Journal of Accounting Research
Vol. 41 No. 2 May 2003
Printed in U.S.A.

ADRs, Analysts, and Accuracy: Does Cross Listing in the United States Improve a Firm's Information Environment and Increase Market Value?

MARK H. LANG,* KARL V. LINS,† AND DARIUS P. MILLER‡

Received 29 December 2001; accepted 13 November 2002

ABSTRACT

This paper investigates the relation between cross listing in the United States and the information environment of non-U.S. firms. We find that firms that cross list on U.S. exchanges have greater analyst coverage and increased forecast accuracy than firms that are not cross listed. A time-series analysis shows that a change in analyst coverage and forecast accuracy occurs around cross listing. We also document that firms that have more analyst coverage and higher forecast accuracy have higher valuations. Furthermore, the change in firm value around cross listing is correlated with changes in analyst following and forecast accuracy, suggesting that cross listing enhances firm value through its effect on the firm's information environment. Our findings support the hypothesis that cross-listed firms have better information environments, which are associated with higher market valuations.

*The University of North Carolina; †The University of Utah; ‡Indiana University. We thank Ray Ball (the editor) and Christian Leuz (the discussant) for detailed suggestions on improving the paper. We also thank Walt Blacconiere, Mike Lemmon, Jim Linck, Laureen Maines, Marlene Plumlee, Lynn Rees, Henri Servaes, Charles Trzcinka, and seminar participants at the 2002 *Journal of Accounting Research* conference, Indiana University, the University of Melbourne, the Eighth Annual Georgia Tech/Fortis International Finance Conference, and the Darden Valuation in Emerging Markets Conference for helpful comments. Alan Montgomery provided valuable research assistance. We are grateful for the analyst forecast information provided by I/B/E/S.

1. Introduction

In this paper, we examine cross listing in the United States and its relation to the information environment of non-U.S. firms.[1] A large literature on international cross listing suggests that information considerations are a key factor in cross listing decisions. However, there is little direct empirical evidence on the relation between cross listing and the information environment of the firm. We document several empirical findings on this subject. First, we show cross-sectionally that non-U.S. firms that are listed on U.S. exchanges have greater analyst coverage and increased forecast accuracy relative to other non-U.S. firms. Second, we perform time-series analyses and find that a change in analyst coverage and forecast accuracy occurs around cross listing. Third, we document that analyst coverage and forecast accuracy are both positively related to firm value. Finally, we show that the change in value around cross listing is correlated with changes in analyst following and forecast accuracy, suggesting that cross listing increases firm value through its effect on the firm's information environment. Our results are robust to adjustments for the potential endogeneity of the listing decision and simultaneity between analyst following and forecast accuracy. Overall, our findings support the hypothesis that important informational effects occur with cross listing and that these effects are positively associated with firm value.

Foerster and Karolyi [1999] and Miller [1999] document positive average abnormal announcement returns for non-U.S. firms that issue exchange-listed American Depositary Receipts (ADRs), whereas Foerster and Karolyi [2000] document positive long-horizon returns for such firms that raise capital. Similarly, Errunza and Miller [2000] find a substantial decline in a firm's cost of capital after an ADR. These studies and others offer a number of explanations for why cross listing on a U.S. stock exchange adds value.[2] However, a crucial component in almost all of these explanations is the firm's information environment. The notion that the information environment should be a function of cross listing is natural because, as discussed in Coffee [2002], cross-listing firms subject themselves to (1) increased enforcement by the Securities and Exchange Commission (SEC), (2) a more demanding litigation environment, and (3) enhanced disclosure and reconciliation to U.S. generally accepted accounting principles (GAAP). In addition, cross-listing firms may face more scrutiny from investors, more pressure to provide guidance than they did in their home markets, and increased scrutiny from their auditors. Firms that list in U.S. markets are, in effect, "bonding" themselves to an increased level of disclosure and scrutiny.

---

[1] Our notion of information environment is similar to the concept of "corporate transparency" in Bushman, Piotroski, and Smith [2001], broadly construed to include the effects of "corporate reporting, private information acquisition and information dissemination."

[2] For references on earlier studies such as Errunza and Losq [1985] and Alexander, Eun, and Janakiramanan [1987, 1988], see the survey by Karolyi [1998].

These changes in transparency could affect firm value by decreasing the cost of capital, increasing the cash flows that ultimately accrue to shareholders, or both.

For the cost of capital, Merton's [1987] investor recognition hypothesis is often used to argue that a U.S. listing creates value because the enhanced disclosure environment reduces the cost of following the firm. This increases the investor base and, therefore, the demand for the firm's securities. Barry and Brown [1985] suggest that cost of capital is a function of "estimation risk" and the better investors are able to assess the prospects for a company, the lower is its expected cost of capital. Along this line of reasoning, Lang and Lundholm [1996] show that analysts' forecasts are more accurate for firms that disclose more, whereas Gebhardt, Lee, and Swaminathan [2001] find that firms with more accurate forecasts enjoy a lower implied cost of capital. Thus, if investors are able to assess more accurately the prospects for a cross-listed firm, its cost of capital should be reduced. Taken together, the investor recognition and estimation risk hypotheses suggest that a firm's information environment can play an important role in determining its cost of capital.

The literature also suggests that the enhanced transparency associated with cross listing may influence value through pure cash-flow effects by reducing agency costs. For example, cross listing may be associated with improved firm-level corporate governance because it bonds a firm to greater transparency, which should reduce the potential diversion of a firm's cash flows to managers and controlling shareholders (Coffee [1999], Stulz [1999]). Consistent with this hypothesis, recent research by Doidge, Karolyi, and Stulz [Forthcoming] documents that a U.S. listing creates the most value for firms with higher growth opportunities located in countries that have poor disclosure and investor protections. Lins, Strickland, and Zenner [2002], Reese and Weisbach [Forthcoming], and Pagano, Roell, and Zechner [Forthcoming] argue that cross listing adds value because the greater transparency increases the willingness of both international and local investors to commit capital.

Diamond and Verrecchia [1991] and Leuz and Verrecchia [2000] emphasize the importance of precommitment in the relation between disclosure and cost of capital. Cross listing provides a credible commitment to increased disclosure because a firm is subject to greater regulatory and investor scrutiny, disclosure requirements, and potential legal exposure. This differentiates cross listing from simply announcing an intention to increase disclosure in the home market because a cross-listed firm cannot easily renege on its commitment if it later turns out to have bad news that it would prefer not to disclose. It is relatively costly from a reputational standpoint to delist, because the firm would risk alienating its international investor base. A similar argument applies to the effect of cross listing on agency problems and cash flows; investors are likely to anticipate better incentive alignment for firms that bond themselves to better investor protection by listing in the United States.

Despite its theoretical importance, surprisingly little direct evidence on the relation between a firm's information environment and cross-listing exists.[3] One factor that makes testing this relation difficult is that it is not possible to measure directly a firm's information environment. For instance, Bailey, Karolyi, and Salva [2002] use price volatility and volume reaction to earnings announcements to conduct tests on the information environment of cross-listed firms. Baker, Nofsinger, and Weaver [2002] test whether "visibility" increases around cross listing using measures of analyst following and media coverage. Our approach is to follow prior research by Lang and Lundholm [1996], Healy, Hutton, and Palepu [1999], and Gebhardt, Lee, and Swaminathan [2001] and use the characteristics of analyst forecasts as a proxy for the information environment. In particular, we focus on two measures: the number of analysts following the firm and the accuracy of analyst forecasts. Previous studies suggest that having more analysts with more accurate forecasts indicates a firm with a better information environment.[4]

Analyst following should be related to cross listing and value for several reasons. First, to the extent that cross listing increases the quantity of information available to the market, either because of explicit disclosure requirements or implicit pressure to provide additional information to analysts and investors, it should reduce the cost of following a firm, which could lead to increased coverage by investment analysts. Second, cross listing widens the potential investment base of a firm. This should be associated with increased analyst activity because analysts are likely to focus on firms that investors find interesting and investors are more likely to consider firms followed by analysts. Similarly, analysts may be attracted to cross-listing firms because they perceive them to be of higher quality and therefore of more potential

[3] There is a substantial body of research examining the nature of reconciling items for cross-listed firms as reported on Form 20-F (summarized in Pownall and Schipper [1999]). Much of that literature examines the association between reconciling items and share prices to infer whether the information in reconciliations is "value relevant" and generally finds a significant association. However, evidence on whether the Form 20-F is the source of the information is mixed because the information-release date is generally not clear. Our interest is different in that we are not concerned about whether specific reconciling items are associated with share price but, rather, about whether cross-listed firms are characterized by generally richer information environments.

[4] Following the prior literature, we view the analyst variables as indicative of, but not necessarily the cause of, changes in a firm's information environment. For example, analyst forecast accuracy is intended as a measure of how well the market understands the firm's economics. This may partially be a result of analyst activity, but it may also reflect disclosure by the firm or information gathering by other investors. Similarly, analyst following is intended to proxy for private information-acquisition activities. Although the analysts' research may indeed enhance the information environment, the same incentives that attract sell-side analysts might also attract buy-side analysts and other investors. Furthermore, our two measures are not intended as mutually exclusive. For example, forecasts for a given firm may be more accurate because there are more analysts following the firm. Similarly, we do not believe the two measures are entirely redundant; forecasts may be accurate because a firm discloses more without more analysts following. Our empirical tests take these two measures into account both separately and simultaneously.

interest to their potential investor base. These predictions are also consistent with the investor relations literature, which suggests that a benefit of increased disclosure is increased analyst following.[5] Finally, additional analyst following should also bring about more scrutiny, which, in the presence of agency costs, should improve firm value by increasing the cash flows that accrue to shareholders (Lang, Lins, and Miller [2002]). In summary, increased analyst following around cross listing can influence aspects of the firm's information environment that are argued to affect both cash flows and the cost of capital.

