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Do Foreign Institutional Investors Improve Price Efficiency?

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We study the impact of foreign institutional investors on price efficiency with firm-level international data. Using additions to the MSCI index and the U.S. Jobs and Growth Tax Relief Reconciliation Act as exogenous shocks to foreign ownership, we show that greater foreign ownership increases stock price informativeness, especially in developed economies. This increase arises from new information that foreign investors bring in and displacement of less-informed domestic retail investors. Finally, we show that foreign ownership, particularly from active investors, increases market liquidity, reduces firms' cost of equity, and increases firms' real investment growth. (*JEL* G11, G12, G14, G15)

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Do foreign investors affect the informational content of asset prices? A large literature on stock market liberalizations has shown that foreign investors can improve domestic investments and economic growth.[1] One channel suggested

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[1] Bekaert and Harvey (2000) offer a seminal contribution on the topic; other notable studies include Bekaert, Harvey, and Lundblad (2005), Gupta and Yuan (2009), and Mitton (2006).

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by that literature is improved risk sharing arising from uninformed foreign investors smoothing fluctuations in asset returns across states. An alternative channel is *informed* foreign investors improving the efficiency of capital allocation by increasing the informational content of prices. Prior work does not disentangle the two channels; hence, the contribution of each is empirically unknown. This paper tests an information-based channel.

Some papers, such as Stiglitz (2000), argue that informational asymmetries prevent foreign capital from being profitably invested, whereas others, such as van Nieuwerburgh and Veldkamp (2009), suggest that informed investors should exhibit a home bias in information acquisition decisions. We argue that certain domestic stocks that have low barriers to investment can attract attention of informed foreign investors. These investors base their entry to domestic markets on their ability to trade profitably off of their information. Because of the frictions associated with foreign entry, their information must be able to compensate them for the cost of entry, similar to the intuition of the trade model of Melitz (2003). But increased holdings by foreign institutions also means reduced holdings by other investors. We conceptualize and test two hypotheses: first, that we should see a big informational effect from foreign institutions if those institutions displace less-informed retail investors as opposed to better-informed domestic institutions; second, that we should observe a large effect on domestic price informativeness if foreign institutions have information that is complementary to, as opposed to substitutable with, the information sets of domestic institutions.

Testing the information hypothesis requires granular, investor-level data and a stock-level measure of price informativeness (or equivalently, for our purposes, price efficiency). One of our main advantages is detailed global portfolio ownership data at the level of individual firms and institutional investors. Our sample covers almost 24,000 firms from 40 countries, both developed and emerging, from 2000 to 2016. These data have been used before (e.g., Ferreira and Matos 2008) but, to our knowledge, we are the first to bring them to bear on the dimensions above. We supplement the data with macroeconomic, market, and accounting information. We show that, contrary to received wisdom, foreign institutions unambiguously increase the informational content of domestic asset prices. Our empirical framework employs a micro-founded stock-level measure of price informativeness, defined as the predicted variation of cash flows using contemporaneous market prices,[2] that has justifications both as a welfare measure and as an informational content measure, but the results are robust to alternative definitions.

[2] Bai, Philippon, and Savov (2016) were the first to show this to be a welfare-based measure, and Kacperczyk, Nosal, and Sundaresan (2017) showed this to be the measure of information extracted by traders when they learn from the price. Similar measures were used in Kothari and Sloan (1992), who analyzed the effect of stock prices on future earnings response. In a recent study, Chen, Kelly, and Wu (2019) use the measure to show the information spillovers between buy-side and sell-side research.

In our first test, we relate foreign institutional ownership to price informativeness at the stock level. We use a regression approach with additional controls and a host of static and time-varying fixed effects. We find that the price informativeness of companies with the highest foreign ownership is significantly greater than that of companies with the lowest ownership. The effect is statistically and economically significant for both short and long horizons. We obtain a similar result for domestic institutional ownership; however, the effect from foreign ownership is larger.

The regression results are difficult to interpret economically due to possible endogeneity. For example, if foreign investors were, in fact, uninformed but rational, we would expect them to target firms with the most informative prices to minimize trading losses against informed domestic traders. While our evidence suggests that the entering foreign investors are on average more informed than their nonentering peers or retail domestic investors, we cannot fully rule out that explanation. Also, the effects could be driven by an omitted time-varying variable correlated with both ownership and price informativeness. To address these concerns, we take advantage of two complementary, quasi-natural experiments that generate cross-sectional variation in foreign ownership. Unlike the stock market liberalization literature where all companies in a country are subject to the same shock, our setting allows us to differentiate between treated and untreated stocks *within* a country. First, stocks added to the Morgan Stanley Capital International (MSCI) index experience strong increases in foreign institutional ownership, not matched by domestic institutional ownership changes.[3] The event generates an economically meaningful and reasonably exogenous variation in foreign ownership, which we use to explain price informativeness using difference-in-differences estimation. The results indicate that stocks added to the index experience a subsequent increase in their price efficiency relative to similar stocks outside the index. About 75% of the effect in price efficiency can be attributed to active investors and the rest to passive investors.

While using the MSCI shock allows us to take advantage of a broad cross-country firm-level variation in the data, a potential concern with using the shock is that including the stock in MSCI increases the incentive for informed foreign investors to enter because the index inclusion increases the liquidity of the stock. For that reason, we complement our empirical identification with another shock that is less subject to this particular concern. We use the passage of the U.S. Jobs and Growth Tax Relief Reconciliation Act (JGTRRA) in 2003, which lowered the dividend tax rate to 15% for U.S. firms and firms domiciled in countries that have tax treaties with U.S. As a result, dividend-paying stocks in treaty countries became more attractive to U.S. institutional investors, which led to plausibly

[3] Hau (2011) uses the MSCI reconstitution of 2000 to analyze benchmark changes in actively managed capital.

exogenous variation in U.S. foreign institutional ownership of non-U.S. firms. We find that dividend-paying stocks from treated countries experience an increase in price informativeness, relative to non-dividend-paying stocks, or stocks from untreated countries.

One of the main reasons for the improvement in economic growth found by the stock market liberalization literature is that capital constraints are relaxed. Given that such constraints bind mostly in emerging markets, much of the literature has focused on such markets. Thus, one might expect to find the biggest effects of foreign capital in developing economies. However, we find that developed economies are actually more sensitive to foreign institutional flows than developing economies are, although all are more sensitive to flows from developed economies than from developing economies. The latter result is consistent with a Melitz-style intuition: developed economies, on average, have more sophisticated investors, who, in turn, have a bigger impact on price informativeness. The former result is a little more surprising. One interpretation is that developed economies are better at incorporating the informational content of trades into prices due to a more sophisticated market microstructure. Further supporting our channel, we find that price informativeness improvements are driven by active investors, while passive investors have a smaller, but still positive effect.

As a robustness test, we consider alternative measures of price informativeness. We look at the effects on post-earnings-announcement drift ($PEAD$) and find that increased levels of foreign institutional investment result in a decrease of $PEAD$ over multiple horizons. We also consider price nonsynchronicity, the variance ratio, and the jump ratio which all respond consistently to our identifying shock. Our results also hold using the measure of Davila and Parlatore (2018).

We further show that investors' activeness is a relevant predictor of price efficiency, especially when capital flows from foreign institutions. To examine whether this predictability implies an information-based channel for efficiency, we test whether greater foreign ownership generates improvements in the stocks' information environment. We conduct a battery of tests that broadly indicate the informational advantage of foreign institutional investors. We first show that foreign institutional investors' portfolio revisions predict future stock returns for assets in their portfolios, indicating skill in investment. Second, we show that foreign institutions' skill is better than that of the retail domestic investors they displace. Finally, we show that better historical performance of foreign institutions strongly predicts their increased presence in domestic markets, consistent with the view that such investors face either greater information benefits or lower costs of entry.

Subsequently, we analyze the specific mechanism through which the information of foreign investors enters domestic markets and improves their informational environment. We show that increased foreign ownership leads to (1) higher market liquidity, or lower asymmetric information in the market;

(2) greater analyst coverage and improvement in information production; and (3) better market risk sharing, manifested by reduced cost of capital of about one percentage point. All three effects are statistically and economically significant. We further show that information contained in foreign investors' trades does not overlap with what is contained in domestic investors' trades and is likely different from what is produced by corporate managers. This finding is consistent with both domestic and foreign institutions significantly contributing to improved price efficiency.[4]

Our paper straddles a number of literatures. Several papers have studied foreign investors' impact on cost of capital, liquidity, and efficiency through the lens of stock market liberalizations. Notable examples include Bekaert and Harvey (2000), Bekaert, Harvey, and Lundblad (2007, 2011) and Bekaert et al. (2007). Henry (2000) looks at liberalizations and documents abnormal returns and private investment growth. Chari and Henry (2004, 2008) show that liberalization improves risk sharing and investment efficiency. Bae, Bailey, and Mao (2006) show that foreign investors increase analyst following, accounting standards, indirectly improving the information environment. The benefits of foreign investors have been shown at the country level (Bekaert, Harvey, and Lundblad 2005), industry level (Gupta and Yuan 2009), and firm level (Mitton 2006). This paper, in contrast, formally tests the information-based explanation of improved price efficiency and explores new dimensions (investor and country specific) of analysis, a larger cross-section of countries, and a more direct measure of price informativeness. We do not endure the confounding factors of economic reforms that came with stock market liberalizations, and we are able to isolate the underlying economic channel.

Our paper also relates to the literature on the price impact of institutional investors (Gompers and Metrick 2001). Though it is not their main focus, Bai, Philippon, and Savov (2016) show a positive relationship between institutional ownership and price informativeness using a simple portfolio-sort approach of U.S. stocks. We extend their analysis internationally and decompose ownership into domestic and foreign components. Finally, we highlight the role of foreign institutional ownership in price informativeness and welfare and test the economic mechanism for efficiency gains.

We contribute to the literature on the information production of financial markets and investment decisions.[5] Bond, Edmans, and Goldstein (2012)

[4] In the Internet Appendix, we show that foreign investors have a greater impact on price efficiency when they are from countries with high financial development or under a common law system, especially when investing in countries with low financial development, under civil law, or with weaker financial controls. We find mixed results for investors' similarity measured in terms of their geographic location, language, colonial background, or the size of bilateral trade.

[5] Examples include Dow and Gorton (1997), Baker, Stein, and Wurgler (2003), Goldstein and Guembel (2008), Bond, Goldstein, and Prescott (2010), Goldstein, Ozdenoren, and Yuan (2013), Kurlat and Veldkamp (2015), and Edmans, Goldstein, and Jiang (2015).

survey the literature, emphasizing the separation of new information produced in markets from what is already known and reflected in prices. Chari, Goldstein, and Jiang (2007) find that two measures of the amount of private information impact the sensitivity of corporate investment to stock prices. In an international setting, Wurgler (2000) finds that financial markets improve the allocation of capital.

Our study is also related to literature on institutional investors and price efficiency, which provides mixed evidence. Campbell, Ramadorai, and Schwartz (2009) find that institutions trade aggressively to exploit mispricing around earnings announcements. Boehmer and Kelley (2009) document a positive relation between institutional shareholdings and the relative informational efficiency of prices. Stein (2009) discusses the potential negative effects of increasing institutional ownership on price efficiency. He et al. (2013) show a positive relation between foreign block shareholdings and stock price informativeness. Our paper differs from the studies that focus on price-based efficiency measures by examining a welfare-based measure of price informativeness. We also extend the literature into an international context, a larger sample and broader institutional types; we also trace down the economic mechanism driving efficiency gains.

A rich accounting literature has documented related patterns in the impact of foreign institutional ownership. Fang, Maffett, and Zhang (2015) show that foreign analyst following increases, and foreign analyst dispersion and error decrease after increases in foreign institutional ownership driven by the JGTRRA shock. Bena et al. (2007) show that greater foreign institutional ownership increases long-term investment in several forms of capital. The authors of both papers posit a monitoring role for foreign institutions in generating these results. This conclusion is consistent with our results, but neither paper directly tests a price informativeness channel.

Finally, our paper is also related to the literature on international capital flows. Hau and Rey (2006) develop an equilibrium model and show that the net equity flows into the foreign market are positively correlated with currency appreciation and financial development. Portes, Rey, and Oh (2001) suggest that informational asymmetries could explain a negative relationship between asset trade and distance. Froot and Ramadorai (2008) find that institutional cross-border flows are linked to fundamentals, while closed-end fund flows are a source of price pressure in the short run. Jotikasthira, Lundblad, and Ramadorai (2012) show that flows to funds domiciled in developed markets force changes in these funds' emerging market portfolio allocations. These forced trades, or fire sales, affect emerging market equity prices, pairwise correlations, and betas. Maggiori, Neiman, and Schreger (2020) show that global portfolio flows are often driven by the asset's currency of denomination. Carpenter, Lu, and Whitelaw (2020) show that Chinese stock market has become as informative as the U.S. market.

1. Conceptual Framework

In this section, we describe the economic foundation for our work. For brevity, we do not provide a full theoretical model, but instead illustrate the key elements bolstering our empirical tests. We posit two mechanisms. First, foreign institutional investors are more likely to participate in domestic markets when faced with lower frictions or higher benefits. This mechanism resonates well with the theoretical framework of Melitz in that firms only enter a market when they are sufficiently productive to overcome the costs of doing so. Second, and special to our research context, the size of the impact of foreign institutions on price informativeness can reveal the uniqueness of information they bring. A high impact of foreign entry would indicate that the information sets of such investors are not subsumed by those of domestic investors. In turn, a relatively small impact would indicate a fairly high correlation across investors' information sets. Observing a nontrivial effect on price informativeness helps one to distinguish our mechanism from alternatives in which entry happens as a rational response to the benefit-cost trade-off, regardless of information quality.[6] We graphically illustrate both mechanisms by analyzing the equilibrium adjustments in the presence of the shocks that constitute our identification strategy.

1.1 Participation constraints

Figure 1 illustrates the entry mechanism of foreign investors. The x-axis represents the quantity of the domestic asset acquired by a particular agent. The curved line represents the expected benefit of holding different quantities of the domestic asset. For simplicity, we assume that the benefit is common to foreign and domestic institutional investors, though this can be easily relaxed. The lowest line represents the total cost to the domestic institutional investor of holding different quantities of the asset. We assume no fixed cost to the domestic institutional investor, and the equilibrium quantity of this investor (Q_D) is determined by equating the marginal cost to the marginal benefit. The highest line represents the total cost to the foreign institutional investors, which has a positive fixed cost, and a higher marginal cost. Initially, foreign investors do not participate in the market ($Q_F = 0$). The remaining supply of the asset (Q_R) is therefore held by domestic retail investors, whose benefit function is omitted here. After a negative shock to the costs of foreign investors, be it MSCI index inclusion or tax reduction, we observe that the cost function for foreign investors shifts down and becomes flatter. The new cost function triggers the entry of foreign investors to the market, crowding out retail investors, and leaving domestic institutional investment unchanged. Algebraically, $Q'_F = Q_R$ and $Q'_R = Q_F = 0$. To the extent that foreign institutional investors are better

[6] An example of such an alternative is the model of Han, Tang, and Yang (2016), in which noise traders enter the market because of improving market liquidity. Their entry, however, does not directly lead to an improvement in information quality.

