Three Essays on International Finance and International Capital Markets

by

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The difficulty lies not in the new ideas but in escaping from the old ones.

- John Maynard Keynes, 1935

To Mom and Dad

Who know unconditional love

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LIST OF ABBREVIATIONS

ADR American Depositary Receipt

EM Emerging Market

DEV Developed Market

GAAP Generally Accepted Accounting Principles

SEC U.S. Securities and Exchange Commission

GDR Global Depositary Receipt

CR Capital-Raising

IPO Initial Public Offering

IFS International Financial Statistics

MSCI Morgan Stanley Capital International

bps Basis Points

ETF Exchange-Traded Fund

ARIMA Autoregressive Integrated Moving Average

ABSTRACT

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This dissertation consists of three essays examining informational and behavioral frictions in international financial and capital markets. Chapter II investigates whether characteristics of the home country capital market environment, such as information disclosure and investor rights protection, continue to affect American Depositary Receipts (ADRs) cross-listed in the U.S. I find that characteristics of the home markets continue to be relevant, especially for emerging market firms. Less transparent disclosure, poorer protection of investor rights and weaker legal institutions are associated with higher levels of information asymmetry. My finding suggests that cross-listing in the U.S. should not be viewed as a substitute for improvement in the quality of local institutions. Chapter III addresses the question of whether it is possible to profit from timing the exchange rate markets by examining foreign firms' decision to issue ADRs. Specifically, we test whether foreign firms consider currency market conditions in their ADR issuance decisions. We find that foreign firms tend to issue ADRs after their local currency has been abnormally strong against the U.S. dollar and before their local currency becomes abnormally weak. Currency market timing is especially significant for companies who are more likely to be affected by higher currency exposure and emerging market companies. It is more pronounced during currency crises and after the market integration, and when the ADR issue raises capital. Currency market timing is also economically significant. These findings suggest that some companies may have private information about foreign exchange market. Chapter III examines behavioral bias in global financial markets. It investigates the relation between lunar phases and stock market returns of 48 countries. The findings indicate that stock returns are lower on the days around a full moon than on the days around a new moon. The magnitude of the return difference is 3% to 5% per annum based on analysis of global portfolios. The return difference is not due to changes in stock market volatility or trading volumes. The lunar effect is not explained away by announcements of macroeconomic indicators, nor is it driven by major global shocks. Moreover, the lunar effect is independent of other calendar-related anomalies.

CHAPTER I

Introduction

Informational frictions have always been a major issue in finance. In our modern financial world, markets expand wider and deeper, information is flowing seemingly costlessly along our INTERNET connections. However, the information that is required to evaluate financial assets is not straightforward and not equally available to all market participants. Research in the areas of home bias (e.g., Tesar and Werner (1995)) and international capital flows (e.g., Brennan and Cao (1997), Portes and Rey (2005)) all points to the evidence that information issues remain important in our increasingly integrated but still segmented global financial markets. The consequence of not fully understanding the information issues involved in our complex financial world, meanwhile, could be very severe, as the global financial crisis in 2007 and 2008 has witnessed. Information asymmetry affects asset prices and capital allocation, exacerbates financial crisis and contagion, and distorts incentives in corporate financing decisions. Also, since international capital markets are segmented by different investment laws and institution qualities of sovereign nations, corporate governance also plays an important role in capital allocation.

Two chapters of this dissertation are related to this broad area. One chapter (Chapter II) studies the impact of the home country capital market environment on the adverse selection cost of the firms' U.S. exchange listed depositary rights, and its implications for government policies aimed at attracting global capital flows. Another chapter (Chapter III) studies firms' ability to time transactions in the currency market when raising capital internationally. In a broader sense, both chapters first try to establish evidence that certain types of information asymmetry persist in international financial markets.[1] Chapter II then focuses on the implications of information asymmetry for capital allocations, while Chapter III focuses more on its implications for firms' capital raising decisions, i.e., when to issue ADRs to raise capital.

Behavioral bias is another reason why financial markets can be less efficient than a classical model predicts. Research evidence suggests that investors are subject to various psychological and behavioral biases. Such biases may play a relatively more important role in international market where informational problems are more pronounced.[2] Chapter IV investigates the lunar effect in asset prices. It presents a test of a particular behavioral hypothesis using an exogenously identifiable variable, motivated by psychological hypothesis. It finds global evidence of a significant correlation between stock returns and lunar phases.

Cross-listing has increasingly become an important form of raising capital abroad. Previous research on cross-listing argues that foreign firms can improve their disclosure and governance by cross-listing in the United States in order to bond themselves to U.S. security regimes. The degree of effectiveness of such bonding, however, remains an open question. In Chapter II, we examine the cross-section of the information asymmetry surrounding ADR firms, and relate it to information disclosure quality, governance, and rule of law measures of the countries in which those firms are domiciled. We find that less transparent disclosure, poorer protection of investor rights, and weaker legal institutions in the countries of origin are associated with higher level of adverse selection in ADRs, especially for firms from emerging markets.

Methodologically, Chapter II can be viewed as a crossover study between corporate

[1] In Chapter II's case, information asymmetry of cross-listed firms in equity markets; and in Chapter III's case, information asymmetry of exchange rate movements in currency markets.

[2] e.g., Schmeling (2009)

governance and market microstructure, with international cross-listing as the subject of the study. It uses a novel approach to proxy for the degree of information asymmetry of ADRs with market microstructure measures such as the bid-ask spread, its adverse selection component, and revealed information in the volume-return relationships. Market microstructure measures of information asymmetry capture adverse selection between informed traders and uninformed traders, as well as the financial market's perception of information advantage held by firm insiders and the resulting adverse selection cost. We measure ADR firms' average quoted spread and effective spread. To extract adverse selection in our spread proxy, we also decompose quoted spreads into an asymmetric-information component and a non-information component. We show that microstructure measures of adverse selection of ADRs are positively correlated with information opaqueness and poor protection of investor rights in the capital market environment of the home countries, and this relation is more pronounced among emerging market ADRs. In other words, the home stigma exists among ADRs traded on the U.S. exchanges. In addition, we provide evidence that this "home stigma" risk appears to be priced by the market.

Cross-sectionally, the difference is significant between firms from emerging markets and firms from more mature capital markets. Among emerging market ADRs, the connection between the home capital market environment and their adverse selection in U.S. trading is much stronger than their developed market counterparts. Moreover, information opacity of emerging markets does not appear to be solely due to the lack of regulation. In fact, survey-based transparency measure fits them better than accounting-rule-based measure, suggesting the lack of enforcement of the existing rules. Among developed market ADRs, the home stigma phenomenon is weaker, but there is evidence that home market accounting standards and quality of legal institution continue to affect their level of adverse section in the U.S. trading.

Robustness of the results is rigorously checked by using alternative measures of adverse selection and statistical methods. We control our empirical regressions of market condition and individual firm characteristics. Given that many firm-level controls such as size, leverage, and book-to-market ratio are possibly correlated with the level of information asymmetry, our tests are actually biased against finding additional adverse selection. We also augment our tests with analyst coverage and forecast dispersion, which, prior research indicates, are associated with a firm's disclosure and information environment. Our results withstand those controls. To alleviate concern about time variation in home-market information transparency, we confirm our primary findings with alternative time-varying measures of information derived from the average R^2 of respective markets. To further substantiate our tests of adverse selection in ADRs, we confirm our results utilizing an alternative proxy that is based on the volume-return relationship. Finally, we also confirm our empirical tests with several statistical methods.

Our finding that the home capital market environment continues to act as a crucial factor in adverse selection of ADR firms makes an important contribution to the bonding literature. The home capital market environment can affect a firm's incentive to adhere to a higher disclosure standard, and therefore affect the level of adverse selection in their U.S. listed securities. Our results suggest that this is especially a concern for emerging market firms. In terms of policy implications, Chapter II highlights the challenges presented to the emerging market firms when they try to attract capital internationally. It illustrates that cross-listing in the United States is not a substitute for improvement of local information transparency and protection of minority shareholders. Governments and policymakers should still focus on promoting information disclosure and improving corporate governance in order to attract capital flows to their countries. Our finding that quality of governance matters for the information environment also suggests that policies aiming to improve domestic disclosure should be complemented with policies promoting better investor rights protection.

Chapter III examines asymmetric information in foreign exchange markets in relation to firms' capital structure decisions. The core notion of the market timing theory is that companies attempt to raise capital by issuing overvalued securities. From an international perspective, the relative overvaluation of the currency may be crucial to firms tapping into the foreign capital market as well. Chapter III examines the firms' timing ability in exchange rate markets. Specifically, we test whether foreign firms consider currency market conditions in their ADR issuance decisions, and thus display some ability to time their local exchange rate market. We find that foreign firms tend to issue ADRs after their local currency has been abnormally strong against the U.S. dollar and before their local currency becomes abnormally weak.

Chapter III employs two methods to investigate currency market timing ability. We start by analyzing the behavior of exchange rate returns in proximity of ADR issuance dates using a standard event study methodology. More specifically, we regress de-trended exchange rate returns on a set of event dates in a window of six months before and after the ADR issuances. De-trending allows us to extract the expected component of exchange rate fluctuations out of our measure of exchange rate returns. Since the expected change of exchange rate returns may be already priced into the ADR offering by the equity market, these changes give the issuing firm little incentive to time the currency market. We find that (the detrended) cumulative abnormal exchange rate returns display a U-shape pattern around ADR issue dates. They decrease before ADR issuance and increase afterward. This suggests that firms tend to issue ADRs when their domestic currency is abnormally strong. In our sub-sample analysis, we split our ADR sample into Capital-Raising (CR) issues and non-Capital-Raising (non-CR) issues. Since non-CR ADRs generate no net revenue for the issuing firm, it provides a natural control group to control for the reason of ADR issues other than market timing. We find that these cumulative returns around CR ADR issue dates display a much more pronounced U-shape profile than for the whole sample, while no such evidence is found in the control sample of non-CR ADRs.

Then we apply an alternative methodology of Poisson regressions. More specifically we estimate the effect of both abnormal currency and local and U.S. stock returns in the holding period on the probability of the ADR issue decision via a Poisson regression model. This approach allows us to control for the effects of stock return dynamics on the ADR issue decision, to take into account the multiple issues in the same month, and to identify firms' best timing horizons. The results are consistent with the patterns we identified with the event study approach. Estimates for the exchange rate return coefficients over the whole sample are negative for windows prior to ADR issuances and mostly statistically significant, and positive for all windows afterward. In the sub-sample regressions, we find that CR ADR issues are significantly more likely than non-CR to occur after an excess appreciation of and before an excess depreciation of the local currency. Our Poisson analysis also shows that extent of the currency market timing varies greatly across regions. It is most pronounced across the sub-sample of emerging market issuers. The more pronounced currency market timing abilities of the emerging markets firms motivate us to look closer into emerging market issuances, in particular, the effects of financial crisis and market integration on the currency market timing. We find that emerging market firms possess superior currency market timing ability in the proximity of crisis periods, which suggests those firms are on average successful in recognizing the symptoms of an impending dramatic depreciation of their local currencies. And this timing ability is more pronounced after financial integration has occurred in emerging markets.

After documenting that firms are able to time foreign exchange market through ADR issues, Chapter III then investigates what kind of issuances and firms are more likely to time the exchange rate market. We investigate effects of relative ADR issue size or issuing firm size, Tobin's q, the firm's industry, and the identity of the underwriting investment bank, respectively. We find that relatively big issues from relatively small firms, firms with lower Tobin's q, i.e., firms with relatively stable market valuations, and firms in the manufacturing industry are more likely to be able to time the exchange market. We do not find evidence that exchange market timing abilities vary across investment bank groupings. This evidence corroborates our intuition that firms whose valuations are more likely to be affected by exchange rate fluctuations are more likely to try to time the market.

The finding in Chapter III suggests that some companies may have, at least occasionally, information advantage about foreign exchanges. It also confirms the prior evidence that emerging market firms have private information about economic fundamentals.[3]

Chapter IV asks a simple question: Do lunar phases affect security markets? If investors make decisions strictly through rational maximization, then the answer is no. However, research evidence suggests that investors are subject to various psychological and behavioral biases. If lunar phases affect mood, by extension, these phases may affect investor behavior and thus asset prices. Since the lunar cycle has little tangible impact on people's economic and social activities in modern societies, investigating lunar effect on security returns is a strong test of whether investor behavior affects asset prices. The causality is obvious.

To investigate the relation between lunar phases and stock returns, Chapter IV first examines the association of lunar phases with the returns of an equally-weighted and a value-weighted global portfolio of 48 country stock indices. The findings indicate that global stock returns are significantly lower during full moon periods than new moon periods. For the equally-weighted global portfolio, the cumulative return difference between the new moon periods and the full moon periods is 40.26 Basis Points (bps) per lunar cycle for the 15-day window specification and 27.48 bps per lunar cycle for the 7-day window specification; both are significant at the 5% level. Es-

[3] Kaufmann, Mehrez, and Schmukler (2005)

timation of value-weighted global portfolio yields similar results. A sinusoidal model of continuous lunar impact is then used to test for the cyclical pattern of the lunar effect. The model suggests that the lunar effect is cyclical.

To fully utilize the cross-sectional and time series data, Chapter IV then estimates a pooled regression with panel corrected standard errors (PCSE). We include in the regression a lunar dummy that takes on a value of one for a full moon period and zero for a new moon period. The coefficient on this variable indicates the difference in the mean daily logarithmic returns between the lunar periods. The PCSE specification adjusts for the contemporaneous correlation and heteroskedasticity among country index returns, as well as for the autocorrelation within each country's stock index returns. The results of pooled regression indicate that stock returns of the new moon periods are significantly higher than those of the full moon periods. Regardless of window specifications, the coefficients on the lunar dummy are negative. When all countries are included in the analysis, the returns of the new moon periods are on average 3.95 bps and 5.93 bps higher than returns of the full moon periods for the 15-day and 7-day windows, respectively. Both estimates are significant at the 5% level. The estimated coefficients on the lunar dummy remain similar when country group dummy variables are included. When country fixed effects are included, the estimated coefficients become larger in magnitude and higher in statistical significance. Interestingly, the extent of the lunar effect seems to vary across different levels of market maturity. The strongest lunar effect is found among the emerging market countries: a 5.60 bps daily difference for the 15-day window and an 11.27 bps daily difference for the 7-day window; both are significant at the 5% level.

To further identify the lunar effect, we examine whether lunar effects are related to stock capitalization, trading volume and market volatility, and macroeconomic events. We find that the estimated lunar effect is stronger for NASDAQ stocks than for NYSE and Amex stocks. Moreover, the lunar effect is in general stronger for smaller size deciles. Since large capitalization stocks have a higher percentage of institutional ownership than small capitalization stocks, these findings are consistent with the hypothesis that stocks with more individual investor ownership display a stronger lunar effect and thus provide further evidence that mood or sentiment may affect asset prices. We do not find evidence that lunar effect is due to patterns of trading volume or stock market volatility. To test the possibility that the return differential between the full moon and the new moon periods reflects the average effect of macroeconomic events or common market shocks, we investigate the lunar effect on the global portfolio, controlling for the following three types of events: macroeconomic announcements, major global shocks, and movements in short-term interest rates. Overall, the results indicate that lunar effect cannot be explained away by these factors.

For further robustness tests, we examine whether the lunar effect can be explained by other calendar anomalies: the January effect, day of week effect, calendar month effect, and holiday effect including lunar holiday effect. We also check the robustness by examining various lunar window lengths, alternative Autoregressive Integrated Moving Average (ARIMA) specifications, and a test of random 30-day cycles. Overall, the findings indicate that lunar effect cannot be explained away by other calendar anomalies and is robust to various lunar window lengths, alternative ARIMA specification of stock returns, and is not a result of some random 30-day cycle. The lunar effect found in Chapter IV provides some evidence of behavioral bias in global stock markets.

CHAPTER II

The Home Stigma: Adverse Selection in ADRs and the Home Capital Market Environment

2.1 Introduction

This chapter examines whether characteristics of the home-country capital market environment, such as information disclosure and investor rights protection, continue to be relevant for foreign firms cross-listed on U.S. exchanges. We find that less transparent disclosure, poorer protection of investor rights, and weaker legal institutions in the countries of origin are associated with higher level of adverse selection in ADRs, especially for firms from emerging markets. Poor home capital market environment increases the average bid-ask spread by 10% with one standard deviation in disclosure quality, and the effect is economically significant. Our finding illustrates that cross-listing in the United States is not a substitute for improvement of local information transparency and protection of minority shareholders.

Foreign issuers are subject to U.S. securities laws concerning disclosures and procedures for equity issuances and are required to reconcile their accounting statements with U.S. Generally Accepted Accounting Principles (GAAP) (using form 20-F). Some studies (e.g., Coffee (1999), Stulz (1999), Reese and Weisbach (2002)) argue that cross-listing in the U.S. serves as a substitute for weak investor protection laws in the home country. Extending that logic, foreign firms can also bond themselves to stricter disclosure requirements and stricter enforcement in order to achieve a better information environment and attract investors who would otherwise be reluctant to invest.

A number of empirical studies have tested the bonding hypothesis, using proxies to capture this effect. Doidge (2004), for example, finds that the premium between voting and non-voting shares declines following cross-listing, an indication that minority investors become better protected. Reese and Weisbach (2002) interpret the frequency and location of equity offerings that follow cross-listing as evidence of firms from weak investor protection countries seeking to bond with the U.S. security regime. Lang, Lins, and Miller (2003) find that cross-listed firms achieve greater analyst coverage with improved forecast accuracy, and Bailey, Karolyi, and Salva (2006) find evidence of higher liquidity. Other studies find that an enhancement in the shareholder protection and information environment, in turn, is associated with higher valuation of the cross-listed firm (Foerster and Karolyi (1999), Doidge, Karolyi, and Stulz (2004), and Bailey, Karolyi, and Salva (2006)).

However, more recently, questions have been raised about the effectiveness of bonding. There are several possible reasons why bonding might not be effective. First, regulation and enforcement of U.S. corporate laws might not be as stringent for foreign firms. Foreign issuers can obtain exemptions from various disclosure requirements of exchanges and regulators. Anecdotal evidence suggests that the U.S. stock exchanges and the U.S. Securities and Exchange Commission (SEC) are not as stringent with foreign firms in enforcing listing standards. Siegel (2005) argues that the lack of effective law enforcement by the SEC and minority shareholders against cross-listed foreign firms adds to the ineffectiveness of regulatory bonding, since it is very difficult to prosecute foreign perpetrators across the border. Further, rules governing corporate bankruptcy and derivative actions against foreign insiders are based on the laws in the companies' home country. Second, mechanisms to credibly commit a firm to higher-quality information disclosure and governance may be unavailable or prohibitively expensive in countries with poor investor protection and insufficient economic development (Doidge, Karolyi, and Stulz (2006), Ball (2001)). Ball (2001) argues that changing accounting standards systems alone is not enough to improve actual financial reporting and disclosure. Third, if firms want to misrepresent their information within the U.S. regulation, they have varieties of means to do so. They can for example, engage in earnings smoothing, aggressive accounting, or loss avoidance.

The question then remains whether bonding is sufficient enough to make U.S. listed foreign firms equally transparent regardless of cross-sectional differences in the home market information environment and corporate governance. This chapter seeks to shed some light on this question.

It is conceivable that the home market information environment and corporate governance still play an important role in cross-sectional differences in information asymmetry of cross-listed foreign firms, even when bonding is effective. First, many disclosure practices are voluntary. SEC and other regulatory rules mean to serve as a floor for necessary reporting. In practice, companies can, and may in fact desire to, disclose more than what SEC regulation requires. Therefore, actual levels of voluntary disclosure may differ for firms from different markets. Second, firms from different countries may have different costs and benefits in implementing measures to improve governance and transparency. Hence their incentives to increase the level of voluntary disclosure and to bond to the U.S. security regime through reputation mechanism may differ. The incentive to improve is low if, for example, there are large controlling blocks of shares and minority investors are not well protected, or when a firm is difficult to effectively differentiate itself because the home market information environment is poor. Doidge et al. (2006), for example, find country characteristics play a dominant role in determining the level of disclosure, and the effect is significant for ADR firms. Third, even when all necessary information is disclosed, it takes effort to decipher and disseminate useful information. This is more challenging for firms from less transparent home markets. Lang, Lins, and Miller (2004) argue that U.S.-based analysts are less likely to follow non-U.S. firms with large family or management-owned controlling blocks of shares, especially for companies with weak legal protections in the home market. It is also possible that the added reporting and disclosure required by regulators for cross-listing could crowd out or substitute for the collection of private information (Kim and Verrecchia (2001)).

Most of the existing empirical research on the bonding theory focuses on finding support for the effectiveness of bonding. However, there is evidence that institutional features of the local environment find their way into U.S. cross-listed firms. Lang, Raedy, and Wilson (2006) show that U.S. cross-listed firms report financial statement information systematically differently than equivalent U.S. firms despite the fact that all firms use the same accounting standards. Foreign firms recognize losses in a less timely manner and generally report more smoothed earnings. Therefore, a cross-listed firm's home environment continues to be relevant in explaining the quality of its U.S. GAAP reported earnings. Doidge et al. (2006) find country characteristics are significant for ADR firms, especially for those from emerging markets. Bailey et al. (2006) report cross-sectional evidence in volume and volatility that suggests cross-listed firms from emerging markets have lower information quality. Ferreira and Fernandes (2008) show the improvement in price informativeness after cross-listing is concentrated in developed market firms; cross-listing is negatively associated with price informativeness in emerging market firms.

In this chapter, we examine the cross-section of the information asymmetry surrounding ADR firms, and relate it to information disclosure quality, governance, and rule of law measures of the countries in which those firms are domiciled. We utilize market microstructure proxies to measure information asymmetry. Market microstructure measures of information asymmetry are designed to capture adverse selection between informed traders and uninformed traders. Since the firm managers and those close to them constitute an important subset of informed traders in the market, market microstructure measures also should capture adverse selection faced by ADR firms, albeit imperfectly. More important, those proxies capture the financial market's perception of information advantage held by firm insiders and the resulting adverse selection cost. Bharath, Pasquariello, and Wu (2008) use microstructure proxies to establish the relationship between information asymmetry and capital structure decisions. Heflin, Shaw, and Wild (2005) find that there is a negative association between disclosure quality and bid-ask-spread-based measures of information asymmetry in the U.S. stock market.

We measure the opacity of a country's information environment with the home country financial opaqueness, since it is a direct concern as a source of adverse selection. A less transparent information environment cannot only lead to a firm's poor disclosure practice, but also downgrade the market's perception about the quality of its disclosure. We also include legal and governance measures in our investigation of adverse selection. Although there are distinctions between financial opaqueness and poor protection of investors, a country's financial opaqueness is often coupled, and mutually reinforced, with imperfect protection of shareholder rights and poor corporate governance. Moreover, as discussed above, investor protection and corporate governance can influence a firm's incentive to adhere to a high level of information disclosure, and ultimately affect adverse selection costs. Thus, we include legal institution and corporate governance variables not only as control variables to investigate whether home country financial transparency remains as a significant factor of adverse selection for ADR firms, but also to examine their direct impact on information asymmetry.

Our primary empirical results indicate that the home market environment still matters for adverse selection of non-U.S. firms cross-listed on U.S. exchanges. ADRs of firms from countries with greater financial opaqueness, and poorer investor protection and corporate governance experience a higher degree of adverse selection, especially when issued by firms from emerging markets. Among those emerging market firms, both quoted bid-ask spreads and effective spreads are negatively correlated with the quality-of-disclosure and governance proxies. Among firms from developed markets, there is weaker evidence that poorer home country information transparency and firm governance lead to bigger quoted spreads. Instead, legal origin of business laws (La Porta, Lopez-de-Silanes, Shleifer, and Vishny (1998)) appears to be a consistent and significant factor in their level of adverse selection. To extract the adverse selection in our spread proxy, we also decompose quoted spreads into an asymmetric-information component and a non-information component. We find that the asymmetric-information component of spreads exhibits a negative relation with domicile market transparency and governance, especially for emerging market firms. It gives additional support that our microstructure measure does indeed capture the adverse selection of ADR firms. Our results complement Ferreira and Fernandes' (2008) event study finding that cross-listing does not improve price informativeness for emerging market firms. We also provide some evidence that the "home country risk" is priced by the market. Abnormal return is larger for firms from countries of less information transparency and weaker institution quality.

Our inferences appear to be quite robust. We control our empirical regressions of the market condition and individual firm characteristics. Given that many firm-level controls such as size, leverage, and the book-to-market ratio are possibly correlated with level of information asymmetry, our tests are actually biased against finding an additional adverse selection component. We also augment our tests with analyst coverage and forecast dispersion, which prior research indicates are associated with a firm's disclosure and information environment (e.g., Lang and Lundholm (1996)). Our results withstand those controls. To alleviate concern about home market information transparency measure, we confirm our primary findings with an alternative time-varying measure of information environment advocated by Morck, Yeung, and Yu (2000) and Jin and Myers (2006). We also confirm our empirical tests with several statistical methods. Finally, to further substantiate our interpretation of spread-based proxies as a measurement of adverse selection, we conduct our tests utilizing an alternative proxy derived by Llorente, Michaely, Saar, and Wang (2002) that is not spread-based. The results of the alternative dependent variable show that the proxy of adverse selection is negatively correlated with home market transparency and better governance for emerging market sub-sample, but not for developed market firms. There are several possible explanations for the finding that developed market firms show less evidence of a relation between the home market environment and the adverse selection in ADRs than emerging market firms. Disclosure measures may not capture capital market characteristics among developed market firms; the cost of adopting better disclosure and governance practice may be small enough for developed market firms to bond themselves more effectively.

Our finding that the home capital market environment continues to act as a crucial factor in the adverse selection of ADR firms, especially for those from emerging markets, has important policy implications. It illustrates that cross-listing in the United States is not a substitute for the improvement of local information transparency and the protection of minority shareholders. In a recent study, Ammer, Holland, Smith, and Warnock (2008) document that if everything else is equal, cross-listing has a smaller impact on U.S. investors' holdings for firms from countries with weaker shareholder protection. Our findings help explain such a home bias in ADR holdings. Therefore, governments and policymakers should still focus on promoting information disclosure and improving corporate governance in order to attract capital flows to their countries.

The rest of the chapter proceeds as follows. Section 2.2 describes the data and provides discussion on some institutional backgrounds. Section 2.3 empirically investigates the relation between ADR information asymmetry and home market environments. Section 2.4 concludes.

2.2 Data

In this section, we describe sample selection and the institutional background, the measurement of firm-specific adverse selection and the home country capital market environment, and the control variables used in this study.

2.2.1 Sample Selection

To construct a uniform testing sample, we need to restrict our firms to similar institutional requirements. Different choices of listing type have different institutional requirements. Non-U.S. companies to be listed in the U.S. are required to file a registration statement with the SEC and furnish an annual report on a Form 20-F with reconciliation with U.S. GAAP. Level I ADRs trade over-the-counter and require only minimal SEC disclosure and no GAAP reconciliation (exempt from Form 20-F by Rule 12g3-2(b)). Level II and III ADRs are exchange-listed securities and they require full SEC disclosure and compliance with the exchange's own listing rules. Level III programs raise capital and must file Form F-2 and F-3 for offerings. Finally, SEC Rule 144a issues raise capital as private placements to qualified institutional buyers and do not require compliance with GAAP.[1] Ordinary listings require an exact replication of settlement facilities as for U.S. securities and go beyond Level II and

[1]it may be that a firm that accesses the U.S. markets by way of a Rule 144a private placement or OTC listing, which require little or no conformity with U.S. GAAP, actually chooses to disclose more because it anticipates a subsequent upgrade to an exchange listing at some later point. Similarly, a firm that chose to list its shares for trading on the NYSE or Nasdaq may have anticipated doing so for some time before and thus had chosen to present its financial statements voluntarily in accordance with IAS or U.S. GAAP in the years before the U.S. listing.

Level III ADRs in requiring full annual and quarterly reports prepared in accordance with U.S. GAAP. We focus on level II and III ADRs to ensure uniform regulatory disclosure requirements. We identify our sample of exchange-listed ADR firms on NYSE, AMEX or Nasdaq from Bank of New York and other sources. To the extent that level II and III ADRs are subject to stricter disclosure requirements than level I and Rule 144a issuances, our sample is biased against finding a home market effect.

We restrict our sample of ADRs to issues that occurred no earlier than 1990, and collect their price and volume information between 1994 and the end of 2006. The purpose of our restriction on issuance and price data is two-fold. First, foreign issuances become more active in the 1990s due to conscious effort by the U.S. to attract foreign issues and financial liberalization in other parts of the world; further, most emerging market ADRs are issued after 1990, and most emerging markets do not have local stock market indices until after 1994. An ADR issued earlier is more likely to be localized by the U.S. market. More important, most of our disclosure measures are time-invariant, and are constructed with data from late 1990s and early 2000s. Our restriction is to keep the data relevant to the question we study. Instead of focusing on a limited period of time (e.g., Lang et al. (2003)), we examine the whole period of time when most price information is available. We also drop ADRs for which institutional data is not available.[2] We end up with 355 firms from 36 countries over 13 years.

To avoid drawing spurious inferences from extreme values and to filter out errors in the data, we winsorize observations in the top and bottom 5% of the asymmetric information measure and firm size over the whole sample period. Also the availability of the specific control variables such as the number of analyst coverage might differ, so the number of observations for each individual regression may vary.

Stock price and return data is from CRSP and Datastream. Non-U.S. Market

[2] We do not include cross-listings from Canada, which are unique in their regulatory requirements.

returns and volatility are calculated based on Datastream total market indices. Firm accounting data is from Compustat. Analyst forecast data is from IBES.

2.2.2 Home Capital Market Environment

We use several measures to characterize the information environment of the capital market of a country. Financial transparency measures capture the intensity and the timeliness of financial disclosures, and their interpretation and dissemination by analysts. We draw our main proxy of disclosure, $GCRSCORE$, from The Global Competitiveness Reports, which includes results from surveys about the level and effectiveness of financial disclosure in different countries. The respondents were asked to assess the statement "The level of financial disclosure required is extensive and detailed" on a scale from 1 (strongly disagree) to 7 (strongly agree), as well as to assess "availability of information" on the same scale. For each country, we take the average over responses to the two questions in the 2000 survey and form a disclosure score $GCRSCORE$. Higher values of $GCRSCORE$ indicate higher degree of information transparency of home markets of ADR issuing firms. The advantage of this survey-based proxy is that it is possible to reflect variations in information disclosure beyond those stipulated in the accounting rules. This is most relevant for emerging markets where existing regulations on the book are not always properly enforced. Jin and Myers (2006) show that this survey-based measure works quite well as a proxy for financial transparency of the markets.

We also use an accounting-based index to augment $GCRSCORE$ disclosure measure, and to ensure the robustness of our results. The Center for International Financial Analysis and Research (CIFAR (1995)) reports for each country a disclosure index that represents a score based on the inclusion or omission of 90 items as required disclosures in the annual reports, with low scores indicating poor accounting standards. These items fall into seven categories: general information, income statements, balance sheets, funds flow statement, accounting standards, stock data, and special items. We denote this index as $CIFAR$.[3]

Figure 2.1 charts the values of $CIFAR$ and $GCRSCORE$ (scaled by 10 times) for countries in our sample. $GCRSCORE$ ranges from 3.8, of China, to 6.4, of Finland, with a median of 5.65, about the level of disclosure for Ireland. $CIFAR$ ranges from the lowest score of 56 for Brazil, to the highest score of 85 for the U.K., with a median of 73, the score for Hong Kong. $CIFAR$ covers slightly fewer countries (31) than $GCRSCORE$ (34), and covers fewer emerging markets. Numbers of ADRs from each country covered in our sample are also reported in Figure 2.1.