The previous arguments suggest that a firm's disclosures and the information produced by analysts complement each other. Consistent with this hypothesis, Lang and Lundholm [1996] and Healy, Hutton, and Palepu [1999] find that, for U.S. firms, increased disclosure is associated with higher analyst following. However, it is important to note that although empirical support exists for a positive relation between analyst activity and increased disclosure, the direction of the association is not obvious. For instance, Lang and Lundholm note that to the extent that extra disclosure levels the playing field among analysts, it could reduce any one analyst's competitive advantage, which would lessen incentives to cover the firm. Botosan [1997] finds that when firms already have a high analyst following, increased disclosure is not associated with a reduction in the cost of capital, suggesting that analysts and disclosure function as substitutes.[6] These latter two studies indicate that additional public disclosure could drive out private information acquisition, resulting in an ambiguous effect on total information in the market.

In an international context, Bushman, Piotroski, and Smith [2001, 2002] provide evidence of a positive correlation between analyst following and disclosure and investor protection, suggesting that analysts might be attracted to firms that cross list on U.S. markets. On the other hand, Chang, Khanna, and Palepu [2000] find that companies in code law countries tend to have greater analyst following than companies in common law countries. This finding suggests that by subjecting itself to U.S. regulatory requirements, a cross-listed firm might discourage analyst following. As a result, the direction of the predicted relation between cross listing and analyst following is an empirical issue.

Similarly, the link between analyst following and value is not necessarily positive. For example, if analysts primarily gather private information, their activities could actually increase cost of capital by raising transactions costs and discouraging uninformed investors from purchasing shares (e.g., see Diamond and Verrecchia [1991]). Although such an effect on valuation might be offset by an increase in investor interest, reduced uncertainty,

---

[5] We focus our discussion on sell-side analysts because the IBES data are primarily based on their forecasts.

[6] However, using a longer period and a larger sample, Botosan and Plumlee [2002] document a negative association between annual report disclosure and cost of capital for highly followed firms.

and reduced agency conflicts within cross-listed firms, the expected relation between analyst following and valuation on net is not clear *ex ante*.

We also examine the *ex post* accuracy of analyst forecasts. To the extent that cross listing directly or indirectly increases the amount of information available about the firm, one can argue that it will improve the accuracy of analyst forecasts. Improved accuracy should reduce the cost of capital through its effect on estimation risk. As noted earlier, Gebhardt, Lee, and Swaminathan [2001] find that firms with lower forecast errors have lower implied costs of capital. Improved accuracy may also lessen agency problems to the extent that managers are held more accountable for the details of their firm's cash flows.

However, as with analyst following, the direction of the association between cross listing and forecast accuracy is not obvious *ex ante*. Although cross-listed firms are not required to change local GAAP reporting when they list on U.S. markets, Lang, Raedy, and Yetman [2003] suggest that firms change in ways that make earnings more volatile and more similar to U.S. firms.[7] Similarly, non-U.S. analysts might find it harder to predict earnings following a movement toward U.S. GAAP because of greater familiarity with the local GAAP. To the extent that earnings become less predictable around cross listing, it should bias against our finding results. We explicitly include controls for earnings surprise in our analysis, which should mitigate the effects of increased volatility.

In summary, although much of the literature suggests that cross-listed firms should have a richer information environment in terms of greater analyst following and forecast accuracy, there are also reasons to believe that the relation between cross listing and analyst following and forecast accuracy might be negative, leaving the issue an empirical question. This paper adds to the literature by examining analyst activity around cross listing to see if, in fact, cross listing is associated with greater analyst following and forecast accuracy. We then investigate whether these changes in the information environment are linked to firm value.

The rest of the paper is organized as follows: Section 2 describes the data. Section 3 examines the impact of cross-listing on analyst following and forecast accuracy. Section 4 details the impact of analyst following and performance on firm value. Section 5 offers concluding remarks.

2. *Data*

To examine the firm's information environment, we focus much of our analysis on the levels of our information variables and consider our changes analysis as supplementary. We take this approach for several reasons. First,

[7] Ball [2001] discusses the evidence on cross-country differences in accounting and concludes that environments like the United States focus on the timely recognition of losses (rather than smoothing them over time), which is consistent with an increase in earnings volatility after cross listing. Similarly, Bailey, Karolyi, and Salva [2002] find that stock return and volume reactions to earnings announcements typically increase once a stock cross lists in the United States, suggesting that earnings becomes less predictable and more informative.

the predictions from the prior literature for the level of the information environment following cross listing are clearest. In particular, depending on the view of cross listing, it is possible to envision situations in which the information environment is important but not necessarily reflected in changes around cross listing. For example, Coffee [1999] and Leuz and Verrecchia [2000] argue that the important aspect of cross listing from a valuation perspective is the commitment to increased disclosure rather than the increase in disclosure itself. Cantale [1998], Fuerst [1998], Moel [1999], and Huddart, Hughes, and Brunnermeier [1999] argue that a firm that has a history of transparency will still have an incentive to list because it signals its commitment to continuing that policy, even when it faces circumstances under which it might wish not to disclose. Following this reasoning, in Doidge, Karolyi, and Stulz [Forthcoming] the decision to cross list indicates a firm with well-aligned incentives that is therefore willing to submit itself to scrutiny. It is important to note, however, that the firm may have also had a rich information environment before the cross listing.

Closely related, even if the information environment explicitly changes because of the cross listing, the timing may not be clear. For example, it is unlikely that a firm anticipating a U.S. listing would increase disclosure suddenly following the listing or that analysts would suddenly increase their activity. Rather, as argued by Bradshaw and Miller [2002] and Lang, Raedy, and Yetman [2003], it is likely that firms increase disclosure gradually in anticipation of the listing by moving their accounting closer to U.S. GAAP, increasing their footnote disclosure, and communicating more freely with analysts and investors. As a result, it is difficult to know the window over which to compute the change. We consider long windows to increase the probability that we capture the entire effect, recognizing that longer windows increase the probability of confounding effects. Finally, evaluating changes substantially reduces both our sample size and our ability to include historical control variables because our data sources begin large-scale coverage only in the early to mid 1990s, and most cross-listed firms became listed before the mid 1990s.

To examine the extent and accuracy of analyst activity, we use data from the Historical I/B/E/S International database. Our main results are for 1996. We choose this year because the year-to-year increase in I/B/E/S coverage of firms begins to slow substantially after 1996 and we want to capture any benefits of ADR listings as far back as possible. We use data from the 11th month of the fiscal year to calculate the number of analysts following a company and the forecast accuracy, as O'Brien and Bhushan [1990] document that analyst activity levels off after the 11th month.[8] Forecast accuracy

[8] Given that it is not clear when the analyst variables should be computed in this context, we replicated the results using earnings forecasts in the seventh month of the fiscal year. Because firms are required to file Form 20-F within six months of their fiscal year-end, this should ensure that the previous year's filing is available to the market. Although we lose some observations because analyst following tends to increase during the year, all results are robust to use of the alternate month.

is defined as the negative of the absolute value of the analyst forecast error, deflated by stock price:[9]

$$\textit{Forecast Accuracy}\ (t) = -\left|\frac{\textit{Actual Earnings}\ (t) - \textit{Estimated Earnings}\ (t)}{\textit{Stock Price}\ (t)}\right|$$

To identify firms listing on U.S. markets, we gather information on ADR listings on the New York Stock Exchange (NYSE), American Stock Exchange (AMEX), and Nasdaq. This information is obtained from the Bank of New York, the NYSE, Nasdaq, and the Center for Research in Security Prices (CRSP) databases. Exchange-listed foreign firms are required to register their offering under the 1933 Securities Act and their reports under the 1934 Exchange Act, and reconcile owners' equity and net income to U.S. GAAP. However, as discussed in Coffee [1999], although cross-listing firms must file with the SEC, requirements for them are limited relative to U.S. firms.[10] We treat direct listings by Canadian and Israeli firms that cross list in the United States as ADRs because these firms must also comply with SEC requirements. We verify that all of our results hold if we remove these two countries from the sample.

Because we are also interested in how the characteristics of analyst forecasts translate into value, we obtain valuation data from Worldscope. We use Tobin's Q as a measure of firm value in regressions that feature analyst activity and accuracy as well as control variables.[11] Tobin's Q is computed as total assets less the book value of equity plus the market value of equity in the numerator and book value of assets in the denominator.

Our potential sample of firms from countries covered by I/B/E/S in 1996 that also have one or more exchange-listed ADR firms contains 8,937 observations. Matching firms from I/B/E/S with Worldscope reduces our sample substantially. Overall, the sample with I/B/E/S analyst forecast data and accounting data for fiscal year 1996 includes 4,859 firms from 28 countries.

