10.01%	1.00	0.84	0.86	1	1
South Africa	939	9	16.16%	0.83	0.58	0.44	0	0
Spain	636	10	10.80%	0.50	0.50	0.78	1	1
Sri Lanka	59	6	16.96%	0.75	0.52	0.19	0	0
Sweden	818	10	12.47%	0.58	0.45	1.00	1	1
Switzerland	948	10	10.90%	0.67	0.48	1.00	1	1
Taiwan	699	10	9.87%	0.75	0.64	0.85	0	0
Thailand	433	10	13.46%	0.92	0.62	0.63	0	0
United Kingdom	4,155	10	10.64%	0.83	0.73	0.86	1	1
United States	4,267	10	10.24%	1.00	0.97	1.00	1	1
Total (Average)	35,118	358	12.97%	0.65	0.56	0.74	21	20

The first two columns provide the number of firm- and country-years by country. r_{AVG} is the mean of the four estimates for the implied cost of equity capital as described in the Appendix. The table reports time-series averages of r_{AVG}. DISREQ measures the level of disclosure regulation based on an index of disclosure requirements in securities offerings. SECREG captures the strength of securities regulation mandating and enforcing disclosures. It is measured as the mean of the disclosure index, the liability standard index and the public enforcement index. All securities regulation variables stem from La Porta et al. [2005]. LAW represents the general quality of the legal environment and is measured as the rule of law index (divided by 10) from La Porta et al. [1997]. The two binary measures of capital market integration are: (1) DEV is equal to one if the country's equity market is classified as developed in the MSCI database, (2) FLOW is equal to one for countries with above-median portfolio inflows and outflows in percent of the GDP, as reported by the IMF for 2001.

TABLE 2

Descriptive Statistics for the Control Variables

Panel A: Distributional Statistics

Variable	*N*	*Mean*	*Min.*	*Percentile* *Q1*	*Q2*	*Q3*	*Max.*	*Standard Deviation*
INFL	358	3.54%	-3.81%	1.52%	2.61%	4.62%	20.02%	3.33%
SIZE	358	709.0	32.5	283.8	458.1	718.3	8,898.6	1,024.0
BETA	358	0.863	-0.653	0.592	0.788	1.060	2.708	0.475
RETVAR	358	0.104	0.046	0.077	0.099	0.118	0.273	0.036
BMR	358	0.600	0.131	0.436	0.573	0.696	1.895	0.234
MACVAR	358	-0.011	-1.307	-0.764	-0.453	-0.154	7.933	1.626
FBIAS	358	0.007	-0.055	0.000	0.002	0.007	0.200	0.018

Panel B: Pearson Correlation Coefficients

Variable	*INFL*	*SIZE*	*BETA*	*RETVAR*	*BMR*	*MACVAR*	*FBIAS*
SIZE	-0.086						
BETA	-0.041	0.010					
RETVAR	0.204 **	-0.244 **	0.367 **				
BMR	0.047	-0.334 **	0.284 **	0.233 **			
MACVAR	0.427 **	-0.236 **	0.298 **	0.381 **	0.211 **		
FBIAS	0.034	-0.021	0.136 **	0.210 **	-0.005	0.188 **	
r_{AVG}	0.457 **	-0.355 **	0.155 **	0.403 **	0.353 **	0.423 **	0.154 **

The sample is based on 35,118 firm-year observations from 40 countries between 1992 and 2001, for which sufficient Worldscope financial data, I/B/E/S forecast and pricing data, and legal institutional data exist. The table reports country-year medians. We have 358 country-years for our analysis, after excluding country-years with less than five individual firm observations or with inflation rates above 25%. INFL is the yearly median of country-specific, one-year-ahead realized monthly inflation rates. SIZE stands for US$ market value of outstanding equity (in millions). We use the natural log of the size variable to compute correlations. BETA is measured relative to the world market index and estimated based on a five-year market model regression. We require at least 24 monthly observations to calculate market beta. RETVAR is the returns variability computed as annual standard deviation of monthly stock returns. BMR is the ratio of accounting book value to market value of equity. Financial data are measured as of the fiscal-year end, BETA, RETVAR and the cost of capital as of month +10. MACVAR is the first principal component of four proxies for macroeconomic variability: (1) the country-year median standard deviation of annual earnings per share over the last five years scaled by total assets per share, (2) the country-year median standard deviation of accounting returns on equity over the last five years, (3) the standard deviation of the residuals from a regression of annual GDP growth rates on a time index over the sampling period, and (4) the coefficient of variation of yearly average exchange rates (US$ to local currency) over the sampling period. FBIAS is the one-year-ahead analyst forecast error (mean forecast minus actual) scaled by forecast-period stock price. r_{AVG} is the mean of the four estimates for the implied cost of equity capital as described in the Appendix. ** indicates statistical significance at the 1% level (two-tailed).