A country's financial opaqueness is often coupled with imperfect protection of shareholder rights and poor corporate governance. Nevertheless, there is a distinction between opaqueness and poor protection of investors, which can affect firms' incentive to voluntary disclose. We thus include legal institution and corporate governance variables, not only to investigate whether those remain as important factors in determining degree of information asymmetry in ADRs, but also to check whether our disclosure variables remain significant after controlling for the shareholder protection.

We use an index of anti-director rights compiled by La Porta, Lopez-De-Silanes, Shleifer, and Vishny (henceforth, LLSV)(1998), denoted $ANTIDIR$, to proxy for strength of corporate governance. It measures the ability of shareholders in a country to challenge managers or dominant shareholders in the corporate decision-making process. It ranges from zero to six, with higher score representing better investor protection. LLSV (1998) show that $ANTIDIR$ is a reasonable proxy for investor rights protection. When applicable, we employ a dummy variable $CIVIL$, which equals one if the home market legal rules are of *Civil Law* origin, and zero if legal rules regarding investor rights are of *Common Law* origin. LLSV (1998) show that, broadly

[3]These variables are from La Porta, Lopez-de-Silanes, Shleifer, and Vishny (1998).

speaking, common law countries afford the best legal protections to shareholders.

Table 2.1 reports the Spearman rank correlations among the disclosure and governance measures, as well as the summary statistics of each measure. It shows that the disclosure measures and the governance proxy are positively correlated with each other and negatively correlated with the civil law dummy. The survey-based disclosure measure $GCRSCORE$ and accounting-based disclosure measure $CIFAR$ have a high correlation of 0.69. The governance measure $ANTIDIR$ and legal origin are more correlated with each other than with disclosure measures in terms of magnitude. Correlation coefficient between $ANTIDIR$ and $CIVIL$ is -0.64, it is 0.13 between $ANTIDIR$ and $GCRSCORE$, and 0.36 between $ANTIDIR$ and $CIFAR$.

Notice that our home environment measures are time-invariant, but our sample spans over several years. Fortunately, institutional quality variables are relatively stable, and are very slow to change. Nevertheless, to check the robustness of our results, we also use a time-varying measure of home country information environment following Jin and Myers (2006) The details of that alternative measure will be explained in the next section.

2.2.3 Asymmetric Information

Information asymmetry measures at the firm level are notoriously hard to estimate and lack accuracy. Researchers have used IBES analyst dispersion of estimates, earning surprises, or firm's accounting discrepancies to proxy for it, with varying degrees of success. We use market microstructure measures to proxy for the degree of adverse selection. Specifically, we first examine bid-ask spreads of ADRs. Presence of traders who possess superior knowledge of the value of a stock can impose adverse selection costs on liquidity traders and market makers. Bid-ask spread compensates liquidity providers for bearing this cost and increases with degree of information asymmetry. Previous research finds the bid-ask spread to be a reasonable proxy for information asymmetry. Lee, Mucklow, and Ready (1993), for example show firms with lower spreads have a lower degree of information asymmetry or adverse selection costs.

We calculate the quoted spread in percentage terms with daily data as

$$S = 100 * (askprice - bidprice)/\frac{1}{2}(askprice + bidprice), \tag{2.1}$$

We then take the yearly average to obtain an average quoted spread, QS, for each stock every year.

The first column of Table 2.2 reports the summary statistics of the average quoted spread across all sample firms as well as across groups that are formed by the measures of disclosure and institutional quality. Quoted spread has a mean of 1.71 (percent), and a wide variation across firms and years with a standard deviation of 1.47. Firms from markets of lower $GCRSCORE$, lower $CIFAR$, lower $ANTIDIR$, and firms from countries of civil law origin and emerging markets all have relatively higher QS. We note that the finding that firms from civil law countries have higher QS is consistent with the findings in LLSV (1998) that civil law origin countries offer weaker legal protection of investor rights. The differences in the means of sub-samples are all significant except that of different $GCRSCORE$s. This gives us a first indication that QS of ADRs, and hence the degree of information asymmetry, is dependent on the information and legal environment of the home countries.

Quoted spread has components other than asymmetric-information reasons such as inventory, transaction, and order-processing costs. George, Kaul, and Nimalendran (1991) show a method of unbiasedly estimating the adverse selection component of a stock's spread by exploiting different stock return autocorrelations of uninformed trading and informational speculation. In light of this consideration, we also estimate the adverse selection component of the quoted spread and substitute that as the dependent variable, following George et al. (1991) and Bharath et al. (2008).

To substantiate our interpretation that quoted spread differences reflect the degree of information asymmetry of ADRs, we also test for the presence of a home environment effect using an alternative non-spread-based dependent variable that measures the relative importance of information-driven trading in stocks' price fluctuations. Llorente et al. (2002) show theoretically and empirically that accounting for the intensity of trading volume accompanying stock return autocorrelation can help identify the extent of informed trading. We develop our alternative proxy following their methodology. Specifically, we estimate

$$r_{it}(\tau) = c_{0it} + c_{1it} r_{it}(\tau - 1) + c_{2it} T_{it}(\tau - 1) r_{it}(\tau - 1) + \varepsilon_{it}(\tau), \tag{2.2}$$

for each firm i in year t, where r_{it} is the stock return of ADR i in year t, τ is an indicator for days in year t, $T_{it}(\tau)$ is the natural logarithm of the daily turnover detrended by its mean over the past 200 observations. The measure of informed trading is given by the coefficients c_{2it}, yielding one observation for each firm-year. For firms with considerable information asymmetry, c_{2it} tends to be positive, as more volume indicates more speculative trading and the stock exhibits positive return autocorrelation. For firms with low information asymmetry, c_{2it} tends to be negative, as more volume indicates liquidity-based (or noise) trading and the return exhibits negative autocorrelation. The lower (higher) is the estimated c_{2it}, the lower (higher) is firm i's degree of adverse selection.

2.2.4 Control Variables

We include a variety of firm characteristics variables to control for the cross-section of individual firm idiosyncrasy, including firm size, asset tangibility, leverage, and the book-to-market ratio.

$SIZE_{it}$ is defined as the logarithmic of firm i's total asset in year t, DTA_{it} is the debt-to-asset ratio, MTB_{it} is the market-to-book ratio, $TANG_{it}$ is the tangibility of firm assets, defined as the ratio of firm's tangible assets (Property, Plant and Equipment Total, Compustat fundamental data item PPENT) to total assets.

Our firm characteristics variables are all possibly related to the degree of information asymmetry of firms. Smaller firms are likely to have more severe information asymmetry problems. A firm with relatively less tangible assets, *ceteris paribus*, is expected to have a higher degree of information asymmetry. Low MTB also may reflect more severe information problem between managers and investors. LLSV (2002) show that firms with better investor protection have higher valuation relative to their assets, i.e., higher MTB. If information asymmetry is an important determinant of a firm's debt issuance decisions, then its cumulative effect, leverage, is likely to be higher for firms facing more severe information problems. Since our main interest is to test whether the home country information disclosure and governance environment have an effect on the information asymmetry of ADRs, controlling for the firm characteristics helps control for the firm level idiosyncrasy. Since firm characteristics is possibly endogenous to country-level characteristics, adding those controls will bias against finding a home market effect.

Table 2.2 reports the summary statistics of the firm-level control variables, along with the means of the firm-level control variables for sub-samples of firms according to whether home market disclosure and governance measures are above or below the medians, or whether firms are from emerging or developed markets. It reports the p-value of T-test of difference in means in parentheses. On average, ADRs from emerging markets, compared with those from developed markets, are slightly smaller (but the difference is not significant), have significantly higher proportion of tangible assets, and have significantly higher leverage ratio and lower MTB. ADRs from more transparent markets and markets of better investor rights protection, as well as ADRs from countries of common law origin, are noticeably smaller in size, have significantly less tangibility, and have comparatively higher MTB ratio. Except for comparison across ANTIDIR subgroups, they also appear to have comparatively lower level of leverages, although the differences are not significant. These features in the data are consistent with the hypothesis that higher disclosure level and better corporate governance of home markets allow ADR issuers to keep a lower level of tangible assets. Relatively lower level of leverage and higher market-to-book ratio of those firms indicate comparatively smaller degree of information asymmetry.

Many authors have used analyst-following data to proxy for the firm-level information environment (e.g., Lang et al. (2003)). We include two variables calculated from IBES data to make for additional firm-level controls. $NUMANALYS$ is measured as the number of analysts following in the annual consensus IBES forecast. $DISPERSION$ is the dispersion of analyst EPS forecasts measured as standard deviation of analysts forecasts in months of calendar year prior to the end of fiscal year, and normalized by absolute value of realized EPS and square root of the number of analysts (following Jin and Myers (2006)). While higher forecast dispersion is generally believed to be associated with higher degree of information asymmetry,[4] the relation between analyst following and information asymmetry is less conclusive in the literature. A priori, we do not have expectations of how $NUMANALYS$ or $DISPERSION$ is related to adverse selection, and control for it only to the extent that it may affect an ADR firm's information environment.

We also include U.S. and local market volatility, measured as standard deviation of daily returns calculated from market indices (in percentages), to control for time-specific variations.

[4]When information production is possible, more potential private information can be discovered and traded upon when there is greater uncertainty regarding future earnings. However, Jiang, Lee, and Zhang (2005) show that the increased uncertainty makes it more costly to discover and profit from private information.

2.3 Empirical Analysis

We are interested in whether the home market information and governance environment continue to be relevant factors in asymmetric information of ADR firms that are subject to SEC regulation and U.S. GAAP reporting. We estimate the following regression model:

$$AS_{ijt} = \gamma_0 + \gamma_1 HME_j + \gamma_2 X_{jt} + \gamma_3 X_{ijt} + \upsilon_{ijt}, \tag{2.3}$$

where AS_{ijt} is the adverse selection measure of firm i from country j in year t, HME_j is the institutional variables of market j where firm i is domiciled, X_{ijt} is a set of firm-level control variables, and X_{jt} is a set of market-level control variables that include volatility of U.S. and home market. All panel test statistics are adjusted for heteroskedasticity and autocorrelation with robust errors. We also check our estimation with cluster regression errors, which yield very similar results.[5]

We pay special attention to information disclosure environment, since intense and timely financial disclosures, including interpretation and coverage by analysts and the media, are crucial for inside managers to have credible ways of conveying hidden firm-specific information to outside investors.

2.3.1 Average Quoted Spreads

This section discusses the results of regressions equation 2.3 with the average quoted spread, QS, as the dependent variable. It provides us with the first evidence to our investigation.

Table 2.3 reports the estimation results of our baseline regression. Examining Table 2.3, we find that both $GCRSCORE$ and $CIFAR$ are negatively correlated with the quoted spread, with similar magnitudes (the mean of $GCRSCORE$ is 5.4 and the

[5] Cluster regression does not change coefficient estimates.

mean of $CIFAR$ is 71), indicating a higher spread for ADRs from less transparent markets. The coefficient estimate of $GCRSCORE$ is -0.06 with a p-value of 0.11. The coefficient estimate of $CIFAR$ is -0.01 with a p-value of 0.02. The significance level of $GCRSCORE$ is relatively weak. This is partly due to the fact that variations of $GCRSCORE$ are smaller relative to variations in spreads throughout the years. It is also explained by the fact, which will transpire more clearly in our later sub-sample and two-stage analysis, that developed markets are less well-characterized by survey measures. The coefficient estimate of $ANTIDIR$ is -0.12. The $CIVIL$ dummy shows a coefficient of 0.42. Both are highly significant. These estimates suggest that firms with better protection of investor rights, measured either by the anti-director rights or by common-law legal origin versus civil-law, on average have lower spreads. The result that the legal and governance environment can affect an ADR firm's information asymmetry is somewhat counterintuitive at the first brush. But recall our discussion that weak investor protection can reduce a firm's incentive to voluntarily disclose, can negatively affect analyst covering of the firm, and can decrease investors' desire to be informed about the firm, then the result isn't all that surprising.

On control variables, as expected, spreads are negatively correlated with firm size and MTB, which are often used as proxies of information asymmetry. Spreads are also positively and significantly correlated with U.S. market volatility. More volatile local markets lead to higher spreads, but the estimates are not significant. Leverage is positively correlated with spread, consistent with the suggestion that higher degree of information asymmetry leads to higher leverage. One surprise is that tangibility is positively correlated with the quoted spreads, since higher tangibility often suggests a lower information asymmetry. We speculate that firms with higher adverse selection, including those from less transparent home markets, tend to elect to have a higher component of tangible assets.

Table 2.4 focuses on the disclosure variable $GCRSCORE$, and further examines the robustness of our baseline results of the quoted-spread regression with various combinations of controls. Columns (1) and (2) estimate regressions without market volatilities and tangibility, respectively. They show that the exclusion of TANG does not affect the significance of our results. Column (3) includes an emerging market dummy variable, EM, to the regression. It shows that $GCRSCORE$ absorbs most of the emerging market factor, but the significance level decreases. Column (4) includes industry fixed effect;[6] it shows that the industry fixed effect reduces home country disclosure factor but the estimate on $GCRSCORE$ is still negative.[7] Columns (6), (7), and (8) add $CIVIL$, $ANTIDIR$, and both to the list of regressors. The coefficient estimates of $GCRSCORE$ remain negative and significant and their magnitude actually increases, indicating that disclosure, although positively correlated with the governance and the civil-law dummy, is a distinctive factor that affects spreads. The estimates on $CIVIL$ and $ANTIDIR$ remain consistent with those of Table 2.3 and are significant. Column (9) shows that home country disclosure effect is robust to controlling of the analyst opinion dispersion and the analysts coverage. The coefficient estimate on $DISPERSION$ is negative and on $NUMANALYS$ is positive, consistent with the argument that analysts dispersion of opinion is an indicator of information asymmetry and the more analysts coverage reduces information asymmetry.

Several authors have argued that ADR firms from markets of different maturity behave differently when being integrated into the new information environment. Ferreira and Fernandes (2008) find that cross-listing has an asymmetric impact on stock price informativeness for emerging-market and developed-market firms. It is therefore of interest to estimate our baseline regressions separately for the Emerging

[6] Industry classification according to 2-digit SIC

[7] Our Fama-Mcbeth estimates show that significance of $GCRSCORE$ is robust to industry fixed effect

Market (EM) and the Developed Market (DEV) firms. Table 2.5 presents the results. For the EM sub-sample, the coefficient estimate on $GCRSCORE$ is slightly bigger (more negative, -0.08 compared to -0.06) than the whole sample, although its significance is a reduced p-value of 0.19, possibly due to the reduced sample size and variation in $GCRSCORE$; the coefficient estimate on $CIFAR$ is negative but not significant. While for the DEV sub-sample, $GCRSCORE$ is much smaller (less negative, -0.01 compares to -0.06) than the whole sample, and insignificant, but the coefficient estimate on $CIFAR$ remains negative and significant. It suggests that survey-based disclosure measure is not as informative an indicator of market transparency for the DEV markets as it is for the EM markets. Conceptually, it is harder to distinguish matured markets by survey respondents. Accounting-based "hard" measures are more appropriate for those markets. While for the EM markets, the accounting-based "hard" measure is not enough to explain the level of market transparency, probably due to the lack of enforcement of the existing rules. The survey-based measure does a better job. Note also for our data, $GCRSCORE$ covers more EM country than $CIFAR$. After examining estimates on $ANTIDIR$ and $CIVIL$, we find that the governance and the legal origin effects are much larger for the EM firms than for the DEV firms. $ANTIDIR$ and $CIVIL$ coefficients are -0.19 and 0.7, respectively, for the EM sub-sample, comparing to -0.02 and 0.2 respectively for the DEV sub-sample. Both $ANTIDIR$ and $CIVIL$ estimates are significant for the EM firms, but $ANTIDIR$ is not significant for the DEV firms, suggesting that corporate governance is less of a concern for the DEV firms than the EM firms. Taken as a whole, results in our sub-sample tests are consistent with the hypothesis that EM ADRs are more subjectable to the home capital market environment than DEV ADRs.

Table 2.5 also shows an interesting difference between the EM and the DEV sub-samples: Estimates on the local market volatility are bigger and significant for the EM firms, and smaller and not significant for the DEV firms. EM ADRs, whose adverse selection costs are more dependent on their home capital market environment, are more affected by their local stock market volatilities. This is especially interesting since many of DEV ADRs have the bulk of their daily trading volume executed in their home markets (Aggarwal, Dahiya, and Klapper (2007)),[8] while EM ADRs tend to have a larger proportion of volume executed in the U.S. exchanges. The fact that EM ADRs have a larger portion of trading in U.S. exchanges makes our finding all the more remarkable. It suggests that the home market effect for emerging market firms is quite strong.

2.3.2 A Time-Varying Measure of Information Environment

One drawback of our disclosure proxies is that they are time invariant. While evidence suggests a country's information disclosure environment and underlying institutional structure are quite stable, we nevertheless try to check the robustness of our results to time-varying degree of the home market information opaqueness with the R^2 measure advocated by Morck et al. (2000) and Jin and Myers (2006).

R_i^2 in a market factor regression can be used as a measure of the extent to which firm-specific information determines stock price movements, because R_i^2 measures precisely the percentage of firm-specific idiosyncratic volatility to total volatility. A high R_i^2 thus indicates a high degree of stock price co-movement and a low observable idiosyncratic risk. Morck et al. (2000) find that stocks have higher R^2s in countries with less-developed financial markets. Jin and Myers (2006) confirm that R^2 is correlated with a country's financial opaqueness. Ferreira and Fernandes (2008) use R^2-based measure to show that emerging-market firms that cross-list in the U.S. achieved less improvement in information environment than their developed-market counterparts.

[8] Lower trading cost, better legal standing due to additional protection from U.S. security law (Aggarwal et al. (2007)) are the possible reasons.

We follow Jin and Myers (2006) to calculate R^2_{it} from the following regression:

$$\begin{aligned} r_{i\tau} &= \alpha_i + \beta_{1i} r_{m,j\tau} + \beta_{2i} r_{exus,j\tau} + \beta_{3i} r_{m,j\tau-1} + \beta_{4i} r_{exus,j\tau-1} + \beta_{5i} r_{m,j\tau-2} \\ &\quad + \beta_{6i} r_{exus,j\tau-2} + \beta_{7i} r_{m,j\tau+1} + \beta_{8i} r_{exus,j\tau+1} + \beta_{9i} r_{m,j\tau+2} + \beta_{10i} r_{exus,j\tau+2} + \epsilon_{i\tau}, \end{aligned} \tag{2.4}$$

for each firm i included in the DATASTREAM stock database of country j in year t, where $r_{i\tau}$ is the firm return, $r_{m,j\tau}$ is the local market return from DATASTREAM total return index of the country, $r_{exus,j\tau}$ is the U.S. market index return adjusted by exchange rate. The leads and lags in the regression are to correct for possible non-synchronous trading. We then average across firms in country j in year t to get our information measure R^2_{jt}, denoted RS, to substitute for disclosure measures in equation 2.3.[9] If quoted spread captures the degree of adverse selection, we expect RS to be positively correlated with QS.

Table 2.6 reports the results of the RS regressions. Through inspection of panel A, we find that RS measure confirms our finding in Table 2.3. RS coefficient estimate is 1.2 when no additional controls are added, and it is significant at the 5% level. After industry, governance, and legal origin are controlled for, respectively, estimates of RS coefficient remain positive. They are significant except when $CIVIL$ dummy is added. Panel B and Panel C compare estimation results for the emerging and the developed market sub-samples. Coefficient estimates of RS for the EM firms are all positive, and are in general much more pronounced than those of the whole sample. Coefficient of RS in the regression without additional control, for example, goes up to 2.18 and is highly significant. In contrast, three of four RS estimates for the DEV firms are negative. We do not find significantly positive correlation between the spreads and RS for the DEV sub-sample. This result is consistent with Ferreira and Fernandes' (2008) finding that DEV firms have improved their information en-

[9] Jin and Myers (2006) show that equally weighted averages perform as well as value-weighted averages.

vironment after cross-listing but EM firms do not. Therefore the EM firms continue to have their degree of adverse selection affected by the home market information environment while the DEV firms do not. We observe in Panel C, however, that the coefficient estimates of $CIVIL$ and $ANTIDIR$ for the DEV firms have the expected signs. That is, consistent with Table 2.5, the coefficient estimate of $CIVIL$ is positive and the coefficient estimate of $ANTIDIR$ is negative. $CIVIL$ stays significant for the DEV firms. Overall, our tests with RS as the home market information opaqueness measure provide supporting evidence of a positive relation between the adverse selection in ADR trading and the home market information opaqueness for the emerging market firms. For the developed market firms, effects of legal institution on adverse selection withstand, but we do not find significant effect of the home market information environment measured by idiosyncratic informativeness in stock prices.

2.3.3 Fama-Macbeth Statistical Method

In the above regression method, we utilize a panel approach by correcting for robust regression error, or adjusting for year and country fixed effects (cluster regression, not reported). If serial correlation in the error or heteroskedasticity is quite big, however, the above statistical methods might produce false significance. An alternative solution is to use a Fama and MacBeth (1973) two-stage procedure, which estimates a separate regression for each cross-section in each year and then takes the time series mean of the coefficients. The relative advantage of the Fama-MacBeth procedure is to minimize the serial correlated error and to maximize the cross-sectional variation in the home country information and institution qualities.

We find that the resulting estimates from the two-stage method largely confirm our earlier results. Table 2.7 reports an example as we reproduce the estimation of Table 2.3 and 2.4 using the two-stage regressing approach. Notice that the U.S. market volatility drops out of the cross sectional regressions. The resulting $GCRSCORE$ coefficient in Table 2.7 is -0.12 with a p-value of 0.08 when no additional controls are added. Its magnitudes increase to around -0.20 when $ANTIDIR$ or $CIVIL$ or both are added, as well as when analyst dispersion and coverage are controlled, with significance levels within 5%. Consistent with results in Table 2.4, magnitude and significance decrease when the industry fixed effects are added. Coefficient estimate on $ANTIDIR$ is negative, and on $CIVIL$ is positive. They are significant when entering regressions separately. Of the control variables, $SIZE$, DTA, and MTB have signs as we expected. One notable difference is that $TANG$ is no longer significant, indicating the existence of autocorrelation in the panel data. Overall, the Fama-MacBeth method confirms our earlier regression results.[10]

The Fama-MacBeth method provides with us an opportunity to examine economic significance of home market effect since it presents direct cross-section comparisons. For a coefficient estimate of -0.20 on $GCRSCORE$, average quoted spread decreases by 0.15 $(0.73 * (-0.2))$[11] if a firm improves its disclosure score by one standard deviation (0.73). This is economically significant, given that the average of these quoted spreads is at 1.71. The hypothetical move almost reduces the average spread by 10%, all else equal. Similarly, difference between a firm from common-law-legal-origin country and civil-law-legal-origin country is 0.2.

Dividing the sample according to emerging and developed markets offers some contrast between the two (not reported here). The resulting estimates for EM firms are consistent with the estimates of the full sample, but $GCRSCORE$ coefficient is estimated positive for DEV firms (CIFAR remains negative but insignificant). $CIVIL$ estimate again remains positive. The effect of the home market information opaqueness is mostly driven by the EM firms; the DEV firms are more affected by their legal

[10]Sub-sample analysis of the Fama-MacBeth approach also confirms our earlier findings. These results are not reported and are available upon request.

[11]Bid-ask spreads are in percentages.

institutional differences.

2.3.4 Effective Spreads and Adverse Selection Component of Spreads

The quoted bid-ask spread is a raw measurement of trading cost and has its own concerns. The quoted bid-ask spread may contain measurement problems, or actual trading may take place mostly within the quotes. Roll (1984) derives a measure of *effective spread* based on negative autocovariance of security returns, under the assumption that market makers face only order processing costs. Moreover, Stoll (1989) and others show that spreads could be decomposed of a non-information component, such as order processing and inventory holding cost, and an asymmetric information component of adverse selection cost. In our context, weak home market institution and information opaqueness can possibly lead to an increase of all three components of bid-ask spreads. Higher (unit) processing cost can be a result of less frequent trading, partly due to less transparent information; higher inventory holding cost also can result from weak institutional protection; higher trading costs of ADRs from certain countries can be a result of the lack of competition among exchanges, which in turn reflects higher local trading costs due to institutional and informational problems. Therefore, we are interested in the effect of the home market environment on both ADRs' effective spreads and the adverse-selection component of these spreads.[12]

We estimate Roll (1984)'s effective spreads and information component of spreads with methods proposed by George et al. (1991). We estimate the former utilizing the difference between returns based on transaction prices and returns calculated using bid-to-bid prices. In the latter case, the adverse selection component of spread is extracted by subtracting the non-information component calculated with auto-covariance of expected return filtered return series from the yearly average quoted

[12]We estimate information component of spreads, not proportional information component, because with big enough spread, information component of a stock can still be relatively high even with a low proportional measure.

spread. Specifically, we calculate the non-information component by estimating $\hat{S}_i = 200\sqrt{-cov(\eta_\tau, \eta_{\tau-1})}$, over a 60-day rolling sample in year t, where η_τ is expectation-filtered transaction return (see George et al. (1991) and Bharath et al. (2008) for details); we then take the average to get the yearly estimate. We are interested in seeing whether effective spreads and/or the information component of spreads are correlated with our disclosure and institutional quality measures, and whether $GCRSCORE$ would remain a significant factor after controlling for other factors.

Table 2.8 and Table 2.9 report the results of regressions in which effective spread and adverse selection component of the spread are used as dependent variables, respectively. Upon examining the tables, we find a consistent pattern of difference between the EM and the DEV firms. The emerging market firms are more sensitive to disclosure, measured either by $GCRSCORE$ or $CIFAR$, and they are also sensitive to governance measures; while the developed market firms are only sensitive to legal origin measures. In fact, the coefficient of $GCRSCORE$ for the developed market firms in the effective spread regression appears to have opposite significance.

In the effective spread regressions, emerging market ADRs' effective spreads are negatively correlated with disclosure, measured either by $GCRSCORE$ or $CIFAR$, and protection of shareholder rights, measured by $ANTIDIR$. Slope estimates on $GCRSCORE$ ranges from -0.10 without additional control to -0.39 when legal origin is controlled. $GCRSCORE$ remains a significant factor after controlling for $CIVIL$, $ANTIDIR$, and both, and after controlling for IBES forecast dispersion and analysts coverage. Slope estimate of $ANTIDIR$ is -0.13 and is highly significant. $CIVIL$, the legal origin variable, however, is not significant, and appears to have negative signs when augmented as a control variable. For the DEV firms, on the other hand, local market opaqueness does not appear to be a significant factor that increases effective spreads, nor does shareholder protection. $GCRSCORE$, in fact, appears to be positively correlated with effective spreads.[13] However, $CIVIL$ is positively and significantly correlated with the effective spread, as expected, indicating that rule of law based on different legal origins is an important home market factor for firms from developed markets. As we find in quoted spreads, effective spreads are more correlated with domestic market volatility for EM firms. These results are consistent with the hypothesis that emerging market ADRs have higher effective trading cost due to local market conditions.

In the adverse-selection component regression results, reported in Table 2.9, the EM firms with better local market financial disclosure have lower spreads attributed to information asymmetry. Slope estimates on $GCRSCORE$ ranges from -0.10 to -0.29. $GCRSCORE$ is significant at 10% level without additional controls, and significant at 5% level when industry and IBES controls are added. It is highly significant when legal origin is controlled. $CIFAR$ has a slope estimate of -0.01, although the p-value is 0.13 and not significant. $CIVIL$ and $ANTIDIR$, in regressions where they are used separately as the main institutional regressors respectively, are significant and have signs as expected. For the DEV firms, $CIFAR$ is negatively and significantly correlated with adverse selection component of spreads. $GCRSCORE$ estimates are negative except when $CIVIL$ is augmented as control variable, but they are not significant. Therefore, unlike results in the effective spread regressions, adverse-selection component of spreads of the DEV firms widens with more opaque home disclosure, albeit less significantly comparing to that of the EM firms. Coefficient estimates of $CIVIL$ are consistently positive and significant across specifications, again indicating legal origin of laws is an important factor in determining trading cost for DEV firms. $ANTIDIR$ has a slope estimate of -0.65 when estimated as the lone institutional

[13] The reason behind a positive relation between effective spreads and $GCRSCORE$ among DEV firms may be due to ADRs from more transparent markets being more actively traded and thus have higher order processing cost and more volatile returns, which both contribute to higher effective spread. When the information asymmetry component of spread is less affected by home market information environment for DEV ADR firms, effects of other components of the bid-ask spread are likely to dominate.

variable, but is not significant, and is otherwise positive. In contrast to results in the effective spread regression, estimates of institutional quality variables for the whole sample all have the expected signs and are significant. This is most likely due to DEV firms' responses of adverse selection component of spreads to disclosure and governance variables are more "well behaved" than that of effective spreads.

Overall, we conclude that better local legal institution qualities and more transparent information environment reduce adverse selection component of trading costs. This effect is more pronounced for emerging market firms, which display significant effects in both effective trading cost and adverse selection component of spreads than for developed market firms. For firms from developed markets, local information disclosure and corporate governance measured by anti-director rights do not appear to adversely affect effective spreads, instead the rule of law based on different legal origins appears to be the dominant institutional factor in affecting trading costs. The adverse selection component of spreads of developed market firms are weakly correlated with domestic institution qualities, but the non-information component is not. Both information and non-information component of spreads of emerging market firms are correlated with domestic institution quality. As a result, the overall sample exhibits correlation between information component of spreads and home quality of institutions.

2.3.5 Alternative Information Measure

So far, we have presented empirical evidence that information-based trading is positively correlated with a more opaque domestic information environment and a poorer institutional quality. More specifically, bid-ask spreads, and their information component, appear to be higher for ADRs from markets of lower disclosure, poorer investor protection, and of civil law legal origin, especially for emerging market firms. To substantiate our interpretation of the relation between information asymmetry of

ADR firms and the home country information and institutional quality, we test for the relation with an alternative adverse selection measure that is not spread-based as a dependent variable.

Llorente at al. (2002) develop an intensity-of-informed-trading measure that accounts for the trading volume accompanying stock return autocorrelations. We follow their method and construct our alternative proxy for each firm-year with resulting coefficient estimates of C_{2it}, denoted C2, from equation 2.2 for all NYSE and AMEX listed firms in our sample.[14] [15] Within each year, higher values of C2 tend to indicate more information-based trading and a higher degree of adverse selection.

