Cost of Capital Effects and Changes in Growth Expectations around U.S. Cross-Listings

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Abstract

This paper examines whether cross-listing in the U.S. reduces foreign firms' cost of capital. While prior studies document that U.S. cross-listings are associated with substantial increases in firm value, the sources of these valuation effects are not well understood. We estimate the cost of capital effects implied by market prices and analyst forecasts, which allows us to explicitly account for changes in growth expectations around cross-listings. We find strong evidence that firms with cross-listings on U.S. exchanges experience a decrease in their cost of capital, which is economically significant and sustained. Consistent with the bonding hypothesis, we document that these effects are larger for firms from countries with weaker institutional structures. Cross-listings in the over-the-counter market are associated with minor reductions in firms' cost of capital, and private placements seem to have adverse effects. We also document positive valuation effects for all types of U.S. cross-listings stemming from changes in (financial analysts') growth expectations. The latter result could reflect that firms seek to cross-list when their growth opportunities happen to expand.

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

Key Words: *Corporate Governance, Cross-listing, Bonding hypothesis, Cost of equity, Disclosure, Law and finance, International finance*

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1. Introduction

There is mounting evidence that countries' institutional frameworks play an important role for access to finance and equity valuations (e.g., La Porta et al, 1997 and 2002). In light of this evidence, cross-listing in the U.S. has been suggested as a way for firms from countries with poor institutions to privately overcome these shortcomings (Coffee, 1999; Stulz, 1999). Consistent with this notion, several studies document that cross-listings have significant effects on firms' market values, using either event-study returns (e.g., Foerster and Karolyi, 1999; Miller, 1999; Lee, 2004) or comparisons with firms that are not cross-listed (e.g., Doidge, 2004; Doidge et al., 2004). This evidence suggests that U.S. cross-listings offer substantial benefits. However, the sources of these benefits are not yet well understood (e.g., Leuz, 2003; Doidge et al., 2004).

One important question is whether and to what extent cross-listing in the U.S. affects firms' cost of capital. The bonding argument suggests that a U.S. cross-listing strengthens outside investor protection making it easier for firms to raise external finance (e.g., Reese and Weisbach, 2002; Benos and Weisbach, 2004; Doidge et al., 2004). Moreover, listings on NASDAQ, NYSE or AMEX require foreign firms to comply with SEC disclosure rules, which typically imply a substantial increase in disclosure and could manifest in a lower cost of capital (e.g., Verrecchia, 2001; Lambert et al., 2006). Similarly, U.S. cross-listings can improve investor recognition and enlarge a firm's investor base (e.g., Merton, 1987; Foerster and Karolyi, 1999).

Alternatively, the documented valuation effects of U.S. cross-listings can reflect concurrent changes in firms' growth opportunities that neither stem from the cross-listing per se nor are the result of changes in the cost of capital lowering the firm's hurdle rate for investments. That is, firms may seek to cross-list when they happen to experience an expansion in their growth opportunities. Moreover, Foerster and Karolyi (1999 and 2000) and Miller (1999) provide evidence of long-run underperformance after cross-listing in the U.S., which raises the question of whether the documented valuation benefits are in fact sustained in the long-run. Similarly, the alleged intention of many foreign firms to delist to escape the burden of the Sarbanes-Oxley Act also questions the existence of sustained and sizeable cross-listing benefits, e.g., stemming from a reduction in the cost of capital.[1] Thus, it is an open and topical question whether U.S. cross-listings persistently reduce the cost of capital.

To shed light on these issues and the mechanism by which cross-listings affects firms' valuations, we analyze the cost of capital effects of U.S. cross-listings. We use a novel approach to compute ex-ante estimates of firms' cost of equity capital implied by market prices and analyst forecasts. This approach explicitly accounts for changes in the market's growth expectations around cross-listings. It also allows us to separately gauge the magnitude of these cash flow (or growth) effects on firms' valuations.

Our analysis is based on a large panel of more than 31,000 firm-year observations from 44 countries over the period from 1992 to 2003. We collect a comprehensive sample of over 1,000 U.S. cross-listings and classify them into exchange listings, over-the-counter (OTC) listings and private placements, accounting for the different regulatory consequences. For an exchange listing, firms have to register with the SEC and file Form 20-F, which requires extensive disclosures and a reconciliation of foreign financial statements to U.S. GAAP. In addition, firms are subject to SEC oversight and bear the threat of U.S. securities litigation. Cross-listings in the OTC markets do not require a 20-F filing, but a registration statement using Form F-6 and home-country disclosures to the SEC. It also subjects firms to the Foreign Corrupt Practices Act, under

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1 For anecdotal evidence on foreign firms' delisting intentions see, e.g., Parker (SEC to review reporting rules for overseas companies, *The Financial Times*, March 14, 2006).

which most SEC enforcement actions are brought (Karpoff et al., 2004). Private placements under Rule 144A do not require SEC registration or any additional (public) disclosures.

Based on this classification, we hypothesize that, if cross-listings reduce firms' cost of capital, the effects are strongest for exchange listings. Moreover, it is not clear that private placements should experience any reduction. Consistent with these hypotheses, we find strong evidence that cross-listings on NYSE, AMEX, or NASDAQ significantly reduce firms' cost of equity capital and that the effects are larger than for the other types of cross-listings. We obtain these results from cross-sectional regressions including firm-fixed effects as well as from difference-in-differences analyses, mitigating concerns about unobserved heterogeneity and selection bias. Our findings suggest an average reduction in the cost of capital between 70 and 120 basis points, which is economically significant, but not too large to be implausible. We also find evidence that cross-listings in the U.S. OTC markets reduce firms' cost of capital. The estimated effect is smaller, i.e., on average between 30 and 70 basis points, and not as robust as the exchange listing effects. Interestingly, in some of our analyses, U.S. private placements are associated with an increase in the cost of capital. This result is consistent with the findings in Miller (1999) and Doidge et al. (2004) as they also document opposite valuation effects for private placements. One possible explanation for our finding is that private placements involve private communication with a small group of institutional investors, which may exacerbate information asymmetries among traders.

The rank order of the cost of capital effects (from exchange listings to private placements) suggests that the regulatory consequences of U.S. cross-listings play an important role, which is consistent with the bonding hypothesis. Further corroborating this notion, we find that the reduction in the cost of capital for exchange listings is larger for firms from countries with weak institutional structures, i.e., less extensive disclosure regulation and weak investor protection.

We conduct extensive robustness checks to validate our findings. We use four different implied cost of capital models and obtain very similar results for each of them. We also gauge the sensitivity of our findings with respect to key model assumptions, in particular those about long-run growth. One potential concern is that cross-listed firms have different long-run growth expectations than non-cross listed firms. To address this issue, we implement our models with long-run growth estimates that vary by cross-listing status, country and year, and obtain very similar results. Further, we check that the results are not unduly affected by differences in the bias and accuracy of analyst forecasts across firms and countries.

Our last set of analyses exploits the fact that we have explicit forecasts for firm growth. Using these forecasts, we decompose the two-year return leading up to the cross-listing into two components, one due to changes in the cost of capital and one due to changes in expected future cash flows. Using this approach, we document substantial valuation effects from changes in growth expectations for all three cross-listing types, indicating that not all valuation effects around cross-listings are attributable to a reduction in the cost of capital. Instead, they could in part reflect that firms seek U.S. cross-listings when their growth opportunities happen to expand.

This study makes several contributions. First, we use a novel approach to provide strong evidence that cross-listings on U.S. exchanges are associated with a statistically significant, yet economically plausible reduction in the cost of capital. Related studies produce estimates based on realized stock returns (or dividend yields) that are too large to be attributable solely to a reduction in the cost of capital. For instance, Errunza and Miller (2000) document an average effect of over 1,000 basis points, and the findings in Sarkissian and Schill (2006) amount to an effect of 590 basis points. Both of these findings are difficult to reconcile with the valuation effects around cross-listings and probably reflect the difficulty of estimating cost of capital effects from realized stock returns or dividend yields. To compute expected returns from realized returns, one needs fairly long time series, yet such data are not available for many cross-listed stocks. Moreover, cross-listings are major corporate events making it particularly difficult to obtain equilibrium estimates for expected returns (e.g., Sarkissian and Schill, 2006), especially considering that cross-listings change firms' exposures to the global market portfolio (Foerster and Karolyi, 1999). Thus, we view our work as a step towards resolving the considerable uncertainty about the existence and magnitude of the cost of capital effects of U.S. cross-listings.

Second, we provide evidence that cross-listings are also associated with valuation effects stemming from changes in cash flow (or growth) expectations. As a result, comparing market values, Tobin's q or realized stock returns around U.S. cross-listings is likely to pick up these growth effects, together with any cost of capital consequences, and hence may overstate the long-run valuation benefits of cross-listing per se. For instance, the valuation benefits are unlikely to be sustained if the documented changes in growth expectations reflect that firms seek cross-listings when they experience a shock to their growth opportunities. Changes in the cost of capital, in contrast, are expected to persist. Our evidence supports this conjecture. Thus, understanding the sources of the cross-listing benefits is important, and it is critical when debating whether additional regulatory costs (e.g., due to the Sarbanes-Oxley Act of 2002) induce foreign firms to delist or cross-list elsewhere.

Third, our analysis is based on one of the largest panels of U.S. cross-listings. Prior studies are generally based on smaller samples, which are typically constructed as of one point in time. Our panel dataset not only mitigates survivorship bias, but also facilitates fixed-effects and difference-in-differences estimation, mitigating concerns about selection bias and endogeneity.

Finally, we provide evidence that the cost of capital effects differ systematically across regulatory consequences of different types of American Depositary Receipts (ADRs) and across firms from different home countries. This cross-sectional evidence corroborates the bonding hypothesis (Coffee, 1999; Stulz, 1999), which several recent studies have questioned (e.g., Siegel, 2005; Gozzi et al., 2006). At the same time, our results also suggest that the benefits of bonding are more subtle and not simply equal to the valuation effects around cross-listing.

The remainder of the paper is organized as follows. Section 2 reviews the literature and develops our hypotheses. In Section 3, we describe the sample and the construction of the implied cost of capital estimates. Section 4 presents the main cross-sectional and difference-in-differences results including robustness checks. In Section 5, we document that cost of capital effects differ depending on firms' home-country legal institutions and analyze the relative magnitude of cash flow and cost of capital effects. Section 6 concludes.

2. Prior Research and Hypothesis Development

Early studies on U.S. cross-listings are built on international asset pricing models and view cross-listing as a mechanism to overcome market segmentation and barriers to international investment (e.g., Karolyi, 1998; Karolyi and Stulz, 2003). The idea is that if a country's capital market is not fully integrated with international capital markets, firms face a higher cost of capital because risk is mostly borne by investors from this country. Cross-listings make it easier for foreign investors to hold shares in these firms and, as a consequence, risk is more widely shared. Thus, cross-listed firms should have a lower cost of capital and higher valuations and the announcement of cross-listings should result in positive stock returns.

Although some of the evidence is consistent with the segmentation hypothesis (e.g., Foerster and Karolyi, 1999; Miller, 1999), several recent studies question the extent to which market integration alone can explain the cross-listing effects (Doidge et al., 2004; Karolyi, 2006). First, as investment barriers have decreased over time, we should see fewer and fewer cross-listings. However, cross-listing behavior exhibits the opposite trend (Karolyi, 2004 and 2006). Similarly, we would expect to see cross-listings primarily from countries where risk sharing benefits and diversification gains are the largest. However, Sarkissian and Schill (2004) and Lee (2004) show that this hypothesis is not supported. Finally, the segmentation argument is – with few exceptions – not specific to U.S. cross-listings and does not predict that different types of ADRs have differential effects. However, cross-listings on U.S. exchanges (i.e., NYSE, AMEX or NASDAQ) generally have much stronger return effects than either cross-listings in other countries (e.g., Sarkissian and Schill, 2006) or cross-listings in other U.S. markets, such as the Pink Sheets or private placements (e.g., Foerster and Karolyi, 1999; Doidge et al., 2004).