TABLE 3

Regression Analysis of the Implied Cost of Equity Capital on Risk and Country Control Variables

$$r_{AVGit} = \alpha_0 + \alpha_1 INFL_{it} + \alpha_2 SIZE_{it} + \alpha_3 BETA_{it} + \alpha_4 RETVAR_{it} + \alpha_5 BMR_{it} + \alpha_6 MACVAR_{it} + \alpha_7 EPSVAR_{it} + \alpha_8 FBIAS_{it} + \sum\alpha_j \text{Industry Controls}_{it} + \sum\alpha_k \text{Year Controls}_t + \varepsilon_{it}$$

Variable	*Pred. Sign*	*Model 1*	*Model 2*	*Model 3*	*Model 4*
N		358	358	358	20,830
Intercept		0.144 *	0.121 #	0.126 *	0.161 **
		(2.16)	(1.82)	(1.97)	(42.55)
INFL	+	0.372 **	0.304 **	0.291 **	0.638 **
		(5.87)	(4.48)	(4.32)	(32.90)
SIZE	–	-0.010 **	-0.007 **	-0.007 **	-0.006 **
		(-4.62)	(-3.27)	(-2.80)	(-26.24)
BETA	+	0.011 **	–	–	–
		(2.60)			
RETVAR	+	–	0.217 **	0.147 *	0.077 **
			(3.71)	(2.30)	(9.87)
BMR	+	0.024 **	0.027 **	0.027 **	0.019 **
		(2.99)	(3.76)	(3.98)	(20.55)
MACVAR	+	–	–	0.003 *	–
				(2.24)	
EPSVAR	+	–	–	–	0.063 **
					(6.01)
FBIAS	+	–	–	0.097	0.168 **
				(1.00)	(20.52)
Industry and Year Controls		included	included	included	included
R^2		57.8%	58.4%	60.0%	36.1%
F-Stat		11.7 **	12.8 **	12.6 **	193.8 **

The sample comprises 358 country-year medians (20,830 firm-years) from 40 countries over the ten-year period from 1992 to 2001. The dependent variable, r_{AVG}, is the mean of the four estimates for the implied cost of equity capital as described in the Appendix. INFL is the yearly median of country-specific, one-year-ahead realized monthly inflation rates. SIZE stands for US$ market value of outstanding equity (in thousands). We use the natural log of size in the analysis. BETA is measured relative to the MSCI world market index and estimated based on a five-year market model regression. We require at least 24 monthly observations to calculate market beta. RETVAR is the returns variability computed as annual standard deviation of monthly stock returns. BMR is the ratio of accounting book value to market value of equity. Financial data are measured as of the fiscal-year end, BETA, RETVAR and the cost of capital as of month +10. MACVAR is the first principal component of four proxies for macroeconomic variability: (1) the country-year median standard deviation of annual earnings per share over the last five years scaled by total assets per share, (2) the country-year median standard deviation of accounting returns on equity over the last five years, (3) the standard deviation of the residuals from a regression of annual GDP growth rates on a time index over the sampling period, and (4) the coefficient of variation of yearly average exchange rates (US$ to local currency) over the sampling period. In the firm-level regression (Model 4), we replace MACVAR by the firm's standard deviation of annual earnings per share over the last five years scaled by total assets per share (EPSVAR). We require at least three yearly observations to calculate earnings variability. FBIAS is the one-year-ahead analyst forecast error (mean forecast minus actual) scaled by forecast-period stock price. Industry controls, i.e., the percentage of firms in an industry per country-year based on the classification in Campbell [1996], and year indicators are included in the regressions but not reported. The table reports OLS coefficient estimates and, in parentheses, t-statistics based on Newey-West heteroscedasticity and autocorrelation corrected standard errors (standard errors are clustered by firm for Model 4). **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 4