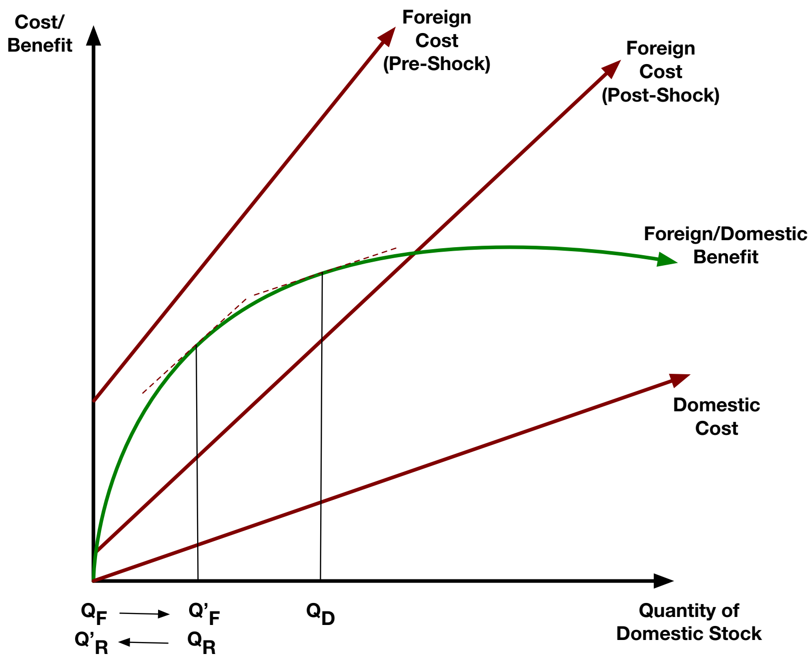

Figure 1
Mechanism for foreign investors to participate in domestic markets
The x-axis shows the quantity of a domestic stock held by domestic institutions Q_D, foreign institutions Q_F, and retail investors Q_R before and after a shock to the costs of investing in the stock for foreign institutions. The straight lines represent the total costs for foreign and domestic institutions, and the curved line represents the benefit (assumed to be common).

informed than domestic retail investors, the result of the change would be an increase in price informativeness. We summarize the essence of these effects in the following prediction.

Prediction 1: *A shock that lowers the cost of investment for foreign institutional participation in domestic markets would cause domestic institutional investment to be unchanged, and foreign investment to increase. As a result, domestic price informativeness will increase.*

Prediction 1 rests on two crucial assumptions. First, the cost function of investment is less concave (more convex) than the benefit function of investment for both foreign and domestic investors. This assumption does not seem tight, as the cost of investing is usually linear (price, taxes, etc.), while the benefit is usually concave. Second, foreign and domestic institutions are (on average) more informed than retail investors. This is a standard assumption in information models that we formally test in Section 4.

1.2 Information uniqueness

Figure 2 illustrates forces related to the *uniqueness* of information that foreign investors bring in to domestic markets. Both plots feature the amount

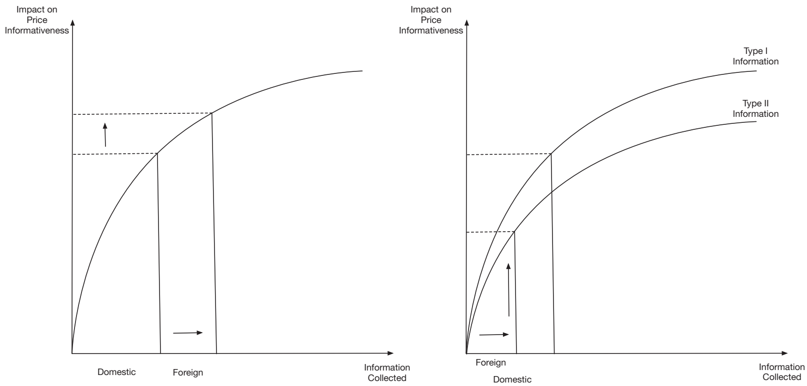

Figure 2
Information sets and price informativeness
The x-axis represents the quantity of information collected by domestic and foreign institutions under two different hypotheses: The left graph shows the impact of foreign information when foreign and domestic institutions collect substitutable information sets (similar to each other). The right graph shows the impact when foreign and domestic institutions collect complementary information sets (different types from each other). Both graphs assume that foreign institutions collect fewer quantities than domestic institutions.

of information collected by different investors on the x-axis, and price informativeness on the y-axis. Under the benchmark model (left plot), foreign and domestic investors have access to information of the same type, or about the same variable, which implies that the entry of foreign traders into the market already explored by domestic investors has a smaller impact on price informativeness. However, under the alternative scenario (right plot), foreign investors could actually have better access to, or pay more attention to a *different source of information*, which would imply a higher marginal impact on domestic markets. As a result, the inclusion of the same amount of foreign information would have a larger impact on domestic price informativeness. We summarize this intuition in the following prediction.

Prediction 2: *If foreign institutional investors have a large impact on price informativeness, it must be because they are using alternative, complementary sources of information to what domestic institutions use, while if they have only a marginal impact on price informativeness, they must be using substitutable sources of information to those of domestic institutions.*

Prediction 2 rests on two instrumental assumptions. First, the impact of any type of information on price informativeness is concave. Second, the overall amount of information from foreign institutions is lower than that from domestic institutions.[7]

[7] If foreign institutions had more information than domestic institutions, our findings could be justified regardless of the type of information had by foreign institutions. Whether complementary or substitutable, their information would have a large effect on price informativeness.

2. Data

Our primary data set is a panel that results from matching several databases. First, we merge FactSet[8] (data on firm-level global institutional ownership) with Datastream (data on firm-level international stock market and accounting data). FactSet reports holdings for a wide range of institutions, such as mutual funds, hedge funds, pension funds, bank trusts, and insurance companies. For non-U.S. firms, FactSet collects ownership data directly from national regulatory agencies, stock exchange announcements, local and offshore mutual funds, mutual fund industry directories, and company proxies and financial reports. We use the last reported value in each calendar year.

We use open-end equity mutual fund return data from Lipper, equity index data from MSCI, and country-level equity market capitalization, gross domestic product (GDP), and industrial production from the World Bank. We merge analyst data from I/B/E/S, and bilateral trade data from the IMF. Our aggregated database has an annual frequency and covers the period 2000–2016. Following previous studies (e.g., Edmans, Jayaraman, and Schneemeier 2017), we exclude financial firms, that is, those with the one-digit Standard Industry Classification (SIC) code 6, and firms with a market capitalization below \$1 million. A firm must have at least 4 successive years of earnings data and a nonzero institutional ownership to be included in our sample. We further limit our sample to countries in which there are at least 20 firms with complete data. The final set consists of 23,811 unique firms for a total of 186,885 firm-year observations; these numbers are similar to those reported in the literature (e.g., Bena et al. 2007).

2.1 Institutional ownership variables

Our data contain 9,449 institutional owners. Foreign institutional ownership (FOR_{it}) is the fraction of firm i's shares held at time t by institutions domiciled in a country other than the one where the stock is listed.[9] The variable FOR_{it} is set to zero if a stock is not held by any foreign institution but is held by at least one domestic institution. Domestic institutional ownership (DOM_{it}) is the fraction of a firm i's shares held at time t by all institutions domiciled in the same country where the stock is listed, relative to the firm's total number of shares outstanding. The variable DOM_{it} is set to zero if a stock is not held by any domestic institution but is held by at least one foreign institution. Total institutional ownership (IO_{it}) is the sum of DOM_{it} and FOR_{it}.

We define active ($ACTIVE_{it}$) and passive ($PASSIVE_{it}$) fractional ownership variables based on two classifications: institutions' investment types and the classification scheme of Bushee (2001). For our first classification, we

[8] We thank Miguel Ferreira and Pedro Matos for making their data available and their useful suggestions.

[9] For multinational companies, we are able to track ownership at the trading desk/subsidiary level. As an example, investments from the Blackrock London office therefore would be considered domestic from the perspective of investing in U.K. companies, but investments from Blackrock U.S. would be considered foreign in the same case.

follow Ferreira and Matos (2008) with one adjustment. In Ferreira and Matos (2008), independent (active) institutions are investment companies (mutual funds), investment advisors, and hedge funds, while other institutions (bank trusts, insurance companies, pension funds and endowments) are considered as grey (passive) investors. Because the mutual fund category includes index funds and exchange-traded funds that invest passively, we adjust the definition to categorize these two types of funds as passive. As an alternative classification, we use the U.S.-based scheme of Bushee (2001). In the fund manager category, he distinguishes three categories: quasi-indexers, with low turnover and high diversification; transient investors, with low turnover and low diversification; and dedicated investors, with low turnover and low diversification. We classify transient and dedicated fund managers as active, while quasi-indexers as passive. For our international sample, we follow Cremers et al. (2016) to use active share measure to gauge how closely the fund is tracking the benchmark index. Then, the new definition of passive investors includes these quasi-indexers (explicit indexers, closet indexers) and other institutions.[10]

We decompose active ownership depending on whether active owners are foreign (FOR_ACTIVE_{it}) or domestic (DOM_ACTIVE_{it}). We separate passive ownership into $FOR\_PASSIVE_{it}$ and $DOM\_PASSIVE_{it}$. For firms listed outside the United States, we define U.S.-based foreign fractional institutional ownership (FOR_US_{it}) and non–U.S.-based foreign ownership (FOR_NUS_{it}). We present all variable definitions and summary statistics in Tables IA.1-IA.4 of the Appendix.

2.2 Stock market and accounting variables

We define the market valuation of firm i at year t as the natural logarithm of market capitalization (M_{it}) to total assets (A_{it}), $log(M/A)_{it}$. Our cash flow variable $(E/A)_{it}$ is earnings before interest and taxes ($EBIT$), divided by total assets. The investment variables include research and development $(R\&D)/A_{it}$, capital expenditures $(CAPEX/A)_{it}$, and total investments $INVESTMENT_{it}$ = $(CAPEX_{it}+R\&D_{it})/A_{it}$, all scaled by total assets. Additional accounting variables include the natural logarithm of sales $log(SALES)_{it}$, measured in thousands of dollars; $LEVERAGE_{it}$, defined as book debt divided by total assets; $CASH_{it}$, defined as cash holdings scaled by total assets; $TANGIBILITY_{it}$, defined as net property, plant, and equipment scaled by total assets; and $FORSALE_{it}$, defined as the fraction of foreign sales in total sales. The variable $CLOSE_{it}$ is the ownership fraction of stock i at time t of all corporate insiders in this firm. $ANALYST_{it}$ is the number of analysts covering a given stock in year t; $TURNOVER_{it}$ is the stock volume divided by total shares outstanding in year t; $Illiquidity_{it}$ is the natural logarithm of

[10] Cremers et al. (2016) use 0.6 as a cutoff point to define closet indexers. Our empirical results are robust if we exclude other institutions and only measure active versus passive investors in the fund manager category or use a different cutoff point for active shares (0.2 or 0.4).

Amihud (2002) illiquidity measure, which is the absolute return over the dollar stock volume using a daily frequency and then averaged within year t. To mitigate the impact of outliers, we winsorize all variables at 1%.

2.3 Country-level variables

We measure the intensity of the connection between any two countries using several different indicators: bilateral trade relations, geographical distance, language, border connections, and colonial origin. The bilateral trade relation between any pair of countries is defined as the sum of their bilateral exports, scaled by the sum of their GDPs. The remaining connection measures are from Mayer and Zignago (2011). Financial system classification data are from Demirguc-Kunt and Levine (1999). Capital control variables for each country are based on the Chinn-Ito index (Chinn and Ito 2006), which measures the country's current account restrictions.

2.4 Price informativeness measure

We use the predicted variation of future cash flows from current market prices as our primary measure of price informativeness (PI). Formally, PI is measured as

$$PI \equiv Corr((M/A)_{i,t},(E/A)_{i,t})\sigma((E/A)_{i,t}) = \frac{Cov((M/A)_{i,t},(E/A)_{i,t})}{\sigma((E/A)_{i,t})}. \quad (1)$$

Scaling the correlation by the standard deviation reflects the fact that a high level of correlation is more meaningful when the asset itself is more volatile. This measure is equivalent to the definition used by Bai, Philippon, and Savov (2016). Similarly, we measure aggregate efficiency as a correlation of investment with earnings, multiplied by the standard deviation of earnings. This term expresses the amount of variation in earnings that can be explained by investment. Internet Appendix IA.D. presents the micro foundation of the informativeness measures. Because we are interested in the contribution of institutional investors to price informativeness, and not just the level of price informativeness for a given stock, our empirical implementation will focus on a slight variant of the above expression, the interaction of institutional investment with price informativeness.

3. Foreign Institutional Capital Flows and PI

In this section, we present our main results on foreign capital flows and price informativeness. We first report the results from the regression model. Next, we discuss the two identification tests that use shocks to foreign institutional ownership. Finally, we provide robustness with respect to various price informativeness measures.

3.1 Regression results

We begin by providing regression evidence on the link between institutional ownership (IO) and price informativeness (*PI*). We estimate the following pooled regression model using firm-level annual-frequency data:

$$E_{i,t+h}/A_{i,t} = a + b_{1,h} log(M/A)_{i,t} + b_{2,h} log(M/A)_{i,t} \times IO_{i,t} + b_{3,h} X_{i,t} + b_{4,h} log(M/A)_{i,t} \times X_{i,t} + e_{i,t+h}, \qquad (2)$$

where $X_{i,t}$ is a vector composed of IO and various controls, including E/A, $log(Asset)$, $CLOSE$, $LEVERAGE$, $TANGIBILITY$, $log(SALES)$, *FORSALES*, and $CASH$. $e_{i,t}$ is measurement error. Further, here and where appropriate thereafter, we interact all control variables in X with $log(M/A)$.[11] We include firm and country $\times$ year fixed effects. To account for possible dependence across firms and years, we cluster standard errors in these two dimensions. The coefficient of interest is $b_{2,h}$, which measures average *PI*, defined as a sensitivity of future earnings to current stock prices, conditional on institutional ownership. Table 1 presents the results. In column 1, we show the results for the 1-year prediction horizon without controls but with all fixed effects. The coefficient $b_{2,h}$ is statistically significant at the 1% level of significance. In column 4, we show that a similar effect holds for *PI* with a prediction horizon of 3 years. We further decompose total institutional ownership into its two components, *FOR* and *DOM*, and estimate the relative contribution to *PI* of each, using the following regression model:

$$E_{i,t+h}/A_{i,t} = a + b_{1,h} log(M/A)_{i,t} + b_{2,h} log(M/A)_{i,t} \times FOR_{i,t} + b_{3,h} log(M/A)_{i,t} \times DOM_{i,t} + b_{4,h} X_{i,t} + b_{5,h} log(M/A)_{i,t} \times X_{i,t} + e_{i,t+h}. \qquad (3)$$

The coefficients of interest are $b_{2,h}$ and $b_{3,h}$, which measure average *PI* conditional on foreign and domestic institutional ownership, respectively. We present the results in columns 2 and 3 for a 1-year horizon with and without stock-level controls. We find that the effect of foreign ownership on *PI* is positive and statistically significant at the 1% level, and is also at least as large as that of domestic ownership. In columns 5 and 6, we report the results for a 3-year horizon. The results remain similar. As a robustness test, we have also estimated the model in which we use changes in both earnings and market values in regression (3). The results, in Table IA.5, are consistent with those obtained using the estimation in levels.

The relative importance of the two ownership measures may be difficult to interpret because the measures of institutional ownership exhibit different variability in the data. Domestic ownership is about three times as variable

[11] Our results remain qualitatively similar in a model without such interactions and are available on request.

