The Informational Efficiency of Cross-Listed Securities and the Quality of Institutions

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Abstract

Keech and Munger (2015) argue that the failures of markets are often preceded by the failures of governments to define the necessary institutions needed for the success of markets. In this paper, we attempt to test this hypothesis by examining whether institutional quality is associated with higher levels of informational efficiency in financial markets. To do so, we use a unique empirical design that holds the structure of financial markets constant while exploiting cross-sectional variation in the level of institutional quality. In particular, we examine the association between the informational efficiency of cross-listed securities on US markets and the home country's level of institutional quality. Results provide some general evidence that the quality of institutions in the home country is associated with greater informational efficiency in the cross-listed securities. These findings are supportive of the Keech and Munger (2015) hypothesis.

JEL codes: G10, G12, G15, G18

Keywords: market efficiency, institutional quality, American depositary receipts, political stability, judicial efficiency, accounting quality, board of director rights.

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1. Introduction

In his seminal work, Fama (1970) contends that the role of capital markets is to create a system of prices that "provide accurate signals for resource allocation" and that when prices fully reflect all available information, the market is said to be "efficient" (383). The presence of excess volatility, bubbles, and crashes, however, has raised some concerns among economists about whether financial markets are efficient. For instance, after outlining several types of financial market failures, Stiglitz (1993) describes a set of government interventions that policymakers could undertake to correct market failures. In contrast, Keech and Munger (2015) posit that while government intervention generally follows market failures, the government may contribute to market failures since the government sets up and defines the institutional framework in which markets are allowed to operate. In other words, since governments establish and oversee the institutions that monitor and affect markets, those institutions play a key role in the success and failure of markets. Thus, the Keech and Munger (2015) argument indicates that institutional quality plays an important role in defining the informational efficiency of financial markets. The objective of our study is not to contribute to the debate about whether or not stock prices are informationally efficient per se. Instead, we seek to determine whether institutional quality is associated with the stock prices that are more informationally efficient.

Determining whether or not the quality of institutions can influence the informational efficiency of stock prices is a difficult task for at least two reasons. First, we need cross-sectional variation in the level of institutional quality. Therefore, examining the institutional quality in the

United States while examining the efficiency of US stocks does not provide variation in institutional quality. We therefore examine a sample of countries that exhibit variation in institutional quality measures. However, the second issue that we face is that institutional quality in a particular country could determine the financial market structure in that country, which undoubtedly influences the informational efficiency of stock prices. To overcome this potential bias, we use a sample of American depositary receipts (ADRs), which are securities that trade on US exchanges but represent the securities of foreign companies. Using ADRs allows us to hold the structure of the financial market constant while allowing for variation in the level of institutional quality in the ADR home country.[1]

Admittedly, we are not the first to use this structure to examine how the quality of institutions in a particular country influences stock characteristics. Eleswarapu and Venkataraman (2006) explore how institutional quality in the ADR home country influences the liquidity of ADRs. Results from their analysis suggest that institutional quality leads to lower bid-ask spreads and higher levels of liquidity. Like Eleswarapu and Venkataraman (2006), we use several measures that capture the various aspects of institutional quality in the ADR home country. We gather data from prior studies that quantify the level of judicial and legal efficiency, the level of political stability, the quality of the country's accounting system, and the liberalness of corporate governance rules. Unlike Eleswarapu and Venkataraman (2006), however, our focus

[1] We note that the ADRs used in this sample are listed on major exchanges. While unsponsored and so-called Level I ADRs trade on over-the-counter markets, this data is unavailable. Therefore, the ADRs in our sample are considered Level II or Level III ADRs. This classification is important when drawing inferences because our sample of ADRs has higher disclosure requirements and must abide by regulatory and exchange requirements. Stated differently, our analysis examines the price efficiency of ADRs that are most similar (in disclosure requirements and other regulatory restrictions) to US stocks listed on major exchanges.

is on the informational efficiency of ADR prices instead of the liquidity of ADRs.[2] We therefore borrow several measures of efficiency from prior studies. From Hou and Moskowitz (2005), we estimate two measures of price delay that capture the delay with which ADRs incorporate market-wide information. These measures of price delay capture the inefficiency of prices. A higher measure of delay equates to prices that are less efficient. From Lo and MacKinlay (1988), we estimate variance ratios that test the random-walk hypothesis. To the extent that prices are efficient, they should follow a random walk. We then adjust the traditional variance ratio to measure the distance from the random-walk benchmark where the variance ratio is equal to one. Again, this adjusted variance ratio inversely measures informational efficiency. Lastly, we use a new measure from Blau, Griffith, and Whitby (2018) that captures the noisiness of prices or the size of intraday temporary price shocks. Using these measures of market inefficiency, we then test whether the institutional quality in the ADR's home country is related to the informational efficiency of ADRs.

We recognize and point out a few limitations of our study. First, while the use of ADRs allows us to avoid the endogenous treatment of heterogeneous (home country) market structures, this research design does not come without a cost. There are several countries that have only a few listed ADRs. Thus, although we attempt to make broad generalizations, we recognize the possibility that this type of unbalanced sample may impact the strength of the conclusions we draw. Accordingly, we are careful with our language when discussing the implications of our findings. Second, measuring the informational efficiency of stock prices is a difficult task. Given

[2] Our study is also somewhat related to Fernandes and Ferreira (2008), who explore the informational environment of cross-listed security prices. They show an important asymmetry in how cross-listing influences the stock price informativeness. From developed markets, cross-listing improves the information environment, while from emerging markets, the effect is the opposite. Our study indirectly attempts to determine whether institutional quality can speak to this asymmetry.

the difficulty of measuring efficiency, we use several proxies that measure different aspects of informational efficiency and remain agnostic as to which is the most important or most proper.

In our first set of tests, we plot the institutional quality for each ADR home country and the efficiency measures for the average ADR from each country. We then report a trend line in the scatterplots. These results provide some general, cross-sectional evidence that as institutional quality in the ADR home countries increases, price delay, the adjusted variance ratios, and the noisiness of prices typically decrease—which indicates that informational efficiency is associated with higher institutional quality. However, these findings are not uniform across all measures of institutional quality. For instance, the results are strongest when we examine political stability and weakest when looking at the liberalness of corporate governance rules.

In the second part of our analysis, we conduct a series of multivariate tests that control for both ADR-specific and country-specific characteristics while allowing us to isolate the effect of institutional quality in the home country on the efficiency of ADRs. When focusing on the efficiency of the judicial or legal system, we find some evidence that judicial and legal efficiency in the ADR home country reduces the level of price delay, although the statistical significance is not robust to various control variables. However, we find stronger evidence that judicial and legal efficiency is associated with a marked reduction in the noisiness of prices and in our adjusted variance ratios. These latter findings suggest that the quality of legal institutions seems to be associated with more efficient ADR prices.

When focusing on political stability, we find strong evidence that the quality of (political) institutions in the home country improves the efficiency of ADR prices. These results are generally robust across our measures of price delay, noise, and the adjusted variance ratio. In economic terms, we find that a 1 percent increase in political stability is associated with a 0.15

percent to 0.89 percent decrease in price delay. Qualitatively similar results are found when we examine noise and adjusted variance ratios. We also find some evidence that the quality of accounting in the ADR home country is associated with our efficiency measures. Our weakest results are found when we examine the association between the liberalness of corporate governance rules and our measures of efficiency. We find very limited evidence that countries with more decentralized power among boards of directors have less noise in their ADR prices.[3]

Combined, our findings show that higher-quality institutions are associated with greater informational efficiency of cross-listed securities specifically and financial markets more generally. The implications of our analysis suggest that improvements in the level of institutional quality—particularly political stability—could lead to more efficient markets. To the extent that regulators are concerned about financial market inefficiencies, our analysis indicates that improvements in institutional quality may help in addressing these concerns. Going forth, section 2 describes the data used throughout the analysis. Section 3 describes the methods used and the results from our empirical tests. Section 4 offers some concluding remarks.

2. Data Description

The data used throughout the analysis come from several sources. From the Center for Research on Security Prices (CRSP), we obtained ADR characteristics, such as prices, returns, market capitalization, trading activity, bid-ask spreads, and exchange listing. From the World Bank, we gathered information about GDP per capita, capital formation, and GDP growth rates in each of the ADR home countries. Our institutional quality measures came from a number of different sources. Like Eleswarapu and Venkataraman (2006), we gathered much of the

[3] In particular, we use as a measure of institutional quality an index that captures the rights of shareholders to challenge the board of directors through a number of different mechanisms.

institutional quality data from La Porta et al. (1998). For instance, the Business International Corporation measured the efficiency of the judicial and legal system on a scale from 1 to 10. La Porta et al. (1998) also quantified the anti-director rights from 0 to 6. This index captures the rights of shareholders to challenge the board of directors through proxy voting, preemptive rights, and so on. La Porta et al. (1998) gathered an index ranging from 1 to 100 that measured accounting standards from the Center for International Financial Analysis and Research (CIFAR). Lastly, we used an index obtained from La Porta et al. (1998) that measures political stability. In particular, the International Country Risk Guide (ICRG) has created an index capturing political stability that ranges from 1 to 100.

