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Do countries matter for voluntary disclosure? Evidence from cross-listed firms in the US

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Abstract
This paper explores the likelihood and consequences of voluntary disclosure (proxied by management earnings forecasts) for a sample of 1005 cross-listed firms in the US from 40 countries over the period 1996–2005. Our study is grounded in a three-tiered conceptual framework that relies on insights from and implications of institutional theory, agency theory and bonding theory to explain the costs and benefits associated with voluntary disclosure. Consistent with institutional theory and agency theory, our results indicate that disclosure likelihood increases with the strength of cross-listed firms' home-country legal institutions, and is also influenced by US listing type, product market internationalization, and ownership structure. Further, our results show that voluntarily committing to US disclosure practice is associated with lower information asymmetry, which supports reputational bonding theory. Overall, our study provides a costs-and-benefits framework to understand voluntary disclosure practices in an international context. Our work also presents evidence that home-country institutions still matter when foreign firms migrate into the US financial market, which highlights the importance of country-level institution development.
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INTRODUCTION

The globalization of financial markets facilitates information flow across national boundaries: this, in turn, leads to convergence in corporate disclosure and governance practices (La Porta, Lopez-de-Silanes, & Shleifer, 2008). For example, Doidge, Karolyi, and Stulz (2007) argue that financial globalization decreases the importance of the home-country legal protection of minority shareholders. In this context, there has been a spirited debate as to whether and how a firm's home-country legal institutions still play a role in determining firm-level disclosure and governance practices once the firm has its own access to the global capital market (e.g., Durnev & Kim, 2005).

This study contributes to the above debate by investigating the determinants and consequences of voluntary disclosures by foreign firms cross-listed in the US. These firms from different countries face a similar information environment in the host country. Thus US cross-listed firms provide an ideal setting in which to examine the issue of whether and how alternative home-country legal

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institutions influence the likelihood and consequences of voluntary disclosures. Better understanding of this issue is important, given that the inflow of foreign capital via cross-listing is a significant contributor to US capital market growth (Foerster & Karolyi, 1993, 1999).[1]

Our proxy for voluntary disclosure is management earnings forecasts (i.e., forecasts related to a firm's anticipated earnings per share). This proxy offers several advantages. First, in the US, management earnings forecasts are one of the primary vehicles through which managers voluntarily disclose private information to outside stakeholders (Healy & Palepu, 2001). Second, management earnings forecasts are a more direct measure of managers' beliefs about future firm performance than other measures, such as analysts' scores (e.g., CIFAR) or self-constructed measures (Ajinkya, Bhojraj, & Sengupta, 2005). Third, since the timing of management forecasts is known, it is possible to evaluate whether changes in economic consequences occur *following* the forecasts. Finally, earnings forecasts by managers of cross-listed firms are released to US investors and are expressed in US dollars. This last feature allows us to test whether home-country institutions carry over to the host country when foreign firms migrate into the US financial market.

Our study is grounded in a three-tiered conceptual framework that relies on insights from and implications of institutional theory, agency theory and bonding theory to explain the costs and benefits associated with voluntary disclosure. First, as will be further elaborated in the next section, institutional theory implies that cross-listed firms face dual pressures from both host and home countries (Hillman & Wan, 2005; Scott, 2001). Given that management earnings forecasts are quite common in the US, normative institutional contagion or governance spillover effects can encourage foreign firms listed in the US to voluntarily converge with US practices (Oxelheim & Randoy, 2005). However, home-country regulative institutions take a long time to form, and thus influence the costs and benefits of voluntary disclosure. For example, managers of firms from weak legal institutions have greater private benefits of control and may face higher proprietary costs when increasing transparency than those from strong legal institutions (Doidge et al., 2007; Durnev, Errunza, & Molchanov, 2009). We therefore expect that firms that originate from countries with a strong legal regime will have a greater propensity to issue forecasts (Hypothesis 1).

Second, agency theory suggests that private benefits of control are more prominent when corporate ownership is concentrated (Shleifer & Vishny, 1997). Agency theory helps to explain that firms with concentrated ownership are more likely to stay opaque, because increased monitoring, litigation, and reputational costs associated with disclosure curb controlling shareholders' opportunities to extract private benefits. Therefore we expect that firms that originate in countries with concentrated ownership are less likely to issue forecasts (Hypothesis 2). Third, institutional duality theory implies that host-country institutions also shape managerial behavior (Hillman & Wan, 2005). We therefore expect that a firm's type of US listing – for example, major US stock exchanges vs over-the-counter (OTC)/Portal – influences its likelihood to issue forecasts (Hypothesis 3). Fourth, the product market exposure effect implies that cross-listed firms would benefit in the product market by releasing more information to the US market (Saudagaran, 1988), so we expect that product market internationalization translates into a higher likelihood that managers will issue forecasts (Hypothesis 4).

Finally, although reputational bonding theory suggests that cross-listed firms would garner benefits by voluntarily committing to US disclosure practices (Siegel, 2005), institutional theory implies that the benefits are conditional on reporting quality, which is influenced by home-country institutions (Ball, Kothari, & Robin, 2000). For these reasons, gauging the economic consequences of voluntary disclosure for cross-listed firms becomes an empirical issue (Hypothesis 5).

Our sample comprises 1005 foreign firms from 40 countries that had their shares cross-listed in the US during the sample period 1996–2005. Our results are broadly consistent with the predictions of our hypotheses, that firms with stronger home-country institutions (Hypothesis 1), lower home-country-level ownership concentrations (Hypothesis 2), listings on a major US stock exchange (Hypothesis 3), and greater levels of product internationalization (Hypothesis 4) are associated with a higher likelihood of releasing management earnings forecasts. Additional analyses reveal that cross-listed firms are more likely to disclose management earnings forecasts when their level of institutional ownership is high. Finally, we find that cross-listed firms that voluntarily release management earnings forecasts are associated with lower information asymmetry, as proxied by analyst forecast dispersion (AFD, Hypothesis 5).

This study contributes to the extant literature in the following ways. First, we add to the debate on whether home-country institutions still matter in today's global business environment. Some scholars argue that the importance of home-country legal institutions decreases when firms have access to global capital markets (Doidge et al., 2007; Durnev & Kim, 2005; Francis, Khurana, & Pereira, 2005; Sundaram & Logue, 1996). For instance, Sundaram and Logue (1996) posit that home-country attributes do not matter in the valuation of foreign firms once they cross-list in the US, suggesting that home-country institutions do not carry over to the US market. Francis et al. (2005) also claim that voluntary disclosure practices seem to function *independently* of country-level factors. By contrast, other studies record that home-country legal institutions play a key role in determining firms' growth options (Tong, Alessandri, Reuer, & Chintakananda, 2008), employment policies (Wu, Lawler, & Yi, 2008), choice of local partners (Luo, Chung, & Sobczak, 2009), corporate profitability (McGahan & Victor, 2010), the use of bond covenants (Qi, Roth, & Wald, 2011), and the quality of financial reporting (Ball et al., 2000). Our study supports the latter view: we find that home-country institutions have a significant bearing on cross-listed firms' incentives to converge fully to US disclosure practices. Our findings also demonstrate that cross-listing cannot be viewed as a substitute for weak home-country institutions, thus highlighting the importance of enhancing a country's institutions.[2]

Second, to our knowledge, our study is one of the few to investigate the likelihood and consequences of voluntary disclosure for cross-listed firms: The extant literature focuses either on the cross-listed firms' *financial reporting quality* or on the *mandatory* reporting requirements by US exchanges and other regulatory bodies (e.g., Bailey, Karolyi, & Salva, 2006; Lang, Raedy, & Wilson, 2006). Moreover, most prior studies on voluntary disclosures adopt a country-specific perspective, with a strong emphasis on US domestic firms (Ajinkya et al., 2005; Healy & Palepu, 2001). However, because the US disclosure environment is already rich, voluntary commitments to increased levels of disclosure have marginal impact (Bailey et al., 2006). In addition, the uncritical generalization of results from the US context to the cross-listing context is questionable.

Finally, Hail and Leuz (2006) find that the cost of raising equity capital via cross-listing is lower when cross-listed firms are from countries with strong legal regimes. In a related vein, Eleswarapu and Venkataraman (2006) find that US cross-listed firms originating from French civil law countries and countries with lower judicial efficiency face higher trading costs. Our findings suggest that less transparent disclosures by these cross-listed firms may actually underlie their higher US trading costs, thus reflecting information asymmetry between managers and investors.