Table 1 provides summary statistics for the sample based on a firm's country of domicile. Of the 4,859 sample firms, 235 have exchange-listed ADRs as

---

[9] To ensure the changes in price around cross-listing do not drive our forecast accuracy results, we performed the analysis again with forecast accuracy deflated by actual earnings (rather than price). Results are robust to this alternative measure.

[10] In particular, requirements for disclosure on items such as compensation, interested director transactions, and quarterly reporting are relaxed. Furthermore, the firm has six months rather than 90 days to file its Form 20-F. In addition, the firm is exempted from filing a proxy statement, complying with Section 16 short-swing trading rules and Regulation FD selective disclosure limitations. Finally, exchanges may waive listing requirements that are not part of normal practice in their home environment, including the use of audit committees and equal voting rights.

[11] Tobin's Q is widely used as a measure of firm value in the academic literature. Research areas in which Q is used to measure firm value include cross listing (Doidge, Karolyi, and Stulz [2002]), corporate diversification (Lang and Stulz [1994]), takeovers (Servaes [1991]), equity ownership (La Porta et al. [2002] and Lins [Forthcoming]), and hedging (Allayannis and Weston [2001]).

TABLE 1

Sample Selection and Descriptive Statistics

	Full Sample						Sample with Three Years of Historical Data					
Country	*N*	ADRs	Analyst	Accuracy	Size	Return	*N*	ADRs	Analyst	Accuracy	Size	Return
Argentina	28	6	13.5	−0.0103	787	0.239	25	5	14.0	−0.0104	819	0.239
Australia	152	15	8.0	−0.0046	481	0.146	132	13	9.0	−0.0047	516	0.141
Brazil	95	1	11.0	−0.0490	942	0.192	91	1	11.0	−0.0449	994	0.182
Canada	215	57	9.0	−0.0065	498	0.240	201	54	10.0	−0.0065	523	0.234
Chile	54	12	4.0	−0.0082	357	−0.086	52	11	4.0	−0.0095	357	−0.086
China	75	2	6.0	−0.0306	218	0.319	58	1	6.0	−0.0290	202	0.341
Columbia	13	2	3.0	−0.0283	871	0.293	11	2	3.0	−0.0283	1,375	0.335
Denmark	60	1	5.0	−0.0124	232	0.218	58	1	5.0	−0.0124	221	0.218
Finland	36	2	8.5	−0.0119	812	0.476	32	1	7.5	−0.0094	784	0.476
France	191	7	7.0	−0.0096	612	0.125	170	7	7.5	−0.0099	697	0.125
Germany	150	2	7.5	−0.0100	691	0.235	140	1	6.0	−0.0115	681	0.216
Hong Kong	191	3	8.0	−0.0099	329	0.190	150	3	12.5	−0.1007	515	0.172
Ireland	36	6	5.0	−0.0046	510	0.268	35	6	5.0	−0.0048	532	0.265
Israel	4	2	4.0	−0.0047	1,861	0.290	–	–	–	–	–	–
Italy	69	1	10.0	−0.0182	2,013	0.091	67	1	9.0	−0.0184	2,013	0.086
Japan	1,984	27	3.0	−0.0031	625	0.382	1,880	27	3.0	−0.0032	650	0.393
Korea	210	3	5.0	−0.0348	1,075	−0.216	208	3	5.0	−0.0348	1,057	−0.214
Mexico	46	15	15.5	−0.0166	105	0.382	44	15	15.5	−0.0166	111	0.382
Netherlands	84	2	15.5	−0.0041	287	0.235	81	2	16.0	−0.0042	292	0.265
New Zealand	39	2	8.0	−0.0076	330	0.021	31	0	8.0	−0.0075	250	0.016
Norway	41	3	7.0	−0.0134	598	0.360	29	2	12.0	−0.0134	839	0.381
Peru	10	2	5.5	−0.0224	140	0.045	10	2	5.5	−0.0224	140	0.045
Philippines	59	1	10.0	−0.0090	272	−0.111	33	1	11.0	−0.0054	307	0.011
Portugal	23	1	5.0	−0.0102	379	0.195	19	1	6.0	−0.0066	269	0.195
South Africa	117	13	5.0	−0.0039	423	−0.006	105	13	5.0	−0.0040	469	−0.011
Spain	35	2	14.0	−0.0132	592	0.375	34	2	15.0	−0.0139	600	0.343
Sweden	81	4	6.0	−0.0118	468	0.349	73	3	6.0	−0.0118	472	0.306
U. Kingdom	761	41	4.0	−0.0037	198	0.082	683	37	5.0	−0.0038	221	0.081
Total	4,859	235	4.0	−0.0052	506	0.242	4,452	215	4.0	−0.0052	540	0.244

ADRs refer to the number firms with cross-listings that trade on U.S exchanges. Analyst is defined as the median number of I/B/E/S analysts that report estimates for each firm. Forecast accuracy is for fiscal year 1996 and is defined as the negative of the absolute value of the deviation of actual earnings per share from the median analyst forecast of earnings per share, deflated by stock price. Forecast accuracy is winsorized at the 5th percentile. Size is the median value of total assets in millions of U.S. dollars. Return is the median stock return over the previous year.

of the end of 1996. Consistent with Chang, Khanna, and Palepu [2000], we find wide variation across counties in analyst following and forecast accuracy. The median earnings forecast error is 0.52% of market value and the median firm is followed by four analysts. Table 1 also reports sample statistics for the 4,452 firms that have three years of historical earnings data. This sample is used to control for historical volatility in earnings and returns-earnings correlation, which may affect our basic regression results. In general, the two samples are similar in terms of forecast accuracy and analyst following.

3. *The Effect of Cross Listing on Analyst Coverage and Performance*

3.1 EMPIRICAL METHODOLOGY

If there are information effects associated with cross-listing, we expect to see a relation between cross listing and the characteristics of a firm's information environment.[12] It is important to control for factors besides cross listing that are also likely to affect the information environment across firms. As such, we follow the models used in Lang and Lundholm [1993, 1996] for our primary specifications and estimate ordinary least squares (OLS) regression models of the following form:

$$\begin{aligned} \textit{Information Variables} = \beta_0 &+ \beta_1 \textit{XLIST} + \beta_2 \textit{Firm size} + \beta_3 \textit{Return STD} \\ &+ \beta_4 (\textit{Return-earnings correlation}) \\ &+ \beta_5 \textit{ Earnings surprise} + \textit{Industry controls} \end{aligned}$$

where:

Information Variables	=	number of analysts, forecast accuracy
XLIST	=	an indicator variable that takes the value 1 if the firm has an ADR traded in the U.S. that requires reconciliation to U.S. GAAP
Firm size	=	the log of total assets converted to millions of U.S. dollars
Return STD	=	the standard deviation of returns over the previous three years (winsorized at the 95th percentile)
Returns-earnings correlation	=	the correlation between returns and earnings over the previous three years
Earnings surprise	=	the absolute value of the difference between current earnings per share and earnings per

[12] Another potential measure of information environment used in studies such as Lang and Lundholm [1996] is the dispersion of forecasts. We do not use that measure in our primary analysis because it is not clear either theoretically or empirically how forecast dispersion affects the information environment of the firm (see Lang and Lundholm [1996], Harris and Raviv [1993], and Kandel and Pearson [1995]). However, as we discuss later, when we rerun our analysis using forecast dispersion, it is not significantly related to cross listing or firm value and its inclusion does not affect any of our other results.

share from the prior year, divided by the firm's stock price

Industry controls = indicator variables for I/B/E/S industry classification (more than 100 classifications that broadly correspond to two-digit Standard Industrial Classification [SIC] codes)

We estimate our first set of regression models with a relatively parsimonious set of control variables to maximize the number of observations and power of our tests. Because we are interested in whether cross listing improves the information environment of a firm, we focus on the coefficient of *XLIST*. Firm size, measured as the log of total assets converted to U.S. dollars, is included in all regressions because larger firms are likely to have more analysts covering them (Bhushan [1989], Brennan and Hughes [1991]) and more forthcoming disclosure policies (Lang and Lundholm [1996]), leading to better accuracy.[13] To control for industry effects, we include I/B/E/S industry classification dummies that broadly correspond to two-digit SIC codes. To control for cross-country effects, we estimate all regression models using country random effects. We verify that the Hausman test does not reject the null that country effects are random.

We estimate a second set of models with control variables suggested by Lang and Lundholm [1996]. These additional controls are the standard deviation of returns, the historical correlation between returns and earnings, and the earnings surprise. These variables are likely to affect forecasts because they affect analysts' incentives to gather information. Lang and Lundholm find that return variability is negatively related to the number of analysts following a U.S. firm, indicating that analysts prefer to follow firms with less performance variability. King, Pownall, and Waymire [1990] find that analyst following is positively related to the returns-earnings correlation for U.S. firms, indicating that the incentives for private information gathering are greater when earnings and returns are highly correlated, whereas Lang and Lundholm find the opposite relation. Finally, Lang and Lundholm include the percentage earnings surprise to control for the fact that forecast characteristics are likely to be affected by the magnitude of the earnings information to be disclosed. As a result, the effect on accuracy can be interpreted as the value that analysts bring in forecasting earnings relative to a naive random walk model.

