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 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 significant decrease in their cost of capital between 70 to 120 basis points. These effects are sustained and still present in recent years and after the passage of the Sarbanes-Oxley Act. Consistent with the bonding hypothesis, we find smaller cost of capital reductions for firms that cross-list in the over-the-counter market and for exchange-listed firms from countries with stronger home-country institutions. For exchange-traded cross-listings, the reduction in cost of capital accounts for more than half of the increase in value around cross-listings, whereas for the other types of cross-listings the valuation effects are primarily attributable to contemporaneous revisions in growth expectations.

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

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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, 2008a). 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., 2007). Similarly, U.S. cross-listings can improve investor recognition and enlarge a firm's investor base (e.g., Merton, 1987; Foerster and Karolyi, 1999).

A potential concern about the documented valuation effects of U.S. cross-listings is that they merely reflect concurrent changes in firms' growth opportunities that do not stem from cross-listing per se. That is, firms may seek cross-listings when they experience an expansion in their growth opportunities, but the decision is unrelated to bonding and the growth expansion does not reflect a reduction in the cost of capital due to cross-listing. 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 debate about delistings from U.S. exchanges and the costs of the Sarbanes-Oxley Act (SOX) questions the existence of sizeable cross-listing benefits, such as a reduction in the cost of capital (Zingales, 2007; Hostak et al., 2007). Thus, it is still 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 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 40,000 firm-year observations from 45 countries over the period from 1990 to 2005. We collect a comprehensive sample of 1,097 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 market do not require a 20-F filing, but a registration statement using Form F-6 and home-country disclosures to the SEC. They are also subject to Rule 10b-5 and the Foreign Corrupt Practices Act, under which most SEC enforcement actions as well as private class action suits are brought (Karpoff et al., 2008). Private placements under Rule 144A do not require SEC registration or any additional (public) disclosures. Given these regulatory consequences, we hypothesize that, if cross-listings reduce firms' cost of capital, the effects are strongest for exchange listings, and it is not clear that private placements should experience any reduction.

Consistent with this hypothesis, we find strong evidence that cross-listings on U.S. exchanges (AMEX, NASDAQ and NYSE) 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 of changes in the cost of capital, mitigating concerns about omitted variables and selection on unobservable characteristics. Most regressions 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 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 effects for exchange listings. U.S. private placements exhibit insignificant changes and, in some of our analyses, an increase in the cost of capital. This result is consistent with the findings in Miller (1999) and Doidge et al. (2004, 2008a) as they also document opposite or insignificant valuation effects for private placements. One possible explanation for the elevated cost of capital 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 weaker disclosure regulation and weaker protection against self-dealing by corporate insiders. We show that the cost of capital effects are sustained for many years after the cross-listing and that they are still present after the passage of SOX. In contrast, we do not find significant cost of capital effects for cross-listings on the London Stock Exchange. Both of these findings are consistent with recent evidence in Doidge et al. (2008a).

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 as well as aggregating (and weighting) the estimates from the four models. 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, even prior to the cross-listing. To address this issue, we implement our models with long-run growth estimates that vary by firm, cross-listing status, country and/or year, and obtain very similar results. Further, we control for differences in analyst forecast bias across firms and countries, and check that the results are not unduly affected by missing or negative earnings forecasts. Lastly, we explore the sources of the cost of capital effects and find little evidence that the results are primarily driven by increases in the shareholder base or market liquidity after cross-listing, strengthening the conclusion that bonding plays an important role for our results.[1]

Our final set of analyses exploits that financial analysts provide explicit forecasts for firm growth. Using these forecasts, we decompose the realized three-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 revisions in growth expectations for all three cross-listing types. However, for cross-listings on U.S. exchanges, the reduction in cost of capital accounts for more than half of the increase in value around cross-listings, whereas for the other types of cross-listings the valuation effects are primarily, if not solely, attributable to an expansion in firms' growth opportunities.

This study makes several contributions. First, we use a novel approach to provide evidence

[1] That said, it should be noted that stronger bonding can lead to an increase in the shareholder base and more market liquidity and, hence, it is difficult to disentangle the bonding hypothesis from alternative cross-listing hypotheses (see also Karolyi, 2006; Stulz, 2008). However, doing so is not the primary goal of this paper. Its main purpose is to first establish the existence of cost of capital effects (which can have multiple sources).

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 Sarkissian and Schill (2008) estimate an effect of 800 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 around cross-listings 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, as cross-listings are major corporate events, it is difficult to obtain equilibrium estimates for expected returns, especially considering that cross-listings change firms' exposures to the global market portfolio (Foerster and Karolyi, 1999). Our approach shows a way to overcome these difficulties.

Second, we provide evidence that cross-listings are associated with significant valuation effects stemming from revisions in cash flow (or growth) expectations. A concern is that analyses using market values, Tobin's q or stock returns pick up these effects, even when they are not related to cross-listing per se, thereby potentially overstating the benefits of U.S. cross-listings. However, shocks to firms' growth opportunities that are concurrent but not incidental to the cross-listing are unlikely to be sustained, whereas changes in growth or the cost of capital that are related to the cross-listing should be persistent. Thus, by documenting a sustained reduction in the cost of capital, we also corroborate earlier findings based on firms' valuations. Moreover, our evidence on large and sustained benefits from U.S. exchange listings contributes to the recent debate about post-SOX regulatory costs and its consequences for the competitiveness of U.S. capital markets (Zingales, 2007; Doidge et al., 2008a; Piotroski and Srinivasan, 2008).

Third, our analysis is based on one of the largest panels of U.S. cross-listings. Prior studies are generally based on smaller samples that are constructed as of one point in time.[2] The panel approach not only mitigates survivorship bias, but also facilitates fixed-effects and difference-in-differences estimation, thereby mitigating concerns about unobserved heterogeneity and (time-invariant) selection bias.

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 home countries with different institutional frameworks. This evidence corroborates earlier evidence on the bonding hypothesis (e.g., Reese and Weisbach, 2002; Doidge et al., 2004), which several recent studies have questioned (e.g., Siegel, 2005; Gozzi et al., 2007).

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 dependent and independent variables. Section 4 presents the main 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. In the Appendix, we delineate the estimation procedure of the implied cost of capital measures.

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

[2] A notable exception is the recent study by Doidge et al. (2008a).

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., AMEX, NASDAQ and NYSE) generally have much stronger return effects than either cross-listings in other countries (e.g., Sarkissian and Schill, 2008) 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).[3] 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,

[3] 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.

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.[4] 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 effects combining all ADR types are stronger for firms from countries where investor protection is weaker.[5] 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. 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. Finally, Doidge et al. (2008a) document that the valuation premium for U.S. cross-listings persists, that it has not fallen in recent years, and that a listing in London does not offer comparable valuation benefits.

The bonding hypothesis suggests that U.S. cross-listings should increase market value but it can do so in several ways. On one hand, effective 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

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4 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).

5 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. (2008b) provide evidence suggesting that private benefits of control are an important reason why controlling insiders shy away from cross-listings on U.S. exchanges.

turn should expand the set of profitable investment opportunities.[6] Thus, the bonding hypothesis predicts both growth and cost of capital effects.

In addition, cross-listing on a U.S. exchange commits foreign firms to disclosure rules that are generally 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 can reduce information asymmetries and lower firms' cost of capital (Verrecchia, 2001; Lambert et al., 2007; Stulz, 2008).[7] A related argument is that cross-listing broadens a firm's investor base. Based on Merton (1987), this effect can also 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., 2008).

Thus, although there are several good economic reasons why U.S. cross-listings should reduce firms' cost of capital, there is relatively little direct empirical evidence thereof. The return-based results in Errunza and Miller (2000) and, more recently, in Sarkissian and Schill (2008) are too large to be solely attributable to cost of capital effects and, while the valuation effects in prior studies are consistent with a reduction in the cost of capital, they neither provide evidence ipso facto that U.S. cross-listings reduce firms' cost of capital nor do they shed much light on the mechanism by which cross-listing affects firms' valuations.

Furthermore, it is possible that firms with new growth opportunities choose to cross-list for product market reasons, and not because they are financially constrained at home or because they

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6 The models in Lombardo and Pagano (2002) and Lambert et al. (2007) also support the notion that less expropriation can manifest in the cost of capital.

7 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; Daske et al., 2008). In addition, Hail and Leuz (2006) provide evidence that strong securities regulation is associated with a lower cost of capital.

need bonding to exploit these opportunities. In this case, the observed valuation effects do not stem from cross-listing per se and they are unlikely to be sustained for a long time. Moreover, in this instance, it is neither clear that the effects should manifest in a decrease in the cost of capital nor that the growth effects exhibit the typical rank order across ADR types predicted by the bonding hypothesis. Thus, by separately analyzing growth and cost of capital effects, we should be able to disentangle the two hypotheses and shed further light on the sources of the valuation benefits.

In sum, we first examine whether U.S. cross-listings have an effect on firms' cost of capital, controlling for changes in growth expectations. 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. As OTC listings and private placements have only weak regulatory consequences, we expect them to have more modest or no cost of capital effects, perhaps due to decreased market segmentation. We also predict that the cost of capital effects are stronger for firms from countries with weaker legal institutions and less extensive disclosure regulation.

Finally, we analyze whether there are significant valuation effects driven by changes in growth expectations. Towards this end, we decompose long-run realized stock returns around the cross-listing into two components, i.e., one measuring the effect of revisions to cash flow (or growth) expectations and one measuring the effect of changes in the cost of capital. However, as a reduction in the cost of capital should expand the set of profitable investment projects, it is difficult to completely disentangle the two components. For this reason, we view this part of our analysis as exploratory in nature and as an attempt to shed further light on the mechanism through which cross-listings affect firm value. With this caveat in mind, we gauge the relative magnitudes of the cash flow and cost of capital effects and also compare them across ADR types. Under the bonding hypothesis, we expect the same rank order for the growth effects across ADR types as for the reduction in cost of capital. However, if the growth effects also reflect other factors unrelated to bonding, the predicted rank order is not clear.

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). Moreover, cross-listings are major corporate events making it particularly difficult to obtain equilibrium estimates of expected returns. 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 accounting-based valuation 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 it does not require any a priori assumptions about the degree of market integration. Furthermore, implied cost of capital models make an explicit attempt to separate cash flow (or growth) effects from cost of capital effects.