CONCEPTUAL FRAMEWORK AND THEORY

Scott (2001: 48) defines an institution as the "cultural-cognitive, normative, and regulative elements that, together with associated activities and resources, provide stability and meaning to social life". Hence institutional theory postulates that institutions that include rules of law, regulations, government, and professional associations delineate what is socially legitimate in different institutional environments, and ultimately influence managers' decision-making (Suchman, 1995). Extending institutional theory, the theory of institutional duality provides a framework for examining how institutions affect international firms (Hillman & Wan, 2005; Luo et al., 2009). More specifically, Hillman and Wan (2005) state that the subsidiaries of multinational enterprises face dual pressures from the host and home countries. For example, foreign firms feel obliged to converge to host-country norms to legitimize their social roles and enhance their performance in the host country (Suchman, 1995). Prior work also shows that the host country's institutional environment is essential for partner selection in joint ventures (Roy & Oliver, 2009), and for organizational diversification (Jandik & Kali, 2009). Given that releasing management earnings forecasts is a normal practice in the US, there may be an institutional contagion or governance spillover effect on foreign firms listed in the US as they strive to gain legitimacy in that market (Oxelheim & Randoy, 2005). This is consistent with the normative mechanism of institutional theory.

However, multinational enterprises' home-country institutions, such as legal origin, investor protection and judicial efficiency, are embedded in their cultures, underlie their development, and thus are critical in their decision-making. Therefore the influence of home-country institutions on organizations, especially in terms of the costs and benefits of voluntary disclosure, is likely to be permanent (Kostova, 1999). For these reasons, it is vital to consider the institutional environment for firms cross-listed in the US, not only in their host countries, but also in their home countries.

Agency theory also guides our investigation. La Porta, Lopez-de-Silanes, and Shleifer (1999) record that, except for the US, ownership structures are concentrated around the world. Ownership concentration has implications for voluntary disclosure. Information asymmetries exist because controlling shareholders are typically better informed than minority shareholders about expected future cash flows. Increased disclosure can lead to higher monitoring costs from outside stakeholders, such as analysts and institutional owners, and higher reputational costs when disclosures are biased (Cumming & Walz, 2010; Healy & Palepu, 2001). Therefore, when private benefits of control are high, controlling shareholders may be less likely to voluntarily disclose information to minority shareholders (Shleifer & Vishny, 1997). However, large owners' concentrated holdings also imply that the benefits of disclosure (e.g., reduced cost of capital) are more important to concentrated owners than to other shareholders. Therefore an alternative perspective inspired by agency theory is that voluntary disclosure is a credible mechanism to reduce financing costs in the presence of domestic managerial agency costs (Barnea, Haugen, & Senbet, 1985: 38).

Finally, our work is related to bonding theory. Siegel (2005) proposes two mechanisms of bonding: legal bonding and reputational bonding. The legal bonding theory posits that the benefits of US cross-listing stem from stricter investor protection and more stringent enforcement of stock market regulations by the US Securities and Exchange Commission (SEC) (Coffee, 1999; Stulz, 1999). However, other studies (e.g., La Porta, Lopez-de-Silanes, Shleifer, & Vishny, 2000; Siegel, 2005) argue that US legal enforcement against foreign firms has been weak. Reputational bonding theory conjectures that foreign firms in the US can voluntarily bond themselves not to expropriate minority shareholders by establishing their reputation (Siegel, 2005, 2009). Consistent with reputational bonding theory, cross-listed firms can signal their own assurances of fair treatment to US minority shareholders by aligning their voluntary disclosure practices with US norms.

The integration of institutional theory, agency theory and bonding theory helps us better understand the trade-offs between costs and benefits associated with voluntary disclosure offered by US cross-listed firms. Voluntary disclosure can lead to many benefits, including reductions in information asymmetries and cost of capital, signaling management talent, reducing litigation costs, and enhancing firm value (Healy & Palepu, 2001). Reputational bonding theory also suggests that by providing voluntary disclosures to US investors, cross-listed firms signal their commitment to improving their corporate governance, with salient benefits to be derived in this context. By contrast, voluntary disclosures are related to potential costs, including proprietary, litigation and monitoring costs to financial analysts and institutional investors (Healy & Palepu, 2001). Hence cross-listed firms strategize their voluntary disclosure after considering the cost–benefit trade-offs, with disclosure being an outcome only when benefits outweigh costs.

HYPOTHESES DEVELOPMENT

Home-Country Legal Institutions

As discussed in the preceding section, institutional theory implies that home-country legal institutions affect cross-listed firms' likelihood of providing voluntary disclosure. In addition, country-specific institutions such as legal origin have pervasive economic consequences, and affect many aspects of a firm's business environment, including access to financing, ownership dispersion, contract enforcement, and information risk (La Porta et al., 2008; Scott, 2001). These institutional factors then determine the supply of information. Also, home-country institutions can influence the costs and benefits associated with improving disclosure transparency and corporate governance (Doidge et al., 2007). For instance, proprietary costs are related to the costs of revealing proprietary information to other stakeholders such as competitors and government (Dye, 1986). In that regard, Durnev et al. (2009) argue that in countries with insecure property rights, greater transparency can increase the likelihood of government expropriation. An important implication of the above discussion is that the cost–benefit trade-offs associated with voluntary disclosures are more favorable to firms from countries with stronger legal institutions. Therefore we hypothesize the following (in alternative form).

> **Hypothesis 1**: *Ceteris paribus*, the likelihood that cross-listed firms in the US will issue management earnings forecasts to US investors is positively associated with the strength of their home-country legal institutions.

Ownership Structure

Leuz (2006) examines the ownership structure of cross-listed firms in the US and shows that such firms exhibit more concentrated ownership than their US counterparts. Similarly, Ayyagari and Doidge (2010) show that foreign firms exhibit reduced ownership concentration once they cross-list in the US, but there is no movement to a fully dispersed ownership structure. One may therefore expect that agency conflicts arise between large controlling shareholders of cross-listed firms and US minority shareholders (Doidge et al., 2009). As discussed before, agency costs influence the trade-off between costs and benefits associated with voluntary disclosures. In addition, previous empirical studies find that private benefits of control associated with agency conflicts influence cost of capital (Boubakri, Guedhami, & Mishra, 2010), IPO underpricing (Boulton, Smart, & Zutter, 2010), financial reporting incentives and quality (Kim & Yi, 2006), and stock price informativeness (Gul, Kim, & Qiu, 2010).

Management earnings forecasts can be verified *ex post* upon the release of actual earnings. Given that foreign firms are subject to the risk of US class action securities litigation related to Rule 10b-5, litigation and reputational costs can limit the opportunities of large controlling owners to extract private benefits of control (Coffee, 2002). Therefore agency costs associated with ownership concentration can discourage large controlling owners from voluntarily disclosing firm-specific information. The above reasoning leads to our second hypothesis (in alternative form).

> **Hypothesis 2**: *Ceteris paribus*, the likelihood that cross-listed firms in the US will issue management earnings forecasts to US investors is inversely associated with the extent of ownership concentration.

Listing Types in the US Security Markets

Institutional theory implies that the financial reporting incentives of cross-listed firms are affected by host-country institutions (Hillman & Wan, 2005). Cross-listed firms in the US can choose to have their shares traded on major exchanges (NYSE, AMEX and NASDAQ) or on the OTC market, as well as via private placements under SEC Rule 144A (Portal). Neither OTC nor Portal firms are required to file Form 20-F with the SEC, or reconcile their financial statements in accordance with US Generally Accepted Accounting Principles (GAAP); they also do not need to incur legal bonding costs by upgrading their disclosures (Coffee, 2002). In contrast, foreign firms listed on a major US exchange are required to file Form 20-F with the SEC and to reconcile their financial statements to US GAAP. They are also subject to enforcement by the SEC and US courts (Coffee, 1999). Therefore foreign firms face different institutional environments in the US according to their listing type, which must affect the likelihood of their issuing voluntary disclosures.

On the one hand, evidence shows that foreign firms listed on the OTC/Portal experience lower cross-listing premiums (Doidge, Karolyi, & Stulz, 2004) and exhibit poorer accounting quality (Lang, Raedy, & Yetman, 2003b) than foreign firms listed on the major US exchanges. Moreover, firms listing on major exchanges often raise capital. One can therefore predict that OTC/Portal firms are less likely to issue management earnings forecasts, since they are not required to bond themselves completely to US governance practices. On the other hand, OTC and Portal firms face lower litigation costs than US exchange-listed firms, and empirical evidence suggests that voluntary forward-looking disclosure is more prevalent in less litigious environments (Baginski, Hassell, & Kimbrough, 2002). Hence one can expect that OTC/Portal listed foreign firms are more likely to issue earnings forecasts. Given these two opposing predictions, we test the following hypothesis with no directional prediction.

> **Hypothesis 3**: *Ceteris paribus*, the likelihood that cross-listed firms in the US will issue management earnings forecasts to US investors is associated with their listing types, namely, whether they are listed on the major US stock exchanges (NYSE/AMEX/NASDAQ) or the non-exchange markets (OTC/Portal).