Results of regressions using C2 (scaled by 100) as dependent variable are reported in Table 2.10. To minimize the error driven by market factor, we estimate with the two-stage regression approach and report the means of the estimates and p-values of the related T-tests. The results for this alternative measure of stock price informativeness support the findings that domestic information environment and institutional qualities matter primarily for emerging market ADRs. The coefficient estimates of $GCRSCORE$ and $ANTIDIR$ for EM firms are negative and significant, while for DEV firms they are positive. We do not find support that information asymmetry is related to legal law of origin with the C2 measure. However, caution should be taken against over-interpreting the C2 measure results, since the goodness-of-fit is relatively poor. R^2 of cross section regressions are at about 5% (consistent with Llorente at al. (2002)).

2.3.6 Is the Home Country Institution Risk Priced?

With the evidence presented above, that domestic information environment and institutional qualities continue to affect adverse selection costs of ADRs, especially

[14]Volume data for NASDAQ firms contains information from after-hour trading and therefore is not as reliable.

[15]We also eliminate firms that have fewer than 60 available observations in a year.

for emerging market firms, it is natural to ask the question: Is the risk posed by home country information opaqueness and inadequate institutional quality priced? Previous research finds evidence that poor shareholder protection (LLSV (2002)) and weak legal institutions (Lombardo and Pagano (1999)) are penalized with lower valuations. We've already seen that poorer home country institutions are associated with higher capital costs in terms of higher bid-ask spreads of ADRs. This section presents some preliminary evidence on the valuation effect of the home country institution risk.

In our context, our task in answering that question is a lot more difficult, because ADR returns also reflect country factors other than the home country capital market environment. It is difficult to differentiate the information factor from other country risk factors in our context. Further, occasional bursts of international financial crisis in emerging markets will likely distort our analysis. Nevertheless, to provide some clue to that question, we analyze monthly excess ADR returns. We use returns after 2002 to avoid complication by the financial crisis. We first consider a hedge portfolio similar to Aboody, Hughes, and Liu (2004). More specifically, we extend the Fama-French three-factor model by using a hedge portfolio based on home capital market environment factor and substituting a world market return for market return. We estimate for each firm-year:

$$r_{p\tau} - r_{f\tau} = \alpha + \beta_1(r_{wd\tau} - r_{f\tau}) + +\beta_2 SMB_\tau + \beta_3 HML_\tau + \zeta_\tau, \tag{2.5}$$

where $r_{p\tau}$ is the (equally weighted) monthly hedge portfolio return of going long in ADRs from markets where $GCRSCORE$ are below the sample median and going short in ADRs from markets where $GCRSCORE$ are above the sample median. $r_{f\tau}$ is the risk free rate, $r_{wd\tau}$ is the U.S. dollar return of world market index; SMB_τ and HML_τ are, respectively, the Fama and French (1992) size and market-to-book factor returns. We are interested in the estimate of α. A positive estimation of α suggests that lower information quality needs to be compensated with higher expected returns. Regression with monthly returns from 2002 to 2006 gives us an estimation of α of 0.67 with a T-stat of 1.5.

To investigate the question in more depth, we next adopt an approach similar to Easley, Hvidkjaer, and OHara (2002). We first estimate factor loadings with returns of 36 months prior to the month. For each firm i at month t we estimate:

$$r_\tau - r_{f\tau} = \alpha + \omega_{wd}(r_{wd\tau} - r_{f\tau}) + \omega_{smb}SMB_\tau + \omega_{hml}HML_\tau + \zeta_\tau. \tag{2.6}$$

We then estimate cross-sectionally (firm subscript is omitted):

$$r_t - r_{ft} = \lambda_{0,t} + \lambda_{1,t}\hat{\omega}_{wd,t} + \lambda_{2,t}\hat{\omega}_{smb,t} + \lambda_{3,t}\hat{\omega}_{hml,t} + \lambda_{4,t}HCME + \varrho_t, \tag{2.7}$$

where $HCME$ are our proxies for home capital market environment. We expect negative estimate of λ_4 on $GCRSCORE$, $CIFAR$, and $ANTIDIR$ if lower information quality requires higher cost of capital. Table 2.11 reports the mean and P-value of T-test of average λ_4 loading from 2002 to 2006.

The resulting estimates show $GCRSCORE$, $CIFAR$, and $ANTIDIR$ are negatively correlated with the risk-adjusted excess return, and $CIVIL$ is positively correlated with the risk-adjusted abnormal return. Lower information transparency, poorer protection of investor rights, and civil law of origin are associated with a higher cost of capital. Therefore, it suggests that risks due to home market information opaqueness and lack of investor protection are taken into account by the market. The significance of coefficient estimates in Table 2.11, however, is relatively weak, with only $GCRSCORE$ being significant at the 10% level. Nevertheless, it provides us with some idea of the answer to the question. As we cautioned above, these excess return analyses should be viewed with extra caution, since factors independent of the home capital market environment effect are difficult to extract. Further research is warranted to draw a more conclusive answer to the question.

2.4 Conclusion

Cross-listing has increasingly become an important form of raising capital abroad. U.S. investors hold ADRs as an alternative way to seek the benefit of international diversification (Ammer et al. (2008)). Previous research on cross-listing finds evidence in support of the bonding theory, which asserts that foreign firms can improve their disclosure and governance by cross-listing in the United States in order to bond themselves to the U.S. security regimes. The degree of effectiveness of such bonding, however, remains an open question. In this chapter, we contribute to the bonding literature by investigating whether the home capital market environment such as information disclosure and the protection of investor rights remains to have effect on the degree of information asymmetry of ADRs cross-listed on the U.S. exchanges. Our evidence supports an affirmative answer to that question, especially for emerging market firms.

We show that both the average quoted bid-ask spread of ADRs and the informational component of the spread are positively correlated with information opaqueness and poor protection of investor rights in the capital market environment of the home countries, and this relation is more pronounced among emerging market ADRs. In other words, the home stigma exists among ADRs traded on the U.S. exchanges. This home stigma is robust to different measures of home market information environment and different statistical estimation methods. Our evidence with spread-based measure implies that adverse selection cost in ADRs is significantly higher for ADRs from less transparent and poorly governed capital markets, and our alternative non-spread-based measure confirms that the pattern reflects an information problem. In addition, we provide evidence that this "home stigma" risk appears to be priced by the market. Cross-sectionally, the difference is significant between firms from emerging markets and firms from more mature capital markets. Among emerging market ADRs, the connection between the home capital market environment and their adverse selection in the U.S. trading is much stronger than their developed market counterparts. Indeed, the "home stigma" problem of emerging market ADRs is so prominent that it also drives their effective spread and overall trading costs. Information opacity of emerging markets does not appear to be solely due to the lack of regulation itself. In fact, the survey-based transparency measure fits them better than the accounting-rule-based measure, suggesting lack of enforcement of the existing rules. Among developed market ADRs, the home stigma phenomenon is weaker, but there is evidence that home market accounting standards and the quality of legal institutions continue to affect their level of adverse section in the U.S. trading.

Timely and intense financial disclosures, including interpretation and coverage by analysts and the media, are crucial for inside managers to have credible ways of conveying hidden firm-specific information to outside investors. Although the regulatory requirements of U.S. listing provide cross-listed firms with an opportunity to bond with the U.S. market regime, much of the disclosure practice is voluntary and may be subject to an individual firm's costs and benefits. Home capital market environments can affect a firm's incentive of adhering to a higher disclosure standard, and therefore affect the level of adverse selection in their U.S.-listed securities. Our results suggest that this is especially a concern for emerging market firms. Given the fact that bonding to the U.S. security regime is often cited as one of the reasons that firms cross-list in the U.S., our finding of home stigma appears a bit puzzling. However, in a recent study of deregister decision by cross-listed foreign firms following the SEC's adoption of Rule 12h-6 in March 2007, Doidge, Karolyi, and Stulz (2008) find evidence that foreign firms are more likely to deregister once the benefit of raising low-cost capital disappears. Their finding suggests another possible channel that home stigma may persist: Once foreign firms are listed in the U.S. and the prospect of requiring new capital diminishes, home capital market characteristics may creep back into ADR firms. This is an interesting aspect of international corporate governance that warrants further research.

Our results have important policy implications. Our results illustrate that cross-listing in the United States is not a substitute for improvement of local information transparency and protection of minority shareholders. Governments and policymakers should still focus on promoting information disclosure and improving corporate governance in order to attract capital flows to their countries. Our finding that quality of governance matters for the information environment also suggests that policies aiming to improve domestic disclosure should be complemented with policies promoting better investor rights protection.

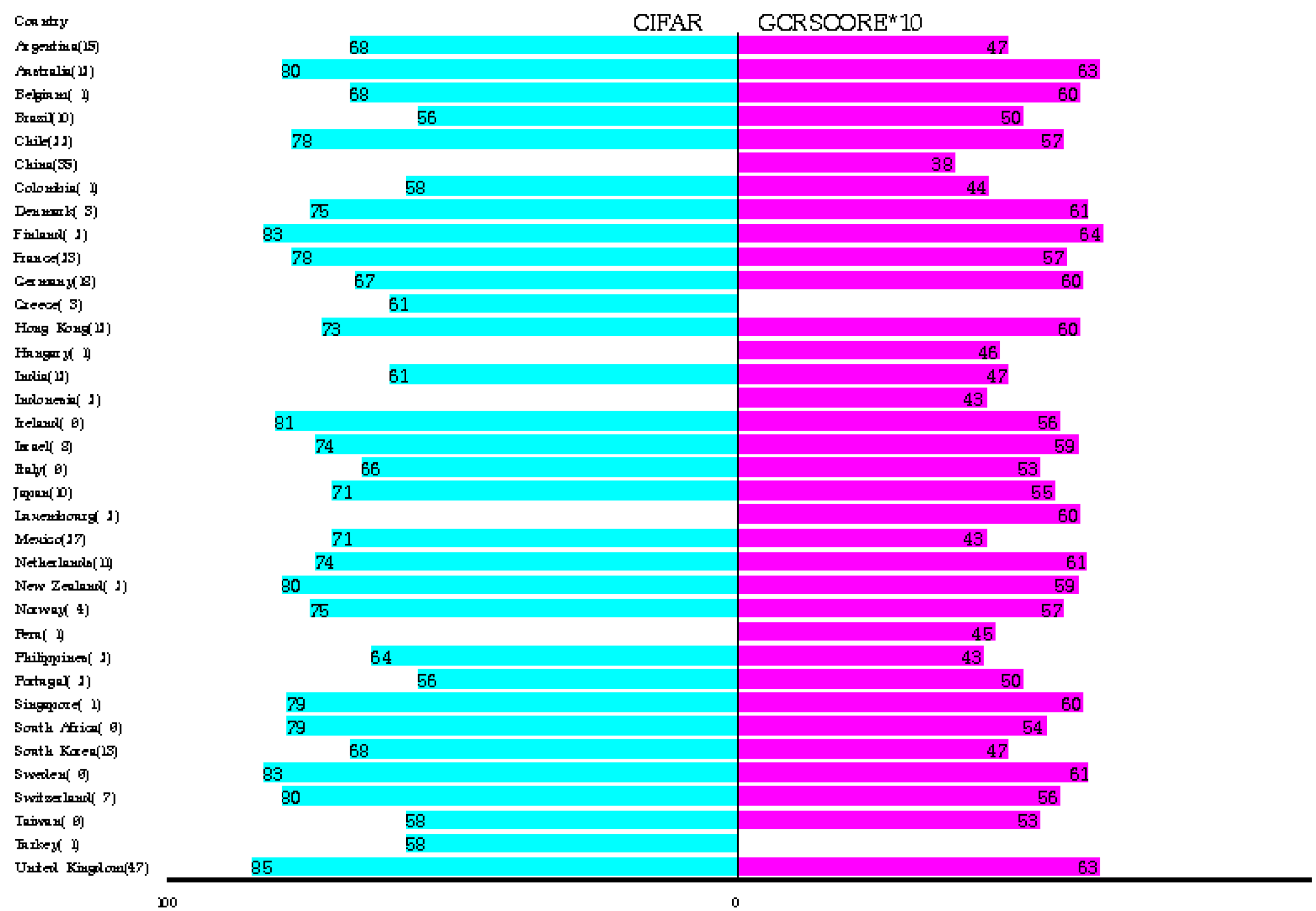

Figure 2.1: Home Market Information Transparency Measures. Left-side bars represent CIFAR scores, and right-side bars represent GCRSCOREs, scale up by 10.

Table 2.1: Correlation Between the Home Capital Market Environment Measures

This table reports the Spearman correlation coefficient of institution quality proxies. Summary statistics of each proxy are also reported. GCRSCORE is a survey-based measure from Global Competitive Report, and proxies for information transparency. CIFAR is an accounting-standard-based measure of disclosure. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin.

	GCRSCORE	CIFAR	ANTIDIR	CIVIL
GCRSCORE	1.00	0.69	0.13	-0.31
CIFAR	0.69	1.00	0.36	-0.43
ANTIDIR	0.13	0.36	1.00	-0.64
CIVIL	-0.31	-0.43	-0.64	1.00
Mean	5.41	71.23	3.03	0.73
Std	0.73	8.91	1.31	0.45
Median	5.65	73.00	3.00	1.00
N	34	31	33	33

Table 2.2: Summary Statistics on ADRs

This table reports the summary statistics for our sample of level II and III ADRs. Means of quoted spread, log size, tangibility, DTA, and MTB of firms from different sub-sample of markets are reported, as well as the means and standard deviations of the whole sample. p-value of T-test of difference in means are reported in parentheses. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio.

	Quoted Spread	SIZE	TANG	DTA	MTB
$GCRSCORE \geq Median$	1.72	7.70	0.32	0.26	1.93
$GCRSCORE < Median$	1.75	7.95	0.46	0.27	1.70
T-test Difference (p-value)	(0.52)	(0.01)	(0.00)	(0.18)	(0.30)
$CIFAR \geq Median$	1.70	7.63	0.34	0.26	1.95
$CIFAR < Median$	1.86	8.05	0.40	0.27	1.68
T-test Difference (p-value)	(0.03)	(0.00)	(0.00)	(0.17)	(0.20)
$ANTIDIR \geq Median$	1.70	7.70	0.38	0.27	2.02
$ANTIDIR < Median$	1.79	8.03	0.40	0.26	1.49
T-test Difference (p-value)	(0.00)	(0.00)	(0.01)	(0.22)	(0.01)
COMMON	1.63	7.28	0.34	0.25	2.68
CIVIL	1.83	8.04	0.38	0.27	1.46
T-test Difference (p-value)	(0.01)	(0.00)	(0.01)	(0.19)	(0.00)
Emerging Market Firms	1.80	7.81	0.49	0.27	1.62
Developed Market Firms	1.67	7.84	0.29	0.25	2.01
T-test Difference (p-value)	(0.05)	(0.65)	(0.00)	(0.03)	(0.02)
Sample Mean	1.71	7.82	0.39	0.26	1.82
Sample Standard Deviation	1.47	1.73	0.27	0.2	3.87
Number of Observation	1909	1909	1905	1909	1879

Table 2.3: Quoted Bid-Ask Spread and Home Capital Market Environment

This table reports the estimates of average quoted spread on institution quality measures of home capital markets. GCRSCORE is a survey-based measure from Global Competitive Report, and proxies for information transparency. CIFAR is a accounting-standard-based measure of disclosure. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio. *P*-values are calculated with White robust errors and reported in parentheses. Error Degree of Freedom and adjusted R-squares are also reported at the bottom. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	GCRSCORE	**CIFAR**	**ANTIDIR**	**CIVIL**
Intercept	2.72^c	3.28^c	3.05^c	2.44^c
	(0.00)	(0.00)	(0.00)	(0.00)
Disclosure and Governance	-0.06	-0.01^b	-0.12^c	0.42^c
	(0.11)	(0.02)	(0.00)	(0.00)
U.S. Market Volatility	1.37^c	1.23^c	1.29^c	1.30^c
	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.01	0.11^c	0.004	0.005
	(0.22)	(0.00)	(0.29)	(0.39)
SIZE	-0.33^c	-0.35^c	-0.35^c	-0.36^c
	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.51^c	0.42^c	0.60^c	0.50^c
	(0.00)	(0.00)	(0.00)	(0.00)
DTA	1.27^c	1.24^c	1.26^c	1.25^c
	(0.00)	(0.00)	(0.00)	(0.00)
MTB	-0.05^c	-0.05^c	-0.04^c	-0.04^b
	(0.01)	(0.00)	(0.01)	(0.02)
EDF	1796	1607	1633	1633
Adj. R-square	0.29	0.29	0.29	0.29

Table 2.4: Quoted Bid-Ask Spread and Home Market Information Transparency

This table reports the estimates of average quoted spreads on disclosure measure and various controls. P-values are calculated with White robust errors and reported in parentheses. Error Degree of Freedom and adjusted R-squares are also reported at the bottom. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	(1)	**(2)**	**(3)**	**(4)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10)**
Intercept	4.35^c	3.04^c	2.82^c	3.25^c	3.34^c	3.87^c	3.41^c	2.57^c	2.75^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GCRSCORE	-0.07	-0.11^c	-0.08	-0.03	-0.14^c	-0.16^c	-0.11^b	-0.13^c	-0.11^a
	(0.11)	(0.01)	(0.16)	(0.51)	(0.01)	(0.00)	(0.05)	(0.00)	(0.08)
U.S. Market Volatility		1.39^c	1.38^c	1.37^c	1.30^c	1.29^c	1.29^c	1.41^c	1.28^c
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility		0.009	0.01	0.01	0.002	0.002	0.002	0.008	0.003
		(0.15)	(0.26)	(0.21)	(0.68)	(0.80)	(0.75)	(0.24)	(0.63)
SIZE	-0.34^c	-0.32^c	-0.33^c	-0.37^c	-0.36^c	-0.36^c	-0.36^c	-0.29^c	-0.32^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG			0.53^c	0.67^c	0.48^c	0.55^c	0.56^c	0.82^c	0.76^c
			(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
DTA	1.64^c	1.44^c	1.27^c	1.13^c	1.29^c	1.30^c	1.30^c	1.19^c	1.18^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
MTB	-0.05^c	-0.05^b	-0.05^c	-0.04^b	-0.04^c	-0.04^c	-0.04^c	-0.05^b	-0.03^b
	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.03)	(0.05)
CIVIL					0.32^c		0.24^c		0.51^c
					(0.00)		(0.01)		(0.00)
ANTIDIR						-0.09^c	-0.06^b		-0.03
						(0.00)	(0.02)		(0.25)
DISPERSION								0.26^b	0.23^b
								(0.02)	(0.02)
NUMANALYS								-0.02^c	-0.02^c
								(0.00)	(0.00)
EM			-0.03						
			(0.70)						
Industry Fixed Effect				Yes					
EDF	1810	1800	1795	1788	1606	1606	1605	1272	1137
Adj. R-square	0.16	0.28	0.29	0.32	0.30	0.29	0.30	0.32	0.37

Table 2.5: Quoted Spread and Home Capital Market Environment: Emerging and Developed Markets

This table reports the estimates of quoted spread on institution quality measures of home capital market for sub-samples of emerging and developed markets. GCRSCORE is a survey-based measure from Global Competitive Report, and proxies for information transparency. CIFAR is an accounting-standard based measure of disclosure. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio. *P*-values are calculated with White robust errors, and reported in parentheses. Error Degree of Freedom and adjusted R-squares are also reported at the bottom. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	Emerging Market Firms				Developed Market Firms			
	GCRSCORE	CIFAR	ANTIDIR	CIVIL	GCRSCORE	CIFAR	ANTIDIR	CIVIL
Intercept	3.27^c	3.64^c	4.12^c	2.88^c	2.36^c	3.37^c	2.42^c	2.31^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Disclosure and Governance	-0.08	-0.003	-0.19^c	0.70^c	-0.01	-0.02^b	-0.02	0.20^b
	(0.19)	(0.59)	(0.00)	(0.00)	(0.92)	(0.02)	(0.41)	(0.02)
U.S. Market Volatility	1.42^c	1.30^c	1.29^c	1.29^c	1.22^c	1.09^c	1.22^c	1.22^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.14^c	0.12^c	0.080^c	0.120^c	0.006	0.20	0.005	0.004
	(0.00)	(0.00)	(0.00)	(0.00)	(0.43)	(0.22)	(0.40)	(0.53)
SIZE	-0.47^c	-0.48^c	-0.49^c	-0.48^c	-0.27^c	-0.28^c	-0.28^c	-0.30^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.67^c	0.23	0.44^b	0.27	0.32^a	0.34^a	0.35^a	0.44^b
	(0.00)	(0.24)	(0.03)	(0.17)	(0.09)	(0.07)	(0.06)	(0.02)
DTA	2.11^c	2.26^c	2.14^c	1.90^c	0.74^c	0.82^c	0.77^c	0.82^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
MTB	-0.04^a	-0.03^a	-0.03^a	-0.03^a	-0.06^b	-0.05^b	-0.05^b	-0.05^b
	(0.08)	(0.06)	(0.07)	(0.08)	(0.04)	(0.05)	(0.04)	(0.04)
EDF	861	709	720	720	935	890	905	905
Adj. R-square	0.36	0.34	0.37	0.36	0.24	0.26	0.25	0.26

Table 2.6: Quoted Spreads and Time-Varying Disclosure Measure

This table reports the estimates of average quoted spreads on time-variant disclosure measure and various controls. Panel A reports the regression results of the whole sample. Panel B reports the regression results of the emerging market firms sub-sample. Panel C reports the regression results of the developed market firms sub-sample. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio. P-values are calculated with White robust errors, and reported in parentheses. Error Degree of Freedom and R-squares are also reported at the bottom. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	Panel A: All Firms				Panel B: Emerging Market Firms				Panel C: Developed Market Firms			
Intercept	1.79^c	2.5^c	2.15^c	1.26^c	1.50^c	2.66^c	1.64^c	2.09^c	2.51^c	2.59^c	2.68^c	2.65^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
RS	1.20^b	0.90^a	0.55	1.15^a	2.18^c	0.92	2.12^c	2.30^c	-0.30	0.03	-0.90	-0.19
	(0.03)	(0.10)	(0.40)	(0.06)	(0.00)	(0.17)	(0.01)	(0.01)	(0.75)	(0.98)	(0.41)	(0.85)
U.S. Market Volatility	1.36^c	1.35^c	1.27^c	1.26^c	1.44^c	1.33^c	1.34^c	1.34^c	1.17^c	1.15^c	1.18^c	1.16^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.01	0.007	0.005	0.006	0.11^c	0.15^c	0.09^c	0.09^c	0.006	0.004	0.004	0.006
	(0.20)	(0.28)	(0.39)	(0.38)	(0.00)	(0.00)	(0.00)	(0.00)	(0.33)	(0.56)	(0.53)	(0.40)
SIZE	-0.30^c	-0.35^c	-0.33^c	-0.32^c	-0.39^c	-0.48^c	-0.42^c	-0.42^c	-0.28^c	-0.33^c	-0.31^c	-0.29^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.53^c	0.59^c	0.48^c	0.53^c	1.05^c	1.76^c	0.81^c	0.80^c	0.16	-0.07	0.31	0.21
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.39)	(0.79)	(0.12)	(0.29)
DTA	0.99^c	0.81^c	0.98^c	0.98^c	1.22^c	0.90^c	0.93^b	1.36^c	0.78^c	0.52^b	0.87^c	0.82^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.02)	(0.00)	(0.00)	(0.02)	(0.00)	(0.00)
MTB	-0.04^c	-0.03^b	-0.04^c	-0.04^c	-0.03^a	-0.01	-0.02^a	-0.03^a	-0.06^b	-0.05^b	-0.05^b	-0.05^b
	(0.01)	(0.02)	(0.01)	(0.01)	(0.09)	(0.16)	(0.10)	(0.07)	(0.04)	(0.04)	(0.04)	(0.04)
CIVIL			0.30^c				0.58^c				0.28^c	
			(0.00)				(0.00)				(0.01)	
ANTIDIR				-0.05^b				-0.03				-0.04
				(0.06)				(0.67)				(0.28)
Industry Fixed Effect		yes				yes				yes		
EDF	1568	1560	1387	1387	699	693	540	540	861	853	838	838
R-square	0.28	0.32	0.28	0.28	0.39	0.45	0.37	0.35	0.24	0.29	0.25	0.25

Table 2.7: Quoted Spread and Home Capital Market Environment: Two-Stage Regression Approach

This table reports the estimates of average quoted spreads on measures of home capital market environment using a two-stage regression approach. Separate regressions are estimated cross-sectionally in each year and the time series means of the coefficients are reported in the table. *P*-values referenced from a T-distribution are reported in parentheses. GCRSCORE is a survey-based measure from Global Competitive Report, and proxies for information transparency. CIFAR is an accounting-standard-based measure of disclosure. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio. NUMANALYS is measured as the number of analysts following of that specific year. DISPERSION is dispersion of analyst EPS forecast measured as standard deviation of analysts forecast in months of calendar year prior to the end of fiscal year, and normalized by absolute value of realized EPS and square root of the number of analysts. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Intercept	5.12^c	5.38^c	5.75^c	5.69^c	5.78^c	4.41^c	4.53^c	5.57^c	4.79^c	4.33^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GCRSCORE	-0.12^a	-0.09	-0.20^b	-0.19^b	-0.20^b	-0.20^c				-0.21^b
	(0.08)	(0.14)	(0.04)	(0.02)	(0.05)	(0.00)				(0.03)
Local Market Volatility	0.018	0.030	0.006	0.009	0.008	0.001	0.018	0.018	0.019	-0.012
	(0.33)	(0.18)	(0.70)	(0.60)	(0.63)	(0.97)	(0.36)	(0.75)	(0.36)	(0.42)
SIZE	-0.34^c	-0.35^c	-0.36^c	-0.35^c	-0.36^c	-0.22^c	-0.36^c	-0.35^c	-0.35^c	-0.23^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	-0.01	-0.02	0.04	0.05	0.04	0.26	0.06	0.02	0.14	0.27
	(0.97)	(0.91)	(0.79)	(0.72)	(0.78)	(0.16)	(0.65)	(0.91)	(0.29)	(0.19)
DTA	1.33^c	1.33^c	1.29^c	1.35^c	0.22^c	0.90^c	1.38^c	1.30^c	1.38^c	0.86^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
MTB	-0.24^c	-0.23^c	-0.24^c	-0.24^c	-0.24^c	-0.15^c	-0.26^c	-0.25^c	-0.25^c	-0.13^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CIVIL			0.03		0.01		0.20^b			0.32^c
			(0.84)		(0.94)		(0.03)			(0.01)
ANTIDIR				-0.03	0.01				-0.07^c	0.01
				(0.27)	(0.71)				(0.00)	(0.63)
DISPERSION						1.16^b				1.207^b
						(0.05)				(0.05)
NUMANALYS						-0.03^c				-0.03^c
						(0.00)				(0.00)
CIFAR								-0.01^c		
								(0.00)		
Industry Fixed Effect		yes								

Table 2.8: Effective Spread and Home Capital Market Environment

This table reports the estimates of effective spreads on measures of home capital market environment. Panel A reports the regression results of the whole sample. Panel B reports the regression results of the emerging market firms sub-sample. Panel C reports the regression results of the developed market firms sub-sample. GCRSCORE is a survey-based measure from Global Competitive Report, and proxies for information transparency. CIFAR is an accounting-standard-based measure of disclosure. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio. NUMANALYS is measured as the number of analysts following of that specific year. DISPERSION is dispersion of analyst EPS forecast measured as standard deviation of analysts forecast in months of calendar year prior to the end of fiscal year, and normalized by absolute value of realized EPS and square root of the number of analysts. *P*-values are calculated with White robust errors and reported in parentheses. Error Degree of Freedom and adjusted R-squares are also reported at the bottom. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**
Panel A: ALL Firms									
Intercept	2.18^c	2.83^c	2.88^c	2.46^c	2.88^c	1.87^c	2.92^c	2.30^c	2.58^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GCRSCORE	0.00	-0.08^a	-0.05	0.01	-0.05	-0.05			
	(0.99)	(0.07)	(0.22)	(0.68)	(0.24)	(0.13)			
CIFAR							-0.01^c		
							(0.01)		
U.S. Market Volatility	1.05^c	0.99^c	0.99^c	1.05^c	0.99^c	1.07^c	0.96^c	0.99^c	0.99^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.003	0.000	0.000	0.002	0.000	0.003	0.055^c	0.002	0.001
	(0.41)	(0.92)	(0.98)	(0.58)	(0.98)	(0.42)	(0.00)	(0.61)	(0.84)
SIZE	-0.26^c	-0.28^c	-0.28^c	-0.27^c	-0.28^c	-0.20^c	-0.28^c	-0.27^c	-0.28^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.20^b	0.15	0.20^b	0.14	0.20^b	0.32^c	0.09	0.15	0.20^b
	(0.03)	(0.12)	(0.04)	(0.24)	(0.05)	(0.00)	(0.36)	(0.12)	(0.04)
DTA	0.60^c	0.60^c	0.61^c	0.52^c	0.61^c	0.40^c	0.57^c	0.57^c	0.59^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)
MTB	-0.03^b	-0.03^b	-0.03^b	-0.03^b	-0.03^b	-0.03	-0.03^b	-0.03^b	-0.03^b
	(0.02)	(0.04)	(0.02)	(0.03)	(0.02)	(0.11)	(0.02)	(0.02)	(0.03)
CIVIL		0.07			0.00			0.13^c	
		(0.23)			(0.98)			(0.01)	
ANTIDIR			-0.05^c		-0.05^c				-0.06^c
			(0.01)		(0.01)				(0.00)
DISPERSION						0.21^b			
						(0.03)			
NUMANALYS						-0.01^c			
						(0.00)			
Industry Fixed Effect				yes					
EDF	1769	1577	1577	1761	1576	1192	1577	1604	1604
Adj. R-square	0.29	0.29	0.29	0.30	0.29	0.34	0.29	0.29	0.29
Panel B: Emerging Market Firms									
Intercept	2.57^c	4.58^c	4.00^c	3.17^c	3.90^c	2.52^c	3.65^c	2.34^c	2.90^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GCRSCORE	-0.10^b	-0.39^c	-0.23^b	-0.14^c	-0.15	-0.12^c			
	(0.02)	(0.00)	(0.04)	(0.00)	(0.24)	(0.01)			
CIFAR							-0.02^c		
							(0.00)		
U.S. Market Volatility	0.97^c	0.84^c	0.84^c	0.93^c	0.83^c	0.96^c	0.87^c	0.85^c	0.87^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.06^c	0.04^a	0.032	0.070^c	0.030	0.042^b	0.045^b	0.058^c	0.027

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Table 2.8 – continued from previous page

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	(0.00)	(0.06)	(0.11)	(0.00)	(0.13)	(0.03)	(0.03)	(0.01)	(0.20)
SIZE	-0.28^{c}	-0.33^{c}	-0.33^{c}	-0.31^{c}	-0.33^{c}	-0.25^{c}	-0.31^{c}	-0.29^{c}	-0.30^{c}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.22	0.24	0.24	0.48^{c}	0.24	0.25^{a}	-0.01	-0.04	0.10
	(0.13)	(0.12)	(0.12)	(0.02)	(0.13)	(0.07)	(0.94)	(0.79)	(0.52)
DTA	1.30^{c}	1.55^{c}	1.47^{c}	1.18^{c}	1.58^{c}	0.99^{c}	1.53^{c}	1.46^{c}	1.36^{c}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
MTB	-0.02^{a}	-0.02^{b}	-0.02^{a}	-0.01	-0.02^{a}	-0.07^{c}	-0.02^{c}	-0.02^{c}	-0.01^{c}
	(0.09)	(0.05)	(0.07)	(0.19)	(0.06)	(0.00)	(0.06)	(0.07)	(0.09)
CIVIL		-0.05			-0.22			0.02	
		(0.65)			(0.11)			(0.85)	
ANTIDIR			-0.07^{a}		-0.11^{b}				-0.13^{c}
			(0.09)		(0.03)				(0.00)
DISPERSION						0.12			
						(0.15)			
NUMANALYS						-0.01			
						(0.13)			
Industry Fixed Effect				yes					
EDF	858	687	687	850	686	595	703	714	714
Adj. R-square	0.30	0.32	0.33	0.33	0.33	0.37	0.28	0.27	0.30

