Foreign Competition for Shares and the Pricing of Information Asymmetry: Evidence from Equity Market Liberalization*

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Abstract

Using the equity market liberalization of 23 emerging market countries between 1996 and 2006, we examine how the degree of competition for a firm's shares affects the price of information asymmetry. We find evidence of a significant decline in the pricing of information asymmetry as countries remove regulatory restrictions on foreign ownership. Our study provides novel evidence on the link between the degree of competitiveness of equity markets and the price of information asymmetry. The work also furthers our understanding of the economic consequences of foreign stock ownership.

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1. Introduction

Theory suggests that whether and to what extent an economy prices information asymmetry amongst traders depends on the level of competition for firms' shares (Kyle, 1989; Lambert et al., 2012). Competition, however, is an abstract concept and measuring it is fraught with controversy. The theoretical literature characterizes competition in the context of risk-bearing capacity – in effect, the total number of investors in the economy. The idea is that as the number of investors in the economy grows, the greater the economy's capacity to bear risk, and the more competition for firms' shares. In the extreme, as risk-bearing capacity becomes very large, an economy approaches perfect competition. [1] In the presence of perfect competition, information asymmetry has no separate effect on cost of capital after controlling for the average precision of information. In contrast, in an economy where competition is imperfect, information asymmetry has a separate effect on firms' cost of capital. However, as the number of investors who participate in the economy grows, the economy's risk-bearing capacity increases – thereby pushing it closer to perfect competition. In conjunction with this, information asymmetry commands a smaller price discount.

Acknowledging that measuring competition is complex, Armstrong et al. (2011) and Akins et al. (2011) in pioneering studies provide important, initial evidence on the effect of competition on the pricing of information asymmetry by employing the number of investors holding a firm's shares as a measure of competition for that firm's shares. While this approach is novel, it raises significant interpretational challenges because of its departure from the theoretical notion of risk-bearing capacity, where risk-bearing capacity measures the economy's potential for share competition *irrespective* of the number of shareholders for a firm

[1] Perfect competition refers to the notion that investors act as price takers and thus can buy or sell any quantity at the market price. Such a characterization implicitly assumes that the number of investors in the economy is very large (countably infinite): see, e.g., Hellwig (1980)

at a particular time.[2] Thus, we believe it is important to re-visit whether and to what extent an economy prices information asymmetry based on its risk-bearing capacity, a construct closer in spirit to the theoretical literature. This study aims to fill this gap by using equity market liberalization of emerging market countries as shocks to the risk bearing capacity of these countries.

We believe this is an important research undertaking given the significance of this question and the current state of the literature. Information asymmetry amongst traders has long been a concern of securities regulators across the globe, with the introduction of many policy initiatives designed to curb this problem. For example, the U.S. Securities and Exchange Commission (SEC) enacted Regulation Fair Disclosure in 2000 to prevent firms from making selective disclosures to investors and analysts. While such initiatives implicitly assume that the costs associated with information asymmetry are large, most recent studies find that asymmetry is not priced on average (Duarte and Young, 2009; Lai et al., 2014; Mohanram and Rajgopal, 2009). The insight originating with Armstrong et al. (2011) and Akins et al. (2011) is to search for evidence of pricing of information asymmetry conditional on the level of competition. We follow-up on this suggestion, but in a fashion that uses risk-bearing capacity to measure competition. The two-fold advantage of our approach is that it positions us closer to the extant theoretical literature and resolves the interpretational challenges associated with using firm-level shareholding patterns to measure competition.[3]

[2] The issue is that this approach implicitly assumes that investors who choose not to hold a stock do not contribute to the overall risk-bearing capacity. However, the mere expectation that these investors might purchase shares if the price of risk falls sufficiently low allows these investors to affect share prices without necessarily holding the stock. That is, it is not necessarily clear that a firm held by, say, 30 investors, has greater risk-bearing capacity than a firm held by 20 investors. The issue of measuring competition becomes even more fraught when one considers that a firm's shareholders are the outcome of endogenous ownership decisions made by shareholders based on investee firm characteristics: there is the possibility that these firm characteristics could explain the observed association between the number of shareholders and the extent to which information asymmetry is priced. For example, it could be that riskier and difficult to understand opaque firms are owned by a few sophisticated blockholders, while less risky and more transparent firms are owned by a larger number of less sophisticated retail investors. These interpretational challenges make it difficult to associate competition at the firm-level with whether information asymmetry is priced.

[3] With regard to the latter, see the discussion in the previous footnote.

Equity market liberalization is a governmental decision to permit foreign ownership of stocks. Therefore, liberalizations essentially represent shocks to country-level risk-bearing capacity and provide a natural setting to study our research question. We identify key liberalization events for our sample of 23 emerging market countries by focusing on large jumps in a time-varying measure of regulatory restrictions on foreign ownership that has been widely employed in prior work.[4] Our sample covers 13 liberalization shocks experienced in these countries between 1996 and 2006. We adopt a difference-in-differences (DiD) methodology to compare the change in the price of information asymmetry for firms in countries that experience liberalization shocks (treatment countries) to the contemporaneous change for firms in countries that experience no such shocks (control countries). We use bid-ask spreads as a proxy for information asymmetry and compare the cost of capital of firms with high bid-ask spreads to firms with low bid-ask spreads to measure the effect of information asymmetry on price. In effect, our DiD design examines changes in the mapping between bid-ask spreads and cost of capital around liberalization shocks for treatment countries relative to control countries.

We measure a firm's cost of capital using an implied cost of capital approach wherein we calculate cost of capital as the discount rate that equates the firm's stock price to the present value of all expected cash flows. As we discuss in detail in Subsection 3.1, use of implied cost of capital (as opposed to stock returns or dividend yields) allows us to address the confounding effect of growth opportunities; this is particularly problematic in a liberalization setting.[5] The implied cost of capital approach provides ex ante estimates at a point in time. This allows us to examine cost of capital changes around relatively narrow liberalization-shock windows,

[4] The measure was developed in Bekaert (1995) and Edison and Warnock (2003). Examples of other studies that use some variant of this measure include Bae et al. (2004), Bae et al. (2012), Bekaert et al. (2007), De Jong and De Roon (2005), and Mitton (2006).
[5] As noted in Henry (2000), policy makers are likely to liberalize when their countries experience expansion in growth opportunities. Second, the liberalization event itself is likely to trigger economic growth.

thereby minimizing concerns about the effect of growth and other confounding events. In addition, this approach enables us to adjust explicitly for forward expectations of growth using analyst forecasts, and, hence, distinguish better between the cost of capital and growth effects (Hail and Leuz, 2009).

We find that the price discount associated with having higher bid-ask spreads declines significantly following liberalization shocks for firms in treatment countries compared to contemporaneous changes in the price discount for firms in control countries. Our estimates imply that following a large liberalization shock, the price discount associated with an increase in bid-ask spreads equivalent to the interquartile range (i.e., increase from the bottom to the top quartile) decreases by about 37 basis points. In support of the parallel trends assumption underlying our DiD design, we find no evidence of this decline in the periods prior to the liberalization shocks. We also find no evidence of such a decline in placebo tests that exploit liberalization of four financial markets other than equity markets: money markets, bond markets, derivative markets, and government securities markets. This suggests that any changes in countries' growth prospects or other reforms that tend to accompany a general opening up of financial markets are unlikely to drive our main results. In further support of our story, we find that the decline in the price of information asymmetry manifests primarily for firms that do not already have access to foreign capital through American Depository Receipts (ADRs). Overall, the results suggest that allowing foreign investors to participate in an economy attenuates the impact of information asymmetry on stock prices.

Finally, we exploit the international nature of our setting to conduct an exploratory analysis of how the quality of home-country institutions and governance infrastructure affects the benefits of liberalizations on the price of information asymmetry. Poor institutional infrastructure and corporate governance typically thwarts local investors from participating in equity markets, thus countries with poor institutions are more likely to approximate imperfect market settings prior to liberalization. Consequently, such countries are more likely to benefit from reductions in the price of information asymmetry resulting from foreign investor participation. Consistent with this prediction, we find that the decline in the price of information asymmetry associated with an increase in risk-bearing capacity is concentrated among countries that have low levels of securities regulation and insider trading enforcement.

2. Background on Liberalization and its Measurement

2.1. *The multistage nature of liberalization and its measurement*

Most emerging market countries initiated the official equity market liberalization process in late 1980s or early 1990s. Liberalization, however, is a multi-step undertaking and this initial liberalization for most countries represents only the first step in the overall liberalization (e.g., Bekaert and Harvey, 1995; Edison and Warnock, 2003). For example, Korea initiated the liberalization process in 1992 by allowing foreign investors to own up to 10% of domestic equity, but a significant part of the liberalization occurred in late 1990s when Korea increased the foreign ownership limit to 55% by 1997; Ownership restrictions were completely eliminated by 1998 in all but a few regulated sectors.

Because of the data requirements for measuring the cost of capital (detailed later) and for the availability of bid-ask spreads, our sample is constrained to begin from 1996, which falls after the initial liberalization dates for our sample countries. As explained above, however, this is not problematic as we are able to exploit the rich cross-sectional and time series variation in the liberalization subsequent to the initial liberalization. We identify the liberalization shocks during our sample period by examining jumps in a measure of regulatory restrictions on foreign ownership that has been widely employed in prior work and constructed based on information provided by the Standard and Poor's/International Finance Corporation (SP/IFC).[6] We obtain

[6] In the Internet Appendix, we present the robustness of our results to two other measures of regulatory restrictions on foreign ownership that have been used in prior work: the Chinn-Ito index and the Schindler (2009) measure, both of which are based on information in the International Monetary Fund's (IMF's) Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER). A significant drawback of these measures is that they

data on this measure from Edison and Warnock (2003), who provide a monthly time series of this measure until the end of 2006. [7]

The methodology for the construction of this measure is as follows. SP/IFC, for each emerging market country, computes two indices: a global index (IFCG), designed to represent the market value of equity in a country, and an investable index (IFCI), designed to represent the subset of the market value of equity in a country that is available to foreign investors. To compute the investable index, SP/IFC first determines the openness of each stock to foreign ownership at the market level by examining country-level restrictions on foreign stock ownership. Next, they determine the extent of industry, corporate by-law, and corporate charter limitations on foreign ownership. Based on the market-wide limits, and industry and firm-specific limitations, an openness (or investability) factor for each stock is determined, which equals the fraction of stock available to foreign investors. This information at the stock level is then aggregated to determine the investible index (IFCI) at the country level, which when scaled by the global index (IFCG) provides a country-level measure of the fraction of equity available for foreign investment. Specifically, one can measure the extent of a country's liberalization as:

$$LIBERALIZATION\ (unadjusted)_{i,t} = \frac{MC_{i,t}^{IFCI}}{MC_{i,t}^{IFCG}}, \qquad (1)$$

where *MC* is the market capitalization at time *t* of country *i*'s IFCI or IFCG index.

The above liberalization measure closely tracks actual changes in regulatory restrictions on foreign ownership. As an example, for Korea the measure increases from 0% in 1991 to 9.5% in 1992, 54.5% by 1997, and 91.2% by 1998. This maps closely into the actual regulatory

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consider equity market liberalization to be an "all-or-nothing" variable and thus do not capture much of the variation in liberalization rules captured by our measure. Despite this issue, we find that our inferences are generally robust to use of these measures.

[7] SP/IFC discontinued this data collection after 2006. Data on investability can be downloaded from Frank Warnock's website at: http://faculty.darden.virginia.edu/warnockf/research.htm.

restrictions which forbade any foreign ownership until 1991, allowed 10% ownership by 1992, 55% by 1997, and 100% by 1998, except for some regulated sectors such as telecommunications, air transportations, and broadcasting. The measure is somewhat lower than the actual regulatory limits because of exceptions for some regulated sectors, which either had lower limits or were completely closed to foreign investors.

One issue with this ratio is that because ownership restrictions vary across sectors, asymmetric shocks to investable and non-investable stocks will lead to relative price changes, causing changes in the ratio of the market capitalizations that are unrelated to restrictions on foreign ownership. To circumvent this issue, Edison and Warnock (2003) also provide an adjusted liberalization measure that approximately removes the effect of asymmetric price shocks by scaling the measure with the relative prices of investable and non-investable stocks in the following way:

$$LIBERALIZATION_{i,t} = \frac{MC_{i,t}^{IFCI} \big/ P_{i,t}^{IFCI}}{MC_{i,t}^{IFCG} \big/ P_{i,t}^{IFCG}}, \qquad (2)$$

where P denotes the price indices. Edison and Warnock (2003) find that the adjusted measure exhibits smoother variation over time and more closely tracks changes in regulatory restrictions on foreign ownership.

We identify liberalization shocks for our sample countries based on jumps in this liberalization measure.[8] Table 1 presents the descriptives for this measure for our sample ranging from 1996 to 2006. For comparison, we also provide the estimated initial liberalization dates from Bekeart and Harvey (2000) and Mitton (2006). [9] First, for all the countries except

[8] We also considered using reverse liberalization shocks (i.e., increases in restrictions on foreign ownership) to test our hypothesis. We are, however, unable to find any economically meaningful reverse liberalization shocks because most emerging market countries have trended toward increasing liberalization during our sample period. Any reverse liberalization shocks that exist are either economically too small or represent only a temporary reversal for a very short-time period.

[9] As Henry (2000, Table II) shows, there is no consensus on the official liberalization date.

South Africa [as per dates outlined in Bekaert and Harvey (2000)] and Israel [as per Mitton (2006)], the official/initial liberalization dates fall well before 1996, the beginning of our sample period. Second, there is considerable time series and cross-sectional variation in the measure that we use to identify changes to liberalization that are staggered across countries over time. To further illustrate this point, in Figure 1 we plot the liberalization measures for two of our sample countries: Brazil and Korea. In each case, the (first) major regulatory reform is indeed associated with a significant increase in this measure, but the measure continues to change. For instance, foreign access to the Korean equity market increased significantly in 1997 and 1998, which results in large jumps in the liberalization measure.