Product Market Internationalization

Global product market diversification can have a significantly positive effect on firm value (e.g., Gande, Schenzler, & Senbet, 2009). Hence a firm's decision to cross-list its shares on a foreign stock exchange can be motivated by product market internationalization, because cross-listing increases corporate visibility and facilitates marketing efforts in the host country (Saudagaran, 1988). Specifically, Saudagaran (1988) finds that firms are more likely to list abroad when they have larger ratios of foreign to total sales. In a similar fashion, product market internationalization can influence cross-listed firms' voluntary disclosure of forward-looking information to US investors, for two reasons. First, prior theoretical work advocates that both financial market and product market factors influence a firm's voluntary disclosure practices (Dye, 1986; Evans & Sridhar, 2002). Khanna, Palepu, and Srinivasan (2004) observe that product market internationalization is associated with higher overall corporate disclosure scores. Given that management earnings forecasts are a common US disclosure practice (Healy & Palepu, 2001), cross-listed companies can face higher costs of doing business if their disclosures do not conform to this US norm.

Second, institutional theory suggests that US firms' corporate governance and disclosure practices can generate a governance spillover effect on foreign firms in the US (Scott, 2001). This notion is well supported by studies such as Oxelheim and Randoy (2005) and Southam and Sapp (2010). For example, Southam and Sapp (2010) show that Canadian firms cross-listed in the US have to offer their chief executive officers a compensation package that is closer to US levels. Here, cross-listed firms' product market internationalization can proxy for the governance spillover effect (Crawley, Ke, & Yu, 2009). Both the product market exposure and governance spillover effects suggest that product market internationalization encourages cross-listed firms to disclose forward-looking information to US investors. This reasoning leads to our fourth hypothesis (in alternative form).

Hypothesis 4: *Ceteris paribus*, the likelihood that cross-listed firms in the US will issue management earnings forecasts to US investors is positively associated with their level of product market internationalization.

Effects of Voluntary Disclosure on Information Asymmetry

From theoretical work by Verrecchia (2001), we can infer that voluntary disclosure reduces information asymmetry by lowering estimation risk and adverse selection costs. US-based empirical evidence largely supports this view (Healy & Palepu, 2001). Reputational bonding theory also suggests that foreign firms benefit from voluntarily committing themselves to more transparent reporting (Siegel, 2005).

However, voluntary disclosure may not be associated with lower information asymmetry for cross-listed firms. First, cross-listed firms in our sample exhibit concentrated ownership structures, which may influence the quality of financial disclosure. Specifically, Leuz, Nanda, and Wysocki (2003) find that firms with concentrated ownership are associated with more earnings management. Second, our sample firms originate from countries with different institutions, and institutional theory suggests that their reporting quality is influenced by home-country institutions (Ball et al., 2000; Tong et al., 2008). Lang et al. (2006) indicate that the financial reporting quality of cross-listed firms is lower than that of their US counterparts, and the negative association is stronger for firms from countries with weaker legal institutions. Cumming and Walz (2010) also argue that, for private equity funds, the information content of disclosure decreases in countries with weak legal systems. These findings imply that voluntary disclosure may not be credible for these firms. Empirically, Bailey et al. (2006) find that increased public disclosures (e.g., annual earnings announcements) by foreign firms cross-listed in the US indeed translate into higher absolute return and trading volume.

Given these conflicting predictions on the impact of disclosure on the information asymmetry, we test the following hypothesis with no directional prediction.

Hypothesis 5: *Ceteris paribus*, for cross-listed firms in the US, there is an association between the issuance of management earnings forecasts to US investors and the level of information asymmetry.

RESEARCH DESIGN

Sample and Data Sources

To construct our sample of foreign firms that are cross-listed in the US, we obtain a complete list of depositary receipts from the Bank of New York website.[3] This list provides information about the names, listing dates, country of origin, and exchanges (i.e., NYSE, AMEX, NASDAQ, OTC, or Portal) of each American Depository Receipt (ADR) as of 2005. We obtain the information on direct-listing Canadian and Israeli firms from the NYSE, NASDAQ, AMEX, OTCBB, and Pink Sheets websites.[4] Although direct-listing and ADR firms have different listing procedures, they have similar financial reporting requirements (Lang, Lins, & Miller, 2003a). More importantly, releasing management earnings forecasts to US investors is not required by the SEC (i.e., it is a voluntary disclosure practice for both direct-listing and ADR firms). Therefore we follow previous cross-listing literature

Table 1 Description of the data set

Panel A: Sample construction process	
Cross-listed firms in the US as of 2005	2050
Firm–year observations of cross-listed firms from 1996 to 2005	11,284
Less	
Firm-level financial variables are unavailable	3564
Observations with extreme values	372
	7348
Less	
Analyst data are unavailable from IBES	3634
Total firm–year observations	3714
Forecasting firm–year observations	616
Non-forecasting firm–year observations	3098
Forecasting firms	252
Non-forecasting firms	753
Total firms	1005

	1	2	3	4	5 or more	Total
Panel B: Distribution of the frequency of earnings forecasts made by a firm in the period 1996–2005						
Number of firms	79	29	27	14	103	252

	Point	Range	Open-ended	Qualitative	Total
Panel C: Distribution of the precision of earnings forecasts made by firms in the period 1996–2005					
Number of forecasts	164	291	56	105	616

(e.g., Bailey et al., 2006; Doidge et al., 2004; Foerster & Karolyi, 1999; Lang et al., 2003a) and include both direct-listing and ADR firms in our sample. To avoid survivorship bias, we include firms that were later delisted. We classify firms that were listed on both OTC markets and NYSE/AMEX/NASDAQ at different points as being listed on NYSE/AMEX/NASDAQ. Following these selection procedures, we obtain a sample of 2050 cross-listed firms as of December 2005.

As shown in Panel A of Table 1, our initial 2050 sample firms represent 11,284 firm–year observations between 1996 and 2005. Firm-specific data such as foreign sales, industry classification, and auditor identity are collected primarily from Worldscope, and are supplemented from firms' annual reports, Form 20-F, and websites when necessary. Country-level data are extracted from La Porta, Lopez-de-Silanes, and Shleifer (2006). After dropping observations with missing firm-level variables and with extreme values, we obtain 7348 firm–year observations. This sample is then matched with IBES data on financial analyst following.[5] This procedure yields our final sample of 3714 final observations representing 1005 firms from 40 countries.[6]

We extract management earnings forecasts data for the period 1996–2005 from the Corporate Investor Guidelines (CIG) database, maintained by First Call. In defining forecasting years, a firm that issues a single forecast and one that releases multiple forecasts are treated the same. For firms issuing multiple forecasts, we select one forecast that is closest to the actual earnings announcement. We use AFD as our proxy for information asymmetry. It is worth noting that every forecast observation is matched with AFD immediately following management earnings forecasts. In this way, we can better capture the effect of management earnings forecasts on information asymmetry, and alleviate the concern of causality. Among our final 3714 firm–year observations (1005 firms), 616 observations representing 252 firms belong to the forecasting years, and 3098 observations representing 753 firms belong to the non-forecasting years.

Panel B of Table 1 reports the forecast frequency for the sample of 252 firms disclosing forecasts during 1996–2005. Panel B shows that 79 firms made only one forecast over the 10-year period, whereas 103 firms released five or more forecasts. Panel C of Table 1 shows the distribution of forecast precision. Among the 616 forecasting observations, 164 are point forecasts, 291 are range forecasts, 56 are open-ended, and 105 are qualitative.

Empirical Models

To test Hypotheses 1 to 4, we estimate the following probit regression, which links the likelihood of managers issuing earnings forecasts with our four test variables and controls:

$$\begin{aligned}\Pr(MF) = {} & \alpha_0 + \alpha_1 LEGAL + \alpha_2 OWNERSHIP \\ & + \alpha_3 LISTTYPE + \alpha_4 FORSALES \\ & + \sum_k \alpha_k CONTROL^k + (Year\ Dummies) \\ & + (Industry\ Dummies) + error \qquad (1)\end{aligned}$$

Empirical definitions of all variables are as provided in the Appendix. In the above equation, the dependent variable, Pr(*MF*), represents the *ex ante* probability that cross-listed firms will issue management earnings forecasts to US investors, which is *ex post* coded as 1 for cross-listed firms that issued the forecasts during the fiscal period, and 0 otherwise.

The term *LEGAL* is our test variable for Hypothesis 1, and represents the efficacy of the legal institution of a cross-listed firm's home country. It is measured using three country-level proxies:

(1) *COMMON*, which takes the value of 1 if a cross-listed firm is from an English common law country, and 0 otherwise[7];
(2) *ANTI-DIRECTOR*, which is an index that aggregates six different shareholder rights and ranges from 0 to 6, with 6 being the highest level of investor protection; and
(3) *JUDICIAL*, which is an assessment score of the efficiency and integrity of a country's legal environment and ranges from 0 to 10, with 10 as the highest standard.