Panel c: Developed Market Firms

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Intercept	1.36^{c}	0.49	1.56^{c}	1.52^{c}	0.33	0.90	2.39^{c}	2.33^{c}	2.23^{c}
	(0.01)	(0.43)	(0.00)	(0.02)	(0.60)	(0.15)	(0.00)	(0.00)	(0.00)
GCRSCORE	0.16^{a}	0.30^{c}	0.12	0.15^{a}	0.25^{c}	0.09			
	(0.06)	(0.00)	(0.21)	(0.08)	(0.02)	(0.38)			
CIFAR							0.00		
							(0.74)		
U.S. Market Volatility	1.05^{c}	1.05^{c}	1.05^{c}	1.07^{c}	1.05^{c}	1.13^{c}	0.91^{c}	1.06^{c}	1.05^{c}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.005	0.007	0.005	0.001	0.006	0.005	0.186	0.000	0.002
	(0.26)	(0.13)	(0.28)	(0.74)	(0.19)	(0.27)	(0.17)	(0.99)	(0.59)
SIZE	-0.26^{c}	-0.28^{c}	-0.26^{c}	-0.26^{c}	-0.29^{c}	-0.18^{c}	-0.26^{c}	-0.27^{c}	-0.26^{c}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.13	0.20	0.10	-0.30	0.25^{a}	0.31	0.14	0.19	0.11
	(0.36)	(0.19)	(0.50)	(0.13)	(0.10)	(0.09)	(0.35)	(0.19)	(0.45)
DTA	0.20	0.24^{a}	0.18	-0.03	0.23	-0.02	0.23^{a}	0.27	0.19
	(0.18)	(0.10)	(0.23)	(0.83)	(0.12)	(0.89)	(0.10)	(0.07)	(0.20)
MTB	-0.04^{a}	-0.04^{a}	-0.04^{a}	-0.05^{a}	-0.04^{a}	-0.02	-0.04	-0.04^{a}	-0.04^{a}
	(0.09)	(0.09)	(0.09)	(0.09)	(0.09)	(0.26)	(0.11)	(0.10)	(0.09)
CIVIL		0.23^{c}			0.50^{c}			0.12^{a}	
		(0.01)			(0.00)			(0.09)	
ANTIDIR			0.02		0.13^{c}				0.03
			(0.50)		(0.00)				(0.18)
DISPERSION						0.45^{b}			
						(0.02)			
NUMANALYS						-0.02^{c}			
						(0.00)			
Industry Fixed Effect				yes					
EDF	903	881	881	895	880	587	866	882	882
Adj. R-square	0.31	0.32	0.31	0.33	0.33	0.37	0.31	0.31	0.31

Table 2.9: Adverse-Selection Component of Spreads and Home Capital Market Environment

This table reports the estimates of adverse selection component of spreads on measures of home capital market environment. Panel A reports the regression results of the whole sample. Panel B reports the regression results of the emerging market firms sub-sample. Panel C reports the regression results of the developed market firms sub-sample. GCRSCORE is a survey-based measure from Global Competitive Report, and proxies for information transparency. CIFAR is an accounting-standard-based measure of disclosure. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio. NUMANALYS is measured as the number of analysts following of that specific year. DISPERSION is dispersion of analyst EPS forecast measured as standard deviation of analysts forecast in months of calendar year prior to the end of fiscal year, and normalized by absolute value of realized EPS and square root of the number of analysts. *P*-values are calculated with White robust errors and reported in parentheses. Error Degree of Freedom and adjusted R-squares are also reported at the bottom. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**
Panel A: ALL Firms									
Intercept	2.22^c	2.34^c	2.98^c	2.48^c	2.34^c	2.05^c	2.55^c	1.66^c	2.10^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GCRSCORE	-0.09^c	-0.11^b	-0.18^c	-0.06	-0.11^b	-0.13^c			
	(0.01)	(0.03)	(0.00)	(0.13)	(0.03)	(0.00)			
CIFAR							-0.01^c		
							(0.01)		
U.S. Market Volatility	1.51^c	1.44^c	1.43^c	1.50^c	1.44^c	1.56^c	1.39^c	1.43^c	1.42^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.008	0.004	0.004	0.008	0.004	0.008	0.064^c	0.006	0.007
	(0.20)	(0.47)	(0.53)	(0.21)	(0.47)	(0.21)	(0.00)	(0.29)	(0.29)
SIZE	-0.27^c	-0.28^c	-0.27^c	-0.29^c	-0.28^c	-0.25^c	-0.27^c	-0.28^c	-0.27^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.35^c	0.35^c	0.37^c	0.58^c	0.35^c	0.68^c	0.32^c	0.36^c	0.44^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
DTA	0.99^c	0.97^c	0.97^c	0.94^c	0.97^c	0.97^c	0.94^c	0.94^c	0.94^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
MTB	-0.06^b	-0.05^b	-0.05^b	-0.05^b	-0.05^b	-0.04^b	-0.06^b	-0.05^b	-0.05^b
	(0.03)	(0.04)	(0.04)	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)
CIVIL		0.32^c			0.32^c			0.39^c	
		(0.00)			(0.00)			(0.00)	
ANTIDIR			-0.04^b		0.00				-0.08^c
			(0.05)		(0.97)				(0.00)
DISPERSION						0.24^b			
						(0.03)			
NUMANALYS						-0.02^c			
						(0.00)			
Industry Fixed Effect				yes					
EDF	1480	1327	1327	1472	1326	1036	1326	1350	1350
Adj. R-square	0.34	0.34	0.33	0.37	0.34	0.38	0.33	0.34	0.33
Panel B: Emerging Market Firms									
Intercept	3.11^c	4.01^c	3.86^c	3.61^c	4.02^c	3.10^c	3.60^c	2.34^c	3.33^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
GCRSCORE	-0.10	-0.29^c	-0.10	-0.13^b	-0.29^b	-0.07			
	(0.10)	(0.00)	(0.48)	(0.03)	(0.05)	(0.33)			
CIFAR							-0.01		
							(0.13)		
U.S. Market Volatility	1.57^c	1.46^c	1.47^c	1.51^c	1.46^c	1.68^c	1.51^c	1.47^c	1.48^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

continued on next page

Table 2.9 – continued from previous page

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Local Market Volatility	0.111^c	0.079^c	0.075^c	0.125^c	0.079^c	0.100^c	0.085^c	0.090^c	0.068^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
SIZE	-0.41^c	-0.43^c	-0.44^c	-0.46^c	-0.43^c	-0.44^c	-0.41^c	-0.40^c	-0.41^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.46^c	0.45^c	0.44^b	1.12^c	0.45^c	0.72^c	0.22	0.23	0.32^a
	(0.01)	(0.01)	(0.02)	(0.00)	(0.01)	(0.00)	(0.24)	(0.21)	(0.07)
DTA	1.30^c	1.12^c	1.35^c	1.06^c	1.12^c	0.95^c	1.28^c	1.04^c	1.28^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)
MTB	-0.14^c	-0.12^c	-0.13^c	-0.06^c	-0.12^c	-0.12^c	-0.15^c	-0.11^c	-0.13^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CIVIL		0.56^c			0.56^c			0.61^c	
		(0.00)			(0.00)			(0.00)	
ANTIDIR			-0.09^b		0.00				-0.12^c
			(0.04)		(0.99)				(0.00)
DISPERSION						0.19^a			
						(0.08)			
NUMANALYS						-0.01			
						(0.30)			
Industry Fixed Effect				yes					
EDF	713	579	579	705	578	505	592	602	602
Adj. R-square	0.42	0.42	0.41	0.47	0.41	0.45	0.39	0.41	0.41
Panel C: Developed Market Firms									
Intercept	1.97^c	1.29	2.09^c	1.77^b	1.20	1.75^b	2.74^c	1.46^c	1.53^c
	(0.00)	(0.11)	(0.00)	(0.02)	(0.13)	(0.02)	(0.00)	(0.00)	(0.00)
GCRSCORE	-0.09	0.03	-0.10	-0.12	-0.01	-0.16			
	(0.42)	(0.83)	(0.42)	(0.26)	(0.93)	(0.17)			
CIFAR							-0.02^c		
							(0.01)		
U.S. Market Volatility	1.34	1.33^c	1.32^c	1.35^c	1.34^c	1.38^c	1.37^c	1.33^c	1.32^c
	(0.39)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Local Market Volatility	0.006^c	0.007	0.006	0.004	0.007	0.007	-0.099	0.007	0.008
	(0.00)	(0.31)	(0.40)	(0.59)	(0.35)	(0.37)	(0.45)	(0.31)	(0.21)
SIZE	-0.20^c	-0.22^c	-0.21^c	-0.22^c	-0.23^c	-0.16^c	-0.20^c	-0.22^c	-0.21^c
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
TANG	0.20	0.28^a	0.21	0.13	0.31^a	0.58^c	0.21	0.28^a	0.20
	(0.22)	(0.10)	(0.21)	(0.50)	(0.07)	(0.00)	(0.20)	(0.10)	(0.23)
DTA	0.57^c	0.63^c	0.58^c	0.36^a	0.62^c	0.64^c	0.60^c	0.63^c	0.57^c
	(0.00)	(0.00)	(0.00)	(0.07)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
MTB	-0.04^a	-0.04^a	-0.04^a	-0.04^a	-0.04^a	-0.02^a	-0.04^a	-0.04^a	-0.04^a
	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.10)	(0.06)	(0.06)	(0.06)
CIVIL		0.20^b			0.36^c			0.19^b	
		(0.04)			(0.00)			(0.02)	
ANTIDIR			0.00		0.08^b				-0.01
			(0.99)		(0.04)				(0.65)
DISPERSION						0.16			
						(0.30)			
NUMANALYS						-0.02^c			
						(0.00)			
Industry Fixed Effect				yes					
EDF	759	739	739	751	738	521	726	740	740
Adj. R-square	0.28	0.29	0.28	0.32	0.29	0.33	0.29	0.29	0.28

Table 2.10: C2 Adverse Selection Measure and Home Capital Market Environment

This table reports the estimates of alternative non-spread based adverse selection measure on measures of home capital market environment. Panel A reports the regression results of the whole sample. Panel B reports the regression results of the emerging market firms sub-sample. Panel C reports the regression results of the developed market firms sub-sample. Dependant Variable C2 is estimated from the coefficients c_{2it} of $r_{it}(\tau) = c_{0it} + c_{1it}r_{it}(\tau-1) + c_{2it}T_{it}(\tau-1)r_{it}(\tau-1) + \varepsilon_{it}(\tau)$, where r_{it} is the stock return of ADR i in year t, τ is an indicator for days in year t, $T_{it}(\tau)$ is the natural logarithm of the daily turnover detrended by its mean over the past 200 observations. GCRSCORE is a survey-based measure from Global Competitive Report, and proxies for information transparency. CIFAR is a accounting-standard based measure of disclosure. ANTIDIR is a measure of minority investor protection. CIVIL is a dummy variable that equals one if a country's business law is of civil law origin, and zero if it is of common law origin. SIZE is logarithmic of firm size, TANG is percentage of tangible assets in total assets, DTA is the debt-to-asset ratio, and MTB is the market-to-book ratio. The regressions are estimated cross-sectionally for each year, and the mean of resulting estimates are reported. P-value in parentheses are referenced from T-distribution. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	GCRSCORE	CIFAR	ANTIDIR	CIVIL
Panel A: All Firms				
Intercept	0.08	-0.08	0.51	1.25
	(0.96)	(0.98)	(0.80)	(0.48)
Disclosure and Governance	0.25	0.02	0.07	-0.47
	(0.12)	(0.43)	(0.59)	(0.33)
SIZE	-0.31	-0.34^{a}	-0.26	-0.28
	(0.12)	(0.10)	(0.18)	(0.15)
DTA	1.11	0.67	1.32	1.47
	(0.50)	(0.70)	(0.48)	(0.43)
MTB	-0.40^{b}	-0.34^{b}	-0.39^{b}	-0.41^{b}
	(0.02)	(0.04)	(0.03)	(0.03)
Panel B: Emerging Market Firms				
	GCRSCORE	CIFAR	ANTIDIR	CIVIL
Intercept	8.41^{c}	5.45	5.37^{c}	3.60^{a}
	(0.01)	(0.25)	(0.01)	(0.07)
Disclosure and Governance	-0.68^{a}	-0.01	-0.31^{a}	0.07
	(0.06)	(0.75)	(0.06)	(0.92)
SIZE	-0.73^{c}	-0.70^{c}	-0.68^{c}	-0.63^{c}
	(0.00)	(0.01)	(0.00)	(0.01)
DTA	0.39	0.77	0.96	1.50
	(0.80)	(0.65)	(0.56)	(0.39)
MTB	-0.70^{a}	-0.61	-0.76^{a}	-0.66
	(0.09)	(0.12)	(0.07)	(0.11)
Panel C: Developed Market Firms				
	GCRSCORE	CIFAR	ANTIDIR	CIVIL
Intercept	-7.29^{a}	1.09	-0.83	-0.14
	(0.07)	(0.81)	(0.74)	(0.96)
Disclosure and Governance	1.29^{a}	0.00	0.45^{a}	-0.20
	(0.08)	(0.99)	(0.08)	(0.68)
SIZE	-0.16	-0.28	-0.22	-0.16
	(0.51)	(0.29)	(0.32)	(0.54)
DTA	2.09	1.78	2.16	2.35
	(0.38)	(0.43)	(0.37)	(0.32)
MTB	-0.55^{c}	-0.33	-0.52^{b}	0.40^{a}
	(0.01)	(0.14)	(0.04)	(0.07)

Table 2.11: Excess ADR Returns and Home Capital Market Environment

This table reports the time series (monthly) means of λ_4, the loadings on home capital market environment. $\lambda_{4,t}$ of each month from 2002 to end of 2006 are estimated from

$$r_t - r_{ft} = \lambda_{0,t} + \lambda_{1,t}\hat{\omega}_{wd,t} + \lambda_{2,t}\hat{\omega}_{smb,t} + \lambda_{3,t}\hat{\omega}_{hml,t} + \lambda_{4,t}HCME + \varrho_t,$$

where $HCME$ are our proxies for home capital market environment. $\hat{\omega}_{wd,t}$, $\hat{\omega}_{smb,t}$, and $\hat{\omega}_{hml,t}$ are estimates for each firm month from

$$r_\tau - r_{f\tau} = \alpha + \omega_{wd}(r_{wd\tau} - r_{f\tau}) + \omega_{smb}SMB_\tau + \omega_{hml}HML_\tau + \zeta_\tau,$$

with 36 months of data prior to the month. $r_{wd\tau}$ is the U.S. dollar return of world market index; SMB_τ and HML_τ are, respectively, the Fama-French (1992) size and market-to-book factor returns. P-value in parentheses are referenced from T-distribution. Superscripts a, b, and c indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

	GRCSCORE	CIFAR	ANTIDIR	CIVIL
Disclosure and Governance	-0.51^a	-0.02	-0.18	0.63
	(0.06)	(0.51)	(022)	(0.18)

CHAPTER III

Is There Timing Ability in Currency Markets? Evidence from ADR Issuances

3.1 Introduction

In this chapter, we examine whether foreign firms issuing American Depositary Receipts (ADRs) have the ability to time their corresponding exchange rate market when doing so.[1] We find that these firms tend to issue ADRs after their local currency has been abnormally strong against the U.S. dollar and before their local currency becomes abnormally weak.[2] This result, to our knowledge novel to the literature, is *prima facie* puzzling since *i)* the currency market is among the largest and most

[1]With the increasing integration of the world financial markets, an increasing number of firms are raising capital abroad (Brian Henderson and Weisbach (2004); Karolyi (2006)). The U.S. ADR market, in particular, has become one of the most important venues for foreign firms to raise equity capital outside their local stock market.

[2]Many recent studies have found that firms are able to exploit temporary mispricings in their *local* capital markets via the issuance of overpriced securities (e.g., Graham and Harvey (2001), Baker and Wurgler (2002), Baker, Greenwood, and Wurgler (2003)). Other studies have raised the possibility that foreign firms may be able to time the world *equity* market by cross-listings. Foerster and Karolyi (1999) and Miller (1999) found a statistically significant run-up and subsequent decline of abnormal stock returns within horizons between one week and one year around ADR announcement and/or listing dates. While Miller (1999) related this phenomenon to market segmentation, Foerster and Karolyi (1999) attributed it to strategic market timing decisions by the management of the issuing firms. Along those lines, in a recent study of security issues on the world capital markets, Henderson, Jegadeesh, and Weisbach (2004) provided evidence that firms successfully time their equity issuances when the corresponding stock markets appear to be overvalued.

liquid financial markets, and *ii)* the inability to predict short to long-term exchange rate fluctuations using macroeconomic fundamentals is one of the profession's most documented empirical facts.[3] Yet, we provide evidence that this result is robust and ultimately plausible. This result is potentially important as well since it suggests that some foreign firms may have, at least occasionally, private information about their local exchange rates.[4] Therefore, our study has important implications for modeling and understanding trading activity in currency markets.

We obtain this result by proceeding in three stages. First, we remove the predictable component of all currency returns in our sample. Second, to motivate our analysis, we conduct a preliminary investigation of the dynamics of cumulative abnormal returns in proximity of ADR issue dates using a standard event study methodology. We find a pattern of increasing cumulative *abnormal* (i.e., unpredictable) local currency returns before ADR issue dates and decreasing cumulative abnormal local currency returns after ADR issue dates. However, the statistical significance of this pattern is not uniform across different event window intervals evidenced by the confidence interval but we find the results are statistically significant for some event windows (e.g., at five percent level for [-6, -1] and [1,6] event windows). This might be attributed to several known shortcomings of event studies, such as event clustering, endogeneity, omitted variable bias, and arbitrary horizon selection. For instance,

[3] Meese and Rogoff (1983a; 1983b) showed that exchange rates and fundamentals are largely disconnected. Later studies failed to dispute these basic results. Frankel and Rose (1995) provide a good survey of the subsequent empirical exchange rate literature through the early 1990s. Mark (1995), Mark and Choi (1997), and Mark and Sul (2001) presented some limited evidence that fundamentals may affect only long-term exchange rate returns, but not their short-term fluctuations.

[4] Accordingly, our findings may also help interpret recent evidence on aggregate order flow explaining and predicting currency fluctuations (Evans and Lyons (2002), (2003), (2004)).

the above approach does not control for the timing ability in local and U.S. stock markets documented by Foerster and Karolyi (1999) and Miller (1999). These considerations motivate us to further investigate the relationship between the likelihood and clustering of ADR issuance activity and past and future abnormal currency returns using Poisson regressions. The Poisson approach allows us to address explicitly those shortcomings in the event study methodology.

The results from the Poisson analysis show that non-U.S. companies display economically and statistically significant timing ability in the corresponding exchange rate markets over and above any timing ability in the corresponding equity markets. Specifically, firms tend to issue ADRs after their local currency has been abnormally strong against the U.S. dollar, and before their local currency becomes abnormally weaker, even when controlling for past and future performance of the local and U.S. stock markets.

The findings from the Poisson analysis is consistent with the idea that firms have private information to take advantage of their temporarily high valuations. When the exchange rate returns for a local currency versus the U.S. dollar have been "abnormally" negative (i.e., when the local currency has been abnormally strong), the valuation of a local firm in terms of the ADR issuing currency (U.S. dollar) is likely to be high as well, ceteris paribus for its valuation in the local currency. In other words, when a local currency is abnormally appreciating versus the U.S. dollar, the existing local shareholders are more likely to gain through an ADR issue, since the latter is conceptually equivalent to a short position not only in the local equity but also in the local currency.

This interpretation is *prima facie* puzzling in light of the current state of the exchange rate literature. After all, what private information could foreign firms possibly have that other market participants wouldn't? Yet, additional investigation provides further support for it. First, currency markets are less efficient than commonly thought. Second, we find that currency market timing ability is strongest exactly when, consistent with our intuition above, foreign firms would possess the greatest potential informational advantage, e.g., when they raise capital through their ADR issuances, during financial crises, in emerging markets, or when their business is most sensitive to exchange rate fluctuations. Lastly, our evidence is robust to a wide array of alternative specifications of our basic methodology.

To begin with, temporary mispricings are rife in international financial markets due to various tangible and intangible frictions and imperfections, such as barriers to capital flows, borrowing and shorting constraints, information asymmetry and heterogeneity, "home bias," market segmentation, etc.[5] In addition, most nominal exchange rates against the U.S. dollar are very volatile. These fluctuations are often driven by political considerations, by the actions of price manipulators like Central Banks and other large speculators (e.g., Pasquariello (2004a; 2004b) and references therein), as well as by the existence of exchange rate regimes. These features may in turn offer foreign firms significant opportunities to time their corresponding currency

[5]There is a vast literature documenting these phenomena (e.g., French and Poterba (1991); Bekaert (1995); Bekaert and Harvey (2002); Tesar and Werner (1995); Bertaut and Kole (2004); Yuan (2005)). Consistently, many empirical studies of exchange rate dynamics suggest that the covered interest parity holds for short-term interest rates (e.g., Clinton (1988)), yet find little or no support for the uncovered interest parity (e.g., Froot and Thaler (1990)) or the covered interest parity at longer maturities (e.g., Popper (1993)). According to Pasquariello (2006), inefficiencies and market segmentation are even more pronounced during financial crises. Shleifer (2000) surveys the literature on why mispricings are not always arbitraged away in capital markets.

markets. Finally, exchange rate fluctuations constitute an important determinant of the revenues stemming from an ADR issuance. For instance, assume that a Brazilian firm planning to raise capital through an ADR issue expects the real to depreciate against the U.S. dollar in the next three months. Ceteris paribus, it is optimal for this firm to issue ADR shares sooner and to convert the proceeds later to maximize the expected real-denominated revenues from the issuance.

Second, we are interested in determining which foreign firms may possess *ex ante*, and display *ex post* greater ability to time the exchange rate market in issuing ADRs. To that purpose, we conduct a natural experiment. Specifically, we split our sample into two subsets made of either capital raising (i.e., Level III) or non-capital raising (i.e., Level II) ADR issues. We find strong evidence of currency market timing ability within the sample made of capital-raising ADRs, but no evidence of currency market timing ability within the control sample made of non-capital raising ADRs. Our evidence on currency market timing is economically significant as well, since it translates into total savings for the issuing firms of about $646 million (or 1.86% of the total capital raised via ADRs) over a one-year horizon surrounding the ADR issue dates.

These results allow us to derive two important conclusions. First, they show that foreign firms issuing ADRs exhibit greater currency timing ability when they have stronger incentives to do so, i.e., when such an ability has the potential to translate into monetary savings. Second, they show that our inference is unaffected by why those firms pursue cross-listings in the first place. There is an extensive literature on this subject. Karolyi (1996; 2006) and Doige, Karolyi, Lins, Miller, and Stulz (1998) provide extensive surveys. Within this literature, ADR issues have been motivated by liquidity, cost of capital, visibility, signaling, and corporate governance considerations, among others. Any of these motivations may interact with exchange rate dynamics around ADR issue dates, thus potentially biasing our inference. However, since those motivations behind cross-listings are potentially relevant for both firms raising capital via ADRs and firms that do not, lack of evidence of currency timing ability only for the latter suggests that our inference is unbiased.

We further divide our sample of firms into different groups based on median issue magnitude, median issuing firm size, Tobin's q, and industry, as well as the identity of the issue underwriter, and study the currency timing of ADR issue decisions for each resulting subset. We find that our market timing result is largely driven by relatively big issues by *relatively* small firms (although those firms are large in absolute terms, especially in emerging markets), issues by firms of relatively low q, issues by firms with higher currency exposure, and issues by manufacturing firms. Relatively large ADR issues are more economically significant for relative small firms, thus exchange rate return timing is more crucial to their capital structure decisions. The investment opportunity set of low q firms is relatively small, and their market valuations relatively more stable. Hence, the effect of the exchange rate on their valuations in the issuing currency is relatively more important, making them more selective in choosing the timing of an ADR issue. Firms whose valuations are more sensitive to local currency fluctuations prior to their ADR issuance may possess a deeper understanding of the currency market, hence may display greater currency timing ability when issuing ADRs. Consistently, manufacturing firms, which are more likely to be export-oriented, may develop a greater understanding of fundamentals driving the relevant exchange rates and use this skill to time the currency market. We also find no evidence that this ability can be attributed to the underwriting investment banks, further suggesting that it is instead intrinsic to the issuing firms.

Finally, we show that our evidence is robust to several alternative specifications of our empirical strategies. For example, we consider event windows and currency holding-period returns of up to six months before and after ADR issuances to account for different firms' timing horizons. We find that firms' market timing ability is generally, albeit not homogeneously, significant across all of those intervals. Our basic evidence is even stronger during the occurrence of financial crises and controlling for the timing of market integration. Intuitively, crisis periods are characterized by more intense mispricing, hence currency market timing skills are more valuable to corporations. We also estimate our models for different groups of countries depending on geographical proximity or stage of economic development. Currency market timing ability reveals to be especially relevant, and especially significant for emerging market companies. This reflects the greater importance of exchange rate fluctuations in their issuance decisions.

Overall, this additional evidence on the relationship between currency timing ability and firm-level, issuance, regional, and economic characteristics make our basic result more intuitively convincing. Yet, this result also raises a further question: If, in fact, managers of these firms had private information about future exchange rates, why wouldn't they simply trade on this information (or even concentrate exclusively on this activity) instead of just issuing ADRs (or keeping pursuing their core business activity)? After all, the total savings reported above would be much greater if these companies could divert more capital to time the exchange rate market. There are several reasons why they may not do so. First, the information advantage that may explain firms' currency timing ability could stem from their core business activity. For example, Evans and Lyons (2005) argued that private information about macroeconomic news originates from micro-level dispersed information about production technologies. Without those production technologies, firms would have no information advantage in the currency market. Second, that information advantage may be occasional, i.e., neither long-lasting nor recurrent enough to warrant its systematic exploitation. Third, in this study we argue neither that foreign firms use their currency timing ability exclusively when issuing ADRs nor that exploiting their ability is the most important benefit of those issuances. The literature has identified many such benefits (e.g., Doige, Karolyi, Lins, Miller, and Stulz (1998)), and some of them are arguably greater than any monetary savings from foreign firms' ability to time the local exchange rate. Instead, in this study we concentrate on those firms' ADR issuances because this activity may allow us to empirically identify that ability. Fourth, currency market timing is inherently risky as compared to riskless arbitrage opportunities. Lastly, there may be several capital market frictions (e.g., transaction costs, borrowing constraints, or taxes) preventing foreign firms from fully exploiting their timing ability.

The rest of the chapter is organized as follows. Section 3.2 describes the data and provides summary statistics on ADR issuances and currency and stock returns. Section 3.3 investigates foreign firms' currency market timing ability by examining their ADR issuance decisions and perform several robustness checks. Section 3.4

studies the relation between issue and firm characteristics and firms' timing ability in the exchange rate market. Section 3.5 concludes.

3.2 Data

3.2.1 Issue Statistics

ADRs are dollar-denominated negotiable certificates representing a specific number of a foreign company's local shares held on deposit in the issuer's domestic market. Consistent with the discussion in the previous section, there are two types of ADR issuances in our sample: Level III and Level II ADRs. Level III ADRs are depositary receipts issued over new local equity, i.e., they raise new capital for the issuing firm. Hence, we label these issuances CR ADRs. Since CR ADRs are sold in a public offering, they have to meet the most stringent regulatory and listing requirements with the Security Exchange Commission (SEC) and the chosen exchange (either NYSE, AMEX, or NASDAQ). In contrast, Level II ADRs are depositary receipts issued over existing local equity, i.e., they do not raise new capital for the issuing firm. We therefore label them non-capital raising (non-CR) ADRs.

To construct the database used in analysis, we start by including all public CR and non-CR ADR issues in the U.S. that were registered with the Security Exchange Commission (SEC) between 1976 and 2003 in Thompson Financial's SDC Platinum tapes. Notably, we do not include Level I and Rule 144A ADRs, as well as Global Depositary Receipts (GDRs) in our sample. This choice is based on the following argument. First, Level I ADRs, Rule 144A ADRs, and GDRs are subject to more diverse registration and reporting requirements than those for Level II and Level III ADR issuances. To begin with, Level I ADRs are non-capital raising issuances traded in the over-the-counter (OTC) market, while Rule 144A ADRs raise capital by being privately placed to sophisticated institutional investors. Further, both programs require little or no review by the Security Exchange Commission (SEC) and the foreign issuing firms are exempt from U.S. reporting requirements under Rule 12g3-2(b). Lastly, GDR programs allow issuers to raise capital in two or more equity markets (including the U.S.) simultaneously, as well as in the Rule 144A private market. Hence, because of their heterogeneous nature, registration and reporting requirements for GDR issuers may vary considerably depending on the specific structure of the U.S. offering.[6] In contrast to Level I ADRs, Rule 144A ADRs, and GDRs, both Level II and Level III ADR programs must comply with a similar set of stringent registration and reporting requirements, such as Form F-6 (registration statement), Form F-20 (financial disclosure), and the timely submission of US GAAP-reconciled financial statements to the SEC.[7] Second, according to the literature, regulatory requirements underlie most of the extant explanations for why foreign firms cross-list in the U.S. Therefore, firms issuing Level II and Level III ADRs are more likely to be motivated by a similar set of considerations while the motivations behind Level I, Rule 144A, or GDR programs are more likely to be diverse. As a result, inference drawn upon comparing currency market timing ability within either Level II or Level III ADR issuances is less likely to be biased than if drawn upon a larger database including

[6]For a detailed description of these features see Karolyi (1996) and Foerster and Karolyi (1999).

[7]Level III ADR programs further require the submission to the SEC of Form F-1 to register the equity securities underlying the ADRs publicly offered in the U.S. for the first time.

Level I ADRs, Rule 144A ADRs, or GDRs.[8]

We further restrict our sample to countries with at least five ADR issues over the sample period, since too few issues from a country may indicate the existence of significant barriers to raising capital in the U.S. stock market. These barriers may in turn hinder the local firms' ability to time the currency market through ADR issues. We also exclude countries adopting fixed exchange rate regimes over the sample period.[9] Nonetheless, the inference that follows is robust to the inclusion of both sets of countries in our sample. In the end we are left with 353 ADR issues from 20 countries.