Table 1
Price informativeness and institutional ownership: Regression evidence

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
		$E_{i,t+1}/A_{i,t}$			$E_{i,t+3}/A_{i,t}$		$E_{i,t+1}/A_{i,t}$	$E_{i,t+3}/A_{i,t}$
$log(M/A)_{i,t}$	0.018***	0.018***	0.004	−0.009***	−0.009***	−0.035***	−0.001	−0.043***
	(0.002)	(0.002)	(0.004)	(0.003)	(0.003)	(0.005)	(0.004)	(0.006)
$IO_{i,t}$	−0.009			−0.085***			−0.001	−0.026*
	(0.011)			(0.013)			(0.011)	(0.014)
$FOR_{i,t}$		−0.040**	−0.031**		−0.178***	−0.104***		
		(0.014)	(0.013)		(0.017)	(0.018)		
$DOM_{i,t}$		0.003	0.008		−0.049***	−0.001		
		(0.011)	(0.011)		(0.015)	(0.015)		
$For\_Ratio_{i,t}$							−0.008***	−0.008**
							(0.002)	(0.003)
$log(M/A)_{i,t} * IO_{i,t}$	0.082***			0.050***			0.066***	0.050***
	(0.005)			(0.008)			(0.005)	(0.008)
$log(M/A)_{i,t} * FOR_{i,t}$		0.105***	0.076***		0.057***	0.053***		
		(0.013)	(0.011)		(0.015)	(0.014)		
$log(M/A)_{i,t} * DOM_{i,t}$		0.077***	0.058***		0.046***	0.039***		
		(0.005)	(0.004)		(0.009)	(0.008)		
$log(M/A)_{i,t} * For\_Ratio_{i,t}$							0.011***	0.016***
							(0.003)	(0.003)
$log(M/A)_{i,t} * CLOSE_{i,t}$			0.007**			0.012***	0.006**	0.011***
			(0.002)			(0.003)	(0.002)	(0.003)
$log(M/A)_{i,t} * LEVERAGE_{i,t}$			−0.045***			−0.029***	−0.045***	−0.029***
			(0.005)			(0.007)	(0.005)	(0.007)
$log(M/A)_{i,t} * TANGIBILITY_{i,t}$			0.019***			0.028***	0.017***	0.027***
			(0.004)			(0.008)	(0.004)	(0.008)

(*Continued*)

Table 1
(Continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$E_{i,t+1}/A_{i,t}$			$E_{i,t+3}/A_{i,t}$			$E_{i,t+1}/A_{i,t}$	$E_{i,t+3}/A_{i,t}$
$log(M/A)_{i,t} * SALES_{i,t}$			0.014***			0.011***	0.014***	0.012***
			(0.001)			(0.002)	(0.001)	(0.002)
$log(M/A)_{i,t} * FORSALES_{i,t}$			0.014***			0.007	0.014***	0.005
			(0.003)			(0.005)	(0.003)	(0.005)
$log(M/A)_{i,t} * CASH_{i,t}$			−0.039***			−0.038***	−0.038***	−0.038***
			(0.007)			(0.010)	(0.007)	(0.010)
$E_{i,t}/A_{i,t}$			0.228***			0.135***	0.228***	0.135***
			(0.016)			(0.014)	(0.016)	(0.014)
$log(Asset)_{i,t}$			−0.009***			−0.044***	−0.009***	−0.044***
			(0.003)			(0.005)	(0.003)	(0.005)
$CLOSE_{i,t}$			0.002			0.007	0.002	0.007
			(0.003)			(0.004)	(0.002)	(0.004)
$LEVERAGE_{i,t}$			0.058***			−0.015	0.059***	−0.015
			(0.010)			(0.017)	(0.010)	(0.017)
$TANGIBILITY_{i,t}$			−0.006			0.032**	−0.006	0.032**
			(0.011)			(0.013)	(0.011)	(0.013)
$SALES_{i,t}$			0.067***			0.036***	0.067***	0.036***
			(0.006)			(0.005)	(0.006)	(0.005)
$FORSALES_{i,t}$			0.007**			0.006	0.007**	0.005
			(0.003)			(0.005)	(0.003)	(0.005)
$CASH_{i,t}$			0.041***			0.023	0.042***	0.023
			(0.009)			(0.015)	(0.009)	(0.014)
Observations	186,714	186,714	186,714	165,344	165,344	165,344	186,714	165,344
R^2	.677	.677	.710	.612	.612	.623	.710	.623

The dependent variable is the ratio of earnings before interest and taxes ($EBIT$), measured at year $t+1$ or $t+3$ to total assets at year t for company i, *E/A*. *log(M/A)* is the natural logarithm of the ratio of a firm's market capitalization to total assets. Foreign institutional ownership (FOR_{it}) is the fraction of firm i's shares held at the end of year t by institutions domiciled in a country other than the one where the stock is listed. Domestic institutional ownership (DOM_{it}) is the fraction of a firm i's shares held at the end of year t by all institutions domiciled in the same country where the stock is listed. Table IA.1 in the Internet Appendix defines the variables. All regression models include firm, and country×year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

Table 2
Price informativeness and institutional ownership: Regional analysis

	A. Developed economies				B. Emerging economies			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$E_{i,t+1}/A_{i,t}$		$E_{i,t+3}/A_{i,t}$		$E_{i,t+1}/A_{i,t}$		$E_{i,t+3}/A_{i,t}$	
$log(M/A)_{i,t}$	0.007***	−0.008	−0.022***	−0.050***	0.046***	0.037***	0.025***	−0.006
	(0.002)	(0.005)	(0.004)	(0.008)	(0.002)	(0.005)	(0.004)	(0.007)
$FOR_{i,t}$	−0.061***	−0.063***	−0.183***	−0.137***	−0.017	0.024	−0.186***	−0.045*
	(0.017)	(0.017)	(0.022)	(0.022)	(0.020)	(0.017)	(0.032)	(0.025)
$log(M/A)_{i,t} * FOR_{i,t}$	0.128***	0.080***	0.079***	0.060***	0.046***	0.038***	−0.005	0.009
	(0.016)	(0.013)	(0.018)	(0.017)	(0.013)	(0.011)	(0.026)	(0.023)
$DOM_{i,t}$	0.005	−0.001	−0.042**	−0.016	0.017	0.050**	−0.107***	0.005
	(0.012)	(0.012)	(0.015)	(0.015)	(0.019)	(0.020)	(0.030)	(0.022)
$log(M/A)_{i,t} * DOM_{i,t}$	0.091***	0.068***	0.063***	0.052***	0.034	0.027	−0.008	−0.011
	(0.005)	(0.005)	(0.009)	(0.009)	(0.021)	(0.018)	(0.017)	(0.014)
Controls	No	Yes	No	Yes	No	Yes	No	Yes
Observations	136,022	136,022	120,124	120,124	50,692	50,692	45,220	45,220
R^2	.677	.712	.610	.620	.602	.630	.581	.626

The dependent variable is *E/A*. All independent variables are the same as those used in Table 1. Columns 1-4 present the results for the sample of stocks in developed economies, and columns 5-8 present the results for the sample of stocks in emerging economies. All regression models include firm, and country×year fixed effects. The coefficients of the control variables are suppressed for brevity. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

as foreign ownership is. To address this issue, we construct another variable, For_Ratio, defined as the ratio of foreign to total ownership, and use it instead of *FOR* and *DOM* in our regression model. We present the results from the estimation in columns 7 and 8. For each of the two prediction horizons, we observe a positive and statistically significant coefficient of the interaction term between For_Ratio and $log(M/A)$, which means that, in our sample, foreign ownership has a stronger effect on *PI* than domestic ownership does.[12] This result supports our hypothesis that the information brought by foreign institutions is complementary to that brought by domestic institutions under the assumption that the total amount of information foreign institutions bring is not greater than that of domestic institutions. We provide evidence that the information is outside to the firm in the following section.

Next, we analyze the impact of institutional ownership separately for firms in developed and emerging markets. For each group, we estimate the regression model in (3), with and without controls. Table 2 presents the results. For brevity, we only report the coefficients of the main variables. Panel A reports the results for developed markets, and panel B do so for emerging markets. We observe striking differences between the two groups. The effects are strong and statistically significant for both types of ownership in developed economies, but they are significant only for a short horizon in emerging economies. For the

12 To account for any nonlinearities in the underlying relationship, we have also performed a portfolio-sort analysis with single and double sorts with regard to different ownership variables. The results paint a qualitatively similar picture to what we observe from the regression analysis and are reported in Tables IA.6–IA.8 of the Internet Appendix.

long horizon, neither type of ownership is statistically important. In Table IA.7 of the Internet Appendix, we also report differences between a subsample of U.S. and non-U.S. firms. For the U.S. sample, we find that domestic ownership has a larger effect on *PI* than foreign ownership does. In all specifications, the coefficients of *FOR* are statistically insignificant. The results become markedly different when we consider a sample of non-U.S. firms. We find that foreign institutions have a much stronger impact on prices at both shorter and longer horizons. Moreover, while domestic ownership is an important *PI* predictor at a 1-year horizon, its significance disappears when we consider a 3-year horizon.[13]

Finally, another dimension we consider is investors' activeness, that is, their focus on information production. To the extent that price informativeness responds to investors' uncovering mispricing in financial markets and properly accounting for risk, one would expect firms with larger shares of active investors to be more informationally efficient. We split domestic and foreign institutional ownership into active and passive using two classification methods from Section 2.1.[14] We generically define active ownership separately for foreign and domestic owners as *FOR_ACTIVE* and *DOM_ACTIVE*, respectively. Similarly, we define the variables related to passive ownership as *FOR_PASSIVE* and *DOM_PASSIVE*. Next, we estimate the following regression model:

$$\begin{aligned} E_{i,t+h}/A_{i,t} &= a + b_{1,h} log(M/A)_{i,t} + b_{2,h} log(M/A)_{i,t} \times FOR\_ACTIVE_{i,t} \\ &\quad + b_{3,h} log(M/A)_{i,t} \times DOM\_ACTIVE_{i,t} + b_{4,h} log(M/A)_{i,t} \\ &\quad \times FOR\_PASSIVE_{i,t} + b_{5,h} log(M/A)_{i,t} \times DOM\_PASSIVE_{i,t} + b_{6,h} X_{i,t} \\ &\quad + b_{7,h} log(M/A)_{i,t} \times X_{i,t} + e_{i,t+h}. \end{aligned} \tag{4}$$

Control variables X are the same as those used in Table 1. All regression models include firm, and country×year fixed effects. We report in parentheses robust standard errors clustered at firm and year levels. Our coefficients of interest are b_2–b_5. Table 3 presents the results. In columns 1 and 2, we present the results corresponding to classification 1, separately for 1-year and 3-year predictability. We find that higher ownership levels of both active and passive investors improve stock price informativeness in the short horizon. The coefficients of the four interaction terms are positive and statistically significant. At the same time, we find that only active investors (domestic and foreign) contribute to improved price informativeness in the longer horizon. In columns 3 and 4, we repeat the

[13] For firms with no matched or missing ownership data, we can simply set the values of *FOR* and *DOM* to zero. In this larger sample, our main results are similar. See Table IA.9.

[14] Our results are robust if we exclude other institutions from passive group and estimate our regression model separately for each group. In untabulated results, we show that both domestic and foreign institutions in the Other category do not affect the price informativeness, in either short or long horizons.

Table 3
Activeness of institutional investors

	(1)	(2)	(3)	(4)
	Classification 1		Classification 2	
	$E_{i,t+1}/A_{i,t}$	$E_{i,t+3}/A_{i,t}$	$E_{i,t+1}/A_{i,t}$	$E_{i,t+3}/A_{i,t}$
$log(M/A)*FOR\_ACTIVE_{i,t}$	0.069***	0.066***	0.065***	0.068***
	(0.013)	(0.014)	(0.013)	(0.017)
$log(M/A)*FOR\_PASSIVE_{i,t}$	0.143***	−0.104	0.115***	−0.036
	(0.044)	(0.084)	(0.032)	(0.051)
$log(M/A)*DOM\_ACTIVE_{i,t}$	0.055***	0.039***	0.050***	0.035***
	(0.006)	(0.009)	(0.005)	(0.010)
$log(M/A)*DOM\_PASSIVE_{i,t}$	0.075**	0.039	0.081***	0.050**
	(0.026)	(0.034)	(0.013)	(0.020)
$log(M/A)_{i,t}$	0.004	−0.034***	0.004	−0.034***
	(0.004)	(0.005)	(0.004)	(0.006)
$FOR\_ACTIVE_{i,t}$	−0.027*	−0.077***	−0.019	−0.082***
	(0.015)	(0.020)	(0.017)	(0.019)
$FOR\_PASSIVE_{i,t}$	−0.067	−0.374***	−0.083**	−0.206***
	(0.052)	(0.093)	(0.029)	(0.042)
$DOM\_ACTIVE_{i,t}$	0.013	0.001	0.014	0.002
	(0.011)	(0.013)	(0.011)	(0.014)
$DOM\_PASSIVE_{i,t}$	−0.041	−0.027	−0.016	−0.010
	(0.033)	(0.045)	(0.024)	(0.027)
Controls	Yes	Yes	Yes	Yes
Observations	186,714	165,344	186,714	165,344
R^2	.710	.623	.710	.623

The dependent variable is *E/A*. Ownership is divided into active and passive groups based on two different classification methods. The first measure is based on institutional types. Active institutions include actively managed mutual funds and hedge funds, while passive ones include ETFs, index funds, and other types (e.g, pension funds, banks, and insurance companies). The second measure is based on the classification of Bushee (2001). Active institutions include transient and dedicated funds, while passive ones include quasi-indexers and explicit indexers and other types. We decompose active ownership depending on whether active owners are foreign (FOR_ACTIVE_{it}) or domestic (DOM_ACTIVE_{it}). We separate passive ownership into $FOR\_PASSIVE_{it}$ and $DOM\_PASSIVE_{it}$. All control variables (omitted for brevity) are the same as those used in Table 1. The coefficients of the control variables are suppressed for brevity. All regression models include firm, and country×year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

same analysis for classification 2. The results largely mimic those based on classification 1. The only difference is that now passive domestic investors also improve price informativeness in the longer horizon. This result highlights a special role of quasi-indexers as potential information intermediaries.

Overall, our results indicate that domestic and foreign institutional ownership are both important predictors of *PI* in the unconditional sample. Further, the effect is much stronger for the sample of developed markets. At the same time, institutions do not improve price efficiency in emerging markets beyond their short-term impact. Finally, active investors contribute significantly to improvements in price efficiency, consistent with them bringing useful information to the market.

3.2 Identification

Our results so far could be interpreted as associations and not causal relations. One of the potential concerns underlying our analysis is that of omitted variables bias. In particular, *PI* may be higher for reasons unrelated to foreign institutional ownership but at the same time correlated with that variable. Even though the multivariate regression framework allows us to control for some observables, and various fixed effects control for time-invariant unobservables, the main coefficients may still be biased due to time-varying unobservables. Further, our identification could be weakened by reverse causality if foreign investors sorted into stocks with higher levels of price informativeness, for example, as a rational response to minimize their trading losses against informed domestic traders. In this section, we address these concerns by taking advantage of two quasi-natural experiments that induce exogenous variation in foreign ownership: MSCI ACWI index inclusion (MSCI shock), and the passage of the Jobs and Growth Tax Relief Reconciliation Act of 2003 (JGTRRA shock). We implement the identification strategies via the difference-in-differences estimation approach.

Evidence from the MSCI shock. Our first strategy is based on a quasi-natural experiment related to MSCI index inclusions. Several foreign institutions only hold indexed stocks and thus an addition to the index is a positive shock to these stocks' foreign ownership levels. We compare the *PI* of firms newly added to the *MSCI* All Country World Index (*ACWI*) to a sample of comparable firms that did not experience the addition. Our identification assumes that firms are added to the index for reasons other than their *PI*; hence, one can consider the shock as being plausibly exogenous.

We require that at least 5 years of accounting and ownership data be available for the tested firms (2 years before and 2 years after the inclusion year). In our sample, 714 firms with complete accounting and market data are affected by the index inclusion treatment. Our treatment is staggered over multiple years and involves different companies and countries; hence, our results are unlikely driven by specific time trends affecting particular groups of stocks.