We adopt four implied-cost-of-capital models that are commonly used in the literature (Claus and Thomas, 2001; Gebhardt et al., 2001; Easton, 2004; Ohlson and Juettner-Nauroth, 2005). All four models are consistent with the discounted dividend valuation model but exploit basic accounting relations to obtain an *equivalent* valuation equation based on residual income or abnormal earnings. We can then substitute the market price and analyst forecasts into the valuation equation and 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 the cost of equity capital, given the 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, explain how we implemented the models in an international setting, and gauge the impact of the numerical procedures used to iteratively solve the models.

Despite their appeal for our setting, implied cost of capital models have methodological and practical limitations. For instance, they require the assumption that consensus analyst forecasts are reasonable proxies for the market's expectations of future earnings, which might not always be the case (e.g., Frankel and Lee, 1998; Easton and Sommers, 2007). The models also limit the sample to firms that are covered by financial analysts and have positive earnings forecasts. We gauge the sensitivity of our results to these assumptions and data requirements. In addition, there are concerns that implied cost of capital estimates suffer from substantial measurement error, e.g., due to noisy or slowly updated analyst forecasts (Easton and Monahan, 2005; Guay et al., 2005).[8]

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

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[8] While the evidence in Guay et al. (2005) raises concerns, it also shows that implied estimates are not worse than realized returns and, in some cases, perform even better. Recently, Pastor et al. (2008) demonstrate that implied estimates are useful in capturing time-series variation in expected returns.

proxies and use the resulting mean estimate as our primary dependent variable (COC). Averaging should reduce idiosyncratic measurement error across the models. But we also conduct analyses using estimates from each model separately, the first principal component of the estimates from the four models, and weighting the estimates by the quality of the input parameters or the fit of the individual models, which should further address measurement error concerns.

Finally, given the debate about the validity of implied estimates, we benchmark the mean COC estimate to a set of alternative, non-accounting based proxies for firms' cost of capital, namely dividend yields and expected returns extracted from realized returns and countries' credit risk ratings (as described in Erb et al., 1996). In each case, the Pearson correlation at the country-year level is substantial, highly significant, and exhibits the predicted sign, suggesting that the implied cost of capital estimates are reasonable (see also Hail and Leuz, 2006).

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 1990 to 2005, except 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.[9]

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

9 See King and Segal (2004) for an analysis of U.S. cross-listings of Canadian firms. If we include Canadian firms in our cross-sectional regressions, the results are similar and the inferences do not change.

or a three-year-ahead earnings forecast, and a contemporaneous share price.[10] 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 well after the fiscal-year end to assure that financial data are publicly available and reflected in the market price at the time of our computations.[11] To calculate our average cost of capital estimate, COC, we require that each firm-year has all four individual implied cost of capital proxies available, and apply a series of common sense filters to ensure that the subsequent results are not driven by a mismatch in the underlying Worldscope and *I/B/E/S* input databases (see Appendix).

We further delete 1% at the top and bottom of all firm-level attributes (dependent and independent variables), except for those with a natural upper or lower bound and firm size where we use the natural logarithm in the analyses. Finally, we eliminate firm-years from countries with no ADR observations or with fewer than 20 firm-year observations (before any data filtering), with market values or total assets smaller than US$ 10 millions, 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 40,497 firm-year observations from 45 countries over the period from 1990 to 2005 representing a total of 9,493 individual firms.

Next, we compile a comprehensive dataset of active and inactive U.S. cross-listings using information from Citibank, JP Morgan, Bank of New York, Datastream and Bloomberg. Including ADRs that have become inactive mitigates concerns about survivorship bias, which would otherwise arise. We manually cross-check the datasets, and construct binary variables indicating

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

11 Note that the exact measurement date should not matter for our analysis (as long as the financial statements are available to investors for all firms). That is, the cost of capital differential across firms with ADRs and those without should not be systematically different at other measurement dates. We also compute cost of capital estimates as of month +7 and obtain very similar results.

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 1,694 unique foreign firms with an ADR that is active at some point between 1990 and 2005, which is a comprehensive panel of ADRs. After merging this dataset with the cost of capital estimates and our firm-year dataset, we have 6,048 ADR firm-years from 1,097 unique firms with U.S. cross-listings in our sample. Due to different 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. Firms with a cross-listing on a U.S. stock exchange have to file Form 20-F with the SEC, requiring extensive disclosures and, during our sample period, 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 may face legal liabilities from shareholder litigation. Cross-listings in the OTC markets do not require a 20-F filing, but have to file a registration statement using Form F-6 and home-country disclosures to the SEC. They are also subject to Rule 10b-5 and the Foreign Corrupt Practices Act, under which SEC enforcement actions and private securities litigation can be brought (Karpoff et al., 2008). 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. 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, companies with private placements 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 average magnitude of the estimates varies across the four individual valuation models, the estimates are all highly and significantly correlated with each other (pairwise correlation of 0.44 or higher) and with the aggregate COC estimates (pairwise correlation of 0.70 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 consumer price indices provided in the Datastream and World Bank 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 variability and financial leverage (e.g., Fama and French, 1992, 1993). We measure SIZE as total assets (in US$ millions) at the end of the fiscal year, 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 total assets as 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 higher valuations.[12] 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

[12] Consistent with this claim, the cross-listing indicators become insignificant or even significantly positive in Table 2 when we include the contemporaneous market value of equity instead of book value of total assets. However, simply lagging the market value by one or two periods restores the results for our firm-fixed effects regressions reported in Table 2 (Model 3), illustrating our point.

the world market portfolio (e.g., Harvey, 1995; Erb et al., 1996).[13] 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).

Next, we note that implied cost of capital estimates can be affected by forecast bias (Easton and Sommers, 2007). 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.[14] 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 control for forecast bias (FBIAS) using the one-year-ahead forecast error (forecasts minus actual values), scaled by lagged total assets, as proxy.

Finally, we use 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 instead. 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). Next, 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 selection bias based on unobservable, time-invariant firm characteristics. As it is a major concern

13 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.

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

that cross-listed firms differ from non-cross-listed firms in ways that are unobservable or at least difficult to measure, we adopt the firm-fixed effects model as our primary specification.

4. Main Results

4.1 Cross-Sectional Comparison of ADR and 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, Panel A, presents OLS coefficient estimates and t-statistics for these regressions. Our inferences are based on heteroscedasticity-corrected standard errors, which, when using industry- and country-fixed effects, are adjusted for firm-level clustering, to account for the fact that the same firm enters the sample multiple times. This assumes that standard errors are independent across firms, but not across time. In the firm-fixed effects regressions, we are even more conservative and cluster the standard errors by country and industry to allow for unspecified correlation in the error terms across time and for firms in the same country and industry.[15] 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

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[15] We prefer a cluster structure that is different from the fixed-effects structure and hence cluster differently across the country/industry-fixed and the firm-fixed effects models (Petersen, 2008). In unreported analyses, we also country-industry clustering for the country/industry-fixed effects model, use two-way clustering by firm and year for both fixed-effects structures, bootstrapped standard errors drawing from the same cluster structure as described above, and Fama-MacBeth (1973) regressions with country/industry-fixed effects. The inferences do not change.

significantly negative, and the latter exceeds the former in magnitude. This pattern continues to hold when we include the controls for inflation, risk, and analyst forecast bias in Model 2, but the magnitude of the OTC and EXCH coefficients substantially decreases to –0.30 and –0.44, 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 at the 10% level. This finding is somewhat surprising and discussed in more detail below. As expected, COC is positively related to the inflation rate, return variability, financial leverage and forecast bias, and negatively associated with firm size. All control variables are highly significant and, together with the year-, industry- and country-fixed effects, Model 2 explains about 35% 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 or at least difficult to measure (Model 3).[16] Together with the year-fixed effects, this specification amounts to a difference-in-differences design around cross-listings. 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 –0.83, suggesting that controlling for unobserved heterogeneity is important. All three ADR type variables are statistically significant and different from each other, and the model has substantial explanatory power (70%).

One concern is that cross-listings take some preparation and are often announced well before the actual ADR issuance (e.g., Foerster and Karolyi, 1999; Errunza and Miller, 2000; Sarkissian and

[16] 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.

Schill, 2008). As a result, markets are expected to price the cross-listings in advance of the actual listing. The last two models address this concern and attempt to account for such anticipation effects. In model 4, we delete the two firm-years immediately preceding the cross-listing and repeat our analysis. Consistent with the existence of anticipation effects, the results become more pronounced, i.e., the OTC and EXCH coefficients increase in magnitude and our inferences become even stronger, suggesting that models 2 and 3 underestimate the cross-listing effects. As an alternative way to capture anticipation effects, we set the three ADR-type variables equal to one in the year before the cross-listing (i.e., in year –1). Again, we find that the estimated cross-listing effects become stronger in magnitude (Model 5).

We gauge the economic significance of the effects using the last three models in Panel A of 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 83 and 132 basis points. The effect of cross-listings in the OTC markets is smaller in magnitude and between 40 and 58 basis points. These reductions in the cost of equity capital are economically significant and translate into a substantial increase in firm value. However, compared to prior return-based evidence, these estimates are rather modest. 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 (2008) 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 annual return differential of about –800 basis points, which is still too large to be plausible.[17] 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 suggests that private placements can be associated with an *increase* in the cost of capital, though it should be noted that the coefficient is not always significant. This result is somewhat surprising, but consistent with prior studies where private placements have insignificant effects or often stand out with opposite findings (e.g., Miller, 1999; Foerster and Karolyi, 2000; Doidge, 2004; Bailey et al., 2006; Doidge et al., 2008a). 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.[18]

4.2 Are the Estimated Cross-Listing Effects Sustained Over Time?

It is interesting to analyze whether the documented cost of capital benefits are sustained, i.e., still present years after the initial cross-listing, or rather short-lived in nature. Given the recent

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[17] Extending the window to 20 years (and excluding 10 years surrounding the cross-listing) reduces the decline to 150 basis points (Sarkissian and Schill, 2008). 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.

[18] 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.

debate about the competitiveness of the U.S. capital markets and their attractiveness to foreign companies (e.g., Interim Report of the Committee on Capital Market Regulation, 2006; Zingales, 2007), it is also interesting to examine whether the effects are still present after the passage of SOX. It has been argued that SOX has imposed substantial costs on foreign firms that are cross-listed in the U.S. without commensurate benefits, prompting firms to delist in the U.S. or to seek listings in other capital markets, in particular the London Stock Exchange. In this section, we perform analyses that relate to these questions (see also Hostak et al., 2007; Doidge et al., 2008a; Piotroski and Srinivasan, 2008).