Summary statistics for ADRs by home country are reported in table 1. Our sample includes ADRs from 35 countries and includes 277 individual securities. We examine ADRs that trade on US exchanges from 2001 to 2006. For each ADR, we estimate price delay and portfolio delay as described in Hou and Moskowitz (2005). In particular, we estimate the following equation using Wednesday-to-Wednesday weekly returns for each ADR during each year.

$$R_{i,t} = \alpha_i + \beta_i Rm_t + \sum_{j=1}^{4} \gamma_{i,t-1\ to\ t-j} Rm_{t-1\ to\ t-j} + \varepsilon_{i,t} \tag{1}$$

Table 1. Summary Statistics—Price Efficiency Measures

	No. of ADRs	Delay	Portfolio Delay	Noise	Variance Ratio
	[1]	[2]	[3]	[4]	[5]
Argentina	9	0.4872	0.3103	0.0169	0.9907
Australia	9	0.5474	0.3874	0.0083	1.0323
Austria	1	0.6181	0.3852	0.0042	1.0175
Belgium	1	0.2473	0.1065	0.0065	1.1032
Brazil	8	0.3806	0.1958	0.0146	1.0522
Chile	14	0.4877	0.3507	0.0059	1.0883
Denmark	2	0.5417	0.3309	0.0095	1.0177
Finland	3	0.2423	0.0943	0.0113	1.0576
France	21	0.3347	0.2551	0.0111	1.0286
Germany	15	0.3284	0.2557	0.0094	1.0270
Greece	4	0.4947	0.3148	0.0134	0.9817

(continued on next page)

	No. of ADRs	Delay	Portfolio Delay	Noise	Variance Ratio
	[1]	[2]	[3]	[4]	[5]
Hong Kong	9	0.5559	0.4088	0.0210	1.0024
India	12	0.4390	0.2803	0.0191	1.0859
Indonesia	2	0.5391	0.2973	0.0098	1.1064
Ireland	7	0.3662	0.1784	0.0177	1.0187
Israel	4	0.5574	0.4476	0.0091	1.0149
Italy	11	0.3394	0.2412	0.0069	1.0329
Japan	22	0.4107	0.2964	0.0082	1.0321
Mexico	19	0.3941	0.2453	0.0123	1.0643
Netherlands	14	0.3205	0.2162	0.0101	1.0341
New Zealand	1	0.5754	0.2997	0.0088	1.0264
Norway	4	0.4524	0.3043	0.0082	0.9904
Peru	1	0.6809	0.3379	0.0159	1.0366
Philippines	2	0.6522	0.3645	0.0213	1.0107
Portugal	2	0.3843	0.1910	0.0090	1.0475
Singapore	2	0.3524	0.3658	0.0152	1.0243
South Africa	9	0.5666	0.3420	0.0157	1.0302
South Korea	11	0.4187	0.2358	0.0116	0.9749
Spain	4	0.1283	0.0572	0.0104	1.0438
Sweden	2	0.2187	0.1020	0.0151	1.0855
Switzerland	9	0.2892	0.1880	0.0091	1.0394
Taiwan	7	0.2672	0.2023	0.0160	0.9903
Turkey	1	0.4468	0.1736	0.0121	1.0852
United Kingdom	34	0.3829	0.2597	0.0132	1.0014
Venezuela	1	0.4694	0.1726	0.0125	1.0248
Average	**7.91**	**0.4262**	**0.2627**	**0.0120**	**1.0343**

Note: The table reports the statistics that summarize the sample of ADRs by home country. Column [1] shows the number of ADRs for each country. Columns [2] and [3] present the results delay and portfolio delay as described in Hou and Moskowitz (2005). Noise measures the intraday change in price that occurs outside of the open and close (temporary price changes), scaled by closing price. The variance ratio is a test of the linearity of the variance across different sampling intervals and is calculated using one-day and five-day intervals (see Lo and MacKinlay 1988). The bottom row of the table presents the average across country for each of the measures of price efficiency.

In equation (1), the dependent variable, $R_{i,t}$, is the weekly return for each ADR i during week t. The independent variables include the contemporaneous (value-weighted) market return Rm_t and the lagged market returns from week $t-1$ to $t-j(\mathrm{Rm}_{\mathrm{i,t-1\ to\ t\text{-}j}})$, where j is equal to 1, 2, 3, or 4. We estimate equation (1) and extract the R^2, which is denoted by Hou and Moskowitz (2005) as the unrestricted R^2. We then estimate equation (1) but restrict $y_{i,t\text{-}j} = 0$ so that the contemporaneous market return is the sole independent variable. From this second pass, we estimate the R^2s for each ADR in each year and denote this as the restricted R^2. Hou and Moskowitz (2005) define this second pass as the restricted R^2. We then estimate delay in the following way.

$$1 - \frac{Restricted\ R^2}{Unrestricted\ R^2} \tag{2}$$

The result from equation (2) is considered first-stage delay. Hou and Moskowitz (2005) describe the potential errors-in-variables problem and attempt to correct for this problem by estimating portfolio delay. Second-stage delay is calculated by sorting all ADRs in the sample into deciles based on market capitalization; then, within each size decile, we sort ADRs into deciles based on price delay. We then estimate delay for each ADR in each of the 100 portfolios and assign the average delay to each ADR in a specific portfolio.[4] The objective in doing so is to avoid the possibility of the error-in-variables problem described in Hou and Moskowitz (2005). We note, however, that we use both first-stage and second-stage (portfolio) delay as our first two measures of market efficiency to provide some robustness.[5]

The second measure of efficiency is a measure that captures the noisiness of prices. Noise is estimated by capturing temporary or transient intraday price shocks as discussed in Blau, Griffith, and Whitby (2018). In particular, if the current day's closing price is greater than the current day's opening price, noise is measured as the difference between the intraday high price and the closing price plus the difference between the opening price and the intraday low price, scaled by the closing price. If, on the other hand, the current day's closing price is less than the

[4] Market capitalization is measured at the ADR level and not the firm level in the home country.

[5] We note a limitation when using delay as a measure of inefficiency. The market return in equation (1) is defined as the market return in the United States given that the ADRs trade in US markets. While it might be more appropriate to use market returns in each of the home countries, data available for these markets is scarce. It is important to note, however, that cross-listed securities become subject to the institutional quality in the country to which the cross is listed. Eleswarapu and Venkantaraman (2006) provide a nice discussion on how cross-listing exposes firms to new types of institutional effects.

intraday opening price, then noise is measured as the difference between the intraday high price and the opening price plus the difference between the closing price and the intraday low price, scaled by the closing price.[6]

The final measure of efficiency is Lo and MacKinlay's (1988) variance ratio. The variance ratio is a test of the linearity of the return variances across different sampling intervals. If the series follows a random walk, the sample variance of a k-period return should equal k times the sample variance of a one-period return. Thus, the variance ratio is the variance of the k-period return divided by k times the one-period return. A series that follows a random walk will have a variance ratio equal to 1. Since we are examining the efficiency of individual securities, our focus is on variance ratios that deviate from 1. Thus, our modified variance ratio measure is simply the absolute value of the difference between the variance ratio and 1, with larger measures being less efficient. Our analysis focuses on tests where $k = 5$. However, the variance ratios when $k = 10$ and when $k = 20$ are similar to the variance ratios when $k = 5$.

Column [1] of table 1 shows the number of ADRs for each country. The average country in our sample has 7.91 ADRs that are traded on US exchanges. Columns [2] and [3] present the average delay and portfolio delay, respectively. Higher measures of delay are interpreted as an inverse measure of informational efficiency. Belgium, Finland, Sweden, and Taiwan have delay measures ranging from 0.2187 to 0.2672, with Spain having the lowest measure of delay at 0.1283. These countries appear to be the most efficient. In contrast, ADRs from countries such as Peru and the Philippines have less efficient prices with delay measures of 0.6809 and 0.6522, respectively. Similar patterns can be seen in portfolio delay, although the magnitudes are smaller.

[6] If more permanent changes in price are owing to information, then temporary changes in price could be attributed to noise traders. For example, if the stock price is increasing during the day and reaches an intraday high of $25 but closes at $23 a share, then the difference between the high and the close ($2) is the temporary price change that can be attributed to noise. The temporary price change during the day is scaled by the closing price.