In Table 3.1 we report summary statistics for these issues. In the sample there are 167 ADR issues from firms in G7 countries, 95 from firms in other developed countries, and 91 from firms in emerging economies. The United Kingdom is the country with the most ADRs issued in the U.S., with 89, followed by Mexico with 40. More than 65% of the ADRs issued from G7 countries are CR ADRs; this percentage is slightly higher for firms from emerging markets. Table 3.1 also shows that the time between the SEC filing date and the issue date, known as the "time spent in registration," varies from firm to firm and from country to country, with a median duration of about a month for most countries.

Table 3.1 further reports, for each country, the total ADR issue volume, the median ADR and CR ADR issue sizes, the relative issue size, firm size, and Tobin's q before the ADR issue. The volume of each ADR issue in U.S. dollars is computed as the

[8] Our sample choice is also consistent with previous research on the costs and benefits of cross-listings, e.g., Doige, Karolyi, Lins, Miller, and Stulz (1998) and references therein.

[9] These countries are Argentina, China, and Hong Kong, whose currencies were all pegged at some point to the U.S. dollar.

number of ADRs issued times their issue price as reported in the SDC Platinum database. Firm size and q values are obtained by matching the issues in the sample with the COMPUSTAT database. Firm size (market capitalization in U.S. dollars) before the issue is calculated by multiplying the firm's average share price over the months prior to the issue (within the same year) with the corresponding total number of shares outstanding and then adjusting for the local exchange rate versus the U.S. dollar; relative issue size is the ADR issue amount normalized by firm size. When the ADR issue coincides with a firm's *initial* public offering (IPO), firm size is instead calculated by multiplying the issue price by the total offering amount in all markets, while the relative issue size is calculated by dividing the amount issued in the U.S. stock market by the total amount issued in all markets. Finally, a firm's q before an ADR offering is computed by dividing the firm's market capitalization before the issue by its corresponding book value. For Initial Public Offering (IPO) issues, we replace the market price with the issuing price, and the book value before the issue with the first available book value afterward in COMPUSTAT.

Not surprisingly, both the total and the median issue size by firms from G7 and other developed countries are bigger than those from emerging market firms, with the United Kingdom being the country with the biggest total ADR issue volume (\$15,724 million) and Germany being the country with the biggest median issue size (\$701.3 million). Among the emerging economies, Asian companies have the largest offerings, especially those from South Korea and Taiwan. Issues from Latin America are generally smaller. Interestingly, ADR issuing firms from emerging countries are bigger on average than their counterparts from G7 and other developed nations.[10] Furthermore, with few exceptions (India and Chile), emerging CR ADR issues are always larger in size than the corresponding non-CR issues. The opposite is true for issues from developed economies (with the exception of Norway and Sweden). Finally, and consistently with recent evidence in Gozzia, Levinea, and Schmukler (2008), the median Tobin's q of ADR issuers is significantly greater than one, albeit heterogeneously so across our sample.

3.2.2 Currency and Equity Returns

We complement the above database with monthly exchange rate data. The adoption of a monthly frequency is not casual. This choice is consistent with the median duration in registration reported in Table 3.1, i.e., with a median delay between SEC filing date and issue date of about a month for most countries in the sample. More important, the monthly frequency allows us to control for market microstructure effects and liquidity considerations in the exchange rate data. Finally, the monthly frequency allows us to examine firms' market timing ability over reasonably long (thus more challenging) periods of time, facilitating the interpretation of the economic significance of our results.

Monthly exchange rates are obtained from the Federal Reserve Bank of New York, which collects average noon market buying prices, with the exception of the Chilean peso and the Israeli shekel. Those exchange rates, often constrained within bands of fluctuations and allowed to float later in the sample, are obtained from International

[10]The ADR issuers from South Korea and Taiwan are quite large compared to ADR issuers from other countries. This explains their small median relative ADR issue size in the sample.

Financial Statistics (IFS). The resulting dataset starts from January 1975 for G7 and other developed countries; for emerging economies, the time series of exchange rates starts from the first month when the local ADR market became officially available to local issuers.[11] The resulting total number of monthly observations for each country is shown in Column C of Table 3.2.

Exchange rates are defined as units of local currency per U.S. dollar. We correct the data for such disruptions as the adoption of the euro for six European Union (EU) countries in 1999 and for Greece in 2001. Hence, exchange rate returns for the euro versus the U.S. dollar are used for these countries after their respective switching dates. Mean and standard deviations of logarithmic exchange rate returns are reported in Column A of Table 3.2, together with first-order autocorrelations ($\rho(1)$). Average monthly exchange rate returns among G7 and other developed countries range from -0.31% for the Japanese yen to 0.26% for the Italian lira, and among developing economies from 0.25% for Asian countries to more than 100 basis points for Brazil and Mexico. There is also some evidence of (weak) persistence in currency fluctuations: First-order autocorrelations are positive and statistically significant, yet never greater than 0.48 (Column A of Table 3.2). These facts suggest the need to control for existing trends in these exchange return series. We do so in the next section.

Finally, our sample includes local and U.S. monthly stock market returns. Logarithmic stock returns are computed from Datastream's Total Market Indices for each country in their respective domestic currencies. Column B of Table 3.2 reports

[11] These dates are from Bekaert, Harvey, and Lumsdaine (2002).

mean and standard deviation of those market returns over the same interval as for the corresponding currency returns. As expected, monthly stock market returns are characterized by significantly lower autocorrelations.

3.3 Timing Ability in Exchange Rate Markets

The core notion of the market timing theory of capital structure is that companies would raise capital by issuing overvalued securities (e.g., Baker and Wurgler (2002)). Within a national context, this argument translates into firms choosing equity over debt and vice versa (Baker and Wurgler (2002)) or among different debt maturities (Baker, Greenwood, and Wurgler (2003)) according to their perceived relative mispricings. From an international perspective, the relative overvaluation or undervaluation of the domestic currency may be crucial as well for firms tapping into foreign capital markets. Hence, the level of the exchange rate at the time of a security issue is going to affect the ensuing proceeds for the issuing firm. Moreover, since there is evidence that security mispricing is more pronounced in international financial markets (e.g., Henderson, Jegadeesh, and Weisbach, (2004)), we would expect those markets to offer greater *ex ante* market timing opportunities.

The U.S. market for ADRs represents one of the most important sources of funding for foreign firms (e.g., Karolyi (1996); Bailey, Chan, and Chung (2002)). Ceteris paribus for its funding needs and valuation in the corresponding local currency, one such firm could maximize the U.S. dollar proceeds of its ADR offering if able to execute the issue around the time when its local currency is or has been "abnormally" strong and/or before its local currency is going to be “abnormally” weak. The first objective of this chapter is to test for the existence of this ability. More specifically, the main hypothesis we test in this study is whether foreign firms consider currency market conditions in their ADR issuance decisions and, in doing so, display some ability to time the exchange rate market. In other words, we intend to test whether exchange rate returns follow a pattern around ADR issue dates consistent with the above considerations, i.e., whether ADR issues can be predicted by exchange rate returns before their occurrence and whether ADR issues can predict exchange rate fluctuations afterward.

We employ two methodologies to investigate the currency market timing abilities of firms. The first is a traditional event study approach where we examine cumulative abnormal exchange rate returns around ADR issue dates. The second is a Poisson analysis where we investigate the relationship between the likelihood and clustering of ADR issues and exchange rate returns over different investment horizons. We describe these methodologies and our ensuing results below.

Before proceeding, a potential concern must be addressed. A firm should be deemed to have timing ability in the exchange rate market only if quickly reacting to or anticipating currency fluctuations which could not be predicted by time trends and/or time-series models. The latter would be the case, for instance, of a currency in a slow but prolonged depreciation/appreciation process against the U.S. dollar (such as in “crawling” managed floating regimes). These exchange rate movements, being already expected, may also be already priced into ADR offerings by the equity market, thus giving the issuing firm little incentive and opportunity to time the currency market.

Therefore, we argue that the effect of exchange rate fluctuations on firms' decisions of when to issue ADRs should be limited to its unexpected components. By removing these trends in exchange rate returns, we attempt to isolate the market timing decision from those considerations, and simultaneously provide a tighter benchmark against which exchange rate market timing ability can be detected.

Thus, we detrend all the exchange rate returns for each country n in our sample according to the following AR(2) model with a time trend:

$$exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}, \tag{3.1}$$

where $exrret_{nt}$ is the logarithmic exchange rate return for the currency of country n against the U.S. dollar over month t. In Table 3.2, we report the corresponding R^2 from the estimation of Eq. (3.1) and the Box-Ljung statistic (computed up to lag 6) for the resulting series of estimated currency and stock return residuals ϵ_{nt}. Overall, Eq. (3.1) appears to be successful in removing the predictable component of exchange rate and equity returns: The null hypothesis that the detrended exchange rate series (ϵ_{nt} in Eq. (3.1)) is white noise cannot be rejected in all cases except Chile, Israel, and New Zealand. Similarly, the detrended local stock return series resemble white noise series as well, with the sole exception of India. In addition, estimates for the first order autocorrelation of ϵ_{nt}, not reported here, are statistically indistinguishable from zero for all detrended currency and stock market returns in our sample. Finally, the R^2 from the estimation of Eq. (3.1) for $exrret_{nt}$ is generally small, ranging from 3% (Israel) to 32% (South Korea). This suggests that the unexpected portion of monthly exchange rate fluctuations, i.e., unexplained by Eq. (3.1), is nonetheless economically significant.[12]

This detrending procedure has no bearing on the results below. In fact, these results are even stronger when we measure firms' currency timing ability with respect to the undetrended exchange rate series. In alternative to Eq. (1), we could have employed a structural model of exchange rate determination. For example, currency dynamics have been related to interest rate differentials, purchasing power parity (PPP), budget and current account deficits/surpluses, or relative GDP growth. Yet, most empirical evidence shows that macroeconomic fundamentals do not explain monthly exchange rate changes (see Meese and Rogoff (1983a)). According to Evans and Lyons (2005), the currency determination puzzle is "the most researched puzzle in international macroeconomics." This motivates our choice of a model-free approach to control for the predictable component of exchange rate dynamics in this study.

3.3.1 Event Study Analysis

We start by analyzing the behavior of exchange rate returns in proximity of ADR issuance dates (as in Foerster and Karolyi (1999)) using a standard event study methodology. More specifically, for any $j \in [-H, H]$ and for any country n, we estimate the following model:

$$\epsilon_{nt} = \alpha + \sum_{j=-H}^{H} \delta_j I_{nt}(j) + \eta_{nt}, \tag{3.2}$$

[12]As expected, the R^2 from the estimation of Eq. (3.1) for local and U.S. stock market returns is instead much lower, ranging between 0.9% (Mexico) and 11% (Chile).

where ϵ_{nt} are the detrended currency returns from Eq. (1) and $I_{nt}(j)$ is a dummy variable defined as:

$$I_{nt}(j) = \begin{cases} 1 & \text{if there is at least one ADR issue in country } n \text{ in month } t+j, \\ 0 & \text{otherwise.} \end{cases}$$

The choice of an appropriate event window (i.e., H) in Eq. (2) is important, yet difficult to make. To help us capture evidence of currency market timing ability, such windows must include foreign firms' investment horizons in timing exchange rate fluctuations. Furthermore, those horizons could be different across firms, nations, or regions. We balance these considerations by adopting a relatively long estimation window of $H = 6$ months prior to and after each ADR issuance event in the sample. However, the results that follow are robust, both qualitatively and quantitatively, to alternative choices for H.[13] The estimated coefficient δ_j in Eq. (2) represents the average marginal (i.e., monthly), *abnormal* exchange rate return j months before (if $j > 0$) or j months after (if $j < 0$) an ADR issue from a firm in country n. Therefore, successive sums of those dummy coefficients can be interpreted as measures of the cumulative abnormal impact of ADR issuances on exchange rates. For example, $\sum_{j=-3}^{6} \delta_j$ is a proxy for the cumulative abnormal impact of ADR issuances on the corresponding exchange rate return from 6 months before the event occurred up to

[13] When implementing the analysis in Section 3 for $H = 12, 24, 36$, we found little or no evidence of currency market timing ability over those longer horizons. This is not surprising, since it is unlikely that firms would display longer-run currency timing ability over detrended exchange rate returns series. These additional estimations, available on request, should nonetheless be interpreted with caution, for longer windows considerably shrink the sample of available ADR issues, especially by emerging market companies.

3 months afterward. We estimate Eq. (2) using the pooled data of ADRs from all countries and report the resulting estimated cumulative coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$ in Figure 3.1, together with their 95% confidence intervals.[14]

In the top panel of Figure 3.1, cumulative abnormal exchange rate returns display a U-shape pattern around ADR issue dates, i.e., they decrease before ADR issuances and increase afterward. This pattern is due to point estimates of the marginal impact of ADR issues on exchange rate returns (δ_j) being first negative and then positive. Before ADR issues, exchange rate returns are below their trend, i.e., local currencies are on average relatively strong against the U.S. dollar; following ADR issues, exchange rate returns are instead above their trend, i.e., the local currencies are on average relatively weak against the U.S. dollar, eventually reverting to pre-event trend levels. Interestingly, the above pattern is centered around one month before the ADR issue month; this is consistent with the average lag between ADR filing dates and issue dates of 28 days reported in Table 3.1.

We further analyze the extent of currency market timing ability across the two subsets of our sample made of only capital raising (CR) and non-capital raising (non-CR) ADR issues, respectively. Recall that CR ADRs represent new equity issued and non-CR ADRs instead represent existing local equity. Ex ante, we expect the former (218 in our sample) to exhibit the greatest timing ability, since CR issues are a crucial source of capital for the issuing corporation and as a result currency movements could have a significant impact on the amount raised. By contrast, non-

[14] In this study, we do not measure currency market timing ability at the country level, since the number of ADR issuances in each of the markets in our sample (in Table 3.1) is often not large enough to allow for meaningful statistical inference.

CR ADRs (the remaining 135 in our sample) generate no net revenue for the issuing firm by definition.

More important, non-CR ADRs allow us to address potential omitted variable biases in our empirical analysis. Specifically, as previously mentioned, foreign companies may issue ADRs for a variety of reasons, such as to expand their shareholders' base, to reduce their cost of capital, to gain greater international visibility, to increase liquidity, to signal quality, or to commit to improve governance (see Doidge, Karolyi, Lins, Miller, and Stulz (1998) for a review). Our empirical methodology does not explicitly control for any of these considerations. However, we account for their presence by estimating the extent of currency market timing ability across CR and non-CR issuances separately, since these considerations should apply to both. Therefore, evidence of timing ability in the CR group and lack thereof in the non-CR group would suggest that these potential biases do not affect our inference.

The resulting patterns (in the bottom panel of Figure 1) are striking: Cumulative abnormal currency returns around CR ADR issue dates display a much more pronounced U-shape profile than for the whole sample while no such evidence is found in the control sample made of non-CR ADRs. This suggests that *i*) firms display the greatest currency market timing ability when raising capital, i.e., when such ability is most valuable and translates into monetary savings for the issuing firms, and that *ii*) such timing ability cannot be attributed to omitted variable biases in our empirical analysis.

The evidence presented so far is consistent with foreign firms successfully attempting to maximize their expected proceeds from ADR issuances by timing issue dates according to exchange rate fluctuations, hence consistent with those firms possessing market timing ability in their local currency markets. Supply imbalance and signaling considerations cannot explain this result. The former, which stem from the imperfect substitutability of assets denominated in different currencies, would cause the local currencies to *appreciate* versus the U.S. dollar in response to the sale of significant U.S. dollar amounts from ADR proceeds (hence a reverse U-shape pattern), contrary to our evidence of post-issuance *depreciation* displayed in Figure 3.1. Moreover, ADR volumes, albeit significant, are much smaller than the average daily volume of trading in most of the currencies under examination (e.g., BIS (2002)). The latter is also incompatible with the observed U-shaped patterns in exchange rate returns, since ADR issuances represent good (rather than bad) news for domestic economies.

It is important to show that our finding of firms' currency market timing ability is not subject to the aggregate *pseudo* market timing bias described by Butler, Grullon, and Weston (2005; 2006). Pseudo market timing, in our context, is the tendency for foreign firms to issue ADR following a run-up in their currencies. In a small sample, pseudo currency market timing could give the appearance of genuine currency market timing. Yet, according to Stambaugh (1986, 1999), this bias, also known as small sample predictive regression bias, is most severe when the sample size is small, predictors are persistent, and their innovations are highly correlated with returns. These conditions do not pertain to our empirical analysis. First, our sample is relatively large (monthly return observations around 353 ADR issues in 20 countries over 28 years). Second, the regressors we employ here are event dummy variables ($I_{nt}(j)$ in Eq. (3.2)) rather than persistent aggregates such as equity and debt issue volumes, dividend initiations, or corporate investments.[15] Third, Baker, Taliaferro, and Wurgler (2004) show that the bias introduced by aggregate pseudo market timing is of small empirical relevance (e.g., only about one percent of the predictive power of the equity share in new issues). Nevertheless, to further analyze the robustness of our results, we compute confidence intervals for cumulative abnormal currency returns using a nonparametric bootstrap procedure in which we randomly draw returns from the sample with replacement. The resulting 90% confidence bands constructed from bootstrapped standard errors and centered around zero (also reported in Figure 3.1) suggest that the most significant pattern is around CR ADR filing dates.

We also test whether firms in different regions or from countries at different stages of economic development may have different ability or incentives to time the exchange rate market. To that purpose, we estimate Eq. (2) for the various subsets of nations specified in Table 3.1: G7 countries, other developed countries, emerging markets, and, within the latter, emerging Asia and Latin America. We report the resulting estimated cumulative coefficients in Figure 3.2, together with their 95% confidence intervals. The plots for our regional groupings reveal some degree of heterogeneity in currency market timing ability. G7 and emerging economies (especially in Latin America) display a cumulative excess exchange rate return pattern similar to the U-shaped one observed for the whole sample. Yet, the sequence of local currency appreciation is much more dramatic for emerging currencies, i.e., up to almost 2.5% over the six months leading to an ADR issue. In contrast, exchange rates of other de-

[15] We further explore the relationship between the cross-sectional and intertemporal dynamics of those dummy variables and both currency and equity returns in the Poisson regressions in the next section.

veloped nations are relatively flat before ADR issues, but then depreciate significantly (by almost 1.5%) in the following months. Finally, emerging Asian currencies display an opposite pattern, for local exchange rates appreciate by about 150 basis points over the last few months before an ADR issue and are relatively stable afterward.

Overall, this evidence suggests that, not only in aggregate but even within different regions of the world, foreign firms may be able to time the foreign exchange market by issuing ADRs following a run-up of their domestic currencies and before a reversion of their trends, especially when issuing unseasoned equity; yet, the extent of this timing ability seems to vary across regions and markets and statistically significance is not uniform across event windows.

3.3.2 Poisson Analysis

The results reported in the previous section suggest the existence of timing ability in the exchange rate market. However, the event study methodology we employed to generate them suffers from several shortcomings ultimately affecting their statistical significance as well as their interpretation. First, the regressions of Eq. (2) are univariate, i.e., do not control for other factors affecting the timing of ADR issuances, such as the dynamics of local and U.S. stock markets. Second, the cumulative abnormal excess currency return estimates implicitly weigh each monthly marginal coefficient equally, hence preventing us from identifying firms' best timing horizons.[16] Third,

[16] Many factors may affect the firm's timing horizon in deciding when to issue, i.e., the horizon over which that firm may time the exchange rate market with an ADR issuance. For example, since the process leading to an ADR issue is lengthy and cumbersome, a firm may not be able to promptly take advantage of every abnormal exchange rate return opportunity. In contrast, a firm has always the option not to issue a registered ADR if its exchange rate expectations are not of its liking.

this approach ignores the possibility that multiple ADRs may be issued in the same month from different firms within the same country. In other words, that information is lost in regressing exchange rate returns on the dummies around ADR issue dates ($I_{nt}(j)$). Most importantly, under the alternative hypothesis that currency market timing ability is present, ADR issue dates are endogenous to past and/or future exchange rate dynamics. A sufficiently large window around each event date, as well as time trends and lagged dependent variables (in Eq. (1)), may attenuate but not eliminate this endogeneity problem.

To address these issues directly, we employ an alternative methodology. More specifically, we estimate the effect of both abnormal currency and (local and U.S.) stock holding-period returns on the probability (thus the timing) of the ADR issue decision via a Poisson regression model. Poisson regressions allow us to test for firms' timing ability over different investment horizons while controlling for the clustering of ADR issues within each month. Consistently with the patterns shown in Figure 3.1, we would expect the likelihood of a firm to issue ADRs to be greater after its local currency abnormally appreciated against the U.S. dollar; we also expect *more* firms to issue ADRs the greater is the past abnormal appreciation of the domestic currency. Along the same lines, we would expect the likelihood of a firm to issue ADRs to be greater before its local currency abnormally depreciates against the U.S. dollar; and similarly we also expect more firms to issue ADRs the greater is the future expected abnormal depreciation of the domestic currency.

We proceed in three steps. First, we compute excess holding period returns over horizons of length $h \in [-H, H]$, labeled $adjexrret_{nt}(h)$, by summing up monthly excess exchange rate returns ϵ_{nt} from Eq. (1) up to and excluding the event month, i.e., $adjexrret_{nt}(h_{<0}) = \sum_{s=t+h}^{t-1} \epsilon_{ns}$ for $|h|$-month horizons before the event month t and $adjexrret_{nt}(h_{>0}) = \sum_{s=t+1}^{t+h} \epsilon_{ns}$ for h-month horizons after the event month t. Along the same lines, we compute excess holding period returns for the local stock markets, $adjmktret_{nt}(h)$, and for the U.S. stock market, $adjusret_t(h)$, for each horizon of length h. Second, we assume that the number of ADR issues from country n in month t, $numissue_{nt}$, follows a Poisson distribution,[17]

$$numissue_{nt} \sim Poisson(\lambda_{nt}). \tag{3.3}$$

Third, we estimate the following Poisson regression model:

$$\begin{aligned} \ln \lambda_{nt} &= \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \\ & \quad \beta_3(h) adjexrret_{nt}(h) + \nu_{nt}(h). \end{aligned} \tag{3.4}$$

Eqs. (3) and (4) are a generalized linear model which we estimate by maximum likelihood for each horizon $h \in [-6, 6]$, along the lines with the analysis of Section 3.3.1, except for the contemporaneous holding-period returns ($h = 0$). Within this model, a positive and significant estimate of $\beta_3(h)$ at horizon $h < 0$ indicates that ADR issues in country n are more likely in month t when realized excess local currency returns have been negative over the period $t + h$ to $t - 1$, i.e., after the local currency has been abnormally appreciating for $|h|$ months prior to the event. Vice versa, a

[17]This assumption is reasonable since the sample average for $numissue_{nt}$ is very close to its sample variance for each of the countries in our database and over the entire set of ADR issue events.

positive and significant estimate of $\beta_3(h)$ at horizon $h > 0$ indicates that ADR issues in country n are more likely in month t when realized excess local currency returns over the period $t+1$ to $t+h$ are positive, i.e., prior to a future abnormal depreciation of the local currency over h months. We report estimates of Eq. (4) for all countries in the sample and over the two subsets made of CR and non-CR ADRs in Table 3.3, and for each regional subset in Table 3.4.

These results provide additional evidence of the existence of currency market timing ability suggested by the event study analysis of Section 3.3.1. Consistent with the patterns presented in Figure 3.1, estimates for the exchange rate return coefficients $\beta_3(h)$ over the whole sample are negative for all windows prior to ADR issuances and mostly statistically significant, and positive for all windows afterward, albeit statistically significant only for three- and six-month horizons (Panel A of Table 3.3). This suggests that firms in our sample are able to issue ADRs neither "too early" nor "too late" relative to the dynamics of the local currency market. Further, and again consistently with the analysis in Section 3.3.1, CR ADR issues are significantly more likely than non-CR ADR issues to occur after an excess appreciation of and before an excess depreciation of the local currency (Panels B and C of Table 3.3). In particular, the estimated coefficients for excess holding-period currency returns, $\beta_3(h)$ in Eq. (3.4), are negative and statistically significant at all horizons prior to ($h < 0$ except when $h = -4$), and positive and statistically significant at all horizons following ($h > 0$) CR ADR issue dates (Panel B of Table 3.3). In contrast, $\beta_3(h)$ is always statistically indistinguishable from zero around non-CR issue dates (Panel C of Table 3.3).[18]

As in Section 3.3.1, we compute non-parametrically bootstrapped p-values for each of the estimated parameters of Eq. (3.4) to analyze the robustness of the above results. These p-values, also reported in Table 3.3, reinforce our earlier inference on the existence of firms' timing ability in currency markets. Additionally, our regressors, i.e., the detrended currency and local and foreign equity returns, do not display persistence, which according to Stambaugh (1986; 1999) would increase the severity of the small-sample predictive-regression bias. Therefore, our Poisson analysis is not susceptible to the aggregate pseudo market timing bias raised by Butler, Grullon, and Weston (2005; 2006).

The evidence in Table 3.3 further suggests that the likelihood of ADR issuances is affected not only by prior abnormal local currency returns but also by the prior abnormal performance of foreign firms' local stock markets. Specifically, we find that estimates for the coefficient $\beta_1(h)$ in Eq. (3.4) for some horizons preceding ADR issuances are positive and statistically significant for both the whole sample and the subsample of NON-CR ADRs (Panels A and C of Table 3.3). These estimates indicate that foreign firms are more likely to issue ADRs following an abnormal run-up in their local stock market. This evidence is weaker for the subsample of CR ADRs (Panel B of Tables 3.3), as well as with respect to the prior abnormal performance of the U.S. stock market (the coefficients $\beta_2(h)$ in all panels of Table 3.3). Hence, foreign firms

[18]Country-specific factors, such as privatizations and political considerations, may have driven some foreign firms' ADR issuance decisions over our sample period, hence may have determined their timing either regardless of or in accordance with the dynamics of the local currencies. Nonetheless, the addition of country-level dummies to the specification of Eq. (3.4) did not meaningfully affect our inference. These results are available from the authors on request.

appear to be much less sensitive to the dynamics of the U.S. stock market prior to their issuance decision. These results are largely consistent with the market timing literature in the U.S. equity market (Baker and Wurgler (2002)) and in international equity markets (Foerster and Karolyi (1999), Miller (1999), Henderson, Jegadeesh, and Weisbach (2004)). Nonetheless, it is important to note that in contrast to our currency market timing results described above, we find no evidence of abnormal dynamics in either the U.S. or any local stock market following Level II and Level III ADR issuances.

Our Poisson analysis also shows that the extent of the currency market timing ability varies greatly across regions, as in Figure 3.2. For example, only short-term run-ups of the local currency (i.e., one and two-month horizons) significantly affect the likelihood of G7 firms to issue ADRs (Panel A of Table 3.4). In contrast, ADR issues from firms in other developed countries appear to be more likely only prior to abnormal local currency depreciations over similarly short windows (Panel B of Table 3.4). Currency market timing ability is even more pronounced when Eq. (4) is estimated across the subsamples of emerging market issuers, although largely limited to the decision to defer the ADR issuance (Panels C, D, and E of Table 3.4). More specifically, the ADR decision of these firms follows past abnormal local currency appreciations, yet appears to be independent from future abnormal currency depreciations, as in Figure 3.2, except over the longest horizon (Panel C of Table 3.4). Intuitively, depreciation risk versus the U.S. dollar is often higher for emerging currencies; thus, valuation risk is often higher for emerging market firms as well, making foreign exchange market timing ability especially crucial for their issuing activity.

Lastly, ADR issues appear to be preceded by an abnormally positive performance of the corresponding local stock markets in most of the regions in our sample (i.e., $\beta_1(h) > 0$ for some $h < 0$ in most panels of Table 3.4), consistent with Table 3.3. A notable exception is represented by Latin American firms, which seem to prefer to issue ADRs following local market downturns ($\beta_1(h) < 0$ for $h < 0$ in Panel E of Table 3.4). This suggests that Latin American companies assign greater weight to currency rather than local equity market dynamics in making their ADR issuance decisions.

Over which horizon is exchange rate market timing more successful? In other words, which of the 12 holding-period returns around the event date t in the corresponding 12 estimations of Eq. (4) across the selected country groupings is the most relevant in explaining the likelihood of ADR issues to take place in month t? To address this question, we could compare the magnitude of the resulting estimated coefficients $\beta_3(h)$ across horizons of different length h. A word for caution is, however, necessary. We should keep in mind that the coefficients $\beta_3(h)$ are estimated for holding-period returns computed over those different windows h. An adequate comparison therefore requires that each coefficient estimate be divided by the corresponding horizon length h. When doing so, we find that the average monthly effects are strongest in the immediate proximity of issues ($|h| = 1$). Hence, foreign firms seem to be most focused on the behavior of their local currencies one month prior to the ADR issuance and most successful in anticipating their reversal within one month afterward.

Interestingly, when examining the estimated coefficient for our set of control variables, we find strong evidence of foreign firms' timing ability in their local stock market, and (more surprisingly, albeit weakly) in the U.S. stock market as well. According to Tables 3.3 and 3.4, ADR issues in the past 28 years were more likely when local and U.S. stock market returns had been abnormally high, i.e., after short or long periods of excessively high market valuations. These results are largely consistent with the market timing literature in the U.S. equity market (Baker and Wurgler (2002)) and for international equity markets (Foerster and Karolyi (1999), Miller (1999), Henderson, Jegadeesh, and Weisbach (2004)). A noteworthy exception is represented by Latin American firms, which seem to prefer to issue ADRs following local market downturns (i.e., $\beta_1(h) < 0$ for $h < 0$ in Panel E of Table 3.4). This suggests that Latin American companies assign greater weight to currency rather than local equity market dynamics in making their ADR issuance decisions.

The evidence in Tables 3.3 and 3.4 nests naturally into the above literature. Generally speaking, these papers suggest that firms should and will take advantage of their relatively high valuations in domestic and international capital markets. Yet, currency timing represents an alternative (and, in some cases, dominant, as in Latin America) set of considerations made by foreign firms when selecting their international capital structure. According to Tables 3.3 and 3.4, when local currencies abnormally appreciate relative to the issuing currency of ADRs, the U.S. dollar, foreign firms expect abnormally high valuations of their assets in U.S. dollars, i.e., abnormally high proceeds from ADRs, and, regardless of prior and expected stock market performance, are more likely to issue them.

3.3.3 Market Timing: Crises and Integration

The evidence presented so far indicates that foreign exchange market timing is especially significant, both economically and statistically, for emerging market firms. Yet, both Figure 3.1 and Table 3.3 also reveal that such ability seems to be limited to the recognition of periods of excess appreciation of the local currency *prior* to ADR issuance events. By contrast, issuers from developed economies display currency market timing ability by expediting their ADR issuances as well. What are the reasons for this apparent dichotomy? Academics and practitioners agree that emerging financial markets differ from their developed counterparts, either for the nature of the trading activity, their institutional features, sensitivity to broad market fluctuations, dependence on foreign investments, or degree of liquidity, just to name a few. Do any of these market characteristics explain the currency timing results described above?