First, we trace out the cost of capital effects in the years surrounding the cross-listing. That is, we replace each of the three ADR-type variables by a series of indicator variables marking the individual years around the respective cross-listing as well as the periods beyond year –3 and beyond +4. Panel B of Table 2 provides the coefficient estimates. In the specification with industry- and country-fixed effects, we find significant cost of capital effects starting the year before ADR, consistent with the idea that ADRs are anticipated and priced before their actual issuance (see also models 4 and 5 in Panel A). For both OTC and exchange listings, the cost of capital effects are significant beyond year +4 after the cross-listing and have a similar magnitude as in Model 2 of Panel A, indicating that the reduction in the cost of capital is sustained over time. Private placements exhibit significantly positive coefficients for only a few years after the listing and then fades, suggesting that the increase in the cost of capital is short-lived and occurs around the capital raising. The time-series pattern in the firm-fixed effects specification is similar, except that the cost of capital effects arise even earlier for exchange listings (year –2) and that the OTC and EXCH coefficients even increase in magnitude over time. Moreover, the effects beyond year +4 are much larger than the coefficients in Model 3 of Panel A. The latter result, however, has to be interpreted cautiously as it stems from a smaller set of firms. In the firm-fixed effects regression, only firms with an uninterrupted time-series of data from at least year –3 to year +4 contribute to the estimation of the time-specific cross-listing indicators, and this subset of firms appears to have stronger effects than the population of ADRs in Model 3 of Panel A. Taken together, the results in Panels A and B show that OTC and exchange cross-listings have significant cost of capital benefits that arise shortly before the listing and are sustained over time.

Second, we examine whether the cross-listing benefits have faded since the passage of SOX and whether cross-listings on the London Stock Exchange confer similar benefits. We define an indicator variable, SOX, that takes on the value of one for all firm-year observations where the cost of capital is measured after September 2002. We introduce this variable as a main effect and an interaction with the ADR-type indicators. Panel C of Table 2 reports the respective coefficient estimates. Our main ADR type variables are largely unaffected by this addition. More importantly, we find no evidence that the cost of capital effects have been attenuated after the passage of Sarbanes-Oxley. If anything, the interaction terms of the ADR type variables with SOX are negative, indicating that U.S. cross-listings, especially on exchanges, still offer significant benefits. Finally, we include an indicator for cross-listings on the main market of the London Stock Exchange. We find that cross-listings in London are not associated with a significant decrease in cost of capital. The results are similar and inferences the same when we include cross-listings on the AIM market (untabulated). These findings emphasize the special role of U.S. cross-listings and the U.S. regulatory system. Both the post SOX and the London results are consistent with recent evidence in Doidge et al. (2008a) based on panel regressions for Tobin's q.

4.3 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 each of the four implied cost of capital models and different ways to aggregate or weight the four estimates. Second, we assess whether the issuance of new equity capital in the U.S. potentially affects our results. Next, we consider alternative model specifications using local risk-free interest rates or risk premiums instead of raw cost of capital estimates, 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, which can be critical and hence may affect our analysis. Table 3 reports only the coefficients of the three ADR type indicators (plus additional 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).

In Panel A, we first report results using estimates from each of the four models. Next, we use the first principal component of the four individual proxies as dependent variable, which is a more sophisticated way to reduce idiosyncratic measurement error across the four models than simply averaging. In the following set of tests, we use proxies for the quality of the cost of capital estimates to compute weights. First, we combine the four estimates to a weighted average based on model fit, i.e., how well the estimates relate to standard controls for risk and expected returns. We separately and by country regress estimates from each model on our set of control variables, and then use the R^2s from these country regressions as weights when averaging across all four models.[19] Second, we use the inverse of the absolute one-year-ahead analyst forecast error as regression weights, giving larger weights to more accurate forecasts. Finally, we use the inverse of the maximum number of iterations needed to obtain a numerical solution across the four implied cost of capital models for a given firm and year as regression weights, giving larger weights to estimates

[19] Our set of control variables consists of size, return variability, leverage and forecast bias. In unreported robustness checks, we also use an alternative, Fama and French (1993) inspired set of control variables (size, return variability and book-to-market ratio), without changing the tenor of the results. The same is true when we use the individual cost of capital estimate from the valuation model with the highest R^2 in a given country for each observation in that country (instead of the weighted average).

where the computational procedures converge faster (see also Appendix). The results for all these refinements are consistent with Table 2 and lead to the same inferences as before. Specifically, EXCH is always negative and significant at the 3% level or better. Moreover, the coefficient has a similar magnitude across models, ranging from –0.63 to –1.02 (except for the principal component regression which is on a different scale and cannot be interpreted in terms of basis points). The OTC coefficient is negative and significant in five out of the eight specifications. PP is positive but generally insignificant. These results demonstrate that our findings do not depend on the choice of a particular cost of capital model.

Our second set of robustness tests addresses the concern that raising new equity capital in the U.S. may affect the 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. 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 Model 3 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. (Level III ADRs) and those that cannot raise capital but are exempt from certain SEC rules (Level II ADRs). Under the bonding hypothesis, we expect the former to experience larger cost of capital benefits. Consistent with this expectation, we find that Level III firms exhibit the largest decrease in cost of capital (about 120 basis points), while the magnitude of the coefficient for Level II ADRs is significantly smaller (at the 12% level, two-sided), but still exceeds the decrease in the cost of capital for OTC firms in magnitude (although the difference is not statistically significant). All three coefficients are statistically significant at the 6% level or better.

In Panel C of Table 3, we report 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 control for inflation with a contemporaneous control for the (nominal) risk-free rate, measured as country-year median of the monthly yields of local treasury bills or, if unavailable, central bank papers, inter-bank loans or money market rates (obtained from Datastream and the World Bank). Second, we check that the results are similar when we use the risk premium as dependent variable, i.e., subtract the local risk-free interest rate 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 it does not force a coefficient of minus one on the proxy for the risk-free interest rate (or expected inflation), even though the construct is likely 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.

In the last row of Panel C, we address concerns about sample selection due to the fact that implied cost of capital models require analyst forecasts for at least two future periods as well as positive realizations for these forecasts and the long-term growth estimate. It is possible that these data requirements differentially screen out ADR and non-ADR firms, thereby biasing our results. 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 63,166 firm-year observations. Despite this substantial increase in sample size, the results are very similar to those reported before.[20]

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 the implied cost of capital models. In using consensus forecasts, implied cost of capital models explicitly account for market (or analyst) expectations about (short-term) growth, which is why they are appealing for our purposes. However, the valuation models also require assumptions about growth beyond the explicit forecast horizon, and these assumptions can have a substantial influence on the estimates. Furthermore, our primary implementation uses a country/year-specific rate for long-term growth and, hence, applies the same assumption across ADR and non-ADR firms. It is possible, if not likely, that ADRs have different long-run growth expectations (even prior to the actual cross-listing), which in turn can bias the cost of capital estimates when using the same long-run growth assumption across all firms in a given country and year. In response to these issues, we vary the long-run growth rate assumption in a number of ways (see also Appendix). First, we make ADR-type/year-specific *adjustments* to our standard long-term growth rate in pre- and post-listing periods, based on either forecasted or realized growth rates. The idea of this adjustment is to account for the possibility that ADR and non-ADR firms grow at different rates (e.g., due to bonding post listing or self-selection pre listing), and that the market prices likely reflect such growth differences, which, in turn, implies that these differences need to be taken into account when imputing the cost of capital. Second, we separately compute growth in terminal values for *each firm*, regardless of its cross-listing status, based on the

20 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 alternative specifications do not materially affect our findings.

realized average growth rate in earnings per share.[21] Third, we use analysts' consensus long-run growth in earnings per share estimates as a firm/year-specific proxy for terminal-value growth, and thereby refrain from using any common growth assumptions across firms or groups of firms. Finally, we also include the one-year-ahead earnings growth rate based on analyst forecasts as additional control variable in the regression, and allow this coefficient to vary across ADR type, i.e., separately interact it with each ADR indicator.

As Panel D of Table 3 shows, the cross-sectional results are remarkably stable across all five modifications. The three coefficients on the cross-listing indicators have the same sign and rank order as before. The coefficients on EXCH and OTC are negative and statistically significant for all modifications, and exhibit a similar magnitude as in Table 2, Panel A. Given these extensive checks, it is unlikely that differential long-run growth expectations drive our findings.

Finally, we acknowledge that cross-listed firms are not randomly drawn. Thus, our cross-sectional tests may suffer from selection bias. Firm-fixed effects address this issue, if the selection is based on time-*invariant* firm characteristics, even if they are unobservable or difficult to measure. However, firms likely choose to cross-list when their financing needs or growth opportunities change and, hence, we are also concerned about time-variant selection, which is not captured by our time-variant control variables. However, addressing this selection problem is very difficult (e.g., Wooldridge, 2001; Larcker and Rusticus, 2008). It requires one or more instruments, which are generally hard to find – the key challenge being the exclusion restriction. In addition, cross-listings in the U.S. are not easily reversed, which implies that the choice in any given period is not

[21] To avoid any hardwiring of systematic biases between ADR and non-ADR firm-year observations, we use only years prior to the cross-listing for the ADR firms to determine average growth.

independent of prior choices. The standard Heckman procedure does not account for such dependencies.[22] Thus, the remaining concerns about self-selection are a caveat to our findings.

4.4 Difference-in-Differences Analysis of Changes in the Cost of Capital

In this section, we focus solely on firms that initiate a U.S. cross-listing over the sample period and examine changes in the cost of capital around cross-listings. By analyzing changes, each firm serves as its own control. Second, we standardize the cost of capital before and after the listing by subtracting the country/year median cost of capital of all non-ADR firms to control for macroeconomic or time trends as well as differences in these trends across countries. This pre- versus post-listing comparison of standardized (i.e., median-adjusted) cost of capital estimates is again in the spirit of a difference-in-differences analysis and alleviates concerns about unobserved heterogeneity across firms or time-invariant selection bias. Based on the results in Table 2, Panel B, we remove all observations from the year of and before the cross-listing (years –1 and 0) to mitigate concerns about announcement effects and to measure long-run changes in the cost of capital. The resulting 2,672 firm-years are the basis of our changes analysis presented in Table 4.