For each firm in the treatment group, we identify five nearest neighbors using the propensity score matching algorithm. These serve as a counterfactual control group. Our matching, with replacement, is based on the following ex ante (1 year before inclusion) characteristics: *FOR*, *FOR_ACTIVE*, *DOM*, *Market Capitalization*, $log(M/A)$, E/A, *Analyst*, *FORSALES*, *INVESTMENT*, *Illiquidity*, and country fixed effects. Panel A of Table 4 shows the quality of the matching by showing the average values of each matched characteristic separately for the treatment and control groups. We find that the characteristics of the treated group are not statistically different from those of the control group. The only statistically significant difference, at the 10% level, is for $log(M/A)$.

To assess the plausibility of the parallel trend assumption that underlies the difference-in-differences methodology, we visually inspect the data around the inclusion period. Figure 3 plots the time series of the differences between the treatment and control groups with respect to domestic and foreign ownership and *PI*. The window between years -1 and 0 refers to the period when the

Table 4
Price informativeness and institutional ownership: Evidence from the MSCI shock

A. Pretreatment comparison

	Treatment group	Control group	t-test (p-value)
FOR %	8.771	8.517	0.51
FOR_ACTIVE %	7.798	7.423	0.19
DOM %	34.672	34.654	0.99
$Market\_Cap(\$Bil)$	6.276	5.987	0.49
$log(M/A)$	0.122	0.071	0.09
E/A	0.109	0.105	0.31
FORSALES	0.272	0.267	0.69
Analyst	19.148	18.291	0.17
$R\&D/A+CAPEX/A$	0.086	0.082	0.16
Illiquidity	−11.069	−11.054	0.86

B. Ownership

	(1)	(2)
	FOR	*DOM*
$Treat*After$	0.019***	−0.001
	(0.002)	(0.004)
Observations	24,474	24,474
R^2	.880	.975

C. Price informativeness

	(1)	(2)
	$E_{i,t+1}/A_{i,t}$	$E_{i,t+3}/A_{i,t}$
$log(M/A)*Treat*After$	0.009**	0.046***
	(0.004)	(0.015)
Observations	24,474	6,727
R^2	.661	.695

The treatment group includes 714 firms added to the MSCI ACWI during the sample period. The control group includes five firms that best match each treated firm using propensity scores matching. Panel A compares average values of the variables in the treatment and control groups in the pretreatment period. The dependent variables in panel B are *DOM* and *FOR*. The dependent variable in panel C is *E/A*. $Treat$ is equal to one if a firm is in the treatment group, and zero otherwise. $After$ is equal to one for every year after the one the treated firm is added to the MSCI ACWI, and zero otherwise. All control variables (omitted for brevity) are the same as in Table 1. All regression models include firm and country×year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

treated firm is added to the index. We do not observe any clear differential pre-trends in both quantities within a 3-year window before the shock. This evidence suggests that any effect we identify is not a continuation of a general differential trend between the two groups of firms. Further, we find that, following the shock, foreign ownership increases for treated firms relative to the control group by about two percentage points. At the same time, the domestic institutional ownership of the same stocks does not change, which indicates, through market clearing, that domestic retail investors are selling their assets. This marketwide rotation should lead to an increase in PI, consistent with Prediction 1. This is indeed what we find.

Next, we validate the significance of the effects using the multivariate regression, which allows us to directly control for any differences in observables across two groups of firms, as well as time-invariant unobservables. Specifically,

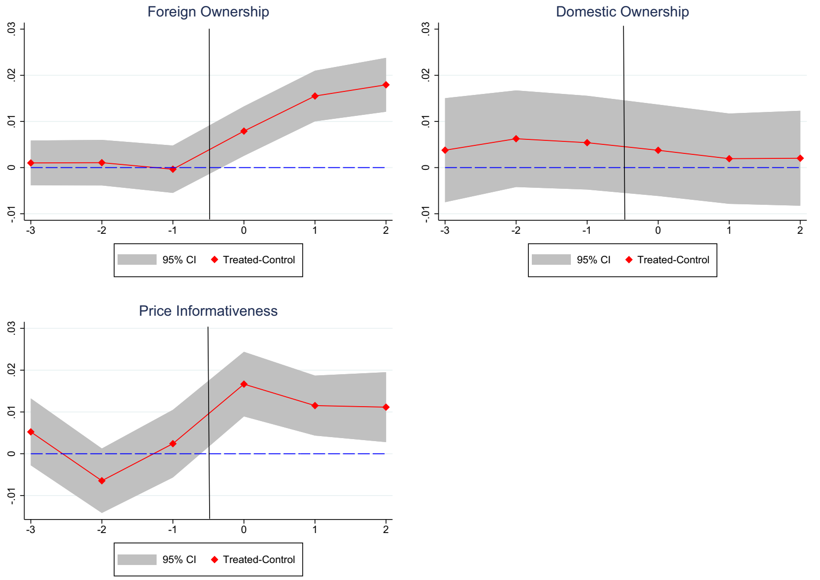

Figure 3
Price informativeness and institutional ownership: MSCI shock
This figure shows point estimates and 95% confidence intervals for the differences in ownership (*FOR* and *DOM*) and price informativeness between treated firms and control firms around stock additions to the MSCI ACWI index. Year 0 is the year when the treated firms were added to the MSCI ACWI index.

for each firm, we define an indicator variable $After$ that is equal to one for the period following the inclusion year and zero for all the years before. We also define an indicator variable $Treat$, equal to one for firms added to the *MSCI ACWI* during our sample period and to zero for all firms in the control group. To zoom in on the shock, we restrict our analysis to the window of 3 years before and 3 years after addition (including the addition year). We estimate the following regression model separately for *FOR* and *DOM*:

$$IO_{i,t} = a + b_1 Treat_i + b_2 After_t + b_3 Treat_i \times After_t + b_4 X_{i,t} + e_{i,t}, \quad (5)$$

where IO is a generic variable for *FOR* and *DOM*. Panel B of Table 4 presents the results. We find that firms added to the index experience an increase in foreign institutional ownership of 1.9 percentage points, on average. The effect is statistically significant at the 1% level and economically large, given that the average firm in the pretreatment sample has an average foreign ownership level of 8.8%. On the other hand, the effect for domestic institutions is economically much smaller and statistically insignificant. In Table IA.9 of the Internet Appendix, we show evidence from foreign institutional stock composition that the foreign institutions are not merely tracking the MSCI index but hold assets that are not part of the index.

Subsequently, we examine the consequence of the shock for *PI* by estimating the following regression model:

$$\begin{aligned} E_{i,t+h}/A_{i,t} = a + b_{1,h} log(M/A)_{i,t} + b_{2,h} Treat_i \times fter_t + \\ b_{3,h} log(M/A)_{i,t} \times Treat_i \times After_t + b_{4,h} log(M/A)_{i,t} \times Treat_i + \\ b_{5,h} log(M/A)_{i,t} \times After_t + b_{6,h} X_{i,t} + b_{7,h} log(M/A)_{i,t} \times X_{i,t} + e_{i,t+h}. \end{aligned} \tag{6}$$

Our coefficient of interest is $b_{3,h}$, which measures the change in *PI* around the shock of the treated group relative to the control group. Panel C of Table 4 presents the results. In column 1, we present the results for a 1-year horizon. We find that, as a result of the shock, *PI* of treated firms increases significantly more on a relative basis. The effect is economically large and statistically significant at the 5% level. In turn, the change in *PI* for the control firms is not statistically different from zero. In column 2, we consider changes in *PI* for a 3-year horizon.[15] Again, we find a statistically significant difference between the treatment and control groups, which is three times as large as that for a short horizon and is economically large.

Evidence from the JGTRRA shock. While using the MSCI shock allows us to take advantage of cross-country firm-level variation in the data, a potential concern with using the shock is that including a stock in MSCI also increases the incentive for informed foreign investors to enter because the event increases the stock's liquidity. For that reason, we consider a complementary shock to foreign ownership that is less subject to this particular concern, namely, the passage of the U.S. Jobs and Growth Tax Relief Reconciliation Act (JGTRRA) of 2003 (e.g., Desai and Dharmapala 2011). JGTRRA lowered the dividend tax rate to 15% for U.S. firms and also extended this tax relief to dividends from firms domiciled in foreign countries that have tax treaties with the United States (nontreaty economies in our sample include Chile, Brazil, Hong Kong, Malaysia, Singapore, and Taiwan). As a result, dividend-paying stocks in treaty countries became more attractive to U.S. investors after the passage of JGTRRA. If U.S. institutions respond to the shock and allocate their portfolios to dividend-paying stocks in treaty countries, the event would create plausibly exogenous variation in U.S. ownership in non-U.S. firms.

Using the difference-in-differences model, we examine whether dividend-paying stocks in the treaty and nontreaty countries experience different patterns in their ownerships in a period of 3 years (2000–2002) before and 3 years (2004–2006) after the shock. Firms in the treaty countries constitute our sample of treated firms, and firms in nontreaty countries make up our control group. Since treated and control firms are from different countries, to mitigate the pure

[15] To avoid overlapping the forecast period before and after addition shock, we only compare observations in the window of 3 years (t-3) before and 2 years (t+2) after the addition at year *t*.

country-level allocation effect, we refine our methodology by first comparing the dividend-paying stocks from treaty and nontreaty countries (dividend DD), and subsequently making a similar comparison for the non-dividend-paying stocks (nondividend DD). A difference between the two differences, or a triple difference (DDD), provides an estimate of the causal effect.

Since our sample has more firms in treaty countries than in nontreaty countries, for each firm in the control group, we identify five nearest neighbors using the propensity score matching algorithm. Our matching, with replacement, is based on the following ex ante characteristics (fixed at 2002 values): *FOR_US*, *FOR_US_ACTIVE*, *FOR_NUS*, *DOM*, *FORSALES*, Market Capitalization, log(M/A), E/A, *INVESTMENT*, and *Illiquidity*. We match the dividend and non-dividend-paying stocks separately based on their payouts in 2002. We show the matching results for both the first and the second difference-in-differences (DD) regressions, and also for the DDD regression. Table 5 shows the quality of the matching by showing the average values of each matched characteristic, separately for the treatment and control groups. The results show that, ex ante, treated firms are not statistically different from control firms for both DD samples. Moreover, even though the values of domestic ownership are, on average, larger for double differences, they are not different using triple differences.

Next, we visually inspect trends in the data around the JGTRRA shock. Figure 4 plots the time series of the differences between the treatment and control groups with respect to foreign investors (*FOR_US*, *FOR_NUS*) and domestic investors. We find that the foreign ownership from U.S. investors increases in the DDD setting around the passage of JGTRRA in 2003. We do not find any evidence of ownership changes for foreign investors from non-U.S. countries and for domestic ownership. This result is consistent with the literature. For *PI*, we repeat the analysis for the DD samples of dividend paying and non-dividend-paying stocks. Figure 4 shows that *PI* increases for treated firms relative to control firms only for the dividend paying DD sample. Further, we do not observe any clear pre trends within the 3 years before the shock. This evidence suggests that any effect we identify is not a continuation of a differential trend between the two groups of firms.

Finally, we validate the significance of the effects using the multivariate regression framework, which allows us to control for any differences in observables across two groups of firms, as well as time-invariant unobservables. We estimate the regression in (6) separately for dividend-paying and non-dividend-paying DD samples.[16] Panel C of Table 5 presents the results. We find a significant improvement in informativeness only for the dividend-paying sample of treated stocks compared to control stocks, while the coefficients

16 The result is robust if we use a quadratic interaction in the DDD regression for PI, $log(M/A)\times Treat\times After\times Dividend$. Figure 4 offers a visualization of the point estimates.

Table 5
Price informativeness and institutional ownership: Evidence from the JGTRRA shock

A. Pretreatment comparison

	DD				DD				DDD	
	Dividend stocks				Nondividend stocks				Diff(Dividend) - Diff(Nondividend)	
	Treated	Control	Diff	Diff (p-value)	Treated	Control	Diff	Diff (p-value)	Diff-Diff	p-value
FOR_US %	1.81	1.85	−0.04	.88	1.07	1.45	−0.38	.25	0.34	.57
FOR_US_ACTIVE %	1.42	1.64	−0.23	.14	0.73	0.92	−0.19	.25	−0.04	.91
FOR_NUS %	1.99	1.82	0.17	.38	0.96	0.70	0.26	.16	−0.09	.83
DOM %	3.80	0.62	3.19	.00	2.69	0.21	2.48	.00	0.71	.33
$Market\_Cap(\$Bil)$	2.30	1.74	0.56	.37	0.47	0.30	0.17	.20	0.39	.73
$log(M/A)$	−0.79	−0.68	−0.11	.04	−1.10	−1.00	−0.09	.24	−0.02	.87
E/A	0.08	0.08	0.00	.85	−0.03	−0.04	0.01	.71	−0.01	.81
FORSALES	0.32	0.28	0.04	.05	0.26	0.30	−0.04	.09	0.08	.05
Analyst	3.08	4.02	−0.94	.00	1.70	1.98	−0.28	.34	−0.66	.45
$R\&D/A+CAPEX/A$	0.06	0.05	0.01	.00	0.06	0.05	0.01	.08	0.00	.90
Illiquidity	−3.60	−3.72	0.12	.19	−2.10	−1.75	−0.35	.01	0.47	.39

(*Continued*)

Table 5
(Continued)

B. Ownership (DDD)

	FOR_US	*FOR_NUS*	*DOM*
$Treat*After*Dividend$	0.011**	−0.002	0.008
	(0.004)	(0.008)	(0.007)
Observations	20,340	20,340	20,340
R^2	.777	.669	.781

C. Price informativeness (DD)

	Dividend stocks		Nondividend stocks	
	$E_{i,t+1}/A_{i,t}$	$E_{i,t+3}/A_{i,t}$	$E_{i,t+1}/A_{i,t}$	$E_{i,t+3}/A_{i,t}$
$log(M/A)*Treat*After$	0.016**	0.030**	−0.028	−0.090***
	(0.008)	(0.012)	(0.020)	(0.021)
Observations	13,455	4,713	6,885	3,427
R^2	.608	.789	.629	.917

The treatment (control) group includes firms in the JGTRRA treaty (nontreaty) countries. Each control firm is then matched to five treatment firms using propensity score matching. We repeat the matching separately for groups of firms paying or not paying dividends in 2002. $Treat$ is equal to one if a firm is in the treatment group, and zero otherwise. $After$ is equal to one after 2003, and zero before 2003. In the DDD regression, $Dividend$ is equal to one if firms pay dividends in 2002 and zero otherwise. Panel A reports the comparison of the variables in the treated and control groups in pre-treatment period. The dependent variables in panel B are *DOM*, *FOR_US*, and *FOR_NUS*. For firms listed outside the United States, we define U.S.-based foreign fractional institutional ownership (FOR_US_{it}) and non–U.S.-based foreign ownership (FOR_NUS_{it}). The dependent variable in panel C is *E/A*. Each regression in panels B and C is a fully saturated models. The regressions in panel C are estimated separately for dividend-paying and dividend non-paying stocks. Control variables (omitted for brevity) are the same as those used in Table 1. All regression models include firm and country×year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

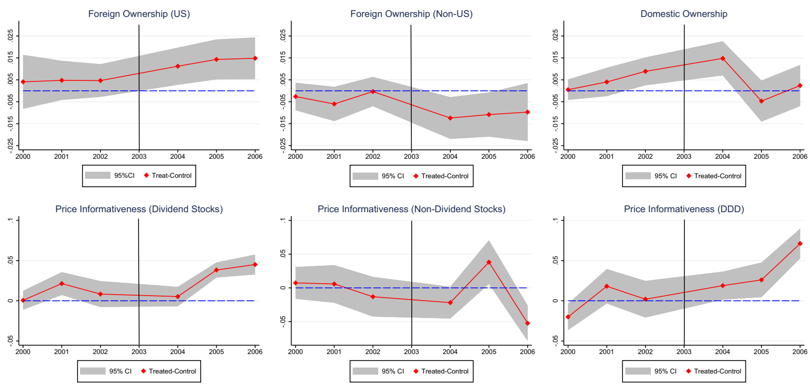

Figure 4
Price informativeness and institutional ownership: JGTRRA shock
The top row shows point estimates and 95% confidence intervals for the differences in price informativeness between treated firms and control firms in a difference-in-differences (DD) setting, for dividend and non-dividend-paying firms, respectively. We also report estimates in a difference-in-differences-in-differences (DDD) setting, namely, between treated firms and control firms, and between dividend and non-dividend-paying firms in a 6-year window around the passage of JGTRRA in 2003. The bottom row shows point estimates and 95% confidence intervals for the differences in ownership (*FOR_US*, *FOR_NUS*, and *DOM*) in the same DDD.

are insignificant in both 1-year and 3-year horizons for the sample of non-dividend-paying stocks.[17]

Active investors. A common perception related to indexing-based shocks is that they largely affect ownership levels of passive institutions, which follow the mandate to hold indexed assets. To the extent that the increase in foreign ownership would come from passive investors, one would expect the information story to have less relevance. If anything, the prediction could be opposite, that is, passive investors would increase their ownership at the expense of active (informationally driven) investors. We test this possibility directly by looking separately at the ownership changes coming from both active and passive investors in the context of our two shocks.[18]

Figure 5 presents the results for the MSCI shock (panel A) and for the JGTRRA shock using U.S. investors only (panel B). While we do observe that passive ownership goes up as a result of MSCI indexation, the effect is even stronger for active ownership. More striking, we find that only active investors respond to the JGTRRA shock. In untabulated results, we also find that the pattern is robust for both developed and emerging economies. Similarly, we

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17 We also consider another difference-in-differences regression comparing stocks that pay dividends in treaty countries as the treatment group and non-dividend-paying stocks as the control group. The result is robust: dividend-paying stocks experience increase in foreign ownership and improvement in price informativeness compared to the control stocks. Table IA.11 in internet appendix shows the result.