A potential concern is that removal of explicit regulatory barriers to foreign ownership may not result in increased competition for firms' shares if foreign investors continue to face significant implicit barriers due to poor local institutional infrastructure. For example, foreign investors may not be willing to invest in stocks in emerging markets if they are concerned about poor investor protection and governance in these countries. Prior studies, however, provide strong evidence that removal of regulatory barriers is associated with a significant increase in foreign capital flows, market valuations, investments, and growth.[10] It would be difficult to explain these findings if the liberalization shocks did not result in a significant increase in the risk-bearing capacity from foreign investors. Although data constraints on foreign ownership significantly limit us, in the Internet Appendix we also provide evidence that removal of regulatory barriers is associated with increased investment from foreign institutional investors.

2.2. *Identifying liberalization shocks*

We identify liberalization events for our sample countries based on economically large changes to the *LIBERALIZATION* measure during our 1996-2006 sample period. Specifically,

[10] See, for example, Bae et al., 2004; Bae and Goyal, 2010; Bekaert, 1995; De Jong and De Roon, 2005; Edison and Warnock, 2003; Henry, 2000; Mitton, 2006.

we classify the changes to the *LIBERALIZATION* measure as "large" if they fall in the top tercile of the distribution of all country-year changes in the continuous measure. We later document that our findings are robust to several alternative definitions of liberalization shocks. In Table 1, we list the years of the first top-tercile liberalization shocks for our sample countries. We find that 13 countries experience a top-tercile liberalization shock, which on average (median) represents a 15% (14%) increase in the regulatory limit on foreign ownership.

We focus on large shocks to liberalization for two reasons. First, by design the Edison and Warnock (2003) measure only approximately captures the restrictions on foreign ownership; the measure also exhibits minor variations even without changes in liberalization rules due to asymmetric price shocks to the numerator and denominator. For example, Korea had a 10% limit on ownership by foreign investors from 1992 to 1994. During this period, the monthly time series of the Edison and Warnock measure varies from 9.4% to 11%. Therefore, while the measure closely tracks the actual limit of 10%, if we used the continuous changes in the measure, we would classify even minor changes resulting from asymmetric price shocks as liberalization events, significantly reducing the power of our tests.

Second, focusing on large changes allows us to construct our analyses around economically significant liberalization events. For example, the liberalization shock dummy for Korea takes the value of 1 for 1997. This was widely viewed as the year in which Korea for the first time indicated its strong commitment to liberalization by increasing the limit on foreign ownership from 20% to 55%. All prior liberalization steps taken by Korea during our sample period (i.e., 1996 onwards) are small and represent no more than 3% increase in foreign ownership limit. Focusing on such key liberalization events is likely to be important because, as argued in Edison and Warnock (2008), foreign investors may hesitate to provide capital if they are unsure about the government's ongoing commitment to liberalization. This methodology helps us identify settings that have statistical and economic power to examine the impact of liberalization on the cost of capital.

3. Research Design

3.1. *Measurement of the cost of capital*

We measure the cost of capital for a firm as the internal rate of return that equates the current stock price with the expected sequence of future (abnormal) earnings (e.g., Gebhardt et al., 2001; Hail and Leuz, 2006). This implied cost of capital approach offers significant advantages in the context of liberalization over other approaches that rely on either realized stock returns or dividend yields to infer the cost of capital effects. As discussed in Stulz (1999), the key issue with using realized stock returns in the context of liberalizations is that one typically requires a long time series of returns to reliably estimate the cost of capital effects using this approach. The multi-stage nature of the liberalization, however, makes this infeasible. For example, as discussed above, Korea underwent multiple liberalization shocks in 1991 and 1998 and the gap between adjacent shocks is not sufficient to permit reliable estimation of cost of capital effects using any approach that requires a long-time series of data. Stulz (1999) illustrates how applying a stock return-based approach in such a situation can lead to dramatically incorrect inferences. Use of dividend yields is also problematic in the context of liberalizations because growth opportunities directly affect dividend yields, making it difficult to disentangle the growth effects from the cost of capital effects (e.g., Bekaert and Harvey, 2000; Hail and Leuz, 2006).

By providing cost of capital estimates at a point in time, the implied cost of capital approach circumvents the issues above that are associated with using realized returns or dividend yields. We exploit this characteristic of the implied cost of capital approach to examine cost of capital changes around relatively narrow liberalization-shock windows, in an attempt to minimize concerns about the effects of growth and other confounding events. Furthermore, this approach allows us to adjust for forward expectations of growth using analyst forecasts, thus providing a better differentiation of cost of capital and growth effects (Hail and Leuz, 2006, 2009).

A key challenge with the use of implied cost of capital measures is the possibility of anticipation effects. If investors learn about upcoming liberalization events in advance, then the implied cost of capital measures could respond prior to the actual liberalization event, making it difficult to detect any changes in cost of capital around the liberalization event. Henry (2000), however, finds that for 7 of the 12 emerging market countries in his sample, the earliest news of liberalization appears in media either on or after the actual liberalization date; and, for 3 of the 5 remaining countries, the news appears only a month in advance. He also finds that the earliest the stock prices begin to move is about 8 months prior to the month of liberalization, with the strongest returns occurring in the month of liberalization itself. We follow a conservative strategy to address this issue by dropping the year prior to the liberalization shocks from our sample. This allows us to compare the cost of capital after the liberalization to the cost of capital in the pre-period when markets are likely to have little knowledge about the upcoming liberalization shock.[11]

One may also be concerned about the possibility that investors may be able to anticipate the liberalization events in our sample at the time of initial reforms that occurred in the late 1980s or early 1990s. We view this possibility as quite remote. As shown in Table 1, the liberalization events in our sample on average occur 9.7 years [8.8 years] after the initial liberalization based on the dates in Bekaert and Harvey (2000) [Mitton, 2006]. It is unlikely that market participants would be able to anticipate and capitalize in stock prices the effect of events almost a decade in advance, especially given the instability and uncertainty that characterizes the policy environments in the emerging market countries. Indeed, Edison and Warnock (2003, 2008), highlight that even after the initial liberalization reforms, emerging

[11] Our inferences are robust if we do not drop the prior year.

market countries exhibit considerable diversity in the speed, timing, and extent to which they pursue their liberalization program, with few countries even exhibiting temporary reversals of prior liberalization events (e.g., Zimbabwe, Malaysia, and Venezuela). Furthermore, anticipation of liberalization events, if it affects our analysis, is only likely to bias us against finding cost of capital effects. Therefore, one can view our estimates as lower bounds on the effects of liberalization.

We adopt four of the most commonly used models to measure the implied cost of capital: Claus and Thomas (2001), Gebhardt et al. (2001), Easton (2004), and Ohlson and Juettner-Nauroth (2005). All four models are consistent with the discounted dividend valuation model, but exploit basic accounting relations to obtain an equivalent valuation equation based on residual income or abnormal earnings. We substitute market price and forecasts of abnormal earnings using analyst forecasts in these models to back out the implied cost of capital. The individual models differ with respect to the use of analyst forecast data, the assumptions regarding short-term and long-term growth, the explicit forecasting horizon, the incorporation of industry effects, and the integration of inflation into the steady-state terminal value. We closely follow the approach in Hail and Leuz (2006, 2009) in making the required assumptions. In the Appendix we summarize the four models, describe the key assumptions, and discuss the data requirements. Following Hail and Leuz (2006, 2009), we use the average of the cost of capital estimates from the four models as our main measure of implied cost of capital (*COC*).

Despite their appeal for our setting, implied cost of capital models have their own limitations. They require the assumption that consensus analyst forecasts are reasonable proxies for the market's expectations of future earnings, which might not always be the case (e.g., Easton and Sommers, 2007; Frankel and Lee, 1998). The models also limit the sample to firms that financial analysts cover and have positive earnings forecasts. We gauge the sensitivity of our results to the model assumptions.

3.2. *Empirical Specification*

We identify the year of the first large jump in equity market liberalization (as explained in Subsection 2.2) and use a standard difference-in-differences (DiD) approach to examine the effect of equity market liberalization on the pricing of information asymmetry. Specifically, we estimate the following two specifications to explore the cost of capital effects of liberalization:

$$COC_{i,c,t} = \alpha_c + \mu_{s,t} + \beta_1 POST - LIBERALIZATION_{c.t} + \Gamma Z + \epsilon_{i,c,t}, \qquad (3)$$

$$COC_{i,c,t} = \alpha_c + \mu_{s,t} + \beta_1 POST - LIBERALIZATION_{c.t} + \beta_2 BIDASK_{i,c,t} + \beta_3 POST - LIBERALIZATION_{c.t} * BIDASK_{i,c,t} + \Gamma Z + \epsilon_{i,c,t}, \qquad (4)$$

where *COC* is firm *i*'s implied cost of capital for year *t*. Consistent with Hail and Leuz (2009), we measure cost of capital as of month+10 after the fiscal-year end *t* to ensure the accounting information for the fiscal year is reflected in stock prices. Subscript *s* denotes the industry to which firm *i* belongs and subscript *c* denotes the country where the firm is located. *POST-LIBERALIZATION* is a dummy variable that takes a value of one starting the year a country first experiences a change in the restrictions of foreign ownership that is in the top tercile across all country-year changes, and zero otherwise. We verify the robustness of our results using several alternative definitions of liberalization shocks. *BIDASK* is a measure of the firm-level information asymmetry; we calculate it as the annual median of the bid-ask spreads in equity prices prevailing during fiscal year *t*. *Z* denotes a vector of control variables that we discuss below.

The specification also includes industry-year interactive fixed effects (i.e., a separate fixed effect, $\mu_{s,t}$, for each industry-year combination) and country-fixed effects (α_c). Inclusion of the industry-year interactive fixed effects ensures that we are comparing treatment and control firms that operate in the same industry, allowing us to "difference away" unobserved time-varying industry shocks (Gormley and Matsa, 2013). This is important because lliberalization events are likely to coincide with expansions in the growth prospects of industries that are crucial to the local economy. We measure these fixed effects using the industry classification in Campbell (1996). We include country-fixed effects to remove any unobserved cross-sectional variation in country characteristics that may drive the results.

With the above fixed effect structure, equation (3) represents a generalized DiD specification (Bertrand and Mullainathan, 2003; Bertrand, et al., 2004). Specifically, the coefficient on *POST-LIBERALIZATION* measures the average change in the cost of capital for firms in countries that experience liberalization shocks relative to the average contemporaneous change in the cost of capital of firms that operate in the same industry, but in countries that do not experience liberalization shocks. As explained in Bertrand and Mullainathan (2003), the above regression framework exploits the staggered timing of liberalization events. The staggered timing implies that the control group includes not only firms in countries that that never experience liberalization shocks during our sample period, but also firms in countries that did not experience liberalization at time *t*, even if liberalization may have taken place for these firms before or after time *t*.

In equation (4), we build upon equation (3) to examine the effect of liberalization on the pricing of information asymmetry. The coefficient on *BIDASK* provides an estimate of the price of information asymmetry in the periods prior to the liberalization shocks. The key coefficient of interest is the coefficient on *POST-LIBERALIZATION*BIDASK,* which is the DiD estimate of the effect of liberalization shocks on the price of information asymmetry. Specifically, the coefficient estimates the change in the price of information asymmetry (as assessed by the mapping of bid-ask spread into the cost of capital) for firms in countries that experience liberalization shocks relative to the change for firms in the same industry but operating in countries that do not experience liberalization shocks. If equity market liberalization reduces the price of information asymmetry, the coefficient on *POST-*

*LIBERALIZATION*BIDASK* should be negative. We estimate the above specifications using ordinary least squares (OLS) and obtain standard errors by clustering at the country level.

A natural issue with any international study such as ours is that the control and treatment observations are drawn from different countries, raising concerns about the validity of the parallel trends assumption underlying our DiD design. To address this issue, we test for parallel trends in periods prior to the liberalization shocks and find supportive evidence. In an additional test, we also obtain within country identification using variation in the benefits of liberalization and obtain similar inferences.

An important point to note is that when we discuss the economic magnitudes of the effect of liberalization, we focus on the DiD coefficient on *POST-LIBERALIZATION*BIDASK*. We avoid reading too strongly into the magnitudes of the *levels* of the price of information asymmetry in either the pre- or post-liberalization period. The key issue is that these estimates (unlike the DiD estimate) are based on simple cross-sectional comparisons of firms with high and low information asymmetry and therefore are likely to be contaminated by measurement error and uncontrolled for cross-sectional differences. The issue of measurement error is particularly acute in our setting because the cost of capital measures are known to contain significant noise. As a result, in the cross-section at any point in time, the mapping between information asymmetry and the cost of capital likely entails significant noise. As a consequence, it could be subject to attenuation bias. This effect of measurement error, however, is likely to be purged out of our DiD estimate, which is estimated based on *changes* in the price of information asymmetry.[12]

3.3. *Control variables*

[12] One can see this with the help of an example. Suppose the true price of a unit of IA is 20 bps in the pre-period and 10 bps in the post-period. However, because of the attenuation bias caused by measurement error in the underlying proxies, allow for the fact that one underestimates the price by 10 bps in both the pre- and post-periods. In this example, our DiD design would correctly estimate the effect of liberalization on the price of IA to be 10 bps. But the fact that a unit of IA is priced at 0 bps in the post-period reflects measurement error and not that information asymmetry is not priced in the post-period.

Our control variables closely follow Hail and Leuz (2009). Our controls for firm-level characteristics include firm size, measured as the natural logarithm of book value of assets (*LOGSIZE*), stock return variability (*RETVAR*), measured as the standard deviation of monthly stock returns over the last 12 months, and leverage (*LEVERAGE*), measured as the ratio of total liabilities to total assets.[13] Based on prior studies, we expect the cost of capital to exhibit a negative association with firm size, and positive associations with return variability and financial leverage (e.g., Fama and French, 1992, 1993; Hail and Leuz, 2009).