As an alternative to delay, we use a measure of noise to proxy for market efficiency at the individual stock level. When focusing on noise in column [4], it appears that ADRs from Hong Kong, India, and the Philippines are the least efficient (have the largest noise measures) while Austria and Belgium seem to be the most efficient (have the smallest noise measures). When examining variance ratios in column [5], we find that ADRs from Belgium, Chile, Indonesia, and India have variance ratios farthest away from 1, indicating that these ADRs are the least efficient. On the other hand, the ADRs from the United Kingdom seem to be the most efficient since the variance ratio in these ADRs is closest to 1. It is important to note that, because of space constraints, we do not report the *p*-values from *F*-tests that determine whether the variance ratio estimates are significantly different from 1. However, significance levels are reported in later tests when we directly examine the Keech and Munger (2015) hypothesis. Since both positive and negative deviations from 1 indicate less efficient prices in variance ratios, in the tests that follow, we use the absolute value of the difference between the variance ratio and 1 in our analysis. Large differences equate to less efficient prices.

Summary statistics related to country-level measures of institutional quality are reported in table 2. Column [1] of table 2 shows the La Porta et al. (1998) measure of judicial and legal efficiency, where higher measures equate to more efficient judicial and legal systems. Measures of judicial and legal efficiency range from 2.50 to 10.00 in our sample. Several countries—mostly developed nations—have measures of 10, while countries like Indonesia, the Philippines, and Turkey have the smallest measures of judicial efficiency. We note that the average across countries is 7.90. Column [2] presents the composite index of political stability obtained from the ICRG. Finland, Ireland, Sweden, and Switzerland are the most politically stable countries in our sample with scores above 92. Indonesia and Venezuela have the least stable political environments with scores below 50. In addition to examining the legal and political environment, we also examine a country's accounting quality from the CIFAR. Although two of the countries in our sample, Indonesia and Ireland, do not have accounting quality measures, the remaining countries range from a low of 36 in Portugal to a high of 83 in Sweden. When examining anti-director rights in column [4], we note that the average anti-director rights score is 2.97 on a scale from 0 to 6.

Table 2. Summary Statistics—Institutional Quality Measures

	La Porta et al. Judicial System	Political Stability	Accounting Quality	Anti-Director Rights
	[1]	[2]	[3]	[4]
Argentina	6.00	62.50	45	4
Australia	10.00	88.50	75	4
Austria	9.50	89.50	54	2
Belgium	9.50	87.00	61	0
Brazil	5.75	62.50	54	3
Chile	7.25	77.50	52	5
Denmark	10.00	91.00	62	2
Finland	10.00	95.00	77	3
France	8.00	80.50	69	3
Germany	9.00	87.50	62	1
Greece	7.00	76.00	55	2
Hong Kong	10.00	80.50	69	5
India	8.00	56.00	57	5
Indonesia	2.50	48.00	n/a	2
Ireland	8.75	92.00	n/a	4
Israel	10.00	58.50	64	3
Italy	6.75	81.00	62	1
Japan	10.00	86.00	65	4
Mexico	6.00	68.00	60	1
Netherlands	10.00	94.00	64	2
New Zealand	10.00	91.00	70	4
Norway	10.00	89.50	74	4
Peru	6.75	65.00	38	3
Philippines	4.75	67.00	65	3
Portugal	5.50	84.50	36	3
Singapore	10.00	90.00	78	4
South Africa	6.00	64.00	70	5
South Korea	6.00	76.00	62	2
Spain	6.25	82.50	64	4
Sweden	10.00	92.00	83	3
Switzerland	10.00	92.50	68	2

(continued on next page)

	La Porta et al. Judicial System	Political Stability	Accounting Quality	Anti-Director Rights
	[1]	[2]	[3]	[4]
Taiwan	6.75	79.50	65	3
United Kingdom	10.00	90.00	78	5
Venezuela	6.50	49.50	40	1
Average	**7.90**	**78.09**	**62.09**	**2.97**

Note: The table reports the statistics that summarize the sample of ADRs by home country. Column [1] shows La Porta et al.'s (1998) measure of judicial efficiency. Column [2] presents the composite index of political stability obtained from the International Country Risk Guide. Accounting quality is an index that measures the quality of accounting standards obtained from the Center for International Financial Analysis and Research. Anti-director rights is an index that captures the corporate governance of various countries. In particular, the index, which is also obtained from La Porta et al. (1998), measures that ability of shareholders to use proxy votes to challenge the incumbent board of directors. The bottom row of the table presents the average across countries for each of the measures of institutional quality.

3. Empirical Results

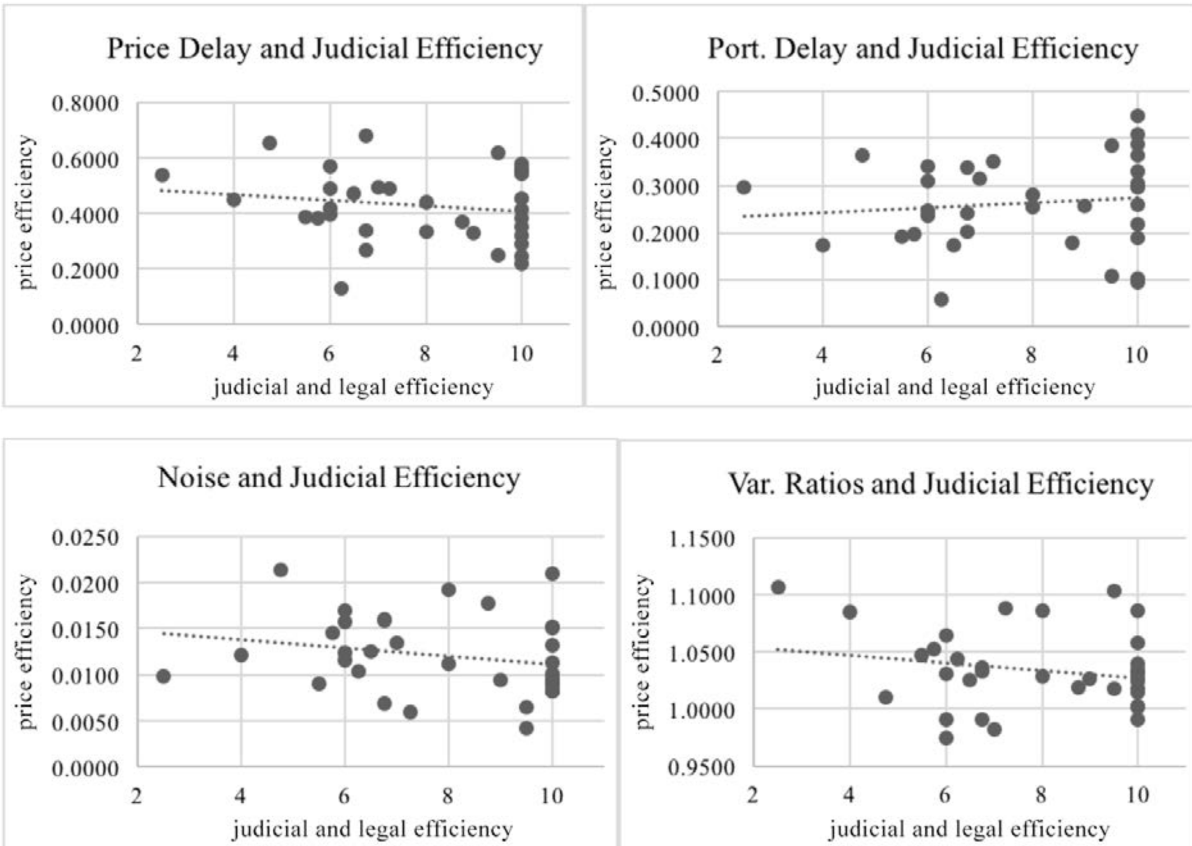
In this section, we test the hypothesis that institutional quality in a particular home country is associated with an improvement in the informational efficiency of ADRs. As mentioned above, we focus our analysis on ADRs that are traded in the United States as a way to control for the potential endogenous relationship between institutional quality and the structure of a particular financial market. Using ADRs allows us to hold the structure of the financial market constant while allowing for variation in the level of institutional quality across home countries. To that end, using ADRs allows us to better isolate the impact of institutional quality on firm-level market efficiency. However, as we have previously cautioned, the universe of ADRs is a fairly small sample and limits the generalizability of our findings. Given our small sample size, we have chosen to include countries that have relatively few US-listed ADRs in our tests.

3.1. Country-Level Scatterplots and Univariate Correlations

We begin by conducting a series of simple tests where we estimate the various measures of efficiency for the average ADR in a particular home country. We then provide a scatterplot of the efficiency measures across our institutional quality measures. We also report a trend line within the scatterplot to ascertain the univariate relationship between the informational efficiency of ADRs and institutional quality. Figures 1 through 4 report the results.