A more promising explanation for the strong effects of U.S. cross-listings – particularly on exchanges – is the bonding hypothesis (Coffee, 1999; Stulz, 1999).[2] It explicitly recognizes the legal consequences of U.S. cross-listings. The idea is that U.S. disclosure requirements, exposure to SEC enforcement, and the threat of shareholder litigation make it harder and more costly for controlling owners and managers to extract private control benefits from outside investors. Thus, cross-listing in the U.S. provides a means to controlling insiders from countries with weak governance structures to credibly commit not to expropriate outside investors. Such bonding is particularly valuable to firms with large growth opportunities and external financing needs.

Several recent studies support this argument.[3] Reese and Weisbach (2002) show that, after cross-listing in the U.S., firms raise more external capital, but mostly in their home markets (also Benos and Weisbach, 2004; Lins et al., 2005). Doidge et al. (2004) document that firms with U.S. cross-listings exhibit a valuation premium relative to non cross-listed firms and that the premium is most pronounced for U.S. exchange listings. They also show that the valuation

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2 Fuerst (1998) provides a related signaling argument based on the idea that U.S. disclosure requirements and legal liability make cross-listing costly, allowing "good" firms to separate from the "bad" firms.

3 But there are also studies that question the effectiveness of cross-listing as a bonding device, or at least, the legal aspects of it. See, e.g., Siegel (2005).

effects are stronger for firms from countries where investor protection is weaker.[4] Doidge (2004) documents that voting premiums, which are a proxy for private control benefits, are lower for cross-listed firms and that this difference is larger for firms from countries with poor outside investor protection. Finally, Lee (2004) shows that announcements of U.S. cross-listings are associated with negative abnormal returns for competitors of the announcing firm, consistent with the notion that cross-listings improve firms' ability to exploit their growth opportunities.

The bonding hypothesis suggests that U.S. cross-listings should increase market value but there are several ways in which it may do so. On one hand, bonding should decrease the amount of outsider expropriation and hence increase investors' expectations about future cash flows. On the other hand, bonding should improve firms' ability to raise capital and lower its cost of capital, which in turn should expand the set of profitable investment opportunities.[5]

In addition, cross-listing on a U.S. exchange generally commits foreign firms to disclosure rules that are more extensive and more strictly enforced than in their home country. Consistent with this claim, empirical studies find significant changes in firms' information environment around U.S. cross-listings indicating an increase in disclosure quality (e.g., Lang et al., 2003a and 2003b; Bailey et al., 2006). These information effects are likely to reduce information asymmetries and lower firms' cost of capital (Verrecchia, 2001; Easley and O'Hara, 2004; Lambert et al., 2006).[6] A related argument is that cross-listing broadens a firm's investor base.

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4 Similarly, La Porta et al. (2002) and Durnev and Kim (2005) document that firms in countries with stronger corporate governance and investor protection enjoy higher valuations. Moreover, Doidge et al. (2006) provide evidence suggesting that private benefits of control are an important reason why controlling insiders shy away from cross-listings on U.S. exchanges.

5 The models in Lombardo and Pagano (2002) and Lambert et al. (2006) also support the notion that less expropriation can manifest in the cost of capital.

6 There are a number of empirical studies supporting the link between disclosure and the cost of capital (e.g., Botosan, 1997; Leuz and Verrecchia, 2000; Hail, 2002; Francis et al., 2004). In addition, Hail and Leuz (2006) provide evidence that strong securities regulation is associated with a lower cost of capital.

Based on Merton (1987), this effect can reduce a firm's cost of capital. Several recent studies provide evidence that a U.S. cross-listing improves a firm's recognition in the media and by financial analysts and that it broadens the investor base (Baker et al., 2002; Lang et al., 2003a; Ammer et al., 2006).

Thus, although there are several good economic reasons why U.S. cross-listings should reduce firms' cost of capital, there is relatively little empirical evidence.[7] The return-based results in Errunza and Miller (2000) and more recently in Sarkissian and Schill (2006) are too large to be solely attributable to cost of capital effects and, while several of the valuation and event studies discussed above are consistent with a reduction in the cost of capital, their results neither provide direct evidence that U.S. cross-listings reduce firms' cost of capital nor do they shed little light on the mechanism by which cross listing affects firms' valuations.

An alternative explanation for valuation effects around cross-listings is that firms choose to cross-list when they happen to experience an increase in their growth opportunities. Under this explanation, the valuation benefits do not stem from cross-listing per se and also should not manifest in a lower cost of capital. It is also not clear whether these valuation benefits are sustained (e.g., Miller, 1999; Gozzi et al., 2006; Sarkissian and Schill, 2006). Moreover, we do not expect the growth effects to exhibit the same rank order as predicted by the bonding hypothesis. If anything, it seems plausible that growth shocks lead to new financing needs and hence that we see larger growth-related valuation effects for cross-listings that typically involve raising new equity capital, i.e., private placements and Level III ADRs.[8]

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7 We do not attempt to separate the different explanations for a decrease in the cost of capital. We pursue the important first step to isolate and document effects on the cost of capital.

8 It is also possible that the growth effects do not differ across ADR types because firms can raise capital in their home country, regardless of ADR type. See Reese and Weisbach (2002). Note also that the bonding hypothesis and the growth explanation are not competing in the sense that they can simultaneously be present.

To shed light on these issues, we first examine whether U.S. cross-listings have an effect on firms' cost of capital, controlling for changes in (analysts') expectations about future cash flows. Based on the bonding and disclosure arguments above, we predict that cross-listing reduces firms' cost of capital and that the effects are most pronounced for cross-listings on U.S. exchanges because they bear the strongest legal and regulatory consequences. As OTC listings and private placements have only weak regulatory consequences, we expect to see more modest cost of capital effects based on the segmentation hypothesis. We also examine cross-sectional differences in the cross-listing effects and predict that the cost of capital effects are stronger for firms from countries with less integrated markets, weaker legal institutions and less extensive disclosure regulation.

Next, we analyze whether there are significant cash flow effects driven by changes in growth expectations. Towards this end, we decompose long-run stock returns leading up to the cross-listing into two components: one measuring the effect of changes in cash flow (or growth) expectations and one measuring the effect of changes in the cost of capital. However, as any reduction in the cost of capital should expand the set of profitable investment projects, it is difficult to completely disentangle the two effects. For this reason, we view this part of our analysis as exploratory in nature and as attempt to shed further light on the mechanism through which cross-listings affect firm value. With this caveat in mind, we examine the relative magnitudes of the cash flow and cost of capital effects and also compare them across ADR types. If the changes in growth expectations are primarily driven by bonding, we expect the same rank order as for the cost of capital effects. However, under the alternative explanation, we do not expect such a rank order and, if anything, it seems plausible that ADRs associated with capital-raising exhibit larger cash flow effects.

3. Research Design and Data

3.1 Estimating the Implied Cost of Capital

Accurately measuring differences and changes in firms' cost of capital is difficult. Prior studies use realized returns or dividend yields as proxies for firms' cost of capital (e.g., Foerster and Karolyi, 1999 and 2000; Errunza and Miller, 2000). However, these proxies also capture changes in market expectations about firms' future cash flows (e.g., Bekaert and Harvey, 2000). Standard techniques to obtain unbiased estimates of expected returns from realized stock returns therefore require fairly long time-series (e.g., Stulz, 1999).[9] Moreover, cross-listings are major corporate events making it particularly difficult to obtain equilibrium estimates of expected returns (Sarkissian and Schill, 2006). Finally, international asset pricing models require assumptions regarding the degree of market segmentation and exposure to the global market portfolio, both of which are likely to change around the cross-listing (e.g., Karolyi, 1998).

Due to these difficulties, we adopt an alternative approach to estimate the cost of capital effects of cross-listings. We employ models that estimate the ex ante rate of return implied in contemporaneous stock price and analyst forecast data. This approach does not rely on extensive time-series of past return data and does not require any a priori assumptions about the degree of market integration. Furthermore, implied cost of capital models are an explicit attempt to separate cash flow (or growth) effects from cost of capital effects.[10]

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9 Separating cash flow and cost of capital effects on dividend yields requires explicit estimates about future growth in dividends.

10 We hasten to add that our results should be viewed as complementing, rather than competing with prior return-based studies. Moreover, implied cost of capital models are also subject to several methodological and practical drawbacks. For instance, we need to assume that consensus analyst forecasts are reasonable proxies for the market's expectations of future earnings. Prior evidence suggests that this might not always be the case (e.g., Frankel and Lee, 1998). For this reason, we provide robustness checks controlling for differences in analyst forecast properties and future expected growth opportunities. Finally, data requirements limit our sample in the main analyses to firms that are covered by financial analysts providing forecasts for at least three periods ahead. But note that we also gauge the impact of this data requirement in our sensitivity analyses (Section 4.2).

We adopt four implied-cost-of-capital models suggested in the literature (Claus and Thomas, 2001; Gebhardt et al., 2001; Easton, 2004; Ohlson and Juettner-Nauroth, 2005, as implemented by Gode and Mohanram, 2003). All four models are consistent with the discounted dividend valuation model but exploit the accounting structure to obtain an equivalent earnings-based representation. The first two models 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 is to substitute market price and analyst forecasts into the 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. This rate of return is an ex-ante estimate of cost of equity capital, and it is obtained controlling for market expectations about future growth.

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, the incorporation of industry effects, and how inflation is integrated into the steady-state terminal value. In the Appendix, we summarize the four models, describe their key assumptions and data requirements, and explain the implementation of our firm-level cost of capital estimates in an international setting. As there is little consensus in the literature, which of the models works best or even how to evaluate the models (e.g., Botosan and Plumlee, 2005; Guay et al., 2005), we average over the four proxies and use the resulting mean estimate as the dependent variable (COC) in our main analyses.[11] In our sensitivity tests, we also use the first principal component of the four COC estimates and report results for the four measures individually.

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[11] Moreover, Easton and Monahan (2005) point to substantial measurement error in implied cost of capital estimates. The evidence in Guay et al. (2005) also raises this concern, but at the same time shows that implied

3.2 Data, Sample Selection and Descriptive Statistics

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 2003, except from Canada and the U.S., and match them to firms covered in I/B/E/S. Canadian firms are excluded because they can directly list their shares, rather than depository receipts, in the U.S. Moreover, Canadian firms are exempted from certain U.S. reporting requirements under the Multi-Jurisdictional Disclosure System.[12]

To be included in the cost of capital computation, 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.[13] All data are measured in local currency and taken as of month +10 after the fiscal-year end. 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.[14] To calculate our average cost of capital estimate, COC, we further require that each firm-year has all four individual implied cost of capital proxies available.

Although both Worldscope and I/B/E/S provide split-adjusted data and hence should be compatible, we conduct further consistency checks. We calculate each of the following three variables involving data from both sources and eliminate the first and 99th percentile of the

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estimates are not worse than realized returns and in some cases perform better. Recently, Pastor et al. (2006) show that implied estimates are useful in capturing time-series variation in expected returns. Given this debate about the validity of implied estimates, our results should be interpreted carefully and as complementing prior return-based studies.

12 See King and Segal (2004) for an analysis of U.S. cross-listings of Canadian firms. If we include Canadian firms in our sample, the results remain largely the same and the inferences do not change.

13 We later relax these data requirements and conduct sensitivity analyses including firms with missing or negative analysts' earnings forecasts.

14 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 the Appendix for further details.

respective distributions: (1) the I/B/E/S price to Worldscope book value of common equity per share ratio, (2) the realized one-year earnings per share growth rate computed using next year's earnings per share from I/B/E/S and this year's earnings per share from Worldscope, and (3) the ratio of book value per share as reported in Worldscope to book value per share calculated using the number of shares outstanding from I/B/E/S. These common sense filters should ensure that the subsequent results are not driven by a mismatch in the underlying input data.