Regression Analysis of the Implied Cost of Capital on Securities Regulation and Legal Quality

$$r_{AVGit} = \alpha_0 + \alpha_1\text{Securities Regulation}_i + \alpha_2\text{LAW}_i + \alpha_3\text{INFL}_{it} + \alpha_4\text{SIZE}_{it} + \alpha_5\text{BETA}_{it} + \alpha_6\text{RETVAR}_{it} + \alpha_7\text{BMR}_{it}$$
$$+ \alpha_8\text{MACVAR}_{it} + \alpha_9\text{FBIAS}_{it} + \sum\alpha_j\text{Industry Controls}_{it} + \sum\alpha_k\text{Year Controls}_t + \varepsilon_{it}$$

		Dependent Variable = r_{AVG}							*Dependent Var.* = r_{WEIGHT}	
Variable	*Predicted Sign*	*DISREQ*	*SECREG*	*LAW*	*DISREQ & LAW*	*SECREG & LAW*	*DISREQ & LAW*	*SECREG & LAW*	*DISREQ & LAW*	*SECREG & LAW*
N		358	358	358	358	358	358	358	358	358
Securities Regulation	–	-0.027 **	-0.023 *	–	-0.034 **	-0.025 *	-0.035 **	-0.023 *	-0.044 **	-0.034 **
		(-2.72)	(-2.24)		(-3.14)	(-2.37)	(-3.17)	(-2.09)	(-4.05)	(-3.17)
Legal Quality	–	–	–	-0.014	-0.028 *	-0.018	-0.044 **	-0.034 **	-0.035 *	-0.023 #
				(-1.06)	(-2.01)	(-1.40)	(-3.53)	(-3.01)	(-2.38)	(-1.64)
Intercept		0.175 **	0.160 *	0.158 *	0.250 **	0.204 **	0.297 **	0.247 **	0.232 **	0.174 *
		(2.57)	(2.52)	(2.22)	(3.09)	(2.89)	(3.59)	(3.44)	(2.69)	(2.26)
INFL	+	0.303 **	0.298 **	0.275 **	0.274 **	0.278 **	0.315 **	0.324 **	0.200 **	0.205 **
		(4.53)	(4.52)	(3.89)	(4.03)	(4.02)	(4.89)	(4.85)	(2.97)	(2.98)
SIZE	–	-0.006 *	-0.006 *	-0.006 *	-0.004	-0.004 #	-0.006 *	-0.007 **	-0.004 #	-0.005 #
		(-2.49)	(-2.28)	(-2.37)	(-1.63)	(-1.72)	(-2.36)	(-2.58)	(-1.69)	(-1.79)
BETA	+	–	–	–	–	–	0.010 **	0.011 **	–	–
							(2.61)	(2.67)		
RETVAR	+	0.191 **	0.174 **	0.117 #	0.143 *	0.138 *	–	–	0.064	0.058
		(3.04)	(2.73)	(1.65)	(2.16)	(1.99)			(0.95)	(0.83)
BMR	+	0.023 **	0.025 **	0.028 **	0.024 **	0.026 **	0.020 **	0.022 **	0.022 **	0.025 **
		(3.26)	(3.60)	(4.08)	(3.35)	(3.74)	(2.66)	(2.86)	(2.91)	(3.38)
MACVAR	+	0.003 *	0.003 *	0.003 *	0.002 #	0.003 *	–	–	0.001	0.002
		(2.08)	(2.43)	(2.13)	(1.81)	(2.32)			(0.71)	(1.44)
FBIAS	+	0.103	0.091	0.102	0.115	0.098	–	–	-0.005	-0.028
		(1.16)	(0.99)	(1.06)	(1.33)	(1.07)			(-0.05)	(-0.25)
Industry and Year Controls		included	included	included	included	included	included	included	included	included
R^2		61.3%	60.9%	60.3%	62.1%	61.3%	60.9%	59.8%	54.6%	53.4%
F-Stat		12.3 **	11.9 **	13.5 **	12.9 **	12.9 **	13.6 **	13.6 **	9.0 **	8.5 **