We begin by focusing on the scatterplots of our efficiency measures and the La Porta et al. (1998) measure of judicial and legal efficiency. In figure 1, and the figures that follow, the efficiency measures are reported on the y-axis while the institutional quality measures are shown on the x-axis. As seen in the upper left panel, the trend line is downward sloping. The correlation coefficient between delay and judicial and legal efficiency is −0.1628. In fact, in three of the four panels, our efficiency measures are negatively related with judicial and legal efficiency. In the bottom two panels, the correlation coefficients are −0.2243 and −0.2021. The only exception is when examining portfolio delay. In the upper right panel, the trend line is slightly upward sloping. The results in figure 1 seem to provide some evidence that institutional quality in the ADR home country—as measured by judicial and legal efficiency—improves the informational efficiency of ADRs.

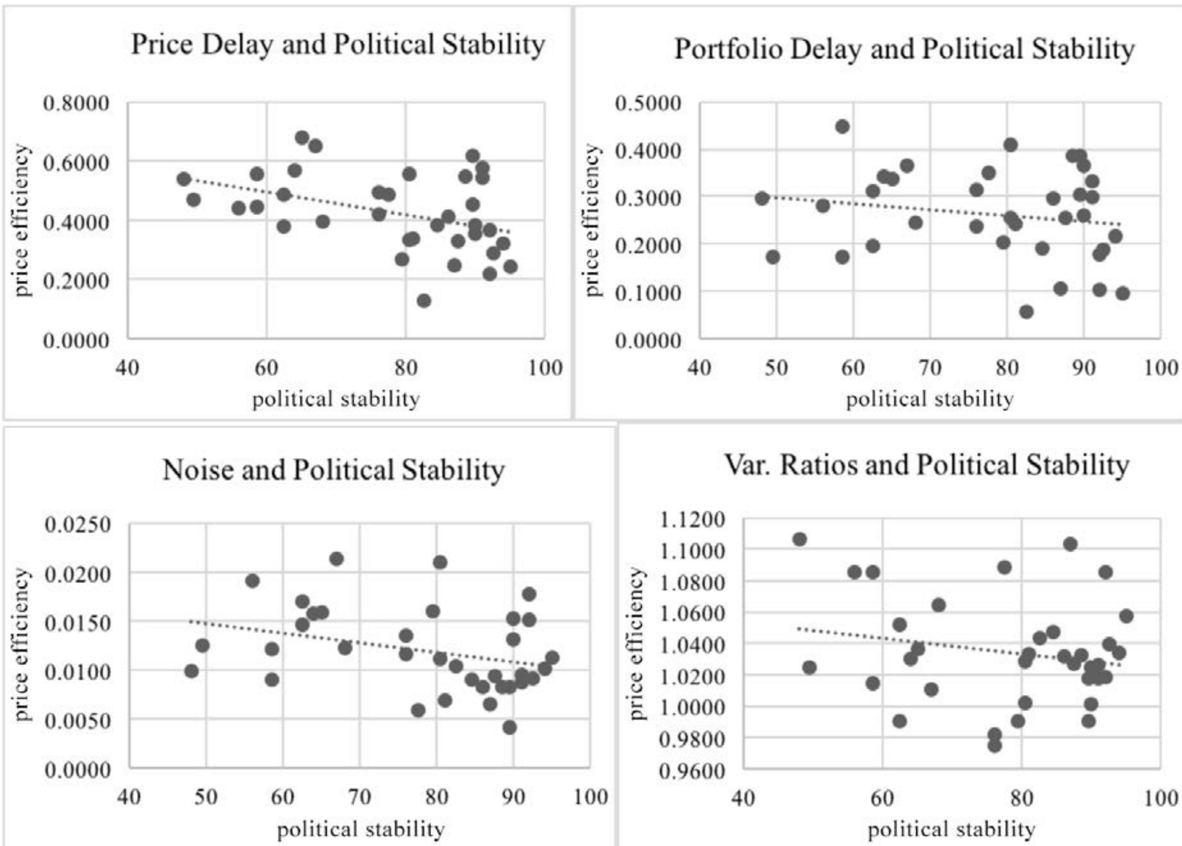
Figure 1. The Price Efficiency of ADRs and Judicial Efficiency



Note: The figure shows a scatterplot with a corresponding trend line of our measures of price efficiency (on the y-axis) and La Porta et al.'s (1998) measure of judicial efficiency (on the x-axis) by country.

Figure 2 shows the results when we use political stability as our measure of institutional quality. In this figure, we find that each of the efficiency measures is negatively associated with our institutional quality measures, as the trend lines in each panel are downward sloping. The correlation coefficients are −0.4091, −0.1855, −0.3256, and −0.1977, respectively. The results in figure 2 provide much stronger support for the idea that institutional quality (i.e., political stability) improves the efficiency of ADRs.

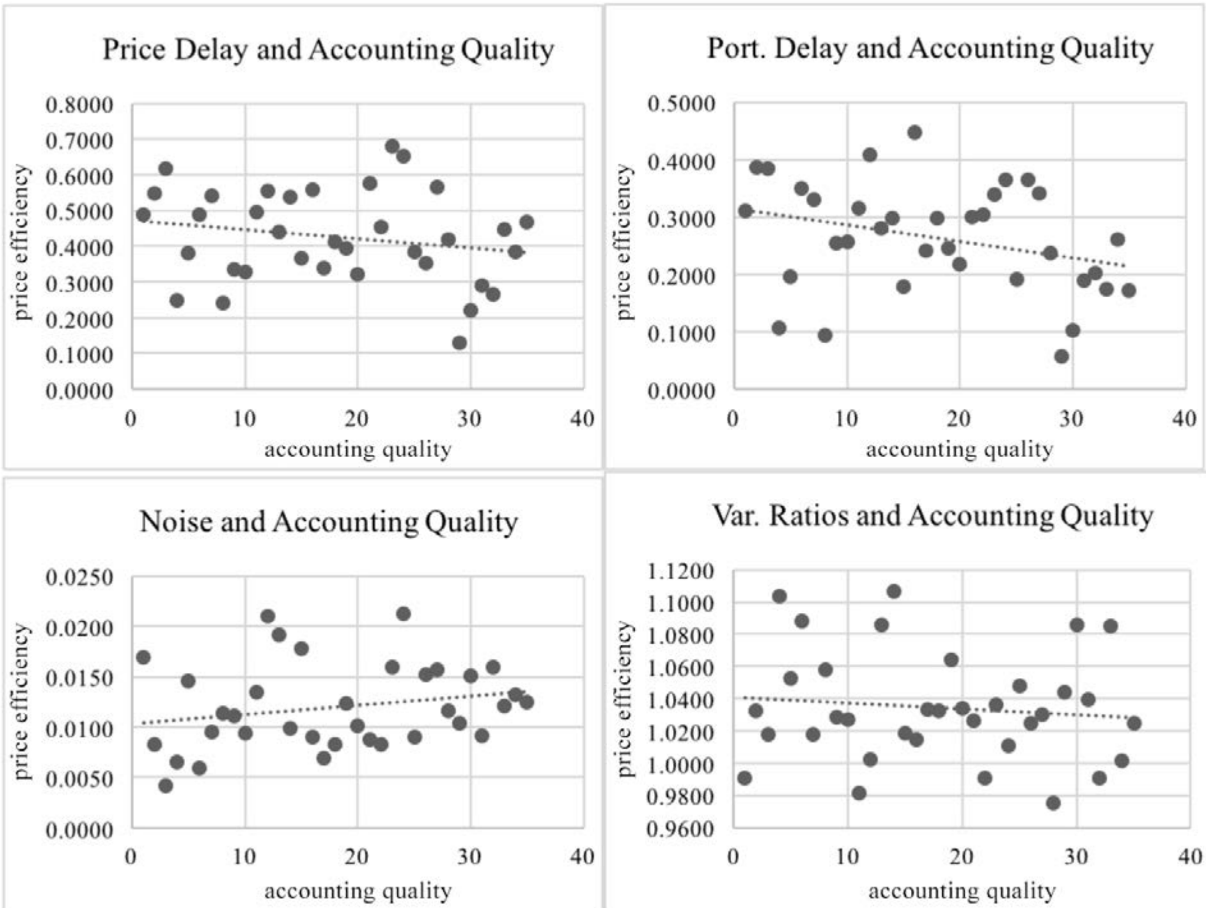
Figure 2. The Price Efficiency of ADRs and Political Stability



Note: The figure shows a scatterplot with a corresponding trend line of our measures of price efficiency (on the y-axis) and the ICRG measure of political stability (on the x-axis) by country.

In general, figure 3 provides some similar inferences: three of the four measures of price efficiency are negatively related to accounting quality. The only exception is in the bottom left panel when we examine the noisiness of prices. Again, figure 3 seems to be somewhat supportive of our hypothesis when focusing on accounting quality.