We include bias in analyst forecasts (*FBIAS*) as an additional control because forecast bias can affect implied cost of capital estimates (Easton and Sommers, 2007). *FBIAS* equals the one year ahead IBES analyst forecast error (mean forecast for the next fiscal year minus actual earnings) scaled by lagged total assets. We measure control variables and bid-ask spreads as of the fiscal-year end. However, as noted in the previous section and, consistent with Hail and Leuz (2009), we measure the cost of capital as of month+10 after the fiscal-year end to ensure accounting information for the fiscal year is reflected in stock prices. In equation (3), we also control for the country-level inflation rate at the end of the fiscal year (*INFLATION*) in all of our regressions. We express analyst forecasts in nominal terms and local currency, which implies that the resulting estimates for the cost of capital reflect countries' expected inflation rates. As investors' expectations for future inflation are only imperfectly observable, we introduce a separate control variable for cross-sectional differences in inflation.

3.4. *Data sources and sample description*

We obtain information on accounting variables from Datastream, Worldscope, and Bloomberg; analyst earnings and growth forecasts are from Institutional Brokers' Estimate System (I/B/E/S); and macroeconomic data are from the World Bank and Datastream. We

[13] We do not use the market value of equity to measure firm size because, as argued in Hail and Leuz (2009), this variable could absorb the hypothesized effect if liberalization indeed decreases the cost of capital and leads to increased valuations.

obtain information on bid-ask spreads from Datastream and Bloomberg. As explained in Subsection 3.1, we control for contamination in our estimates resulting from anticipation of liberalization events by dropping the year prior to the liberalization shocks. Our final sample contains firms for which information on all required variables is available and comprises a maximum of 8,474 firm-year observations; this represents over 2,927 firms from 23 emerging market countries from the 1996 to 2006 period. *BIDASK* spreads are widely available only from 1996 and not for all firms. Accordingly, we base our tests involving *BIDASK* on a relatively smaller sample of 5,694 observations. More details about our sample construction are in the Appendix.

Table 2 presents summary statistics for our measures of the cost of capital and control variables. The mean (median) cost of capital is 15.5% (13.6%). Because our sample is comprised of emerging market countries that tend to be riskier than developed countries examined in prior studies, not surprisingly these estimates are somewhat higher than the estimates in Hail and Leuz (2006, 2009).

4. Results

4.1. *Main results*

As a baseline, we first examine the effect of liberalization on the cost of capital by presenting OLS estimates of equation (3) in Table 3. Our key coefficient of interest is the coefficient on *POST-LIBERALIZATION*. The results in column (1) show that the coefficient on *POST-LIBERALIZATION* is negative and significant (coefficient = -0.006; *t*-stat = -2.965), suggesting that post-liberalization firms experience a reduction in cost of capital. In terms of economic magnitude, the results suggests that firms on average experienced a 60 bps reduction in cost of capital following a top-tercile shock to liberalization, which on average is equivalent to a 15% increase in the limit on foreign ownership. Consistent with prior literature, we also find a positive and significant coefficient on *INFLATION*, a negative coefficient on *LOGSIZE*, a positive coefficient on *RETVAR*, positive coefficient on *LEVERAGE*, and a positive coefficient on *FBIAS* (e.g., Fama and French, 1992, 1993; Hail and Leuz, 2009). This model explains about 30% of the international variation in the cost of equity capital, which is consistent with prior work (e.g., Botosan and Plumlee, 2005; Hail and Leuz, 2006, 2009).

We also re-estimate the specification for a smaller subsample of observations that have non-missing *BIDASK* spreads. Given that the main focus of the remainder of this study is the pricing of information asymmetry, it is important to ensure that the results are not sensitive to the sample attrition. The results in column (2) in Table 3 indicate that the results that liberalization reduces the cost of capital continues to hold with similar statistical and economic significance (coefficient = -0.007; *t*-stat =-3.811).

We next turn our attention to the main specification of interest, equation (4), which examines how liberalization affects the price of information asymmetry. Column (3) of Table 3 presents the results. Our key coefficient of interest is the coefficient on the interaction term between *POST-LIBERALIZATION* and *BIDASK*, which provides the DiD estimate of the effect of liberalization on the pricing of information asymmetry. The coefficient on *BIDASK* is an estimate of the price of information asymmetry prior to the liberalization shock. It can be seen that the coefficient on *BIDASK* is positive and significant (coefficient = 0.413; *t*-stat = 3.450), suggesting that information asymmetry is priced on average prior to the liberalization shocks. Most importantly, we find that the coefficient for *POST-LIBERALIZATION*BIDASK* is negative and significant (coefficient = -0.312; *t*-stat = -2.874), suggesting that the price of information asymmetry declines following liberalization shocks.

We next discuss the economic significance of the effect of liberalization shocks on the price of information asymmetry. For reasons explained in Subsection 3.2, we avoid reading too strongly into the magnitudes of the *levels* of the price of information asymmetry either in the pre- or post-liberalization period. These results are based on a cross-sectional comparison of firms with high and low information asymmetry, and therefore are likely to be contaminated by measurement error and uncontrolled for cross-sectional differences. Instead, we focus our discussion around the DiD estimate of the effect on the price of information asymmetry, which is designed to mitigate the above issues that are typically associated with estimates based on simple cross-sectional comparisons. The estimate of coefficient on *POST-LIBERALIZATION*BIDASK* implies that the price discount associated with an increase in bid-ask spreads equivalent to the interquartile range (i.e., increase from the bottom to the top quartile) decreases by about 37 bps following a liberalization shock.[14] We caution readers from comparing our estimates to the ones obtained in prior work (Armstrong et al., 2011; Akins et al., 2011) because of the interpretational challenges associated with using firm-level shareholding patterns to measure competition (see footnote 2). Furthermore, while prior studies focus on U.S. capital markets, our estimates are based on emerging markets.

4.2. *Assessing parallel trends and the timing of cost of capital changes*

In this Subsection, we explore the timing of cost of capital changes to test the parallel trends assumption underlying our DiD design and to examine the persistence of the effects. We do so by estimating modified versions of the regression specifications in which we replace the *POST-LIBERALIZATION* indicator with separate indicator variables for years leading up to the shock and after the shock. Specifically, we include indicators (*POST-LIBERALIZATION (-2)* and *POST-LIBERALIZATION (-1)*) for the periods two-years and one-year prior to the liberalization announcement. We also include *POST-LIBERALIZATION (0), POST-LIBERALIZATION (1)*, *POST-LIBERALIZATION (2)*, and *POST-LIBERALIZATION (3)*, which are indicator variables for the year of, one-year after, two-years after, and three-years after a country experiences a liberalization shock.[15] Finally, *POST-LIBERALIZATION (>4)* is

[14] This is estimated as the (interquartile range of *BIDASK*)*(coefficient on *BIDASK*POST-LIBERALIZATION*) = 0.012*0.312 = 0.37%.

[15] Because our sample is constrained to begin from 1996 (due to unavailability of bid-ask spread data) and 10 (out of total 13) of the liberalization shocks occur in years 1999 or before, we do not have sufficient pre-period

a dummy that takes a value of one if a country experienced a liberalization shock four or more years ago.

Table 4, Panel A presents evidence on the timing of average cost of capital changes in a tabular format and Figure 2 also presents a time-series plot of the coefficient estimates. In support of the parallel trends assumption, the estimates show that the coefficients on *POST-LIBERALIZATION (-1)* and *POST-LIBERALIZATION (-2)* are both statistically and economically insignificant. Furthermore, the coefficients on liberalization shock dummies for the year of the shock and after the shock continue to be significant, suggesting that liberalization shocks are associated with persistent declines in the cost of capital.

Finally, Table 4, Panel B and Figure 3 presents evidence on the timing of changes in the price of information asymmetry around the liberalization shocks. It can be seen that the coefficients on *POST-LIBERALIZATION(-2)*BIDASK* and *POST-LIBERALIZATION(-1)*BIDASK* are both statistically and economically insignificant. This suggests that treatment and control firms do not exhibit any differential trends in their pricing of information asymmetry prior to a liberalization shock. It can also be seen that our DiD estimate of the decline in the price of information asymmetry is persistent and remains statistically and economically significant in the years after the liberalization. Specifically, the coefficient on *BIDASK *POST-LIBERALIZATION(0)* is -0.376 (*t*-stat; = -2.678), which is very similar to the coefficient on *BIDASK*POST-LIBERALIZATION(>4)* of -0.311 (*t*-stat = -2.395). These results help rule out the possibility that temporary price pressures drive our results (Harris and Guerel, 1986; Shleifer, 1986).

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data to construct reliable tests of parallel trends using dummies for years prior to *POST-LIBERALIZATION (-2)*. Consider a liberalization shock to country *A* in 1999. For country *A*, we would have only 2 years of pre-shock data: 1997 (*POST-LIBERALIZATION(-2)*) and 1996 (*POST-LIBERALIZATION(-1)*). As a result, shocks in 1999 or before, cannot contribute to the identification of coefficient on *POST-LIBERALIZATION(-3)*. Coefficient on *POST-LIBERALIZATION (-3)*, if included in the regression, would only be informed by 3 of the 13 shocks that occur after 1999. As an imperfect solution, in the internet appendix, we examine pre-trends for two additional years in cost of capital movements (and not pricing of information asymmetry) by expanding our sample to include years 1995 and 1994. We find that the coefficient on all pre-liberalization dummies is not significantly different from zero in this analysis.

4.3. *Sensitivity to key research design choices*

We next explore the robustness of our main results on the pricing of information asymmetry to four key research design choices: fixed effects structure, measurement of *BIDASK*, measurement of liberalization shocks, and measurement of cost of capital. Table 5, Panel A presents the robustness of our results to use of three alternative fixed effect structures. The results in columns (1) and (2) show that our findings are robust to use of two relatively simpler fixed structures that combine (i) country and year fixed effects and (ii) country, year, and industry fixed effects. We next examine the robustness of our findings to augmenting the fixed structure we employ in our main analyses (i.e., country and industry-year interactive fixed effects) with firm fixed effects to absorb the effect of any time-invariant firm characteristics. The results in column (3) show that our findings are robust under these specifications.

In Panel B of Table 5, we explore the robustness of our findings to different ways of measuring *BIDASK*. First, we control for skewness in the *BIDASK* measure by examining the robustness of our results to the use of ranked *BIDASK* measures (median, quartile, quintile, and decile ranks). The results in columns (1) - (4) reveal that our results continue to hold to the use of these alternative measures. We next attempt to better isolate the information asymmetry component of the *BIDASK* spread in addition to taking the skewness into account. We do so employing a two-step approach. In the first-step, we regress the natural logarithm of *BIDASK* spread on the natural logarithm of lagged values of market capitalization, stock return volatility, and share turnover.[16] The natural logarithm of the *BIDASK* accounts for the skewness while regressing on measures of market capitalization, stock return volatility, and share turnover helps isolate the information asymmetry component of the *BIDASK*. We then use the residual from this first-step regression (*RESIDUAL_BIDASK*) as an alternative measure of information

[16] Consistent with prior literature (e.g., Christensen et al. 2017), the first stage regression produces a negative and significant coefficient on the logged values of lagged market value of equity and share turnover and positive (although insignificant) coefficient on the logged value of lagged stock return volatility.

asymmetry (Daske et al., 2008) and re-estimate Equation (4) as a second stage. *RESIDUAL_BIDASK* has a mean of zero (by construction) and has a standard deviation of 0.017. The results in column (5) from the second stage show that our inferences are robust to the use of this alternative measure of information asymmetry.

Next, we examine the robustness of our results to several alternative ways of measuring the liberalization shock. The results in Panel C of Table 5 show that our inferences are robust to using many alternative measures of the shock including a top quintile shock (column (1)), top median shock (column (2)), and a shock that is twice the median shock (column (3)). We also examine the sensitivity of our results to using a decile ranked measure of the changes in the continuous liberalization measure and the results in column (4) show that our inferences hold.

Finally, we present several robustness checks that are related to the cost of capital estimations. First, we examine whether our findings hold for each of the four individual implied cost of capital models. Table 5, Panel D presents the results using estimates from each of the four models. The results are consistent with our findings in Table 3. Specifically, across all four models, the coefficients for *BIDASK* are positive and significant, while the coefficients for *BIDASK*POST-LIBERALIZATION* are negative and significant. Moreover, the coefficients have similar magnitudes to those in Table 3 across all models. These results suggest that our findings do not depend on the choice of a particular cost of capital model.

Next, we gauge the sensitivity of our results to the use of risk premiums instead of raw cost of capital estimates. We calculate risk premium as *COC* minus the risk-free rate on the U.S. Treasury bill. The results in column (1) of Table 5, Panel E show that using a risk premium does not alter our findings from Table 3. Finally, we consider the robustness of our results to the use of alternative assumptions about the long-run growth rate that affects terminal value. Changing assumptions about growth beyond the explicit forecasting horizon affects only two of the four valuation models: Claus and Thomas (2001) and Ohlson and Juettner-Nauroth (2005). In line with Hail and Leuz (2006, 2009), we consider two alternative assumptions about long-run growth. First, we set the long-run growth rate equal to a country's annual real GDP growth rate plus its long-run inflation rate, where we measure the latter as the median inflation over the sample period. Second, we assume a constant inflation rate of 3% across countries. While the first specification allows for perpetual differences in growth rate across countries, the second specification assumes that growth rates in equilibrium converge to zero real growth rate (or 3% nominal growth rate). The results in columns (2) and (3) of Table 5, Panel E show that changing the long-run growth rate assumptions makes little difference to our results.

Overall, the results in Table 5 provide robust evidence that removal of restrictions on foreign ownership results in a decrease in the pricing of information asymmetry in our sample.

4.4. *Could concurrent growth shocks and/or other reforms drive our results?*

We show that the treatment and control countries exhibit parallel movements in their price of information asymmetry in the years leading up to the liberalization shocks. Nevertheless, an important concern in any DiD setting is whether the parallel trends would have continued in the post-treatment period absent any changes in restrictions on foreign ownership. This concern affects our setting because liberalizations typically accompany expansions in growth prospects, as well as other reforms such as macroeconomic stabilization, easing of exchange controls, and trade liberalization. Therefore, to the extent our specifications do not fully absorb the effect of growth opportunities and other concurrent reforms, possibility exists that our inferences could be confounded by these other concurrent events. In this subsection, we provide two analyses to address this concern.