We address this issue in this section. More specifically, we examine the robustness of our market timing results to two crucial events affecting the economic and financial well-being of both emerging and developed countries: Financial crises and market integration. We do so because a majority of the emerging countries in our sample are exposed to these events over a significant portion of our sample period. We start by focusing on the effect of financial turmoil on our inference. To do so, we first amend the event study model of Eq. (2) to control for crisis periods as follows:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \sum_{j=-6}^{6} \delta_j^* I_{nt}^*(j) + \eta_{nt}, \tag{3.5}$$

where $I_{nt}^*(j)$ is a dummy variable equal to one if any firm in country n issued ADRs in month $t+j$ and month $t+j$ is within a crisis period, and zero otherwise. We define our crisis periods as December 1994 to January 1995 for the Mexican Peso Crisis, July 1997 to November 1997 for the Asian Crisis, and August 1998 to January 1999 for the Russian Crisis.[19] In Figure 3.3 we plot the resulting cumulative abnormal currency returns in proximity of ADR issues within and outside the crisis periods for each of the regional subsets listed in Table 3.1. In particular, the dotted lines represent estimates for $\Sigma_{j\in[-6,6]}^{6}\delta_j$, i.e., the cumulative abnormal currency returns around ADR issues occurring over the portion of the sample period privy of financial crises, while the solid lines represent estimates for $\Sigma_{j\in[-6,6]}^{6}\delta_j + \delta_j^*$, the cumulative abnormal currency returns around ADR issues occurring during financial crises.

Figure 3.3 reveals that cumulative abnormal exchange rate returns around ADR issuances are of much greater absolute magnitude during periods of financial turmoil. More interestingly, especially in comparison with Figure 3.1, the U-shape patterns of those return aggregations are more important during crisis periods than during stable periods. Cumulative abnormal currency returns are now downward sloping prior to ADR issues and upward sloping afterward for emerging markets in general, but especially for Latin America. Hence, foreign firms' currency market timing ability, far from disappearing, is actually stronger in correspondence with periods of financial turmoil. This is plausible since crisis periods are exactly when this skill is most valuable to a corporation and mispricing is generally deemed to be most intense. For example, Pasquariello (2006) found that arbitrage violations are most frequent

[19]The use of two sets of dummies in Eq. (3.5) is necessary since these crisis periods do not span the 13-month event window around ADR issuances.

during periods of international financial instability. Figure 3.3 seems to suggest that most foreign companies, but especially those based in Latin America (and, to a lesser extent, Asia), have been able to effectively account for the likelihood of a currency crisis in choosing their international capital structure.

To confirm these findings, we modify the Poisson regression model of Eq. (4) by adding a term capturing the interaction between cumulative abnormal exchange rate holding-period returns and the occurrence of a crisis. Specifically, we estimate

$$\begin{aligned} \ln \lambda_{nt} &= \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \\ &\quad \beta_3(h) adjexrret_{nt}(h) + \beta_4(h) adjexrret_{nt}(h) \cdot Crisis_t + \nu_{nt}(h) \end{aligned} \tag{3.6}$$

where $Crisis_t$ is a dummy variable equal to one if month t is within a financial crisis period, and zero otherwise. Table 3.5 reports estimates for the parameters of the above equation.[20] As compared with Table 3.3, the coefficients measuring the effect of excess holding period currency returns before (after) ADR issuances on the likelihood of these issuances to take place during the event month are still negative (positive), mostly (seldom) significant, and of generally smaller absolute magnitude. Yet, more interestingly (and consistently with Figure 3.3), estimates for the additional impact of currency returns on λ_{nt} during financial crises, $\beta_4(h)$, are mostly negative before ADR issues, mostly positive afterward, and of generally greater absolute magnitude than the corresponding estimates for $\beta_3(h)$, regardless of the selected timing horizon h.

[20] Estimates for the intercept in all the Poisson regressions that follow are similar in sign, magnitude, and statistical significance to those reported in Table 3.3. Therefore, these estimates are omitted for economy of exposition.

Again, foreign firms appear to display better currency market timing ability in times of crisis. Not surprisingly, this is especially true for emerging market companies. The estimated sum of $\beta_4(h)$ and $\beta_3(h)$ for both emerging Asian and Latin American firms is often negative prior to, and often positive following ADR issues. This suggests that local currencies of emerging market firms possess a superior currency market timing ability in proximity of crisis periods, i.e., that those firms are on average successful in recognizing the symptoms of an impending dramatic depreciation of their local currencies and in raising capital accordingly.

Next, we examine the relevance of another important feature of the international financial market, the ongoing process of financial integration among local economies, for the evidence on currency market timing ability established above. Over the course of the last three decades, most of the emerging market countries in our sample have experienced not only those official capital market liberalizations making ADR issuances possible, but also significant regulatory changes that have furthered their effective financial integration with the rest of the world. The process of market integration would clearly have a significant impact on the international capital structure decisions of a firm. The same process also may reasonably affect the likelihood of foreign companies to issue ADRs, therefore altering the dynamics of the relation between exchange rate returns and ADR issuances described so far. Hence, we need to test for the robustness of our evidence of firms' foreign exchange market timing ability to these implications of market integration. To that purpose, we amend again the

Poisson regression model of Eq. (4) by estimating instead

$$\begin{aligned}\ln \lambda_{nt} &= \alpha(h) + \beta_1(h)\, adjmktret_{nt}(h) + \beta_2(h)\, adjusret_t(h) + \\ &\quad \beta_3(h)\, adjexrret_{nt}(h) + \beta_5(h)\, INTEG_{nt} + \nu_{nt}(h),\end{aligned} \tag{3.7}$$

where $INTEG_{nt}$ is a dummy variable equal to one if, on date t, country n has already experienced a significant financial integration regime shift, according to the endogenous chronology reported in Bekaert, Harvey, and Lumsdaine (2002, Table 3), and zero otherwise.

The resulting coefficient estimates, in Table 3.6, reveal that, as expected, foreign firms become more active in the ADR market following the integration of their domestic equity market with the rest of the world: $\beta_5(h)$ is positive and strongly significant (at the 1% level or less) in most cases.[21] Yet, evidence of timing ability in the foreign exchange ($\beta_3(h)$) and local stock ($\beta_1(h)$) markets is unaffected. The introduction of integration dummies does not alter, but rather often magnifies either sign or economic and statistical significance of both sets of coefficients over different investment horizons h, namely negative and significant coefficients prior to, and positive and significant coefficients following ADR issuances. To test the robustness of these findings, we also amend the event study regression of Eq. (2) to account for

[21]Eq. (7) is estimated only for the subset of the countries in the sample whose market integration dates are later than the official liberalization dates, i.e., do not overlap with our sample period (e.g., South Korea and Taiwan).

financial integration by estimating the following model:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \sum_{j=-6}^{6} \delta_j^I I_{nt}^I(j) + \eta_{nt}, \tag{3.8}$$

where I_{nt}^I is a dummy variable equal to one if at least one firm in country n issued ADRs in month $t+j$ and month $t+j$ is past the endogenous financial integration date for country n estimated by Bekaert, Harvey, and Lumsdaine (2002, Table 3), and zero otherwise. We report the resulting estimates in Figure 3.4, where the dotted lines represent estimates for $\sum_{j\in(-6,6)}^{6} \delta_j$, i.e., the cumulative abnormal currency returns around ADR issues occurring before financial integration took place, while the solid lines represent estimates for $\sum_{j\in(-6,6)}^{6} \delta_j + \delta_j^I$, the cumulative abnormal currency returns around ADR issues occurring after financial integration. Figure 3.4 reveals a distinct U-shape pattern for the latter but not for the former. These dynamics confirm the evidence in Table 3.6: Currency market timing ability is more pronounced after financial integration has occurred, especially in emerging markets. Intuitively, fewer barriers to international capital markets facilitate a company's efforts at maximizing its proceeds from the issuance of securities to the public. Therefore, Table 3.6 and Figure 3.4 suggest that market integration strengthens, rather than weakens, the basic finding of our study: Foreign firms display currency market timing ability in issuing ADRs.

Finally, we explore the economic significance of the currency market timing results described above. In particular, we want to gauge the impact of firms' currency market timing ability on their bottom line. To that purpose, we employ the estimated cumulative coefficients from Eq. (3.2) for capital raising ADR issuances, plotted in the bottom left-hand panel of Figure 3.1. We focus on CR ADRs since any currency market timing ability exhibited in their issuance translates into monetary savings for the issuing firms. Specifically, for each subset of countries under consideration, we compute the negative of the cumulative abnormal returns from 6 months before to 1 month before an ADR issue, $-\sum_{j=1}^{6}\delta_j$, and the cumulative abnormal returns from 1 month after to 6 months after an ADR issue, $\sum_{j=-6}^{-1}\delta_j$. We then multiply the resulting estimates by the corresponding median ADR issue size and total ADR issue volume (both from Table 3.1). The ensuing numbers, reported in Table 3.7, represent the average and the total U.S. dollar amounts foreign companies saved by selling ADRs neither "too early" (if $\sum_{j=1}^{6}\delta_j < 0$) nor "too late" (if $\sum_{j=1}^{6}\delta_j > 0$), respectively. Table 3.7 shows that this market timing ability is economically significant. Over the sample period, foreign firms have saved on average about \$0.65 million each (i.e., \$330 million in total) by deferring their ADR issuances and \$0.62 million each (i.e., \$315 million in total) by expediting them. This amounts to economically and statistically significant savings of about 1.86% of the total capital raised via ADRs over the sample (i.e., \$646 million). Not surprisingly, emerging market firms are the biggest beneficiaries, especially in Latin America, where savings averaged \$2.21 million per issue (i.e., for a total of \$203 million) over the five-month period before and \$0.98 million per issue (i.e., for a total of \$90 million) over the five-month period after their ADR issuances. These savings are of even greater magnitude when measured during financial crises (Figure 3.3) or after controlling for endogenous market integration (Figure 3.4).

3.4 Who Times the Exchange Rate Market?

In the previous section, we documented that firms are able to time foreign exchange market through ADR issues. The evidence is stronger after controlling for the occurrence of financial crises and the timing of market integration. Moreover, we found that the foreign exchange market timing ability is especially relevant for emerging market companies. In this section, we investigate further what kind of issuances and firms are more likely to time the exchange rate market.

We first examine whether the relative size of an ADR issuance or the size of the ADR issuing firm lead to differential market timing ability. Specifically, we first divide our sample into four size groups based on the relative ADR issue size and the relative firm size: (1) *BigBig*, which includes all large ADR issues (i.e., above the median relative ADR issue size) from large firms (i.e., above the median issuing firm size) in a country; (2) *BigSmall*, which includes all large ADR issues (i.e., above median relative ADR issue size) from small firms (i.e., below the median issuing firm size); (3) *SmallBig*, which includes all small ADR issues (i.e., below the median relative ADR issue size) from large firms (i.e., above the median issuing firm size) in a country; and (4) *SmallSmall*, which includes all small ADR issues (i.e., below the median relative ADR issue size) from small firms (i.e., below the median issuing firm size) in a country. We then re-estimate both the event study model of Eq. (2) and the Poisson regression model of Eqs. (3) and (4) for each of these four size groups across all countries in the sample. The results are reported in Figure 3.5 and Table 3.8 respectively.

When comparing Panel A of Table 3.3 to the corresponding results in Table 3.8, it is clear that the market timing result documented in Section 3 is driven mostly by relatively big issues from *relatively* small issuers, although those firms are large in absolute terms, especially in emerging markets (see Table 3.1). In particular, the likelihood of a relatively large ADR issue by a relatively small firm is significantly higher after an abnormal appreciation ($\beta_3(h) < 0$ when $h < 0$ in Panel B of Table 3.8) and prior to a future abnormal depreciation of the local currency ($\beta_3(h) > 0$ when $h > 0$ in Panel B of Table 3.8). Intuitively, large ADR issues are more economically significant for small issuers, thus exchange rate return timing is more crucial to their capital structure decision. The dynamics of cumulative abnormal returns around ADR issuances across issue and firm size groups (Figure 3.5) are consistent with these results, with the *BigSmall* grouping displaying the most significant U-shape patterns.[22]

We then test whether a firm's investment opportunity set (proxied by Tobin's q) is an indicator of its foreign exchange timing ability. We do so by first re-estimating the Poisson regression model of Eq. (4) separately for firms with above median Tobin's q, i.e, growth firms, and for firms with below median Tobin's q, i.e., value firms, in each of the countries in our sample. The results reported in Table 3.9 suggest that, on aggregate, the currency market timing evidence of Section 3 is largely driven by firms with low q. Intuitively, the investment opportunity set of low q firms is relatively small, and their market valuations relatively more stable. Hence, the effect of the exchange rate on their valuations in the issuing currency is relatively more

[22] We also estimate both Eqs. (2) and (4) for each of the subsets of countries described in Table 3.2. The results, available upon request from the authors, are qualitatively consistent with those reported in Figure 3.5 and Table 3.8.

important, making them more selective in choosing the timing of an ADR issue.[23] Again, similar results are obtained from the estimation of the event study model of Eq. (2) across value (low q) and growth (high q) firms, reported in Figure 3.6.

Next, we test whether the currency market timing ability of firms issuing CR ADRs is related to greater sensitivity of these firms' business activities to currency fluctuations. To that purpose, we first compare the currency exposure of issuers of CR ADRs versus non-CR ADRs. We estimate this exposure by the absolute value of the "currency γ," $|\gamma_i|$, from the following regression:

$$ret_{int} = a_i + b_i mktret_{nt} + \gamma_i exrret_{nt} + \epsilon_{int} \tag{3.9}$$

where ret_{int} is the stock return of the ADR issuing firm i from country n at time t, $mktret_{nt}$ is the return of the corresponding local stock market, and $exrret_{nt}$ is the corresponding exchange rate return versus the U.S. dollar. Eq. (3.9) is estimated over a period of no less than two and no more than five years prior to the ADR issuing month. This restriction leaves us with a subset of 59 CR ADR and 34 non-CR ADR issuing firms. We find that the median $|\gamma_i|$ for CR ADR issuers (0.52 with a standard error of 0.10) is almost 50% larger than the corresponding median for non-CR issuers (0.32 with a standard error of 0.23). Hence, the valuation of CR ADR issuing firms in the two-to-five year period prior to their decision to issue appears to

[23] When estimating Eq. (4) for low and high q firms across each of the regional groups in Table 3.2, we further find that this dichotomy in currency market timing ability disappears within emerging markets. This is not surprising, since (as suggested in Section 3) depreciation risk represents an overriding concern for Latin American and Asian companies issuing ADRs. These results are available on request from the authors.

be more sensitive to fluctuations of their local currency than the valuation of firms eventually issuing non-CR ADRs. This suggests that CR ADR issuing firms may have developed greater understanding of the currency market than non-CR ADR issuers prior to their issuance decision. This may in turn translate into greater currency timing ability when issuing ADRs.

As a further test of this hypothesis, we examine whether there is differential currency market timing ability across firms in different industries. To that purpose, we divide our sample into the following eight industries according to SIC codes: Agriculture, Construction, Mining, Manufacturing, Utility, Sales, Financial, and Service. Then we estimate both the event study regression model of Eq. (2) and the Poisson model of Eq. (4) over each resulting industry subset of our sample. We report the corresponding estimates in Figure 3.7 and Table 3.10, respectively.[24] Both sets of results indicate that currency market timing ability is most pronounced among firms in the Manufacturing industry. For instance, only the likelihood of a manufacturing firm to issue an ADR is both negatively related to past abnormal currency returns ($\beta_3(h) < 0$ when $h < 0$ in Panel B of Table 3.10) and positively related to future abnormal currency returns ($\beta_3(h) > 0$ when $h > 0$ in Panel B of Table 3.10). Intuitively, the revenues of these firms are more likely to be generated in foreign markets. Therefore, their management is more likely either to develop a deeper understanding of the relevant currency markets, to affect the exchange rate through their business activities, or to lobby for a more favorable currency policy with the corresponding

[24]Neither model could be estimated for the Agriculture and Construction groupings since they covered a total of only five ADR issues.

local government.

Finally, we consider the possibility that the currency market timing ability of foreign firms originates from the investment banks underwriting the issuances rather than from the foreign firms themselves. To do so, we first divide our sample of ADR issue firms into subsets according to the identity of the underwriting institution; then we estimate both Eq. (2) and Eq. (4) across the subsets made of issues managed by the top six underwriting firms in the U.S.: Credit Suisse First Boston (CSFB), Goldman Sachs, Lehman Brothers, Merrill Lynch, Morgan Stanley, and Salomon Smith Barney.[25] The results, reported in Figure 3.8 and column $\beta_3(h)$ of Table 3.11, show little or no evidence of currency market timing ability across investment bank groupings.[26] This fact, together with our previous findings, indicates that currency market timing ability is intrinsic to the issuing firms and not to their advisors.

Overall, the ability of a foreign firm to time the exchange rate market while issuing ADRs appears to be related to important firm and issue characteristics like size, Tobin's q, and industry, as well as to the relative magnitude of the proceeds at stake, but not to the identity of the underwriting investment bank. This evidence corroborates our basic conclusions from Section 3, since it anchors them to intuitive corporate finance grounds.

[25] We did not include in the analysis ADRs underwritten by other investment banks (representing less than one third of the sample) because of the insufficient number of issuances in our sample for each of them separately.

[26] Specifically, and consistently with the cumulative plots in Figure 3.8, only ADRs underwritten by Morgan Stanley (55 in our sample) are more likely to be issued prior to an abnormal depreciation of the local currency: $\beta_3(h) > 0$ and statistically significant for $h = 1, 3, 4, 6$ in the corresponding panel of Table 3.11. Interestingly, Table 3.11 also suggests that ADRs are more likely to be issued following an unexpected run-up of the local equity market when underwritten by Goldman Sachs and Merrill Lynch.

3.5 Conclusion

In this chapter, we assess whether foreign firms can time their corresponding local currency markets by studying the relationship between exchange rate returns and all ADR issuances in the U.S. in the last 28 years. We provide economically and statistically significant evidence of foreign firms' timing ability in the exchange rate market, especially when these firms raise capital through an ADR program. We further show that currency market timing ability is most pronounced for companies with higher currency exposure, value companies, manufacturing firms, relatively small (yet large in absolute terms) companies issuing relatively large amounts of ADRs, and emerging market companies, and especially during currency crises and following the integration of their domestic market with the rest of the world; yet, this ability cannot be attributed to the investment banks underwriting the issues.

Our study is the first to document the existence of currency market timing ability. In addition, our findings also suggest that some market participants in the global foreign exchange market (selected foreign firms issuing ADRs) may have, at least occasionally, private information about currency movements. Thus, timing ability in the exchange rate markets may contribute to interpret recent evidence on the order flow explaining and predicting exchange rate fluctuations (Evans and Lyons (2002), (2003), (2004)). Foreign exchange market timing ability in the ADR market entails foreign firms either possessing private information about the fundamentals driving the long-term dynamics of their local currencies, or being able to affect directly those fundamentals. Therefore, any order flow aggregate containing these companies' trading activity in the local exchange rate markets, and information about it, would play such an important role in exchange rate determination.

Table 3.1: Summary Statistics on ADR Issues

This table reports the summary statistics for our sample of ADR issues. In particular, it displays, for each country in the sample, the first ADR issue date, the total number of ADR issues, the number of capital raising (CR) issues among those, the total ADR issue volume, the total CR ADR issue volume, the median ADR issue size, the median CR ADR issue size, the median firm size, the median relative ADR issue size, and the median Tobin's q before the ADR issue. Firm size and ADR issue sizes are in millions of U.S. dollars. Relative issue size is the ADR issue size normalized by the firm size. Tobin's q is computed by dividing each firm's market capitalization before the ADR issue by its corresponding book value (obtained from COMPUSTAT). Duration in registration is the number of days from the ADR filing date to the ADR issue date.

Country	First Issue	No. of Issues	No. of CR Issues	Total Volume	Total CR Vol.	Median Issue Size	Median CR Issue Size	Median Firm Size	Median Rel. Issue Size	Median Tobin's q	Median Duration
G7 Countries	Jan-76	167	109	$36,237.30	$15693.9	$84.00	$64.00	$242.20	0.43	2.76	31
France	Jun-84	31	22	$4,565.10	$2714.7	$99.40	$96.10	$242.42	0.33	2.61	26
Germany	Jan-94	9	7	$6,478.00	$4167.9	$701.30	$429.80	$5,201.40	0.19	3.63	29
Italy	Jun-89	14	9	$3,122.50	$1362.8	$96.65	$93.60	$312.77	0.18	2.87	35
Japan	Jan-76	24	11	$6,347.40	$835.7	$83.35	$49.40	$1,446.93	0.04	3.17	65
UK	Jun-77	89	60	$15,724.30	$6612.8	$64.00	$57.80	$121.10	0.54	2.37	32
Other Developed	Jul-81	95	45	$14,531.10	$4873.3	$79.80	$70.10	$515.87	0.14	3.97	28
Australia	May-87	16	5	$3,043.20	$717.4	$84.15	$66.10	$2,090.81	0.08	2.22	31
Denmark	Jul-81	6	0	$1,473.80	$0	$53.55	$0	$790.78	0.26	3.68	52
Ireland	Jan-84	22	15	$975.80	$640.6	$43.15	$39.00	$74.05	0.78	6.78	34
Netherlands	Oct-84	5	2	$2,169.70	$561.1	$302.10	$280.55	$1,6133	0.06	1.77	26
New Zealand	Jul-91	7	5	$944.10	$467.4	$79.20	$79.20	$123.60	0.66	3.89	26
Norway	May-83	12	6	$889.10	$446.8	$56.45	$65.65	$352.17	0.14	6.61	30
Spain	Jun-87	18	6	$3,759.50	$1093.4	$187.95	$167.60	$5,548.47	0.03	2.33	24
Sweden	May-83	9	6	$1,275.90	$946.6	$57.20	$101.25	$1,079.34	0.13	4.42	26
Emerging Markets	Aug-87	91	64	$18,419.10	$14232.3	$44.00	$73.45	$310.05	0.18	2.92	40
Israel	Aug-87	7	4	$419.40	$223.4	$44.00	$53.75	$166.36	0.18	2.72	34
Emerging Asia	Oct-94	29	21	$10,412.60	$8242.7	$202.70	$395.80	$7,221.40	0.04	3.4	31
South Korea	Oct-94	14	7	$5,166.90	$3252.7	$201.35	$430.80	$9,287.71	0.03	1.72	32
Taiwan	May-96	9	9	$4,540.10	$4540.1	$497.50	$497.50	$22,743.27	0.03	4.73	41
India	Sep-99	6	5	$705.60	$449.9	$106.00	$75.00	$211.95	0.5	8.95	14
Emerging Latin	Jul-90	55	39	$7,587.10	$5766.2	$62.60	$62.60	$272.50	0.31	2.17	27
Brazil	Oct-95	11	9	$2,828.40	$2711.4	$134.55	$152.10	$1,012.49	0.11	1.52	31
Chile	Jul-90	26	19	$2,075.00	$1143.9	$56.40	$50.90	$143.15	0.51	3.01	26
Mexico	Feb-94	40	11	$2,683.70	$1910.9	$60.80	$67.50	$484.41	0.2	1.9	27
All Sample	Jan-76	353	218	$69,187.50	$34799.5	$79.20	$68.20	$353.73	0.24	2.91	28

Table 3.2: Summary Statistics of Equity and Currency Market Returns

This table reports mean (μ), standard deviation (σ), and first-order autocorrelation ($\rho(1)$) of exchange rate (Panel A) and local stock market returns (Panel B), and the number of available monthly observations (Panel C) for each country in the sample. p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate significance at the 10%, 5%, and 1% level, respectively. This table also reports R^2 from estimating the following AR(2) model with a time trend: $exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}$, where $exrret_{nt}$ is the logarithmic exchange rate return for the currency of country n against the U.S. dollar over month t. We also report R^2s from estimating the AR(2) model above for local stock returns, and the Box-Ljung statistics (computed up to lag 6) for the resulting series of estimated currency and stock return residuals ϵ_{nt}.

Country	**A: Exchange Rate Return**					**B: Local Market Return**					**C**
	μ	σ	$\rho(1)$	R^2	LB(6)	μ	σ	$\rho(1)$	R^2	LB(6)	N
G7 Countries											
France	0.06 %	2.62%	0.30***	9.58%	5.19	0.91%***	6.19%	0.07*	1.05%	6.12	348
	(0.68)		(0.00)		(0.52)	(0.01)		(0.10)		(0.41)	
Germany	-0.11%	2.67 %	0.31***	10.20%	1.58	0.55%**	5.30%	0.09**	0.95%	4.33	348
	(0.42)		(0.00)		(0.95)	(0.05)		(0.04)		(0.63)	
Italy	0.26%*	2.61%	0.39***	16.58%	0.87	0.89%**	7.18%	0.10**	1.04%	5.53	348
	(0.06)		(0.00)		(0.99)	(0.02)		(0.03)		(0.48)	
Japan	-0.29%*	5.28%	0.34***	12.67%	5.90	0.43%	5.28%	0.07*	1.63%	1.04	348
	(0.05)		(0.00)		(0.43)	(0.13)		(0.08)		(0.98)	
UK	0.08%	2.50%	0.35***	14.08%	4.94	1.04%***	5.60%	0.09**	5.17%	6.49	348
	(0.54)		(0.00)		(0.55)	(0.00)		(0.05)		(0.37)	
US						0.81%***	4.29%	0.05	0.76%	2.00	348
						(0.00)		(0.18)		(0.92)	
Other Developed Countries											
Australia	0.17%	2.34%	0.28***	8.93%	4.03	0.85***	5.67%	0.06	2.96%	2.84	348
	(0.19)		(0.00)		(0.67)	(0.01)		(0.11)		(0.83)	
Denmark	0.01%	2.59%	0.32***	10.75%	2.47	0.99%***	5.15%	0.08**	1.29%	5.67	348
	(0.92)		(0.00)		(0.87)	(0.00)		(0.05)		(0.46)	
Ireland	0.12%	2.59%	0.33***	11.38%	3.17	1.16%***	6.60%	0.11***	2.17%	0.22	348
	(0.37)		(0.00)		(0.79)	(0.00)		(0.01)		(0.99)	
Netherlands	-0.09%	2.65%	0.33***	11.57%	1.35	0.77%***	4.99%	0.04	1.17%	2.56	348
	(0.52)		(0.00)		(0.97)	(0.00)		(0.23)		(0.86)	
New Zealand	0.01%	2.29%	0.28***	9.47%	13.63**	0.37%	5.26%	-0.03	0.16%	10.11	194

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Table 3.2(continued)

Country	A: Exchange Rate Return					B: Local Market Return					C
	μ	σ	$\rho(1)$	R^2	LB(6)	μ	σ	$\rho(1)$	R^2	LB(6)	N
	(0.94)		(0.00)		(0.03)	(0.34)		(0.66)		(0.12)	
Norway	0.10%	2.39%	0.37***	15.22%	0.56	0.84%**	7.31%	0.13***	1.91%	7.20	290
	(0.49)		(0.00)		(0.99)	(0.05)		(0.01)		(0.30)	
Spain	0.06%	2.71%	0.45***	14.38%	5.11	0.66%	6.43%	0.13**	3.09%	2.84	205
	(0.77)		(0.00)		(0.53)	(0.15)		(0.03)		(0.83)	
Sweden	0.08%	2.62%	0.40***	17.52%	2.51	1.09%**	2.71%	0.15***	2.86%	1.97	266
	(0.61)		(0.00)		(0.87)	(0.02)		(0.01)		(0.92)	
Emerging Asian											
India	0.35%***	1.30%	0.25***	10.33%	4.25	0.56%	9.00%	0.11*	2.15%	10.56*	134
	(0.00)		(0.00)		(0.64)	(0.47)		(0.06)		(0.10)	
South Korea	0.27%	3.49%	0.48***	31.87%	4.23	0.30%	9.93%	0.13**	2.42%	4.42	184
	(0.30)		(0.00)		(0.65)		(0.68)	(0.04)		(0.62)	
Taiwan	0.15%	1.34%	0.35***	12.00%	9.58	0.27%	9.71%	0.03	0.09%	2.73	156
	(0.18)		(0.00)		(0.14)		(0.73)	(0.35)		(0.84)	
Emerging Latin American											
Brazil	1.13%***	4.63%	0.42***	25.76%	3.66	1.14%	8.83%	-0.02	0.92%	3.04	105
	(0.01)		(0.00)		(0.72)	(0.19)		(0.52)		(0.80)	
Chile	0.42%***	2.12%	0.25***	7.90%	46.20***	1.57%***	6.44%	0.28***	11.19%	3.96	168
	(0.01)		(0.00)		(0.00)	(0.00)		(0.00)		(0.68)	
Mexico	1.08%***	4.50%	0.19**	7.26%	6.81	0.93%	7.53%	-0.02	0.05%	8.41	119
	(0.01)		(0.02)		(0.34)	(0.18)		(0.53)		(0.21)	
Other Emerging Market											
Israel	0.36%**	1.93%	0.06	2.75%	36.15***	0.72%	7.36%	0.03	0.57%	2.40	135
	(0.04)		(0.24)		(0.00)	(0.27)		(0.38)		(0.88)	

Table 3.3: Poisson Regressions: ADRs

This table presents the estimates of the following Poisson regressions for 12 event windows of length $h \in [-6, 6]$ except $h = 0$: $\ln \lambda_{nt} = \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \beta_3(h) adjexrret_{nt}(h) + \nu_{nt}(h)$, where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; and $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t+h\ ,t]$ when $h < 0$), or as h-month after the observation month t (i.e., $[t,\ t+h]$ when $h > 0$). To compute excess dollar exchange rate returns, we adjust for autocorrelation and time trends by estimating $exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}$, where $exrret_{nt}$ is the dollar exchange rate return of country n at month t. Then we compute the excess holding period currency return from month $t+h$ to month t, $adjexrret_{nt}(h_{<0})$, as $\sum_{s=t+h}^{t-1} \epsilon_{ns}$, and the excess holding period currency return from month t to month $t+h$, $adjexrret_{nt}(h_{>0})$, as $\sum_{s=t+1}^{t+h} \epsilon_{ns}$. Excess holding period local stock returns, $adjmktret_t(h)$, and excess holding period U.S. stock returns, $adjusret_{nt}(h)$, are similarly defined. Panel A reports estimates for the whole sample; Panels B and C report estimates for the two subsamples made of either capital raising (CR) ADRs or non-capital raising (non-CR) ADRs. p-values (rounded to two decimal places) are in parentheses. For each estimate, we also report bootstrapped p-values. The bootstrap procedure consists of randomly drawing returns from the observed time series with replacement and estimating the aforementioned model with the resulting sample. After repeating this procedure 1000 times, we locate our Poisson estimates in this simulated distribution with two-tailed p-values. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

Event Window	$\beta_1(h)$	P-value	Boot p-value	$\beta_2(h)$	p-value	Boot p-value	$\beta_3(h)$	p-value	Boot p-value
Panel A: All Countries									
6 months before	1.51	<.0001***	0.01***	0.99	0.10*	0.08*	-1.50	0.09*	0.16
5 months before	1.64	<.0001***	0.01***	1.09	0.10*	0.08*	-1.91	0.05**	0.11
4 months before	1.52	0.00***	0.01***	0.79	0.29	0.17	-1.50	0.18	0.21
3 months before	1.47	0.01***	0.01***	0.72	0.40	0.23	-1.99	0.13	0.14
2 months before	1.66	0.01***	0.01***	1.14	0.28	0.16	-2.94	0.06*	0.08*

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Table 3.3 (continued)