We estimate an additional regression model that has specific control variables found in Alford and Berger [1999] and Lev and Thiagarajan [1993]. These studies postulate that the amount of new equity raised during the year might affect analyst coverage of a firm. We proxy for this measure,

[13] We use assets rather than the market value of equity as our size control because we hypothesize that stock market valuation is a function of the firm's information environment. Inclusion of industry controls should mitigate the effects of differences in tangible asset intensity across firms. In a prior version of the paper, we used the market value of equity as a size control with similar results. Results are also consistent if we use the log of sales.

called "stock," in our analyst regression using the difference in book equity between years. Because of Worldscope data limitations, the sample is smaller in this specification. Both studies also use a fundamental variable intended to capture signals plausibly related to forecast accuracy. Following Alford and Berger, we include in our accuracy regression a composite fundamental variable that sums the number of fundamental variables in the extreme quartiles of the distribution in year $t-1$ and is divided by the number of available variables. The fundamental variables are taken from Lev and Thiagarajan and comprise an inventory signal (the percentage change in inventory minus the percentage change in sales), an accounts receivable signal (the percentage change in accounts receivable minus the percentage change in sales), a gross margin signal (the percentage change in sales minus the percentage change in gross margin), a sales efficiency signal (the percentage change in selling and administration expenses minus the percentage change in sales), and a tax rate signal.[14]

Finally, we recognize that an important concern regarding these specifications is endogeneity. Suppose that firms with high analyst following or high forecast accuracy tend to cross list for reasons unrelated to their information environment and that our controls or country and industry effects do not capture this. Then, we might infer a link between information variables and cross-listing when none exists. For example, suppose that

$$Y = \beta x + \delta C + \varepsilon,$$

where C is the indicator variable that takes the value 1 if the firm lists in the United States. Because firms decide whether to cross-list based on various factors, we can model this decision as

$$C^* = \gamma' w + u$$
$$C = 1 \text{ if } C^* > 0, 0 \quad \text{otherwise.}$$

If the typical firm selects to cross-list because of some expected benefit in Y, OLS estimates of δ will not correctly measure the effect of cross-listing. This problem of self-selection is often handled empirically with a treatment effect model (e.g., see Greene [1990]).

To mitigate this potential endogeneity issue, we apply a self-selection model that controls for this bias. Similar to Doidge, Karolyi, and Stulz [Forthcoming], we model the decision to cross-list as a function of country-level variables, including the legal origin and the aggregate liquidity ratio (dollar value of shares traded divided by the average market capitalization in 1997). We also include firm-specific determinants of the cross-listing decision. Specifically, we include firm size as an explanatory variable because

[14] Our computation of the "fundamental" variable follows closely the procedure used in table 2 of Alford and Berger [1999], except that we do not make industry adjustments for changes in capital expenditures and our effective tax computation uses taxes paid as a percentage of earnings before tax.

larger firms are more likely to cross list, and we include three-year sales growth because high-growth firms are more likely to need capital and, hence, to cross list. We also include indicator variables corresponding to broad industry groupings as defined in Campbell [1996] because capital intensity (and, hence, financing needs), as well as other determinants of cross listing choice, is likely to be a function of industry. We obtain consistent estimates via full maximum likelihood estimation.[15]

3.2 THE RELATION BETWEEN CROSS LISTING AND ANALYST COVERAGE

To examine the relation between cross listing and analyst coverage, we estimate variants of our basic equation with the number of analysts as the dependent variable. Table 2 reports the coefficient estimates. Model 1 shows that after controlling for firm size, country, and industry, firms that cross list in the United States and reconcile to U.S. GAAP have an average of 3.28 more analysts covering the firm.

Model 1 does not include any historical earnings data, which increases sample size. In models 2 and 3 we investigate whether our results are robust when we require data on these additional control variables. In all specifications, the coefficient on *XLIST* is significant and similar in magnitude to the full-sample results. For example, model 2 indicates that firms that cross list in the United States have an average of 3.27 more analysts covering the firm. Model 3 shows that 2.32 more analysts cover a cross-listed firm and that new equity issuances also significantly explain analyst coverage. Overall, the control variables have the expected signs and are generally significant. In all models, analyst coverage relates positively to firm size and negatively to the standard deviation of returns. Model 2 also shows that analyst coverage relates positively to the correlation between returns and earnings.

The fourth and fifth columns of table 2 report results for the analyst following regression after controlling for potential selection bias. The explanatory variables are generally significant in the primary model. Consistent with expectations, the probit model shows that cross-listing firms tend to be larger firms from countries with an English legal tradition (the indicator variable excluded from the regression). The sales growth variable is not significant, although several of the industry controls (not tabulated) are. More important, with the selection-bias correction, the *XLIST* coefficient is similar in magnitude and significance to the model 3 result. Subject to the limitations of this self-selection approach, it does not appear that endogeneity drives our regression results.

Overall, the results contained in table 2 suggest that cross listing is associated with greater analyst following. This finding is consistent with the notion that cross listing lowers the cost of following the firm, which, in turn, leads to increased coverage of the firm by investment analysts. In addition,

[15] A Heckman [1979] two-step estimation procedure produces similar results.

TABLE 2
Multivariate Tests of Analyst Coverage

Model Specification	Random Effects Model 1	Random Effects Model 2	Random Effects Model 3	Treatment Effects Model 4 Probit	Treatment Effects Model 4
XLIST	3.2819	3.2671	2.3171		2.5441
	(0.00)	(0.00)	(0.00)		(0.00)
Firm size	1.9765	1.9830	1.1143	0.5477	2.0130
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Stock			1.2126		0.3289
			(0.00)		(0.00)
Return STD		−1.3810	−2.5048		−1.0577
		(0.00)	(0.00)		(0.02)
Returns-earnings correlation		0.6122	0.1836		0.1534
		(0.00)	(0.25)		(0.32)
Earnings surprise		−0.0024	−0.0032		−0.0129
		(0.49)	(0.65)		(0.65)
Industry dummies	*	*	*	*	*
Country dummies	*	*	*		*
French law				−0.5157	
				(0.02)	
German law				−1.6775	
				(0.00)	
Scandinavaian law				−0.7670	
				(0.03)	
Liquidity ratio				−0.3742	
				(0.23)	
Sales growth				0.0001	
				(0.992)	
Intercept	−22.7159	−22.0790	−1.8784	−8.5237	−32.1106
	(0.00)	(0.00)	(0.75)	(0.00)	(0.00)
N	4,859	4,452	2,600		1,545
Adj. R^2	0.30	0.32	0.40		

Analyst coverage is for fiscal year 1996 and is defined as the number of I/B/E/S analysts that report estimates for each firm. *XLIST* is an indicator variable that takes the value 1 if the firm has securities traded on a U.S. exchange. Firm size is the log of total assets in thousands of U.S. dollars. Stock is the log of the amount of new equity issued during year t–1 (in millions of U.S. dollars). Returns-earnings correlation is the correlation between returns and earnings over the previous three years. Return STD is the standard deviation of returns over the previous three years. Earnings surprise is the absolute value of the difference between current earnings per share and earnings per share from the prior year, divided by the firm's stock price. Legal variables are from La Porta et al. [1998]. Liquidity ratio is the ratio of the dollar value of shares traded divided by the average market capitalization in 1997. Sales growth is over the previous three years. Dummy variables for I/B/E/S industry classification (more than 100 classifications, broadly corresponding to two-digit SIC codes) are included but not reported. Country dummy variables are also included but not reported. Standard deviation of return on equity is winsorized at the 95th percentile. Regressions are estimated with robust standard errors. The p-value of the two-tailed t-test of equality of the coefficient to zero is reported in parentheses. ∗ denotes controls that are included but not reported.

the finding of increased analyst activity is consistent with an increase in the firm's shareholder base, because Bhushan and O'Brien [1990] document a strong correlation between institutional investors and analyst following. Therefore, our findings provide support for the hypothesis that cross listing improves the information environment of the firm, at least as measured by analyst following. Furthermore, because changes in the shareholder base have been linked to changes in stock returns around cross listing (Foerster and Karolyi [1999]), the increase in analyst activity may indeed lead to an increase in shareholder value. We return to this possibility later.

3.3 THE RELATION BETWEEN CROSS-LISTING AND FORECAST ACCURACY

In this section, we investigate the association between cross listing and the information environment of the firm by examining analysts' forecast accuracy. If cross listing increases the amount of information available about the firm, analysts should be able to predict more accurately the earnings of non-U.S. firms. We test this by reestimating our basic equation (1) with forecast accuracy as the dependent variable.

Model 1 of table 3 shows that after controlling for firm size, country, and industry effects, the coefficient on *XLIST* is positive and significant (0.0056, p-value = .04). This finding is consistent with the hypothesis that the increased transparency associated with cross listing provides analysts with more complete and reliable information with which to predict earnings.

Models 2 and 3 confirm that these results are robust when we include the additional historical earnings controls. Again, signs on the control variables are generally consistent with the prior literature. Forecast accuracy is higher for larger firms and firms with a greater correlation between returns and earnings, and is lower for firms with more volatile returns and in years in which there are large earnings surprises. Finally, as with table 2, the selection-bias correction applied in columns 4 and 5 of table 3 indicates that the positive relation between cross-listing and accuracy is robust to endogeneity controls.