4.4.1. Placebo tests based on opening of other financial markets

We conduct a series of placebo tests by examining the effect of liberalization of four financial markets other than equity markets: money markets, bond markets, derivative markets, and government securities markets. The liberalization of these other markets provides compelling placebos for two reasons. First, similar to equity markets, the desire to attract foreign capital flows drives the liberalization of these markets. Therefore, similar to equity markets, the liberalization for these markets would also be expected to be accompanied by expansions in growth prospects and other reforms that tend to accompany general liberalizations of financial markets. Second, there is no reason to expect that increased external capital flows in bonds, money market instruments, derivatives, and/or government securities markets will have a bearing on the pricing of information asymmetry in equity markets. To identify liberalization shocks in these alternative markets, we use data from Fernández et al. (2016) and Schindler (2009). These authors build a binary measure of liberalization in various capital markets based on information in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER). For each financial market, the data provides a binary outcome variable based on the authors' assessment of the regulatory restrictions. We create dummy variables for the opening up of these alternative financial markets based on these binary measures and re-estimate equation (4) using these placebo shocks.[17]

We present the results in columns (1)-(4) of Table 6. The results show that the effect of opening of these markets on price of information asymmetry is economically and statistically indistinguishable from zero.[18] This suggests that any changes in countries' growth prospects

[17] We identify the liberalization shock based on the first year in our sample period when the binary variable corresponding to purchase restrictions by non-residents switches from restricted (one) to unrestricted (zero) in the respective markets.

[18] In the Internet Appendix, we also present results of a falsification test in the spirit of Altonji et al. (2005) to gauge the extent to which concurrent local growth/economic shocks could be driving our results. Christensen et al. (2016) and Christensen et al. (2017) have employed this strategy to rule out concurrent confounds. Specifically, under this approach, we first regress our implied cost of capital measures on a set of three proxies based on prior work (e.g., Bekaert and Harvey, 2000) that are expected to be strongly correlated with local growth/economic shocks: GDP growth, stock market valuation as a fraction of GDP, and trade as a percentage of GDP. We use this first-stage regression to create a predicted value for the implied cost of capital based on these variables. We use the predicted value as the dependent variable in our main DiD specification used to measure the effect of liberalization on the cost of capital. We find that the coefficient on *POST-LIBERALIZATION*BIDASK* is statistically and economically indistinguishable from zero, which suggests that our main specifications are effective at absorbing the confounding effect of concurrent growth/other economic shocks.

or other reforms that tend to accompany a general opening up of financial markets are unlikely to drive our main results.

4.4.2. Within country identification using ADRs

In our second analysis to address this concern, we exploit purely within-country variation to estimate the effect of liberalization. By holding the country constant, this analysis mitigates concerns that omitted country-level factors could be driving our results. This analysis is based on the observation that the effect of liberalization within a country should be relatively muted for firms that already have access to foreign capital through listing of American Depositary Receipts (ADRs). For this test, we first identify firms that have outstanding ADRs at the beginning of a fiscal year. We obtain information on cross listing from a variety of sources including the Bank of New York, J.P. Morgan, Citibank, the NYSE, the NASDAQ, and CRSP database. We include both active and inactive listings using the data provided by Citibank and CRSP to mitigate concerns related to survivorship bias. Our sample of cross-listings includes exchange listings on the NYSE, NASDAQ, and AMEX, as well as over-the-counter (OTC) listings in the Pink Sheets or the OTC Bulletin Board, and private placements under Rule 144A.

Columns (5) and (6) of Table 6 present the results of equation (4) separately for subsamples of firms that are cross-listed and those that are not. It can be seen that the liberalization shocks reduce the price of information asymmetry only for firms without any ADRs. Specifically, while the coefficient on the interaction term between *BIDASK* and the liberalization shock dummy is negative and significant for the subgroup of firms without an ADR (coefficient = -0.367; *t*-stat = 2.480), it is positive and not significantly different from zero for the ADR group (coefficient = 0.185; *t*-stat = 1.184). A potential concern is that lack of statistical power due to the smaller sample size drives the insignificant results for the ADR subgroup. Inconsistent with this explanation, however, we find that the coefficient on *BIDASK* is positive and significant at 1% level for both subsamples and it is only the effect of liberalization that is insignificant in the ADR group. Furthermore, if it is purely statistical power that is driving the difference, then we would expect the magnitude of the DiD estimates to be similar across the two groups; we, however, do not find that to be the case with a coefficient of 0.185 for the ADR group and -0.367 for the group without ADRs.[19]

4.5. *Role of home country institutions*

In this subsection, we explore how the quality of home country institutions affects the benefits of liberalization on the pricing of information asymmetry. The impact of institutional quality is not obvious. On the one hand, because poor institutional infrastructure and corporate governance can thwart local investor participation in equity markets, countries with poor institutions are more likely to approximate imperfect capital markets prior to the liberalization. Consequently, such countries may benefit more from the addition of foreign investors following liberalization. One the other hand, poor quality local institutions can act as implicit barriers to economic integration even if a country removes explicit regulatory barriers through equity market liberalization (e.g., Bae and Goyal, 2010; Carrieri et al., 2013; Stulz, 2005). For example, foreign investors may not feel comfortable investing in countries with poor foreign investor protection even if the rules permit their participation. In this scenario, poor local institutional quality may reduce the benefits of liberalization by preventing foreign investor participation.

In Table 7, we explore the role of several features of a country's institutional environment. We partition the sample into subgroups based on institutional characteristics, and

[19] We discuss an additional within country test in the Internet Appendix. This analysis is based on the observation that most country-level liberalization policies explicitly leave out regulated industries (Bekaert et al., 2007). Thus, we expect the effect of liberalization shocks to be muted for firms in regulated industries. However, an important empirical issue is in identifying these regulated industries, which tend to differ across countries. Nevertheless, using two alternative definitions of regulated industries, we find that the decline in the pricing of information asymmetry following liberalization is concentrated only in unregulated industries and is insignificant in the subsample of regulated industries.

run separate regressions for each subsample. The regressions generating the results in Table 7 have the same control variables as for the regressions for Table 3, but we do not report the coefficient estimates on these variables for brevity. We focus on how the coefficients for *BIDASK* and *POST-LIBERALIZATION*BIDASK* vary across subsamples with different country characteristics.

We focus on the following variables to measure variation in institutional quality and governance. First, following Hail and Leuz (2006), we measure the *strength of securities regulation* as the mean of the disclosure index, the liability standard index, and the public enforcement index. We use a dummy variable, which that takes a value of one if the country has a value on the index that is above the median across all countries, and zero otherwise. All securities regulation variables stem from La Porta et al. (2006). Second, we measure the *extent of insider trading* activities using a dummy variable that equals one following the year when insider trading laws were first enforced (Bhattacharya and Daouk, 2002; Bushman et al., 2005). In countries where insider trading laws are not enforced, the likelihood that insiders have an informational advantage is higher. Third, we measure the *extent of investor protection* using the investor protection index from La Porta et al. (2006). Fourth, we measure the *extent of insider ownership* 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 (Bae et al., Tan, 2008; La Porta et al., 1999). We define a dummy variable that takes a value 1 (*High*) if the ownership by insiders is above-median and 0 (*Low*), otherwise. Higher insider ownership implies there is more potential for poor corporate governance and information-based trades. Fifth, we examine the effect of the *origin of the legal system* of a country. La Porta et al. (1998) point out the importance of legal origin in explaining the economic and financial institutions in a country and argue that English common law provides better protection of individual rights against the state. Thus, we employ origin of the legal system as a dummy variable that is equal to one if the country is of English law origin and zero otherwise. Finally, we employ an overall measure of *accountability of institutions* from Kurtzman et al. (2004), which is a country level index reflecting a country's level of corruption, efficacy of the legal system, deleterious economic policies, inadequacy of accounting and governance practices, and detrimental regulatory structures We construct this index following the methodology of PricewaterhouseCoopers (2001).[20]

Estimates in Table 7 show that across all specifications the benefits of liberalization through reductions in the pricing of information asymmetry are concentrated in countries with poor institutions and corporate governance prior to the liberalization shocks. For example, the coefficient on the *POST-LIBERALIZATION*BIDASK* is negative and significant for countries where securities regulations are poor (coefficient = -0.384) and there is low enforcement of insider trading laws (coefficient = -0.270), while the coefficient is negative but economically small and statistically insignificant for the countries where the securities regulations are good and insider trading laws are better enforced. Further, in countries where there is civil law the impact of a large liberalization event (coefficient = -0.407) is about 60% greater than countries with common law (coefficient = -0.255). These findings suggest that the benefits of liberalization in terms of the pricing of information asymmetry accrue largely to countries that have imperfectly competitive capital markets due to poor institutional infrastructures.

5. Conclusion

Using the equity market liberalization of 23 emerging market countries between 1996 and 2006, we examine the effect of expanding an economy's investor base to include foreign

[20] As noted in Kurtzman et al. (2004), the accountability index draws upon 65 objective variables from 41 sources including the World Bank, the International Monetary Fund, the International Securities Services Association, the PRS group, the International Country Risk Guide, and regulators in individual countries. The index uses the variables to create five sub-indices measuring a country's level of corruption, efficacy of the legal system, deleterious economic policies, inadequacy of accounting and governance practices, and detrimental regulatory structures. The accountability index is the simple average of these five sub-indices. Other studies that use this measure include Boubakri et al. (2013), Brouthers et al. (2008), and Guedhami and Pittman (2006).

investors on the pricing of information asymmetry. We find evidence of a significant decline in the pricing of information asymmetry as countries remove regulatory restrictions on foreign ownership. We also find that this benefit of liberalization accrues primarily to firms that do not already have access to foreign capital through ADRs and to economies that have poor institutional infrastructures and corporate governance.

Our study addresses a significant challenge associated with measuring aggregate competition in trying to assess its effect on the pricing of information asymmetry. The firm-level measures of competition used in extant studies implicitly assume that investors who decline to hold stock do not contribute to the overall risk-bearing capacity and, hence, do not influence the price of risk. However, the mere expectation that these investors might step in to purchase shares if the price of risk were to fall sufficiently low allows these investors to affect share prices without necessarily holding the stock. We address this challenge by measuring shocks to risk-bearing capacity at the economy level. Furthermore, the plausibly exogenous nature of the liberalization shocks also allows us to mitigate concerns that are associated with using endogenous variation in competition based on firm-level shareholding patterns. Finally, our findings also enhance our understanding of the economic consequences of increased foreign investor participation by examining its effect on pricing of information asymmetry.

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Appendix

A. Implied cost of equity capital models

All of the four methods we examine are derived from same underlying valuation model, i.e., the dividend discount model. The Claus and Thomas (2001) and Gebhardt et al. (2001) models are based on the residual income valuation models and specify valuation using return on equity. Residual income is defined as forecasted earnings per share less a cost of capital charge for the beginning of the fiscal year book value of equity. Both models assume clean surplus accounting which requires that earnings are fully allocated between dividends and retained earnings (i.e., whatever portion of earnings that is not paid out in dividends) is added to the book value of equity.

Claus and Thomas (2001)

The valuation equation in Claus and Thomas (2001) is given by:

$$P_t = bv_t + \sum_{k=1}^{T} \frac{(eps_{t+k} - r_{CT} * bv_{t+k-1})}{(1+r_{CT})^k} + \frac{(eps_{t+T} - r_{CT} * bv_{t+T-1})(1+g)}{(r_{CT} - g)(1+r_{CT})^T},$$

where P_t is the market price of a firm's stock at time t, bv is the book value per share at the beginning of the fiscal year, eps is the expected future earnings per share using either explicit analyst forecasts or derived from growth forecasts, g is the expected perpetual future growth rate, and r_{CT} is the implied cost of equity capital that is calculated as the internal rate of return of the above equation. In the model, we use actual book values per share and forecasted earnings per share up to five years ahead to derive the expected future residual income series. Dividends are set equal to a constant fraction of forecasted earnings. At time $T = 5$, it is assumed that the nominal residual income grows at rate g. As a proxy for g, we use the annualized median of country-specific one-year ahead realized monthly inflation rates.

Gebhardt et al. (2001)

The valuation equation in Gebhardt et al. (2001) is given by:

$$P_t = bv_t + \sum_{k=1}^{T} \frac{(eps_{t+k} - r_{GLS} * bv_{t+k-1})}{(1+r_{GLS})^k} + \frac{(eps_{t+T+1} - r_{GLS} * bv_{t+T})}{r_{GLS} * (1+r_{GLS})^T},$$

where P_t is the market price of a firm's stock at time *t*, *bv* is the book value per share at the beginning of the fiscal year, *eps* is the expected future earnings per share using either explicit analyst forecasts or derived from growth forecasts, *g* is the expected perpetual future growth rate, and r_{GLS} is the implied cost of equity capital that is calculated as the internal rate of return of the above equation. In this model, we use explicit forecasts of earnings for the first three years, followed by a nine-year period in which the return on equity (ROE) linearly reverts to the industry median ROE [based on Campbell (1996)]. The industry median ROE is calculated using the past three years of data in a given country. From T = 12, the residual income is assumed to be constant.

Ohlson and Juettner-Nauroth (2005)

The valuation equation in Ohlson and Juettner-Nauroth (2005) is given by:

$$P_t = \frac{d_{t+1}}{(r_{OJN} - g_l)} + \frac{eps_t(g_s - g_l)}{r_{OJN}(r_{OJN} - g_l)},$$

where P_t is the market price of a firm's stock at time *t*, *eps* is the expected future earnings per share using either explicit analyst forecasts or derived from growth forecasts, g_s is the expected short-term future growth rate, *d* is the expected future net dividends per share derived from the dividend payout ratio times the earnings per share, g_l is the expected long-term future growth rate, and r_{OJN} is the implied cost of equity capital that is calculated as the internal rate of return of the above equation. Gode and Mohanram (2003) implement this theoretical model of Ohlson and Juettner-Nauroth (2005) by assuming that the short-term growth rate is equal to the average of the forecasted growth rate between year one and year two and the average five-year growth rate provided by analysts. We adopt the same approach. Furthermore, Gode and Mohanram assume that the long-term growth rate is equal to expected inflation for all firms. We use the annualized median of country-specific one-year ahead realized monthly inflation rates.