We further delete 1% of extreme realizations for all firm-level attributes (dependent and independent variables), except firm size where we use the natural logarithm in the analyses. Finally, we eliminate firm-years from countries with no ADR observations, and if the inflation rate for the country in a particular year is above 25%. Taken together, these restrictions result in a final sample of 31,825 firm-year observations from 44 countries over the twelve-year period from 1992 to 2003 representing a total of 7,968 individual firms.

Next, we compile a comprehensive dataset of U.S. cross-listings using information from Citibank, JP Morgan, Bank of New York, Datastream and Bloomberg. Citibank and Datastream provide information on both, active and inactive ADRs. Including ADRs that have become inactive in the meantime mitigates concerns about survivorship bias, which would otherwise arise. We manually cross-check the various datasets, and construct binary variables indicating the existence and type of a U.S. cross-listing in a given year. This coding also accounts for changes in the ADR types and hence the sequence of U.S. cross-listings for a given firm.

We originally identify 1,310 unique foreign firms with an ADR that is active at some point between 1992 and 2003, which is a comprehensive panel of ADRs. After merging this dataset with the cost of capital estimates, we have 5,083 ADR firm-years from 1,001 unique firms with U.S. cross-listings in our sample. Due to differences in their legal and regulatory consequences, we differentiate between exchange listings on NYSE, NASDAQ, and AMEX, over-the-counter listings in the Pink Sheets or the OTC Bulletin Board, and private placements under Rule 144A. By cross-listing on a U.S. stock exchange, firms have to file Form 20-F with the SEC, requiring extensive disclosures and a reconciliation of foreign financial statements to U.S. GAAP. Moreover, by virtue of filing with the SEC, firms are subject to SEC enforcement and face legal liabilities from shareholder litigation. Cross-listings in the OTC markets do not require a 20-F filing, but a registration statement using Form F-6 and home-country disclosures to the SEC. It also subjects firms to the Foreign Corrupt Practices Act, under which most SEC enforcement actions are brought (Karpoff et al., 2004). Private placements under Rule 144A do not require any registration with the SEC or any disclosures, but are limited to a closely defined set of qualified institutional investors.

Table 1, Panel A, reports information on the sample composition, descriptive statistics on the cost of capital estimates, inflation rates and firm characteristics used in subsequent analyses by country. It shows considerable variation in the mean cost of equity capital across countries, ranging from a high of 22.4% for non-ADR firms in Colombia to a low of 7.0% for ADR firms in Japan. Panel B provides descriptive information on the aggregate and the individual cost of capital estimates by ADR type. The mean COC estimate is similar for ADR and non-ADR firms. However, this result is obtained because companies with private placements under Rule 144A exhibit on average a higher cost of capital than non-ADR firms, while firms with a U.S. exchange listing display a lower average cost of capital. The individual cost of capital estimates exhibit the same rank order. Although the magnitude of the estimates varies across the four individual valuation models, the estimates are all highly correlated with each other (pairwise correlation of 0.55 or higher) and with the aggregate COC estimates (pairwise correlation of 0.77 or higher).

3.3 Control Variables

As simple comparisons across countries and types of cross-listings can be misleading because they do not control for firm and country characteristics known to affect firms' cost of capital, we conduct regression analyses controlling for traditional risk and country factors.

First, we control for differences in inflation. 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. As the market's expectation for future inflation is only imperfectly observable, we 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 rather than forcing a coefficient of minus one when using inflation-adjusted cost of capital estimates. We expect the coefficient to be positive but smaller than one, as measurement error in the expected inflation proxy 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 monthly inflation rates as a proxy for the expected future inflation.

Another factor is time-series variation in risk-free interest rates. It is common in international studies to convert local returns into US$ and then to 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 time preferences are the same across countries. Thus, one needs to control only for time-series variation in the risk-free rate. We make a similar assumption and use year-fixed effects to control for such variation, recognizing that the T-bill rate is a yearly constant in our firm-year analysis. However, we also conduct sensitivity analyses using local risk-free rates.

Next, we include a number of traditional controls for risk at the firm level. 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 return volatility and financial leverage (e.g., Fama and French, 1992, 1993). We measure SIZE as the firm's US$ book value of total assets as of the fiscal-year end, return variability (RVAR) as the standard deviation of monthly stock returns over the last twelve months, and financial leverage (LEV) as the ratio of total liabilities to total assets.

We use an accounting-based measure of firm size in our analyses because including contemporaneous market values in the panel regressions would absorb the hypothesized effect if cross-listing indeed leads to a lower cost of capital and hence higher valuations.[15] We use return variability instead of market beta as a control for risk 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 it is not clear that our sample firms trade in integrated markets. In fact, one reason for using the implied cost of capital approach is that it does not require choosing a market portfolio and hence 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., 1996).[16] We include financial leverage as a third firm-level control because it is known to have a systematic effect on the equity cost of capital (Modigliani and Miller, 1958).

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15 To confirm this conjecture, we include contemporaneous market value of equity instead of book value of total assets and find that the cross-listing indicators become insignificant or even significantly positive. However, when we run firm-fixed effects regressions lagging market values by one or two periods, the results are again very similar to those reported in Table 2, which illustrates the issue that contemporaneous market values already reflect the effects of cross-listing.

16 Using a global beta instead of return variability or adding beta to the full model does not materially affect our results. Similarly, we find that adding earnings volatility as control does not materially change our results.

In addition to the three firm-level controls, we include industry- and country-fixed effects in our regression models. Fama and French (1997) document substantial variation in factor loadings across industries. We use the industry classification in Campbell (1996) to construct industry indicators. However, the results are also robust to using two-digit SIC codes to construct industry-fixed effects. The country-fixed effects are intended to control for differences in firms' cost of capital that stem from countries' economic and institutional environments (e.g., Erb et al., 1996; Hail and Leuz, 2006).

A major concern about the cross-sectional analysis is that cross-listed firms differ from non-cross-listed firms in ways that are unobservable or at least difficult to measure. To address this concern, we exploit the panel structure of our data and estimate regressions with firm-fixed effects (instead of country- and industry-fixed effects). These regressions should mitigate concerns about correlated omitted variables and unobserved heterogeneity across cross-listed and non-cross-listed firms.

4. Main Results

4.1 Comparison of ADR to non-ADR Firms

In our first set of analyses we compare firms that are cross-listed in the U.S. to local firms without cross-listing. To test our hypotheses delineated in Section 2, we regress the implied cost of capital measure on a panel of three binary variables indicating three ADR types (i.e., PP for private placements, OTC for shares traded over-the-counter, and EXCH for exchange listings) and an extensive set of risk and country control variables. The ADR panel also accounts for the fact that firms often change their cross-listing type once they have initiated an ADR program. Table 2 presents OLS coefficient estimates and t-statistics for these regressions. Our inferences are based on heteroscedasticity-corrected standard errors, which are clustered at the firm level—

except when including firm-fixed effects—to account for the fact that the same firm enters the sample multiple times (see, e.g., Petersen, 2006). For expositional purposes, we multiply all coefficients by 100.

Model 1 serves as our starting point in that we include only the ADR type indicator variables as well as year-, industry- and country-fixed effects. The coefficients on OTC and EXCH are significantly negative, and the latter exceeds the former in magnitude. This pattern continues to hold when we include the controls for inflation and risk in Model 2, but the magnitude of the OTC and EXCH coefficients substantially decreases to –0.32 and –0.35, respectively. Both variables are significant at the 5% level or better (two-tailed), but the coefficients are not statistically distinguishable from each other. The coefficient on PP is now positive and significant. This finding is somewhat surprising and discussed in more detail below. As expected, COC is positively related to the inflation rate, return variability and financial leverage, and negatively associated with firm size. All control variables are highly significant and, together with the year-, country- and industry-fixed effects, Model 2 explains about 36% of the international variation in firms' cost of capital, which is consistent with prior work (e.g., Botosan and Plumlee, 2005; Hail and Leuz, 2006).

Next, we exploit the panel structure of our data set and introduce firm-fixed effects as controls for time-invariant firm characteristics that are unobservable and difficult to measure (Model 3).[17] Introducing firm-fixed effects does not materially change the inferences, but increases the magnitude of the ADR coefficients. In particular, the EXCH coefficient increases to –0.85, suggesting that controlling for unobserved heterogeneity is important. All three ADR type

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[17] As a robustness check, we also introduce country-year-fixed effects in Model 2. Using interactions between country and year dummies accounts for time-specific country effects (e.g., market liberalizations). Our results are robust to the introduction of these effects and very similar to those reported for Model 2.

variables are statistically significant and different from each other, and the model has substantial explanatory power (72%). The last model addresses the concern that markets often price cross-listings well in advance of the actual ADR issuance or even its announcement (e.g., Foerster and Karolyi, 1999; Errunza and Miller, 2000; Sarkissian and Schill, 2006). To account for such anticipation effects, we delete the two firm-years immediately preceding the cross-listing and repeat our analysis (Model 4). Consistent with the notion of anticipation effects, the results become more pronounced, i.e., the coefficients increase in magnitude, suggesting that, if anything, our prior models underestimate the cross-listing effects.

We gauge the economic significance of the effects using the last two models in Table 2, as they are conceptually the most appealing ones. We find that exchange listings are associated with an average reduction in the cost of capital between 85 and 160 basis points. The effect of cross-listings in the OTC markets is smaller in magnitude and between 30 and 50 basis points. These reductions in the cost of equity capital are economically significant and translate into a substantial increase in firm value.

However, they are rather modest in comparison to prior return-based evidence. For instance, Errunza and Miller (2000) estimate a reduction in the cost of capital on the order of 1,100 basis points comparing long-run, realized returns relative to a matched sample of non-cross-listed firms. But as discussed before, long-run return comparisons are rather difficult and further complicated by the fact that it is hard to find appropriate local matches without cross-listings, as cross-listed firms are generally the largest firms in the industry or even economy. In addition, Sarkissian and Schill (2006) show that it is difficult to estimate equilibrium returns around the cross-listing event, requiring long time-series to obtain stable results. Using a 10-year window around the cross-listing, they estimate a pre- versus post-cross-listing return differential of about –900 basis points, which is still too large to be plausible.[18] Our findings are more in line with an estimate provided in Karolyi (1998) based on a small sample of representative companies from five regions. He gauges the effect based on changes in local and global beta factors around the cross-listing and estimates it to be on the order of –126 basis points, primarily due to reductions in local market betas.

In contrast to exchange- and OTC-traded cross-listings, the positive coefficient on PP indicates that private placements are associated with an *increase* in the cost of capital, even after including firm-fixed effects. This result is somewhat surprising, but consistent with prior studies where private placements often stand out with opposite findings as well (e.g., Miller, 1999; Foerster and Karolyi, 2000; Doidge, 2004; Bailey et al., 2006). One potential explanation for our result is that private placements entail giving a group of selected investors privileged access to information. If such private communications lead to greater information asymmetries in capital markets, an increase in the cost of capital would not be surprising. An alternative explanation is that private placements are negatively received by outside investors because they reveal that, despite a need for new capital, the firm shuns the legal consequences associated with a cross-listing in U.S. public equity markets.[19]

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[18] Extending the window to 20 years (and excluding 10 years surrounding the cross-listing) reduces the estimate to –200 basis points (Sarkissian and Schill, 2006). This estimate is closer to ours. But requiring a 20-year time series excludes many firms from the analysis. Moreover, it is not clear that a firm can still be compared to itself after such a long time period.

[19] Related to this explanation, Baek et al. (2006) show that Korean firms at times use private securities for the purpose of tunneling, and Leuz and Oberholzer-Gee (2006) provide evidence suggesting that the most connected (and least transparent) firms often have foreign private securities but shun publicly-traded foreign securities.