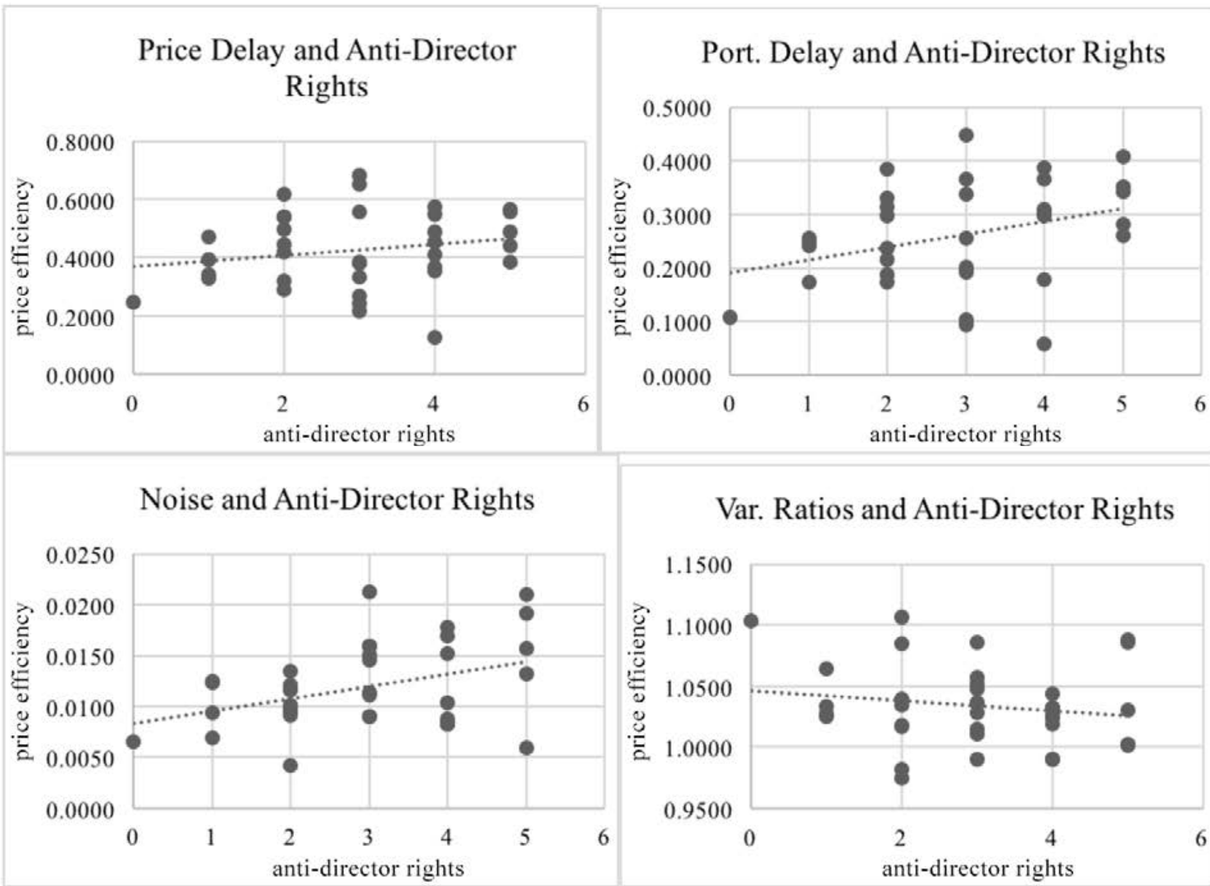
Figure 3. The Price Efficiency of ADRs and Accounting Quality



Note: The figure shows a scatterplot with a corresponding trend line of our measures of price efficiency (on the y-axis) and the CIFAR measure of accounting quality (on the x-axis) by country.

In our final set of univariate tests, we examine the association between our measures of efficiency and anti-director rights. Figure 4 shows the results from scatterplots that provide mixed findings. While portfolio delay and noise are positively related to anti-director rights, the opposite is true when looking at adjusted variance ratios. In the upper left panel, the relationship between delay and anti-director rights is relatively close to zero. Combined with earlier results, these results suggest that, in general, institutional quality is directly associated with our measures of efficiency. The strongest relation appears to be between informational efficiency and political stability, while the weakest association is between efficiency and anti-director rights.

Figure 4. The Price Efficiency of ADRs and Director Rights



Note: The figure shows a scatterplot with a corresponding trend line of our measures of price efficiency (on the y-axis) and La Porta et al.'s (1998) measure of anti-director rights (on the x-axis) by country.

3.2. Multivariate Tests

Thus far, the scatterplots have provided some support for our hypothesis. We recognize the need to control for other factors when attempting to isolate the association between institutional quality and the efficiency of ADRs. To do so, we estimate the following equation using a pooled sample of ADRs from 2001 to 2006.

$$Ln(Inefficiency_{i,t}) = \beta_0 + \beta_1 Ln(Institutional\ Quality_i) + \beta_2 NASDAQ_i + \beta_3 Ln(Size_{i,t}) + \beta_4 Ln(Turn_{i,t}) + \beta_5 Spread_{i,t} + \beta_6 Beta_{i,t} + \beta_7 IdioVolt_{i,t} + \beta_8 Ln(GDP/Cap_{i,t}) + \beta_9 Ln(CapForm)_{i,t} + \beta_{10} GDPgrowth_{i,t} + \varepsilon_{i,t} \quad (3)$$

Here, the dependent variable is the natural log of inefficiency, where inefficiency is one of our four measures of market inefficiency (delay, portfolio delay, noise, and the adjusted variance ratio). The independent variable of interest is the natural log of institutional quality, where institutional quality is one of the four measures discussed above as in Eleswarapu and Venkantaraman (2006). Given that a variety of known factors have been shown to influence the efficiency of stocks, we also include a number of control variables. *NASDAQ* is an indicator variable equal to one if ADR_i is listed on NASDAQ and zero otherwise. *Ln(Size)* is the natural log of market capitalization for each ADR in each year. *Ln(Turn)* is the natural log of share turnover. *Spread* is the relative bid-ask spread, which is the difference between the closing ask price and the closing bid price, scaled by the spread midpoint. *Beta* is the slope coefficient from a regression of daily returns for each ADR on the value-weighted CRSP index, and *IdioVolt* is the standard deviation of daily residual returns where residual returns are obtained from estimating a daily market model. *Ln(GDP/Cap)* is the natural log of GDP per capita. *Ln(CapForm)* is the natural log of capital formation in each home country. *GDP growth* is the growth rate of GDP in each home country. We note that the country-level data from the World Bank are not available for every home country. Specifically, data from Taiwan and Hong Kong are unavailable. Therefore, we report the results from estimating equation (3) with and without controls for macrolevel factors. We note that when estimating equation (3), we include year fixed effects and cluster standard errors across both ADRs and ADR home countries.

We begin our multivariate analysis by including *Ln(Judicial)* as the independent variable of interest in table 3. Columns [1] and [2] of table 3 estimate equation (3) with the *Ln(Delay)* as the dependent variable. With respect to the control variables, we find that delay is larger for smaller-cap stocks and for stocks with lower bid-ask spreads, lower betas, and higher idiosyncratic volatility. We also find the delay is negatively associated with GDP per capita and directly related to GDP growth rates. When focusing on the independent variable of interest, we find a negative association between the *Ln(Delay)* and judicial quality, with a coefficient of −0.3037 (*t*-statistic = −3.25) that is significant at the 0.01 level. However, once we control for macrolevel factors, the coefficient on *Ln(Judicial)*, while negative, loses its significance, as seen in column [2]. A similar pattern is repeated in columns [3] and [4], which use the natural log of portfolio delay as the dependent variable. So, although the quality of the legal system appears negatively related to efficiency when measured by delay, the relation does not appear to be robust to controls for GDP per capita, capital formation, and GPD growth rates in the ADR home country.

Columns [5] and [6] of table 3 estimate equation (1) when the natural log of noise is used as the dependent variable. We find that NASDAQ-listed stocks, larger-cap stocks, higher turnover stocks, high beta stocks, and stocks with higher idiosyncratic volatility have higher measures of noise. We also find some evidence that spreads and capital formation are negatively associated with noise, while GDP growth rates are positive related to noise. When focusing on judicial and legal efficiency, we find a negative relation between the variable of interest and noise. The coefficients on *Ln(Judicial)* are −0.3488 and −0.3117 (*t*-statistics = −6.66 and −2.87)—both of which are significant at the 0.01 level. In economic terms, a 1 percent increase in *Ln(Judicial)* is associated with a 0.35 percent (from column [5]) to a 0.31 percent (from column [6]) reduction in noise. These results seem to strongly support our hypothesis, as we find that judicial efficiency has a robust, negative relationship with noise.