Easton (2004)

The valuation equation in Easton (2004) for the modified price-earnings growth (PEG) ratio model is:

$$P_t = \frac{r_{MPEG} * d_{t+1} + eps_{t+2} - eps_{t+1}}{r_{MPEG} * r_{MPEG}},$$

where P_t is the market price of a firm's stock at time *t*, *eps* is the expected future earnings per share using either explicit analyst forecasts or derived from growth forecasts, g_s is the expected short-term future growth rate, *d* is the expected future net dividends per share derived from the dividend payout ratio times the earnings per share, and r_{MPEG} is the implied cost of equity capital that is calculated as the internal rate of return of the above equation. It uses one-year ahead and two-year ahead earnings per share forecasts, as well as expected dividends per share in period $t + 1$ to derive a measure of abnormal earnings growth. Dividends are set equal to a constant fraction of forecasted earnings. In this model, we assume that the growth rate in the change in dividends is equal to zero so that dividends grow by the same dollar amount every year into perpetuity.

Data requirements and implementation

We closely follow Hail and Leuz (2006, 2009) in the implementation of these implied cost of capital models for an international sample of firms. We obtain information on all dead/inactive stocks and include them in the sample. Inclusion of inactive stocks ensures that we do not have survivorship bias. We obtain firm-specific stock price and analyst earnings per share forecasts and long-term growth forecasts from IBES. All estimates are mean analyst consensus forecasts and measured in local currency. For an observation to be included in our sample, we follow Hail and Leuz (2009) and apply the same filters. We require the current stock price data, analyst earnings per share for two periods ahead, and either a forecasted earnings per share for the three-years ahead or an estimate of long-term earnings growth.

Further, we use firms that only have positive earnings forecasts and growth rates.

We gather financial data from Worldscope. Each of the four valuation models require an estimate of future dividends. Following Hail and Leuz (2009), we assume that the net dividends are a constant fraction of expected future earnings per share for all periods. The dividend ratio required for this calculation is obtained by averaging the ratios over the last three fiscal years. If this ratio is missing or outside the range of zero and one, we replace it by the median payout ratio for the country-year. The Gebhardt et al. (2001) model requires an industry-specific target return on equity. Following Hail and Leuz (2009), we first compute the average firm-level return on equity ratios over the last three years and then choose the median of these values for a given industry, country, and year. We replace missing or negative target ratios by the country/industry median and, if still missing or negative, by the country/year median.

These models also require assumptions about long-term growth. Following Hail and Leuz (2009), we assume that, in the long-run, firms grow at the country's inflation rate and use next year's country-specific median of the realized monthly percentage changes in the Consumer Price Index as a proxy for future inflation. We replace negative values by the country's historical inflation rate, computed as the median of the monthly inflation rates over the sample period. Similarly, we replace values exceeding 10% by the country's historical inflation rate.

We measure the cost of capital as of month +10 after the end of the fiscal year. This ensures that the market price and analyst expectations reflect financial statement information at the time of measurement. Valuation models, however, assume full year discounting. Therefore, following Hail and Leuz (2009), for consistent discounting, we first move month+10 prices to the beginning of the fiscal year and then use full year discounting. Since the valuation models do not have a unique closed-form solution, we use an iterative procedure to determine the annual firm-specific discount rate that equates market price to the respective valuation models. We restrict our sample to firm-year observations where we were able to obtain estimates for all the four models. Moreover, to be included in the regression models in a given year, we require that a country has at least ten firms with all non-missing data.

Figure 1: Liberalization and the Investability Measure

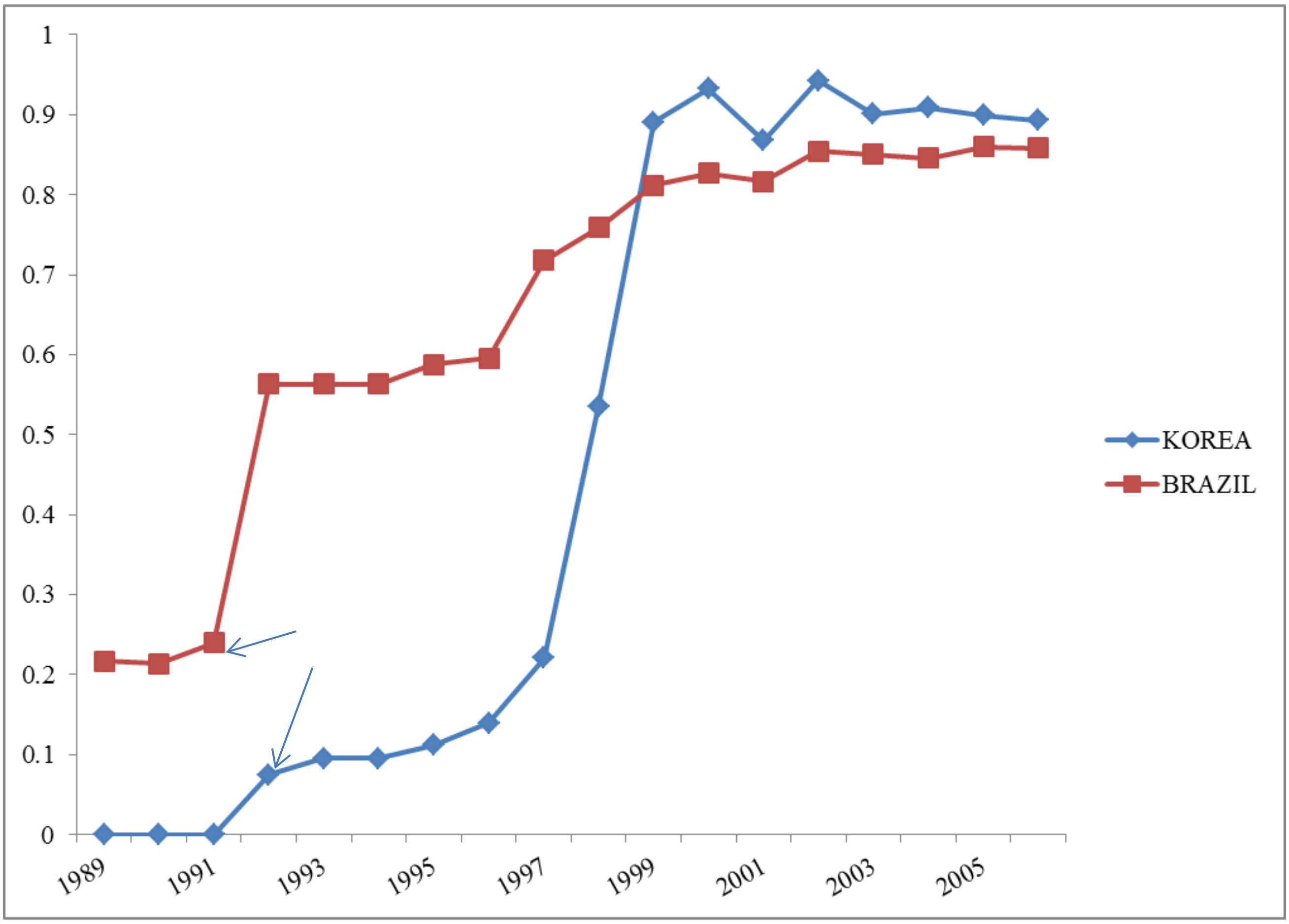

Figure 2: Timing of changes in cost of capital

This figure presents the timing of changes in cost of capital around liberalization shocks. ● represents coefficients that are insignificant while ■ represents coefficients that are significant. The bars represent the 90% confidence interval around the estimated coefficients. The horizontal axis presents the years relative to the shock year (t=0).

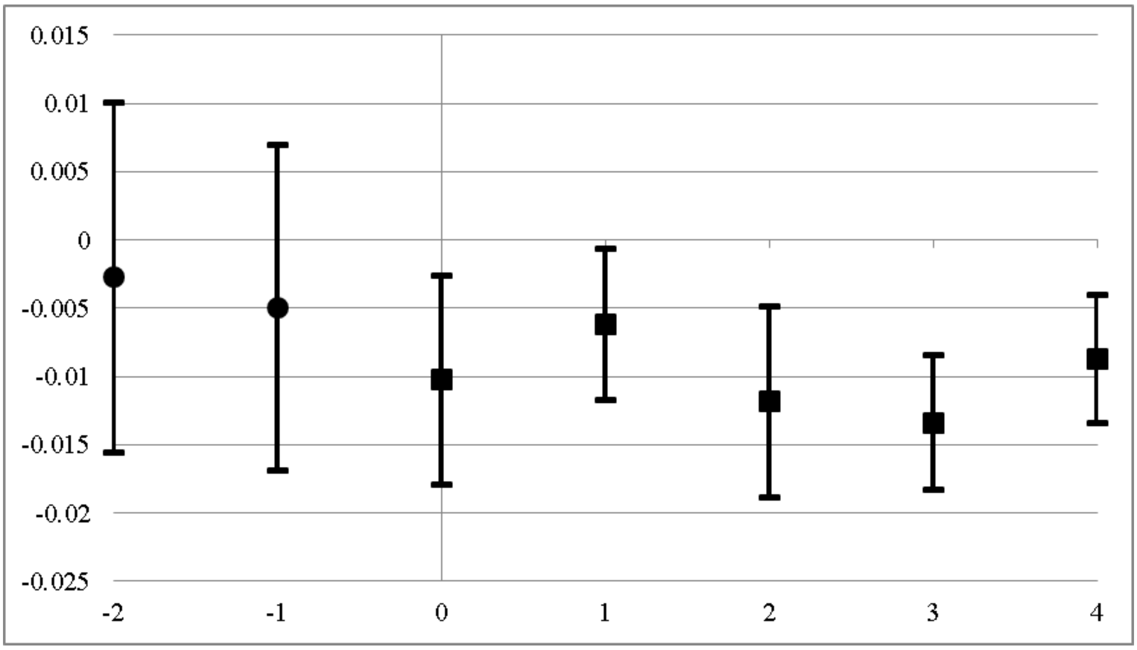

Figure 3: Timing of changes in pricing of information asymmetry

This figure presents the timing of changes in pricing of information asymmetry around liberalization shocks. ● represents coefficients that are insignificant while ■ represents coefficients that are significant. The bars represent the 90% confidence interval around the estimated coefficients. The horizontal axis presents the years relative to the shock year (t=0).

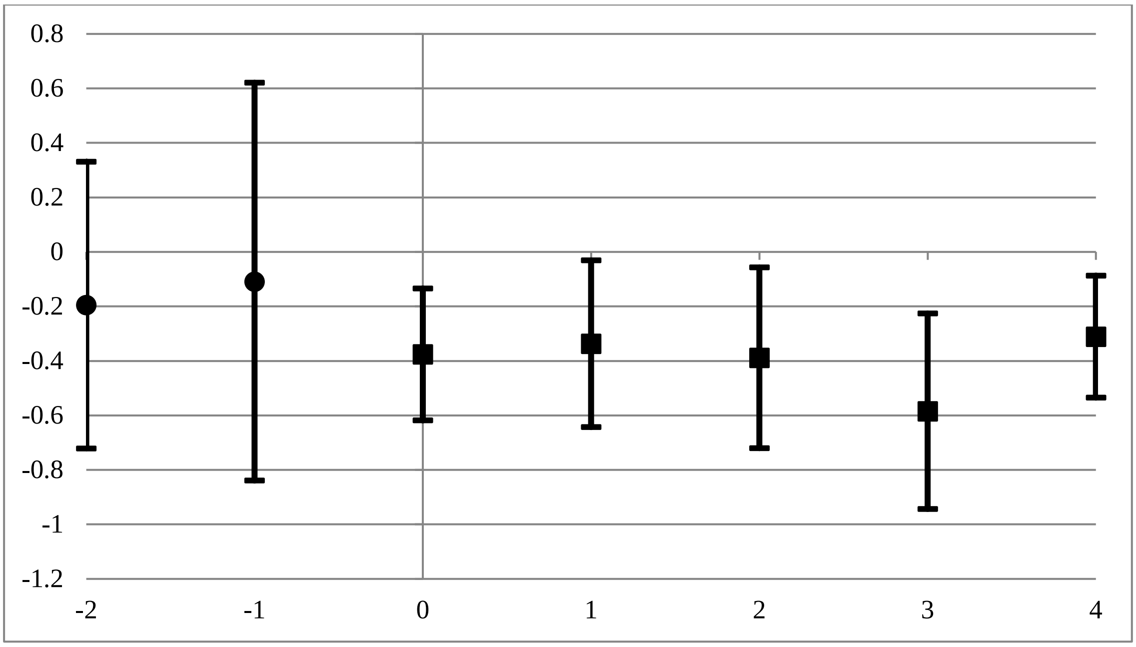

Table 1: Descriptive Statistics on Liberalization Measures

Country	Median	Mean	Std. Dev.	Official Liberalization Year based on Bekaert & Harvey (2000)	First firms investable from Mitton (2006)	First Year of Top Tercile Shock in *BIDASK* Pricing Sample (1996-2006)
Argentina	0.9	0.9	0.001	1989	1988	None
Brazil	0.86	0.85	0.014	1991	1988	1999
Chile	0.86	0.87	0.033	1992	1988	None
China	1.39	1.36	0.25	n/a	1992	1997
Czech Republic	0.38	0.44	0.107	n/a	1994	1999
Egypt	0.65	0.67	0.093	1992	n/a	1998
Greece	1	0.96	0.077	1987	1988	2001
Hungary	0.69	0.66	0.035	n/a	1992	1996
Indonesia	0.8	0.69	0.188	1989	1990	1997
India	0.29	0.42	0.212	1992	1992	2002
Israel	1	0.99	0.006	1993	1997	None
Korea, Republic of	0.9	0.84	0.206	1992	1992	1998
Sri Lanka	0.28	0.28	0.01	1991	1992	None
Mexico	0.69	0.68	0.023	1989	1988	None
Malaysia	0.91	0.84	0.215	1988	1988	1997
Pakistan	0.65	0.51	0.279	1991	1991	None
Philippines	0.53	0.57	0.092	1991	1988	2004
Poland	0.98	0.98	0.013	n/a	1992	None
Portugal	1	0.9	0.125	1986	1988	1998
Taiwan	0.55	0.59	0.231	1991	1991	None
Thailand	0.61	0.56	0.112	1987	1988	1998
Turkey	0.97	0.96	0.018	1989	1989	None
South Africa	1.01	1.01	0.007	1996	1992	None

Table 2: Descriptive Statistics

This table presents descriptive statistics for the measures of cost of capital and control variables. The sample comprises a maximum of 8,474 firm-year observations representing 23 emerging market countries over the 1996 to 2006 period. r_{GLS} is the implied cost of equity capital derived based on the Gebhardt et al. (2001) model. r_{CT} is the implied cost of equity capital derived based on the Claus and Thomas (2001) model. r_{OJN} is the implied cost of equity capital derived based on the Ohlson and Juettner-Nauroth (2005) model. r_{MPEG} is the implied cost of equity capital derived based on the Easton (2004) model. *COC* is the average cost of capital estimate implied by the four models. *INFLATION* is the rate of inflation in the firm's home country. *BIDASK* represents the annual average of the monthly median bid-ask spreads in equity prices. *SIZE* is the firm's total book value of assets measured at the end of the fiscal year. *RETVAR* is the stock return variability measured as the standard deviation of monthly stock returns over the last 12 months. *LEVERAGE* is the ratio of total liabilities to total assets. *FBIAS* equals the one year ahead IBES analyst forecast error (mean forecast for the next fiscal year minus actual earnings) scaled by lagged total assets.