Event Window	$\beta_1(h)$	p-value	Boot p-value	$\beta_2(h)$	p-value	Boot p-value	$\beta_3(h)$	p-value	Boot p-value
1 month before	1.47	0.11	0.08*	1.80	0.21	0.28	-5.41	0.02**	0.03**
1 month after	-0.89	0.34	0.20	2.21	0.12	0.09*	2.84	0.16	0.12
2 months after	0.27	0.69	0.38	-0.14	0.89	0.45	2.24	0.12	0.06*
3 months after	0.12	0.82	0.45	-0.17	0.84	0.45	2.00	0.09*	0.04**
4 months after	0.12	0.81	0.45	-0.18	0.81	0.43	1.24	0.23	0.09*
5 months after	0.03	0.94	0.49	-0.15	0.82	0.42	1.18	0.20	0.07*
6 months after	-0.08	0.83	0.41	0.24	0.69	0.37	2.12	0.01***	0.01***
Panel B: CR ADRs									
6 months before	1.60	0.00***	0.01***	0.55	0.48	0.23	-2.41	0.04**	0.05*
5 months before	1.51	0.00***	0.02**	0.74	0.38	0.19	-2.69	0.04***	0.04***
4 months before	1.44	0.01***	0.04**	0.50	0.60	0.28	-2.20	0.12	0.09*
3 months before	1.05	0.13	0.11	0.76	0.48	0.24	-3.27	0.05*	0.04**
2 months before	1.30	0.12	0.10	1.37	0.31	0.16	-4.35	0.03**	0.03**
1 month before	1.01	0.38	0.22	1.60	0.39	0.19	-5.26	0.07*	0.04**
1 month after	-0.24	0.84	0.37	2.64	0.15	0.09*	4.72	0.05*	0.07*

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Table 3.3 (continued)

Event Window	$\beta_1(h)$	p-value	Boot p-value	$\beta_2(h)$	p-value	Boot p-value	$\beta_3(h)$	p-value	Boot p-value
2 months after	0.57	0.50	0.33	-0.12	0.92	0.49	3.87	0.02**	0.06*
3 months after	0.40	0.56	0.38	-0.10	0.92	0.48	3.54	0.01**	0.04**
4 months after	0.37	0.54	0.37	-0.11	0.90	0.48	2.40	0.06*	0.12
5 months after	0.26	0.62	0.44	-0.35	0.68	0.46	2.11	0.07*	0.13
6 months after	0.13	0.79	0.48	0.37	0.63	0.31	3.19	0.00***	0.04*
Panel C: NON-CR ADRs									
6 months before	0.99	0.11	0.06*	1.70	0.09*	0.05*	-0.13	0.93	0.47
5 months before	1.71	0.01***	0.01***	1.31	0.24	0.11	-0.70	0.67	0.35
4 months before	1.48	0.05*	0.02**	1.14	0.36	0.18	-0.33	0.86	0.45
3 months before	1.81	0.04**	0.02**	0.60	0.67	0.33	-0.59	0.78	0.40
2 months before	2.00	0.06*	0.02**	0.75	0.66	0.31	-1.27	0.63	0.35
1 month before	2.25	0.12	0.06*	1.89	0.43	0.19	-5.85	0.12	0.07*
1 month after	-1.65	0.28	0.10*	1.83	0.43	0.21	-0.90	0.80	0.41
2 months after	-0.02	0.99	0.50	0.11	0.95	0.47	-0.58	0.82	0.41
3 months after	-0.05	0.95	0.46	-0.16	0.91	0.48	-0.25	0.91	0.46

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Table 3.3 (continued)

Event Window	$\beta_1(h)$	p-value	Boot p-value	$\beta_2(h)$	p-value	Boot p-value	$\beta_3(h)$	p-value	Boot p-value
4 months after	-0.24	0.76	0.35	-0.14	0.91	0.50	-0.56	0.75	0.40
5 months after	-0.25	0.73	0.34	0.23	0.83	0.38	0.00	1.00	0.49
6 months after	-0.27	0.68	0.31	0.05	0.96	0.44	0.78	0.58	0.33

Table 3.4: Poisson Regressions: Regional ADRs

This table presents the estimates of the following Poisson regressions for 12 event windows of length $h \in [-6, 6]$ except $h = 0$: $\ln \lambda_{nt} = \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \beta_3(h) adjexrret_{nt}(h) + \nu_{nt}(h)$, where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; and $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t+h\ ,t]$ when $h < 0$), or as h-month after the observation month t (i.e., $[t,\ t+h]$ when $h > 0$). To compute excess dollar exchange rate returns, we adjust for autocorrelation and time trends by estimating $exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}$, where $exrret_{nt}$ is the dollar exchange rate return of country n at month t. Then we compute the excess holding period currency return from month $t+h$ to month t, $adjexrret_{nt}(h_{<0})$, as $\sum_{s=t+h}^{t-1} \epsilon_{ns}$, and the excess holding period currency return from month t to month $t+h$, $adjexrret_{nt}(h_{>0})$, as $\sum_{s=t+1}^{t+h} \epsilon_{ns}$. Excess holding period local stock returns, $adjmktret_t(h)$, and excess holding period U.S. stock returns, $adjusret_{nt}(h)$, are similarly defined. Panel A reports estimates for the whole sample; Panels B to F report estimates for country groups. p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
	Panel A: G-7			**Panel B: Other Dev.**			**Panel C: Emerging**			**Panel D: Emg. Asia**			**Panel E: Emg. Latin**		
6 months before	1.51***	0.29	-0.31	0.91	3.52***	0.98	1.26**	-0.32	-6.87***	2.83***	-2.92	-0.66	0.25	0.14	-6.13***
	(0.01)	(0.75)	(0.80)	(0.28)	(0.00)	(0.58)	(0.02)	(0.77)	(0.00)	(0.00)	(0.12)	(0.90)	(0.76)	(0.92)	(0.00)
5 months before	1.74***	0.03	-1.26	1.37	3.44***	1.85	1.13*	0.32	-7.39***	2.93***	-1.66	-6.72	-0.37	0.57	-6.11***
	(0.01)	(0.97)	(0.35)	(0.14)	(0.01)	(0.34)	(0.06)	(0.80)	(0.00)	(0.00)	(0.44)	(0.31)	(0.68)	(0.72)	(0.01)
4 months before	1.64**	-0.25	-0.77	1.84*	2.59*	2.51	0.75	0.18	-7.85***	3.24***	-2.92	-13.98**	-1.71*	0.85	-5.83**
	(0.03)	(0.82)	(0.62)	(0.08)	(0.09)	(0.24)	(0.26)	(0.89)	(0.00)	(0.00)	(0.22)	(0.03)	(0.08)	(0.63)	(0.02)
3 months before	1.41	-0.59	-1.63	3.07***	1.58	3.08	0.26	1.13	-8.95***	2.19*	-0.93	-16.22**	-2.23**	1.46	-7.20***
	(0.12)	(0.64)	(0.38)	(0.01)	(0.38)	(0.22)	(0.74)	(0.48)	(0.00)	(0.06)	(0.73)	(0.03)	(0.05)	(0.48)	(0.01)
2 months before	1.54	0.24	-3.79*	3.49**	-0.22	4.81	0.36	3.20	-8.77***	2.14*	1.33	-21.48***	-2.40*	3.94	-6.49**
	(0.15)	(0.88)	(0.09)	(0.02)	(0.92)	(0.11)	(0.70)	(0.11)	(0.01)	(0.09)	(0.70)	(0.01)	(0.09)	(0.13)	(0.05)
1 month before	-0.25	2.48	-5.73*	3.50*	0.29	-0.82	1.48	2.35	-8.26*	3.35*	-2.80	-27.79***	-0.68	4.26	-4.00
	(0.87)	(0.24)	(0.07)	(0.09)	(0.92)	(0.85)	(0.26)	(0.39)	(0.07)	(0.06)	(0.56)	(0.00)	(0.76)	(0.23)	(0.37)
1 month after	-1.53	3.17	-0.99	-0.16	3.91	8.83**	-0.56	-0.73	2.74	0.85	-6.65	-12.23	-1.23	1.25	3.35
	(0.32)	(0.13)	(0.75)	(0.94)	(0.18)	(0.04)	(0.68)	(0.78)	(0.40)	(0.68)	(0.14)	(0.31)	(0.56)	(0.71)	(0.29)
2 months after	-0.09	-0.76	2.36	2.11	1.46	5.96**	-0.60	-1.00	-1.99	0.70	-4.17	-16.40*	-2.09	1.00	-1.02
	(0.94)	(0.61)	(0.29)	(0.17)	(0.51)	(0.05)	(0.54)	(0.58)	(0.51)	(0.64)	(0.18)	(0.07)	(0.14)	(0.67)	(0.72)
3 months after	0.76	-0.98	1.41	2.09*	1.76	2.70	-1.28	-1.40	0.14	0.59	-5.68**	-4.13	-3.15***	1.22	-0.38
	(0.41)	(0.45)	(0.44)	(0.09)	(0.33)	(0.28)	(0.11)	(0.32)	(0.95)	(0.64)	(0.02)	(0.59)	(0.00)	(0.50)	(0.85)
4 months after	0.63	-0.43	1.01	1.67	0.23	2.27	-1.09	-1.00	-1.05	0.51	-4.19**	-9.41	-2.79***	0.78	-1.39
	(0.43)	(0.70)	(0.52)	(0.12)	(0.88)	(0.29)	(0.11)	(0.41)	(0.60)	(0.64)	(0.05)	(0.18)	(0.00)	(0.62)	(0.46)
5 months after	0.55	-0.49	1.18	1.31	0.39	1.55	-1.05*	-0.82	-0.80	0.62	-3.92**	-11.38*	-2.62***	0.55	-0.98
	(0.45)	(0.64)	(0.39)	(0.16)	(0.78)	(0.42)	(0.09)	(0.45)	(0.65)	(0.53)	(0.03)	(0.09)	(0.00)	(0.69)	(0.56)
6 months after	0.48	-0.15	1.12	1.08	0.24	1.85	-0.86	0.29	2.22*	0.38	-2.21	-12.22**	-2.33***	1.40	1.76
	(0.46)	(0.87)	(0.37)	(0.21)	(0.85)	(0.29)	(0.13)	(0.77)	(0.09)	(0.67)	(0.20)	(0.04)	(0.00)	(0.25)	(0.17)

Table 3.5: Poisson Regressions: ADRs and Financial Crises

This table presents the estimates of the following Poisson regressions for 12 event windows of length $h \in [-6,6]$ except $h = 0$:$\ln \lambda_{nt} = \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \beta_3(h) adjexrret_{nt}(h) + \beta_4 adjexrret_{nt}(h) \cdot Crisis_t + \nu_{nt}(h)$ where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t; and $Crisis_t$ is a dummy variable equal to one if month t is within a financial crisis period and zero otherwise. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t+h\ ,t]$ when $h<0$), or as h-month after the observation month t (i.e., $[t,\ t+h]$ when $h>0$). To compute excess dollar exchange rate returns, we adjust for autocorrelation and time trends by estimating $exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}$, where $exrret_{nt}$ is the dollar exchange rate return of country n at month t. Then we compute the excess holding period currency return from month $t+h$ to month t, $adjexrret_{nt}(h_{<0})$, as $\sum_{s=t+h}^{t-1} \epsilon_{ns}$, and the excess holding period currency return from month t to month $t+h$, $adjexrret_{nt}(h_{>0})$, as $\sum_{s=t+1}^{t+h} \epsilon_{ns}$. Excess holding period local stock market returns, $adjmktret_t(h)$, and excess holding period U.S. stock returns, $adjusret_{nt}(h)$, are similarly defined. Crisis periods are defined as December 1994 to January 1995 for the Mexican Peso Crisis, July 1997 to November 1997 for the Asian Crisis, and August 1998 to January 1999 for the Russian Crisis. Panel A reports estimates for the whole sample; Panels B to F report estimates for country groups. p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_4(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_4(h)$
	Panel A: All Countries				**Panel B: G-7 Countries**			
6 months before	1.52***	1.01*	-1.37	-4.33	1.52***	0.30	-0.28	-1.84
	(0.00)	(0.09)	(0.13)	(0.38)	(0.01)	(0.74)	(0.82)	(0.83)
5 months before	1.66***	1.13*	-1.66*	-8.33	1.75***	0.06	-1.16	-4.93
	(0.00)	(0.09)	(0.10)	(0.13)	(0.01)	(0.95)	(0.40)	(0.58)
4 months before	1.53***	0.82	-1.31	-5.46	1.65**	-0.24	-0.70	-2.72
	(0.00)	(0.27)	(0.24)	(0.35)	(0.03)	(0.83)	(0.66)	(0.76)
3 months before	1.48***	0.74	-1.71	-6.65	1.46*	-0.56	-1.39	-6.47
	(0.01)	(0.39)	(0.20)	(0.27)	(0.10)	(0.66)	(0.46)	(0.45)
2 months before	1.70***	1.22	-2.16	-12.64**	1.67	0.37	-3.02	-11.63*
	(0.01)	(0.25)	(0.18)	(0.02)	(0.12)	(0.81)	(0.18)	(0.09)
1 month before	1.51*	1.83	-5.03**	-6.84	-0.20	2.53	-5.51*	-4.78
	(0.10)	(0.21)	(0.03)	(0.47)	(0.89)	(0.23)	(0.09)	(0.73)
1 month after	-0.62	1.94	-0.83	9.23***	-1.77	2.86	-2.30	24.17**
	(0.51)	(0.18)	(0.72)	(0.00)	(0.25)	(0.18)	(0.48)	(0.05)
2 months after	0.44	-0.29	0.92	5.23**	-0.15	-0.78	1.94	9.15
	(0.52)	(0.78)	(0.57)	(0.05)	(0.89)	(0.60)	(0.40)	(0.36)
3 months after	0.25	-0.29	1.01	4.19*	0.75	-0.97	1.32	1.95
	(0.64)	(0.73)	(0.44)	(0.07)	(0.42)	(0.45)	(0.48)	(0.82)
4 months after	0.15	-0.22	0.86	2.44	0.64	-0.43	1.04	-0.56
	(0.75)	(0.76)	(0.44)	(0.34)	(0.43)	(0.70)	(0.51)	(0.94)
5 months after	0.06	-0.19	0.93	1.86	0.54	-0.51	1.07	2.84
	(0.89)	(0.77)	(0.35)	(0.45)	(0.45)	(0.62)	(0.45)	(0.66)
6 months after	-0.07	0.21	1.99**	1.11	0.47	-0.18	1.04	2.22
	(0.85)	(0.72)	(0.02)	(0.63)	(0.47)	(0.85)	(0.41)	(0.71)

continued on next page

Table 3.5 (continued)

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_4(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_4(h)$
	Panel C: Other Developed				**Panel D: Emerging Markets**			
6 months before	0.87	3.42***	0.49	9.99	1.31**	-0.41	-6.07***	-9.28*
	(0.31)	(0.01)	(0.79)	(0.11)	(0.02)	(0.71)	(0.00)	(0.09)
5 months before	1.37	3.31**	1.49	9.44	1.25**	0.19	-6.38***	-13.31**
	(0.14)	(0.02)	(0.45)	(0.24)	(0.03)	(0.88)	(0.01)	(0.03)
4 months before	1.82*	2.39	2.16	11.73	0.79	0.14	-7.06***	-9.99
	(0.09)	(0.12)	(0.33)	(0.23)	(0.23)	(0.92)	(0.01)	(0.17)
3 months before	3.06***	1.35	2.74	11.49	0.26	1.03	-8.28***	-7.10
	(0.01)	(0.46)	(0.28)	(0.34)	(0.74)	(0.52)	(0.01)	(0.38)
2 months before	3.48**	-0.39	4.53	9.03	0.42	2.83	-7.72**	-10.06
	(0.02)	(0.86)	(0.14)	(0.55)	(0.66)	(0.16)	(0.03)	(0.25)
1 month before	3.41*	0.11	-1.54	16.43	1.57	2.10	-6.91	-12.95
	(0.10)	(0.97)	(0.73)	(0.41)	(0.23)	(0.44)	(0.16)	(0.34)
1 month after	-0.18	3.80	8.50**	6.29	-0.40	-1.87	-10.19**	17.68***
	(0.93)	(0.19)	(0.05)	(0.71)	(0.77)	(0.48)	(0.04)	(0.00)
2 months after	2.13	1.37	5.59*	7.17	-0.39	-1.86	-8.49***	13.07***
	(0.16)	(0.53)	(0.07)	(0.55)	(0.69)	(0.32)	(0.01)	(0.00)
3 months after	2.09*	1.80	3.10	-8.06	-1.04	-2.28	-5.42*	9.61***
	(0.09)	(0.32)	(0.22)	(0.47)	(0.19)	(0.12)	(0.08)	(0.01)
4 months after	1.67	0.21	2.18	1.72	-1.02	-1.35	-3.35	5.64
	(0.12)	(0.89)	(0.32)	(0.85)	(0.14)	(0.28)	(0.20)	(0.11)
5 months after	1.34	0.31	1.33	4.13	-1.02*	-0.97	-1.71	2.79
	(0.16)	(0.82)	(0.51)	(0.58)	(0.10)	(0.37)	(0.43)	(0.42)
6 months after	1.11	0.16	1.66	3.58	-0.85	0.32	2.37	-0.58
	(0.20)	(0.90)	(0.36)	(0.60)	(0.13)	(0.74)	(0.11)	(0.84)
	Panel E: Emerging Asian				**Panel F: Emerging Latin**			
6 months before	2.92***	-2.99	0.63	-9.28	0.12	-0.10	-5.71***	-14.34**
	(0.00)	(0.12)	(0.91)	(0.47)	(0.89)	(0.95)	(0.01)	(0.03)
5 months before	2.96***	-1.64	-6.13	-5.23	-0.52	0.19	-5.49**	-18.78***
	(0.00)	(0.45)	(0.37)	(0.78)	(0.57)	(0.90)	(0.02)	(0.01)
4 months before	3.15***	-3.49	-18.10***	52.79*	-2.03**	0.45	-4.92**	-26.85***
	(0.00)	(0.13)	(0.01)	(0.07)	(0.04)	(0.80)	(0.05)	(0.00)
3 months before	2.04	-0.70	-19.76**	14.21	-2.33**	0.95	-6.46**	-20.01**
	(0.08)	(0.80)	(0.02)	(0.45)	(0.04)	(0.64)	(0.03)	(0.05)
2 months before	1.94	1.84	-25.26***	13.86	-2.41*	3.29	-5.75*	-19.52
	(0.14)	(0.61)	(0.01)	(0.47)	(0.09)	(0.20)	(0.09)	(0.13)
1 month before	3.18	-2.44	-31.46***	18.51	-0.56	3.77	-2.89	-29.05
	(0.08)	(0.62)	(0.00)	(0.49)	(0.80)	(0.29)	(0.52)	(0.18)
1 month after	1.01	-7.29	-21.85*	26.09*	-1.23	-0.06	-6.55	13.69**
	(0.62)	(0.11)	(0.06)	(0.08)	(0.56)	(0.99)	(0.19)	(0.02)
2 months after	0.82	-4.53	-23.80***	25.94**	-1.84	-0.54	-6.05*	11.32**
	(0.56)	(0.16)	(0.01)	(0.02)	(0.19)	(0.83)	(0.08)	(0.02)
3 months after	0.54	-5.59**	-12.99	15.80	-2.61**	-0.40	-4.51	8.02**
	(0.67)	(0.02)	(0.15)	(0.13)	(0.02)	(0.84)	(0.13)	(0.04)
4 months after	0.54	-4.09**	-12.32	10.13	-2.77***	-0.10	-3.62	7.77**
	(0.62)	(0.05)	(0.11)	(0.40)	(0.00)	(0.95)	(0.13)	(0.05)
5 months after	0.69	-3.80**	-14.71**	13.61	-2.57***	0.19	-1.73	3.74
	(0.48)	(0.04)	(0.03)	(0.21)	(0.00)	(0.89)	(0.37)	(0.34)
6 months after	0.46	-2.12	-14.46***	12.83	-2.34***	1.34	1.63	0.70
	(0.61)	(0.22)	(0.01)	(0.22)	(0.00)	(0.29)	(0.25)	(0.82)

Table 3.6: Poisson Regressions: ADRs and Market Integration

This table presents estimates of the following Poisson regressions for 12 event windows of length $h \in [-6, 6]$ except $h = 0$:$\ln \lambda_{nt} = \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \beta_3(h) adjexrret_{nt}(h) + \beta_4 INTEG_{nt} + \nu_{nt}(h)$, where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t; and $INTEG_{nt}$ is a dummy variable equal to one if, on month t, country n has already experienced a significant financial integration regime shift, according to the endogenous chronology reported in Bekaert, Harvey, and Lumsdaine (2002, Table 3), and zero otherwise. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t+h\ ,t]$ when $h < 0$), or as h-month after the observation month t (i.e., $[t,\ t+h]$ when $h > 0$). To compute excess dollar exchange rate returns, we adjust for autocorrelation and time trends by estimating $exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}$, where $exrret_{nt}$ is the dollar exchange rate return of country n at month t. Then we compute the excess holding period currency return from month $t+h$ to month t, $adjexrret_{nt}(h_{<0})$, as $\sum_{s=t+h}^{t-1} \epsilon_{ns}$, and the excess holding period currency return from month t to month $t+h$, $adjexrret_{nt}(h_{>0})$, as $\sum_{s=t+1}^{t+h} \epsilon_{ns}$. Excess holding period local stock returns, $adjmktret_t(h)$, and excess holding period U.S. stock returns, $adjusret_{nt}(h)$, are similarly defined. Panel A reports results for the whole sample; Panel B reports estimates for emerging market countries. p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_4(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_4(h)$
	Panel A: All Countries				**Panel B: Emerging Markets**			
6 months before	1.42***	1.12*	-1.41	0.44***	1.22**	-0.29	-7.10***	0.72**
	(0.00)	(0.06)	(0.11)	(0.00)	(0.02)	(0.79)	(0.00)	(0.04)
5 months before	1.54***	1.23*	-1.80*	0.44***	1.09*	0.32	-7.63***	0.71**
	(0.00)	(0.06)	(0.07)	(0.00)	(0.06)	(0.79)	(0.00)	(0.04)
4 months before	1.41***	0.93	-1.41	0.44***	0.72	0.17	-8.08***	0.70**
	(0.00)	(0.21)	(0.20)	(0.00)	(0.27)	(0.90)	(0.00)	(0.05)
3 months before	1.35***	0.85	-1.90	0.43***	0.23	1.10	-9.23***	0.68**
	(0.01)	(0.31)	(0.14)	(0.00)	(0.77)	(0.49)	(0.00)	(0.05)
2 months before	1.51**	1.28	-2.82*	0.43***	0.31	3.17	-8.99***	0.66*
	(0.02)	(0.22)	(0.08)	(0.00)	(0.74)	(0.11)	(0.00)	(0.06)
1 month before	1.34	1.89	-5.30**	0.43***	1.41	2.35	-8.39*	0.63*
	(0.13)	(0.19)	(0.02)	(0.00)	(0.28)	(0.39)	(0.07)	(0.07)
1 month after	-0.79	2.24	2.65	0.45***	-0.57	-0.72	2.40	0.65*
	(0.38)	(0.11)	(0.18)	(0.00)	(0.67)	(0.78)	(0.46)	(0.07)
2 months after	0.28	-0.05	2.09	0.45***	-0.63	-1.00	-2.36	0.67*
	(0.66)	(0.96)	(0.13)	(0.00)	(0.52)	(0.58)	(0.44)	(0.06)
3 months after	0.14	-0.08	1.84	0.45***	-1.29*	-1.40	-0.19	0.67*
	(0.79)	(0.92)	(0.11)	(0.00)	(0.10)	(0.32)	(0.93)	(0.06)
4 months after	0.12	-0.09	1.11	0.45***	-1.10*	-1.02	-1.40	0.69**
	(0.79)	(0.90)	(0.28)	(0.00)	(0.10)	(0.40)	(0.49)	(0.05)
5 months after	0.04	-0.07	1.05	0.45***	-1.07*	-0.84	-1.18	0.69*
	(0.91)	(0.91)	(0.25)	(0.00)	(0.08)	(0.43)	(0.51)	(0.05)
6 months after	-0.04	0.30	1.96**	0.45***	-0.86	0.26	1.89	0.63*
	(0.91)	(0.60)	(0.02)	(0.00)	(0.12)	(0.79)	(0.15)	(0.08)

Table 3.7: Economic Significance of Currency Market Timing Abilities

This table reports the average and total U.S. dollar amounts (in millions) foreign companies have saved by selling ADRs neither "too early" (by deferring) nor "too late" (by expediting), respectively. More specifically, we estimate these amounts in two steps. First, we compute the negative of the cumulative abnormal returns from 6 months before to 1 month before a CR ADR issue, $-\sum_{j=1}^{6}\delta_j$, reported in column (4), and the cumulative abnormal returns from 1 month after to 6 months after a CR ADR issue, $\sum_{j=-6}^{-1}\delta_j$, reported in column (6), using estimates for δ_j coefficients from the bottom left-hand panel of Figure 3.1. Their corresponding p-values (rounded to two decimal places) are in columns (5) and (7), respectively. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level. Then, we multiply the resulting estimates by the corresponding median CR ADR issue size (in millions of U.S. dollars in column (1), from Table 3.1) to obtain median saving amounts (reported in columns (8) and (9)), and by the corresponding total CR ADR volume (in column (3), again from Table 3.1) to obtain total saving amounts (reported in columns (10) and (11)), for the whole sample and for selected regional subsets. The total saving percentage (column 12) is the ratio of the sum of total savings by selling ADRs neither "too early" nor "too late" (i.e., column (10) + column (11)) and the total CR ADR volume (column (3)).

Region	Median CR Issue Size	Number of Issues	Total CR ADR Volume	Cumulated Abnormal Currency Returns				Median Saving by		Total Saving by		Total Saving %
				$-\sum_{j=1}^{6}\delta_j$	p-value	$\sum_{j=-6}^{-1}\delta_j$	p-value	Deferring	Expediting	Deferring	Expediting	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
All Countries	$68.20	218	$34,799.50	0.95%**	0.03	0.91%**	0.04	$0.65	$0.62	$330.26	$315.39	1.86%
G-7	$64.00	109	$15,693.90	0.30%	0.65	0.62%	0.35	$0.19	$0.40	$46.67	$96.92	0.91%
Other Dev.	$70.10	45	$4,873.30	0.60%	0.51	2.23%**	0.01	$0.42	$1.56	$29.20	$108.57	2.83%
Emerging	$73.45	64	$14,232.30	2.02%**	0.03	0.53%	0.56	$1.48	$0.39	$287.16	$76.14	2.55%
Emg. Asia	$395.80	21	$8,242.70	1.29%	0.25	-1.05%	0.38	$5.09	-$4.16	$106.02	-$86.63	0.24%
Emg. Latin	$62.60	39	$5,766.20	3.53%**	0.03	1.56%	0.33	$2.21	$0.98	$203.43	$89.84	5.09%

Table 3.8: Poisson Regression: ADRs Across Firm and Issue Size

This table presents the estimates of the Poisson regression model across the following four subsets of all firms issuing ADR in our sample: "BigBig" includes all large ADR issues (i.e., above the median relative ADR issue size) from large firms (i.e., above the median issuing firm size) in a country; "BigSmall" include all large ADR issues (i.e., above the median relative ADR issue size) from small firms (i.e., below the median issuing firm size) in a country; "SmallBig" includes all small ADR issues (i.e., below the median relative ADR issue size) from large firms (i.e., above the median issuing firm size) in a country; and "SmallSmall" includes all small ADR issues (i.e., below the median relative ADR issue size) from small firms (i.e., below the median issuing firm size) in a country. The Poisson regression models for 12 event windows of length $h \in [-6, 6]$ except $h = 0$: $\ln \lambda_{nt} = \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \beta_3(h) adjexrret_{nt}(h) + \nu_{nt}(h)$, where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; and $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t. Excess returns are comupted with detrended returns as in Table 3.6. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t+h\ ,t]$ when $h < 0$), or as h-month after the observation month t (i.e., $[t,\ t+h]$ when $h > 0$). p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
	Panel A: BigBig			**Panel B: BigSmall**		
6 months before	2.35**	-0.77	0.67	1.45**	1.71**	-4.33***
	(0.04)	(0.68)	(0.81)	(0.01)	(0.04)	(0.01)
5 months before	2.54**	0.59	0.28	1.34**	1.78**	-5.10***
	(0.04)	(0.78)	(0.93)	(0.02)	(0.05)	(0.00)
4 months before	1.17	0.44	1.34	0.97	1.40	-4.14***
	(0.43)	(0.85)	(0.69)	(0.14)	(0.17)	(0.01)
3 months before	1.80	-1.08	1.41	0.91	1.09	-4.88***
	(0.29)	(0.69)	(0.72)	(0.22)	(0.35)	(0.01)
2 months before	2.03	0.62	-0.41	1.07	2.19	-6.24***
	(0.32)	(0.85)	(0.93)	(0.24)	(0.13)	(0.01)
1 month before	2.57	-0.25	-0.62	0.54	3.24	-7.28**
	(0.36)	(0.96)	(0.93)	(0.67)	(0.11)	(0.02)
1 months after	-2.82	3.74	-2.20	-1.22	3.60*	5.18**
	(0.340)	(0.40)	(0.75)	(0.34)	(0.07)	(0.04)
2 months after	-2.38	4.74	0.39	0.13	-0.33	3.62**
	(0.26)	(0.15)	(0.93)	(0.89)	(0.81)	(0.05)
3 months after	-0.41	0.78	1.28	-0.36	1.51	3.58**
	(0.81)	(0.77)	(0.74)	(0.64)	(0.20)	(0.02)
4 months after	-0.59	2.93	2.8	-0.12	1.06	2.33*
	(0.69)	(0.21)	(0.35)	(0.85)	(0.30)	(0.09)
5 months after	-0.47	1.70	2.25	-0.56	0.95	2.23*
	(0.73)	(0.42)	(0.42)	(0.33)	(0.30)	(0.07)
6 months after	-0.49	2.21	1.13	-0.41	1.60**	4.41***
	(0.69)	(0.25)	(0.67)	(0.43)	(0.05)	(0.00)
	Panel C: BigSmall			**Panel D: SmallSmall**		
6 months before	1.37***	1.34	-1.25	2.22*	-1.23	0.07
	(0.01)	(0.12)	(0.33)	(0.10)	(0.57)	(0.98)
5 months before	1.36**	1.26	-1.30	3.02**	-2.02	0.61

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Table 3.8 (continued)