Columns [7] and [8] estimate equation (1) with the natural log of the adjusted variance ratio, which is the absolute value of the difference between the variance ratio and 1, as our dependent variable. Since both positive and negative deviations from 1 indicate less market efficiency for the variance ratio, we take the absolute value of those differences as our measure of inefficiency. Thus, larger differences equal less efficient prices. Of all the independent variables, only *Ln(Judicial)* produces a coefficient that is significantly different from zero. Once again, we find that the quality of the judicial and legal system is negatively related to efficiency. We note, however, that when controlling for macrolevel factors, the coefficient on *Ln(Judicial)* is no longer reliably different from zero. Although our measures of delay are not robust to country-specific control variables, we find, at a minimum, some support for the notion that better legal institutions translate into more efficient prices.

Table 3. Price Efficiency and the Quality of the Legal System

	Ln(Delay)		Ln(P_Delay)		Ln(Noise)		Ln(|VarRatio − 1|)	
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
Intercept	0.4156	0.6489	2.0707***	3.4766***	−6.4874***	−5.0346***	−2.8728***	−2.1326**
	(1.30)	(1.12)	(6.11)	(5.36)	(−19.54)	(−8.89)	(−5.39)	(−2.19)
Ln(Judicial)	−0.3037***	−0.0234	−0.2337**	0.0776	−0.3488***	−0.3117***	−0.3526*	−0.1270
	(−3.25)	(−0.19)	(−2.13)	(0.54)	(−3.71)	(−2.87)	(−1.93)	(−0.59)
NASDAQ	0.0814	0.0866	0.0053	−0.0095	0.3143***	0.2968***	0.0670	0.0324
	(1.30)	(1.36)	(0.08)	(−0.15)	(5.04)	(4.75)	(0.63)	(0.30)
Ln(Size)	−0.0660***	−0.0688***	−0.2822***	−0.2803***	0.1163***	0.1164***	0.0231	0.0237
	(−3.88)	(−3.95)	(−15.85)	(−15.92)	(7.03)	(7.10)	(0.92)	(0.94)
Ln(Turn)	−0.0103	−0.0148	−0.0465*	−0.0461*	0.0840***	0.0850***	0.0221	0.0219
	(−0.45)	(−0.67)	(−1.78)	(−1.91)	(3.22)	(3.39)	(0.61)	(0.59)
Spread	−6.0814***	−5.2402***	−10.0204***	−8.2038***	−3.1429*	−2.2928	1.8679	2.7183
	(−3.32)	(−3.01)	(−4.59)	(−4.01)	(−1.78)	(−1.29)	(0.57)	(0.81)
Beta	−0.8133***	−0.8220***	−0.4933***	−0.4857***	0.2102***	0.2532	−0.0329	0.0131
	(−13.58)	(−13.92)	(−8.12)	(−8.37)	(4.93)	(6.24)	(−0.41)	(0.16)
IdioVolt	16.9557***	15.2591***	12.8507***	10.3486***	29.4202***	28.4518***	3.9345	2.7167
	(6.28)	(5.62)	(5.52)	(4.50)	(10.97)	(10.66)	(1.16)	(0.80)
Ln(GDP/Cap)		−0.0979***		−0.0965***		0.0092		−0.0664
		(−3.13)		(−2.74)		(0.40)		(−1.28)
Ln(CapForm)		0.0439		−0.4036**		−0.5969***		−0.2216
		(0.29)		(−2.53)		(−4.18)		(−0.95)
GDP growth		0.0262**		0.0442***		0.0348***		0.0254
		(2.50)		(3.20)		(3.94)		(1.35)
R^2	0.3446	0.3589	0.4599	0.4719	0.5494	0.5688	0.0290	0.0363
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	1,506	1,471	1,506	1,471	1,506	1,471	1,506	1,471

Note: The table reports the results from estimating the following equation using a sample of ADRs from 2001 to 2006.

$$Ln(Delay_{i,t}) = \beta_0 + \beta_1 Ln(Judicial_i) + \beta_2 NASDAQ_i + \beta_3 Ln(Size_{i,t}) + \beta_4 Ln(Turn_{i,t}) + \beta_5 Spread_{i,t} + \beta_6 Beta_{i,t} + \beta_7 IdioVolt_{i,t} + \beta_8 Ln(GDP_{i,t}) + \beta_9 Population_{i,t} + \varepsilon_{i,t}$$

The dependent variables are the natural log of each of our four measures of price efficiency (delay, portfolio delay, noise, and |VarRatio − 1|). The independent variables include the following: *Ln(Judicial)* is the natural log of La Porta et al.'s (1998) measures of judicial efficiency; *NASDAQ* is an indicator variable equal to 1 if ADR *i* is listed on NASDAQ—zero otherwise; *Ln(Size)* is the natural log of market capitalization for each ADR in each year; *Ln(Turn)* is the natural log of share turnover; *spread* is the relative bid-ask spread, which is the difference between the closing ask price and the closing bid price, scaled by the spread midpoint; *beta* is the slope coefficient from a regression of daily returns for each ADR on the value-weighted CRSP index; *IdioVolt* is the standard deviation of daily residual returns where residual returns are obtained from estimating a daily market model; *Ln(GDP)* is the natural log of gross domestic product; and *population* is the population in billions. In parentheses, we report corresponding *t*-statistics obtained from robust standard errors. We also include year fixed effects. *,**, and *** reflect statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Next, we turn our attention to the quality of political institutions across ADR home countries. To do so, we estimate a variation of equation (3) but include *Ln(PolStab)*, which is the natural log of political stability, as the independent variable of interest. We use the ICRG index of political stability that was used in Eleswarapu and Venkataraman (2006) and La Porta et al. (1998) as our measure of the quality of political systems. Once again, we include a number of stock-specific and country-specific control variables that could influence the efficiency of ADR prices. Columns [1] through [4] examine the relation between our delay measures and political stability. In general, the control variables produce coefficients that are similar in sign to the corresponding coefficients in the previous table. When focusing on the independent variable of interest, we find a strong negative relation between political stability and price delay with coefficients that range from −0.1534 in column [4] to −0.8872 in column [1]. Moreover, three of the four specifications on the delay measures are significant at the 0.01 level. The less stable the political system, the longer it takes ADR prices to incorporate relevant, market-wide information.

In columns [5] and [6], we examine the relation between political stability and noise. Here, we find consistency with our hypothesis because the estimates for *Ln(PolStab)* are negative and reliably different from zero. Without controls for country-specific factors (GDP and population), we find a coefficient of −0.5026. Once we control for country-specific variables, we find a coefficient of −0.5568. The results in these two columns are both statistically and economically significant. For instance, the estimate in column [6] indicates that a 1 percent increase in political stability is associated with a 0.56 percent decrease in noise. Similar to our findings with delay, noise appears to be negatively associated with political stability—the more stable a country's political system, the less noise we find in ADR prices. Columns [7] and [8] report results for our final dependent variable, which is the natural log of the adjusted variance ratio. Here, we again find support of our hypothesis because political stability is negatively associated with the adjusted variance ratio, indicating that the quality of political institutions is associated with an improvement in the informational efficiency of ADRs. More specifically, we find that the coefficients on $Ln(|VarRatio - 1|)$ are −0.8320 (t-statistic = −3.22) and −0.8969 (t-statistic = −1.71). These results seem to suggest that ADRs from countries that have less political stability are less likely to follow a random walk.