	N	Mean	Median	25th Percentile	75th Percentile	Std. Dev
r_{GLS}	8474	0.120	0.111	0.078	0.151	0.064
r_{CT}	8474	0.149	0.130	0.101	0.172	0.081
r_{OJN}	8474	0.168	0.152	0.122	0.194	0.075
r_{MPEG}	8474	0.162	0.145	0.111	0.190	0.085
COC	8474	0.155	0.136	0.111	0.173	0.059
BIDASK	5694	0.014	0.008	0.004	0.016	0.018
INFLATION	8474	0.038	0.036	0.020	0.053	0.023
SIZE (US$ MN)	8474	3412	509	183	1822	9128
RETVAR	8474	0.114	0.101	0.071	0.142	0.060
LEVERAGE	8474	0.514	0.509	0.358	0.654	0.220
FBIAS	8474	0.000	0.000	0.000	0.000	0.002

Table 3: Equity Market Liberalization and the Pricing of Information Asymmetry

This table presents evidence on the effect of equity market liberalization on cost of capital. The sample comprises 23 emerging market countries for the period 1996 to 2006. The dependent variable COC is the mean of four estimates for the implied cost of equity capital (see Appendix). *POST-LIBERALIZATION* is a dummy that takes a value one starting the year a country experiences a change in the restrictions on foreign ownership that is in the top tercile of changes across all country-year changes and zero otherwise. The results in column (1) are for the full sample while in column (2) the results are for a subsample of firms that have non-missing values for *BIDASK*. All other variables are defined in Table 2. Specifications include industry-year interactive fixed effects and country fixed effects. Standard errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5%, and 1% level is denoted by *, **, and ***, respectively.

VARIABLES	Predicted Sign	(1) *COC*	(2) *COC*	(3) *COC*
POST-LIBERALIZATION	**-**	**-0.006*** **	**-0.007*** **	**-0.00417** **
		(-2.965)	**(-3.811)**	**(-2.170)**
BIDASK	**+**			**0.413*** **
				(3.450)
BIDASK*POST-LIBERALIZATION	**-**			**-0.312*** **
				(-2.874)
INFLATION	+	0.248**	0.205*	0.215**
		(2.732)	(1.935)	(2.119)
LOGSIZE	-	-0.006***	-0.00692***	-0.00634***
		(-3.422)	(-3.779)	(-4.083)
RETVAR	+	0.102***	0.0756***	0.0747***
		(4.101)	(4.286)	(4.278)
LEVERAGE	+	0.0469***	0.0487***	0.0480***
		(6.418)	(6.011)	(6.170)
FBIAS	+	1.873**	1.487*	1.495**
		(2.326)	(2.080)	(2.095)
Hypothesis testing:				
*BIDASK+BIDASK*POST-LIBERALIZATION*	*Difference*			*0.101**
	t-stat			*(1.797)*
Observations		8,474	5,694	5,694
R-squared		0.296	0.327	0.334
Country Fixed Effects		Included	Included	Included
*Year*Industry Fixed Effects*		Included	Included	Included

Table 4: Timing of Changes in Cost of Capital

This table presents evidence on the timing of the cost of capital changes around liberalization shocks. *POST-LIBERALIZATION (-2)* and *POST-LIBERALIZATION (-1)* are dummy variables that take a value one in the period two-year prior and one-year prior to year of liberalization announcement and zero otherwise. *POST-LIBERALIZATION (0)*, *POST-LIBERALIZATION (1)*, *POST-LIBERALIZATION (2)*, and *POST-LIBERALIZATION (3)* are dummy variables that take a value one in the year of, one-year after, two-year after, and three-year after a country experiences a liberalization shock. *POST-LIBERALIZATION (>4)* is a dummy that takes a value one if a country has experienced a liberalization shock four or more years ago. All other variables have been defined in the caption of Table 2. Specifications include industry-year interactive fixed effects as well as country fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

Panel A: Timing of Changes of Baseline Result

VARIABLES	Predicted Sign	*COC* Coefficient	*t-stat*
POST-LIBERALIZATION (-2)	n/a	-0.003	(-0.371)
POST-LIBERALIZATION(-1)	n/a	-0.005	(-0.720)
POST-LIBERALIZATION(0)	-	-0.010**	(-2.317)
POST-LIBERALIZATION(+1)	-	-0.006*	(-1.922)
POST-LIBERALIZATION (+2)	-	-0.012***	(-2.910)
POST-LIBERALIZATION(+3)	-	-0.013***	(-4.696)
POST-LIBERALIZATION(>4)	-	-0.009***	(-3.190)
Controls		Included	
Country Fixed Effects		Included	
*Year*Industry Fixed Effects*		Included	

Panel B: Timing of Changes of Pricing of Information Asymmetry

VARIABLES	Predicted Sign	*COC* Coefficient	*t-stat*
BIDASK	+	0.418***	(3.327)
*BIDASK*POST-LIBERALIZATION (-2)*	n/a	-0.195	(-0.638)
*BIDASK*POST-LIBERALIZATION(-1)*	n/a	-0.109	(-0.258)
*BIDASK*POST-LIBERALIZATION(0)*	-	-0.376**	(-2.678)
*BIDASK*POST-LIBERALIZATION(+1)*	-	-0.337*	(-1.900)
*BIDASK*POST-LIBERALIZATION (+2)*	-	-0.388*	(-2.019)
*BIDASK*POST-LIBERALIZATION(+3)*	-	-0.585**	(-2.810)
*BIDASK*POST-LIBERALIZATION(>4)*	-	-0.311**	(-2.395)
Main effects and Controls		Included	
Country Fixed Effects		Included	
*Year*Industry Fixed Effects*		Included	

Table 5: Robustness Checks

This table presents evidence on the robustness of the findings to research design choices. All variables have been defined in the captions of Tables 2 and 3 and in Section 4.3. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

Panel A: Robustness to Alternative Fixed Effects

	(1)	(2)	(3)
VARIABLES	*COC*	*COC*	*COC*
POST-LIBERALIZATION	-0.00284	-0.00308	0.00581
	(-0.957)	(-1.041)	(0.466)
BIDASK	0.398***	0.406***	0.316***
	(3.229)	(3.501)	(4.776)
*BIDASK*POST-LIBERALIZATION*	-0.312**	-0.320***	-0.340***
	(-2.427)	(-2.874)	(-2.946)
Controls	Included	Included	Included
Country Fixed Effects	Included	Included	No
Year Fixed Effects	Included	Included	No
Industry Fixed Effects	No	Included	No
*Industry*Year Fixed Effects*	No	No	Included
Firm Fixed Effects	No	No	Included

Panel B: Robustness to alternative measurements of BIDASK

	(1)	(2)	(3)	(4)	(5)
	BIDASK QUARTIILE	*BIDASK MEDIAN*	*BIDASK QUINTILE*	*BIDASK DECILE*	*RESIDUAL LOG_BIDASK*
VARIABLES	*COC*	*COC*	*COC*	*COC*	*COC*
POST-LIBERALIZATION	0.005*	0.002	0.005	0.003	-0.003***
	(1.796)	(0.700)	(1.710)	(1.303)	(-5.044)
BIDASK	0.008***	0.012***	0.007***	0.004***	0.182***
	(5.577)	(3.064)	(5.897)	(5.949)	(3.057)
BIDASK POST-LIBERALIZATION*	-0.006***	-0.007**	-0.005***	-0.002***	-0.121**
	(-4.507)	(-2.507)	(-4.513)	(-4.307)	(-2.133)
Controls	Included	Included	Included	Included	Included
Country Fixed Effects	Included	Included	Included	Included	Included
*Year*Industry Fixed Effects*	Included	Included	Included	Included	Included

Table 5 (Contd.)

Panel C: Robustness to alternative measurement of liberalization shock

	(1)	(2)	(3)	(4)
	QUINTILE SCHOCK	*MEDIAN SHOCK*	*2*MEDIAN SHOCK*	*DECILERANK SHOCK*
VARIABLES	*COC*	*COC*	*COC*	*COC*
POST-LIBERALIZATION	-0.003	-0.000	0.000	0.000
	(-1.667)	(-0.096)	(0.107)	(0.320)
BIDASK	0.355**	0.429***	0.396***	0.524***
	(2.665)	(4.274)	(3.256)	(5.900)
*BIDASK*POST-LIBERALIZATION*	-0.258**	-0.323***	-0.283**	-0.051***
	(-2.457)	(-2.913)	(-2.323)	(-4.045)
Controls	Included	Included	Included	Included
Country Fixed Effects	Included	Included	Included	Included
*Year*Industry Fixed Effects*	Included	Included	Included	Included

Panel D: Robustness to individual cost of capital estimates

	(1)	(2)	(3)	(4)
VARIABLES	r_{GLS}	r_{CT}	r_{OJN}	r_{MPEG}
POST-LIBERALIZATION	-0.002	-0.005**	-0.004	-0.005
	(-0.560)	(-2.321)	(-1.379)	(-1.092)
BIDASK	0.465***	0.358***	0.453***	0.533***
	(3.911)	(3.261)	(3.755)	(3.136)
*BIDASK*POST-LIBERALIZATION*	-0.280**	-0.254**	-0.311***	-0.334**
	(-2.449)	(-2.435)	(-2.948)	(-2.374)
Controls	Included	Included	Included	Included
Country Fixed Effects	Included	Included	Included	Included
*Year*Industry Fixed Effects*	Included	Included	Included	Included

Panel E: Robustness to the use of risk premium measure and long-run growth assumptions

	(1)	(2)	(3)
VARIABLES	*Risk-Premium*	*Alt_COC1*	*Alt_COC2*
POST-LIBERALIZATION	-0.004**	-0.007***	-0.006**
	(-2.214)	(-3.390)	(-2.281)
BIDASK	0.413***	0.393***	0.421***
	(3.460)	(2.863)	(3.461)
*BIDASK*POST-LIBERALIZATION*	-0.311***	-0.226*	-0.273**
	(-2.875)	(-1.756)	(-2.518)
Controls	Included	Included	Included
Country Fixed Effects	Included	Included	Included
*Year*Industry Fixed Effects*	Included	Included	Included

Table 6: Placebo Tests and within Country Identification

The sample comprises 23 emerging market countries for the period 1996 to 2006. The dependent variable *COC* is the mean of four estimates for the implied cost of equity capital (see the Appendix). Columns (1) through (4) present results from employing placebo shocks. *POST-SHOCK* is a dummy variable based on changes in the binary measures related to the opening up of these alternative financial markets as provided by Schindler (2009) and Fernández et al. (2015). Columns (5) and (6) present results of a within-country analysis of subsample for firms with and without ADRs. In Columns (5) and (6) *POST-SHOCK* represents the *POST-LIBERALIZATION* dummy. All control variables are defined in Table 2. All specifications include industry-year interactive fixed effects as well as country fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

	(1)	(2)	(3)	(4)	(5)	(6)
	Money Market Shock	Bond Market Shock	Derivative Market Shock	Government Securities Shock	Non-ADR Sample	ADR Sample
VARIABLES	*COC*	*COC*	*COC*	*COC*	*COC*	*COC*
POST-SHOCK	-0.012***	-0.005	-0.006	0.003	-0.005**	-0.004
	(-3.793)	(-1.630)	(-1.717)	(0.401)	(-2.255)	(-0.533)
BIDASK	0.252*	0.300**	0.267**	0.295***	0.446***	0.241***
	(2.002)	(2.116)	(2.126)	(3.448)	(2.878)	(3.138)
*BIDASK*POST-SHOCK*	0.284	-0.0590	-0.0193	-0.043	-0.367**	0.185
	(0.821)	(-0.692)	(-0.113)	(-0.224)	(-2.480)	(1.184)
Controls	Included	Included	Included	Included	Included	Included
Country Fixed Effects	Included	Included	Included	Included	Included	Included
*Year*Industry Fixed Effects*	Included	Included	Included	Included	Included	Included
Observations	5,694	5,694	5,694	5,694	4,820	874
R-squared	0.332	0.332	0.331	0.330	0.340	0.432

Table 7: Country Characteristics and the Effect of Equity Market Liberalization on the Pricing of Information Asymmetry

This table presents the effect of country characteristics on the relation between liberalization and the pricing of information asymmetry. Country characteristics include variables related to the degree of investor protection, institutional environment and political risk. The sample comprises 23 emerging market countries for the period 1996 to 2006. The dependent variable in all the models is *COC*, which is the mean of four estimates for the implied cost of equity capital (see the Appendix). The regressions include the same control variables as in Table 3 and not reported for brevity. All variables have been defined in the caption of Table 2. All specifications include industry-year interactive fixed effects as well as country fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

		High		**Low**	
Model	Country Characteristics	*BIDASK*	*POST-LIBERALIZATION*BIDASK*	*BIDASK*	*POST-LIBERALIZATION*BIDASK*
1)	Securities Regulation	0.259	-0.0903	0.436**	-0.384**
		(1.881)	(-0.144)	(3.093)	(-2.725)
2)	Enforcement of Insider Trading Laws	0.190	-0.165	0.474***	-0.270**
		(0.997)	(-0.743)	(4.976)	(-2.626)
3)	Investor Protection	0.161*	0.0703	0.461***	-0.399***
		(3.393)	(0.115)	(3.557)	(-3.943)
4)	Insider Ownership	0.465**	-0.392*	0.281**	-0.152
		(2.733)	(-2.380)	(2.678)	(-1.036)
5)	Legal Origin	0.335*	-0.255	0.465**	-0.407**
	(High=Common Law; Low=Civil Law)	(2.208)	(-1.362)	(2.495)	(-2.897)
6)	Accountability of Institutions	0.376*	-0.269	0.376***	-0.291**
		(2.595)	(-1.809)	(3.977)	(-2.589)

INTERNET APPENDIX

The following Appendix provides additional results that supplement the analysis in *"Foreign Competition for Shares and the Pricing of Information Asymmetry: Evidence from Equity Market Liberalization"*

This Appendix is in three sections. The first section provides some validation tests of the liberalization measure we employ. In the second section, we examine alternative liberalization measures. Additional robustness checks are provided in the last section.