4.2 Sensitivity Analyses

This section presents several robustness checks for the cross-sectional results as well as our cost of capital estimation. First, we examine whether our findings hold for estimates from each of the four implied cost of capital models. Second, we assess whether the issuance of new equity capital in the U.S. potentially affects our results. Next, we consider several alternative model specifications, i.e., we include local risk-free interest rates, control for differences in analyst forecast properties (e.g., forecast bias), and check whether restricting the analysis to firms with existing and positive earnings forecasts biases our results. Finally, we gauge the sensitivity of our COC estimates to the models' assumptions about long-run growth. Systematic differences in long-run growth expectations across countries and/or ADR types may affect our estimates and bias the tests. Table 3 reports only the coefficients of the three ADR type indicators (plus the capital issuance indicators in Panel B), but the results stem from regressions including the full set of controls and firm-fixed effects (i.e., Model 3 in Table 2).

Panel A of Table 3 reports the results when we replace the aggregate COC measure with the individual estimates. Regardless of the valuation model, EXCH is always negative and significant at the 4% level or better. Moreover, the coefficient has a similar magnitude across models ranging from –0.66 to –1.04. The coefficient on OTC is negative and significant in two out of the four specifications. PP is positive and generally significant. These results demonstrate that our findings do not depend on the choice of a particular cost of capital model. However, they also indicate that aggregating across models helps in reducing noise in the estimates. Consistent with this assertion, we find that using the first principal component of the four individual proxies as dependent variable (not tabulated) further increases the power of our tests. PP, OTC and EXCH assume the same signs as in Table 2 and are all highly significant.

Our second set of tests addresses the concern that raising new equity capital in the U.S. may affect our analyses. The issuance of equity generally results in negative market reactions (e.g., Myers and Majluf, 1984). To address this issue, we include two separate indicator variables into our model to control for the capital raising activity of ADR firms. ISS_PP and ISS_EXCH take on the value of one in the year of a private placement or a public equity offering by an exchange-listed firm, respectively. OTC firms are not allowed to raise new equity capital within the U.S. and hence do not require an indicator variable. The first row in Panel B shows that the estimated coefficients for the ADR-type indicators are very similar to those reported in Table 2, and neither capital issuance indicator is significant. In the next row, we distinguish between *exchange*-listed ADRs that face stricter SEC rules in return for their ability to raise capital in the U.S. (i.e., Level III ADRs) and those that cannot raise capital but are exempt from certain SEC rules (i.e., Level II ADRs). We expect the former cross-listings to have larger cost of capital effects. Consistent with this expectation and the bonding hypothesis, we find that Level III firms exhibit the largest decrease in cost of capital (about 110 basis points), while the magnitude of the coefficient for Level II ADRs is smaller (–0.74) but still exceeds the decrease in the cost of capital for OTC cross-listings (–0.34). All coefficients are statistically significant at the 6% level or better.

Panel C of Table 3 reports ADR-type coefficient estimates from several alternative model specifications. First, we allow for the possibility that real risk-free interest rates differ across countries, reflecting among other things saving rates or interest rate regimes. We therefore replace the expected inflation proxy with the country-year median of the monthly risk-free interest rates drawn from yields of local treasury bills, central bank papers or inter-bank loans provided by Datastream. Second, we check that the results are similar when we use a risk premium as dependent variable, i.e., subtract the local risk-free interest rate (or, alternatively, inflation) from COC, rather than using raw COC together with the local risk-free rate as control on the right-hand side. We prefer the latter approach as using risk premiums forces a coefficient of minus one on the proxy for the risk-free interest rate (or expected inflation), despite the fact that the proxies are measured with noise. However, as shown in Panel C, neither controlling for local risk-free interest rates nor using risk premiums significantly alters our findings.

Next, we address concerns related to differences in analyst forecasts across countries and firms. Implied cost of equity capital estimates can be affected by forecast bias. For instance, if forecasts tend to be overly optimistic but market participants understand this bias and properly adjust prices, implied cost of capital models yield upwardly biased estimates. To the extent that such differences in forecasting behavior and the resulting "mechanical" effects are country- or industry-specific, the fixed effects in our regression models are likely to pick them up.[20] But forecasting differences could also be concentrated in particular sets of firms, e.g., they could be different for cross-listed firms, or change over time. Thus, we include a proxy for forecast bias calculated as the one-year-ahead forecast error. Alternatively, we include the two-year-ahead forecast error or forecast accuracy (both not tabulated). In all cases, our results are not materially affected. In addition, we estimate regressions giving more weight to COC estimates that stem from more accurate forecasts, which should reduce the influence of noisy inputs. We use the inverse of the absolute one-year-ahead analyst forecast error as weighing variable. Again, the tenor of the results does not change (see row 4 in Panel C).[21]

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[20] The same logic applies to differences in the accounting rules, which due to the short explicit forecast horizon in all implied cost of capital models might have a country effect. See Hail and Leuz (2006) for a discussion.

[21] In untabulated analyses we further (1) control for analyst following and changes therein around the cross-listing by including the number of one-year-ahead forecasts in the model, (2) restrict the sample to observations with at least three analysts following the firm, and (3) account for sluggishness in analyst forecasts, as suggested in Guay et al. (2005), by including the stock return over the three-month period immediately preceding the estimation of our cost of capital proxies as additional control. Again, these alternate specifications do not materially affect our findings.

In the last row of Panel C, we address concerns about sample selection stemming from the fact that implied cost of capital models require analyst forecasts for at least two periods ahead as well as positive realizations for these forecasts and the long-term growth estimate. It is possible that these data requirements screen out cross-listed and non-cross-listed firms differently. To gauge the effect of these data requirements, we replace all negative or missing earnings forecasts for up to three years with imputed forecasts using a firm's beginning book value of equity and the historic three-year median return on equity in a given country, industry and year. This imputation allows us to expand the sample by about a third to 42,558 firm-year observations. Despite this substantial increase in the sample, the results are very similar to those reported before, suggesting that sample selection due to the data requirements does not significantly bias our results.

Our final set of analyses, reported in Panel D, addresses concerns about the sensitivity of our COC estimates to the long-run growth assumptions in implied cost of capital models. In using analyst forecasts, these models make an explicit attempt to capture market expectations about (short-term) growth, which is why implied cost of capital estimates are appealing for our purposes. However, the underlying valuation models have to make assumptions about growth in residual income or abnormal earnings beyond the explicit forecast horizon, which could have a substantial influence on our estimates and hence the cross-sectional results.[22] In particular, we are concerned that cross-listed firms have different growth expectations, either because bonding extends their investment opportunity set or because they choose to cross-list when they happen to

[22] To gauge the general sensitivity of our COC estimates to the terminal value assumptions, we re-estimate the models using two alternative growth assumptions: (1) we use a constant 3% inflation rate for all countries; (2) we use a country's real GDP growth rate plus its long-run inflation rate. Whereas the latter assumes that countries' growth differences persist in perpetuity, the former assumes that growth rates converge to a global competitive equilibrium with zero real growth. The two assumptions represent opposite ends of a spectrum and our primary implementation (as detailed in the Appendix) falls somewhere in the middle. We find that our results are essentially the same using either assumption. See Hail and Leuz (2006) for similar sensitivity checks.

experience a shock to their growth opportunities. Differential growth expectations, if ignored, could bias our COC estimates and mechanically produce the cross-sectional results. As explained in more detail in the Appendix, we vary the long-run growth rate across ADR types in the following four ways: (1) we adjust the long-run growth rate in the terminal value computation by an ADR-type-specific component; (2) and (3) we use a correction of the long-run growth rate that not only varies by ADR type, but also by year, based on forecasted and real growth rates, respectively; (4) we derive a proxy for terminal-value growth on the firm-year level, and thereby refrain from using any common growth assumptions across firms or groups of firms.

As Panel D of Table 3 shows, the cross-sectional results, particularly for exchange-listed firms, are remarkably stable across all modifications. The three coefficients on the cross-listing indicators have the same sign and rank order as before. The coefficient on EXCH is negative and statistically significant for all modifications and exhibits a similar magnitude as in Table 2. Given these extensive checks, it is unlikely that differential long-run growth expectations are responsible for our findings, particularly for the cost of capital effects of U.S. exchange listings.[23]

4.3 Difference-in-Differences in Cost of Capital (Pre- vs. Post-Listing)

In this section, we examine changes in the cost of capital for firms that initiate an ADR program during the sample period. Analyzing changes pre- versus post-listing allows us to alleviate concerns that the differences in cost of capital are driven by unobserved heterogeneity across firms or selection bias. First, by looking at changes, each firm serves as its own control. Second, we standardize the pre- and post-listing cost of capital by subtracting the country-year median cost of capital of all non-ADR firms to control for time and macroeconomic trends as well

[23] Including the forecasted, one-year-ahead earnings growth rate or the analysts' long-term growth estimates as additional control variable (as well as introducing interaction effects between the growth rates and the cross-listing indicators) does not materially affect our results, and the inferences remain the same.

as differences in these trends across countries. This pre- versus post-listing comparison of standardized (i.e., median-adjusted) cost of capital estimates is in the spirit of a difference-in-differences analysis. The changes sample comprises only firms that initiated a U.S. cross-listing over the twelve-year period from 1992 to 2003. We remove all observations from the year of and before the cross-listing to mitigate concerns about announcement effects and to measure long-run changes in the cost of capital. The resulting 1,911 firm-years are the basis of our difference-in-differences analysis presented in Table 4.

Panel A reports the mean and median of the standardized cost of capital before and after cross-listing. The sample is limited to companies where pre- and post-listing data are available and also excludes firm-years where higher-level ADR types exist at the same time (e.g., a private placement is initiated but an exchange listing is already in place). The first two columns compare the mean and median standardized cost of capital in the years –2 and +1 around the cross-listing. The next two columns compare all available firm-years before and after ADR initiation, except those from years 0 and –1. By using all other firm-year observations, we verify that the changes in the cost of capital are in fact sustained. As cost of capital changes are expected to be permanent, it also increases the power of our tests. The changes in the first two columns are not significant, but they are similar in direction and magnitude to the cross-sectional analysis, suggesting that the small sample size and noise in the estimates are responsible for the insignificance. The changes in the last two columns are statistically significant. Moreover, the changes in the standardized cost of capital exhibit the hypothesized rank order. Exchange listings are associated with a decrease in the cost of capital by about 125 basis points. The table also highlights that firms with exchange listings already have a lower cost of capital than non-ADR firms *prior* to initiating the cross-listing (i.e., the pre-listing median is negative), illustrating the issue of selection bias in comparing ADR and non-ADR firms. Firms initiating an OTC cross-listing experience a significant reduction in the cost of capital between 60 and 90 basis points. As before, private placements are associated with a significant increase in the cost of capital.

Panel B reports coefficient estimates from regressing the standardized COC on ADR-type indicators, industry-fixed effects, and three firm-level controls for risk (i.e., size, return variability and leverage). Controls for inflation, country- and year-fixed effects are no longer needed as they have been subtracted by standardizing the cost of capital estimates with the country-year median COC of non-ADR firms. The regression analysis offers two advantages. First, it controls for changes in firm size and leverage, which is important as firms often change their investment and financing policies around cross-listings. Second, it allows us to simultaneously estimate the effects of different ADR types. The results from this difference-in-differences analysis confirm and nicely tie into our cross-sectional findings presented in Section 4.1. The negative effect of an exchange listing is significant at the 2% level or better in all specifications and on the order of 70 to 120 basis points. The reduction in the cost of capital is less pronounced but still significant for cross-listings in the OTC markets (–60 to –85 basis points). Private placements experience a significantly positive effect, but only if we include all firm-years around the cross-listing.

5. Additional Analyses

In this section, we extend our analyses of the relation between firms' cost of equity capital and U.S. cross-listings in two ways. First, we examine whether the cross-listing effect on the cost of capital differs by home-country legal institutions. Under the bonding hypothesis, we expect cross-listing effects to be stronger for firms originating from countries with weaker legal institutions. Second, we explore whether cross-listings in the U.S. exhibit significant cash flow effects stemming from changes in growth expectations. We exploit that we have analysts' growth estimates before and after the cross-listing and that we can use these estimates to decompose changes in firm value around the cross-listing into valuation effects from changes in growth expectations and those from changes in the cost of capital.