Hypothesis 2 predicts an inverse relation between Pr(*MF*) and ownership concentration, as proxied by a *country-level* variable, *OWNERSHIP*, which represents the average percentage of common shares held by the top three shareholders in the 10 largest non-financial, privately owned domestic firms in a given country. To obtain further insight into the effect of ownership structure on the likelihood of disclosing forward-looking information, we also consider two additional *firm-level* ownership variables, *OWNCON* and *INST*. Here *OWNCON* is a firm-level variable that represents the extent to which a cross-listed firm has a concentrated ownership structure: this variable is measured as the percentage of cash flow rights held by the largest shareholder in the forecast year, as defined in Claessens, Djankov, Fan, and Lang (2002). *INST* is a firm-level variable representing the percentage of a cross-listed firm's common shares held by institutions.[8] Consistent with Hypothesis 2, we expect an inverse relation between Pr(*MF*) and *OWNCON*. We predict a positive relation between Pr(*MF*) and *INST*.

Hypothesis 3 focuses on whether Pr(*MF*) is associated with the type of exchange used for cross-listing (*LISTTYPE*). The variable *LISTTYPE* equals 1 for a firm listed on the major US exchanges (NYSE/AMEX/NASDAQ), and 0 for a firm listed on OTC/Portal.[9]

Hypothesis 4 is concerned about whether Pr(*MF*) is greater for firms with more product market internationalization, and smaller for firms with less product market internationalization, as proxied by *FORSALES* (i.e., the dollar values of foreign sales, deflated by total sales). Hypothesis 4 translates into a positive coefficient for *FORSALES*.[10] We use foreign sales in total rather than country-specific foreign sales, since firms do not disclose country-by-country breakdowns of foreign sales.

We include in our probit regression a set of control variables that are deemed to influence Pr(*MF*): *REGFD*, *LITIGATE*, *BIG4*, *SIZE*, *LOSS*, *ANALYST*, *ANALYST* × *LISTTYPE*, *GROWTH*, and *NEWS*. Here *REGFD* is an indicator variable that equals 1 if the observation is related to the post-Regulation Fair Disclosure period (after October 2000), and 0 otherwise. Consistent with Bailey, Li, Mao, and Zhong (2003), we expect the number of forecasts to have increased in the post-Regulation Fair Disclosure period. The term *LITIGATE* is an indicator variable that equals 1 if a cross-listed firm pertains to the biotechnology (Standard Industrial Classification, or SIC, 2833–2836 and 8731–8734), computers (SIC 3570–3577 and 7370–7374), electronics (SIC 3600–3674), or retail (SIC 5200–5961) industries, and 0 otherwise. Evidence shows that the litigation risk is higher for firms active in these industries (Ajinkya et al., 2005; Skinner, 1994). We expect a positive sign for the *LITIGATE* coefficient.

The variable *BIG4* is an indicator variable that equals 1 if a cross-listed firm has its financial statement audited by one of the Big Four auditors, and 0 otherwise. Previous research provides evidence that Big Four auditors are associated with better disclosure (Lang & Lundholm, 1996), leading us to expect a positive coefficient. The variable *SIZE* is measured by the natural log of total sales in US dollars. Previous research provides mixed evidence about the relation between Pr(*MF*) and *SIZE*. Kasznik and Lev (1995) find that firm size is positively associated with the occurrence of management earnings forecasts. By contrast, larger firms can incur higher political costs (Watts & Zimmerman, 1990), which can lower the likelihood of issuing management forecasts. We therefore make no prediction on the sign of this coefficient. The term *LOSS* is an indicator variable that equals 1 for cross-listed firms that report a loss in the current period, and 0 otherwise. Prior research suggests that earnings are less value-relevant for loss-making firms (Hayn, 1995). Managers of loss-making firms also face greater uncertainty about future prospects, thus reducing their ability, and willingness, to forecast future earnings.

Here *ANALYST* represents the number of financial analysts following a firm. Lang and Lundholm (1996) show that firms with more analyst followings are associated with a higher quality of corporate disclosure. Therefore we predict a positive coefficient for *ANALYST*. Further, we include an interaction between *ANALYST* and *LISTTYPE*. Previous studies show that firms listing on the major US stock exchanges have a better information environment than those traded on OTC and Portal (Doidge et al., 2004). Hence the role of financial analysts may be particularly important for OTC/Portal firms, and we predict a negative coefficient for *LISTTYPE* × *ANALYST*. The variable *GROWTH* is sales growth over the past two years, and we expect it to have a negative coefficient. Finally, *NEWS* is an indicator variable that equals 1 if the current-period earnings per share is greater than that for the previous period, and 0 otherwise. Consistent with Skinner's (1994) finding that firms with bad news are more likely to issue forecasts, we predict a negative coefficient for *NEWS*.

To test Hypothesis 5, the effect of management earnings forecasts on information asymmetries between managers and outside investors, we estimate the following regression:

$$DISP = \beta_0 + \beta_1 MF + \sum_k \beta_k CONTROL^k + (Year\ Dummies) + (Industry\ Dummies) + error \qquad (2)$$

The Appendix provides detailed definitions of all variables considered in Eq. (2). The dependent variable, *DISP*, represents the standard deviation of analyst forecasts following the issuance of management earnings forecasts.[11] For non-forecasting firms, we assume that management earnings forecasts occur 90 days (the median number of days in the forecasting samples) before the last day of the accounting period to measure AFD.[12] Our main test variable is *MF*, which is a dummy variable (equal to 1 if a firm issues a forecast, and 0 otherwise).

We also include a set of control variables – *LEGAL*, *LIQUIDITY*, *LISTTYPE*, *PROFIT*, *LIABILITY*, *SIZE*, and *MKBK* – that are expected to influence information asymmetry. We expect the strength of the home country's legal regime (*LEGAL*) to have an inverse relation with information asymmetry, because firms from countries with stronger legal institutions are associated with lower information asymmetry (La Porta et al., 2008). Here *LIQUIDITY* is the average ratio of the dollar value of shares traded as a percentage of country gross domestic product. We expect a negative coefficient for it. Foreign firms listing on major US stock exchanges are associated with better information environments (Bailey et al., 2006). We therefore expect a negative sign for the coefficient of *LISTTYPE*. The variable *PROFIT* is operating income deflated by total assets. We predict a negative coefficient for this variable (Lang et al., 2003b). The variable *LIABILITY* is the ratio of total liabilities to total assets, and *SIZE* is the log of total sales at the beginning of the period. We predict positive coefficients for both variables, following prior studies (e.g., Kothari, Li, & Short, 2009). Finally, *MKBK* is the market value of equity divided by the book value of equity. Firms with higher *MKBK* ratios are viewed as more successful, and carry lower risks. We expect a negative coefficient for *MKBK* (e.g., Kothari et al., 2009). We include *Industry Dummies* and *Year Dummies* to control for year and industry effects.

Descriptive Statistics

Table 2 reports descriptive statistics for country- and firm-specific variables for all years, forecasting years, and non-forecasting years, and provides the results of univariate tests for the mean and median differences for each variable between forecasting and non-forecasting years. As shown in Table 2, 56% of our sample firms are from English common law countries. Likewise, 19% of our sample firms report a loss, 39% report good news earnings, and foreign sales account for 37% of the overall sales of our sample firms. In addition, 93% of our sample firms use Big Four auditors. The mean analyst

Table 2 Descriptive statistics and univariate comparisons between management forecast years and non-forecast years

	All years	Forecast years	Non-forecast years	*t*-test	Wilcoxon *z*-test
Sample size	3714	616	3098		
Country-level variables					
COMMON	0.56	0.87	0.5	17.85***	17.13***
	1	1	0		
ANTI-DIRECTOR	3.92	4.47	3.81	11.78***	13.46***
	4	5	4		
JUDICIAL	8.86	9.25	8.78	7.43***	4.38**
	9.25	9.25	9.25		
LIQUIDITY	0.61	0.57	0.61	−1.72*	−5.54***
	0.58	0.58	0.58		
OWNERSHIP	0.29	0.14	0.32	−9.29***	−9.18***
	0	0	0		
Firm-level variables					
DISP	0.23	0.13	0.26	−8.10***	−10.98***
	0.13	0.08	0.15		
LISTTYPE	0.66	0.72	0.64	3.47***	3.46***
	1	1	1		
ANALYST	5.31	9.57	4.47	20.73***	21.03***
	3.00	8	2		
LITIGATE	0.18	0.42	0.14	16.77***	16.17***
	0	0	0		
FORSALES	0.37	0.54	0.34	13.14***	12.15***
	0.30	0.66	0.25		
PROFIT	0.06	0.05	0.06	1.1	1.72**
	0.05	0.06	0.05		
LIABILITY	0.56	0.47	0.57	−9.11***	−9.73***
	0.56	0.45	0.58		
SIZE	6.19	5.68	6.41	−13.17***	−14.68***
	6.30	5.76	5.98		
LOSS	0.19	0.38	0.39	−0.69	−0.69
	0	0	0		
GROWTH	0.23	0.23	0.23	0.03	5.67***
	0.09	0.14	0.09		
NEWS	0.39	0.45	0.38	3.38***	3.39***
	0	0	0		
MKBK	1.20	1.77	1.08	12.01***	11.03***
	0.84	1.15	0.78		
BIG4	0.93	0.91	0.94	−2.72***	−2.71***
	1	1	1		

All variables are defined in the Appendix. The mean value for each variable is provided in the top row and the median value in the bottom row. Here *, **, and *** indicate significance at the 10, 5 and 1% levels (two-tailed test), respectively.

coverage for our sample is 5.31, while it is 9.57 for the forecasting group and 4.47 for the non-forecasting group.