I. Validation tests of the liberalization measure using foreign ownership data

As discussed in Subsection 2.1 of the paper, a potential concern with employing the liberalization shock is that removal of explicit regulatory barriers to foreign ownership may not result in increased competition for firms' shares if foreign investors continue to face significant implicit barriers due to poor local institutional infrastructure. For example, foreign investors may not be willing to invest in stocks in emerging markets if they are concerned about poor investor protection and corporate governance in these countries. Prior studies, however, provide strong evidence that removal of regulatory barriers is associated with a significant increase in foreign capital flows, market valuations, investments, and growth (e.g., Bae et al., 2004; Bae and Goyal, 2010; Bekaert, 1995; De Jong and De Roon, 2005; Edison and Warnock, 2003; Henry, 2000; Mitton, 2006). It would be difficult to explain these findings in prior work if the liberalization shocks did not result in significant increase in risk-bearing capacity from foreign investors. Still, we present some analysis below on how the foreign ownership changes around liberalization.

We obtain data on foreign ownership from FactSet. FactSet data are only available from 2000, which falls after the bulk of the major liberalization steps were taken in the vast majority of the sample countries. Thus, this analysis is significantly constrained by the availability of ownership data. Out of a total 13 shocks used in our main analysis, only 3 occur in our sample period from 2000 onwards (i.e., years 2000 to 2006). This significantly limits the extent to which we can provide a statistical analysis of comparison of firm-level foreign ownership pre- and post-liberalization shocks. Given this limitation, the purpose of this analysis is not to provide causal evidence but rather to provide some descriptive evidence on the construct validity of our liberalization measure.

First, we examine how the aggregate dollar investment levels by foreign investors in a given country vary with liberalization. Changes in aggregate ownership better approximates the notion of changes in risk-bearing capacity in a given stock market. Results presented in column (1) of Table A.I. Panel A suggest that an increase in liberalization is associated with an increase in investments by foreign investors. We next examine the effect of including country- and time-specific effects. Given the constraints in data availability, a significant portion of the variation in foreign ownership restrictions after 2000 would be cross-sectional in nature, which would be absorbed by inclusion of country fixed effects. In the presence of country fixed effects, therefore, we would primarily capture the effects of small changes in liberalization, which may not be representative of the effects of the key liberalization changes we exploit in our main analysis. Similarly, because of the limited data pre- and post-liberalization in this analysis, inclusion of time dummies can soak the effect of liberalization shocks. Nevertheless, we present results after the inclusion of country and year fixed effects. Results presented in columns (2) and (3) of Table A.I. Panel A suggest that our findings continue to hold.

Next, we examine firm-level holdings. Table A.I. Panel B presents the results of a OLS regression of firm-level holdings on *LIBERALIZATION*. Across specifications, we find that an increase in liberalization is associated with an increase in firm-level equity holdings by foreign investors. We also examine changes in the foreign investor ownership around the top tercile liberalization shocks that we use in our main analysis of the price of information asymmetry. We are able to provide only descriptive evidence using this approach because only 3 of the total 13 top tercile liberalization shocks used in our main analysis occur from 2000 onwards (i.e., years 2000 to 2006). We find that firm-level ownership on average is 1% in the three years prior to a shock, rising to 2% in the year of shock, to 3% in the year following the shock, and to 4% three years subsequent to the shock.

Although the above results on firm-level ownership provide a useful construct validity test, the interpretation of the findings and economic magnitudes requires significant care. The key interpretational issue is that the changes in the *actual* ownership levels around the liberalization shocks do not capture increases in risk-bearing capacity, which is the underlying construct of interest in our study. Even if some foreign investors do not hold the stock, the fact that after removal of regulatory restrictions these investors have the ability to step in to purchase shares if the price of risk were to fall sufficiently low allows these investors to affect share prices.

Therefore, while the above results suggest that actual ownership levels increases by 3%, it is important to keep in mind that on average a top tercile liberalization shock in our sample increases the limit on foreign ownership by a much larger fraction of 15%, suggesting that the effect of competition from foreign investors waiting on the sidelines is likely to be quite large. We believe it is because of this reason that even with these levels of actual foreign ownership prior work has found the effects of liberalizations on market valuations, cost of capital, and growth to be quite large (e.g., Bekaert et al., 2005; Chari and Henri, 2004; De Jong and De Roon, 2005; Henry, 2000).

Table A.I: Validation of the liberalization measure using foreign institutional holdings data

This table presents least squares regressions of foreign institutional ownership on liberalization. The dependent variable is a measure of foreign equity characteristic. In Panel A, *IO_Foreign_Value* is the aggregate dollar value of shares held by foreign institutions in a country. The data on Foreign and US Institutional Holdings is obtained from Thomson Reuter's FactSet database and is based on Ferrera and Matos (2008). This measure is available from 2000 to 2006. In Panel B, *IO_Foreign_Pct* is the firm-level foreign institutional ownership and is calculated as the fraction of outstanding shares of a firm held by institutions from FactSet database. This measure is available from 2000 to 2006. *LIBERALIZATION* is the measure of the intensity of restrictions on foreign ownership in a country's stock market as defined in Edison and Warnock (2003). Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

Panel A: Value of foreign institutional ownership holdings

	(1)	(2)	(3)
VARIABLES	*IO_Foreign_Value*	*IO_Foreign_Value*	*IO_Foreign_Value*
LIBERALIZATION	2.515**	3.077**	2.140*
	(2.578)	(2.310)	(1.768)
Observations	168	168	168
R-squared	0.105	0.768	0.268
Fixed Effects	None	Country	Year

Panel B: Firm-level foreign institutional ownership

	(1)	(2)	(3)
VARIABLES	*IO_Foreign_Pct*	*IO_Foreign_Pct*	*IO_Foreign_Pct*
LIBERALIZATION	0.0274**	0.0312**	0.0239**
	(2.660)	(4.896)	(2.124)
Observations	18,204	18,204	18,204
R-squared	0.023	0.126	0.030
Fixed Effects	None	Country	Year

II. Alternative measures of liberalization

In this section, we examine the sensitivity of our inferences to the use of two alternative measures that, like the Edison and Warnock (2003) measure, are also explicitly based on regulations.

There are two key alternative rule-based measures of equity market openness that have been used: the Chinn-Ito index and the Schindler (2009) measure, both of which are constructed based on information in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER). The AREAER is the primary source of the rules and regulations that countries use to govern capital transactions, as well as the proceeds arising from them, between residents and non-residents.

The Chinn-Ito index [introduced in Chinn and Ito (2006)] is based on four binary dummy variables that codify the tabulation of restrictions on cross-border financial transactions reported in the AREAERs. Specifically, the four dummy variables relate to: (1) the openness of a country's capital account, (2) the openness of the current account, (3) the stringency of requirements for the repatriation and/or surrender of export proceeds, and (4) the existence of multiple exchange rates for capital account transaction. Chinn-Ito index take the principal component of these four binary indicators that were reported in the AREAER. A major drawback of this measure is that it is unclear to what extent these indicators are measures of equity market openness in a narrow sense, given that three of the four indices underlying these indicators represent information that is not directly related to capital account transactions. Furthermore, the binary measure related to capital account openness encompasses all types of assets including equity, bonds, real estates, derivatives and government securities and not just equity. Thus, this measure is a crude approximation of equity market liberalization.

Schindler (2009) [subsequently expanded by Fernández et al. (2015) and Klein (2012)] builds a measure of liberalization based on the de jure information from the AREAER as well. The key difference between the Schindler (2009) measure and the Chinn-Ito measure is the level of granularity. In particular, as mentioned above, the Chinn-Ito index codes capital account liberalization as a binary outcome variable based on the authors' read of the country's capital market regulations as a whole and is not asset-type specific. In contrast, the Schindler (2009) index is disaggregated both by whether the controls are on inflows or outflows, and by different categories of assets, such as equities and bonds. Thus, Schindler (2009) allows for a more detailed analysis of capital controls. For each granular category, Schindler (2009) provides a binary outcome variable based on the authors' assessment of the regulatory restrictions.

The key issue with both of these measures is that they ignore rich variation in the liberalization process by coding the measures as an "all-or-nothing" variable. As a result, these measures do not capture all shocks and reduce the statistical power of tests especially in the sample period post-1995. For example, as per these measures, India did not experience any liberalization shock during our sample period. However, as per the Edison and Warnock (2003) measure, India experienced a liberalization shock in 2002. As per the Reserve Bank of India's 2002 Annual Report, India indeed relaxed restrictions on foreign ownership during this year.[21] Based on the Chinn and Ito (2006) and Schindler (2009) measures, we are able to identify only 6 and 9 shocks (respectively) compared to 13 shocks identified by the Edison and Warnock measure.

Nevertheless, we explore the robustness of our inferences to use of the alternative measures. Similar to our main analyses, we identify the first year when the restrictions are relaxed in our sample year and examine the pricing of information asymmetry subsequent to these shocks. Specifically, for the Chinn-Ito index, which is a continuous measure representing the principal component of the four binary measures as discussed above, we follow an approach similar to our main analysis and

[21] See the section on Development and Regulation of Financial Markets in the Reserve Bank of India's 2002 Annual Report.

identify a top tercile shock within our sample period. Since Schindler (2009) provides granular binary variables, we identify the liberalization shock based on the first year in our sample period when the binary variable corresponding to equity purchase restrictions by nonresidents switches from restricted (one) to unrestricted (zero).

Panel A of Table A.II presents the results for Schindler (2009) measure and Panel B presents the results for Chinn and Ito measure. We find that liberalization reduces the price of information asymmetry in 8 out of 10 specifications.

Table A.II: Equity Market Liberalization and the Pricing of Information Asymmetry – Alternative Measures of Equity Market Liberalization

This table presents evidence on the effect of equity market liberalization on the pricing of information asymmetry. The sample comprises 23 emerging market countries for the period 1996 to 2006. The dependent variable *COC* is the mean of four estimates for the implied cost of equity capital (see the Appendix of the paper). Panel A presents results for analyses based on the Schindler (2009) and Fernandez et al. (2015) measure of liberalization while Panel B presents evidence from analyses based on Chinn-Ito (2006) measure of liberalization. For the Chinn-Ito index which is a continuous measure, we follow an approach similar to our main analysis and identify a top tercile shock within our sample period. Since Schindler (2009) provides granular binary variables, we identify the liberalization shock based on the first year in our sample period when the binary variable corresponding to equity purchase restrictions by nonresidents switches from restricted (one) to unrestricted (zero). *BIDASK* represents the annual average of the monthly median bid-ask spreads in equity prices. In Columns (1) through (5) of both panels, *BIDASK* represents the continuous, quartile, median, quintile and decile measure of the annual average of the monthly median bid-ask spreads in equity prices respectively. All control variables have been defined in the caption of Table 2. All specifications include industry-year interactive fixed effects as well as country fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

Panel A: Schindler (2009) measure related to purchase of equity by foreigners

VARIABLES	(1) BIDASK CONT *COC*	(2) BIDASK QUARTILE *COC*	(3) BIDASK MEDIAN *COC*	(4) BIDASK QUINTILE *COC*	(5) BIDASK DECILE *COC*
POST-Schindler-SHOCK	-0.0117***	-0.002	0.000809	-0.00213	-0.00323
	(-3.799)	(-0.615)	(0.195)	(-0.560)	(-0.861)
BIDASK	0.311**	0.006***	0.0116***	0.00526***	0.00261***
	(2.484)	(4.927)	(4.972)	(5.057)	(5.345)
*BIDASK*POST-Schindler-SHOCK*	-0.006	-0.004***	-0.00832***	-0.00324***	-0.00157***
	(-0.069)	(-3.129)	(-4.069)	(-3.359)	(-3.418)
INFLATION	0.240**	0.249**	0.241**	0.251**	0.252**
	(2.245)	(2.390)	(2.285)	(2.410)	(2.416)
LOGSIZE	-0.00628***	-0.00553***	-0.00594***	-0.00544***	-0.00542***
	(-3.848)	(-3.242)	(-3.448)	(-3.187)	(-3.185)
RETVAR	0.0726***	0.0749***	0.0744***	0.0744***	0.0740***
	(3.651)	(3.807)	(3.737)	(3.736)	(3.728)
LEVERAGE	0.0535***	0.0519***	0.0523***	0.0516***	0.0517***
	(8.401)	(8.427)	(8.329)	(8.442)	(8.562)
FBIAS	1.526**	1.556**	1.564**	1.565**	1.562**
	(2.364)	(2.362)	(2.376)	(2.392)	(2.389)
Country Fixed Effects	Included	Included	Included	Included	Included
*Year*Industry Fixed Effects*	Included	Included	Included	Included	Included
Observations	5,794	5,794	5,794	5,794	5,794
R-squared	0.376	0.379	0.376	0.380	0.380