The sample comprises 358 country-year medians from 40 countries over the ten-year period from 1992 to 2001. The dependent variable, r_{AVG}, is the mean of the four estimates for the implied cost of equity capital as described in the Appendix. In the last two columns, we replace r_{AVG} with r_{WEIGHT}, the country-year mean of the four accuracy-weighted cost of capital estimates. We use the absolute one-year-ahead analyst forecast error scaled by forecast-period stock price as weights. DISREQ measures the level of disclosure regulation based on an index of disclosure requirements in securities offerings. SECREG captures the strength of securities regulation mandating and enforcing disclosures. It is measured as the mean of the disclosure index, the liability standard index and the public enforcement index. All securities regulation variables stem from La Porta et al. [2005]. LAW represents the overall quality of the legal system and is measured as the rule of law index (divided by 10) from La Porta et al. [1997]. INFL is the yearly median of country-specific, one-year-ahead realized monthly inflation rates. SIZE stands for US$ market value of outstanding equity (in thousands). We use the natural log of the size variable in the analysis. BETA is measured relative to the MSCI world market index and estimated based on a five-year market model regression. We require at least 24 monthly observations to calculate market beta. RETVAR is the returns variability computed as annual standard deviation of monthly stock returns. BMR is the ratio of accounting book value to market value of equity. Financial data are measured as of the fiscal-year end, BETA, RETVAR and the cost of capital as of month +10. MACVAR is the first principal component of four proxies for macroeconomic variability, i.e., country-level earnings variability and ROE variability, volatility in GDP growth rates, and exchange rate variability. See Table 3 for details. FBIAS is the one-year-ahead analyst forecast error (mean forecast minus actual) scaled by forecast-period stock price. Industry controls, i.e., the percentage of firms in an industry per country-year based on the classification in Campbell [1996], and year indicators are included in the regressions but not reported. The table reports OLS coefficient estimates and t-statistics based on Newey-West heteroscedasticity and autocorrelation corrected standard errors (in parentheses). **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 5

Influence of Long-Run Growth Differences across Countries on the Relation between the Implied Cost of Capital and Securities Regulation

		Alternate Growth in Terminal Value Assumptions		*Accounting ROA as Control*	*Simultaneous Estimation of Cost of Capital and Growth*			*Growth as Control*	*Non-Accounting-Based Cost of Capital Estimates*	
Variable	*Predicted Sign*	Model 1: r_{AVG}	Model 2: r_{AVG}	Model 3: r_{AVG}	Model 4: r_{Easton}	Model 5: r_{ETSS}	Model 6: r_{ETSS}	Model 7: r_{AVG}	Model 8: r_{DIV}	Model 9: r_{CRED}
N		358	352	358	330	330	330	358	358	355
DISREQ	–	-0.034 **	-0.021 *	-0.045 **	-0.028	-0.014	-0.023 *	-0.051 **	-0.019 *	-0.040 **
		(-3.03)	(-2.15)	(-3.89)	(-1.49)	(-1.53)	(-2.20)	(-4.29)	(-2.28)	(-4.37)
LAW	–	-0.026 #	-0.037 **	-0.026 #	-0.014	-0.007	-0.027 *	-0.015	0.008	-0.155 **
		(-1.89)	(-2.68)	(-1.87)	(-0.63)	(-0.67)	(-2.52)	(-1.18)	(1.03)	(-13.31)
INFL	+	0.155 *	0.246 **	0.238 **	0.250 *	0.101 #	0.132 *	0.211 **	-0.060 #	0.048
		(2.32)	(3.46)	(3.62)	(2.30)	(1.79)	(2.25)	(3.00)	(-1.67)	(0.72)
SIZE	–	-0.004 #	-0.003	-0.003	-0.010 *	-0.006 **	-0.008 **	-0.002	0.000	-0.007 **
		(-1.79)	(-1.34)	(-1.05)	(-2.24)	(-2.60)	(-3.44)	(-0.82)	(-0.01)	(-2.72)
RETVAR	+	0.140 *	0.112 #	0.129 *	0.120	0.151 *	–	0.087	0.036	0.135 *
		(2.10)	(1.67)	(2.02)	(0.98)	(2.39)		(1.39)	(0.87)	(2.28)
BMR	+	0.026 **	0.016 *	0.035 **	0.001	-0.002	-0.009	0.038 **	0.008	-0.006
		(3.60)	(2.45)	(4.38)	(0.09)	(-0.33)	(-1.40)	(4.72)	(1.50)	(-0.70)
MACVAR	+	0.002	0.004 **	0.002 #	0.007 **	0.003 #	–	0.000	0.000	0.000
		(1.62)	(2.88)	(1.68)	(2.96)	(1.81)		(0.00)	(0.53)	(-0.03)
FBIAS	+	0.122	0.124	0.142	0.042	-0.086	–	0.145 #	–	–
		(1.37)	(1.43)	(1.59)	(0.28)	(-1.27)		(1.76)		
ROA		–	–	0.228 **	–	–	–	–	–	–
				(3.10)						
BETA		–	–	–	–	–	0.014 **	–	–	–
							(4.48)			
g_{ETSS}		–	–	–	–	0.479 **	0.522 **	0.453 **	–	–
						(6.61)	(7.29)	(4.31)		
g_{Easton}		–	–	–	0.280 **	–	–	–	–	–
					(12.38)					
$g_{Forecasts}$		–	–	–	–	–	–	–	-0.026	–
									(-1.35)	
Intercept, Industry and Year Controls		included	included	included	included	included	included	included	included	included
R^2		57.7%	65.2%	63.8%	74.9%	77.9%	78.1%	68.0%	35.0%	87.8%
F-Stat		11.0 **	15.0 **	14.3 **	21.8 **	25.6 **	29.0 **	14.3 **	6.1 **	52.7 **