Panel A reports the mean and median of the standardized costs 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 years –2

[22] Nonetheless, we assess the potential impact of self-selection with two-step treatment-effects models for each ADR type for the country/industry-fixed and the firm-fixed effects specification using the full panel. In the first stage, we use firm size, leverage, average asset growth over the past two years, return on assets, foreign sales, percentage of closely held shares, English legal origin, and year-fixed effects, and has significant explanatory power. The two variables serving as potential instruments are foreign sales and the fraction of closely held shares. Both variables predict U.S. cross-listings and it is not immediately obvious that they have a direct impact on the cost of capital, other than through cross-listing. The second stage produces a significantly negative coefficient on EXCH and negative but insignificant coefficients on OTC and PP. Based on these findings, which should be interpreted cautiously, our inferences do not appear to suffer from a major selection problem.

and +1 around the cross-listing. The next two columns compare all available firm-years before and after ADR initiation, except those from years –1 and 0. By using all other firm-year observations, we check again whether 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 differences in the differences in the first two columns (post – pre) are not significant, but except for OTC they point in the same direction as the cross-sectional analysis, suggesting that small sample size and greater noise in the changes are responsible for the insignificance. Using all years (except -1 and 0), the difference-in-differences estimates for EXCH and OTC are statistically significant. Moreover, they exhibit the hypothesized rank order and are of similar magnitude as in the cross-sectional analyses. Exchange listings are associated with a decrease in the cost of capital by about 110 basis points and ADRs initiated in the OTC markets experience a reduction around 50 basis points. The table also highlights that firms that later cross-list in the U.S. already have a lower cost of capital than non-ADR firms in the pre-listing period (i.e., the pre-listing median and average is generally negative), illustrating the selection issue in comparing ADR and non-ADR firms.

Panel B reports coefficient estimates from regressing the standardized COC on ADR-type indicators, industry-fixed effects, and our standard firm-level controls. Variables 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 cost of capital 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 analysis tie nicely into our cross-sectional findings presented in Section 4.1. The negative effect of an exchange listing is significant at the 8% level or better in all specifications and between 55 and 95 basis points. The reduction in cost of capital is less pronounced but still significant in three out of four cases for OTC cross-listings and between 40 and 60 basis points. Private placements exhibit a positive coefficient as before but the effect is never significant.

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 important 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 U.S. cross-listings also exhibit significant valuation effects stemming from changes in growth (or cash flow) expectations. We exploit analysts' forecasts for growth before and after the cross-listing and use these estimates to decompose changes in firm value around the cross-listing into a growth and a cost of capital component.

5.1 Cross-sectional Differences by Home-Country Institutions

We begin by analyzing whether institutional and market characteristics of firms' home countries are associated with cross-sectional differences in the benefits of cross-listing as suggested by the bonding hypothesis. To test this argument, we partition the sample into 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 legal tradition (La Porta et al., 1997; Ball et al., 2000), (2) disclosure regulation is set equal to high for countries with above-median index values of disclosure requirements in securities offerings (La Porta et al., 2006), (3) regulation of self-dealing is equal to high for countries with above-median values of the anti-self-dealing index representing the legal protection of minority shareholders against expropriation by controlling insiders (Djankov et al., 2008), (4) insiders' private control benefits are assumed to be low in countries with below-median average block premiums paid in corporate control transactions (Dyck and Zingales, 2004), and (5) equity market development is set equal to high for countries with above-median percentage ranks averaging across three La Porta et al. (1997) market development variables (i.e., the ratio of the aggregate stock market capitalization held by minorities to gross national product, the number of listed domestic firms per capita, and the number of initial public offerings per capita). With all these proxies, we 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 being subject to U.S. regulation and markets.

Table 5 provides coefficient estimates, t-statistics and significance tests across sub-samples. The regressions are based on Model 2 (see Table 2).[23] As before, the control variables behave as predicted and are highly significant. Also in line with expectations, the coefficient on EXCH is negative and highly significant for firms from countries with code-law origin, low disclosure requirements, weak regulation limiting insiders' self-dealing, high block premiums, and with less developed equity markets. Moreover, the estimated magnitude of the cost of capital reduction is always larger for firms in these countries than for firms from countries with strong institutions, and the EXCH coefficient is statistically different across sub-samples for two partitions at the 5% level (two-sided) and close to conventional levels for another partition. These findings support the bonding hypothesis.

A slightly different picture emerges with respect to ADRs in the OTC markets. In countries with weak institutions, the OTC coefficient is negative and smaller in magnitude than the EXCH coefficient, as expected and consistent with our earlier findings. But OTC firms seem to experience a larger reduction in cost of capital when they are from countries with common-law origin, high

[23] Using the firm-fixed effects specification for this analysis is quite demanding on the data as the ADR indicators are estimated solely from firms that experience changes over our sample period and so splitting the ADR sample considerably reduces power. Hence, it is not surprising that the differences across the two sub-samples (high vs. low) are generally insignificant.

disclosure regulation and high protection against self-dealing. Although the difference is never statistically significant across sub-samples and, hence, should be viewed cautiously, it is not inconsistent with the bonding hypothesis. As bonding is of lesser value to firms from home countries with 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 may serve this purpose but are less costly than exchange listings, making them attractive to firms with these other motives.[24] However, OTC cross-listings provide less assurance to outsiders in countries with weak institutions, implying that firms from these countries benefit less than firms from countries with a strong institutions, which in turn can explain the smaller OTC coefficient for the low protection sub-samples.

The PP coefficient is significantly positive and different from the other sub-sample for firms from common law countries and high disclosure and anti-self-dealing regulation. 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, which as explained is not inconsistent with the bonding hypothesis but suggests a role for other motives to cross-listing.

To shed light on some of these other cross-listing motives and other sources for the cost of

[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.

capital effects, we conduct two additional tests. First, we compute the change in ownership concentration around firms' exchange listings in the U.S. (using the percentage of closely held shares in the home market as proxy), and then split the ADR sample by the median. We find that the EXCH coefficient is very similar across both groups and, if anything, larger for firms with larger changes in ownership concentration. This finding suggests the decrease in the cost of capital is not primarily driven by shareholder base effects. Second, we split the ADR sample by the median level of average trading volume in the U.S. after the cross-listing. The idea is that firms with high U.S. trading volume (and hence larger liquidity increases) should experience larger cost of capital reductions. We find that the EXCH coefficients are similar across the two sub-samples, and if anything, slightly larger for the low U.S. trading volume group, inconsistent with liquidity being the primary driver behind our results. Thus, these two additional tests strengthen the conclusion that bonding likely plays an important role in our cost of capital results.

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 stock return around the cross-listing into a cash flow and a cost of capital component. We start with calculating the realized buy-and-hold return, ΔP, around the issuance of the ADR. Based on the observation that markets anticipate the cross-listing effects, we compare prices before the run-up has started, say in year t = –3, to prices that have already impounded the cross-listing effects, say in year t = 0 (i.e., the year when the firm starts to cross-list in the U.S.). This leads to the following continuously compounded three-year return:

$$\Delta \mathrm{P} = \ln\left(P_0 / P_{-3}\right), \tag{1}$$

where P_t is a firm's stock price in the forecast period month of year t. Based on the dividend discount model (which is also at the core of the accounting-based valuation models), we next note that, at each point in time, price is a function of the market's cash flow expectations and the cost of capital. Thus, similar to Campbell (1991) or Vuolteenaho (2002), we decompose total returns into a cash flow component and a cost of capital component:[25]

$$\Delta \mathrm{P} = \Delta \mathrm{P}_{\mathrm{CF}} + \Delta \mathrm{P}_{\mathrm{COC}} = \ln\left(P_{\mathrm{CF},0}/P_{-3}\right) + \ln\left(P_{\mathrm{COC},0}/P_{-3}\right). \quad (2)$$

$\Delta \mathrm{P}_{\mathrm{CF}}$ represents the change in price from P_{-3} to P_{CF} (measured at time 0) due solely to changes in cash flow expectations. Correspondingly, $\Delta \mathrm{P}_{\mathrm{COC}}$ represents the change in price from P_{-3} to P_{COC} (measured at time 0) due to revisions to the discount rate holding cash flow expectations constant.

Empirically, P_{CF} and P_{COC} are unobservable, and we rely on the accounting-based valuation models to infer "as-if" prices that let us approximate the cash flow and the cost of capital components. That is, while P_{-3} equals the actual market price at time $t = -3$, we use the implied cost of capital estimate from the same time period ($t = -3$) along with analysts' earnings forecasts from $t = 0$ to compute $P_{\mathrm{CF},0}$ and, consequently, $\Delta \mathrm{P}_{\mathrm{CF}}$. The idea of this imputation is to hold the cost of capital constant (and unaffected by the cross-listing), and to estimate the price effect that would have resulted if only the growth expectations had changed. The result is a three-year return reflecting (among other things) the cash flow (or growth) effects of cross-listing in the U.S.[26]

Similarly, we derive the as-if price change that is attributable to revisions in the cost of capital around the cross-listing, $\Delta \mathrm{P}_{\mathrm{COC}}$. We use our estimate of the implied cost of capital at time $t = 0$,

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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. Second, empirically, 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 an initial step to shed light on the complex mechanisms through which cross-listings affect firm value.

26 To illustrate the computation of $P_{\mathrm{CF},0}$ consider the modified PEG ratio model as described in the Appendix. Here, one- and two-year-ahead earnings forecasts, *fy1* and *fy2*, are taken from the year of the actual cross-listing ($t = 0$), while the discount rate is set equal to the implied cost of capital estimate, r_{PEG}, from year $t = -3$, leading to the following valuation formula: $P_{\mathrm{CF},0} = (fy2_0 + r_{\mathrm{PEG},-3} \cdot d_0 - fy1_0)/r^2_{\mathrm{PEG},-3}$, where d stands for dividends paid.

along with forecasted earnings as of $t = -3$ to compute $P_{COC,0}$ and then compare it to P_{-3}. Again, the underlying idea is that the updated cost of capital proxy already incorporates the effect of the cross-listing, while we deliberately select analyst forecasts that are not yet updated to reflect those changes. The result is a three-year return capturing (among other things) the cost of capital effects of cross-listing in the U.S.[27]

In aggregating as-if price changes across firms and in event time, we attempt to "wash out" other effects and extract the growth and cost of capital effects of cross-listing. In the spirit of a difference-in-differences design, 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 year. This adjustment is akin to market-adjusting returns. The issue is that all firms in the economy likely experience changes in the growth expectations and in the cost of capital over a three-year period and, hence, by standardizing ADR firms' as-if price changes, we attempt to strip out those general time-period effects (e.g., inflation).[28] With respect to timing, we build on the results in Table 2, Foerster and Karolyi (1999) and Errunza and Miller (2000), and assume that market prices reflect the cross-listing decision in year –1 or 0. Thus, we flag the three-year returns beginning in years –4 and –3 (and ending in years –1 and 0), and compare them to all the other three-year returns of ADR firms.[29]

Table 6 provides descriptive statistics for the total market-adjusted returns around the cross-listing as well as the as-if cash flow and cost of capital components. As we cannot compute returns for the last three years of the sample period, we lose the most recent observations. We further

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27 Applying the modified PEG ratio model, we compute $P_{COC,0}$ based on one- and two-year-ahead earnings forecasts, *fy1* and *fy2*, measured at t = –3, and combine them with the implied cost of capital estimate, r_{PEG}, derived in the year of the cross-listing ($t = 0$): $P_{COC,0} = (fy2_{-3} + r_{PEG,0} \cdot d_0 - fy1_{-3})/r^2_{PEG,0}$.