Table 4. Price Efficiency and the Quality of the Political System

	Ln(Delay)		Ln(P_Delay)		Ln(Noise)		Ln(|VarRatio – 1|)	
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
Intercept	3.6464***	2.7650***	4.7688***	3.9322***	−5.0671***	−3.6664***	−0.0004	0.2525
	(5.74)	(3.28)	(6.35)	(4.09)	(−8.42)	(−4.27)	(−0.01)	(0.16)
Ln(PolStab)	−0.8872***	−0.7801***	−0.7306***	−0.1534	−0.5026***	−0.5568**	−0.8320***	−0.8969*
	(−6.66)	(−3.18)	(−4.73)	(−0.56)	(−4.05)	(−2.15)	(−3.22)	(−1.71)
NASDAQ	0.0902	0.0858	0.0143	−0.0046	0.2994***	0.2782***	0.0676	0.0253
	(1.49)	(1.39)	(0.22)	(−0.07)	(4.70)	(4.40)	(0.65)	(0.24)
Ln(Size)	−0.0627***	−0.0647***	−0.2797***	−0.2803***	0.1216***	0.1223***	0.0276	0.0294
	(−3.71)	(−3.75)	(−15.83)	(−15.94)	(7.38)	(7.34)	(1.10)	(1.14)
Ln(Turn)	−0.0069	−0.0112	−0.0436*	−0.0447*	0.0843***	0.0849***	0.0246	0.0251
	(−0.32)	(−0.52)	(−1.72)	(−1.86)	(3.23)	(3.36)	(0.68)	(0.68)
Spread	−5.2549***	−5.0913***	−9.3521***	−8.2645***	−2.5657	−1.8628	2.6829	2.9998
	(−2.95)	(−2.97)	(−4.39)	(−4.02)	(−1.41)	(−1.03)	(0.80)	(0.89)
Beta	−0.8207***	−0.8281***	−0.4992***	−0.4860***	0.2051***	0.2456***	−0.0401	0.0049
	(−14.18)	(−14.06)	(−8.52)	(−8.43)	(4.74)	(5.97)	(−0.49)	(0.06)
IdioVolt	15.2350***	15.3456***	11.4004***	10.4329***	28.8362***	28.2713***	2.4730	2.7335
	(5.64)	(5.71)	(4.80)	(4.59)	(10.35)	(10.44)	(0.73)	(0.80)

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	Ln(Delay)	Ln(P_Delay)	Ln(Noise)	Ln(|VarRatio – 1|)		Ln(Delay)	Ln(P_Delay)	Ln(Noise)
	[1]	[2]	[3]	[4]		[1]	[2]	[3]
Ln(GDP/Cap)		0.0063		−0.0634		0.0385		0.0381
		(0.16)		(1.41)		(0.95)		(0.45)
Ln(CapForm)		0.0874		−0.3924**		−0.5557***		−0.1750
		(0.58)		(−2.42)		(−3.94)		(−0.74)
GDP growth		0.0276**		0.0451***		0.0333***		0.0261
		(2.57)		(3.26)		(3.73)		(1.40)
R^2	0.3578	0.3623	0.4658	0.4718	0.5481	0.5663	0.0346	0.0387
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	1,506	1,471	1,506	1,471	1,506	1,471	1,506	1,471

Note: The table reports the results from estimating the following equation using a sample of ADRs from 2001 to 2006.

$$Ln(Delay_{i,t}) = \beta_0 + \beta_1 Ln(PolStab_i) + \beta_2 NASDAQ_i + \beta_3 Ln(Size_{i,t}) + \beta_4 Ln(Turn_{i,t}) + \beta_5 Spread_{i,t} + \beta_6 Beta_{i,t} + \beta_7 IdioVolt_{i,t} + \beta_8 Ln(GDP_{i,t}) + \beta_9 Population_{i,t} + \varepsilon_{i,t}$$

The dependent variables are the natural log of each of our four measures of price efficiency (delay, portfolio delay, noise, and |VarRatio − 1|). The independent variables include the following: *Ln(PolStab)* is the natural log of the ICRG index of political stability; *NASDAQ* is an indicator variable equal to 1 if ADR *i* is listed on NASDAQ—zero otherwise; *Ln(Size)* is the natural log of market capitalization for each ADR in each year; *Ln(Turn)* is the natural log of share turnover; spread is the relative bid-ask spread, which is the difference between the closing ask price and the closing bid price, scaled by the spread midpoint; *Beta* is the slope coefficient from a regression of daily returns for each ADR on the value-weighted CRSP index; *IdioVolt* is the standard deviation of daily residual returns where residual returns are obtained from estimating a daily market model; *Ln(GDP)* is the natural log of gross domestic product; and *population* is the population in billions. In parentheses, we report corresponding *t*-statistics obtained from robust standard errors. We also include year fixed effects. *,**, and *** reflect statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

Now we turn our attention to the quality of accounting systems in different ADR home countries. Since accounting standards affect the information environment of securities, it is natural to associate the standards companies use when reporting their financial details with market efficiency. Following Eleswarapu and Venkataraman (2006) and La Porta et al. (1998), we use the CIFAR index of accounting standards as our measure of accounting system quality. The index measures the quality of accounting standards in each of the home countries. Table 5 reports the results when we include *Ln(AcctQual)* as our measure of institutional quality. As before, the control variables produce coefficients that are similar in sign to the corresponding coefficients found in tables 3 and 4. Similar to our analysis of judicial or legal institutions, columns [1] through [4] show that delay is negatively related to accounting standards but is not significantly different from zero when we control for country-specific factors. In contrast, the relation between noise and accounting quality is reliably different from zero whether or not we include macrolevel controls. For instance, column [5] reports the noise results without country-level controls, and column [6] reports the noise results with country-level controls. The coefficient on *Ln(AcctQual)* goes from −0.4172 to −0.3371, respectively (*t*-statistics = −4.66 and −2.05). In economic terms, a 1 percent increase in accounting quality is associated with a 0.42 percent decrease in noise in column [5]. We note that, similar to noise, the negative relation between adjusted variance ratios and accounting quality appears to be robust to country-level controls. For instance, the coefficients on *Ln(AcctQual)* are −0.8194 and −0.7096 (*t*-statistics = −2.88 and −2.19), respectively. Combined, the findings in table 5 provide some evidence that accounting institutions of higher quality are associated with more efficient stock prices.

Table 5. Price Efficiency and the Quality of the Accounting System

	Ln(Delay)		Ln(P_Delay)		Ln(Noise)		Ln(|VarRatio – 1|)	
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
Intercept	2.1660***	1.4464*	3.0201***	3.5473***	−5.6733***	−4.0038***	−0.4184***	−0.0524
	(3.19)	(1.85)	(3.93)	(4.18)	(−7.92)	(−4.94)	(−0.34)	(−0.04)
Ln(AcctQual)	−0.5863***	−0.2505	−0.3421*	0.0353	−0.4172***	−0.3371**	−0.8194***	−0.7096**
	(−3.66)	(−1.46)	(−1.92)	(0.19)	(−2.64)	(−2.05)	(−2.88)	(−2.19)
NASDAQ	0.0977	0.0972	−0.0038	−0.0263	0.2966***	0.2786***	0.1098	0.0704
	(1.55)	(1.52)	(−0.06)	(−0.41)	(4.51)	(4.31)	(1.03)	(0.65)
Ln(Size)	−0.0616***	−0.0696***	−0.2841***	−0.2892***	0.1288***	0.1260***	0.0371	0.0356
	(−3.68)	(−4.01)	(−16.62)	(−17.34)	(7.72)	(7.54)	(1.49)	(1.41)
Ln(Turn)	−0.0071	−0.0151	−0.0524**	−0.0546**	0.0952***	0.0961***	0.0317	0.0307
	(−0.32)	(−0.69)	(−1.99)	(−2.26)	(3.60)	(3.80)	(0.88)	(0.84)
Spread	−5.8036***	−5.2888***	−9.7857***	−8.2315***	−2.8578	−1.9005	2.5453	3.1031
	(−3.15)	(−3.07)	(−4.51)	(−4.07)	(−1.54)	(−1.03)	(0.77)	(0.92)
Beta	−0.8140***	−0.8215***	−0.4844***	−0.4736***	0.2070***	0.2536***	−0.0270	0.0246
	(−13.67)	(−13.97)	(−7.87)	(−8.24)	(4.54)	(5.91)	(−0.34)	(0.31)
IdioVolt	16.5791***	14.9785***	12.3291***	9.6702***	30.4305***	28.9170***	3.8917	3.2559
	(6.05)	(5.47)	(5.37)	(4.46)	(10.47)	(10.13)	(1.13)	(0.97)
Ln(GDP/Cap)		−0.0770***		−0.0851***		−0.0148		−0.0131
		(−2.84)		(−2.68)		(−0.68)		(−0.24)

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	Ln(Delay)	Ln(P_Delay)	Ln(Noise)	Ln(|VarRatio – 1|)		Ln(Delay)	Ln(P_Delay)	Ln(Noise)
	[1]	[2]	[3]	[4]		[1]	[2]	[3]
Ln(CapForm)		0.0406		−0.4256***		−0.6348***		−0.2741
		(0.27)		(−2.67)		(−4.58)		(−1.16)
GDP growth		0.0301***		0.0458***		0.0318***		0.0316*
		(2.82)		(3.23)		(3.70)		(1.70)
R^2	0.3500	0.3634	0.4649	0.4773	0.5460	0.5693	0.0346	0.0412
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	1,456	1,421	1,456	1,421	1,456	1,421	1,456	1,421

Note: The table reports the results from estimating the following equation using a sample of ADRs from 2001 to 2006.

$$Ln(Delay_{i,t}) = \beta_0 + \beta_1 Ln(AcctQual_i) + \beta_2 NASDAQ_i + \beta_3 Ln(Size_{i,t}) + \beta_4 Ln(Turn_{i,t}) + \beta_5 Spread_{i,t} + \beta_6 Beta_{i,t} + \beta_7 IdioVolt_{i,t} + \beta_8 Ln(GDP_{i,t}) + \beta_9 Population_{i,t} + \varepsilon_{i,t}$$