18 Here and also in the next sections, we use the first classification method (by investment type) to classify active and passive investors. The results are similar if we use the classification of Bushee (2001).

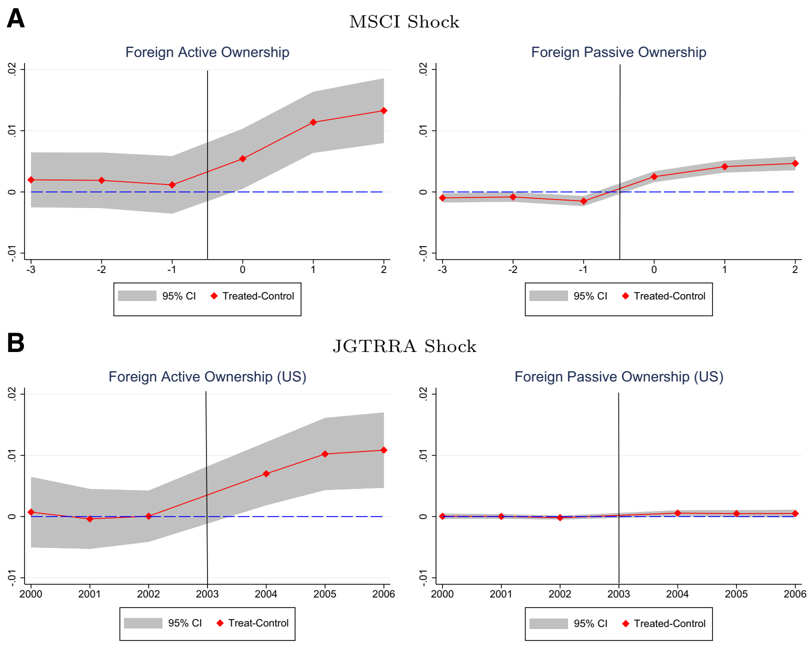


Figure 5
Foreign institutional ownership: Active and passive
This figure depicts point estimates and 95% confidence intervals for the differences in ownership (*FOR_ACTIVE* and *FOR_PASSIVE*) between treated firms and control firms. The first two rows are around stock additions to the MSCI ACWI index. Year 0 is the first year after the treated firms are added to the MSCI ACWI index. The last two rows show results in a 6-year window around the passage of JGTRRA in 2003.

observe that the effect of greater ownership of dividend stocks by U.S. investors is largely due to active investors.

We provide additional statistical verification of the findings in Table 6, separately for MSCI and JGTRRA shocks. Consistent with our univariate evidence, we observe that the increase in foreign institutional ownership is primarily driven by the effect due to active investors. Moreover, for the JGTRRA shock, the foreign ownership goes up only through the increase in active, not passive, ownership. This finding is intuitive since JGTRRA does not necessarily impose binding commitment on the side of indexers.

Notably, an important amplification also happens through the change in composition of ownership, implied by market clearing. In particular, apart from an increase in foreign institutional ownership, our results indicate no change in ownership levels by domestic institutions, which implies that retail investors are the likely sellers of the newly acquired assets.[19] In light of the large literature on relative performance of institutional and retail investors, one could reasonably

[19] We have also confirmed this result using the unconditional sample. In untabulated results, we find that the average correlation between the change in foreign institutional ownership and domestic institutional ownership is a mere

Table 6
Active and passive ownership

A. MSCI shock

	(1)	(2)
	FOR_ACTIVE	*FOR_PASSIVE*
$Treat*After$	0.013***	0.006***
	(0.002)	(0.001)
Observations	24,474	24,474
R^2	.868	.793

B. JGTRRA shock

	(1)	(2)
	FOR_US_ACTIVE	*FOR_US_PASSIVE*
$Treat*After*Dividend$	0.010**	0.0004
	(0.004)	(0.0003)
Observations	20,340	20,340
R^2	.772	.698

Panel A reports results for the MSCI shock and panel B reports the results for the JGTRRA shock (around the passage of the Jobs and Growth Tax Relief Reconciliation Act (JGTRRA) in 2003). In panel A, the dependent variables are *FOR_ACTIVE* and *FOR_PASSIVE*. In panel B, the dependent variables are *FOR_US_ACTIVE* and $FOR\_US\_PASSIVE$, representing the fraction of active and passive foreign investors originating from the United States, respectively. In both panels, $Treat$ is equal to one if a firm is in the treatment group, and zero otherwise. $After$ is equal to one after 2003, and zero before 2003. Control variables (omitted for brevity) are the same as those used in Table 1. All regression models include firm and year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

expect that such retail investors are less informed and thus the reduction in their ownership may amplify the information effects we hypothesize.

Overall, we find that as a result of the ownership shocks, affected companies benefit in terms of growth in ownership from both active and passive foreign institutional investors and, at the same time, from the reduced ownership by less-informed retail investors.

3.3 Alternative efficiency measures

One of the possible concerns with any analysis based on a covariance-based informativeness measure is that it may also capture effects other than changes in price efficiency. For example, the addition to an index may reflect differential exposure of individual stocks to risk factors. To address this concern, we consider several alternative measures of price efficiency.

Post-earnings-announcement drift. First, we consider post-earnings-announcement drift ($PEAD$) reflected in the sensitivity of abnormal stock returns to the value of earnings surprises. Notably, the $PEAD$ is not subjected to risk-based explanations and is a standard way to capture deviations from price efficiency. In a fully efficient market, prices immediately adjust to any

0.01. Also, even though the coefficient in the regression model relating the two quantities is positive, it is not statistically significant. Detailed results are available on request.

earnings surprises and the drift should be zero. To the extent that the presence of foreign investors improves price efficiency, one would expect the magnitude of the drift to decrease as foreign ownership increases.

To construct the variable $PEAD$, we need to define unexpected earnings surprises. We consider two different measures of standardized unexpected earnings (SUE): a time-series SUE and a consensus-based SUE. The time-series SUE is based on a seasonal random walk model with drift (e.g., Bernard and Thomas 1989),

$$SUE_{i,t}^{TS} = \frac{E_{i,q} - E_{i,q-4} - U_{i,t}}{\sigma_{i.t}}, \tag{7}$$

where $E_{i,q}$ measures quarterly earnings per share in quarter q, $E_{i,q-4}$ is earnings per share four quarters before, $U_{i,t}$ and $\sigma_{i.t}$ are the mean and standard deviation of $(E_{i,q} - E_{i,q-4})$ over the preceding eight quarters.

Livnat and Mendenhall (2006) argue that institutional traders react more to analysts' consensus-based earnings surprises rather than to time-series–based earnings surprises. We follow their methodology and compute them as the quarter's actual earnings minus the average of the most recent analyst forecasts, divided by the standard deviation of those forecasts: $SUE_{i,t}^{CB} = \frac{E_{i,q} - \bar{E}_{i,q}}{\sigma_{i.t}}$.

We hypothesize that the magnitude of the $PEAD$ should decrease for the treated firms after the shock. Figure 6 depicts the evolution of the consensus-based $PEAD$ around the MSCI and the JGTRRA shocks for a trading horizon of 60 days into the future. Consistent with our hypothesis that increased foreign ownership improves price efficiency, we find that stocks added to the MSCI index and dividend-paying stocks in countries with JGTRRA benefit experience a drop in $PEAD$ relative to stocks in the control group.

We further assess the robustness of this result to any confounders by estimating the following multivariate regression model:

$$\begin{aligned} CAR\_d1\_dn = a + b_{1,h} SUE_{i,t} + b_{2,h} Treat_i \times After_t + \\ b_{3,h} SUE_{i,t} \times Treat_i \times After_t + b_{4,h} SUE_{i,t} \times Treat_i + \\ b_{5,h} SUE_{i,t} \times After_t + b_{6,h} X_{i,t} + b_{7,h} SUE_{i,t} \times X_{i,t} + e_{i,t+h}, \end{aligned} \tag{8}$$

where CAR_d1_dn denotes the cumulative abnormal return (net of market return) from the first day to the nth day after a quarterly earnings announcement. We consider $n = 60$ or 75. Our coefficient of interest is b_3, which measures the response of abnormal returns to earnings surprises for treated stocks relative to the control group. For the JGTRRA shock, we only analyze a consensus-based $PEAD$ since the quarterly earnings data have limited availability for the early sample.[20]

[20] Our definition of day-0 announcements reflects the timing of the information release. Any release that occurs before market close is measured as of time t, and any release that occurs after market close is measured as of time $t+1$.

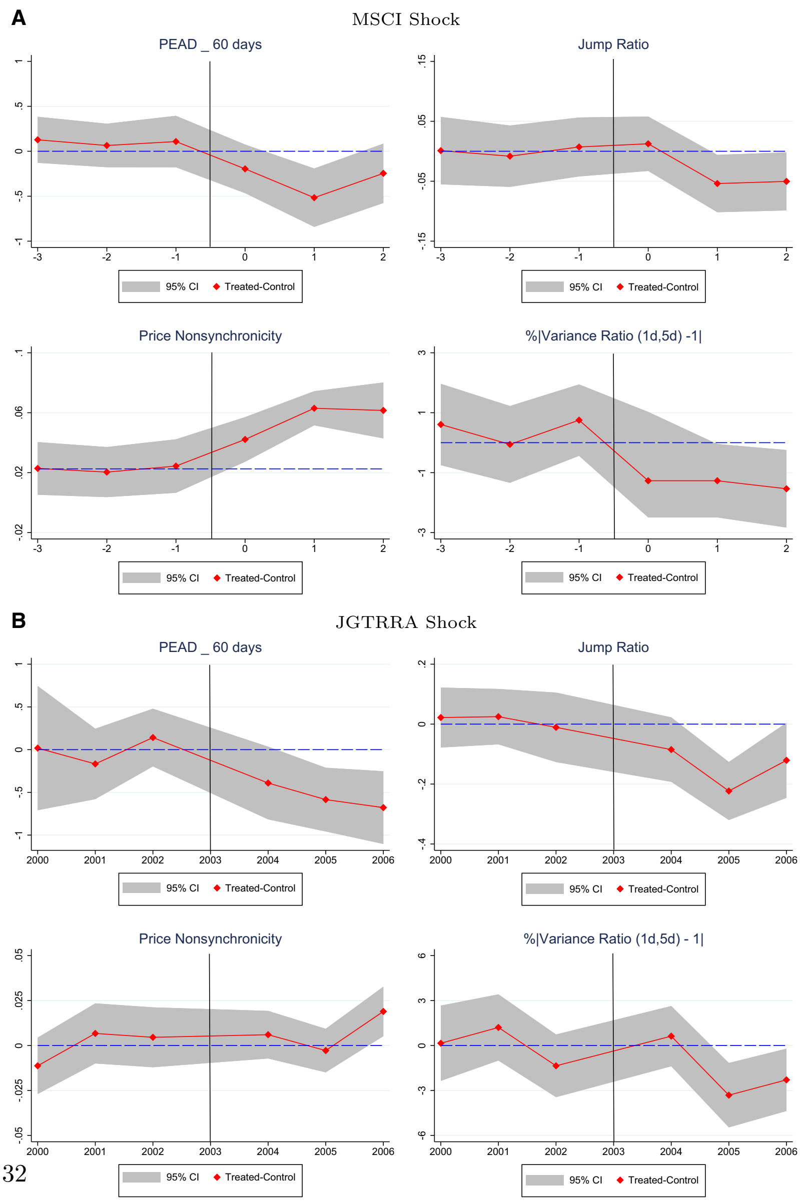


Figure 6
Alternative efficiency measures
This figure shows point estimates and 95% confidence intervals for the differences in post-earnings-announcement drift, the jump ratio, price nonsynchronicity, and the variance ratio between treated firms and control firms. The first two rows are around stock additions to the MSCI ACWI index. Year 0 is the first year after treated firms are added to the MSCI ACWI index. The last two rows show the results for a 6-year window around the passage of JGTRRA in 2003.

Table 7
Post-earnings-announcement drift

A. MSCI shock

	(1)	(2)	(3)	(4)
	SUE^{CB}		SUE^{TS}	
	CAR_60days	CAR_75days	CAR_60days	CAR_75days
$SUE*Treat*After$	−0.351***	−0.361***	−0.354***	−0.312***
	(0.105)	(0.130)	(0.106)	(0.113)
Observations	37,866	37,866	37,806	37,806
R^2	.354	.384	.336	.357

B. JGTRRA shock

	(1)	(2)	(3)	(4)
	Dividend stocks		Nondividend stocks	
	CAR_60days	CAR_75days	CAR_60days	CAR_75days
$SUE*Treat*After$	−0.552**	−0.671**	−0.562	−0.696
	(0.249)	(0.300)	(0.435)	(0.527)
Observations	7,767	7,767	2,146	2,146
R^2	.352	.345	.476	.436

The dependent variables are CAR_d1_d60 and CAR_d1_d75, where CAR_d1_dn denotes the cumulative abnormal return (net of market return) from the first day to the nth day after a quarterly earnings announcement (n=60, or n=75). Panels A and B report the results for the MSCI and JGTRRA shocks, respectively. In both panels, $Treat$ is equal to one if a firm is in the treatment group, and zero otherwise. $After$ is equal to one after 2003, and zero before 2003. Control variables are the same as those used in Table 1. All regression models include firm and year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

Table 7 presents the results. We find that the response of abnormal returns to earnings surprises becomes relatively smaller for stocks added to the MSCI index (panel A) and stocks paying dividends in treaty countries affected by the JGTRRA shock (panel B). The result holds for two different specifications of abnormal returns and is statistically significant at the 5% level. Further, it is robust to alternative specifications of unexpected earnings surprises. Overall, we conclude that an exogenous shock to foreign institutional ownership has a significant positive effect on price efficiency and is unlikely to be due to spurious comovement between prices and earnings.