5.1 Cross-sectional Differences by Home-Country Institutions

In this section, we focus on whether and how the institutional characteristics of firms' home countries are associated with the cost of capital effects of cross-listing. The bonding hypothesis suggests that there are cross-sectional differences in the benefits of cross-listing, which depend on home-country characteristics and institutions. To test this argument, we partition the sample into several sub-samples by home-country institutions and run separate regressions for each sub-sample. We use the following partitioning variables: (1) countries are selected into two groups depending on their code or common law legal tradition (La Porta et al., 1997; Ball et al., 2000), (2) disclosure regulation is equal to high for countries with above-median index values of disclosure requirements in securities offerings (La Porta et al., 2006), (3) securities law enforcement is equal to high for countries with above-median levels of private and public enforcement of laws governing securities offerings (measured as the mean of the liability standard index and the public enforcement index from La Porta et al., 2006), and (4) investor protection is equal to high for countries with above-median anti-director rights index values (La Porta et al., 1998). All these proxies have in common that they attempt to capture the extent to which the institutional framework in the home country protects outside investors and, consequently, the extent to which firms benefit from U.S. cross-listings under the bonding hypothesis.

Table 5 provides coefficient estimates, t-statistics and significance tests across sub-samples. The regressions are based on Model 2 (see Table 2). As before, the control variables behave as predicted and are highly significant. Also in line with expectations, the coefficient on EXCH is negative and significant with a two-sided p-value of 4% or better for firms from countries with code-law origin, low disclosure requirements, weak securities law enforcement or weak investor protection. Moreover, the magnitude of the EXCH coefficient in these sub-samples is always larger than its counterpart for firms from countries with strong-institutions. Tests across the sub-samples indicate that the difference in the EXCH coefficients is statistically significant for all institutional variables, except legal tradition. These findings support the bonding hypothesis.

With respect to OTC cross-listings, the reverse picture emerges. In countries with weak institutions, the OTC coefficient is negative and generally smaller in magnitude than the EXCH coefficient, as expected. While the coefficient itself is not significant, it is significantly different from the EXCH coefficient for disclosure regulation and investor protection. Interestingly, OTC is significantly negative and larger in magnitude in countries with common-law origin, high disclosure regulation and high investor protection. Although the OTC coefficients are not significantly different, this pattern across sub-samples can be interpreted within the context of the bonding hypothesis. As bonding is of lesser value to firms from strong institutions, we expect firms from these countries to seek cross-listings for other reasons, e.g., to reduce trading barriers or to enlarge their shareholder base, which are likely to have more modest effects. OTC cross-listings serve this purpose and are less costly than exchange listings, which should make them attractive to firms with these other motives.[24] However, OTC cross-listings provide less assurance to outsiders in countries with weak institutions, which implies that firms from these countries benefit less than firms from strong institutions, consistent with a smaller coefficient.

The PP coefficient is significantly positive for countries with strong institutions across all four partitioning variables and significantly different across sub-samples in countries with high

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[24] In countries with strong institutions, the OTC coefficient also appears to be larger than the EXCH coefficient, which is surprising, given that exchange listings in these countries should offer some diversification and shareholder base effects. However, the differences between these coefficients are never significant.

disclosure regimes and high investor protection. While we do not have an explanation for the latter result, the rank order of the cross-listing coefficients across ADR types is consistent with our expectations and the earlier findings.

Taken together, the results suggest that the cost of capital effects of U.S. cross-listings vary by home-country characteristics and venue of cross-listing. Exchange-listed firms from countries with weaker institutions enjoy a larger reduction in cost of equity capital than firms from countries with strong outsider protection, consistent with the bonding argument. OTC cross-listings from countries with strong institutions show modest cost of capital effects, consistent with cross-listing motives other than bonding.

5.2 Decomposition of Valuation Effects around U.S. Cross-Listings

In this section, we exploit that analysts provide explicit estimates of firm growth. We use these forecasts, the (implied) cost of capital estimates and the four valuation models to derive "as-if" price changes that decompose the total return around the cross-listing into two components: a price effect due to revisions in the cost of capital and a price effect due to revisions in cash flow (or growth) expectations.[25] Below, we illustrate the basic idea with a perpetuity valuation model. However, our empirical estimations reported in Table 6 are based on the more complex valuation models described in the Appendix.

We calculate the realized buy-and-hold return, ΔP, around the ADR issuance starting at a point when market prices do not yet reflect the cross-listing effects, say $t = -2$, and up to $t = 0$

[25] We hasten to acknowledge that this procedure is not without problems. First, firms' growth opportunities and investment policies are not independent of the cost of capital. We therefore use a logarithmic decomposition but still do not expect the decomposition to be perfect. Second, we use the data first to estimate the implied cost of capital and then use it again to decompose the price effects. It is possible that this procedure introduces correlations into the estimates that we do not take into account in our analyses. Third, all our estimates rely on analysts' ability to predict the cash flow consequences of cross-listings. Nonetheless, we see our approach as a first step to shed light on the complex mechanisms through which cross-listings affect firm value.

when the firm starts to cross-list in the U.S. We note that, at each point in time, price is a function of the market's cash flow expectations and the concurrent cost of capital:

$$\Delta\text{P} = \left(P_0 / P_{-2}\right) = \left(\frac{\overline{CF_0}}{r_0} \middle/ \frac{\overline{CF_{-2}}}{r_{-2}} \right), \tag{1}$$

where P_t is a firm's current stock price in the forecast period month of year t, $\overline{CF_t}$ represents the stream of expected future cash flows in perpetuity and r_t a firm's (implied) cost of capital. Rearranging terms and expanding this equation yields:

$$\Delta\text{P} = \left(P_0 / P_{-2}\right) = \left(\frac{\overline{CF_0}}{r_{-2}} \middle/ \frac{\overline{CF_{-2}}}{r_{-2}} \right) \times \left(\frac{\overline{CF_{-2}}}{r_0} \middle/ \frac{\overline{CF_{-2}}}{r_{-2}} \right). \tag{2}$$

Next, we log the price effect to linearly decompose the two multiplicative effects:

$$\ln\left(P_0 / P_{-2}\right) = \ln\left(\frac{\overline{CF_0}}{r_{-2}} \middle/ \frac{\overline{CF_{-2}}}{r_{-2}} \right) + \ln\left(\frac{\overline{CF_{-2}}}{r_0} \middle/ \frac{\overline{CF_{-2}}}{r_{-2}} \right). \tag{3}$$

Based on equation (3), we compute the as-if price change attributable to revisions in cash flow expectations around the cross-listing, $\Delta\text{P}_{\text{CF}}$, as follows:

$$\Delta\text{P}_{\text{CF}} = \ln\left(\frac{\overline{CF_0'}}{r_{-2}'} \middle/ \frac{\overline{CF_{-2}}}{r_{-2}} \right) = \ln\left(\frac{\overline{CF_0'}}{r_{-2}'} \right) - \ln\left(P_{-2}\right). \tag{4}$$

A prime indicates estimates, rather than realized values. In essence, to decompose the total price effect ΔP into a cash flow and cost of capital component, we use the cost of capital estimate from t = –2 along with cash flow expectations from t = 0. The idea of these replacements is to hold the cost of capital constant and to estimate the price effect of changes in growth expectations. The result is a two-year-ahead return estimate reflecting (among other things) the cash flow (or growth) effects of cross-listing in the U.S.

Similarly, we derive the as-if price change that is attributable to revisions in the cost of capital around the cross-listing, $\Delta\text{P}_{\text{COC}}$. We use our estimate of the implied cost of capital at time

$t = 0$, r_0', along with the estimate of expected cash flows at time $t = -2$, $\overline{CF'_{-2}}$, essentially holding cash flow expectations constant:

$$\Delta P_{COC} = \ln\left(\frac{\overline{CF'_{-2}}}{r_0'} \Big/ \frac{\overline{CF_{-2}}}{r_{-2}} \right) = \ln\left(\frac{\overline{CF'_{-2}}}{r_0'} \right) - \ln(P_{-2}). \qquad (5)$$

Again, the underlying assumption is that r_0' captures the effect of the cross-listing on the cost of capital and that $\overline{CF'_{-2}}$ does not yet anticipate the cross-listing effects. The result is a two-year-ahead return reflecting (among other things) the cost of capital effects of cross-listing in the U.S.

In aggregating as-if price changes across firms and event time, we attempt to "wash out" other effects and to extract the cash flow or cost of capital effects of cross-listing. In addition, we subtract from each as-if price change of a cross-listed firm the median as-if price change of all non-ADR firms in the same country and from the same year. This adjustment should control for market-wide trends in the stock price changes and account for country- and year-specific effects, such as inflation. With respect to timing, we assume that market prices reflect the cross-listing decision in year –1 or 0 and, hence, we compute two-year price changes in years –3 and –2, respectively. This design choice is based on prior studies documenting a stock price run-up starting 12 months prior to the listing (e.g., Foerster and Karolyi, 1999; Errunza and Miller, 2000), but it might still fall short of capturing the full effect (e.g., Sarkissian and Schill, 2006).

Table 6 provides descriptive statistics for the realized two-year returns around the cross-listing as well as the estimated cash flow and cost of capital effects. We cannot compute two-year (ahead) returns for the last two years of the sample interval and hence lose some observations. We further eliminate the top and bottom one percent of the price change distributions to mitigate the influence of outliers. These choices leave us with price changes for a maximum of 3,875 ADR firm-year observations. Panel A reports the median, mean and standard deviation of the price change variables in years –3 and –2, and assesses whether ΔP_{CF} and ΔP_{COC} are statistically different from zero and from each other. The median-adjusted changes in realized prices around the cross-listing range from 6% (mean for OTC listed firms) to over 19% (median for exchange listings) and are comparable to prior studies (e.g., Errunza and Miller, 2000; Doidge et al., 2004). The high standard deviations indicate substantial variation in the two as-if price change components, likely indicating a fair amount of measurement error.

The mean cash flow effects are relatively similar in magnitude across all three ADR types, but they are significantly different from zero only for OTC and exchange-listed firms. Exchange-listed ADRs exhibit by far the largest cost of capital effects, which is consistent with our earlier findings. For these firms, both cash flow and cost of capital effects appear to contribute equally to the overall price increase around the cross-listing. Based on the decomposition, the positive total price change around OTC cross-listings stems primarily from improved growth expectations, rather than from a decrease in the cost of capital. The same appears to be true for private placements. However, the price change components for private placements are based on a relatively small sample and not statistically significant.

Since these univariate comparisons do not control for differences in firm characteristics associated with firms' returns and changes in those characteristics around cross-listings, we next employ regression analysis using the same model specifications as in Panel B of Table 4. We re-code the cross-listing indicator variables to align them in time with the cash flow and cost of capital effects, which are measured over two years beginning in year –3 and year –2 before the actual cross-listing. In essence, our regression estimates whether there are any systematic cash flow and cost of capital effects *around cross-listings*, relative to all the other ADR firm-years and relative to all firm-years of non-ADR firms. To account for contemporaneous correlations in the error terms across the two price-effect models, we estimate them simultaneously using seemingly unrelated regressions and assess the significance of the coefficients using z-statistics.

We find that firms initiating an exchange-listed ADR exhibit significantly positive cash flow *and* cost of capital effects in the years surrounding the cross-listing. The magnitude of the coefficients suggests that both effects contribute equally to the positive valuation effects around cross-listings. We also find significantly positive cash flow effects for OTC-traded firms and private placements, which are similar in magnitude to the cash flow effects of exchange listings. The cost of capital effects of OTC and PP firms, however, are insignificant.