Panel B: Chinn-Ito (2006) Index

VARIABLES	(1) BIDASK CONT *COC*	(2) BIDASK QUARTILE *COC*	(3) BIDASK MEDIAN *COC*	(4) BIDASK QUINTILE *COC*	(5) BIDASK DECILE *COC*
POST-Ito-Chinn-SHOCK	-0.00811	0.00657	0.00865	0.00463	0.00324
	(-1.369)	(0.969)	(1.089)	(0.729)	(0.544)
BIDASK	0.298**	0.00544***	0.00954***	0.00439***	0.00219***
	(2.374)	(4.464)	(3.961)	(4.571)	(4.848)
*BIDASK*Post-Ito-Chinn-Shock*	0.0919	-0.00609**	-0.0115**	-0.00434**	-0.00212**
	(0.279)	(-2.718)	(-2.391)	(-2.671)	(-2.679)
INFLATION	0.239**	0.250**	0.241**	0.252**	0.253**
	(2.214)	(2.365)	(2.257)	(2.361)	(2.374)
LOGSIZE	-0.00744***	-0.00663***	-0.00702***	-0.00654***	-0.00652***
	(-4.825)	(-4.022)	(-4.237)	(-3.944)	(-3.945)
RETVAR	0.0750***	0.0762***	0.0764***	0.0758***	0.0753***
	(3.762)	(3.880)	(3.823)	(3.826)	(3.813)
LEVERAGE	0.0553***	0.0534***	0.0538***	0.0532***	0.0533***
	(8.232)	(8.367)	(8.218)	(8.330)	(8.452)
FBIAS	1.572**	1.621**	1.624**	1.631**	1.634**
	(2.439)	(2.448)	(2.466)	(2.467)	(2.470)
Country Fixed Effects	Included	Included	Included	Included	Included
*Year*Industry Fixed Effects*	Included	Included	Included	Included	Included
Observations	5,794	5,794	5,794	5,794	5,794
R-squared	0.371	0.374	0.371	0.374	0.374

III. Additional Robustness Analysis

In this section, we provide findings from:

(i) A falsification test based on Altonji et al. (2005) (Table A.III.)

(ii) A test of differential effect on regulated industries (Table A.IV.)

(iii) A sensitivity test of our results to the use of unsmooth liberalization measure (Table A.V.).

(iv) Extended pre-periods for tests of parallel trends (Table A.VI.)

(i) Altonji et al. (2005) based falsification tests

We conduct falsification tests in the spirit of Altonji et al. (2005) to gauge the extent to which concurrent local growth/economic shocks could be driving our results. This strategy has been employed in recent studies such as Christensen et al. (2016) and Christensen et al. (2017) to rule out concurrent confounds. Specifically, under this approach, we first regress our implied cost of capital measures on a set of three proxies based on prior work (e.g., Bekaert and Harvey, 2000) that are expected to be strongly correlated with local growth/economic shocks: GDP growth, stock market valuation as a fraction of GDP, and trade as a percentage of GDP. We use this first-stage regression to create a predicted value for implied cost of capital based on these variables. We use the predicted value as the dependent variable in our main DiD specification used to measure the effect of liberalization on the cost of capital. If our main results are driven by local growth/other economic shocks, we would expect the coefficient on *BIDASK*POST-LIBERALIZATION* to be similar to what we find in our main tests. The results in Table A.III. show that the coefficient on *BIDASK*POST-LIBERALIZATION* is statistically and economically indistinguishable from zero. This suggests that our main specifications are effective at absorbing the confounding effect of concurrent growth/other economic shocks.

Table A.III: Assessing Identification Assumptions – Altonji et al. (2005)-based test

This table provides a falsification test in the spirit of Altonji et al. (2005). In untabulated first stage, we regress our implied cost of capital measures on a set of three proxies that are expected to be correlated with local growth/economic shocks: (i) GDP growth, (ii) stock market valuation as a fraction of GDP, and (iii) trade as a percentage of GDP. We use this first-stage regression to create a predicted value for implied cost of capital based on these variables (*Predicted COC*). We use the predicted value as the dependent variable in our main specification. Table presents the results of this second-stage regression. All variables have been defined in the caption of Table 4. All specifications include industry-year interactive fixed effects as well as country fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

	(1) Altonji et al. (2005) test
VARIABLES	*Predicted COC*
POST- LIBERALIZATION	-0.001
	(-1.647)
BIDASK	0.006
	(0.928)
*BIDASK*POST-LIBERALIZATION*	0.009
	(1.157)
INFLATION	-0.050**
	(-2.554)
LOGSIZE	0.001***
	(3.584)
RETVAR	0.003
	(1.579)
LEVERAGE	-0.001
	(-1.462)
FBIAS	-0.110
	(-1.604)
Country Fixed Effects	Included
*Year*Industry Fixed Effects*	Included
Observations	5,694
R-squared	0.922

(ii) Differential effects on regulated industries

We next report evidence from a test that exploits within-country variation in estimating the effects. Similar to the ADR test in the study, by holding the country constant, this analysis mitigates concerns that omitted country-level factors could be driving our results. The analysis is based on the observation that most country-level liberalization policies explicitly leave out regulated industries (Bekaert et al., 2007). Thus, we expect the effect of liberalization shocks to be muted for firms in regulated industries. However, an important empirical issue is in identifying these regulated industries. Although academic studies based on U.S. data typically classify banking and utilities as regulated industries, there is considerable heterogeneity across countries in the notion of "regulated" industries and, hence, in the industries that they choose to liberalize. For example, by 1998, Korea completely eliminated any restrictions on foreign ownership except for some regulated sectors, such as telecommunications, air transportations, and broadcasting. Accordingly, we explore two definitions of regulated industries. The first definition of regulated follows the broad classification adopted in studies such as Bekaert et al. (2007) and Bertrand and Kramarz (2002) and includes industries related to utilities, food products, agriculture, tobacco, alcohol, defense, health, mining, banking, insurance, and transportation. The second definition of regulated industries exclusively focuses on banking and utilities.

We present the results in Table A.IV. Columns (1) and (2) for the first definition and columns (3) and (4) for the second definition of regulated industries, respectively. These columns present results of equation (4) separately for subsamples of firms that are in regulated industries (*REGULATED*) and those that are not. We observe that information asymmetry is priced across both subsamples (coefficient on *BIDASK*). However, consistent with expectations, the results show that the decline in the pricing of information asymmetry following liberalization (coefficient on *POST-LIBERALIZATION*BIDASK*) is concentrated only in unregulated industries and is insignificant in the subsample of regulated industries. We arrive at a similar inference using both definitions of regulated industries.

Table A.IV: Differential Effects on Regulated Industries

This table presents results of subsample analyses for regulated and unregulated industries. In Columns (1) and (2), regulated firms represents firms related to utilities, food products, agriculture, tobacco, alcohol, defense, health, mining, banking, insurance and transportation (Bekaert et al., 2007; Bertrand and Kramarz, 2002). In Columns (3) and (4), regulated firms represents firms related to utilities and banking. All variables have been defined in the caption of Table 4. All specifications include country-year interactive fixed effects as well as firm fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

	(1)	(2)	(3)	(4)
	Unregulated Industries	Regulated Industries	Unregulated Industries (Alternative)	Regulated Industries (Alternative)
VARIABLES	*COC*	*COC*	*COC*	*COC*
POST-LIBERALIZATION	N/A	N/A	N/A	N/A
BIDASK	0.357**	0.293***	0.331**	0.158
	(2.623)	(3.341)	(2.670)	(1.035)
*BIDASK*POST-LIBERALIZATION*	-0.439***	-0.191	-0.305*	-0.207
	(-3.119)	(-0.788)	(-2.060)	(-1.424)
LOGSIZE	0.00399	-0.000240	0.003	-0.008
	(0.933)	(-0.0666)	(0.784)	(-1.141)
RETVAR	0.0115	-0.0169	0.001	-0.014
	(0.489)	(-0.412)	(0.0799)	(-0.395)
LEVERAGE	0.0365*	0.00518	0.0251	0.0457
	(2.080)	(0.246)	(1.561)	(1.116)
FBIAS	1.997*	2.271	2.368***	-0.126
	(2.076)	(1.655)	(2.883)	(-0.407)
*Country*Year Fixed Effects*	Included	Included	Included	Included
Firm Fixed Effects	Included	Included	Included	Included
Observations	3,538	2,156	4,838	856
R-squared	0.783	0.786	0.775	0.846

(iii) *Robustness of results to the use of the unsmooth liberalization measure of Edison and Warnock (2003)*

In all our analysis, we estimate shocks to liberalization measure based on the measure presented in Equation (2). We next examine the sensitivity of our results to using the measure presented in Equation (1). Recall that the difference between the two measures is that Equation (1) is not smoothened while Equation (2) is adjusted for the effect of asymmetric price shocks by scaling the measure in Equation (1) with the relative prices of investable and non-investable stocks. Results presented in Table A.V Panel B suggest that our results continue to hold for this alternative measure.

Table A.V: Using Unsmooth Measure of Restrictions on Foreign Ownership

This table presents evidence on the effect of equity market liberalization on the pricing of information asymmetry. The sample comprises 23 emerging market countries for the period 1996 to 2006. The dependent variable *COC* is the mean of four estimates for the implied cost of equity capital (see the Appendix of the paper). *POST-LIBERALIZATION (Unsmooth)* is constructed as a dummy variable based on changes in the *unsmooth* measure of the intensity of restrictions on foreign ownership in a country's stock market as defined in Edison and Warnock (2003). Panel A presents the descriptive statistics for the unsmooth measure. In Panel B, *POST-LIBERALIZATION (Unsmooth)* is a dummy that takes a value one following the first year a country experiences a change in the restrictions of foreign ownership that is in the top tercile of changes across all country-year changes and zero otherwise. *BIDASK* represents the annual average of the monthly median bid-ask spreads in equity prices. All other control variables have been defined in the caption of Table 2. Specification includes industry-year interactive fixed effects and country fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

Panel A: Descriptive Statistics

Variable	*N*	Mean	Median	Std. Dev.
Unsmooth EW Measure	5676	0.867	0.951	0.169

Panel B: Pricing of Information Asymmetry

VARIABLES	(1) *COC*
POST-LIBERALIZATION (Unsmooth)	-0.00476*
	(-1.809)
BIDASK	0.458***
	(4.820)
*BIDASK*POST-LIBERALIZATION (Unsmooth)*	-0.387***
	(-4.344)
INFLATION	0.232**
	(2.436)
LOGSIZE	-0.006***
	(-4.072)
RETVAR	0.077***
	(4.437)
LEVERAGE	0.048***
	(6.328)
FBIAS	1.515**
	(2.133)
Observations	5,676
R-squared	0.338
Country Fixed Effects	Included
*Year*Industry Fixed Effects*	Included

(iv) *Extended pre-periods for tests of parallel trends*

Next, we examine pre-trends for two additional years in cost of capital movements (and not price of IA) by expanding our sample to include years 1995 and 1994. We are able to do so because we do not require bid-ask spread data for this test and only need to obtain implied cost of capital and control variables for the additional years. Table A.VI below presents the findings from this analysis. It can be seen that the coefficient on all pre-liberalization dummies is not significantly different from zero. We, however, strongly recommend caution while interpreting these results because of the rather limited IBES coverage of international firms during these years, which allows us to obtain implied cost of capital estimates for much fewer firms during these additional years. For example, the number of observations in 1994 is about half of those in 1996.

Table A.VI: Timing of Changes of Baseline Result – Extended Pre-Period

This table presents evidence on the timing of the cost of capital changes around liberalization shocks. *POST-LIBERALIZATION (-4), POST-LIBERALIZATION (-3), POST-LIBERALIZATION (-2)* and *POST-LIBERALIZATION (-1)* are dummy variables that take a value one in the period four-year prior, three-year prior, two-year prior and one-year prior to year of liberalization announcement and zero otherwise. *POST-LIBERALIZATION (0), POST-LIBERALIZATION (1), POST-LIBERALIZATION (2),* and *POST-LIBERALIZATION (3)* are dummy variables that take a value one in the year of, one-year after, two-year after, and three-year after a country experiences a liberalization shock. *POST-LIBERALIZATION (>4)* is a dummy that takes a value one if a country has experienced a liberalization shock four or more years ago. All other variables have been defined in the caption of Table 2. Specifications include industry-year interactive fixed effects as well as country fixed effects. Standard Errors have been obtained by clustering at the country level. Statistical significance (two-sided) at the 10%, 5% and 1% level is denoted by *, **, and ***, respectively.

	Predicted	*COC*	
VARIABLES	Sign	*Coefficient*	*t-stat*
POST-LIBERALIZATION (-4)	n/a	-0.005	(-0.170)
POST-LIBERALIZATION (-3)	n/a	0.000	(0.010)
POST-LIBERALIZATION (-2)	n/a	-0.003	(-0.371)
POST-LIBERALIZATION(-1)	n/a	-0.005	(-0.720)
POST-LIBERALIZATION(0)	-	-0.010**	(-2.317)
POST-LIBERALIZATION(+1)	-	-0.006*	(-1.922)
POST-LIBERALIZATION (+2)	-	-0.012***	(-2.910)
POST-LIBERALIZATION(+3)	-	-0.013***	(-4.696)
POST-LIBERALIZATION(>4)	-	-0.009***	(-3.190)
Controls		Included	
Country Fixed Effects		Included	
*Year*Industry Fixed Effects*		Included	