Price nonsynchronicity, jump ratio, and variance ratio. We also consider three other popular alternatives. The first one is the price nonsynchronicity of Roll (1988), calculated as $1-R^2$, where R^2 is the R-squared from a regression of individual stock returns on the market factor. We estimate the market model using weekly stock returns for each stock-year pair. Higher levels of nonsynchronicity indicate greater information revelation in prices. The second measure is the price jump ratio of Weller (2017). Specifically, this ratio quantifies the share of information acquired and incorporated into prices before earnings announcement by looking at the ratio between announcement returns and the return before and including the announcement. Higher values of jump ratio indicate less information in prices relative to the post-announcement information set. The third measure is the variance ratio (e.g., Lo and McKinlay 1988). In a random walk process, the ratio of long- to short-term return variances equals one. Any deviation from one reflects less informative prices. We compute the standardized variance ratio as |1âŁ“$VR(nday,mday)$|, where *VR(nday, mday)* is the ratio of the return variance over m days to the return variance over n days, divided by the period length.[21]

We begin by inspecting patterns in the three measures for our two shocks. Figure 6 reports the results. In panel A, we present the results for the MSCI shock. In line with our earlier results, we find that the shock induces a significant increase in nonsynchronicity. We also find a relatively strong decline in the variance ratio. In turn, the results for the jump ratio move in the expected direction but are statistically less significant. In panel B, we show the results for the JGTRRA shock. All three efficiency measures significantly respond to the shock, though the response for nonsynchronicity is slightly delayed in time. For both shocks, we observe no visible differences in pre-trends between the treatment and control groups, which lends support to our identification.

We further corroborate the findings using a difference-in-differences regression model for the three measures and report the results in Table 8 for the MSCI shock (panel A) and the JGTRRA shock (panel B). Column 1 reports the results for price nonsynchronicity. We find that the measure increases significantly for stocks added to the index relative to those in the control group. A similar test for the JGTRRA shock reveals a statistically insignificant result (in panel B). Next, we present the results for jump ratio in column 2. We find that the jump ratio weakly decreases for stocks added to the MSCI index relative to those in the control group. For the JGTRRA shock, we find a significant reduction in jump ratio for the dividend-paying stocks, while the reduction is insignificant for the DD sample including stocks that do not pay dividends. For variance ratio, we use the (1 day, 5 days) version of the measure in our difference-in-differences estimation model. We present the results in column 3. We find that the standardized variance ratio decreases for stocks added to the index relative to those in the control group; that is, their prices become more informative. The results are qualitatively similar for the JGTRRA shock, but we only find significant reduction in variance ratio for the dividend-paying stocks' DD regression, while the reduction is insignificant for the DD sample including stocks that do not pay dividends.[22]

It is noteworthy that Bai, Philippon, and Savov (2016) measure *PI* in the cross-section of stocks for each year, whereas our tests utilize a panel setting.

[21] For price nonsynchronicity, we use Wednesday prices to calculate returns; the result is robust when using other days. For the jump ratio, we follow the selection rule from Weller (2017); the result is robust if we apply a different threshold level to the select sample. The number of observations varies across different measures because of the data availability.

[22] Some values for the alternative efficiency measures in the MSCI shock test are missing. The result is robust if we restrict the analysis to a smaller sample of 426 addition firms and their control counterparty firms with more balanced data. See Table IA.20 in the Internet Appendix.

Table 8
Institutional ownership and alternative efficiency measures

A. MSCI shock

	(1)	(2)	(3)
	Price nonsynchronicity	Jump ratio	$\|VR-1\|(\%)$
$Treat*After$	0.030**	−0.041*	−1.119**
	(0.013)	(0.021)	(0.535)
Observations	18,165	18,715	17,744
R^2	.396	.091	.195

B: JGTRRA shock

	(1)	(2)	(3)
	Price nonsynchronicity	Jump ratio	$\|VR-1\|(\%)$
	Dividend stocks		
$Treat*After$	0.003	−0.225***	−2.823**
	(0.007)	(0.072)	(1.178)
Observations	13,199	4,150	13,244
R^2	.571	.465	.323
	Nondividend stocks		
$Treat*After$	0.014	0.035	−5.500
	(0.017)	(0.104)	(4.893)
Observations	5,767	1,264	5,886
R^2	.580	.557	.439

The dependent variables are price nonsynchronicity, the jump ratio, and the variance ratio. Price nonsynchronicity of Roll (1988) is calculated as $1-R^2$, where R^2 is the R-squared from a regression of individual stock returns on the market factor. We estimate the market model using weekly stock returns for each stock-year pair. The price jump ratio of Weller (2017) quantifies the share of information acquired and incorporated into prices before earnings announcement by looking at the ratio between announcement returns and the return before and including the announcement. The variance ratio (Lo and McKinlay 1988) is calculated as $|1VR(nday,mday)|$, where *VR(nday, mday)* is the ratio of the return variance over *m* days to the return variance over *n* days, divided by the period length. Panels A and B report the results for the MSCI and JGTRRA shocks, respectively. Control variables (omitted for brevity) are the same as those used in Table 1. All regression models include firm and year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

In a recent work, Davila and Parlatore (2018) show how to recover absolute and relative *PI* in dynamic environments with rich heterogeneity across investors (regarding signals, private trading needs, or preferences), minimal distributional assumptions, and multiple risky assets. Because their measure requires a long time series of data, we demonstrate empirical results using portfolio sorts. Specifically, we use quarterly earnings and price data in our estimation and require firms to have at least 10 years of data. We estimate absolute and relative *PI* for each firm; then, we sort firms into five portfolios according to firms' sample average foreign ownership levels within their own country. The results, presented in Table IA.12 of the Internet Appendix, indicate a strongly increasing pattern in *PI* for both measures across the five portfolios: low-ownership firms have less informative stock prices than high-ownership firms do. We further isolate the effect by stocks' country of origin and find that the conditional sort preserves the results.

4. Testing the Economic Mechanism

In this section, we examine various economic mechanisms behind our results. We consider two different channels through which foreign ownership can affect capital allocation efficiency, one based on information and another one based on corporate governance.

4.1 The information channel

The premise of our empirical framework is that foreign institutional investors generate useful information, which, if traded on, would result in more information entering domestic stock prices. We posit that decisions to enter the domestic market are rational and depend on the benefit-cost trade-off, which favors investors with better information and lower transaction costs. Conditional on foreign investors entering domestic markets, the degree of improvement in stock-level price informativeness depends on two factors. The first one is whether new information is a substitute or a complement to information generated by other market participants. The second one is whether foreign institutions are better informed than domestic retail investors. In this section, we go through a number of empirical tests to investigate these arguments.

Are foreign institutional investors informed? In our first test, we evaluate foreign investors by their ability to forecast future returns. If foreign investors are informed, we should expect that any changes in their (aggregate) demand for individual stocks should positively predict subsequent changes in the stock returns. To test this hypothesis, we regress future quarterly returns on individual stocks on changes in institutional ownerships, foreign and domestic, of the stocks:

$$Return_{i,q+t} = b_1 \Delta FOR_{i,q} + b_2 \Delta DOM_{i,q} + b_3 X_{i,q} + e_{i,q+t}, \tag{9}$$

where $t=1,4$, and $\text{Return}_{i,q+4}$ measures the return on stock i from quarter $q+1$ to $q+4$. Control variables $X_{i,q}$ include log(size), book to market ratio, past 12-month volatility, and momentum return. We also include firm and time fixed effects and use country $\times$ time fixed effects to absorb the time-varying country-level effects. Table 9, columns 1 and 4, reports the results.

We find that changes in foreign ownership are positively correlated with future stock returns. Two observations are particularly noteworthy. First, the predictive power is stronger for longer-term returns of four quarters ahead. Second, relative to domestic institutions, the effect for foreign institutions is at least as large for short-term returns and much stronger for long-term returns. Thus, if anything, foreign institutions appear to be better informed than domestic institutions.

Given that private information is likely to rest with active investors, we expect to find that the ability of foreign institutions to forecast future stocks returns

Table 9
Performance measure of institutional investors

	(1)	(2)	(3)	(4)	(5)	(6)
	$Ret_{i,q+1}$			$Ret_{i,q+1\to q+4}$		
$\Delta FOR_{i,q}$	0.132			0.572***		
	(0.085)			(0.199)		
$\Delta DOM_{i,q}$	0.013			−0.100		
	(0.044)			(0.101)		
$\Delta FOR\_Active_{i,q}$		0.143*	0.146*		0.708***	0.697***
		(0.084)	(0.087)		(0.212)	(0.215)
$\Delta DOM\_Active_{i,q}$		−0.003	0.005		−0.044	−0.075
		(0.043)	(0.046)		(0.107)	(0.109)
$\Delta FOR\_Passive_{i,q}$			0.045			−0.098
			(0.142)			(0.289)
$\Delta DOM\_Passive_{i,q}$			0.069			−0.266**
			(0.072)			(0.122)
$Log(Size)_{i,q}$	−0.043***	−0.043***	−0.043***	−0.195***	−0.196***	−0.196***
	(0.002)	(0.002)	(0.002)	(0.008)	(0.008)	(0.008)
$Momentum_{i,q}$	0.022***	0.022***	0.022***	0.027***	0.026***	0.026***
	(0.005)	(0.005)	(0.005)	(0.008)	(0.008)	(0.008)
$B/M_{i,q}$	0.054***	0.054***	0.054***	0.132***	0.132***	0.132***
	(0.004)	(0.004)	(0.004)	(0.010)	(0.010)	(0.010)
$Volatility_{i,q}$	−0.456***	−0.456***	−0.456***	−0.831***	−0.829***	−0.830***
	(0.142)	(0.142)	(0.142)	(0.307)	(0.307)	(0.307)
Observations	790,147	790,147	790,147	790,147	790,147	790,147
R^2	.330	.330	.330	.459	.459	.459

The dependent variables are Ret_{q+1} and $Ret_{q+1\to q+4}$, which measure the total returns one quarter ahead (q+1) and 1 year ahead (from q+1 to q+4). Control variables include the natural logarithm of a firm's stock market capitalization, book-to-market ratio, past 12-month volatility, and momentum return. The Internet Appendix defines the variables. All regression models include firm, and country×year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

should be stronger for active investors. Thus, we explore the performance of different types of institutions and decompose each group of institutional owners into active and passive subgroups and estimate the following regression model:

$$Return_{i,q+t} = b_0 + b_1 \Delta FORActive_{i,q} + b_2 \Delta FORPassive_{i,q} + b_3 \Delta DOMActive_{i,q} + b_4 \Delta DOMPassive_{i,q} + b_5 X_{i,q} + e_{i,q+t}. \tag{10}$$

Our results in columns 2 and 5 indicate that active investors' demand moves in the direction of future stock returns. The result is particularly strong for foreign investors whose demand strongly predicts stock returns both one quarter and 1 year ahead. In columns 3 and 6, we additionally include changes in demand related to passive investors. We find that the effect coming from active investors is significantly stronger than that coming from passive investors, consistent with the view that the superior information of active institutions drives the results. Figure 7 further shows the long-term performance of institutional investors. The results indicate that active foreign ownership predicts future returns up to 12 quarters ahead. They are also much weaker for passive foreign investors and domestic institutional ownership.

As a robustness check, we have also explored the same set of tests using our two experiments. The results from these experiments, in Table IA.13,

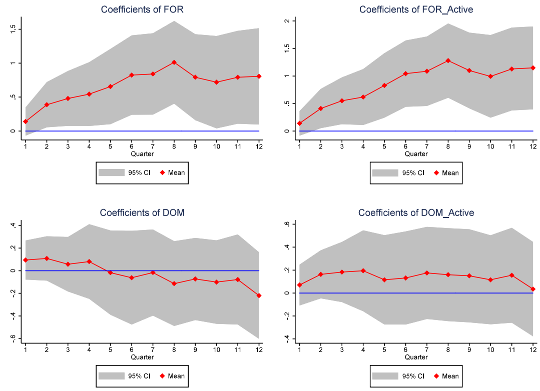


Figure 7
Long-term performance of institutional investors
This figure shows results from regression models of future return on institutional ownership from 1 quarter to 12 quarters ahead. The *FOR* and *DOM* are foreign and domestic institutional ownership, respectively. *FOR_Acitve* and *DOM_Acitve* are foreign and domestic active institutional ownership, *FOR_Passive* and *DOM_Passive* are foreign and domestic passive institutional ownership.

are qualitatively similar. We have also compared the effect for investors in developed and emerging markets. Our results indicate that foreign institutions, especially from developed markets, are better able to predict future returns in emerging markets. Finally, we find that the ability of foreign investors to predict returns is particularly strong when domestic investors are not able to do so at the same time, which suggests that their information sets may be complementary to each other. All the results are available on request.

Which foreign institutions enter the domestic market? Our analysis so far implicitly assumes that all foreign investors participate in the domestic market. However, the decision to enter the domestic market for each institution results from a trade-off of costs and benefits to enter, and not all foreign institutions decide to do so. Our economic framework implies that institutions facing lower costs or greater benefits of information should be more likely to enter; that implication should also guide the intensive margin of the entry. While it is generally difficult to measure the costs and benefits precisely, one can argue that institutions with greater information quality would be more likely to enter. Thus, we study the effect of foreign institutional investors' performance on their entry to domestic stock markets. We estimate the following regression model:

$$\Delta FORShare_{j,q+t} = b_1 Performance_{j,y} + b_2 X_{j,q} + e_{j,q+t}, \qquad (11)$$

where Performance$_{j,y}$ measures the return or Sharpe ratio for institution j over the previous year y (quarter $q-3$ to quarter q). X is a vector of institution-specific control variables, including a natural logarithm of assets under management, a natural logarithm of stocks held by an institution, and

Table 10
Entry of foreign investors: Institutional level

	All Institutional Investors		Active		Passive	
	(1)	(2)	(3)	(4)	(5)	(6)
	$\Delta FORshare_{j,q+1}$		$\Delta FORshare_{j,q+1}$		$\Delta FORshare_{j,q+1}$	
$Ret_{j,q-3\rightarrow q}$	0.699**		0.782**		−2.556	
	(0.328)		(0.339)		(2.061)	
$SharpeRatio_{j,q-3\rightarrow q}$		0.060***		0.065***		−0.083
		(0.016)		(0.017)		(0.059)
$log(AUM)_{j,q}$	−0.037***	−0.037***	−0.038***	−0.038***	−0.017	−0.018
	(0.004)	(0.004)	(0.004)	(0.004)	(0.018)	(0.013)
$log(\#Stocks)_{j,q}$	0.032***	0.031***	0.032***	0.032***	0.018	0.020
	(0.006)	(0.006)	(0.006)	(0.006)	(0.021)	(0.017)
$log(Age)_{j,q}$	−0.004	−0.004	−0.004	−0.004	−0.012	−0.011
	(0.009)	(0.009)	(0.009)	(0.009)	(0.025)	(0.033)
Observations	311,579	311,579	295,940	295,940	15,269	15,269
R^2	.490	.490	.491	.491	.555	.554

$FORshare_j$ is the dollar value of foreign stocks scaled by total values of all stocks under management by institution j. The dependent variable $\Delta FORshare_{j,q+1}$ measures the change of $FORshare_j$ from quarter q to quarter $q+1$. $Ret_{j,q-3\rightarrow q}$ and $SharpeRatio_{j,q-3\rightarrow q}$ measure the return and Sharpe ratio for institutional investor j from quarter $q-3$ to quarter q. Control variables include the natural logarithm of: the dollar value under management $log(AUM)$, the total number of stocks under management $log(\#Stocks)$, and the institutional investor's age $log(Age)$. The sample in columns 1 and 2 is all foreign institutional investors; the sample in columns 3 and 4 is foreign active investors; and the sample in columns 5 and 6 is foreign passive investors. All regression models include institution, and country×quarter fixed effects. Robust standard errors, clustered at the institution level, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

a natural logarithm of the institution's age. The $FORShare_j$ is the dollar value of foreign stocks scaled by total values of all stocks under management by institution j. The change of $FORShare_j$ is measured by the following formula:

$$\Delta FORShare_{j,q+t} = FORShare_{j,q+t} - FORShare_{j,q}\frac{r^f_{j,q+t}}{1+r_{j,q+t}}, \qquad (12)$$

where $r_{j,q+t}$ is return of institution j between quarter q and $q+t$, and $r^f_{j,q+t}$ is return of the foreign component for institution j between quarter q and $q+t$, $t=1,4$.