(continued)

The sample comprises a maximum of 358 country-year medians from 40 countries over the ten-year period from 1992 to 2001. We report country-level regressions of various cost of equity capital estimates on the full set of controls plus the institutional factors (except in Model 6, where we use a reduced model including beta factors, and Models 8 and 9, where we exclude forecast bias). See Table 4 for details. We use the following dependent variables: (1) the mean of the four estimates for the implied cost of equity capital, r_{AVG}, as described in the Appendix, (2) r_{ETSS} derived from regressing forecasted return on equity on the price-to-book ratio as described in Easton et al. [2002], (3) r_{Easton} derived from regressing forecasted two-year-ahead cum-dividend earnings on one-year-ahead earnings as described in Easton [2004], (4) dividend yield, r_{DIV}, measured as actual dividends paid during the last fiscal year scaled by stock price as of month +10, and (5) expected returns , r_{CRED}, derived from regressing country index returns on country credit-risk ratings, as described in Erb et al. [1996a]. We assess the influence of countries' long-run growth differences as follows: in Model 1 the implied cost of capital estimates are computed based on a uniform perpetual growth rate of 3%. In Model 2 we use the country-specific annual change in real GDP plus the country's long-run inflation rate instead. In Model 3 we include accounting profitability, ROA, measured by return on total assets as additional control for accounting differences. In Models 4 through 7 we include long-run estimates in residual income growth, g_{ETSS}, or in abnormal earnings growth, g_{Easton}, as additional control. The two growth measures are estimated on the country-year level (country-level for Model 7) as described in Easton et al. [2002] and Easton [2004], respectively. See also Section 4.2 for further explanations. In Models 8 and 9 we employ alternative, non-accounting-based cost of equity capital models (Model 8 also includes earnings growth, $g_{Forecasts}$, measured as country-year median of one-year-ahead percentage changes in analyst earnings per share forecasts as additional control). DISREQ measures the level of disclosure regulation based on an index of disclosure requirements in securities offerings (La Porta et al. [2005]). LAW represents the overall quality of the legal system and is measured as the rule of law index (divided by 10) from La Porta et al. [1997]. The table reports OLS coefficient estimates and t-statistics based on Newey-West heteroscedasticity and autocorrelation corrected standard errors (in parentheses). **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 6

Sensitivity Analyses for the Relation between the Implied Cost of Capital, Securities Regulation and Legal Quality