28 We note that standardization is not crucial for our results. When we re-estimate the regressions from Table 6 using raw price changes, the coefficient pattern is very similar, and none of the inferences changes.

29 Results flagging three-year returns beginning in years –4 through –2 relative to the cross-listing, or using two-year instead of three-year price changes (and flagging years –3 and –2) are similar, albeit slightly weaker for the EXCH coefficient in the ΔP_{COC} specification, consistent with earlier evidence that the effects begin before year –1.

eliminate the top and bottom one percent of the price change distributions to mitigate the influence of outliers. These choices leave us with returns for 4,342 ADR firm-year observations. Panel A reports the median, mean and standard deviation of the price change variables beginning in years –4 and –3. We also assess whether ΔP_{CF} and ΔP_{COC} are statistically different from zero and from each other. The market-adjusted total returns around the cross-listing range from 8.5% (median for private placements) to 15.7% (mean for exchange listings). The corresponding raw returns (i.e., without market-adjustments) are about 30% over three years (9% annualized) for all three ADR types, and slightly lower compared to prior studies (e.g., Foerster and Karolyi, 1999; Errunza and Miller, 2000). The high standard deviations indicate substantial variation in the price changes.

The average cash flow effects are similar in magnitude across all three ADR types, and significantly different from zero for all but one case (median PP). Exchange-listed ADRs exhibit 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 total return around OTC cross-listings stems primarily from improved growth expectations, rather than from a decrease in the cost of capital. The relation is even more extreme for private placements, for which the cost of capital effects are close to zero. However, the PP sample is small and we have to interpret these results cautiously.

Since the univariate comparisons do not control for differences in firm characteristics that are known to be associated with returns and changes in those characteristics around cross-listings, we employ regression analysis using the same model specifications as in Table 4. We re-code the ADR indicator variables to mark the three-year, market-adjusted cash flow and cost of capital effects beginning in years –4 and –3. 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, by market adjusting, relative to all non-ADR observations. To account for contemporaneous correlations in the error terms across the two price components, we estimate the two models simultaneously using seemingly unrelated regressions, and assess the significance of the coefficients using z-statistics.

We find that firms with an U.S. exchange listing exhibit significantly positive cash flow *and* cost of capital effects. The magnitude of the coefficients suggests that both effects contribute 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 for OTC and PP firms are negligible and insignificant.[30] Thus, our decomposition approach documents substantial valuation effects from revisions in growth expectations for all three cross-listing types. For cross-listings on U.S. exchanges, the reduction in cost of capital accounts for more than half of the increase in value around cross-listings, whereas for the other types of cross-listings the valuation effects are primarily, if not solely, attributable to an expansion in firms' growth opportunities.

6. Conclusion

In this paper, we examine the cost of capital effects of U.S. cross-listings for a large panel of ADR firms from 45 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 or disclosure theory would suggest. Alternatively, it is possible that the valuation effects merely reflect that firms cross-list when they experience an expansion in their growth opportunities, even though the latter is unrelated to the

[30] Due to lack of power, most coefficients between the two models and across ADR types do not differ from each other in statistical terms (except for the PP and OTC coefficients, which are significantly smaller than the EXCH coefficient in the ΔP_{COC} model).

cross-listing per se. 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 or dividend yields, as this approach makes an explicit attempt to account for revisions 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, analyst forecast bias, and firm-fixed effects. The magnitude and the ranking of the effects seem plausible with exchange listings experiencing a reduction in cost of capital between 70 and 120 basis points, followed by OTC listings with about 30 to 70 basis points. Firms that access U.S. markets via private placements exhibit no change or a small increase in cost of capital. These cost of capital effects are sustained over time, and we find no evidence that Sarbanes-Oxley has diminished the benefits of U.S. cross-listings. We do not find similar cost of capital effects for listings on the London Stock Exchange. When we investigate cross-sectional differences in the cost of capital effects, we find that firms from countries with weak disclosure regulation and weak protection of outside investors against self-dealing by corporate insiders benefit the most from cross-listing on U.S. exchanges. Overall, the ranking of the cost of capital effects across ADR types (from Level III to Level I) and the cross-sectional results lend support to the bonding hypothesis. Our evidence of sustained cost of capital effects also mitigates concerns that valuation effects documented by prior studies stem merely from growth shocks that are concurrent but unrelated to cross-listing per se.

That said there are significant valuation effects stemming from changes in (analysts') growth expectations for all three cross-listing types, indicating that not the entire effect is attributable to a reduction in the cost of capital. For cross-listings on U.S. exchanges, the reduction in cost of capital accounts for more than half of the increase in value around cross-listings, whereas for the other ADR types the valuation effects around cross-listings are primarily attributable to an expansion in firms' growth opportunities.

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, we cannot preclude 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 and countries in the documented way. There also remain concerns about self-selection, despite our efforts to control for the fact that cross-listed firms differ from non-cross-listed firms in observable and unobservable ways. Lastly, cost of capital and 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):

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

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

Gebhardt, Lee, and Swaminathan (2001):

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

Model-specific assumptions:
This is a special case of the residual income valuation model. It uses actual book values per share and forecasted earnings per share up to three years ahead to impute future expected residual income for an initial three-year period. We assume clean surplus, i.e., future book values are imputed from current book values, forecasted earnings and dividends. Dividends are set equal to a constant fraction of forecasted earnings. After the explicit forecast period of three years, the residual income series is derived by linearly fading the forecasted accounting return on equity to the 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):

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

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):

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

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

In this section, we provide a detailed description of the data requirements, our assumptions implementing the models and in case any data items are missing, a time line for measurement, and information on the reliability and speed of convergence of the numerical procedures.

We obtain stock price and analyst earnings per share forecasts including long-term earnings growth estimates from the *I/B/E/S* summary database. All estimates are mean analyst consensus forecasts and measured in local currency. 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). 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. Later, we gauge the sensitivity of our findings to this requirement.

Next, we gather financial data from the Worldscope database. Each of the four valuation models requires an estimate of future dividends. We assume net dividends ($\hat{d}_{t+\tau}$) as a constant fraction of expected future earnings per share for all periods. For each firm, we define k_t, the dividend payout ratio, as dividends per share (Field 05101) divided by earnings per share (Field 05201), averaged over the fiscal years –2 through 0. If k_t is missing or outside the range of zero and one, we replace it by the country/year median payout ratio. For the residual income valuation models we also need book value per share at the beginning of the valuation year (bv_t). We compute bv_t from Worldscope common equity (Field 03501) divided by *I/B/E/S* shares outstanding (multiplied by 1,000). Moreover, r_{GLS} requires an industry-specific target return on equity. We first compute firm-level returns on equity as earnings per share (Field 05201) divided by beginning book value per share (Field 05476), averaged over the fiscal years –2 through 0, and then take the median in a given country, industry and year as target ratio. To avoid sample attrition, we replace missing or negative target ratios by the country/industry median and, if still missing or negative, by the country/year median.

Even though Worldscope and *I/B/E/S* provide split-adjusted data and, hence, should be compatible, we apply three common sense filters to ensure that our results are not driven by a mismatch in input data. For each of the following variables involving data from both sources, we eliminate the first and 99th percentile of the respective distributions: (1) the *I/B/E/S* price to Worldscope book value per share (Field 05476) ratio, (2) the realized one-year earnings per share growth rate computed using next year's actual earnings per share from *I/B/E/S* and this year's earnings per share from Worldscope (Field 05201), and (3) the ratio of book value per share as reported in Worldscope (Field 05476) to book value per share calculated from Worldscope's common equity (Field 03501) divided by the number of shares outstanding from *I/B/E/S*.

The final ingredient is the long-run growth expectation (g or g_{lt}) in the terminal value. We assume that, in the long-run, firms grow at the country's inflation rate and use next year's country-specific median of the realized monthly percentage changes in the consumer price index (obtained from Datastream or, if unavailable, the World Bank) as a proxy for future inflation. As deflation cannot persist forever, we replace negative values by the country's historical inflation rate, computed as the median of the monthly inflation rates over the entire sample period. Similarly, we replace values exceeding 10% by the country's historical inflation rate.[31]

To ensure that, at the time of measurement, the market price and analyst expectations reflect financial statement information that we are using in the valuation models, we measure the cost of capital as of month +10 after the end of the fiscal year (*I/B/E/S* provides updates every third Thursday of each month). As a consequence, analyst forecasts collected in month +10 represent consensus estimates for fiscal years ending in 2, 14, 26, etc. months. The valuation models, on the other hand, assume discounting for a full year, i.e., they start at the beginning of the fiscal year. For consistent discounting, we first move month +10 prices (which contain the information available at the time of forecasting) back to the beginning of the fiscal year using the imputed cost of capital and then use full-year discounting.[32] This procedure merely shifts prices over time for proper discounting and is equivalent to applying partial-year discount factors to the forecasted valuation attributes (see e.g., Francis et al., 2000; Botosan and Plumlee, 2005).

Since most of the valuation models do not have a unique 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 (Table A1). We stop iterating if the imputed price falls within a 0.001 difference of its actual value. This cut-off value is chosen such that the remaining deviation from price is

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31 In untabulated tests, we gauge the sensitivity of the COC estimates to general changes in the assumptions about long-run growth using (1) a constant 3% inflation rate for all countries, or alternatively (2) a country's real GDP growth rate plus its long-run inflation rate. The latter assumes that countries' growth differences persist in perpetuity, while the former assumes that growth rates converge to a global competitive equilibrium with zero real growth. Our primary implementation (using countries' inflation rates) falls between these two extremes. Using either of the two extreme assumptions does not materially affect our results.

32 We discount the month +10 prices using $[1+r]^{-10/12}$ as discount factor, where r equals r_{CT}, r_{GLS}, r_{OJ} and r_{PEG}, respectively. As the imputed cost of capital is yet to be determined, this procedure requires an iterative process to solve the valuation equations provided in Table A1.

negligible with respect to a cost of capital measured with two decimal places. On average, the four models required 16 to 18 iterations to achieve convergence. In addition, the case of non-convergence was extremely rare with fewer than two out of 100 cases for r_{CT}, and fewer than three out of 1,000 cases for the other models. We further investigate whether the speed of convergence differs systematically across firms in a way that affects our analyses. That is, we regress the number of iterations on our standard set of controls and the three ADR-type indicators. None of the coefficients on PP, OTC, and EXCH are significant, suggesting that the iterative procedure has no undue influence on our coefficients of interest.