Our results clearly indicate that not all valuation effects around cross-listings are attributable to a reduction in the cost of capital. There are substantial valuation effects from changes in growth expectations. These cash flow effects could reflect that firms choose to cross-list when they experience (unrelated) shocks to their growth opportunities or that firms are able to expand their growth opportunities due to the cross-listing. However, the fact that the cash flow effects are fairly similar across ADR types and, in particular, do not exhibit the rank order that we observe for the cost of capital effects, suggests that the former explanation is more likely than the latter.[26] More generally, these results caution us to compare valuation changes (or returns) around cross-listings as these measures likely capture growth effects, together with any cost of capital consequences, and hence may overstate the long-run valuation benefits of cross-listings per se.

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[26] In untabulated tests, we find that ADRs associated with capital-raising activities exhibit the largest cash flow effects, which is also more consistent with the explanation that firms seek cross-listings when their growth opportunities happen to expand.

6. Conclusion

In this paper, we examine the cost of capital effects of U.S. cross-listings for a large panel of 1,001 unique ADR firms from 44 countries. Prior research documents significant valuation effects of cross-listings in the U.S. However, it provides little evidence on the mechanism(s) by which cross-listing affects firm value. One important question is whether and to what extent these benefits stem from a reduction in firms' cost of capital, as the bonding hypothesis and disclosure theory would suggest. Alternatively, it is possible that the valuation effects reflect that firms choose to cross-list when they experience an expansion in their growth opportunities that is unrelated to the cross-listing. Thus, understanding the sources of the valuation benefits is important. This issue is, for instance, at the heart of the debate about the Sarbanes-Oxley Act and its potential consequences on foreign firms' cross-listing or delisting decisions.

We use cost of capital estimates implied by current market prices and analyst forecasts, rather than estimates based on realized returns, as this approach makes an explicit attempt to account for changes in growth expectations around cross-listings. We find strong evidence that cross-listings on U.S. exchanges are associated with a significant decrease in firms' cost of equity capital, after controlling for traditional proxies for risk and firm-fixed effects, or conducting difference-in-differences tests. The magnitude and the ranking of the effects seem plausible with exchange listings experiencing a reduction in cost of capital of about 70 to 120 basis points, followed by OTC listings with about 30 to 70 basis points. Firms that access U.S. markets via private placements exhibit an increase in cost of capital. While this finding is unexpected, it is consistent with several prior studies documenting opposite effects for private placements in their analyses. When we investigate cross-sectional differences in the cost of capital effects, we find that firms from countries with weak disclosure regulation and investor protection benefit the most from cross-listing on U.S. exchanges. Overall, these findings lend support to the bonding hypothesis.

We also document significant valuation effects stemming from changes in (analysts') growth expectations around cross-listings, indicating that not the entire valuation effect is attributable to a reduction in the cost of capital. The cash flow effects could partly reflect that firms seek U.S. cross-listings when their growth opportunities happen to expand. Consistent with this explanation, we document that the effects are fairly similar across ADR types and, if anything, are largest for cross-listings that involve raising new equity capital.

Finally, several caveats are in order. First, cross-listings represent major corporate events making it difficult to compute equilibrium cost of capital estimates. While our methodology should be well equipped to estimate the long-run consequences of cross-listings for firms' cost of equity capital, it is possible that the events leading up to the cross-listing affect the calculation of our proxies. Second, while we interpret our evidence as indicating that cross-listings reduce the cost of capital, it is possible that causality runs the other way (i.e., firms choose to cross-list after they experience a reduction in their cost of capital). In this case, however, we would not expect the cost of capital effects to differ across ADR types in the documented way. Lastly, cost of capital and cash flow (or growth) effects are likely to be intertwined in a non-trivial way. Thus, the results from our decomposition of the two effects should be interpreted cautiously.

APPENDIX: Implied Cost of Equity Capital Models

A.1 Overview and Model-specific Assumptions

Claus and Thomas (2001):

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

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 industry-specific median return. We compute the historic three-year average return on equity in a given country and year based on the industry classification in Campbell (1996). Negative yearly target returns are replaced by country-industry 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. In our main tests, 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). 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.[27] However, this implies that the one-year ahead forecast ($\hat{x}_{t+1}$) is for a fiscal year that ends just two months later.

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[27] We repeat our 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$ 1 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.[28] 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.[29]

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 our proxy for long-run growth expectations (g or g_{lt}) in the terminal value computations. Negative values are 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.

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

29 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 cost of capital estimates.

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.

In our sensitivity analyses in Table 3, Panel D, we vary the assumptions about long-run growth beyond the explicit forecast horizon, where applicable (i.e., for r_{CT} and r_{OJ}). We use the following four ways to create ADR-specific adjustments to the long-run growth rate (g or g_{lt}), which in our base version assumes growth equal to the expected inflation rate and varies by country and year: (1) We extract an ADR-specific growth component from regressing non-negative analyst long-run growth estimates, *ltg*, on ADR-type indicators, industry-, year- and country-fixed effects. The coefficients on the ADR-type variables represent the growth differentials relative to non-cross-listed firms, and we add them to the growth rate of the base version. (2) In a further refinement, we use adjustments that not only vary by ADR type, but also by year. That is, we compute the differences in analysts' long-run growth estimates across ADR types and non-ADR firms on a yearly basis (i.e., median $ltg_{ADR\ type,t}$ – median $ltg_{non\text{-}ADRs,t}$ where ADR type is equal to PP, OTC, and EXCH, respectively). (3) As *ltg* reflects analysts' expectations about the future growth, which may not necessarily translate into actual growth differences, we repeat the calculations in (2) using ex-post realized growth rates, based on historic three-year average growth rates in earnings per share. (4) We directly use analysts' long-run growth estimates (*ltg*), which are *firm*- and *year*-specific, as proxy for terminal-value growth.

As the computation of r_{GLS} is not affected by the above adjustments to g, we instead vary the industry-specific target accounting return on equity, which serves a similar purpose in the GLS model as the long-run growth assumptions in the other models. We use the following two approaches to create ADR-specific adjustments that are analogous to those for the long-run growth rate: (i) We extract ADR-specific corrections from regressing historic three-year average returns on equity on ADR-type indicators, industry-, year- and country-fixed effects. The coefficients on the ADR-type variables are return on equity differentials relative to non-cross-listed firms, which we add to the existing target return on equity, *roe*, that varies by country, industry and year. (ii) We use adjustments that not only vary by ADR type, but also by year. We compute the differences in historic three-year average returns on equity across ADR types and non-ADR firms on a yearly basis (i.e., median $roe_{ADR\ type,t}$ – median $roe_{non\text{-}ADRs,t}$). For the results presented in Table 3, Panel D, we combine approach (i) of correcting *roe* with approaches (1) and (4) of correcting *g*, while (ii) is combined with (2) and (3).

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TABLE 1

Descriptive Statistics for the Implied Cost of Capital Estimates and Control Variables

Panel A: Sample Information, Cost of Capital Estimates and Firm Attributes by Country

	Firm-Years		*COC*			*Firm Characteristics*		
Country	*Total*	*ADRs*	*Non-ADRs*	*ADRs*	*INFL*	*SIZE*	*RVAR*	*LEV*
Argentina	142	69	13.45%	12.70%	0.64%	2,999	0.125	0.492
Australia	1,775	375	11.86%	10.67%	2.69%	4,742	0.085	0.500
Austria	235	89	13.30%	11.47%	1.74%	5,825	0.092	0.622
Belgium	509	33	11.65%	11.35%	1.71%	16,700	0.081	0.583
Brazil	284	152	21.34%	20.67%	6.76%	6,321	0.143	0.502
Chile	143	94	14.56%	13.02%	4.92%	2,909	0.094	0.482
China	343	80	12.19%	14.35%	0.35%	2,161	0.129	0.427
Colombia	13	3	22.39%	20.50%	11.77%	2,945	0.102	0.641
Czech Republic	56	14	15.60%	10.87%	6.06%	3,007	0.114	0.533
Denmark	585	26	12.32%	9.05%	2.13%	2,955	0.084	0.578
Egypt	22	13	21.47%	21.11%	4.60%	1,822	0.109	0.728
Finland	534	66	14.35%	12.67%	1.40%	1,730	0.101	0.528
France	2,329	297	11.70%	10.72%	1.51%	13,010	0.105	0.618
Germany	1,686	209	11.96%	10.52%	1.60%	21,760	0.114	0.627
Greece	281	35	12.33%	11.32%	3.74%	3,691	0.132	0.525
Hong Kong	1,280	337	16.57%	14.22%	1.26%	3,154	0.129	0.422
Hungary	94	60	15.92%	17.38%	10.27%	1,404	0.119	0.372
India	765	251	15.81%	16.77%	5.94%	2,201	0.131	0.525
Indonesia	348	23	18.20%	16.89%	8.23%	1,126	0.144	0.549
Ireland	162	34	14.47%	12.84%	3.65%	7,235	0.094	0.633
Israel	73	15	11.77%	11.78%	2.38%	11,630	0.102	0.572
Italy	793	137	11.42%	9.98%	2.71%	20,440	0.097	0.697
Japan	4,421	281	8.42%	7.00%	-0.25%	7,953	0.108	0.546
Korea (South)	809	104	17.07%	17.06%	3.39%	5,776	0.155	0.594
Malaysia	1,296	32	11.92%	12.25%	2.49%	1,533	0.115	0.474
Mexico	184	147	16.21%	15.59%	8.68%	4,616	0.108	0.479
Netherlands	1,112	238	13.88%	11.78%	2.39%	13,190	0.091	0.621
New Zealand	355	29	11.71%	10.87%	2.40%	612	0.080	0.452
Norway	437	101	14.48%	12.89%	1.92%	2,379	0.108	0.582
Pakistan	62	2	21.30%	15.86%	6.64%	507	0.138	0.681
Peru	37	14	16.21%	17.19%	5.62%	2,034	0.113	0.552
Philippines	287	78	15.78%	13.91%	6.37%	1,534	0.138	0.523
Poland	154	47	15.83%	15.43%	7.94%	2,091	0.134	0.533
Portugal	198	41	12.18%	11.29%	3.07%	7,215	0.086	0.647
Singapore	888	77	12.59%	11.23%	1.10%	1,291	0.120	0.438
South Africa	933	214	19.13%	16.91%	6.15%	2,303	0.121	0.503
Spain	712	102	11.81%	11.27%	3.19%	11,560	0.083	0.610
Sri Lanka	59	6	18.56%	12.67%	5.99%	227	0.124	0.607
Sweden	875	112	13.17%	12.08%	1.17%	5,985	0.103	0.574
Switzerland	998	128	11.99%	9.58%	0.84%	16,110	0.087	0.595
Taiwan	723	166	12.34%	11.65%	0.93%	2,909	0.135	0.438
Thailand	527	39	15.30%	13.85%	2.88%	913	0.129	0.535
Turkey	5	3	11.92%	16.58%	8.42%	5,837	0.108	0.621
United Kingdom	4,301	710	12.20%	10.79%	2.55%	8,977	0.096	0.563
Total (Average)	31,825	5,083	12.49%	12.47%	2.24%	7,919	0.108	0.550

(continued)

TABLE 1 (continued)

Panel B: Sample Information, Summary and Individual Cost of Capital Estimates by Listing Type

		COC		r_{CT}		r_{GLS}		r_{OJ}		r_{PEG}	
Type	*Firm-Years*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*
Non-ADR Firms	26,742	12.49%	5.50%	12.05%	6.80%	9.48%	5.11%	14.18%	6.12%	14.24%	6.95%
ADR Firms	5,083	12.47%	5.16%	11.64%	5.90%	9.88%	4.87%	14.17%	5.90%	14.20%	6.86%
Private Placements 144A	1,297	14.01%	5.51%	13.32%	6.07%	11.43%	5.32%	15.72%	6.34%	15.56%	7.37%
OTC Listing	2,051	12.48%	5.18%	11.63%	5.64%	9.80%	4.65%	14.24%	6.07%	14.25%	7.14%
Exchange Listing	2,033	11.54%	4.74%	10.62%	5.77%	9.00%	4.56%	13.18%	5.40%	13.34%	6.25%
All Firms	31,825	12.48%	5.45%	11.99%	6.66%	9.54%	5.07%	14.18%	6.09%	14.23%	6.94%

The sample comprises 31,825 firm-year observations from 44 countries between 1992 and 2003, for which sufficient Worldscope financial data, I/B/E/S forecast and pricing data exist. We include only observations from countries with at least one ADR, and years with inflation rates below 25%. ADR observations comprise placements under Rule 144A, traded shares in the over-the-counter (OTC) markets, and NYSE, Nasdaq or AMEX exchange listings of non-U.S. firms. The table reports the number of firm-year observations and average values for the variables by country (Panel A), and number of firm-year observations and means and standard deviations by listing type (Panel B). COC is the average cost of capital estimate implied by the mean analyst consensus forecasts and stock prices using (1) the Claus and Thomas (2001) model, r_{CT}, (2) the Gebhardt, Lee, and Swaminathan (2001) model, r_{GLS}, (3) the Ohlson and Juettner-Nauroth (2005) model, r_{OJ}, and (4) the Easton (2004) model, r_{PEG}. These models are described in more detail in the Appendix. INFL is the yearly median of country-specific, one-year-ahead realized monthly inflation rates. SIZE stands for US$ total assets (in millions). RVAR is the return variability computed as annual standard deviation of monthly stock returns. Financial leverage (LEV) is measured as the ratio of total liabilities to total assets. Accounting data are measured as of the fiscal-year end, RVAR and the cost of capital as of month +10 after the fiscal-year end. With the exception of SIZE (where we use the natural logarithm), we delete 1% of extreme realizations for all firm-level attributes.