The dependent variables are the natural log of each of our four measures of price efficiency (delay, portfolio delay, noise, and |VarRatio − 1|). The independent variables include the following: *Ln(AcctQual)* is the natural log of the CIFAR index of accounting standards; *NASDAQ* is an indicator variable equal to 1 if ADR *i* is listed on NASDAQ—zero otherwise; *Ln(Size)* is the natural log of market capitalization for each ADR in each year; *Ln(Turn)* is the natural log of share turnover; *spread* is the relative bid-ask spread, which is the difference between the closing ask price and the closing bid price, scaled by the spread midpoint; *beta* is the slope coefficient from a regression of daily returns for each ADR on the value-weighted CRSP index; *IdioVolt* is the standard deviation of daily residual returns where residual returns are obtained from estimating a daily market model; *Ln(GDP)* is the natural log of gross domestic product; and *population* is the population in billions. In parentheses, we report corresponding *t*-statistics obtained from robust standard errors. We also include year fixed effects. *,**, and *** reflect statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

In our final set of tests, we replicate our multivariate analysis performed in tables 3 through 5 but examine the effect of corporate governance quality in the ADR home countries (*Ln(Director)*) on the efficiency of ADR prices. Table 6 reports the results from estimating equation (3) with *Ln(Director)* as the independent variable of interest. La Porta et al. (1998) document which countries have the most liberal corporate governance rules by creating an index numbered 0 to 6, where countries with a 6 are those that have a framework that allows shareholders to challenge incumbent boards of directors. The weakest of our empirical tests, table 6 shows the results from the analysis. *Ln(Director)* produces negative estimates in five of the eight columns. However, the coefficients are only significant in column [6], suggesting that the noise of ADRs is lower when home countries have more liberal corporate governance rules. We recognize, however, that these results are not robust to measures of price delay (in columns [1] through [4]) and specifications that do not include country-level control variables.

Table 6. Price Efficiency and the Quality of Corporate Governance

	Ln(Delay)		Ln(P_Delay)		Ln(Noise)		Ln(|VarRatio – 1|)	
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
Intercept	−0.3808	0.7045	1.4148	3.6331	−7.2232***	−4.8102***	−3.5821***	−1.7842*
	(−1.58)	(1.14)	(5.47)	(5.28)	(−27.63)	(−8.01)	(−8.84)	(−1.81)
Ln(Director)	0.0635	−0.0266	0.0783	0.0474	−0.0498	−0.1491**	−0.0741	−0.1629
	(0.99)	(−0.42)	(1.20)	(−0.67)	(−0.81)	(−2.40)	(−0.70)	(−1.48)
NASDAQ	0.0368	0.0895	−0.0339	0.0029	0.2834***	0.3018***	0.0396	0.0509
	(0.58)	(1.39)	(−0.53)	(0.04)	(4.36)	(4.80)	(0.38)	(0.47)
Ln(Size)	−0.0599***	−0.0688***	−0.2773***	−0.2816***	0.1226***	0.1179***	0.0293	0.0231
	(−3.61)	(−3.98)	(−15.72)	(−15.96)	(7.27)	(7.13)	(1.16)	(0.91)
Ln(Turn)	−0.0149	−0.0143	−0.0509**	−0.0441*	0.0823***	0.0862***	0.0212	0.0251
	(−0.65)	(−0.63)	(−1.97)	(−1.83)	(3.14)	(3.45)	(0.59)	(0.68)
Spread	−5.6506***	−5.2939***	−9.5810***	−8.4301***	−3.1118*	−2.3955	1.8103	2.3718
	(−3.09)	(−3.05)	(−4.43)	(−4.12)	(−1.67)	(−1.34)	(0.54)	(0.71)
Beta	−0.8078***	−0.8250***	−0.4855***	−0.4898***	0.2024***	0.2343***	−0.0436	−0.0053
	(−13.31)	(−14.01)	(−7.93)	(−8.53)	(4.58)	(5.76)	(−0.54)	(−0.07)
IdioVolt	17.3788***	15.3050***	13.0673***	10.5282***	30.3773***	28.5612***	0.4531***	3.0115
	(6.27)	(5.61)	(5.59)	(4.62)	(10.83)	(10.59)	(3.71)	(0.89)
Ln(GDP/Cap)		−0.1017***		−0.0848***		−0.0395*		−0.0870*
		(−4.18)		(−3.06)		(−1.95)		(−1.85)

(continued on next page)

	Ln(Delay)	Ln(P_Delay)	Ln(Noise)	Ln(|VarRatio – 1|)		Ln(Delay)	Ln(P_Delay)	Ln(Noise)
	[1]	[2]	[3]	[4]		[1]	[2]	[3]
Ln(CapForm)		0.0335		−0.4182***		−0.6411***		−0.2847
		(0.22)		(−2.61)		(−4.57)		(−1.25)
GDP growth		0.0274**		0.0473***		0.0401***		0.0330*
		(2.57)		(3.20)		(4.54)		(1.72)
R^2	0.3402	0.3589	0.4586	0.4719	0.5389	0.5681	0.0255	0.0380
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	1,506	1,471	1,506	1,471	1,506	1,471	1,506	1,471

Note: The table reports the results from estimating the following equation using a sample of ADRs from 2001 to 2006.

$$Ln(Delay_{i,t}) = \beta_0 + \beta_1 Ln(Director_i) + \beta_2 NASDAQ_i + \beta_3 Ln(Size_{i,t}) + \beta_4 Ln(Turn_{i,t}) + \beta_5 Spread_{i,t} + \beta_6 Beta_{i,t} + \beta_7 IdioVolt_{i,t} + \beta_8 Ln(GDP_{i,t}) + \beta_9 Population_{i,t} + \varepsilon_{i,t}$$

The dependent variables are the natural log of each of our four measures of price efficiency (delay, portfolio delay, noise, and |VarRatio − 1|). The independent variables include the following: *Ln(Director)* is the natural log of the LLSV (1998) measure of anti-director rights; *NASDAQ* is an indicator variable equal to 1 if ADR *i* is listed on NASDAQ—zero otherwise; *Ln(Size)* is the natural log of market capitalization for each ADR in each year; *Ln(Turn)* is the natural log of share turnover; *spread* is the relative bid-ask spread, which is the difference between the closing ask price and the closing bid price, scaled by the spread midpoint; *beta* is the slope coefficient from a regression of daily returns for each ADR on the value-weighted CRSP index; *IdioVolt* is the standard deviation of daily residual returns where residual returns are obtained from estimating a daily market model; *Ln(GDP)* is the natural log of gross domestic product; and *population* is the population in billions. In parentheses, we report corresponding *t*-statistics obtained from robust standard errors. We also include year fixed effects. *,**, and *** reflect statistical significance at the 0.10, 0.05, and 0.01 levels, respectively.

In summary, we are left to conclude that the quality of institutions seems to provide some effect on the informational efficiency of security prices. The strongest results are found when examining political institutions, while the weakest results occur when examining the rights of shareholders relative to boards of directors.

4. Conclusion

Much of the finance literature has focused on determining whether markets are efficient, or if market prices reflect all available information. As an anecdote, the presence of bubbles and subsequent crashes seems to question the efficiency of markets. Furthermore, the real effect of such crashes on the macroeconomy has initiated calls by some for more government intervention. Keech and Munger (2015) point out, however, that while the call for government intervention in response to so-called market failures is common, government failures may have preceded the market failures, given that governments establish and oversee the institutions necessary for the success of markets. In finance, the call for government intervention to correct for financial market inefficiency seems misguided, if indeed the institutional framework is associated with market inefficiency. This study tests the Keech and Munger (2015) hypothesis by examining the association between institutional quality and the informational efficiency of security prices in financial markets.

Tests of this hypothesis are difficult to conduct given the need for cross-sectional variation in the level of institutional quality. While institutional quality will no doubt vary across countries, the financial market structure, which will heavily influence the efficiency of prices, is likely to be endogenously determined by the countries' institutional framework. To overcome these issues, we follow Eleswarapu and Venkataraman (2006) and use a sample of ADRs, which are certificates traded in the United States but representing shares of foreign stocks. This research design allows us to overcome this potential endogeneity while still allowing for variation in the level of institutional quality across the ADR home countries.

Results show a significant, cross-sectional relationship between institutional quality in the ADR home country and the efficiency of ADR prices. Using the measures of institutional quality in Eleswarapu and Venkataraman (2006), we find that these measures are negatively associated with various measures of market inefficiency. The results are strongest for political stability and weakest for the liberalness of corporate governance rules. These results seem to indicate that security prices are the most efficient when institutional quality is highest.

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