Results of *t*- and *z*-tests show that the mean and median differences between forecasting and non-forecasting years are significant for most variables. With respect to country-level variables, we find that cross-listed firms are more likely to issue management earnings forecasts when they are from countries with common law origins (*COMMON*), stronger investor protection (*ANTI-DIRECTOR*), higher judicial efficiency (*JUDICIAL*), and more diffuse ownership structure (*OWNERSHIP*). This is consistent with Hypotheses 1 and 2. With respect

Table 3 Probability of management earnings forecast

	Sign	(1)	(2)	(3)	(4)	(5)	(6)
		Probit	Weighted	Probit	Weighted	Probit	Weighted
Intercept	±	−0.401	−0.239	−0.132	−0.234	−0.437	−0.110
		(0.13)	(0.37)	(0.67)	(0.45)	(0.22)	(0.77)
COMMON (H1)	+	0.614***	0.490***				
		(0.00)	(0.00)				
ANTI-DIRECTOR (H1)	+			0.123***	0.056*		
				(0.00)	(0.10)		
JUDICIAL (H1)	+					0.117***	0.073**
						(0.00)	(0.02)
OWNERSHIP (H2)	−	−0.253***	−0.144***	−0.179***	−0.102	−0.267***	−0.130
		(0.00)	(0.00)	(0.00)	(0.25)	(0.04)	(0.13)
LISTTYPE (H3)	±	0.492***	0.505***	0.479***	0.500***	0.492***	0.511***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
ANALYST	+	0.156****	0.153***	0.167***	0.162***	0.175***	0.166***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
LISTTYPE × *ANALYST*	−	−0.110***	−0.108***	−0.116***	−0.113***	−0.119***	−0.114***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
FORSALES (H4)	+	0.006***	0.006***	0.007***	0.006***	0.006***	0.006***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GROWTH	−	−0.001**	−0.001**	−0.001**	−0.001**	−0.001**	−0.001**
		(0.04)	(0.04)	(0.05)	(0.04)	(0.05)	(0.05)
BIG4	+	−0.150	−0.173	−0.171	−0.195*	−0.205*	−0.208*
		(0.18)	(0.13)	(0.12)	(0.08)	(0.07)	(0.07)
SIZE	±	−0.320***	−0.316***	−0.391***	−0.379***	−0.422***	−0.389***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
LOSS	−	−0.528***	−0.552***	−0.517***	−0.556***	−0.534***	−0.562***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
NEWS	−	0.204***	0.198***	0.190***	0.189***	0.200***	0.190***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
RegFD	+	−0.099	−0.096	−0.064	−0.065	−0.051	−0.061
		(0.11)	(0.13)	(0.30)	(0.29)	(0.40)	(0.32)
LITIGATE	+	0.758***	0.771***	0.735***	0.750***	0.665***	0.716***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Pseudo R^2		0.23	0.24	0.22	0.23	0.22	0.23
No. of observations		3714	3714	3714	3714	3714	3714

This table reports probit regression and country-weighted probit regression (i.e., weighted) results of the determinants of management earnings forecasts. The dependent variable is *MF*, with a value of 1 for forecasting firms, and 0 otherwise. All other variables are defined in the Appendix. Coefficient estimates are provided in the top row and *p*-values in the bottom row. Here *, ** and *** indicate significance at the 10, 5 and 1% levels (two-tailed test), respectively.

to firm-specific control variables, we find that cross-listed firms are more likely to issue management earnings forecasts when their shares are listed on major US exchanges (i.e., NYSE/AMEX/NASDAQ) as opposed to the OTC/Portal (*LISTTYPE*). Moreover, management earnings forecasts are more common for firms that have greater product market internationalization (*FORSALES*), which is consistent with Hypothesis 4. We also find that firms with greater analyst followings and operating in industries with higher litigation risk (*LITIGATE*) are more likely to issue management earnings forecasts. Further, forecasting firms are smaller in size (*SIZE*) and more likely to experience a loss (*LOSS*) and to have good earnings news (*NEWS*) than non-forecasting firms. Finally, we find that forecasting firms are associated with lower information asymmetry (*DISP*).

RESULTS OF MAIN REGRESSIONS

Table 3 reports the results of our probit regressions in Eq. (1). To alleviate concerns about potential problems that can arise from the unequal distribution of sample firms across different countries, we also estimate Eq. (1) by applying the weighted least squares (WLS) procedure with an equal weight assigned to each sample country (Choi, Kim, Liu, & Simunic, 2009).[13] In estimating Eq. (1), we include the three proxies for legal regimes (i.e., *COMMON*, *ANTI-DIRECTOR* and *JUDICIAL*) one by one because they are highly correlated with each other. In Table 3, columns 1, 3 and 5 report the results by estimating probit models; in columns 2, 4 and 6, we present the results adopting country-weighted probit models. Unless otherwise mentioned, *p*-values are adjusted using standard errors corrected for clustering at the firm level (Gow, Ormazabal, & Taylor, 2010).[14]

As shown in Table 3, the coefficients of *COMMON*, *ANTI-DIRECTOR* and *JUDICIAL* are significantly positive across all six models. This evidence is consistent with Hypothesis 1, and suggests that US cross-listed firms originating from countries with English common law (*COMMON*), stronger investor protection (*ANTI-DIRECTOR*), and more impartial judicial systems (*JUDICIAL*) are more likely to issue earnings forecasts to US investors.

Consistent with Hypothesis 2, the coefficient of *OWNERSHIP* is negative across all six columns of Table 3, and significant in four out of six columns, thus suggesting that cross-listed firms from countries with high ownership concentration are less likely to disclose management earnings forecasts to US investors. However, our measure of country-level ownership concentration may be noisy. Therefore we construct a reduced sample using *firm-level* ownership concentration data. The results using this reduced sample are discussed further in the next section.

The coefficient of *LISTTYPE* is significantly positive ($p<0.01$) across all six columns of Table 3. Hence foreign firms listed on major US stock exchanges are more likely to issue earnings forecasts than those listed on the OTC/Portal, an outcome that is consistent with greater bonding to US corporate governance practices (one perspective in Hypothesis 3). The coefficient of *FORSALES* is significantly positive across all columns in Table 3. This finding is consistent with Hypothesis 4, suggesting that firms with more product market internationalization are more likely to release earnings forecasts.

With respect to the estimated coefficients on our control variables, the following are noteworthy. First, the coefficient of *LITIGATE* is positive and significant in all six specifications. This is consistent with the view that managers of US cross-listed firms in the more litigious industries tend to use voluntary disclosure as a means to mitigate future litigation risk. Second, the coefficient of *SIZE* is significantly negative in all six specifications, which is in line with the view that large firms are less likely than small firms to disclose management forecasts, because large firms tend to have high agency costs and political costs that deter them from voluntarily disclosing more firm-specific information (Durnev & Kim, 2005; Watts & Zimmerman, 1990). Third, consistent with our expectations, the coefficient of *ANALYST* is significantly positive across all cases. Finally, the coefficient of *NEWS* is significantly positive across all cases, suggesting that good news firms are more likely to make earnings forecasts. Skinner (1994) posits that managers release good news forecasts to distinguish themselves from those doing less well, while disclosing bad news forecasts to reduce litigation concerns. Our results imply that the disclosure incentives of cross-listed firms are more associated with establishing reputations through signaling their good performance.[15]

The results of our ordinary least squares (OLS) regressions are presented in Table 4 for Eq. (2).[16] In columns 1, 3 and 5 of Table 4 we report the OLS results, and in columns 2, 4 and 6 we present the WLS results. As shown, the coefficients of *MF* are negative and significant in all six columns ($p<0.01$). This finding suggests that, for US cross-listed firms, the issuance of management forecasts leads to lowering information asymmetry (one perspective in Hypothesis 5). The coefficients of the three *LEGAL* variables are all negative and highly significant ($p<0.01$), consistent with the view that firms from countries with stronger home-country institutions are associated with lower information asymmetries. We also find that the information asymmetry is lower for firms from countries with more liquid stock markets, and for firms with higher market-to-book ratios (*MKBK*).