TABLE 2

Cross-Sectional Cost of Capital Effects of U.S. Cross Listings

$$COC_{it} = \alpha_0 + \alpha_1 PP_{it} + \alpha_2 OTC_{it} + \alpha_3 EXCH_{it} + \alpha_4 INFL_{it} + \alpha_5 SIZE_{it} + \alpha_6 RVAR_{it} + \alpha_7 LEV_{it} + \sum\alpha_j \text{Year Controls}_t + \sum\alpha_k \text{Industry Controls}_i + \sum\alpha_l \text{Country Controls}_i + \varepsilon_{it}$$

Variable	*Pred. Sign*	*Model 1* *(all Years)*	*Model 2* *(all Years)*	*Model 3* *(all Years)*	*Model 4* *(except Years -1/0)*
N		31,825	31,825	31,825	30,806
PP	+/–	-0.159	0.412 #	0.640 *	0.995 *
		(-0.71)	(1.90)	(2.09)	(2.33)
OTC	–	-1.113 **	-0.319 *	-0.339 #	-0.494 *
		(-6.81)	(-2.00)	(-1.87)	(-2.26)
EXCH	–	-1.478 **	-0.348 *	-0.854 **	-1.607 **
		(-10.24)	(-2.34)	(-3.85)	(-5.94)
p-values:	$\alpha_1=\alpha_2$	(0.001) **	(0.008) **	(0.006) **	(0.002) **
	$\alpha_2=\alpha_3$	(0.084) #	(0.887)	(0.056) #	(0.001) **
Intercept		9.434 **	12.350 **	9.074 **	9.000 **
		(43.68)	(31.65)	(8.37)	(8.06)
INFL	+	–	19.847 **	19.170 **	19.341 **
			(9.94)	(10.93)	(10.47)
SIZE	–	–	-0.558 **	0.014	0.026
			(-21.20)	(0.18)	(0.32)
RVAR	+	–	16.732 **	12.089 **	12.099 **
			(22.46)	(16.12)	(15.86)
LEV	+	–	4.816 **	2.518 **	2.449 **
			(23.42)	(7.67)	(7.25)
Fixed Effects:					
Year		included	included	included	included
Industry & Country		included	included	–	–
Firm		–	–	included	included
R^2		29.4%	35.8%	72.2%	72.7%
F-Stat		87.8	119.1	126.8	122.8

The sample comprises 31,825 firm-year observations from 44 countries over the twelve-year period from 1992 to 2003 with data available. The dependent variable, COC, is the mean of four estimates for the implied cost of equity capital (see Appendix). Three binary variables indicate ADR observations: (1) PP is equal to one if the firm has a private placement under Rule 144A, (2) OTC is equal to one if firm shares trade in the over-the-counter markets, and (3) EXCH is equal to one if firm shares are listed on the NYSE, Nasdaq or AMEX. For Model 4, we assume that the market learns about the ADR in the two years immediately preceding the listing, and therefore exclude the years -1 and 0 from the analysis. INFL is the yearly median of country-specific, one-year-ahead realized monthly inflation rates. SIZE stands for US$ total assets (in thousands). RVAR is the return variability computed as annual standard deviation of monthly stock returns. Financial leverage (LEV) is measured as the ratio of total liabilities to total assets. Accounting data are measured as of the fiscal-year end, RVAR and the cost of capital as of month +10 after the fiscal-year end. We use the natural log of SIZE in the analysis and delete 1% of extreme realizations for all other firm-level attributes. Industry-fixed effects based on the classification in Campbell (1996), year-, country- and firm-fixed effects are included in the regressions where indicated, but not reported. The table reports OLS coefficient estimates and t-statistics (in parentheses) based on robust standard errors that are clustered by firm in Models 1 and 2, while using firm-fixed effects in Models 3 and 4. It also reports p-values from Wald tests comparing the magnitude of the ADR variables. For expositional purposes we multiply all coefficients by 100. **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 3

Sensitivity Analyses for the Cross-Sectional Cost of Capital Effects

Panel A: Individual Cost of Capital Estimates as Dependent Variables

Variable	*N*	*PP*	*OTC*	*EXCH*
r_{CT}	31,825	0.703 #	-0.287	-0.940 **
		(1.82)	(-1.25)	(-3.08)
r_{GLS}	31,825	0.703 **	-0.040	-1.036 **
		(2.91)	(-0.24)	(-5.01)
r_{OJ}	31,825	0.692 #	-0.371 #	-0.777 **
		(1.86)	(-1.68)	(-3.12)
r_{PEG}	31,825	0.461	-0.660 *	-0.662 *
		(1.05)	(-2.40)	(-2.06)

Panel B: Additional Controls for Capital Issuance

Variable	*N*	*PP*	*OTC*	*EXCH*	*ISS_PP*	*ISS_EXCH*
Capital issuance indicator	31,825	0.582 #	-0.332 #	-0.817 **	0.375	-0.342
		(1.87)	(-1.83)	(-3.67)	(1.14)	(-1.13)
				LEVEL II	*LEVEL III*	
Exchange level indicator	31,825	0.639 *	-0.337 #	-0.739 **	-1.103 **	
		(2.09)	(-1.86)	(-3.06)	(-3.40)	

Panel C: Alternative Model Specifications

Variable	*N*	*PP*	*OTC*	*EXCH*
Use local risk-free rates instead of inflation	31,825	0.660 *	-0.436 *	-0.799 **
		(2.17)	(-2.40)	(-3.58)
Use risk premia instead of COC estimates	31,588	1.477 **	-0.464 *	-0.712 **
		(4.22)	(-2.37)	(-3.00)
Control for analyst forecast bias	29,886	0.592 #	-0.344 #	-0.936 **
		(1.91)	(-1.88)	(-4.27)
Accuracy-weighted COC estimates	29,769	0.458	-0.354 *	-0.922 **
		(1.46)	(-2.03)	(-4.29)
Replace missing/negative earnings forecasts	42,558	0.804 **	-0.282	-0.911 **
		(2.61)	(-1.44)	(-4.04)

Panel D: Influence of (Long-Run) Growth Differences across ADR Type

Variable	*N*	*PP*	*OTC*	*EXCH*
ADR-specific correction of terminal-value growth	31,901	0.458	-0.644 **	-1.075 **
		(1.50)	(-3.52)	(-4.79)
ADR- & year-specific correction of terminal-value growth (forecasted)	31,786	1.252 **	-0.233	-0.686 **
		(4.04)	(-1.26)	(-3.06)
ADR- & year-specific correction of terminal-value growth (realized)	31,482	1.581 **	-0.128	-0.500 *
		(4.81)	(-0.66)	(-2.16)
Firm- & year-specific growth rates in terminal values	19,415	0.088	-0.574 *	-0.831 **
		(0.20)	(-2.07)	(-3.26)

(continued)

The sample comprises a maximum of 31,825 firm-year observations from 44 countries over the twelve-year period from 1992 to 2003 with data available. Except for Panel A where we use the individual cost of capital estimates (r_{CT}, r_{GLS}, r_{OJ}, and r_{PEG}), the dependent variable, COC, is the mean of four estimates for the implied cost of equity capital (see Appendix). The panels report only the coefficients (t-statistics) of the three ADR indicator variables (and the capital issuance variables in Panel B), but regressions include the full set of controls (see Model 3 in Table 2). Regressions in Panel B either include ISS_PP and ISS_EXCH, which are set equal to one in the years of a private placement and the capital issuance on the NYSE, Nasdaq or AMEX, respectively, or we replace EXCH by two distinct indicator variables for firms whose shares are listed with (LEVEL III) or without (LEVEL II) U.S. public offering during the sample period. Regressions in Panel C are based on the following alternative model specifications: (1) we replace INFL by r_F, the yearly median of country-specific, monthly risk-free interest rates, (2) we use risk premia (COC minus one-year-ahead realized r_F) instead of the raw cost of capital estimates as the dependent variable, (3) we include the one-year-ahead analyst forecast error (mean forecast minus actual) scaled by lagged total assets as additional control, (4) we estimate accuracy-weighted regressions with the inverse of the absolute one-year-ahead analyst forecast error scaled by lagged total assets as the weights, and (5) we replace all negative or missing forecasted earnings per share for up to three years using beginning book values, historic three-year median returns on equity in a given country, industry and year, and assuming clean surplus. In Panel D, we control for differences in projected (or realized) growth rates across ADR types. We compute the implied cost of capital estimates (1) by adding an ADR-type-specific component to the growth rate in the terminal value (equal to the coefficient estimates from regressing analyst long-run growth estimates on ADR-type-, industry-, year- and country-fixed effects), (2) by adding an ADR-type- and year-specific component to the terminal-value growth rate, using the yearly median differences in analyst long-run growth estimates across ADR types and non-ADR firms, (3) same as (2) but using realized, three-year average growth rates in earnings per share, and (4) by using (firm- and year-specific) analysts' long-run growth estimates as perpetual growth rates in the terminal value. For the computation of r_{GLS}, we also vary the industry-specific target accounting return on equity across ADR types using similar techniques as described above (see Appendix). The table reports OLS coefficient estimates and (in parentheses) t-statistics based on robust standard errors. For expositional purposes we multiply all coefficients by 100. **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 4

Difference-in-Differences Analysis for the Cost of Capital Effects of U.S. Cross Listings

$$COC_{it} = \alpha_0 + \alpha_1 PP_{it} + \alpha_2 OTC_{it} + \alpha_3 EXCH_{it} + \alpha_4 SIZE_{it} + \alpha_5 RVAR_{it} + \alpha_6 LEV_{it} + \sum \alpha_k \text{Industry Controls}_i + \varepsilon_{it}$$

Panel A: Univariate Analysis of the Implied Cost of Capital (pre vs. post ADR Listing)

		Years -2/+1 only			*all Years (except Years -1/0)*		
Variable	*Time*	*N*	*Median*	*Mean*	*N*	*Median*	*Mean*
PP	pre	66	0.63%	0.50%	237	0.18%	0.02%
	post		2.00%	1.63%		0.43%	1.67%
	post - pre		1.37%	1.13%		0.25% #	1.65% **
OTC	pre	178	-0.48%	-0.34%	773	-0.63%	-0.63%
	post		-0.90%	-0.51%		-1.50%	-1.21%
	post - pre		-0.43%	-0.17%		-0.87% **	-0.58% *
EXCH	pre	180	-0.67%	-0.67%	828	-0.64%	-0.28%
	post		-1.33%	-1.10%		-1.88%	-1.59%
	post - pre		-0.67%	-0.43%		-1.23% **	-1.31% **