Table 10 reports the results. In columns 1 and 2, we consider all foreign institutions. The results indicate a strong positive relationship between past performance and the share in foreign stocks. We further estimate the model separately for active and passive institutions. The results are quite striking. While the past performance of active institutions strongly predicts their participation in foreign stocks, the effect is not significant for passive institutions. These results are consistent with the hypothesis that entry into other markets is likely related to the information advantage of such investors.

In addition to the portfolio-level analysis, we also conduct a fund-stock level analysis, which allows us to exploit the cross-sectional variation in investing across all stocks a given institution invests in abroad. We estimate the following regression model:

$$\Delta FOR_{i,j,q+t} = b_1 Performance_{j,y} + b_2 X_{j,q} + e_{j,q+t}. \qquad (13)$$

Table 11
Entry of foreign investors: institutional stock level

	All institutional investors				Active				Passive			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	$\Delta FOR_{i,j,q+1}$		$\Delta FOR_{i,j,q+1\to q+4}$		$\Delta FOR_{i,j,q+1}$		$\Delta FOR_{i,j,q+1\to q+4}$		$\Delta FOR_{i,j,q+1}$		$\Delta FOR_{i,j,q+1\to q+4}$	
$Ret_{j,q-3\to q}$	0.048***		0.157***		0.045***		0.139***		0.097		0.497	
	(0.009)		(0.032)		(0.007)		(0.025)		(0.089)		(0.375)	
$SharpeRatio_{j,q-3\to q}$		0.151***		0.624***		0.154***		0.534***		0.216		2.260
		(0.037)		(0.165)		(0.036)		(0.123)		(0.325)		(2.048)
$log(AUM)_{j,q}$	−0.148***	−0.148***	−0.878***	−0.878***	−0.135***	−0.136***	−0.847***	−0.848***	−0.268***	−0.271***	−1.141**	−1.139**
	(0.016)	(0.016)	(0.061)	(0.061)	(0.015)	(0.015)	(0.057)	(0.057)	(0.100)	(0.101)	(0.449)	(0.442)
$log(\#Stocks)_{j,q}$	0.117***	0.117***	0.813***	0.811***	0.096***	0.095***	0.732***	0.730***	0.207*	0.208*	1.066**	1.057**
	(0.026)	(0.026)	(0.107)	(0.106)	(0.021)	(0.021)	(0.080)	(0.080)	(0.115)	(0.114)	(0.510)	(0.506)
$log(Age)_{j,q}$	−0.008	−0.008	−0.047	−0.049	0.017	0.017	0.039	0.037	−0.243	−0.246	−0.939	−0.957
	(0.031)	(0.031)	(0.121)	(0.122)	(0.022)	(0.022)	(0.090)	(0.090)	(0.242)	(0.245)	(0.920)	(0.929)
Observations	35,232,501	35,232,501	35,232,501	35,232,501	30,731,061	30,731,061	30,731,061	30,731,061	4,387,657	4,387,657	4,387,657	4,387,657
R^2	.058	.058	.073	.073	.062	.062	.074	.074	.151	.151	.186	.186

$\Delta FOR_{i,j,q+t}$ is the change of ownership of institution j in its foreign holding stock i from quarter q to quarter $q+t$, $t=1,4$. $Ret_{j,q-3\to q}$ and $SharpeRatio_{j,q-3\to q}$ measure the return and Sharpe ratio for institutional investor j from quarter $q-3$ to quarter q. Control variables include the natural logarithm of the dollar value under management $log(AUM)$, the total numbers of stocks under management $log(\#Stocks)$, and the institutional investor's age $log(Age)$. The sample in columns 1-4 is all foreign institutional investors; the sample in columns 5-8 is foreign active investors; and the sample in columns 9-12 is foreign passive investors. All regression models include institution, and stock×quarter fixed effects. Robust standard errors, clustered at the institution level, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

$\Delta FOR_{i,j,q+t}$ is the change of ownership of institution j in its foreign holding stock i from quarter q to quarter $q+t$, $t=1,4$. A Stock × Time fixed effect is included to remove any time-varying firm-specific characteristics.

Table 11 reports the results, which are qualitatively similar to those reported for the aggregated data. Past performance strongly predicts future participation in foreign stocks and the results are confined to a sample of active investors only.

Do foreign institutions have different information than domestic institutions? The contribution of foreign ownership to an individual stock's price efficiency vastly depends on the amount of new information the investors bring to the market, relative to what is contributed by domestic investors and firm managers (as in Chari, Goldstein, and Jiang 2007). While independently interesting, in our paper, we do not study explicitly the real feedback effects due to managerial information. Instead, we focus on the relative overlap of information sets between domestic and foreign investors. Our discussion in Section 2 provided testable implications dependent on the overlap between the two sets, and our empirical results in Section 4 largely support the view that information in possession of foreign investors is fairly unique. But this evidence is relatively indirect. To provide more support for this hypothesis, we present two sets of results that provide more direct evidence on the issue. Overall, our results suggest that foreign institutions improve firm-level price efficiency because they have information that is unique to them.

In our first test, we build on the empirical framework of Edmans, Jayaraman, and Schneemeier (2017). Although the authors use a data set that cannot be directly mapped into our framework, we conduct a test that is similar in spirit. We use their measure of Forecasting Price Efficiency (FPE)—price nonsynchronicity—and estimate two regression models: one with Revelatory Price Efficiency (RPE) only, and another one in which we additionally control for FPE. We assess RPE based on the predictability of future investments using current market prices. We consider three measures of investment: CAPEX, R&D, and the sum of the two. We implement the test using a full sample of firms, as well as subsamples based on our two shocks. Specifically, we estimate the following regression model (including FPE):

$$\begin{aligned}\text{Investments}_{i,t+1} =& b_1 log(M/A)*FOR_{i,t} + b_2 log(M/A)*DOM_{i,t} + \\ & b_3 log(M/A)*FPE_{i,t} + b_4 X_{j,q} + b_{5,h} log(M/A)_{i,t} \times X_{i,t} + e_{j,q+t}.\end{aligned} \tag{14}$$

Table 12 reports the results. In panel A, we show the results for the full sample. We find that our main coefficient, b_1, remains positive for all measures of investments. Moreover, it is positive and statistically significant for total investments and CAPEX. In panel B, we report the results for MSCI shock. Again, we find no material change in the value of the coefficient for all three measures of investments. In panel C, we report the results for JGTRRA shock. The same result holds here though the statistical significance of the effects is

Table 12
Information uniqueness of foreign investors

A. Unconditional sample			
	(1) $Invest_{i,t+1}/A_{i,t}$	(2) $CAPEX_{i,t+1}/A_{i,t}$	(3) $R\&D_{i,t+1}/A_{i,t}$
$log(M/A)*FOR_{i,t}$	0.011**	0.012***	−0.003
	(0.005)	(0.004)	(0.002)
$log(M/A)*FPE_{i,t}$	0.449	0.062	0.294***
	(0.298)	(0.152)	(0.078)
Controls and Other interactions	Yes	Yes	Yes
Observations	168,052	168,052	168,052
R^2	.709	.645	.894
B. MSCI shock			
	$Invest_{i,t+1}/A_{i,t}$	$CAPEX_{i,t+1}/A_{i,t}$	$R\&D_{i,t+1}/A_{i,t}$
$log(M/A)*Treat*After$	0.454**	0.373**	−0.004
	(0.214)	(0.189)	(0.054)
$log(M/A)*FPE_{i,t}$	0.679*	0.760**	0.084
	(0.350)	(0.314)	(0.098)
Controls and Other interactions	Yes	Yes	Yes
Observations	18,638	18,638	18,638
R^2	.798	.773	.965
C. JGTRRA shock			
	$Invest_{i,t+1}/A_{i,t}$	$CAPEX_{i,t+1}/A_{i,t}$	$R\&D_{i,t+1}/A_{i,t}$
$log(M/A)*Treat*After$	0.003	0.002	0.001
	(0.004)	(0.003)	(0.001)
$log(M/A)*FPE_{i,t}$	−0.002	−0.005	0.002*
	(0.006)	(0.007)	(0.001)
Controls and Other interactions	Yes	Yes	Yes
Observations	13,178	13,178	13,178
R^2	.532	.478	.865

The dependent variables are $Invest$, $CAPEX$, and $R\&D$. FPE is measured by price nonsynchronicity. Panel A is the entire sample of firms. Panels B and C report the results for the MSCI and JGTRRA shocks, respectively. Control variables (omitted for brevity) are the same as those used in Table 1. All regression models include firm and year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

smaller. This positive coefficient is consistent with the hypothesis that foreign investors likely contribute information to prices that is outside of the manager's information set.

In our second test, we exploit cross-sectional variation with respect to unique information environment foreign investors are associated with. Specifically, foreign investors may acquire unique information through direct linkages that firms they invest in have with their country of domicile. These linkages could be based on firm-level exposure to foreign markets or industry-level expertise of investors that aggregates information beyond that originating in the domestic country. If such linkages are vital, we should expect that the impact of foreign investors on price efficiency should be stronger the stronger such linkages are.

We use three different ways to assess the degree of linkages between foreign investors and firms they invest in. The first one is based on the percentage of a firm's foreign sales in its total sales (*FORSALE*). We argue that firms with high values of *FORSALE* are those in which foreign investors may have

Table 13
Information uniqueness of foreign investors (conditional)

A. Foreign sales ratio (extensive margin)

	$E_{i,t+1}/A_{i,t}$			$E_{i,t+3}/A_{i,t}$		
	High	Low (Ratio=0)	High-Low	High	Low (Ratio=0)	High-Low
$log(M/A)_{i,t} * FOR_{i,t}$	0.077***	0.057***	0.019*	0.037**	0.029	0.008
	(0.011)	(0.016)	(0.011)	(0.012)	(0.024)	(0.016)
$log(M/A)_{i,t} * DOM_{i,t}$	0.055***	0.076***	−0.021***	0.034***	0.062**	−0.028***
	(0.005)	(0.006)	(0.004)	(0.008)	(0.011)	(0.007)
Observations	80,784	102,381		70,950	90,986	
R^2	.666	.589		.746	.666	

B. Foreign sales ratio (intensive margin)

	High	Low	High-Low	High	Low	High-Low
$log(M/A)_{i,t} * FOR_{i,t}$	0.078***	0.058***	0.020	0.049**	0.016	0.033*
	(0.017)	(0.016)	(0.013)	(0.018)	(0.015)	(0.018)
$log(M/A)_{i,t} * DOM_{i,t}$	0.053***	0.053***	0.000	0.031***	0.030**	0.001
	(0.007)	(0.005)	(0.009)	(0.009)	(0.011)	(0.008)
Observations	39,275	39,425		34,493	34,762	
R^2	.675	.707		.592	.659	

C. Market component in return

	High	Low	High-Low	High	Low	High-Low
$log(M/A)_{i,t} * FOR_{i,t}$	0.095***	0.062***	0.033***	0.075***	0.034	0.042**
	(0.014)	(0.015)	(0.011)	(0.018)	(0.022)	(0.016)
$log(M/A)_{i,t} * DOM_{i,t}$	0.051***	0.057***	−0.006	0.046***	0.040***	0.006
	(0.005)	(0.006)	(0.005)	(0.009)	(0.011)	(0.007)
Observations	88,349	91,954		78,630	79,985	
R^2	.726	.742		.642	.669	

D. Industry foreign ownership

	High	Low	High-Low	High	Low	High-Low
$log(M/A)_{i,t} * FOR_{i,t}$	0.077***	0.055**	0.022	0.046**	0.013	0.033
	(0.011)	(0.024)	(0.014)	(0.017)	(0.041)	(0.022)
$log(M/A)_{i,t} * DOM_{i,t}$	0.057***	0.065***	−0.008	0.030**	0.048***	−0.018**
	(0.005)	(0.006)	(0.005)	(0.010)	(0.011)	(0.007)
Observations	90,207	90,982		81,114	78,844	
R^2	.741	.722		.649	.653	

The dependent variable is *E/A*. All independent variables are the same as those used in Table 1. Each panel reports coefficients from estimating the model in (3) for samples of observations sorted within each year and country according to different characteristics. In panel A, we sort firms into those with positive and zero foreign sales ratio (Foreign Sales/Total Sales). In panel B, we sort firms with nonzero foreign sales into high and low foreign sales ratio (Foreign Sales/Total Sales>0). In panel C, we sort firms into high and low R^2 from CAPM model using within-year weekly return data. In panel D, we split firms based on their respective four digit SIC code industries into those with high and low foreign ownership. We also report the coefficients of the difference (**High-Low**) in levels across different subsamples. All control variables (omitted for brevity) are the same as those used in Table 1. All regression models include firm, and country×year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

informational advantage and they should benefit more from foreign investors' ownership. In our sample, around 60% of firms have the value of zero. To represent this feature of the data properly, we consider two versions of the test. On the extensive margin, within each country and year we allocate firms with zero and nonzero *FORSALE* to two bins. Similarly, on the intensive margin, we compare firms with nonzero high and low *FORSALE* separated according to the median value of the sales number. However, since foreign sales are highly correlated with foreign ownership, we impose an additional conditional sort. We first sort firms in each country and year into high and low foreign ownership based on the median value and next within each group we sort sample into high vs. low *FORSALE*, also based on the median split. Then all high *FORSALE* subgroups are combined together, as are all low *FORSALE* subgroups. Within each group of companies described above, we estimate the regression model of (3). Panels A and B of Table 13 present the results, which indicate that foreign investors have a greater impact on price efficiency in companies with a greater fraction of revenues derived from foreign sales. Notably, the relative advantage of foreign investors is particularly visible with respect to longer-term (3-year horizon) efficiency measure, even though the difference between the two coefficients is statistically significant only for the intensive margin results.

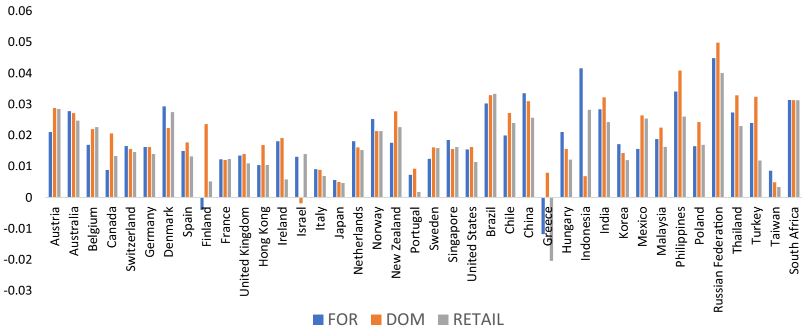
The second way in which we capture the information advantage of foreign investors is based on the percentage of total stock return that is explained by market return using R^2 from the index model. The idea is that foreign investors are less likely to benefit from unique information that is driven by domestic firm-level shocks but more so from information that is due to aggregate macro shocks. To evaluate this hypothesis, in each year and country, we sort firms according to their individual R^2 values calculated from market model with weekly returns for that year. We then split the sample into high and low R^2 according to the median value in that year and country. Again, we estimate the regression model in (3) separately for each group. Panel C presents the results. Consistent with our hypothesis, we find that the advantage of foreign investors is indeed with companies that have higher systematic components of stock returns. As in the previous case, the relative advantage of foreign investors is particularly visible with respect to a longer-term horizon. The differences between the two samples of firms are statistically significant for both short and long horizons.