ROBUSTNESS CHECKS

Effect of Firm-Level Ownership

This section further examines the relation between Pr(*MF*) and *OWNERSHIP* using firm-level ownership data, as measured by *OWNCON* (largest shareholder's ownership stake) and *INST* (institutional ownership) rather than country-level ownership data.[17] We obtain firm-level data on *OWNCON*

Table 4 Effect of management earnings forecasts on analyst forecast dispersion

	Sign	(1)	(2)	(3)	(4)	(5)	(6)
		OLS	WLS	OLS	WLS	OLS	WLS
Intercept	±	0.137*	0.153*	0.091	0.109	0.182**	0.206**
		(0.10)	(0.07)	(0.29)	(0.21)	(0.04)	(0.02)
MF (H5)	±	−0.064***	−0.065***	−0.081***	−0.082***	−0.079***	−0.081***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
COMMON	−	−0.095***	−0.103***				
		(0.00)	(0.00)				
ANTI-DIRECTOR	−			−0.017***	−0.020***		
				(0.00)	(0.00)		
JUDICIAL	−					−0.026***	−0.029***
						(0.00)	(0.00)
LIQUIDITY	−	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
LISTTYPE	−	0.010	0.010	0.013	0.013	0.008	0.008
		(0.54)	(0.52)	(0.43)	(0.41)	(0.62)	(0.61)
PROFIT	−	−0.029	−0.025	−0.028	−0.024	−0.032*	−0.028
		(0.13)	(0.21)	(0.15)	(0.21)	(0.10)	(0.16)
LIABILITY	+	0.060*	0.058*	0.051	0.049	0.046	0.042
		(0.06)	(0.07)	(0.12)	(0.13)	(0.15)	(0.19)
SIZE	+	0.020**	0.019**	0.032***	0.031***	0.042***	0.044***
		(0.03)	(0.04)	(0.00)	(0.00)	(0.00)	(0.00)
MKBK	−	−0.019***	−0.019***	−0.021***	−0.020***	−0.020***	−0.019***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
SIC and year dummies		Yes	Yes	Yes	Yes	Yes	Yes
Pseudo R^2		0.10	0.10	0.10	0.10	0.10	0.11
No. of observations		2569	2569	2569	2569	2569	2569

This table reports the OLS and country WLS regression results of the effects of management earnings forecasts on AFD. The dependent variable is *DISP*, measured by the standard deviation of analyst forecasts following management earnings forecasts. All other variables are defined in the Appendix. Coefficient estimates are provided in the top row and *p*-values in the bottom row. Here *, ** and *** indicate significance at the 10, 5 and 1% levels (two-tailed test), respectively.

for 3126 firm–years and on *INST* for 2286 firm–years. Using this reduced sample data set, we re-estimate Eq. (1) with *LISTTYPE* excluded. We exclude the *LISTTYPE* variable in these estimations because most firm–years with firm-level ownership data for *OWNCON* and *INST* come from foreign firms listed on the major US exchanges.

For the sake of brevity, Table 5 reports the estimated coefficients of the test variables only. The coefficients of *COMMON*, *ANTI-DIRECTOR* and *JUDICIAL* are significantly positive across all columns ($p<0.01$); we also find that the coefficient of *FORSALES* is significantly positive across all columns ($p<0.01$). These results are consistent with Hypotheses 1 and 4. In terms of Hypothesis 2, as shown in columns 1, 2 and 3 of Table 5, when firm-level ownership concentration (i.e., *OWNCON*) is used in lieu of country-level ownership concentration (i.e., *OWNERSHIP*), the coefficient of *OWNCON* is significantly negative across all three cases ($p<0.01$), suggesting that the probability of managers issuing earnings forecasts to US minority investors is lower when ownership is highly concentrated in the hands of the largest shareholder.

Prior evidence shows that institutions are more likely to purchase shares of firms with persistent disclosure improvement (Bushee & Noe, 2000), and the home bias literature shows that foreign firms with better disclosure attract more US institutional investors (Aggarwal, Klapper, & Wysocki, 2005; Bradshaw, Bushee, & Miller, 2004; Kang & Stulz, 1997). For example, Bradshaw et al. (2004) show that US institutional investors are more likely to be attracted to foreign firms exhibiting a higher level of US GAAP conformity. US investors are likely to exhibit a similar home bias in their preference for foreign firms that voluntarily conform to US disclosure practices by providing management earnings

Table 5 Analysis of firm-level concentrated and institutional ownership

	Sign	(1)	(2)	(3)	(4)	(5)	(6)
Intercept	±	−4.67***	−4.05***	−3.801***	−5.723***	−4.818***	−4.497***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
COMMON (H1)	+	0.874***			0.888***		
		(0.00)			(0.00)		
ANTI-DIRECTOR (H1)	+		0.439***			0.409***	
			(0.00)			(0.00)	
JUDICIAL (H1)	+			0.201***			0.159**
				(0.00)			(0.02)
OWNCON (H2)	−	−0.009***	−0.008***	−0.009***			
		(0.00)	(0.01)	(0.00)			
INST (H2)	+				0.034***	0.033***	0.037***
					(0.00)	(0.00)	(0.00)
FORSALES (H4)	+	0.011***	0.013***	0.012***	0.009***	0.010***	0.010***
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Control variables		Yes	Yes	Yes	Yes	Yes	Yes
Pseudo R^2		0.14	0.13	0.12	0.21	0.20	0.19
No. of observations		3126	3126	3126	2286	2286	2286

This table reports probit regression results of the determinants of management earnings forecasts on firm-level ownership data. The dependent variable is *MF*, with a value of 1 for forecasting firms, and 0 otherwise. Here *OWNCON* is defined as the share of cash flow rights held by the largest shareholders in the forecast year, and *INST* is the percentage of a company's aggregate common stock held by institutions. Coefficient estimates are provided in the top row and *p*-values in the bottom row. Here *, ** and *** indicate significance at the 10, 5 and 1% levels (two-tailed test), respectively.

forecasts, because such practices are more familiar to US investors, reduce their information processing costs, and thus help them better predict the future cash flow of these foreign firms. Therefore we predict that the likelihood of a cross-listed firm providing earnings forecasts in the US is positively associated with the level of institutional ownership. Consistent with our prediction, as shown in columns 4, 5 and 6 of Table 5, we find that the coefficient of *INST* is significantly positive across all cases. Overall, our results suggest that foreign firms listed on the major US exchanges provide voluntary disclosure in response to institutional investors' demand.

Granger Lead–Lag Regressions

With respect to the inverse relation observed between management forecasts and information asymmetries, as shown in Table 4, one cannot rule out the possibility that the observed relation is driven by potential endogeneity problems, or reverse causality. Specifically, the negative association between the presence of management forecasts and AFD could stem from the fact that managers of firms with higher AFD (information asymmetry) find it less cost-effective to voluntarily release forward-looking information.

To address the endogeneity, or reverse causality, issue, we adopt the approach used by Ajinkya et al. (2005) and estimate a Granger lead–lag regression:

$$DISP = \beta_0 + \beta_1 MF + \beta_2 LAGDISP + \sum_k \beta_k CONTROL^k + (Year\ Dummies) + (Industry\ Dummies) + error \qquad (3)$$

where *LAGDISP* is *DISP* lagged one period. As explained by Hamilton (1994), in this condition, *MF* lags the dependent variable *DISP*, and *LAGDISP* in turn lags *MF*, so the time sequence is $LAGDISP \rightarrow MF \rightarrow DISP$. The objective of this model is to control for the potential effect of prior AFD on management forecasts so that we can separate the incremental power of *MF* in the subsequent AFD.

Table 6 presents the results of Granger lead–lag regressions. The coefficient of *LAGDISP* is positive and significant, suggesting that past AFD explains future AFD. Nonetheless, *MF* continues to load with a negatively significant coefficient, implying that the presence of voluntary disclosure in the current period is associated with lower future AFD after controlling for the correlation between past AFD and disclosure. Results of the Granger lead–lag estimation lend more credence to our main regression results, reported in Table 4.

Table 6 Granger lead–lag estimations of the relation between analyst forecast dispersion and management earnings forecast occurrence

	Sign	(1)	(2)	(3)
Intercept	±	0.054	0.016	0.128
		(0.57)	(0.87)	(0.21)
LAGDISP	+	0.058***	0.068***	0.064***
		(0.01)	(0.00)	(0.00)
MF (H5)	±	−0.070***	−0.084***	−0.081***
		(0.00)	(0.00)	(0.00)
COMMON	−	−0.081***		
		(0.00)		
ANTI-DIRECTOR	−		−0.015**	
			(0.02)	
JUDICIAL	−			−0.026***
				(0.00)
LIQUIDITY	−	−0.001***	−0.001***	−0.001***
		(0.00)	(0.00)	(0.00)
LISTTYPE	−	0.007	0.010	0.006
		(0.72)	(0.58)	(0.74)
PROFIT	−	−0.041	−0.039	−0.044
		(0.15)	(0.18)	(0.13)
LIABILITY	+	0.052	0.046	0.044
		(0.15)	(0.20)	(0.22)
SIZE	+	0.027***	0.035***	0.043***
		(0.01)	(0.01)	(0.01)
MKBK	−	−0.017***	−0.017***	−0.016***
		(0.00)	(0.00)	(0.00)
SIC and year dummies		Yes	Yes	Yes
Adjusted R^2		0.10	0.10	0.10
No. of observations		1743	1743	1743

This table reports the Granger lead–lag regression results of the effects of management earnings forecasts on analyst forecast dispersion. The dependent variable is *DISP*, measured by the standard deviation of analyst forecasts following management earnings forecasts. Here *LAGDISP* is *DISP* lagged one period. The Granger lead–lag estimation is to control for potential causality between management earnings forecasts and analyst forecast dispersion. All other variables are defined in the Appendix. Coefficient estimates are provided in the top row and *p*-values in the bottom row. Here *, ** and *** indicate significance at the 10, 5 and 1% levels (two-tailed test), respectively.