3.4 OTHER ANALYSES

Our previous regressions control for self-selection among cross-listing firms. Another potential concern with our analysis is that we have misspecified our models by estimating the analyst following regression separately from the forecast accuracy model. In particular, Alford and Berger [1999] suggest that analyst forecast accuracy and analyst following might be simultaneously determined. To assess this issue, we specify two models similar to those contained in Alford and Berger. We model analyst coverage as a function of forecast accuracy, cross listing, total assets, and stock, and we model forecast accuracy as a function of analyst coverage, cross listing, the standard deviation of returns, total assets, and fundamentals. The results (not tabulated) are generally consistent with those reported in tables 2 and 3. In particular, the coefficient on the cross listing variable is 1.26 (significant at the .04 level) in the analyst following regression and 0.0076 (significant at the .06 level) in the forecast accuracy regression.

We also test whether the effect of cross listing is larger for some countries than for others. In particular, Ball, Kothari, and Robin [2000] suggest that

TABLE 3

Multivariate Tests of Forecast Accuracy

Model Specification	Random Effects Model 1	Random Effects Model 2	Random Effects Model 3	Treatment Effects Model 4 Probit	Treatment Effects Model 4
XLIST	0.0056	0.0062	0.0082		0.0124
	(0.04)	(0.03)	(0.01)		(0.00)
Firm size	0.0023	0.0022	0.0018	0.4406	0.0016
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Fundamentals			−0.0013		−0.0043
			(0.56)		(0.24)
Return STD		−0.0076	−0.0074		−0.0098
		(0.00)	(0.01)		(0.01)
Returns-earnings correlation		0.0029	0.0025		0.0019
		(0.00)	(0.01)		(0.05)
Earnings surprise		−0.0002	−0.0002		−0.0039
		(0.00)	(0.00)		(0.00)
Industry dummies	*	*	*	*	*
Country dummies	*	*	*		*
French law				−0.6493	
				(0.00)	
German law				−1.4955	
				(0.00)	
Scandinavian law				−0.6939	
				(0.01)	
Liquidity ratio				−0.2236	
				(0.37)	
Sales growth				−0.0112	
				(0.11)	
Intercept	−0.1326	−0.0562	−0.1272	−6.8568	−0.0190
	(0.00)	(0.15)	(0.00)	(0.00)	(0.79)
N	4,859	4,452	3,819		2,293
Adj. R^2	0.09	0.13	0.13		

Analyst coverage is for fiscal year 1996 and is defined as the number of I/B/E/S analysts that report estimates for each firm. *XLIST* is an indicator variable that takes the value 1 if the firm has securities traded on a U.S. exchange. Firm size is the log of total assets in thousands of U.S. dollars. Fundamentals is a score equal to the number of fundamental variables (see text for list of variables) in the extreme quartiles of the distribution in year t-1 divided by the number of available variables. Returns-earnings correlation is the correlation between returns and earnings over the previous three years. Return STD is the standard deviation of returns over the previous three years. Earnings surprise is the absolute value of the difference between current earnings per share and earnings per share from the prior year, divided by the firm's stock price. Legal variables are from La Porta et al. [1998]. Liquidity ratio is the ratio of the dollar value of shares traded divided by the average market capitalization in 1997. Sales growth is over the previous three years. Dummy variables for I/B/E/S industry classification (more than 100 classifications, broadly corresponding to two-digit SIC codes) are included but not reported. Country dummy variables are also included but not reported. Standard deviation of return on equity is winsorized at the 95th percentile. Regressions are estimated with robust standard errors. The p-value of the two-tailed t-test of equality of the coefficient to zero is reported in parentheses. $*$ denotes controls that are included but not reported.

developed common law countries such as the United Kingdom, Canada, and Australia tend to be much more similar institutionally to the United States than code (civil) law countries and countries with emerging markets. To examine whether institutional differences might be important, we use classifications in *The Economist* magazine and in La Porta et al. [1998] to split the sample into two subsamples: developed common law countries and emerging market or code law countries.[16]

The results (not tabulated) are consistent with the prediction that cross-listing effects are stronger among emerging, code law countries. In particular, emerging, code law firms with an ADR have 4.10 more analysts than firms that do not have an ADR, whereas among developed market common law countries, firms with an ADR have 1.36 more analysts. The difference between the ADR effect on analyst coverage for emerging, code law countries compared with developed common law countries is significant at the .01 level. Results are generally consistent, but weaker, for forecast accuracy. Firms from emerging, code law countries with an ADR have 0.080 higher forecast accuracy than firms that do not have an ADR, whereas among developed market common law countries the difference is only 0.010. Although this difference is large in magnitude, standard errors are also large and the difference is statistically insignificant (p-value = .29).

We also analyze whether cross listing affects the dispersion of forecasts. As noted earlier, we do not rely on forecast dispersion as a primary measure of a firm's information environment because it is difficult to interpret clearly. For example, Barron et al. [1998] assume that all analysts share the same likelihood function (i.e., they interpret shared information in the same way) and argue that additional information could either increase or decrease the dispersion of forecasts depending on whether it is private or public information. Harris and Raviv [1993] and Kandel and Pearson [1995] demonstrate that if analysts differ in terms of their interpretation of public information, increased public disclosure could either increase or decrease forecast dispersion. Similarly, Lang, Raedy, and Yetman [2003] suggest that firms appear to change local accounting choices around cross listing, reducing smoothing, and increasing the variability and skewness of reported earnings. To the extent that such changes make earnings more difficult to predict and increase uncertainty about current earnings, they could also increase dispersion even though the market is better informed about the firm's economic prospects.

The relation between dispersion and the cost of capital is also ambiguous. Although there are theoretical and empirical reasons to believe that the cost of capital is lower for firms that have greater analyst following and forecast accuracy, for forecast dispersion the relation is not as clear. Conceptually, if new information causes forecast dispersion, firms with less consensus may actually have lower cost of capital. Consistent with this, Gebhardt, Lee, and Swaminathan [2001] find that their estimates of implied cost of capital are decreasing in forecast dispersion after controlling for other factors,

[16] We combine emerging and code law countries because there are relatively few emerging market observations and the most striking comparison should be between developed common countries (i.e., such as the United States) and other countries.

suggesting that the market views firms with higher forecast dispersion as less risky.[17]

Despite the difficulties inherent in interpreting forecast dispersion, we reestimate our models with forecast dispersion, rather than forecast accuracy, as our dependent variable (not tabulated). Across all specifications, the relation between dispersion and cross listing is negative, but generally it is not significant. Coupled with the results for forecast accuracy, the dispersion results suggest that, although cross listing increases investor understanding of the firm, it does not appreciably affect dispersion. This could potentially be the result of public information being interpreted differently across analysts or of increased private information acquisition accompanying increased public disclosure.

Although we do not know of a way to disentangle convincingly these interpretations, we use the Barron et al. [1998] approach in an attempt to decompose information into public and private components. Barron et al. propose that "private information" can be measured as a ratio with dispersion in the numerator and 1 minus the number of analysts multiplied by dispersion plus the squared mean forecast error, quantity squared, in the denominator. Public information can be measured as a ratio with the squared mean forecast error less the ratio of dispersion to the number of analysts in the numerator and the same denominator as the private information variable.

To get a sense for what these measures tell us about our cross-listed firms, we compare total, private, and public information for our cross-listed sample with our non-cross-listed sample. Results (not tabulated) suggest that, by this measure, total information available to the market is higher for the cross-listed firms, as expected. Comparing the components of information, it is the "private information" component that drives the empirical result; the measure of "public information" is positive, but not significant. We hesitate to draw strong conclusions, however, because our interpretation is sensitive to the assumption of homogenous interpretation of information on the part of analysts. To the extent that analysts differ in how they interpret information, increased public information could produce similar results.

3.5 TIME-SERIES ESTIMATION

Our cross-sectional analysis establishes a link between cross listing and a firm's information environment. Furthermore, the results are robust to our adjustments for potential endogeneity and simultaneity. However, a cross-sectional analysis does not shed light on the extent to which the information

---

[17] Gehardt, Lee, and Swaminathan [2001] find that the correlation between dispersion and implied cost of capital is positive, but that the sign changes when other variables are included. They argue that a negative relation between dispersion and cost of capital is consistent with the predictions in Miller [1977] in the presence of short-selling constraints and heterogeneous beliefs.

environment changes around cross-listing. For example, it could still be the case that only firms with high analyst following and accurate forecasts choose to cross list and our controls have not successfully mitigated that effect. An alternate approach that allows us to assess to what extent observed differences between ADR and non-ADR firms are due to self-selection is to use time-series data for ADR firms to examine how the information environment changes before and after cross-listing. If we are simply documenting self-selection, we should not observe significant changes in information production around the cross-listing period.

Our time-series tests consist of regressions estimated using panel data on analyst activity around the time of the ADR listing. As noted earlier, a problem with this approach is determining the window over which to expect the change. For instance, it can be argued that cross-listing firms will incrementally change their reporting strategies before the formal adoption of U.S. GAAP. We choose to use a long event window, including the three years before and after cross listing, to increase our confidence that we do, in fact, capture the entire effect and have enough observations to ensure reliable measures of our dependent variables. However, this approach reduces our sample size to 59 firms, which lessens the power of our statistical tests.