		Model 1		*Model 2*	
Variable	*N*	*DISREQ*	*LAW*	*SECREG*	*LAW*
Panel A: Potential Correlated Omitted Variables					
Per-capita GDP	358	-0.033 **	-0.031	-0.023 *	-0.027
		(-2.62)	(-1.55)	(-2.30)	(-1.39)
Equity market size	357	-0.040 **	-0.029 *	-0.031 **	-0.018
		(-3.44)	(-2.07)	(-2.81)	(-1.38)
Disclosure practice	358	-0.029 #	-0.024	-0.019 #	-0.014
		(-1.94)	(-1.33)	(-1.88)	(-0.97)
Tax rates	352	-0.042 **	-0.026 *	-0.029 **	-0.015
		(-4.03)	(-2.05)	(-2.84)	(-1.23)
Panel B: Risk-free Rates and Risk Premia					
Local risk-free rates	358	-0.025 *	-0.027 #	-0.018 #	-0.019
		(-2.18)	(-1.77)	(-1.74)	(-1.38)
Risk premia instead of r_{AVG}	358	-0.034 **	0.005	-0.028 *	0.014
		(-2.56)	(0.25)	(-2.16)	(0.79)
Panel C: Alternative Specifications and Estimation Methods					
Accounting-based controls only	358	-0.044 **	-0.054 **	-0.035 **	-0.041 **
		(-4.37)	(-4.05)	(-3.39)	(-3.23)
Accounting-based controls & 2SLS	358	-0.028 **	–	-0.038 **	–
		(-2.59)		(-2.64)	
Panel-corrected standard errors	358	-0.047 #	-0.045 *	-0.039 *	-0.034 **
		(-1.94)	(-2.37)	(-2.19)	(-2.59)
Years 2000 & 2001 only	73	-0.071 **	-0.054 #	-0.057 *	-0.028
		(-2.87)	(-1.67)	(-2.02)	(-1.04)
Panel D: Country-Fixed Effects from Firm-Level Regressions					
Full model	40	-0.030 #	-0.055 **	-0.034 #	-0.057 **
		(-1.71)	(-4.06)	(-1.77)	(-4.18)
R^2			34.5%		34.8%
Full model plus ΔGDP	40	-0.035 #	-0.052 **	-0.038 #	-0.054 **
		(-1.87)	(-3.53)	(-1.87)	(-3.73)
R^2			35.6%		35.6%

The sample comprises 358 country-year medians from 40 countries over the ten-year period from 1992 to 2001. The dependent variable, r_{AVG}, is the mean of the four estimates for the implied cost of equity capital as described in the Appendix. Panel A reports only the coefficients (Newey-West t-statistics) of the securities regulation (DISREQ or SECREG) and legal quality (LAW) variables, but the full set of controls (see Model 3 in Table 3) plus one of the following variables are included: (1) the natural log of per-capita GDP (in constant 1995 US$), (2) the size of equity markets as measured by the overall market capitalization in percent of GDP, (3) a country's disclosure practice using the country means of the 1995 CIFAR firm-level disclosure scores, (4) the average personal top income tax rate and the average top corporate tax rate over the sampling period. Panel B reports regressions (1) replacing the inflation control with the yearly median of the nominal local yields on short-term treasury bills, central bank papers or inter-bank loans, as reported in Datastream, and (2) using inflation-adjusted risk premia instead of the raw cost of capital estimates as the dependent variable. Panel C presents the coefficients (Newey-West t-statistics) using (1) accounting-based controls only, i.e., the natural log of US$ total assets (in thousands) for SIZE and the ratio of total liabilities to total assets (financial leverage) instead of RETVAR and BMR, (2) accounting-based controls and two-stage-least-squares estimation with the English legal origin as instrument (where appropriate), (3) Prais-Winsten regressions with panel-corrected standard errors (Beck and Katz [1995]), and (4) observations from the years 2000 and 2001 only. For Panel D, we first estimate firm-level regressions of r_{AVG} on the full set of firm-specific controls and inflation (see Model 4 in Table 3) plus country-fixed effects. We subsequently regress these country-fixed effects on the institutional variables (and ΔGDP, growth in per-capita GDP). **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 7