The estimation procedure described above yields 48,199 firm-year observations, for which all four cost of capital estimates are available (before any further data filtering, matching with the control variables, and outlier deletion).

As a sensitivity check, we vary our assumption about long-run growth beyond the explicit forecast horizon, where applicable (i.e., for r_{CT} and r_{OJ}), allowing for differential growth rates across ADR types and non-ADR firms: (1) We compute *ADR-type/year*-specific growth adjustments and add them to the long-run growth rate of the base version, which is next year's inflation rate. The adjustments are computed on a yearly basis as the difference in the median of analysts' long-run growth estimates (*ltg*) for each of the three ADR types and the median *ltg* for non-ADR firms. The adjustments apply only to ADR firms, but adjust the entire time-series, i.e., pre and post cross-listing periods. (2) As *ltg* reflects analysts' expectations about the future growth, which may not necessarily translate into actual growth differences, we repeat the calculations in (1) using realized yearly percentage changes in earnings per share (averaged over the fiscal years –2 through 0). (3) We separately construct a terminal value growth rate for *each firm*, regardless of its cross-listing status, and set it equal to the mean yearly percentage change in realized earnings per share over the entire sample period. For ADR firms, we only consider the pre-cross-listing years in the means' computation. (4) We directly use analysts' consensus long-run growth in earnings per share estimates (*ltg*), which are *firm*- and *year*-specific, as terminal-value growth proxy for all sample firms, and therefore refrain from using any common growth assumptions across firms or groups of firms. If *ltg* is missing or negative, we use the terminal value assumption from the base version.

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 Gebhardt et al. (2001) model as the long-run growth assumptions in the other models. Analogous to the growth adjustments, we extract *ADR-type/year*-specific adjustments, and add them to the target return on equity in the base version that varies by country, industry and year. We define these adjustments as yearly median differences in historic three-year average return on equity (computed as previously described) between each of the three ADR types and non-ADR firms. The adjustments apply to ADR firms only, but affect their entire time-series.

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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*	*FBIAS*
Argentina	141	68	12.17%	11.77%	0.75%	2,721	0.124	0.475	0.008
Australia	2,599	496	11.02%	10.26%	2.69%	4,672	0.081	0.492	0.005
Austria	272	98	12.63%	11.15%	1.77%	6,749	0.093	0.611	0.010
Belgium	630	32	11.23%	11.21%	1.88%	17,984	0.076	0.589	0.006
Brazil	315	158	17.63%	16.71%	6.62%	7,544	0.136	0.497	0.003
Chile	176	107	13.88%	12.69%	4.68%	3,634	0.087	0.510	0.011
China	414	73	11.91%	13.30%	0.52%	2,636	0.125	0.429	0.001
Czech Republic	68	14	15.37%	9.50%	6.43%	3,048	0.106	0.519	0.005
Denmark	767	30	11.57%	8.75%	2.11%	3,037	0.082	0.571	0.005
Egypt	32	15	19.71%	21.05%	4.54%	1,848	0.114	0.721	0.004
Finland	575	74	13.37%	12.05%	1.27%	1,959	0.100	0.528	0.001
France	3,048	339	11.17%	10.36%	1.65%	14,291	0.100	0.623	0.005
Germany	2,070	247	11.03%	10.23%	1.92%	22,757	0.107	0.629	0.008
Greece	264	36	13.14%	11.22%	4.16%	4,415	0.135	0.543	0.003
Hong Kong	1,262	273	14.96%	12.88%	3.37%	2,869	0.114	0.420	0.010
Hungary	104	64	14.71%	16.47%	10.61%	1,536	0.121	0.368	0.010
India	899	289	14.52%	15.46%	5.82%	2,716	0.133	0.527	0.004
Indonesia	393	27	15.95%	16.32%	8.26%	1,360	0.136	0.560	0.016
Ireland	263	66	12.90%	11.94%	3.30%	12,679	0.091	0.639	-0.002
Israel	70	20	10.52%	11.21%	2.34%	13,333	0.080	0.631	0.005
Italy	967	149	11.12%	9.90%	2.77%	20,660	0.097	0.689	0.003
Japan	5,772	370	8.31%	7.18%	-0.16%	8,541	0.107	0.541	0.004
Korea (South)	897	134	15.94%	15.95%	3.43%	7,082	0.159	0.606	0.008
Luxembourg	30	10	10.53%	17.28%	2.06%	8,890	0.121	0.530	-0.001
Malaysia	1,623	40	11.14%	10.75%	2.72%	1,582	0.108	0.470	0.006
Mexico	212	162	15.19%	14.67%	7.72%	5,247	0.107	0.482	0.009
Netherlands	1,417	286	13.16%	11.33%	2.38%	12,845	0.087	0.618	0.005
New Zealand	511	28	11.49%	10.52%	2.11%	683	0.079	0.459	0.006
Norway	585	114	13.65%	12.63%	2.05%	2,379	0.107	0.583	0.014
Pakistan	46	2	18.64%	16.84%	7.35%	567	0.147	0.693	0.003
Peru	30	13	15.89%	16.72%	6.09%	2,334	0.117	0.607	0.010
Philippines	312	72	14.46%	12.19%	6.58%	1,626	0.136	0.527	0.011
Poland	168	58	14.77%	13.71%	6.85%	2,807	0.126	0.555	0.014
Portugal	243	45	11.80%	11.34%	3.13%	7,209	0.086	0.649	0.001
Russia	10	7	17.79%	18.65%	12.95%	8,240	0.110	0.519	0.000
Singapore	1,207	103	11.50%	10.10%	1.22%	2,749	0.112	0.455	0.006
South Africa	1,231	271	16.99%	15.71%	6.10%	2,467	0.110	0.479	0.006
Spain	904	118	11.59%	11.29%	3.46%	13,448	0.085	0.614	0.003
Sri Lanka	44	5	17.31%	11.88%	6.17%	222	0.120	0.579	0.000
Sweden	1,105	152	12.53%	11.74%	1.18%	5,822	0.100	0.574	0.007
Switzerland	1,187	151	11.47%	9.37%	0.95%	20,533	0.084	0.595	0.004
Taiwan	783	186	12.05%	11.53%	1.11%	2,939	0.131	0.439	0.007
Thailand	638	44	14.27%	13.35%	3.42%	962	0.133	0.539	0.012
Turkey	21	11	14.66%	13.89%	8.06%	4,913	0.118	0.546	-0.017
United Kingdom	6,192	991	11.48%	10.48%	2.75%	10,273	0.092	0.562	0.003
Total (Average)	40,497	6,048	11.67%	11.71%	2.37%	8,760	0.103	0.551	0.005

(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}	
Listing Type	*Firm-Years*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*	*Mean*	*Std. Dev.*
Non-ADR Firms	34,449	11.67%	4.49%	11.46%	5.12%	7.63%	4.20%	13.72%	5.04%	13.77%	5.69%
ADR Firms	6,048	11.71%	4.36%	11.25%	4.74%	8.01%	4.10%	13.72%	4.97%	13.78%	5.71%
Private Placements 144A	1,475	13.27%	4.96%	12.90%	5.29%	9.32%	4.47%	15.29%	5.50%	15.20%	6.38%
OTC Listing	2,471	11.59%	4.27%	11.20%	4.73%	7.87%	3.89%	13.62%	4.99%	13.65%	5.70%
Exchange Listing	2,454	10.89%	3.83%	10.32%	4.11%	7.39%	3.89%	12.88%	4.46%	13.07%	5.26%
All Firms	40,497	11.67%	4.47%	11.43%	5.06%	7.69%	4.19%	13.72%	5.03%	13.78%	5.69%

The sample comprises 40,497 firm-year observations from 45 countries between 1990 and 2005 for which sufficient Worldscope financial data, I/B/E/S forecast and pricing data exist. We require 20 firm-years per country and include only observations from countries with at least one ADR, with market values or total assets larger than US$ 10 millions, 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 mean values for the variables by country (Panel A), and the number of firm-year observations, 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 et al. (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. FBIAS equals the I/B/E/S analyst forecast error (mean forecast for the next fiscal year minus actual earnings) scaled by lagged total assets. Accounting data are measured as of the fiscal-year end, RVAR, FBIAS and the cost of capital as of month +10 after the fiscal-year end. Except for variables with natural lower or upper bounds and SIZE (where we use the natural logarithm), we truncate all firm-level attributes at the first and 99th percentile.

TABLE 2

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

$$COC_{it} = \alpha_0 + \alpha_1 PP_{it} + \alpha_2 OTC_{it} + \alpha_3 EXCH_{it} + \alpha_4 INFL_{it} + \alpha_5 SIZE_{it} + \alpha_6 RVAR_{it} + \alpha_7 LEV_{it}$$
$$+ \alpha_8 FBIAS_{it} + \sum\alpha_j \text{Year Controls}_t + \sum\alpha_k \text{Industry Controls}_i + \sum\alpha_l \text{Country Controls}_i + \varepsilon_{it}$$

Panel A: Cross-Sectional Panel Regressions

Variable	*Pred. Sign*	*Model 1* *(all years)*	*Model 2* *(all years)*	*Model 3* *(all years)*	*Model 4* *(except -1/0)*	*Model 5* *(Anticipation by one year)*
N		40,497	40,497	40,497	39,340	40,497
PP	+/–	-0.094	0.340 #	0.595 #	0.728	0.333
		(-0.49)	(1.85)	(1.66)	(1.59)	(0.84)
OTC	–	-0.874 **	-0.295 *	-0.395 #	-0.577 *	-0.521 *
		(-6.94)	(-2.41)	(-1.95)	(-2.37)	(-2.47)
EXCH	–	-1.292 **	-0.436 **	-0.834 **	-1.317 **	-1.083 **
		(-10.57)	(-3.61)	(-3.32)	(-4.22)	(-3.98)
p-values:	$\alpha_1=\alpha_2$	(0.001) **	(0.005) **	(0.017) *	(0.013) *	(0.062) #
	$\alpha_2=\alpha_3$	(0.013) *	(0.373)	(0.095) #	(0.026) *	(0.048) *
INFL	+	–	20.991 **	23.023 **	22.747 **	22.983 **
			(12.90)	(7.58)	(7.37)	(7.53)
SIZE	–	–	-0.418 **	0.013	0.007	0.021
			(-20.40)	(0.12)	(0.07)	(0.20)
RVAR	+	–	6.867 **	2.839 **	2.951 **	2.831 **
			(13.87)	(3.83)	(3.87)	(3.82)
LEV	+	–	4.114 **	2.469 **	2.483 **	2.459 **
			(26.04)	(5.42)	(5.38)	(5.39)
FBIAS	+	–	21.540 **	17.562 **	17.735 **	17.568 **
			(24.00)	(15.51)	(16.13)	(15.54)
Fixed Effects:						
Year		included	included	included	included	included
Industry & Country		included	included	–	–	–
Firm		–	–	included	included	included
R^2		28.3%	35.3%	70.0%	70.5%	70.0%
F-Stat		109.8	155.7	55.1	54.2	55.0