The third way we evaluate information uniqueness is by looking at the concentration of foreign investors' portfolios in certain industries. The idea is that foreign investors may be more informed about companies that come from industries in which these investors have already expertise. To generate such industry variation, we classify industries by their four-digit SIC codes and calculate the total foreign ownership for each industry. Next, we use this value to split industries into high and low foreign ownership industries within each country-year pair based on the median value. Finally, we estimate the regression model in (3) for each subgroup using all country-year observations. Panel D presents the results. We find that foreign investors improve price efficiency more in companies that belong to industries in which foreign investors have greater stock ownership. However, the respective differences between pairs of coefficients are borderline insignificant.

Are foreign institutions better informed than domestic retail investors? Foreign institutions can also improve price efficiency if the domestic retail investors they displace are less informed. To evaluate this hypothesis, we compare the aggregate performance of different types of investors—foreign institutions, domestic institutions, and retail investors—based on the aggregate holdings of each stock and the country in which they operate. Specifically, for each stock i in each quarter, we generate group-level dollar ownership variables (FOR, DOM, Retail$= 1-FOR-DOM$). Next, we generate country-specific value-weighted returns for each institutional type using the stock returns they hold and their weights in the stocks. We also obtain aggregate returns for each institution type across groups of countries using equal weights and value weights proportional to each country's market capitalization. Figure 8 offers a summary of the results. Our aggregate results indicate that foreign institutions generate higher returns than retail investors, both one quarter ahead and 1 year ahead. The difference in returns is economically large. In the disaggregated data, foreign investors outperform retails investors in 27 of 40 countries one quarter ahead, and in 29 of 40 countries 1 year ahead.

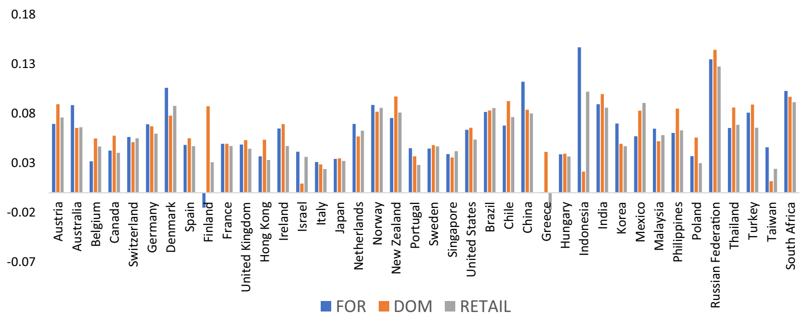
The impact of information. Apart from taking into consideration exogenous factors pertinent to their entering financial markets, foreign investors should also account for their expected impact on the information environment of the target market. We argue that such investors can affect that environment in at least three ways. First, they can affect market liquidity and thus reduce asymmetric information in the market. Further, they can affect the decision of sell-side analysts to cover the target markets, improving information production. Finally, they can improve risk sharing and thus reduce the cost of capital in the market. In all three cases, one would expect price efficiency to improve. In this section, we evaluate the possibilities in the context of our two experiments.

We first present the time-series evolution in market illiquidity and analyst coverage in the context of our two shocks. We consider two measures of market liquidity: turnover (trading volume over shares outstanding) and Amihud (2002) illiquidity measure. Figure 9 shows a significant increase in turnover and analyst coverage and a decrease in illiquidity, all results consistent with our hypothesis. Next, we estimate a regression model akin to that in formula (6), with the same measures as dependent variables. Table 14 (columns 1—3) presents the results for the MSCI (panel A) and JGTRRA shocks (panel B). We find that stocks that are added to an index, on average, experience a significant increase in their market liquidity, relative to a comparable group of control stocks. For instance, for the MSCI shock, average turnover increases by about 16% of one standard deviation, while illiquidity decreases by close to 12% of one standard deviation. Both effects are economically and statistically significant.



Quarterly Average Return

FOR	DOM	RETAIL
1.682%	1.772%	1.400%



Yearly Average Return

FOR	DOM	RETAIL
6.311%	6.213%	5.452%

Figure 8
Total performance of different types of investors
This figure shows the aggregate performance based on the total holdings of didderent investors: foreign institutions, domestic institutions, and retail investors. For each stock i in each quarter, we use the levels of ownership (FOR, DOM, Retail (1- FOR-DOM)) to calculate the dollar value for each type of investor. Scaled by country market capitalization, we create pseudo-index weights based on weights $w_{i,f}$, $w_{i,d}$, $w_{i,re}$ in each country, and the total returns are the value-weighted average returns for different types of investors. Value-weighted average returns across countries for different types of investors are reported in the table below the figure.

Next, we evaluate the impact of the shocks on equity analyst coverage. Our measure of coverage is based on the number of sell-side analysts issuing firm-level forecasts in a given year.[23] We present the result in column 3 of Table 14. Stocks added to the index experience a relatively greater increase in analyst coverage of about three analysts per stock, or 20% of a standard deviation. The effect is significant both economically and statistically for the MSCI shock.

[23] Our data do not allow us to distinguish between domestic and foreign analysts. Being able to do so could be a useful additional dimension of our test.

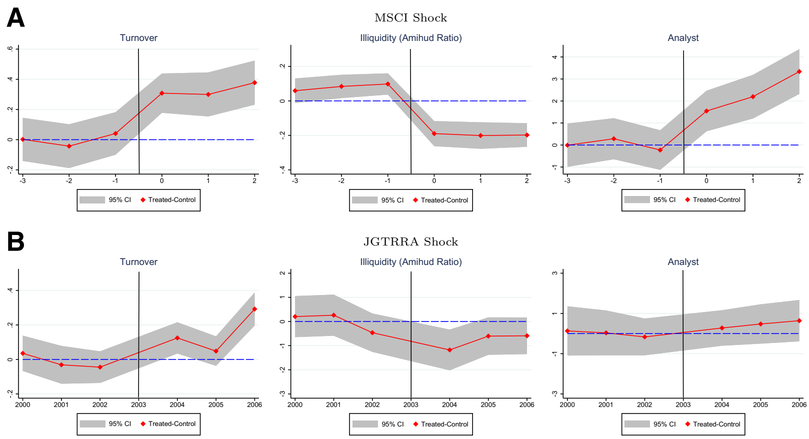


Figure 9
Turnover, illiquidity, and analyst coverage
This figure shows point estimates and 95% confidence intervals for the differences in the turnover ratio, illiquidity (Amihud ratio), and analyst coverage. The first row is around stock additions to the MSCI ACWI index. Year 0 is the first year after treated firms are added to the MSCI ACWI index. The second row shows the results for a 6-year window around the passage of JGTRRA in 2003.

For the JGTRRA shock, the estimated coefficient is insignificant, though it is positive. Hence, stock inclusion could lead to the greater production of relevant information coming from increased analyst coverage. Following the evidence in Hong and Kacperczyk (2010), one could further argue such information should be, on average, less biased thus enhancing its quality.

Further, we examine the price effects of the changing composition of asset ownership by looking at the cost of equity ($ICOE$). Following Gebhardt, Lee, and Swaminathan (2001), we calculate $ICOE$ from the residual income model. Figure 10 graphically shows the differences in the values of cost of capital around index inclusion and the JGTRRA shock. The patterns in this figure show a significant reduction in the cost of equity, consistent with our hypothesis. We further assess the statistical significance of the results using the difference-in-differences model. Table 14 (column 4) presents the results for the MSCI (panel A) and JGTRRA (panel B) shocks. We find negative coefficients, which are significant at the 1% level, for both settings. The results are also economically large: as a result of the shock, treated firms for the MSCI (JGTRRA) shock experience a reduction in their costs of equity of about 1.2 percentage points (2.6 percentage points) relative to firms in the control group.

The reduction in the cost of equity of treated firms suggests that these firms should invest more as a result, since the threshold for accepting profitable projects drops, holding investment opportunities constant. This mechanism leads to a testable hypothesis of changes in investment levels. We assess this hypothesis separately for investments in capital expenditures and $R\&D$ and report the results in columns 5 and 6. We find a positive effect on

Table 14
Economic consequences of the information channel

A Liquidity, analyst coverage, ICOE and investment (MSCI shock)

	(1)	(2)	(3)	(4)	(5)	(6)
	Turnover	Illiquidity (Amihud)	Analyst	ICOE	CAPEX	R&D
$Treat*After$	0.306***	−0.245***	3.561***	−0.012***	0.005**	0.0012*
	(0.059)	(0.036)	(0.343)	(0.003)	(0.002)	(0.0007)
Observations	19,150	19,749	24,474	14,667	23,687	23,687
R^2	.743	.939	.903	.571	.713	.894

B. Liquidity, analyst coverage, ICOE and investment (JGTRRA shock)

	(1)	(2)	(3)	(4)	(5)	(6)
	Turnover	Illiquidity (Amihud)	Analyst	ICOE	CAPEX	R&D
$Treat*After*Dividend$	0.291***	−0.453***	0.453	−0.026***	0.004**	−0.004
	(0.060)	(0.159)	(0.396)	(0.007)	(0.002)	(0.008)
Observations	13,571	15,773	20,325	15,050	20,325	7,156
R^2	.694	.883	.921	.917	.623	.877

C. Governance index (MSCI shock)

	(1)	(2)
	Governance index	E_index
$Treat*After$	−0.007	−0.063
	(0.006)	(0.076)
Observations	7,887	3,442
R^2	.831	.784

The dependent variables in panels A and B are Amihud's illiquidity, analyst coverage, implied cost of equity (ICOE), CAPEX, and R&D investments. Panels A and B report the results for the MSCI and JGTRRA shocks, respectively. The dependent variables in panel C, for the MSCI shock, are the Governance index and E-index. Control variables are the same as those used in Table 1. All regression models include firm and year fixed effects. Robust standard errors, clustered at the firm and year levels, are reported in parentheses. *p < .1; **p < .05; ***p < .01.

investment for each shock, which is also consistent with the findings in Bena et al. (2007). The result is statistically more significant for changes in capital expenditures.

Finally, we evaluate the information-based mechanism separately for developed and emerging markets in the context of the MSCI shock. Table IA.14 reports the results. We find that cost of equity and liquidity go up as a result of the shock in both markets. At the same time, only firms in developed markets benefit from the shock in terms of higher investment levels, both $CAPEX$ and $R\&D$. This finding suggests that despite improvements in risk sharing and better market environment, firms in emerging markets, on average, are not able to capitalize on their foreign portfolio flows. Overall, our results indicate economically significant welfare gains associated with the increased foreign stock ownership, a novel result in the international finance literature.

4.2 The governance channel

An alternative channel through which institutional ownership could affect price efficiency is improved corporate governance through better monitoring. To the

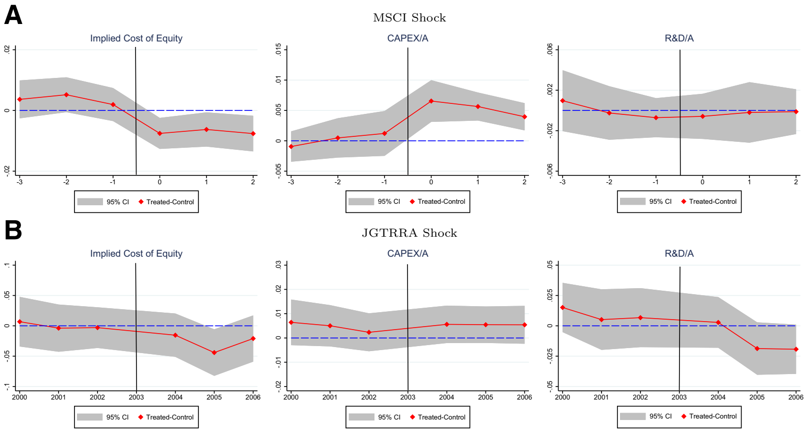


Figure 10
Implied cost of equity and investment
This figure shows point estimates and 95% confidence intervals for the implied cost of equity (ICOE), CAPEX, and *R&D* between treated firms and control firms. The top row shows the results around stock additions to the MSCI ACWI index. Year 0 is the first year after treated firms are added to the MSCI ACWI index. The second row shows the results for a 6-year window around the passage of JGTRRA in 2003.

extent that increased ownership boosts incentives to monitor, one could expect better efficiency as a result. This function is often facilitated by large passive owners, as suggested in the literature. Our results thus far indicate that both types of ownership increase because of our shocks. Given that passive investors increase their presence, one could expect they could improve monitoring inside the firms they hold.

We test this hypothesis formally by using two corporate governance measures: the composite governance index of Albuquerque et al. (2018), and a global version of E-index from Homanen and Liang (2018). The composite governance index is based on 16 attributes obtained from the governance category of the Bloomberg ESG database. The attributes are divided into four subcategories: board, audit, antitakeover provisions, and compensation and ownership. Homanen and Liang (2018) construct a global version of E-index from Bebchuk, Cohen, and Ferrell (2008) by using data from MSCI Governance Metrics. We estimate the difference-in-differences regression model with both governance indexes as the dependent variables. Panel C of Table 14 reports the results, and Figure IA.4 of the Internet Appendix graphically reports the results.

We do not find a significant relation between the MSCI index inclusion and either measure of governance. The lack of statistical and economic significance is true for both developed and emerging markets.[24] In the analysis

[24] Similar to Aggarwal et al. (2011), we find a positive relation between foreign ownership and corporate governance in the full sample. We further test whether the interaction term of the governance level and foreign ownership

of the JGTRRA shock, we find that only active ownership experiences a significant increase, while passive investors enter with statistically insignificant magnitudes.[25]

In sum, our results indicate that institutional owners are more likely to improve price efficiency through their impact on the information environment than through their effect on governance. At the same time, we want to caution that the absence of statistically significant results for governance channel may be a result of the inherent difficulty in measuring its effect, especially in the international data. Further, governance index we use may be less suitable to capture monitoring activities and it is more a measure of changes in governance structure. Since our interpretation is more focused on monitoring, this may be another reason why our results are insignificant.[26] In sum, we cannot conclusively rule out the possibility that the governance mechanism operates independently of the information channel we focus more on in this paper.

5. Concluding Remarks

The growing presence of institutional investors has resulted in greater integration of financial markets. Using a large data set and a new micro-founded measure of price informativeness, we examine the role of foreign institutional capital flows on the price informativeness of domestic stocks. We obtain several results. First, at the aggregate level, foreign investors have a large impact on the informational content of prices. The effect on price informativeness is driven by active investors, while passive investors have a smaller, but still positive effect. Further, developed markets are more sensitive to informational foreign flows than developing markets are, but investors from developed economies have a bigger impact on foreign markets than do investors from developing economies. Overall, our results underscore the significant role of foreign institutional investors in affecting price efficiency. To the extent that foreign investors improve price informativeness, the common view that such investors cause financial market instability may require additional scrutiny. Also, informational disadvantages faced by foreign investors may induce a positive selection among the institutions entering foreign markets. In this regard, any policy intervention that forces foreign financial players to enter other markets may be unnecessary. We leave these and other issues for future research.

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has a compounded effect on price informativeness. If the monitoring channel is at play, we should expect foreign investors to have a larger effect conditional on a lower level of corporate governance. We do not find such an effect.

25 The composite governance index data are only available after 2004, and the E-index is available after 2008. We also estimate the difference-in-differences regression model for the MSCI shock only using data after 2004 or 2008 and find that all the results (ownership, price informativeness, liquidity) are qualitatively similar.

26 We thank an anonymous referee for pointing out this distinction to us.

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