Cost of Capital Effects of Securities Regulation Conditional on the Integration of Capital Markets

$$r_{AVGit} = \alpha_0 + \alpha_1\text{Securities Regulation}_i + \alpha_2\text{Securities Regulation}_i\text{*Integration}_{it} + \alpha_3\text{Integration}_{it} + \alpha_4\text{LAW}_i + \alpha_5\text{INFL}_{it} + \alpha_6\text{SIZE}_{it} + \alpha_7\text{RETVAR}_{it} + \alpha_8\text{BMR}_{it} + \alpha_9\text{MACVAR}_{it} + \alpha_{10}\text{FBIAS}_{it} + \Sigma\alpha_j\text{Industry Controls}_{it} + \Sigma\alpha_k\text{Year Controls}_t + \varepsilon_{it}$$

		Integration measured by			
		MSCI Developed Markets Index (DEV)		*Portfolio In- and Outflows in Percent of GDP (FLOW)*	
Variable	*Pred. Sign*	*DISREQ*	*SECREG*	*DISREQ*	*SECREG*
Panel A: Country-Year Regressions (N = 358)					
Securities Reg.	–	-0.050 **	-0.100 **	-0.053 *	-0.091 **
		(-2.80)	(-4.06)	(-2.38)	(-3.70)
Securities Reg.* Integration	+	0.029	0.102 **	0.034	0.093 **
		(1.39)	(3.90)	(1.20)	(3.44)
Integration	–	-0.027	-0.063 **	-0.022	-0.052 **
		(-1.53)	(-3.69)	(-0.86)	(-2.65)
Legal Quality	–	-0.018	-0.022	-0.028	-0.026 #
		(-1.03)	(-1.32)	(-1.37)	(-1.64)
Risk, Industry, and Year Controls		included	included	included	included
H_0: $\alpha_1+\alpha_2 = 0$ (p-value)		0.056	0.821	0.128	0.850
Panel B: Country-Fixed Effects Regressions (N = 40)					
Securities Reg.	–	-0.050 *	-0.127 **	-0.050 *	-0.092 **
		(-2.00)	(-4.12)	(-2.01)	(-3.06)
Securities Reg.* Integration	+	0.039	0.133 **	0.046	0.103 *
		(1.11)	(3.61)	(1.20)	(2.45)
Integration	–	-0.022	-0.070 **	-0.024	-0.049 #
		(-0.84)	(-3.10)	(-0.87)	(-1.98)
Legal Quality	–	-0.060 *	-0.065 **	-0.060 **	-0.066 **
		(-2.40)	(-2.99)	(-2.96)	(-3.51)
H_0: $\alpha_1+\alpha_2 = 0$ (p-value)		0.665	0.785	0.889	0.658

The sample comprises 358 country-year medians from 40 countries over the ten-year period from 1992 to 2001. The dependent variable, r_{AVG}, is the mean of the four estimates for the implied cost of equity capital as described in the Appendix. In Panel A, the table reports only the main and interaction effects of market integration (DEV or FLOW) and the securities regulation variables (DISREQ or SECREG) and the legal quality variable (LAW), but the full set of controls is included. See Model 3 in Table 3 for details. For Panel B, we first estimate firm-level regressions using the full set of firm-specific controls and inflation (see Model 4 in Table 3) plus country-fixed effects. We subsequently regress these country-fixed effects on the main and interaction effects of market integration and the institutional variables. We adopt two binary measures of capital market integration: (1) DEV is equal to one if the country's equity market is classified as developed according to the MSCI database, (2) FLOW is equal to one for countries with above-median portfolio inflows and outflows in percent of GDP, as reported by the IMF for 2001. DISREQ measures the level of disclosure regulation based on an index of disclosure requirements in securities offerings. SECREG captures the strength of securities regulation mandating and enforcing disclosures. All securities regulation variables stem from La Porta et al. [2005]. LAW represents the general quality of the legal environment and is measured as the rule of law index (divided by 10) from La Porta et al. [1997]. The table reports OLS coefficient estimates and, in parentheses, t-statistics (based on Newey-West heteroscedasticity and autocorrelation corrected standard errors for Panel A). It also reports p-values from a Wald test indicating statistical significance of the effects of securities regulation in integrated countries. **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.