Effects of Management Earnings Forecast Precision

Management earnings forecasts are not limited to point forecasts but include range (i.e., closed-interval), open-ended (i.e., minima and maxima) and qualitative forecasts of general impressions about firms' earnings prospects (Baginski & Hassell, 1997). Analytical work shows that more precise signals lead to greater belief revision (Kim & Verrecchia, 1991). To obtain further insights, we now examine: (1) the determinants of forecast precision; and (2) the effects of forecast precision on AFD.

Table 7 summarizes our results: The three precision models (columns 1, 3 and 5) estimate which factors determine the precision of management earnings forecasts released by cross-listed firms. The dependent variable of the precision models is the log of 1 plus forecast precision, which takes the value of 4 for point forecasts, 3 for range forecasts, 2 for open-ended forecasts, 1 for qualitative forecasts, and 0 otherwise. The results of regressions with forecast precision as the dependent variable reveal that firms with stronger home-country legal institutions, from countries with less concentrated ownership, with listings on major US stock exchanges, and with more foreign sales tend to release more precise forecasts.

The AFD regressions presented in columns 2, 4 and 6 of Table 7 investigate the effects of forecast precision on the information asymmetry proxied by AFD. The coefficients of *PRECISION* are all negative and highly significant ($p<0.01$), suggesting that firms releasing more precise forecasts are associated with lower AFD. Overall, the results using forecast precision are consistent with our primary results in Tables 3 and 4, lending further support to Hypothesis 5.

Table 7 Management earnings forecast precision and analyst forecast dispersion

	(1)	(2)	(3)	(4)	(5)	(6)
	Precision	AFD	Precision	AFD	Precision	AFD
Intercept	0.573***	0.140*	0.639***	0.092	0.690***	0.186**
	(0.00)	(0.09)	(0.00)	(0.28)	(0.00)	(0.04)
PRECISION		−0.047***		−0.059***		−0.058***
		(0.00)		(0.00)		(0.00)
COMMON	0.154***	−0.095***				
	(0.00)	(0.00)				
ANTI-DIRECTOR			0.032***	−0.017***		
			(0.00)	(0.00)		
JUDICIAL					0.019**	−0.026***
					(0.03)	(0.00)
LIQUIDITY		−0.001***		−0.001***		−0.001***
		(0.00)		(0.00)		(0.00)
PROFIT		−0.029		−0.028		−0.032*
		(0.14)		(0.15)		(0.10)
LIABILITY		0.060*		0.051		0.046
		(0.06)		(0.12)		(0.15)
MKBK		−0.020***		−0.019***		−0.020***
		(0.00)		(0.00)		(0.00)
OWNERSHIP	−0.056**		−0.049*		−0.069**	
	(0.02)		(0.07)		(0.02)	
LISTTYPE	0.088***	0.010	0.081***	0.013	0.084***	0.008
	(0.00)	(0.53)	(0.00)	(0.42)	(0.00)	(0.61)
ANALYST	0.048***		0.050***		0.052***	
	(0.00)		(0.00)		(0.00)	
LISTTYPE × *ANALYST*	−0.034***		−0.035***		−0.036***	
	(0.00)		(0.00)		(0.00)	
FORSALES	0.002***		0.003***		0.002***	
	(0.00)		(0.00)		(0.00)	
GROWTH	−0.0001*		−0.0001		−0.0001	
	(0.10)		(0.11)		(0.11)	
BIG4	−0.025		−0.031		−0.039	
	(0.57)		(0.49)		(0.38)	
SIZE	−0.111***	0.019**	−0.131***	0.032***	−0.144***	0.042***
	(0.00)	(0.04)	(0.00)	(0.00)	(0.00)	(0.00)
LOSS	−0.140***		−0.139***		−0.146***	
	(0.00)		(0.00)		(0.00)	
NEWS	0.052***		0.048**		0.051***	
	(0.01)		(0.02)		(0.01)	
RegFD	0.024		0.035*		0.040**	
	(0.25)		(0.10)		(0.05)	
LITIGATE	0.258***		0.252***		0.237***	
	(0.00)		(0.00)		(0.00)	
SIC and year dummies	No	Yes	No	Yes	No	Yes
Adjusted or pseudo R^2	0.24	0.10	0.23	0.09	0.23	0.10
No. of observations	3714	2569	3714	2569	3714	2569

The precision models estimate which factors determine the precision of management earnings forecasts released by our sample of cross-listed firms. The dependent variable is the log form of 1 plus forecast precision, which takes the value of 4 for point forecasts, 3 for range forecasts, 2 for open-ended forecasts, 1 for qualitative forecasts, and 0 otherwise. The AFD models estimate the impact of management forecast precision on AFD. The dependent variable is *DISP*, measured by the standard deviation of analyst forecasts following management earnings forecasts. All models are estimated by OLS regressions. Coefficient estimates are provided in the top row and *p*-values in the bottom row. Here *, ** and *** indicate significance at the 10, 5 and 1% levels (two-tailed test), respectively.

CONCLUSIONS

This study explores the likelihood and consequences of voluntary disclosure, as proxied by management earnings forecasts, for a large sample of US cross-listed firms originating from 40 countries, where different home-country institutions generate differing trade-offs between the costs and benefits associated with voluntary disclosure. Our results reveal that the strength of home-country legal institutions is an important factor in increasing the likelihood that cross-listed firms will release forecasts. This result is in line with the prediction of institutional duality theory, that home-country institutions influence the likelihood of international firms engaging in voluntary disclosure. In addition, we show that cross-listed firms are more likely to release management earnings forecasts to US market participants when their shares are listed on major US exchanges (NYSE/AMEX/NASDAQ) than when they are listed on the OTC/Portal. We also find that the likelihood of cross-listed firms releasing management forecasts to US investors is positively associated with the percentage of foreign sales to total sales. Moreover, we find that forecast likelihood is negatively associated with the percentage of cash flow rights held by the largest shareholder, suggesting that agency costs deter controlling shareholders of US cross-listed firms from voluntarily disclosing firm-specific information to US market participants, to hide expropriation activities from US minority investors. Finally, we show that firms releasing management earnings forecasts are associated with lower AFD, suggesting that forecasting firms garner benefits from reducing information asymmetry. Our results are generally robust to a variety of sensitivity checks.

Our research is subject to several caveats. For instance, our results on ownership concentration become weaker after we drop Canadian sample firms. This highlights the importance of exploring another perspective of agency theory, namely, that in some countries controlling shareholders may be more likely to disclose because voluntary disclosure is a signaling mechanism to reduce financing costs. This issue warrants further investigation in an international context. In addition, when examining the economic consequences of voluntary disclosure, our study focuses only on its effect on information asymmetry proxied by AFD. However, the cost of equity capital and the cost of debt capital could likewise be influenced by voluntary disclosure and agency problems (Barnea et al., 1985). We recommend further research in this direction.

Despite these limitations, our study provides a cost–benefit framework for understanding voluntary disclosure practices in an international context. Our work also presents evidence that home-country institutions still matter when foreign firms migrate into the US financial market, and highlights the importance of country institution development. Future research may explore the question of whether, given the current global business climate, country-level factors or firm-level factors matter more in defining firm disclosure and corporate governance.

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NOTES

[1]In 2009, more than 2200 foreign firms from 80 countries were cross-listed on US markets (data from Citigroup Corporate and Investment Banking, www.citigroup.com/adr, and the Bank of New York Global Equity Investing Depositary Receipt Services, www.adrbny.com).

[2]Our research purpose is not to evaluate which (country-level or firm-level) factors matter more in defining firms' voluntary disclosure practices. However, to gain further insights, we calculate the R^2 values for two models: one with country-level factors only, and one with firm-level factors only. Our preliminary results show that both country- and firm-level variables provide significant incremental explanatory

powers, but firm-level variables explain more of the model variations.

[3]For more details, please refer to http://www.adrbny.com.

[4]Sources of information on direct-listing Canadian and Israeli firms include http://www.nyse.com/international/nonuslisted/int_listed.html, http://www.nasdaq.com/asp/NonUsOutput.asp, http://www.amex.com, http://www.pinksheets.com/companysearch/ps_list.jsp, and http://www.otcbb.com.