(continued)

TABLE 2 (continued)

Panel B: Sustainability of U.S. Cross-Listing Effects over Time

Years Relative to Cross-Listing	*≤ -3*	*-2*	*-1*	*0*	*+1*	*+2*	*+3*	*≥ +4*
Specification with Year-, Industry- and Country-Fixed Effects (N = 40,497)								
PP	-0.129	-0.038	-0.453	0.275	0.603 *	0.826 **	0.481	0.079
	(-0.31)	(-0.09)	(-1.35)	(0.88)	(2.21)	(2.57)	(1.61)	(0.39)
OTC	-0.039	-0.005	-0.649 **	-0.319	-0.299	-0.259	-0.126	-0.337 *
	(-0.19)	(-0.02)	(-2.84)	(-1.43)	(-1.50)	(-1.23)	(-0.59)	(-2.39)
EXCH	0.340	-0.271	-0.492 #	-0.381 #	-0.302	-0.552 *	-0.531 *	-0.454 **
	(1.35)	(-0.84)	(-1.81)	(-1.64)	(-1.34)	(-2.54)	(-2.33)	(-3.21)
Specification with Year- and Firm-Fixed Effects (N = 40,497)								
PP	–	-0.226	-0.429	0.556	0.457	0.624	0.427	0.362
		(-0.35)	(-0.59)	(0.79)	(0.68)	(0.95)	(0.57)	(0.52)
OTC	–	-0.043	-0.683 *	-0.548 #	-0.703 *	-0.704 *	-0.473	-0.907 **
		(-0.15)	(-2.35)	(-1.73)	(-2.18)	(-2.23)	(-1.38)	(-2.68)
EXCH	–	-0.729 *	-1.275 **	-1.087 **	-1.076 **	-1.423 **	-1.407 **	-1.870 **
		(-2.38)	(-4.16)	(-3.19)	(-2.85)	(-4.26)	(-3.83)	(-5.27)

Panel C: Impact of Sarbanes-Oxley and London Stock Exchange Listing on U.S. Cross-Listing Effects

Variable (N = 40,497)	*PP*	*PP*SOX*	*OTC*	*OTC*SOX*	*EXCH*	*EXCH*SOX*	*SOX*
Passage of the Sarbanes-Oxley Act	0.846 *	-0.769 **	-0.332	-0.145	-0.767 **	-0.161	2.017 **
	(2.19)	(-2.68)	(-1.59)	(-0.65)	(-2.85)	(-0.70)	(9.26)
							LSE
Listing on London Stock Exchange	0.620 #		-0.393 #		-0.826 **		-0.309
	(1.70)		(-1.93)		(-3.26)		(-0.51)

The sample comprises 40,497 firm-year observations from 45 countries over the period from 1990 to 2005. 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 a description of the remaining variables see notes to Table 1. In Panel A, we report the results from cross-sectional panel regressions. For Models 1 to 3, we use all 40,497 firmyears. In Model 4, we account for the possibility that the market learns about the ADR one or two years prior to the actual listing, and therefore exclude the years -1 and 0 from the analysis. Alternatively, in Model 5, we already set the respective ADR indicator variables equal to one in the year leading up to the cross-listing (i.e., beginning in year -1 instead of year 0). Each model includes an intercept, year-fixed effects, and, depending on the specification, industry-fixed effects based on the industry classification in Campbell (1996), country- or firm-fixed effects, but we do not report the coefficients. The panel reports OLS coefficient estimates and t-statistics (in parentheses) based on robust standard errors that are clustered by firm in Models 1 and 2 and by country-industry in Models 3 to 5. It also reports p-values from Wald tests comparing the magnitude of the ADR coefficients. In Panel B, we replace the unique ADR listing type variables by a series of indicator variables marking the time periods relative to the respective cross-listing (i.e., from years -3 and earlier to +4 and later). In Panel C, we split the timeline into a period before and after the passage of the Sarbanes-Oxley Act and introduce a SOX variable that takes on the value of one for all firmyear observations with COC measured after September 2002. We then interact SOX with the three ADR type variables. Finally, we include an LSE indicator variable in the regressions, which marks non-U.K. firms' listing on the main market of the London Stock Exchange. Panels B and C report only the coefficients (t-statistics) of the variables of interest, but regressions include the full set of controls and fixed effects (see Model 3 in Panel A if not indicated otherwise). 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: Alternative Cost of Capital Estimates as Dependent Variable

Variable	*N*	*PP*	*OTC*	*EXCH*
r_{CT}	40,191	0.574	-0.114	-0.628 *
		(1.31)	(-0.52)	(-2.17)
r_{GLS}	39,979	0.346	-0.317	-0.741 **
		(1.01)	(-1.58)	(-2.78)
r_{OJ}	40,193	0.556	-0.655 *	-1.018 **
		(1.34)	(-2.25)	(-2.86)
r_{PEG}	40,293	0.759 #	-0.398	-0.824 **
		(1.85)	(-1.64)	(-2.83)
First principal component of	40,472	0.181 #	-0.119 #	-0.253 **
individual COC measures		(1.67)	(-1.95)	(-3.33)
Model fit-weighted	39,234	0.484	-0.357 #	-0.778 **
average COC estimates		(1.37)	(-1.80)	(-3.19)
Accuracy-weighted	40,148	0.450	-0.379 #	-0.874 **
COC regression		(1.20)	(-1.93)	(-3.40)
Iteration-weighted	40,497	0.472	-0.358 #	-0.900 **
COC regression		(1.35)	(-1.77)	(-3.49)

Panel B: Additional Controls for Capital Issuance

Variable	*N*	*PP*	*ISS_PP*	*OTC*	*EXCH*	*ISS_EXCH*
Capital	40,497	0.553	0.353	-0.390 #	-0.813 **	-0.348
issuance indicator		(1.52)	(0.96)	(-1.92)	(-3.18)	(-1.15)
					LEVEL II	*LEVEL III*
Exchange	40,497	0.591 #		-0.393 #	-0.692 *	-1.210 **
level indicator		(1.67)		(-1.94)	(-2.42)	(-3.88)

Panel C: Alternative Model Specifications and Sample Restrictions

Variable	*N*	*PP*	*OTC*	*EXCH*
Use local risk-free rates	40,497	0.629 #	-0.520 *	-0.796 **
instead of inflation		(1.77)	(-2.31)	(-3.18)
Risk premiums instead	40,038	1.025 *	-0.274	-0.634 *
of COC estimates		(2.48)	(-1.02)	(-2.33)
Replace missing/negative earnings	63,166	0.428	-0.298	-0.833 **
forecasts in COC computation		(0.97)	(-1.46)	(-2.79)

Panel D: Influence of (Long-Run) Growth on Cost of Capital Estimates

Variable	*N*	*PP*	*OTC*	*EXCH*
COC with ADR-type/year-specific	40,413	0.530	-0.552 **	-0.976 **
correction to terminal-value growth (forecasted)		(1.43)	(-2.66)	(-3.85)
COC with ADR-type/year-specific	39,661	0.301	-0.779 **	-1.227 **
correction to terminal-value growth (realized)		(0.73)	(-3.75)	(-4.83)
COC with firm-specific	31,707	0.578	-0.557 *	-0.974 **
terminal-value growth		(1.20)	(-2.29)	(-3.17)
COC with firm/year-specific	24,759	0.595	-0.546 *	-0.649 *
terminal-value growth		(1.18)	(-2.49)	(-2.23)
Control for firm/year-specific	37,572	0.569	-0.413 *	-0.809 **
forecasted growth in regression		(1.50)	(-2.04)	(-3.27)

(continued)

The sample comprises a maximum of 40,497 firm-year observations from 45 countries over the period from 1990 to 2005. If not indicated otherwise, we use COC, the mean of four estimates for the implied cost of equity capital (see Appendix), as the dependent variable. The four 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, year- and firm-fixed effects (see Model 3 in Panel A of Table 2). In Panel A, we report results for the following alternative dependent variables: the four individual cost of capital estimates, the first principal component of the four individual estimates, and the model fit-weighted average of the four estimates (where the R^2 from country regressions of the individual cost of capital estimates on the firm attributes and year-fixed effects serve as weights). In addition, we also report the coefficients from weighted OLS regressions with one plus the log of (1) the inverse of the absolute one-year-ahead analyst forecast error scaled by lagged total assets (accuracy-weighted), or (2) the inverse of the maximum number of iterations used to impute the individual cost of capital estimates (iteration-weighted) as the weights. 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 the contemporaneous local risk-free or money-market rate, r_f, measured as country/year median of monthly values (source: Datastream and World Bank), (2) we use the risk premium (COC minus r_f) instead of the raw cost of capital estimates as the dependent variable, and (3) 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 assess the sensitivity of the cost of capital estimates to our long-run growth assumptions. We compute the cost of capital estimates by (1) adjusting ADR firms' perpetual growth rate in all years by the yearly median difference in analyst long-run growth estimates between the respective ADR type and non-ADR firms, (2) same as (1) but using realized, three-year average growth rates in earnings per share, (3) using (firm-specific) mean realized annual growth in earnings per share over the entire period (the pre cross-listing period for ADR firms) as perpetual growth rate, and (4) setting perpetual growth equal to (firm- and year-specific) analysts' consensus long-run growth estimates. For specifications (1) to (4), we also vary the industry-specific target accounting return on equity across ADR types when computing r_{GLS} (see Appendix for details). Finally, we include the (firm- and year-specific) expected growth in earnings per share based on actual values and the one-year-ahead mean analyst consensus forecasts as additional control variable in the regression, and allow this coefficient to vary across the three ADR types, i.e., we include interactions with all ADR indicators. The table reports OLS coefficient estimates and t-statistics (in parentheses) based on robust standard errors that are clustered by country-industry. 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 of the Cost of Capital Effects around U.S. Cross-Listings