We first estimate models in which the dependent variable in our regressions is either the number of analysts or forecast accuracy. To capture whether analysts or accuracy change around the ADR, we construct a *Post-ADR* dummy variable set equal to 1 for the years in the panel data set that fall after the firm has issued its exchange-listed ADR. Because the event centers on the year of the ADR, this year is removed from the regressions. All regressions include firm fixed effects and are reported with robust standard errors with firm clusters that account for a lack of independence between the observations of each firm. Firm fixed effects implicitly control for the calendar year of each ADR firm, which helps to ensure that our results are not driven by a subset of years in which analyst following or forecast accuracy increased for other reasons. The time-series distribution of our sample ADRs is as follows (number of ADRs per year is in parentheses): 1990 (5), 1991 (10), 1992 (7), 1993 (9), 1994 (14), 1995 (10), 1996 (4).

It would be optimal to control also for contemporaneous changes in the historical earnings variables, the standard deviation of returns, and returns-earnings correlation in our time-series models. Unfortunately, we cannot include these controls from our cross-sectional regressions because each variable requires three years of historical data to compute, and our time-series analysis already requires three years of data before the ADR listing date. The I/B/E/S data do not go back far enough to make these controls feasible. Furthermore, we cannot include total assets because doing so requires matching to Worldscope, which substantially reduces the sample size. We include the firm's prior-year stock return as a simple contemporaneous control to capture the possibility that analysts may be drawn to firms with higher returns to shareholders. We also include earnings surprise in these models because it, too, requires only one year of historical earnings data. Including these two control variables causes us to lose the initial pre-ADR observation for 14 of our firms. We also lose 3 observations because of fiscal month changes. Thus, our panel for the 59 firms is slightly unbalanced (337 observations compared with an expected 354 observations). For robustness, we reestimate our models using only the first two post-ADR years when the initial pre-ADR year is missing because of data omissions (thus obtaining a balanced panel) and find that our conclusions are unchanged (not tabulated).

A related concern is whether we might be capturing the effects of general changes in forecast accuracy or analyst following for the population of firms as a whole. To address this possibility, we adopt two approaches. First, for each ADR firm we adjust our change in analyst following and accuracy variables for the median change in analyst following and accuracy for all firms with a market capitalization of $100 million or more from the ADR firm's home country over the same period. This procedure should capture general countrywide trends in analyst coverage and forecast accuracy. In our second approach, we compare results for the cross-listed firms with a sample of firms matched by country, year, industry, and size. This matched procedure should more closely control for changes in analyst following and accuracy specific to a firm's country, industry, and size, but unrelated to cross-listing. We lose six firms because of an inability to match on industry in a given country and year.

Table 4 presents the results of our time-series analyses. Regression models 1, 3, and 5 are estimated for analyst coverage. In terms of the unadjusted results contained in model 1, ADR firms experienced an increase of 3.81 analysts around their cross listing. Model 3 reports that when adjusted for country median analyst coverage, by year, ADR firms still pick up 2.75 more analysts around their cross listing. Together, models 1 and 3 show that the median non-ADR home-country firm had an increase in analyst coverage of 1.06 analysts (computed as 3.81 − 2.75). Finally, our matched sample regression (model 5) indicates that firms similar to our ADR firms experience an increase in analyst coverage of 1.37 analysts around a typical cross listing date, which corresponds to an ADR firm's picking up 2.44 more analysts around cross listing (3.81 − 1.37). All of these differences are significant at the .01 level. Overall, our results are consistent with Baker, Nofsinger, and Weaver [2002], who also document that analyst coverage increases substantially following cross listing.

In terms of accuracy, models 2 and 4 show that the cross-listed firms experience an increase of 0.015, relative to an increase of 0.004 (computed as 0.015 − 0.011) for median firms from their country, a difference that is significant at the .01 level. As reported in model 6 of table 4, the matched sample experiences a decrease in accuracy of −0.007. This result is also significantly different from the cross-listing firm result at the .01 level. The time-series results indicate that the cross-listing firms experience an improvement in their information environment beyond that experienced by other firms in their

TABLE 4

Time-Series Tests of Analyst Coverage and Forecast Accuracy

Independent Variables	ADR Firms: Analyst Coverage Model 1	ADR Firms: Forecast Accuracy Model 2	ADR Firms: Adjusted Analyst Coverage Model 3	ADR Firms: Adjusted Forecast Accuracy Model 4	Matched Sample of Non-ADR Firms: Analyst Coverage Model 5	Matched Sample of Non-ADR Firms: Forecast Accuracy Model 6
Post-ADR dummy	3.814	0.015	2.750	0.011		
	(0.00)	(0.00)	(0.00)	(0.03)		
Post-pseudo ADR dummy					1.374	−0.007
					(0.02)	(0.46)
Prior-year return	−0.427	0.004	−0.674	0.006	−0.562	0.018
	(0.07)	(0.12)	(0.00)	(0.15)	(0.06)	(0.08)
Earnings surprise	−1.393	−0.023	−1.312	−0.023	−0.043	−0.081
	(0.00)	(0.01)	(0.00)	(0.01)	(0.03)	(0.00)
Intercept	12.931	−0.027	6.893	−0.018	11.053	−0.028
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Number of observations	337	337	323	323	292	292
Number of firms	59	59	59	59	53	53
Adj. R^2	0.78	0.29	0.77	0.28	0.90	0.91

Time-series regression estimates of dependent variables for analyst coverage and analyst forecast accuracy on disclosure variables of interest and controls. The table contains results for two samples and four dependent variables. Models 1 through 4 are estimated on a sample consisting of ADR firms for which three years pre- and post-ADR data for control variables are available. Post-ADR is a dummy variable equal to 1 for the years in the panel data set that fall after the year the firm listed its ADR. Models 5 and 6 are estimated on a non-ADR matched sample for these ADR firms. Matches are based first on industry groupings and then on size. Industry matches are unavailable for six of the ADR firms. A pseudo-ADR date for each non-ADR firm is set equal to the year of the ADR for the respective matching ADR firm. Post-pseudo ADR is a dummy variable equal to 1 for the years in the panel data set that fall after the year of the pseudo-ADR. The data for the non-ADR matched sample span at least two years before and after the pseudo-ADR date with a maximum of three years pre- and post-pseudo ADR. Dependent variables used previously are Forecast accuracy, defined as the negative of the absolute value of the analyst forecast error deflated by stock price, and Analyst coverage, defined as the number of I/B/E/S analysts that report estimates for each firm. Dependent variables new to this table are Adjusted analyst coverage and Adjusted forecast accuracy. These adjusted measures are computed by subtracting median analyst coverage and forecast accuracy, by country and year for all firms over $100 million in size, from the analyst coverage and forecast accuracy values for an ADR firm in a given year. Prior-year return is the total stock return over the previous year. All regressions include firm fixed effects and do not include data for the year of the listing. Regressions are estimated with robust standard errors by firm clusters that account for a lack of independence between the observations of each firm. The *p*-value of the two-tailed *t*-test of equality of the coefficient to zero is reported in parentheses.

home markets. Overall, the evidence contained in table 4 indicates that self-selection is not exclusively driving the previous cross-sectional results. We interpret these findings as additional support for our cross-sectional analysis that the information environment, as represented by analyst coverage and forecast accuracy, improves around a U.S. cross listing.

Taken together, the evidence in the preceding sections suggests that firms that cross list in the United States have greater analyst coverage and increased forecast accuracy relative to other firms not cross listed, providing empirical support for the notion that the cross listing is associated with a significantly improved informational environment. We next investigate whether this increase in information production translates into a higher firm value.

4. The Relation Between Analyst Coverage and Accuracy and Firm Value

4.1 CROSS-SECTIONAL ESTIMATION

Tobin's Q is our proxy for firm value and these data are obtained for our sample firms from Worldscope. There are 5,539 firms with 1996 IBES analyst data for which fiscal-year 1996 Tobin's Q can be computed. Table 5 reports the results of regression models in which Tobin's Q is regressed on both forecast accuracy and the number of analysts as well as controls. Because country-level factors such as investor protections can have important implications for firm value (La Porta et al. [2002], Doidge, Karolyi, and Stulz [Forthcoming]), we employ country controls. Again, we implement the treatment effect model to account for potential endogeneity using the same decision criteria as in tables 2 and 3.

Table 5 shows that after controlling for firm size, industry, and country of domicile, the coefficient on forecast accuracy is positive and significant (2.3670, p-value = .00). This finding suggests that when analyst forecasts are more accurate, firm values are higher. This higher valuation is consistent with the Gebhardt, Lee, and Swaminathan [2001] results suggesting that firms with lower forecast errors tend to have a lower implied cost of capital or higher expected cash flows. It also supports the hypothesis that firm value is a function of estimation risk and that the better investors are able to assess the prospects of a company, the greater is firm value.

Table 5 also shows that the coefficient on the number of analysts is positive and significant (0.0287, p-value = .00). This evidence indicates that firm value is increasing in analyst following. Because analyst following is likely to be correlated with investor following, the higher valuation is consistent with a lower cost of capital or with reduced agency costs. Because both forecast accuracy and analyst coverage are positively and significantly related to firm value, this suggests that cross listing may indeed increase value by improving the information environment of the firm. For example, if firms that cross-list pick up an additional 2.44 analysts and a 0.022 increase in accuracy (from the time-series analysis in table 4), this would translate into an increase of 12.2% in firm value.[18] These results are robust to the inclusion of the cross listing variable, suggesting that it is not simply that we are capturing the higher valuations of cross-listed firms documented in Doidge, Karolyi, and Stulz [Forthcoming]. On the other hand, the cross listing variable is also significant in the presence of forecast accuracy and the number of analysts, indicating that information effects (at least as we measure them) do not alone explain the higher valuations of cross-listed firms.