Panel B: Regression Analysis of the Implied Cost of Capital (pre vs. post ADR Listing)

	Pred.	*Years -2/+1 only*		*all Years (except Years -1/0)*	
Variable	*Sign*	*Model 1*	*Model 2*	*Model 1*	*Model 2*
N		454	440	1,911	1,845
PP	+/–	0.327	0.382	0.951 *	0.826 *
		(0.65)	(0.78)	(2.28)	(2.14)
OTC	–	-0.577 #	-0.580 #	-0.852 **	-0.692 **
		(-1.78)	(-1.83)	(-3.08)	(-2.66)
EXCH	–	-1.212 **	-1.068 *	-1.140 **	-0.690 *
		(-3.09)	(-2.49)	(-4.13)	(-2.38)
Intercept		-0.848 #	2.009	-1.636 **	2.585 *
		(-1.81)	(1.16)	(-3.06)	(2.17)
SIZE	–	–	-0.385 **	–	-0.463 **
			(-3.19)		(-5.39)
RVAR	+	–	7.752	–	5.427 *
			(1.38)		(2.09)
LEV	+	–	4.183 **	–	4.351 **
			(3.59)		(5.59)
Industry-Fixed Eff.		included	included	included	included
R^2		10.0%	16.3%	8.6%	13.7%
F-Stat		3.3	3.6	5.8	6.9

The sample for the difference-in-differences analysis comprises only observations from newly listed ADR firms over the twelve-year period from 1992 to 2003. COC is the mean of four estimates for the implied cost of equity capital (see Appendix). We subsequently subtract the country-year medians of non-ADR firms from all COC estimates. PP stands for firms with private placements, OTC for firms trading in the over-the-counter markets, and EXCH for firms listed on the NYSE, Nasdaq or AMEX. We assume that the market learns about the ADR in year -1 or 0. SIZE is the natural log of US$ total assets (in thousands). RVAR is the return variability computed as annual standard deviation of monthly stock returns. Financial leverage (LEV) is measured as the ratio of total liabilities to total assets. Accounting data are measured as of the fiscal-year end, RVAR and the cost of capital as of month +10 after the fiscal-year end. Industry-fixed effects based on the classification in Campbell (1996) are included in the regressions but not reported. Panel A reports median (mean) values before and after a firm's first ADR and its upgrading to an exchange listing. We evaluate differences (post minus pre) using Wilcoxon rank sum tests (t-tests). Panel B reports OLS coefficient estimates and (in parentheses) t-statistics based on robust standard errors that are clustered by firm. For expositional purposes we multiply all coefficients by 100. **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 5

Differences in the Cost of Capital Effects by Home-Country Institutions

$$COC_{it} = \alpha_0 + \alpha_1 PP_{it} + \alpha_2 OTC_{it} + \alpha_3 EXCH_{it} + \alpha_4 INFL_{it} + \alpha_5 SIZE_{it} + \alpha_6 RVAR_{it}$$
$$+ \alpha_7 LEV_{it} + \sum\alpha_j \text{Year Controls}_t + \sum\alpha_k \text{Industry Controls}_i + \sum\alpha_l \text{Country Controls}_i + \varepsilon_{it}$$

	Predicted	*Legal Tradition*		*Disclosure Regulation*		*Securities Law Enforcement*		*Investor Protection*	
Variable	*Sign*	*code law*	*common law*	*low*	*high*	*low*	*high*	*low*	*high*
N		19,349	12,476	10,785	20,393	17,383	13,795	13,506	18,319
PP	+/–	0.185	0.919 *	-0.133	0.701 *[#]	0.003	0.713 *	-0.039	0.935 **[*]
		(0.77)	(2.05)	(-0.43)	(2.27)	(0.01)	(2.09)	(-0.15)	(2.63)
OTC	–	-0.109	-0.487 *	-0.212	-0.410 *	-0.335	-0.330	-0.107	-0.404 *
		(-0.50)	(-2.14)	(-0.80)	(-2.05)	(-1.51)	(-1.45)	(-0.41)	(-2.04)
EXCH	–	-0.380 *	-0.140	-0.737 **	-0.128 [*]	-0.645 **	-0.116 [#]	-0.766 **	0.023 [*]
		(-2.06)	(-0.57)	(-3.21)	(-0.68)	(-3.22)	(-0.53)	(-3.36)	(0.12)
Intercept		14.130 **	12.678 **	13.474 **	12.588 **	14.562 **	12.620 **	12.484 **	13.428 **
		(6.82)	(20.87)	(6.06)	(26.88)	(7.01)	(21.33)	(5.70)	(27.31)
INFL	+	23.209 **	12.804 **	23.751 **	16.115 **	16.765 **	20.302 **	19.282 **	15.630 **
		(8.11)	(4.49)	(6.18)	(5.99)	(5.49)	(6.75)	(5.26)	(6.52)
SIZE	–	-0.576 **	-0.569 **	-0.520 **	-0.576 **	-0.572 **	-0.542 **	-0.514 **	-0.612 **
		(-17.64)	(-12.75)	(-11.52)	(-17.55)	(-16.82)	(-12.84)	(-12.59)	(-17.49)
RVAR	+	14.709 **	19.277 **	14.521 **	17.533 **	14.386 **	18.709 **	14.601 **	18.066 **
		(15.62)	(15.94)	(10.87)	(19.04)	(14.10)	(16.44)	(12.29)	(18.86)
LEV	+	5.324 **	4.006 **	5.633 **	4.487 **	5.638 **	3.856 **	5.406 **	4.473 **
		(20.36)	(12.05)	(14.47)	(18.22)	(20.49)	(12.05)	(15.17)	(17.94)
Industry-, Country- & Year-Fixed Effects		included	included	included	included	included	included	included	included
R^2		38.7%	30.8%	32.2%	38.3%	42.5%	25.8%	31.6%	39.2%
F-Stat		103.9	55.6	44.9	143.3	119.0	46.4	56.5	117.3

The sample comprises a maximum of 31,825 firm-year observations from 44 countries over the twelve-year period from 1992 to 2003. We partition the sample into sub-samples according to the following institutional characteristics in the country of origin: (1) countries with code versus common law legal tradition (La Porta et al., 1997; Ball et al., 2000), (2) disclosure regulation is equal to high for countries with above-median index values of disclosure requirements in securities offerings (La Porta et al., 2006), (3) securities law enforcement is equal to high for countries with above-median levels of private and public enforcement of laws governing securities offerings (measured as the mean of the liability standard index and the public enforcement index from La Porta et al., 2006), and (4) investor protection is equal to high for countries with above-median anti-director rights index values (La Porta et al., 1998). The dependent variable, COC, is the mean of four estimates for the implied cost of equity capital (see Appendix). Three binary variables indicate ADR observations: (1) PP is equal to one if the firm has a private placement under Rule 144A, (2) OTC is equal to one if firm shares trade in the over-the-counter markets, and (3) EXCH is equal to one if firm shares are listed on the NYSE, Nasdaq or AMEX. INFL is the yearly median of country-specific, one-year-ahead realized monthly inflation rates. SIZE stands for US$ total assets (in thousands). RVAR is the return variability computed as annual standard deviation of monthly stock returns. Financial leverage (LEV) is measured as the ratio of total liabilities to total assets. Accounting data are measured as of the fiscal-year end, RVAR and the cost of capital as of month +10 after the fiscal-year end. We use the natural log of the size variable in the analysis and delete 1% of extreme realizations for all other firm-level attributes. Industry-fixed effects based on the classification in Campbell (1996), country- and year-fixed effects are included in the regressions but not reported. The table reports OLS coefficient estimates and (in parentheses) t-statistics based on robust standard errors that are clustered by firm. The table also indicates [in square brackets] whether the ADR coefficients across the two subsamples are statistically different from each other. For expositional purposes we multiply all coefficients by 100. **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

TABLE 6

Decomposition of Two-Year Price Changes around U.S. Cross Listings

$$\Delta P_{CFit} \text{ or } \Delta P_{COCit} = \alpha_0 + \alpha_1 PP_{-3/-2it} + \alpha_2 OTC_{-3/-2it} + \alpha_3 EXCH_{-3/-2it} + \alpha_4 SIZE_{it} + \alpha_5 RVAR_{it} + \alpha_6 LEV_{it} + \sum \alpha_k \text{Industry Controls}_i + \varepsilon_{it}$$

Panel A: Univariate Analysis of Cash Flow Effects and Cost of Capital Effects

	Price	*Years -3/-2*			
Variable	*Effect*	*N*	*Median*	*Mean*	*Std. Dev.*
PP	ΔP	63	15.93%	18.02%	42.01%
	ΔP_{CF}		2.02%	8.93%	45.43%
	ΔP_{COC}		-0.95%	3.70%	48.57%
OTC	ΔP	173	8.12%	6.38%	44.03%
	ΔP_{CF}		5.78% #	5.81% #	42.13%
	ΔP_{COC}		-3.13%	-0.08%	39.48%
EXCH	ΔP	175	19.13%	17.93%	42.84%
	ΔP_{CF}		7.60% **	10.18% **	33.04%
	ΔP_{COC}		8.87% **	10.10% **	39.80%

Panel B: Regression Analysis of Cash Flow Effects and Cost of Capital Effects

	Cash Flow Effects (ΔP_{CF})		Cost of Capital Effects (ΔP_{COC})	
Variable	*Model 1*	*Model 2*	*Model 1*	*Model 2*
N	3,875	3,772	3,875	3,772
$PP_{-3/-2}$	7.858 #	10.781 *	3.471	3.072
	(1.74)	(2.30)	(0.74)	(0.63)
$OTC_{-3/-2}$	6.661 *	6.993 *	-0.355	-0.803
	(2.35)	(2.42)	(-0.12)	(-0.27)
$EXCH_{-3/-2}$	10.979 **	9.822 **	7.548 *	6.419 *
	(3.68)	(3.23)	(2.43)	(2.02)
Intercept	-3.381	-5.895	2.937	2.635
	(-1.30)	(-0.86)	(1.08)	(0.37)
SIZE	–	0.271	–	0.049
		(0.60)		(0.10)
RVAR	–	-32.153 *	–	-9.117
		(-2.40)		(-0.65)
LEV	–	3.616	–	1.121
		(0.88)		(0.26)
Industry-Fixed Eff.	included	included	included	included
R^2	1.0%	1.2%	0.7%	0.7%
χ^2	40.0	46.6	28.4	26.4

The sample comprises ADR observations from 1992 to 2001 only because we decompose the two-year-ahead price change into a cash flow- and cost of capital-component. There is a maximum of 3,875 firm-years with both, cost of capital and cash flow effects available. We calculate three different price change variables: (1) ΔP is the two-year-ahead stock return using actual stock prices, (2) ΔP_{CF} is the simulated two-year-ahead percentage change in price due to changes in cash flow expectations using COC estimates in year t and earnings forecasts in year t+2, and (3) ΔP_{COC} is the simulated two-year-ahead percentage change in price due to changes in cost of capital expectations using earnings forecasts in year t and COC estimates in year t+2. All price change variables are computed as the difference between the natural log of price in t+2 and t. We subtract the country-year medians of non-ADR firms from all price changes, and delete the first and 99th percentiles of the median-adjusted price changes. PP stands for firms with private placements, OTC for firms trading in the over-the-counter markets, and EXCH for firms listed on the NYSE, Nasdaq or AMEX. We assume that the market learns about the ADR in year -1 or 0. Hence, we compute the two-year-ahead price changes for the years -3 and -2 as indicated by the subscripts. See Table 4 for a description of the remaining controls. Panel A reports the mean, median and standard deviation of the price change variables. Panel B reports coefficient estimates and z-statistics (in parentheses) from seemingly-unrelated regressions of the cash flow and cost of capital effects on the ADR indicators and controls. For expositional purposes we multiply all coefficients by 100. **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.