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

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

		Years -2/+1 only			*all Years (except Years -1/0)*		
Variable	*Time*	*N*	*Median*	*Mean*	*N*	*Median*	*Mean*
PP	pre	82	0.10%	0.35%	280	-0.02%	0.01%
	post		0.28%	0.92%		-0.08%	0.49%
	post - pre		0.17%	0.57%		-0.06%	0.48%
OTC	pre	220	-0.58%	-0.32%	1,235	-0.61%	-0.38%
	post		-0.46%	-0.24%		-1.15%	-0.88%
	post - pre		0.12%	0.08%		-0.55% **	-0.50% **
EXCH	pre	170	-0.87%	-0.59%	1,052	-0.42%	-0.07%
	post		-1.15%	-1.09%		-1.47%	-1.20%
	post - pre		-0.28%	-0.49%		-1.05% **	-1.13% **

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

	Pred.	*Years -2/+1 only*		*all Years (except Years -1/0)*	
Variable	*Sign*	*Model 1*	*Model 2*	*Model 1*	*Model 2*
N		510	510	2,672	2,672
PP	+/–	0.577	0.648	0.531	0.451
		(1.24)	(1.49)	(1.59)	(1.45)
OTC	–	-0.530 #	-0.416	-0.590 **	-0.510 **
		(-1.70)	(-1.44)	(-2.94)	(-2.62)
EXCH	–	-0.967 **	-0.695 #	-0.749 **	-0.545 *
		(-2.61)	(-1.74)	(-3.31)	(-2.32)
SIZE	–	–	-0.400 **	–	-0.253 **
			(-3.67)		(-3.58)
RVAR	+	–	8.975 *	–	4.722 *
			(2.00)		(2.33)
LEV	+	–	4.295 **	–	3.127 **
			(3.93)		(5.34)
FBIAS	+	–	14.913 *	–	14.726 **
			(2.44)		(4.40)
Industry-Fixed Effects		included	included	included	included
R^2		5.0%	13.1%	6.0%	10.5%
F-Stat		2.0	3.4	5.2	7.4

The sample for the difference-in-differences analysis of cost of capital changes comprises only observations from newly listed ADR firms over the period from 1990 to 2005. 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 years -1 or 0. For a description of the remaining variables see notes to Table 1. We include an intercept and industry-fixed effects in the regressions, but do not report the coefficients. Panel A reports median (mean) values before and after firms' first ADR and any subsequent upgrading to an exchange-listing. We evaluate differences 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

$$COC_{it} = \alpha_0 + \alpha_1 PP_{it} + \alpha_2 OTC_{it} + \alpha_3 EXCH_{it} + \alpha_4 INFL_{it} + \alpha_5 SIZE_{it} + \alpha_6 RVAR_{it} + \alpha_7 LEV_{it} + \alpha_8 FBIAS_{it} + \sum\alpha_j \text{Year Controls}_t + \sum\alpha_k \text{Industry Controls}_i + \sum\alpha_l \text{Country Controls}_i + \varepsilon_{it}$$

	Predicted	*Legal Tradition*		*Disclosure Regulation*		*Regulation of Self-Dealing*		*Private Benefits of Control*		*Equity Market Development*	
Variable	*Sign*	*code law*	*common law*	*low*	*high*	*low*	*high*	*high*	*low*	*low*	*high*
N		23,912	16,585	13,196	26,507	15,834	24,663	23,815	13,978	14,967	24,736
PP	+/–	0.091	0.860 *[#]	-0.100	0.550 *[#]	0.002	0.643 *[#]	0.105	-0.183	0.148	0.466
		(0.43)	(2.40)	(-0.36)	(2.19)	(0.01)	(2.34)	(0.45)	(-0.64)	(0.66)	(1.39)
OTC	–	-0.141	-0.413 **	-0.193	-0.353 *	-0.224	-0.333 *	-0.307 #	-0.179	-0.345 #	-0.280 #
		(-0.77)	(-2.59)	(-0.86)	(-2.41)	(-1.02)	(-2.32)	(-1.91)	(-0.95)	(-1.66)	(-1.88)
EXCH	–	-0.428 **	-0.309	-0.764 **	-0.260 #[*]	-0.680 **	-0.190 [*]	-0.498 **	-0.215	-0.607 **	-0.392 **
		(-2.89)	(-1.51)	(-4.05)	(-1.65)	(-3.68)	(-1.19)	(-3.01)	(-1.16)	(-3.01)	(-2.64)
INFL	+	21.667 **	18.062 **	17.350 **	20.273 **	23.285 **	18.011 **	23.795 **	11.516 **	17.768 **	24.962 **
		(8.77)	(7.94)	(5.15)	(9.37)	(8.04)	(8.79)	(11.25)	(2.89)	(7.26)	(9.88)
SIZE	–	-0.428 **	-0.428 **	-0.345 **	-0.455 **	-0.355 **	-0.465 **	-0.415 **	-0.488 **	-0.321 **	-0.481 **
		(-16.43)	(-12.96)	(-9.53)	(-18.05)	(-10.27)	(-18.33)	(-14.75)	(-16.44)	(-8.82)	(-19.80)
RVAR	+	4.818 **	9.976 **	6.168 **	7.422 **	5.039 **	8.402 **	5.654 **	8.348 **	5.215 **	8.029 **
		(8.03)	(11.78)	(6.67)	(12.29)	(6.35)	(13.20)	(8.88)	(9.85)	(6.52)	(12.40)
LEV	+	4.663 **	3.295 **	4.664 **	3.845 **	4.597 **	3.800 **	4.189 **	4.083 **	3.723 **	4.323 **
		(23.48)	(13.12)	(16.29)	(20.05)	(16.05)	(20.28)	(19.38)	(16.91)	(13.26)	(22.64)
FBIAS	+	21.381 **	21.293 **	20.780 **	21.633 **	20.835 **	21.834 **	21.359 **	22.167 **	21.082 **	21.495 **
		(19.16)	(14.37)	(14.85)	(18.40)	(16.03)	(17.70)	(20.13)	(11.70)	(15.74)	(17.67)
Year-, Industry- & Country-Fixed Effects		included	included	included	included	included	included	included	included	included	included
R^2		38.4%	31.1%	30.4%	37.9%	31.1%	38.2%	29.3%	44.2%	33.7%	32.8%
F-Stat		138.4	74.1	62.1	178.6	70.8	160.3	83.2	144.2	78.7	131.7

The sample comprises a maximum of 40,497 firm-year observations from 45 countries over the period from 1990 to 2005. 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 for the disclosure requirements in securities offerings (La Porta et al., 2006), (3) regulation of self-dealing is equal to high for countries with above-median values of the anti-self-dealing index representing the legal protection of minority shareholders against insider expropriation (Djankov et al., 2008), (4) private benefits of control are assumed to be low for countries with below-median average block premia paid in control transactions (Dyck and Zingales, 2004), and (5) equity market development is high for countries with above-median average percentage rank across the three La Porta et al. (1997) variables (i) ratio of the aggregate stock market capitalization held by minorities to gross national product, (ii) number of listed domestic firms relative to the population, and (iii) number of IPOs relative to the population. We compute all median values to partition the sample based on the firmyears available. The dependent variable, COC, is the mean of four estimates for the implied cost of equity capital (see Appendix). 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. For a description of the remaining variables see notes to Table 1. We include an intercept, year-, industry-, and country-fixed effects in the regressions, but do not report the coefficients. 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 Three-Year Price Changes around U.S. Cross-Listings

$$\Delta P_{CFit} \text{ or } \Delta P_{COCit} = \alpha_0 + \alpha_1 PP_{-4/-3it} + \alpha_2 OTC_{-4/-3it} + \alpha_3 EXCH_{-4/-3it} + \alpha_4 SIZE_{it}$$
$$+ \alpha_5 RVAR_{it} + \alpha_6 LEV_{it} + \alpha_7 FBIAS_{it} + \sum \alpha_k \text{Industry Controls}_i + \varepsilon_{it}$$

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

	Price	*Years -4/-3*			
Variable	*Effect*	*N*	*Median*	*Mean*	*Std. Dev.*
PP	ΔP	43	8.54%	11.42%	44.21%
	ΔP_{CF}		9.20%	14.23% #	47.37%
	ΔP_{COC}		0.87%	3.12%	43.75%
OTC	ΔP	164	12.76%	12.94%	39.68%
	ΔP_{CF}		9.59% **	9.36% **	43.00%
	ΔP_{COC}		5.12% #	4.88%	39.85%
EXCH	ΔP	155	12.84%	15.73%	40.99%
	ΔP_{CF}		8.19% **	9.19% **	37.75%
	ΔP_{COC}		9.39% **	12.90% **	40.29%

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*
$PP_{-4/-3}$		11.426 *	11.265 *	0.208	0.923
		(2.02)	(2.00)	(0.04)	(0.17)
$OTC_{-4/-3}$		7.872 *	7.054 *	1.026	0.682
		(2.51)	(2.25)	(0.33)	(0.22)
$EXCH_{-4/-3}$		8.958 **	8.878 **	11.491 **	11.561 **
		(2.58)	(2.57)	(3.36)	(3.38)
p-values:	$\alpha_1=\alpha_2$	(0.580)	(0.719)	(0.086) #	(0.105)
	$\alpha_2=\alpha_3$	(0.814)	(0.692)	(0.022) *	(0.017) *
SIZE		–	-0.330	–	-1.092 *
			(-0.69)		(-2.33)
RVAR		–	-24.838 #	–	-48.280 **
			(-1.83)		(-3.60)
LEV		–	-3.829	–	5.691
			(-0.89)		(1.34)
FBIAS		–	-202.101 **	–	42.686
			(-6.43)		(1.37)
Industry-Fixed Effects		included	included	included	included
R^2		1.2%	2.3%	0.8%	1.2%
χ^2		53.0	99.9	35.5	53.8

The sample comprises only ADR observations from 1990 to 2002 because we decompose the three-year-ahead price change into a cash flow- and cost of capital-component. There is a maximum of 4,342 firm-years with both, cost of capital and cash flow effects available. We calculate three different price change variables: (1) ΔP is the three-year-ahead stock return using actual stock prices, (2) ΔP_{CF} is the implied three-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+3, and (3) ΔP_{COC} is the implied three-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+3. All price change variables are computed as the difference between the natural log of price in t+3 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 years -1 or 0 and hence compute three-year-ahead price changes ending in these years, i.e., starting in the years -4 and -3 as indicated by the subscripts. For a description of the remaining variables see notes to Table 1. 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. It also reports p-values from Wald tests comparing the magnitude of the coefficients across ADR types. For expositional purposes we multiply all coefficients by 100. **, *, and # indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.