Our goal in table 5 is simply to show that higher levels of analyst following and forecast accuracy are generally associated with higher values of Tobin's Q. Our concern is that, in the absence of showing a relation, it is more

[18] Computed as $(2.3670 * 0.022) + (0.0287 * 2.44) = 12.21\%$.

TABLE 5

Regressions of Tobin's Q on Accuracy and Analyst Coverage, Selection Bias Corrected

	Probit	Tobin's Q
Forecast accuracy		2.3670
		(0.00)
Number of analysts		0.0287
		(0.00)
XLIST		0.3542
		(0.04)
Firm size	0.4464	−0.1370
	(0.00)	(0.00)
Profit		2.0279
		(0.15)
Sales growth	−0.0105	0.0033
	(0.12)	(0.01)
Industry dummies	*	*
Country dummies		*
French law	−0.6321	
	(0.00)	
German law	−1.4920	
	(0.00)	
Scandinavian law	−0.7347	
	(0.01)	
Liquidity ratio	−0.2166	
	(0.39)	
Intercept	−6.9514	2.7929
	(0.00)	(0.00)
N		2,325
Psuedo-R^2		0.21

Regression estimates of Tobin's Q on disclosure variables of interest and controls. Tobin's Q is computed as total assets less the book value of equity plus the market value of equity in the numerator and book value of assets in the denominator. Tobin's Q is computed using data predominantly from fiscal year 1996. Firms with financial services industry classifications are excluded from the regressions. Forecast accuracy is for fiscal year 1996 and is defined as the negative of the absolute value of the analyst forecast error, deflated by stock price. Number of analysts is defined as the median number of I/B/E/S analysts that report estimates for each firm. Firm size is the log of total assets in thousands of U.S. dollars. Profit is operating income deflated by total assets, in U.S. dollars. Sales growth is over the previous three years. Legal variables are from La Porta et al. [1998]. Liquidity ratio is the ratio of the dollar value of shares traded divided by the average market capitalization in 1997. All regressions are estimated using unreported dummy variables for I/B/E/S industry classification (more than 100 classifications, broadly corresponding to two-digit SIC codes) and country of domicile. Regressions are estimated with heteroskedastic-consistent robust standard errors. The p-value of the two-tailed t-test of equality of the coefficient to zero is reported in parentheses. * denotes included but not reported.

difficult to argue that analyst following and forecast accuracy are likely to be associated with firm value. A related question is whether the relation between firm value and information should be more pronounced for cross-listing firms—testing this would require interaction variables. However, we do not have strong priors on whether there should, indeed, be an incremental effect. For example, one might argue that analyst following is more important for a non-cross-listed firm because governance issues are likely to be more pronounced for these firms.

Absent a strong prediction, we nevertheless estimate (unreported) regression models with interactions between cross listing and our information variables. We find that the coefficient on the accuracy and *XLIST* interaction is significant and positive and the one on the analyst coverage and *XLIST* interaction is marginally significant and negative (p-value = .08). The negative coefficient on the analyst following interaction is consistent with the notion that cross listing and analyst following are substitutes; therefore, an increase in analyst following is relatively less important for valuation if the firm is taking the other steps associated with cross listing. That said, it is also possible that cross-listing firms are picking up some analysts who are free riding on other analysts because of increased investor interest. The accuracy result suggests that the increased accuracy is particularly important for cross-listing firms.

4.2 TIME-SERIES ESTIMATION

To document an incremental effect of analyst following and forecast accuracy on firm value around the cross listing, we need tests that control for other factors associated with cross listings and increased disclosure, which may, in turn, affect firm value. For instance, our prior analysis indicates that analyst following and forecast accuracy increase around cross listing, and we know from prior research that cross-listing firms, in general, have positive returns. Therefore, we are concerned that a documented positive relation between changes in information variables and changes in value could simply capture that phenomenon.

The crux of our valuation question is whether firms that experience the greatest improvement in their information environment enjoy the greatest valuation benefits to cross listing. To answer this question, we implement a two-stage approach. In the first stage, we use our models that explain analyst following and forecast accuracy to predict the expected levels of these measures for each firm for each year from 1990 to 1998. Next, for our cross-listed sample firms, we obtain the residuals from these models, which correspond to the levels of excess analysts and excess forecast accuracy for each cross-listed firm by year. Because the coefficients of these models are determined from the full sample of firms, we are able to determine the cross-listed firm's expected information environment (in terms of analyst following and forecast accuracy) based on its overall firm attributes.

Our approach to obtain predicted and residual values in each year is to regress either analyst coverage or forecast accuracy on firm size, the standard deviation of returns, the return to earnings correlation, and the earnings surprise. The models used correspond to model 2 of tables 2 and 3, except for our size control, which uses the market value of equity from I/B/E/S rather than a firm's total assets. This substitution avoids a substantial loss of observations in the early years due to poor Worldscope coverage.

We next compute Tobin's Q values for our cross-listed firms for each year from 1990 to 1998. Finally, we conduct a regression analysis (estimated with heteroskedastic-consistent standard errors) in which the dependent variable is the change in Tobin's Q from the year before the ADR to the year

TABLE 6

Regressions of Changes in Tobin's Q on Changes in Residual Accuracy and Analyst Coverage Measures

	Model 1	Model 2	Model 3
Excess analyst following	0.049		0.051
	(0.02)		(0.01)
Excess forecast accuracy		4.943	5.255
		(0.04)	(0.04)
Intercept	0.392	0.280	0.362
	(0.01)	(0.02)	(0.01)
N	66	66	66
Adj. R^2	0.05	0.04	0.08

The dependent variable in all regression models is the change in Tobin's Q from the year before the ADR to the year after the ADR (a two-year change). For each year, Tobin's Q is computed as total assets less the book value of equity plus the market value of equity in the numerator and book value of assets in the denominator. The independent variables are the changes in the residual levels of analyst coverage and forecast accuracy from the year before the ADR to the year after the ADR. To obtain residuals in each year, we regress either analyst coverage or forecast accuracy on firm size, the standard deviation of returns, the returns to earnings correlation, and the earnings surprise for 1990 to 1998. Forecast accuracy is defined as the negative of the absolute value of the analyst forecast error, deflated by stock price. Number of analysts is defined as the median number of I/B/E/S analysts that report estimates for each firm. The models used in computing the residual values correspond exactly to model 2 of tables 2 and 3, with the exception of our size control, which uses the market value of equity rather than a firm's total assets. This substitution avoids a substantial loss of observations in the early years because of poor Worldscope coverage. We refer to the residual values each year as measures of a firm's excess analyst following and excess forecast accuracy. Our regressions of changes in Tobin's Q on changes in excess analyst and accuracy are estimated with heteroskedastic-consistent robust standard errors. The p-value of the two-tailed t-test of equality of the coefficient to zero is reported in parentheses.

after the ADR. This two-year window allows us to include 66 firms in the analysis—enlarging the window beyond two years is not feasible, as it would reduce the sample by half. The independent variables in the regression are the changes in the residual levels of analyst coverage and forecast accuracy from the year before the ADR to the year after the ADR. Our argument is that if changes in the information environment are important benefits to cross listing, we should see a positive association; firms that benefit the most from cross listing should be those for which cross listing more significantly improves their information environment.

Table 6 presents results from this analysis. Our results show that an improved information environment around cross-listing is associated with an increase in firm valuation. In particular, there is a positive association between the change in Tobin's Q and the change in excess analysts and excess forecast accuracy. Models 1 and 2 in the table estimate the effects separately, and model 3 indicates that changes in both excess analysts and excess forecast accuracy are important in explaining changes in firm value. We conclude from this valuation analysis that cross listing does indeed enhance firm value through its effect on the firm's information environment.

5. *Conclusion*

In this paper, we examine whether cross listing on U.S. stock exchanges improves the information environment and, ultimately, the valuation of non-U.S. firms. We document several interesting findings. First, we show that non-U.S. firms that cross list enjoy greater analyst coverage and increased forecast accuracy relative to other firms that are not cross listed. Second, a time-series analysis shows that the change in analyst coverage and forecast accuracy occurs around the cross listing period. Third, we document that firms that have more analyst coverage and higher forecast accuracy have a higher valuation. Finally, we show that ADR firms with greater improvements in their information environment around cross listing also experience larger increases in valuations, which is consistent with these firms enjoying a lower cost of capital or improved corporate governance.

These findings have important implications for several strands of research. The large literature on international cross listings suggests that information disclosure plays a key role in the cross listing decision. Although theory predicts firms that cross list on a more transparent exchange should be more highly valued, there has been little direct empirical evidence regarding the role of the information environment and its impact on cross listing. Our findings provide evidence that important changes occur in the information environment of firms around cross listing and that these changes are rewarded with higher valuations by the market. In addition, because other factors such as investor protection and agency problems are argued to be important to the cross listing decision, our findings suggest control variables that may allow for a more detailed examination of other benefits to cross listing.

Our findings also are consistent with the literature that links disclosure to the cost of capital. Theoretical research such as Diamond and Verrecchia [1991] and Baiman and Verrecchia [1996] indicates that a commitment to increase transparency will be rewarded with a reduced cost of capital. Our findings that link analyst activity and firm value around cross listing are consistent with these theories and complement recent empirical work such as Leuz and Verrecchia [2000] and Gebhardt, Lee, and Swaminathan [2001].