[5]Our database does not allow us to identify whether a financial analyst is local (i.e., from the home country) or foreign (i.e., from the US or another country). Bae, Stulz, and Tan (2008) find that local analysts have an information advantage compared with foreign analysts. Therefore future research may try to investigate how management earnings forecasts released by cross-listed firms influence local and foreign financial analysts differently.

[6]In the regressions testing the effect of management earnings forecasts on AFD, 2569 observations are employed, because we need observations that have at least two analysts following to calculate the analyst forecast dispersion. To check the robustness of our results, we also look at firms with three or more analysts following, and our results are not sensitive to this correction.

[7]As a robustness check, in all models, we also use three dummies (English common law, French civil law, and German civil law) as proxies for legal origins. However, our results for Hypothesis 1 are not sensitive to this modification.

[8]Data on *OWNCON* and *INST* are collected from Worldscope and Compact Disclosure, respectively.

[9]As a sensitivity check, in all models, we also use two dummies (major US exchanges and OTC) as proxies for listing types. However, the results are not sensitive to this correction.

[10]In our primary tests, we include *FORSALES* as a proxy for cross-border product market integration. We also consider another proxy for a cross-listed firm's interaction with foreign product markets, that is, *FOROP*, which equals 1 if a cross-listed firm has foreign operations, and 0 otherwise. The use of this alternative proxy does not alter the results of our primary tests.

[11]As sensitivity checks, we also use: (1) analyst forecast standard deviations scaled by beginning-of-period stock prices; (2) analyst forecast standard deviations scaled by the absolute value of mean/median analyst forecast estimates; and (3) changes in analyst forecast standard deviations as our dependent variables. Our major conclusion on Hypothesis 5 (i.e., the negative association between the issuance of management forecasts and analyst forecast dispersion) is not sensitive to these robustness checks.

[12]Similarly, Bamber and Cheon (1998) employ the number of days between the management forecast release and the end of the accounting period to estimate the variables of the non-forecasters in their matched sample design.

[13]Canadian firms constitute about 31% of all firms in our sample. To check whether our results reported in Table 3 are driven by their presence, we exclude them from our sample and re-estimate Eqs. (1) and (2) using this reduced sample. Our primary findings generally hold, except that the coefficients of *ANTI-DIRECTOR* (Hypothesis 1) and *OWNERSHIP* (Hypothesis 2) become statistically insignificant at the conventional levels ($p>0.1$). In an additional test using the reduced sample excluding Canadian observations, we replace the country-level ownership variable with the country-level index for risk of expropriation as our proxy for agency costs. This country-level index is an evaluation of the risk of minority shareholder expropriation, and is scaled from 0 to 10, with lower scores for higher risks. Our additional results suggest that: (1) our primary findings hold, whether or not Canadian firms are included in our sample; and (2) firms from countries with higher agency cost are less likely to make earnings forecasts, which is consistent with Hypothesis 2.

[14]In a sensitivity check, we also correct for country-level standard error clustering, and the results are generally consistent with our primary findings, except that the coefficient of *OWNERSHIP* (Hypothesis 2) has its expected sign but is insignificant.

[15]As a sensitivity check, we partition our sample firms into good news and bad news groups and re-run Eq. (1). Our results are robust to this correction. Following Skinner (1994), we also classify good/bad news according to the difference between management earnings forecast and mean analyst forecast consensus, and the results from the robustness checks are generally consistent with our primary results.

[16]Since management earnings forecasts are voluntary, our regression results reported in Table 4 may suffer from a self-selection bias. Hence we re-estimate Eq. (2) using the Heckman (1979) two-stage procedure. In the first stage, we estimate a probit model of a firm's probability to release management earnings forecasts (i.e., Eq. (1)), and thus obtain inverse Mills ratios. In the second stage, we estimate Eq. (2), with the inverse Mills ratio as a control variable. Results using the Heckman two-stage regressions are qualitatively similar to those reported in Table 4. In all models, the coefficients of *MF* are negative and significant ($p<0.01$). It is worth noting that it

is difficult to identify true instrumental variables, so the results are not tabulated or presented in our main text.

[17]Ideally, we would like to use the percentage of shares held by US institutional owners. However, the database on US institutional ownership (i.e., Form 13(f)) covers only a small portion of our sample firms and thus may cause selection bias. We therefore use the percentage of shares held by both US and non-US institutions.

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Appendix

Table A1 Variables definitions and data sources

Variable	Definition	Data source
Dependent variable		
Pr(*MF*)	*Ex ante* probability that a cross-listed firm issues management earnings forecasts, which is *ex post* coded 1 for firms that issued the forecast during the fiscal year, and 0 otherwise	First Call Corporate Investor Guideline (CIG) database
PRECISION	Log form of 1 plus the precision of management earnings forecasts, which takes a value of 4 for point forecasts, 3 for range forecasts, 2 for open-ended forecasts, 1 for qualitative forecasts, and 0 for non-forecasters	First Call Corporate Investor Guideline (CIG) database
DISP	Standard deviation of analyst forecasts following management earnings forecasts	IBES
Country-level variables		
LEGAL	The strength of cross-listed firms' home-country legal regime proxied by one of three proxies: *COMMON*, *ANTI-DIRECTOR*, or *JUDICIAL*	La Porta et al. (2006), Allen, Qian, and Qian (2005)
COMMON	A dummy variable that equals 1 for a cross-listed firm from an English common law country, and 0 otherwise	La Porta et al. (2006)
ANTI-DIRECTOR	An index that aggregates six different shareholder rights and ranges from 0 to 6, with 6 as the highest level of investor protection	La Porta et al. (2006), Allen, Qian, and Qian (2005)
JUDICIAL	An assessment of the efficiency and integrity of a country's legal environment, ranging from 0 to 10, with 10 as the highest standard	La Porta et al. (2006)
OWNERSHIP	An indicator variable equal to 1 if the given country's concentration of ownership is equal to or above the median level, where the ownership concentration is measured as the average percentage of common shares owned by the top three shareholders in the ten largest non-financial, privately owned domestic firms in a given country	La Porta et al. (2006)
LIQUIDITY	Represents the average ratio of the dollar value of shares traded as a percentage of gross domestic product for the period 1996–2000	La Porta et al. (2006)
Firm-level variables		
OWNCON	Ownership concentration of a cross-listed firm, measured by the percentage of cash flow rights held by the largest shareholder in the forecast year	Worldscope, Mergent Online, Form 20-F, company website
INST	Percentage of the company's aggregate common stock held by institutions	Worldscope, Compact Disclosure, Mergent Online, Form 20-F

Table A1 *Continued*

Variable	Definition	Data source
LISTTYPE	An indicator variable that equals 1 for firms cross-listed on the major US exchanges (NYSE/AMEX/NASDAQ), and 0 for firms cross-listed on the OTC and the Portal	Bank of New York; websites of NYSE, AMEX, and NASDAQ; Pink Sheet
ANALYST	Number of analysts following the firm	IBES
FORSALES	Dollar values of foreign sales, deflated by total sales	Worldscope
PROFIT	Operating income deflated by total assets	Worldscope
LIABILITY	Ratio of total liabilities to total assets	Worldscope
GROWTH	Sales growth over the past two years	Worldscope
MKBK	Market value of equity divided by the book value of equity	Worldscope
REGFD	Equals 1 if the observation is related to the post-Regulation Fair Disclosure period, and 0 otherwise	
LITIGATE	An indicator variable that equals 1 for firms in the biotechnology (SIC 2833–2836 and 8731–8734), computers (SIC 3570–3577 and 7370–7374), electronics (SIC 3600–3674), and retail (SIC 5200–5961) industries, and 0 otherwise	Worldscope and firms' annual reports
BIG4	An indicator variable that equals 1 if a cross-listed firm is audited by one of the Big Four auditors, and 0 otherwise	Worldscope, and firms' annual reports
SIZE	Log of the total sales of a firm at the beginning of the fiscal period	Worldscope
LOSS	An indicator variable that equals 1 if the firm reported a loss in the current period, and 0 otherwise	Worldscope
NEWS	An indicator variable that equals 1 if the current-period earnings per share is greater or equal to that of the previous period, and 0 otherwise	Worldscope and IBES
Other variables		
YearDummies	Year dummies	
IndustryDummies	Industry dummies, where industries are as defined in Durnev and Kim (2005): petroleum (SIC 13, 29), consumer durables (SIC 30, 36, 37, 50, 55, 57), basic industry (SIC 8, 10, 12, 14, 24, 26, 28, 33), food and tobacco (SIC 20, 21, 54), construction (SIC 15, 16, 17, 32), capital goods (SIC 34, 35, 38, 39), transportation (SIC 40–42, 44, 45, 47), textiles and trade (SIC 22, 23, 51, 53, 56, 59), services (SIC 7, 73, 75, 80, 82, 83, 87, 96), leisure (SIC 27, 58, 70, 79), unregulated utilities (SIC 48), regulated utilities (SIC 49), and financials (SIC 60–63, 65, 67)	Worldscope and firms' annual reports

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