REFERENCES

ALEXANDER, G.; C. S. EUN; AND S. JANAKIRAMANAM. "Asset Pricing and Dual Listing on Foreign Capital Markets: A Note." *Journal of Finance* 42 (1987): 151–58.

ALEXANDER, G.; C. S. EUN; AND S. JANAKIRAMANAM. "International Listings and Stock Returns: Some Empirical Evidence." *Journal of Financial and Quantitative Analysis* 23 (1988): 135–52.

ALFORD, A., AND P. BERGER. "A Simultaneous Equations Analysis of Forecast Accuracy, Analyst Following, and Trading Volume." *Journal of Accounting, Auditing, and Finance* (1999): 219–46.

ALLAYANNIS, Y., AND J. WESTON. "The Use of Foreign Currency Derivatives and Firm Market Value." *Review of Financial Studies* 14 (2001): 243–76.

BAILEY, W.; G. A. KAROLYI; AND C. SALVA. "The Economic Consequences of Increased Disclosure: Evidence from International Cross-listings." Working paper, Cornell University, 2002.

BAIMAN, S., AND R. VERRECCHIA. "The Relation Among Capital Markets, Financial Disclosure, Production Efficiency, and Insider Trading." *Journal of Accounting Research* (1996): 1–22.

BAKER, K.; J. R. NOFSINGER; AND D. G. WEAVER. "International Cross-Listing and Visibility." *Journal of Financial and Quantitative Analysis* 37 (2002): 495–521.

BALL, R. "Infrastructure Requirements for an Economically Efficient System of Public Financial Reporting and Disclosure." *Brookings-Wharton Papers on Financial Services* (2001): 127–82.

BALL, R.; S. P. KOTHARI; AND A. ROBIN. "An Extension of the Effect of International Institutional Factors on the Properties of Accounting Earnings." *Journal of Accounting Economics* 29 (2000): 1–51.

BARRON, O.; O. KIM; S. LIM; AND D. STEVENS. "Using Analysts' Forecasts to Measure Properties of Analysts' Information Environment." *The Accounting Review* 73 (1998): 421–33.

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

BHUSHAN, R. "Firm Characteristics and Analyst Following." *Journal of Accounting and Economics* 11 (1989): 255–74.

BHUSHAN, R., AND P. O'BRIEN. "Analyst Following and Institutional Ownership." *Journal of Accounting Research* (1990): 55–82.

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

BOTOSAN, C., AND M. PLUMLEE. "A Re-examination of Disclosure Level and the Expected Cost of Equity Capital." *Journal of Accounting Research* (2002): 21–40.

BRADSHAW, M. T., AND G. S. MILLER. "Are Detailed Accounting Standards Sufficient to Ensure Compliance? Evidence from Non-U.S. Firms Implementing US GAAP." Working paper, Harvard Business School, 2002.

BRENNAN, M., AND P. HUGHES. "Stock Prices and the Supply of Information." *Journal of Finance* 46 (1991): 1665–91.

BUSHMAN, R.; J. PIOTROSKI; AND A. SMITH. "What Determines Corporate Transparency?" Working paper, University of North Carolina, 2001.

BUSHMAN, R.; J. PIOTROSKI; AND A. SMITH. "Does Insider Trading Crowd Out Information Intermediaries?" Working paper, University of North Carolina, 2002.

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

CANTALE, S. "The Choice of a Foreign Market as a Signal." Working paper, Tulane University, 1998.

CHANG, J.; T. KHANNA; AND K. PALEPU. "Analyst Activity Around the World." Working paper, Harvard University, 2000.

COFFEE, J. "The Future as History: The Prospects for Global Convergence in Corporate Governance and its Implications." *Northwestern University Law Review* (1999): 641–708.

COFFEE, J. "Racing Towards the Top?: The Impact of Cross-Listings and Stock Market Competition on International Corporate Governance." Working paper, Columbia Law School, 2002.

DIAMOND, D. W., AND R. E. VERRECCHIA. "Disclosure, Liquidity, and the Cost of Capital." *Journal of Finance* 46 (1991): 1325–59.

DOIDGE, C.; G. A. KAROLYI; AND R. STULZ. "Why are Foreign Firms Listed in the U.S. Worth More?" *Journal of Financial Economics* (Forthcoming).

ERRUNZA, V., AND E. LOSQ. "International Asset Pricing Under Mild Segmentation: Theory and Test." *Journal of Finance* 40 (1985): 105–24.

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

FOERSTER, S. R., AND G. A. KAROLYI. "The Effects of Market Segmentation and Investor Recognition on Asset Prices: Evidence from Foreign Stocks Listing in the U.S." *Journal of Finance* 54 (1999): 981–1013.

FOERSTER, S. R., AND G. A. KAROLYI. "The Long Run Performance of Global Equity Offerings." *Journal of Financial and Quantitative Analysis* 35 (2000): 499–528.

FUERST, O. "A Theoretical Analysis of the Investor Protection Regulations: Argument for Global Listing of Stocks." Working paper, Yale School of Management, 1998.

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

GREENE, W. *Econometric Analysis*. New York: MacMillian Press, 1990.

HARRIS, M., AND A. RAVIV. "Differences in Opinion Make a Horse Race." *Review of Financial Studies* 6 (1993): 473–94.

HEALY, P. M.; A. P. HUTTON; AND K. G. PALEPU. "A Stock Performance and Intermediation Changes Surrounding Sustained Increases in Disclosure." *Contemporary Accounting Research* 16 (1999): 485–20.

HECKMAN, J. "Sample Selection Bias as a Specification Error." *Econometrica* 47 (1979): 153–61.

HUDDART, S.; J. HUGHES; AND M. BRUNNERMEIER. "Disclosure Requirements and Stock Exchange Listing Choice in an International Context." *Journal of Accounting and Economics* 26 (1999): 237–69.

KANDEL, E., AND N. PEARSON. "Differential Interpretation of Public Signals and Trade in Speculative Markets." *Journal of Political Economy* 103 (1995): 831–53.

KAROLYI, G. A. "Why Do Companies List Their Shares Abroad? A survey of the evidence and Its Managerial Implications." *Salomon Brothers Monograph Series*, New York University, 1998.

KING, R.; G. POWNALL; AND G. WAYMIRE. "Expectations Adjustments Via Timely Management Forecasts: Review, Synthesis, and Suggestions for Future Research." *Journal of Accounting Literature* 9 (1990): 113–44.

LANG, L., AND R. STULZ. "Tobin's *q*, Diversification, and Firm Performance." *Journal of Political Economy* (1994): 1248–80.

LANG, M. H.; K. V. LINS; AND D. P. MILLER. "Do Analysts Matter Most When Investors Are Protected Least? International Evidence." Working paper, University of North Carolina-Chapel Hill, University of Utah, and Indiana University, 2002.

LANG, M. H., AND R. LUNDHOLM. "Cross-Sectional Determinants of Analyst Ratings of Corporate Disclosures." *Journal of Accounting Research* (1993): 246–71.

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

LANG, M.; J. S. RAEDY; AND M. H. YETMAN. "How Representative Are Firms That Are Cross-Listed? An Analysis of Accounting Quality." *Journal of Accounting Research* 41 (2003): 363–86.

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

LA PORTA, R.; F. LOPEZ-DE-SILANES; A. SHLEIFER; AND R. W. VISHNY. "Investor Protection and Corporate Valuation." *Journal of Finance* 58 (2002): 1147–70.

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

LEV, B., AND R. THIAGARAJAN. "Fundamental Information Analysis." *Journal of Accounting Research* 31 (1993): 190–215.

LINS, K. "Equity Ownership and Firm Value in Emerging Markets." *Journal of Financial and Quantitative Analysis* (Forthcoming).

LINS, K.; D. STRICKLAND; AND M. ZENNER. "Do Non-U.S. Firms Issue Stock on U.S. Equity Markets to Relax Capital Constraints?" Working paper, University of Utah, 2002.

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

MILLER, D. P. "The Market Reaction to International Cross-listings: Evidence from Depositary Receipts." *Journal of Financial Economics* 51 (1999): 103–23.

MILLER, E. "Risk, Uncertainty, and Divergence of Opinion." *Journal of Finance* 32 (1977): 1151–68.

MOEL, A. "The Role of Information Disclosure on Stock Market Listing Decisions: The Case of Foreign Firms Listing in the U.S." Working paper, Harvard Business School, 1999.

O'BRIEN, P., AND R. BHUSHAN. "Analyst Following and Institutional Holding." *Journal of Accounting Research* (1990): 55–76.

PAGANO, R.; A. ROELL; AND J. ZECHNER. "The Geography of Equity Listings: Why Do European Companies List Abroad?" *Journal of Finance* (Forthcoming).

POWNALL, G., AND K. SCHIPPER. "Implications of Accounting Research for the SEC's Consideration of International Accounting Standards for U.S. Securities." *Accounting Horizons* 13 (1999): 259–80.

REESE, W., AND M. WEISBACH. "Protection of Minority Shareholder Interests, Cross-listings in the United States, and Subsequent Equity Offerings." *Journal of Financial Economics* (Forthcoming).

SERVAES, H. "Tobin's *Q* and the Gains from Takeovers." *Journal of Finance* 46 (1991): 409–19.

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