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
	(0.02)	(0.19)	(0.35)	(0.04)	(0.40)	(0.86)
4 months before	1.13*	1.13	-0.90	3.66**	-3.63	-3.43
	(0.09)	(0.29)	(0.57)	(0.02)	(0.17)	(0.41)
3 months before	1.42*	1.20	-2.72	3.11*	-3.15	-5.76
	(0.07)	(0.33)	(0.15)	(0.10)	(0.30)	(0.24)
2 months before	1.68*	1.30	-4.42**	4.12*	-3.7	-2.85
	(0.07)	(0.39)	(0.05)	(0.06)	(0.31)	(0.62)
1 month before	2.82**	1.45	-6.54**	2.97	-3.31	-7.09
	(0.03)	(0.49)	(0.04)	(0.36)	(0.52)	(0.38)
1 month after	0.78	1.83	0.73	-0.73	-4.65	1.57
	(0.56)	(0.37)	(0.81)	(0.83)	(0.35)	(0.84)
2 months after	1.94**	-0.78	0.60	-2.39	-1.73	-2.29
	(0.04)	(0.60)	(0.78)	(0.32)	(0.61)	(0.69)
3 months after	1.53**	-1.64	0.11	-2.59	-3.06	2.70
	(0.05)	(0.17)	(0.95)	(0.18)	(0.28)	(0.49)
4 months after	0.80	-1.21	-0.77	-1.60	-3.59	2.82
	(0.25)	(0.25)	(0.62)	(0.35)	(0.15)	(0.40)
5 months after	0.8	-0.95	1.02	-0.99*	-3.83	4.23
	(0.19)	(0.31)	(0.45)	(0.51)	(0.08)	(0.14)
6 months after	0.25	-0.77	4.50***	-0.25	-2.64	2.59
	(0.65)	(0.35)	(0.00)	(0.86)	(0.20)	(0.39)

Table 3.9: Poisson Regression: ADRs and Tobin's q

This table presents estimates of the Poisson regression model across the following two subsets of all firms issuing ADR in our sample: "High Tobin's q Firms" includes all ADR issues from firms with above median Tobin's q in a country, and "Low Tobin's q Firms" includes all ADR issues from firms with below median Tobin's q in a country. The Poisson regression models of 12 event window of length $h \in [-6, 6]$ except $h = 0$:$\ln\lambda_{nt} = \alpha(h) + \beta_1(h)adjmktret_{nt}(h) + \beta_2(h)adjusret_t(h) + \beta_3(h)adjexrret_{nt}(h) + \nu_{nt}(h)$, where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; and $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t+h, t]$ when $h < 0$), or as h-month after the observation month t (i.e., $[t, t+h]$ when $h > 0$). To compute excess dollar exchange rate returns, we adjust for autocorrelation and time trends by estimating $exrret_{nt} = \phi_{0n} + \phi_{1n}exrret_{nt-1} + \phi_{2n}exrret_{nt-2} + \phi_{3n}t + \epsilon_{nt}$, where $exrret_{nt}$ is the dollar exchange rate return of country n at month t. Then we compute the excess holding period currency return from month $t+h$ to month t, $adjexrret_{nt}(h_{<0})$, as $\sum_{s=t+h}^{t-1}\epsilon_{ns}$, and the excess holding period currency return from month t to month $t+h$, $adjexrret_{nt}(h_{>0})$, as $\sum_{s=t+1}^{t+h}\epsilon_{ns}$. Excess holding period local stock returns, $adjmktret_t(h)$, and excess holding period U.S. stock returns, $adjusret_{nt}(h)$, are similarly defined. p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
Event Window	**Panel A: High q Firms**			**Panel B: Low q Firms**		
6 months before	1.37***	2.02***	-2.22*	1.69***	0.34	-2.48**
	(0.00)	(0.01)	(0.06)	(0.00)	(0.66)	(0.03)
5 months before	1.29**	2.18***	-2.20*	1.80***	0.29	-3.20***
	(0.02)	(0.01)	(0.09)	(0.00)	(0.73)	(0.01)
4 months before	0.93	1.78*	-1.75	1.47***	0.03	-2.77*
	(0.13)	(0.06)	(0.23)	(0.01)	(0.98)	(0.06)
3 months before	0.94	2.00*	-1.99	1.70***	-0.59	-5.00***
	(0.19)	(0.07)	(0.24)	(0.01)	(0.59)	(0.00)
2 months before	1.26	2.84**	-2.94	1.91**	-0.15	-6.59**
	(0.14)	(0.04)	(0.15)	(0.02)	(0.91)	(0.00)
1 month before	2.57**	2.61	-4.14	1.00	1.03	-8.55***
	(0.03)	(0.17)	(0.16)	(0.40)	(0.58)	(0.00)
1 month after	-0.47	1.13	2.25	-0.58	3.54*	3.16
	(0.70)	(0.54)	(0.40)	(0.63)	(0.06)	(0.23)
2 months after	0.76	-1.67	1.78	0.25	1.24	1.92
	(0.39)	(0.20)	(0.35)	(0.78)	(0.36)	(0.31)
3 months after	0.27	-0.98	1.97	0.2	0.61	2.11
	(0.71)	(0.37)	(0.20)	(0.78)	(0.58)	(0.16)
4 months after	-0.19	-0.72	1.17	0.37	0.65	1.26
	(0.77)	(0.44)	(0.38)	(0.55)	(0.50)	(0.35)
5 months after	-0.15	-0.34	1.73	0.07	0.16	2.08*
	(0.78)	(0.69)	(0.14)	(0.89)	(0.85)	(0.07)
6 months after	-0.35	0.39	4.16***	0.05	0.36	4.02***
	(0.48)	(0.61)	(0.00)	(0.92)	(0.64)	(0.00)

Table 3.10: Poisson Regressions: ADRs Across Industries

This table presents estimates of the Poisson regression model for six major industry groups across all ADR issuing firms in our sample. We use SIC codes to classify firms into 8 industries: Agriculture, Mining, Manufacturing, Utility, Sales, Financial, Construction, and Service. The Poisson regression models of 12 event windows of length $h \in [-6, 6]$ except $h = 0$: $\ln \lambda_{nt} = \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \beta_3(h) adjexrret_{nt}(h) + \nu_{nt}(h)$, where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t + h\ ,t]$ when $h < 0$), or as h-month after the observation month t (i.e., $[t,\ t + h]$ when $h > 0$). To compute excess dollar exchange rate returns, we adjust for autocorrelation and time trends by estimating $exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}$, where $exrret_{nt}$ is the dollar exchange rate return of country n at month t. Then we compute the excess holding period currency return from month $t + h$ to month t, $adjexrret_{nt}(h_{<0})$, as $\sum_{s=t+h}^{t-1} \epsilon_{ns}$, and the excess holding period currency return from month t to month $t + h$, $adjexrret_{nt}(h_{>0})$, as $\sum_{s=t+1}^{t+h} \epsilon_{ns}$. Excess holding period local stock returns, $adjmktret_t(h)$, and excess holding period U.S. stock returns, $adjusret_{nt}(h)$, are similarly defined. The Poisson model could not be estimated for Agriculture and Construction industries since less than 5 ADR issues were available for each. We also report number of ADR issuances within each industry, in parentheses next to the corresponding industry. p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
	Panel A: Mining (20)			**Panel B: Manufacturing (147)**		
6 months before	0.62	1.15	-0.53	1.77**	0.21	-0.85
	(0.70)	(0.65)	(0.88)	(0.00)	(0.81)	(0.50)
5 months before	0.18	1.44	-3.56	2.03***	-0.08	-1.18
	(0.92)	(0.60)	(0.39)	(0.00)	(0.94)	(0.41)
4 months before	-0.64	1.29	-4.07	2.02***	-0.53	-1.06
	(1.29)	(0.67)	(0.39)	(0.00)	(0.62)	(0.51)
3 months before	-0.87	1.66	-3.95	2.10***	-0.85	-1.57
	(0.71)	(0.64)	(0.47)	(0.01)	(0.49)	(0.40)
2 months before	0.75	1.11	-1.26	2.13***	-0.01	-1.97
	(0.79)	(0.80)	(0.85)	(0.02)	(0.99)	(0.38)
1 month before	-1.26	3.18	-2.03	1.64	0.21	-6.46**
	(0.76)	(0.60)	(0.83)	(0.21)	(0.92)	(0.05)
1 month after	-1.90	2.96	2.25	-1.07	1.69	6.16***
	(0.64)	(0.63)	(0.80)	(0.43)	(0.41)	(0.01)
2 months after	-0.59	1.46	6.54	-0.72	0.01	2.24
	(0.84)	(0.74)	(0.22)	(0.46)	(0.99)	(0.27)
3 months after	0.72	0.31	5.33	-0.35	-1.07	3.50 **
	(0.76)	(0.93)	(0.25)	(0.65)	(0.37)	(0.02)
4 months after	1.88	1.13	4.01	-0.15	-1.37	2.65*
	(0.34)	(0.72)	(0.32)	(0.83)	(0.19)	(0.06)
5 months after	1.32	0.57	3.69	0.07	-1.26	2.26*
	(0.45)	(0.84)	(0.32)	(0.91)	(0.18)	(0.08)
6 months after	0.73	-1.13	4.22	0.04	-0.60	2.66**
	(0.65)	(0.64)	(0.20)	(0.95)	(0.48)	(0.02)
	Panel C: Utility (76)			**Panel D: Sales (13)**		

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Table 3.10 (continued)

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
6 months before	1.69**	1.54	-2.46	0.96	5.86*	-7.23*
	(0.04)	(0.24)	(0.20)	(0.67)	(0.08)	(0.10)
5 months before	1.85**	2.31	-2.23	1.51	7.79**	-6.68
	(0.04)	(0.12)	(0.29)	(0.50)	(0.03)	(0.17)
4 months before	1.42	2.14	-1.30	2.64	3.81	-9.47*
	(0.17)	(0.19)	(0.58)	(0.30)	(0.34)	(0.09)
3 months before	0.66	3.00	-2.66	3.62	2.73	-6.32
	(0.59)	(0.11)	(0.34)	(0.22)	(0.56)	(0.34)
2 months before	0.99	3.98*	-3.72	4.71	-0.97	-5.71
	(0.50)	(0.07)	(0.28)	(0.20)	(0.86)	(0.46)
1 month before	1.45	6.40**	-4.00	7.87	-4.59	-4.31
	(0.48)	(0.05)	(0.41)	(0.12)	(0.54)	(0.69)
1 month after	-0.35	-0.23	3.15	-1.88	1.27	-0.12
	(0.87)	(0.94)	(0.46)	(0.73)	(0.87)	(0.99)
2 months after	1.23	-1.87	4.67*	-1.64	2.54	1.88
	(0.41)	(0.40)	(0.09)	(0.68)	(0.64)	(0.78)
3 months after	0.99	-1.33	3.71	-4.19	9.81**	1.06
	(0.42)	(0.47)	(0.12)	(0.16)	(0.03)	(0.83)
4 months after	0.91	-0.20	3.08	-6.06***	6.81*	-0.97
	(0.39)	(0.90)	(0.14)	(0.01)	(0.07)	(0.81)
5 months after	0.57	0.17	3.27*	-6.20***	9.94***	-2.22
	(0.54)	(0.91)	(0.08)	(0.00)	(0.01)	(0.57)
6 months after	0.14	0.72	2.44	-6.17***	9.29***	-0.66
	(0.87)	(0.58)	(0.16)	(0.00)	(0.01)	(0.85)
	Panel E: Financial (42)			**Panel F: Service (50)**		
6 months before	1.51	0.70	-1.13	1.51	1.67	-1.61
	(0.17)	(0.69)	(0.66)	(0.18)	(0.30)	(0.48)
5 months before	1.38	0.49	-1.77	1.59	1.14	-2.12
	(0.25)	(0.80)	(0.54)	(0.20)	(0.52)	(0.41)
4 months before	0.91	0.26	-0.27	1.92	1.58	-1.99
	(0.51)	(0.90)	(0.93)	(0.17)	(0.43)	(0.49)
3 months before	1.04	-0.03	-0.93	1.90	1.23	-3.22
	(0.52)	(0.99)	(0.80)	(0.24)	(0.59)	(0.35)
2 months before	0.77	-0.37	-5.10	2.68	1.15	-4.38
	(0.69)	(0.90)	(0.27)	(0.17)	(0.68)	(0.29)
1 month before	0.47	-0.68	-2.70	0.24	2.99	-8.82
	(0.86)	(0.87)	(0.67)	(0.93)	(0.44)	(0.13)
1 month after	-1.26	3.82	-5.37	-0.40	8.29**	1.72
	(0.64)	(0.36)	(0.41)	(0.88)	(0.03)	(0.76)
2 months after	1.42	-0.07	-1.23	1.25	2.16	2.87
	(0.47)	(0.98)	(0.78)	(0.54)	(0.45)	(0.50)
3 months after	0.53	-0.40	-1.66	1.33	2.18	2.43
	(0.75)	(0.87)	(0.66)	(0.43)	(0.38)	(0.44)
4 months after	-0.78	0.42	-2.19	1.69	0.70	1.25
	(0.58)	(0.84)	(0.49)	(0.24)	(0.73)	(0.65)
5 months after	-0.90	0.41	-2.01	1.49	-0.17	0.88
	(0.48)	(0.83)	(0.47)	(0.25)	(0.93)	(0.73)
6 months after	-0.26	0.09	1.15	1.06	1.39	3.28
	(0.82)	(0.96)	(0.63)	(0.37)	(0.41)	(0.13)

Table 3.11: Poisson Regressions: ADRs Across Underwriters

This table presents estimates of the Poisson regression model for the subsets of ADRs underwritten by each of the top six major ADR underwriting investment banks during our sample period: Credit Suisse First Boston (CFSB), Goldman Sachs, Lehman Brothers, Merrill Lynch, Morgan Stanley, and Salomon Smith Barney. The Poisson regression models of 12 event windows of length $h \in [-6, 6]$ except $h = 0$:$\ln \lambda_{nt} = \alpha(h) + \beta_1(h) adjmktret_{nt}(h) + \beta_2(h) adjusret_t(h) + \beta_3(h) adjexrret_{nt}(h) + \nu_{nt}(h)$, where the number of ADR issues from country n in month t follows a Poisson distribution, $Poisson(\lambda_{nt})$; $adjmktret_{nt}(h)$ is the excess holding period local stock market return of country n in month t for an event window h; $adjusret_t(h)$ is the excess holding period U.S. stock market return in month t for an event window h; $adjexrret_{nt}(h)$ is the excess holding period dollar exchange rate return of country n at month t. An event window is defined either as $|h|$-month before the observation month t (i.e., $[t+h, t]$ when $h < 0$), or as h-month after the observation month t (i.e., $[t, t+h]$ when $h > 0$). To compute excess dollar exchange rate returns, we adjust for autocorrelation and time trends by estimating $exrret_{nt} = \phi_{0n} + \phi_{1n} exrret_{nt-1} + \phi_{2n} exrret_{nt-2} + \phi_{3n} t + \epsilon_{nt}$, where $exrret_{nt}$ is the dollar exchange rate return of country n at month t. Then we compute the excess holding period currency return from month $t+h$ to month t, $adjexrret_{nt}(h_{<0})$, as $\sum_{s=t+h}^{t-1} \epsilon_{ns}$, and the excess holding period currency return from month t to month $t+h$, $adjexrret_{nt}(h_{>0})$, as $\sum_{s=t+1}^{t+h} \epsilon_{ns}$. Excess holding period local stock returns, $adjmktret_t(h)$, and excess holding period U.S. stock returns, $adjusret_{nt}(h)$, are similarly defined. We also report number of ADR underwritten by each investment bank, in parentheses next to the corresponding investment bank. p-values (rounded to two decimal places) are in parentheses. (*), (**), and (***) indicate the estimate is significant at the 10%, 5%, and 1% level, respectively.

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
	CSFB (18)			**Goldman Sachs (81)**		
6 months before	0.67	1.90	3.43	2.08***	0.31	-1.31
	(0.68)	(0.47)	(0.36)	(0.01)	(0.80)	(0.48)
5 months before	0.20	1.46	2.77	2.83***	0.16	-1.73
	(0.91)	(0.62)	(0.50)	(0.00)	(0.91)	(0.40)
4 months before	-0.65	1.30	2.01	2.61***	1.55	-1.64
	(0.75)	(0.69)	(0.67)	(0.01)	(0.32)	(0.48)
3 months before	-2.31	2.69	-2.52	2.60**	1.32	-2.23
	(0.34)	(0.47)	(0.67)	(0.02)	(0.46)	(0.41)
2 months before	-1.94	0.20	-4.70	3.00**	1.46	-3.98
	(0.51)	(0.96)	(0.52)	(0.02)	(0.51)	(0.23)
1 month before	-1.68	8.45	3.59	3.82**	1.00	-3.75
	(0.68)	(0.18)	(0.69)	(0.03)	(0.74)	(0.42)
1 month after	-6.68*	8.56	0.63	-1.07	1.94	3.67
	(0.10)	(0.16)	(0.95)	(0.58)	(0.51)	(0.37)
2 months after	0.87	-0.62	2.64	0.99	-0.31	4.46*
	(0.76)	(0.89)	(0.70)	(0.48)	(0.89)	(0.10)
3 months after	-1.46	0.01	-2.95	0.09	-0.92	3.64
	(0.55)	(1.00)	(0.62)	(0.94)	(0.60)	(0.11)
4 months after	-1.45	2.23	0.90	0.08	-0.57	1.85
	(0.49)	(0.49)	(0.85)	(0.94)	(0.71)	(0.38)
5 months after	-1.94	2.24	1.17	0.04	-0.06	1.99
	(0.30)	(0.44)	(0.78)	(0.97)	(0.97)	(0.29)
6 months after	-1.93	3.01	0.37	-0.24	0.26	2.66
	(0.26)	(0.27)	(0.92)	(0.77)	(0.83)	(0.11)
	Lehman Brothers (21)			**Merrill Lynch (55)**		
6 months before	1.48	0.20	0.65	1.53*	-0.25	-4.31*

continued on next page

Table 3.11 (continued)

Event Window	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$	$\beta_1(h)$	$\beta_2(h)$	$\beta_3(h)$
	(0.29)	(0.94)	(0.85)	(0.09)	(0.87)	(0.07)
5 months before	1.79	2.59	-0.96	2.16**	-0.53	-3.72
	(0.23)	(0.35)	(0.81)	(0.02)	(0.75)	(0.15)
4 months before	1.62	2.56	-1.66	1.94*	-1.11	-2.66
	(0.33)	(0.41)	(0.71)	(0.08)	(0.54)	(0.36)
3 months before	2.14	3.84	-2.38	2.61**	-1.88	-2.69
	(0.26)	(0.28)	(0.65)	(0.03)	(0.37)	(0.43)
2 months before	1.58	6.90	1.40	3.20**	-2.26	-6.25
	(0.50)	(0.13)	(0.81)	(0.03)	(0.37)	(0.13)
1 month before	2.84	2.82	2.11	4.54**	-3.16	-11.04
	(0.38)	(0.64)	(0.80)	(0.02)	(0.37)	(0.05)
1 month after	-1.48	5.78	-1.69	1.89	0.03	-6.92
	(0.68)	(0.34)	(0.85)	(0.39)	(0.99)	(0.23)
2 months after	0.25	-2.08	-3.38	1.66	-2.88	-2.28
	(0.92)	(0.61)	(0.60)	(0.30)	(0.24)	(0.57)
3 months after	-0.30	0.18	-2.30	0.84	-1.11	-3.40
	(0.89)	(0.96)	(0.66)	(0.52)	(0.59)	(0.31)
4 months after	-1.06	1.36	-1.95	0.89	-1.89	-2.69
	(0.55)	(0.65)	(0.66)	(0.44)	(0.29)	(0.35)
5 months after	-0.95	0.77	-1.22	0.91	-2.68*	-2.48
	(0.56)	(0.77)	(0.75)	(0.37)	(0.09)	(0.33)
6 months after	-0.57	0.70	-2.56	0.66	-1.90	-2.48
	(0.70)	(0.77)	(0.49)	(0.48)	(0.18)	(0.29)
	Morgan Stanley (55)			**Salomon Smith Barney (21)**		
6 months before	2.20**	0.34	0.01	1.24	-0.17	-0.83
	(0.02)	(0.82)	(0.99)	(0.38)	(0.95)	(0.81)
5 months before	1.55	0.49	0.75	0.97	1.43	-1.42
	(0.13)	(0.77)	(0.75)	(0.53)	(0.60)	(0.71)
4 months before	1.69	0.49	2.28	1.90	-1.66	-4.82
	(0.15)	(0.80)	(0.36)	(0.27)	(0.58)	(0.29)
3 months before	0.97	0.25	3.64	1.62	0.74	-8.96*
	(0.48)	(0.91)	(0.18)	(0.41)	(0.84)	(0.10)
2 months before	1.92	0.07	2.50	-0.49	3.03	-8.27
	(0.25)	(0.98)	(0.48)	(0.85)	(0.49)	(0.19)
1 month before	2.52	0.85	-2.88	-3.67	4.16	-11.64
	(0.27)	(0.82)	(0.61)	(0.31)	(0.48)	(0.19)
1 month after	-1.24	0.50	8.12**	4.73	-1.89	-16.60**
	(0.59)	(0.89)	(0.03)	(0.15)	(0.76)	(0.04)
2 months after	-0.59	1.53	3.50	2.49	-1.57	-10.25*
	(0.73)	(0.56)	(0.28)	(0.31)	(0.72)	(0.09)
3 months after	1.16	-0.46	6.60***	1.73	-4.29	-6.87
	(0.39)	(0.83)	(0.00)	(0.42)	(0.20)	(0.21)
4 months after	1.03	-0.16	4.49**	3.24**	-5.18*	-7.15
	(0.38)	(0.93)	(0.04)	(0.06)	(0.08)	(0.12)
5 months after	1.34	-0.21	3.27	2.95*	-3.91	-5.40
	(0.20)	(0.90)	(0.13)	(0.05)	(0.15)	(0.18)
6 months after	1.14	0.30	5.01***	2.67**	-2.58	0.02
	(0.22)	(0.84)	(0.01)	(0.05)	(0.29)	(1.00)

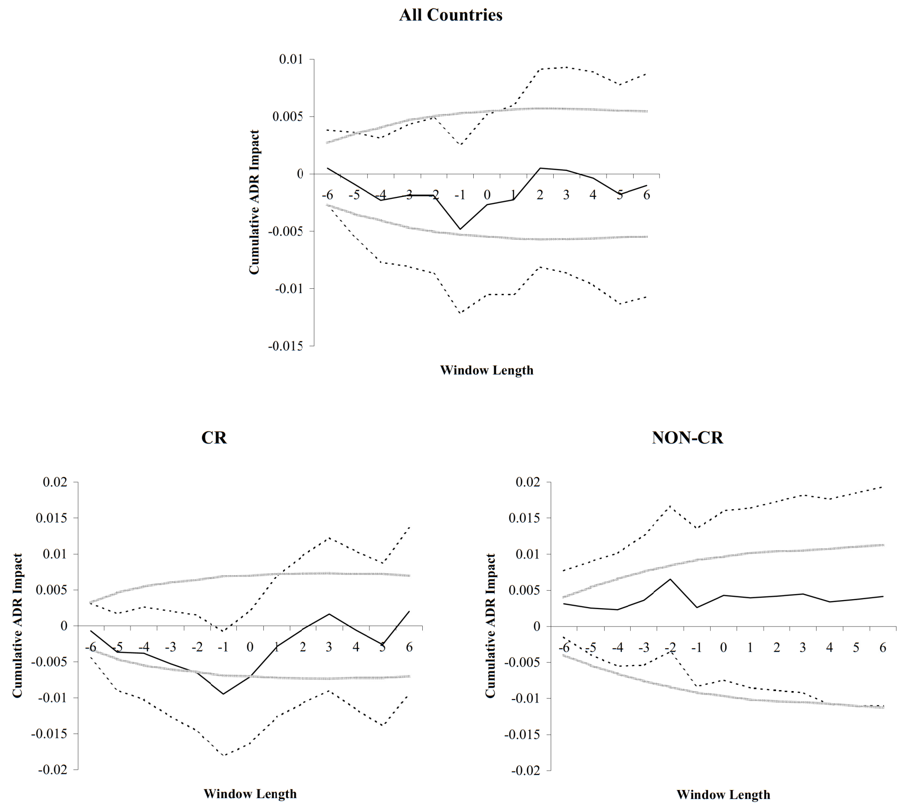

Figure 3.1: Cumulative Abnormal Exchange Rate Returns Around ADR Issues

This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue dates, i.e., the cumulative sum of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$ in the following equation:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return and $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise. Time 0 is the ADR issuance month. The estimated cumulative impact of ADR issues are plotted in solid lines. Their 95% confidence intervals are plotted in dotted lines. We also display 90% confidence intervals, centered around zero, constructed using bootstrapped standard errors. The bootstrap procedure consists of randomly drawing returns from the observed time series with replacement and estimating the aforementioned model with the resulting sample. After repeating this procedure 1000 times, we compute the corresponding standard errors. The plots are displayed for the whole ADR sample as well as two subsets of ADRs: Capital raising (Level III) ADRs and non-capital raising (Level II) ADRs, i.e., CR ADRs and non-CR ADRs, respectively.

Figure 3.2: Cumulative Abnormal Exchange Rate Returns Around ADR Issues: Regional Groups

This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue dates, i.e., the cumulative sum of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$ in the following equation:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return and $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise. Time 0 is the ADR issuance month. The estimated cumulative impact of ADR issues are plotted in solid lines. Their 95% confidence intervals are plotted in dotted lines.

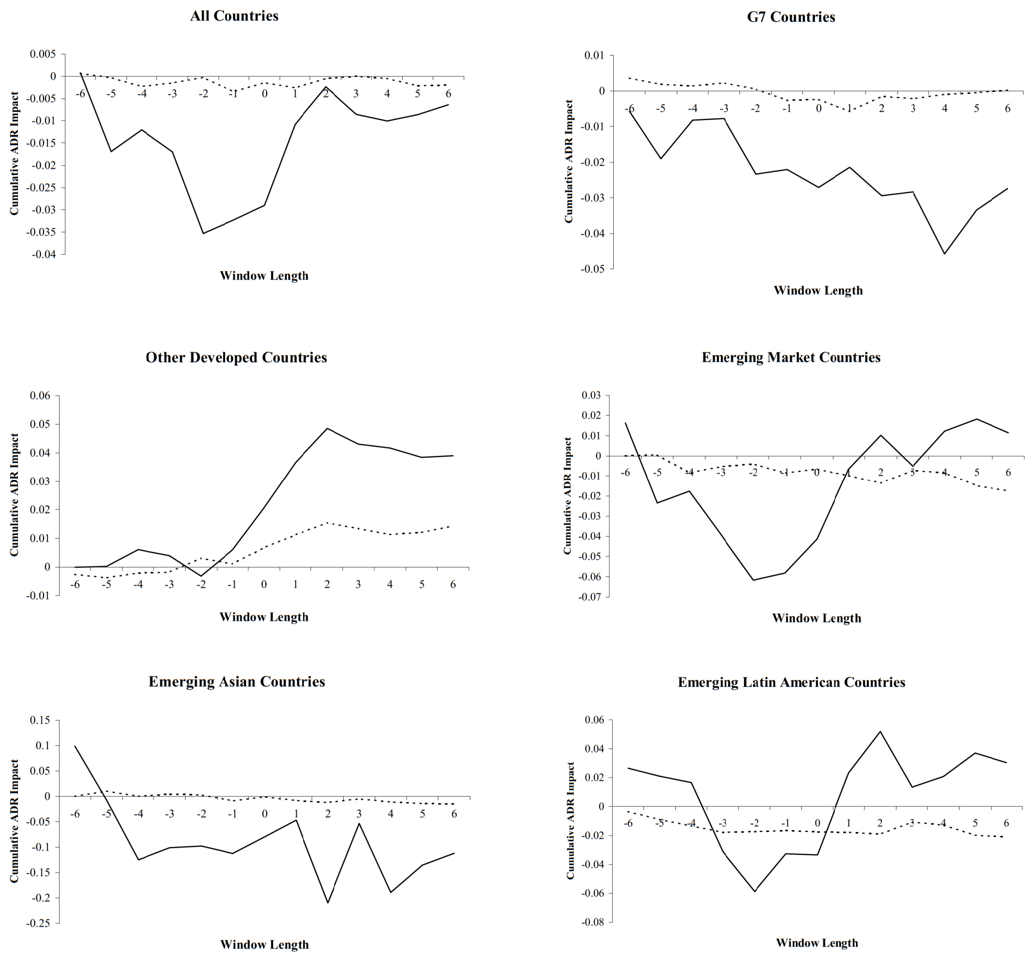

Figure 3.3: Cumulative Abnormal Exchange Rate Returns Around ADR Issues: Financial Crises

This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue dates. More specifically, it plots the cumulative sum of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$, i.e., estimates for the cumulative abnormal currency returns around ADR issues occurring over the portion of the sample period privy of financial crises (in dotted lines), and $\sum_{j\in[-6,6]}^{6} \delta_j + \delta_j^*$, i.e., estimates for the cumulative abnormal currency returns around ADR issues occurring during financial crises (in solid lines). Coefficients δ_j and δ_j^* are obtained from estimating the following regression:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \sum_{j=-6}^{6} \delta_j^* I_{nt}^*(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return, $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise, and $I_{nt}(j)$ is a dummy variable equal to 1 if there is at least one ADR issue in country n in month $t+j$ and month $t+j$ is considered a crisis period. Crisis periods are defined as December 1994 to January 1995 for the Mexican Peso Crisis, July 1997 to November 1997 for the Asian Crisis, and August 1998 to January 1999 for the Russian Crisis.

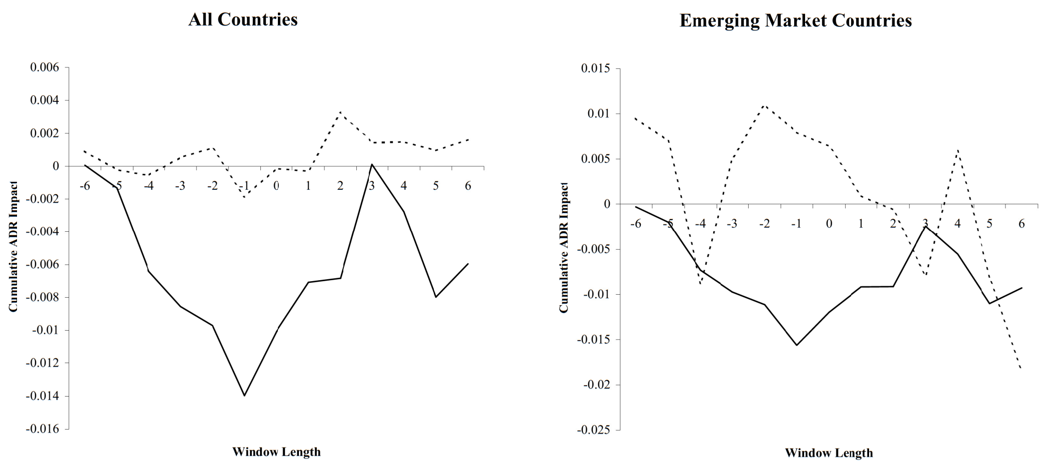

Figure 3.4: Cumulative Abnormal Exchange Rate Returns Around ADR Issues: Market Integration

This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue date for the whole sample and the subset of emerging markets. More specifically, it plots the cumulative sum of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$, i.e., estimates for the cumulative abnormal currency returns around ADR issues occurring over the portion of the sample period before market integration (in dotted lines), and $\sum_{j\in[-6,6]}^{6} \delta_j + \delta_j^I$, i.e., estimates for the cumulative abnormal currency returns around ADR issues occurring after market integration (in solid lines). Coefficients δ_j and δ_j^* are obtained from estimating the following regression:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \sum_{j=-6}^{6} \delta_j^I I_{nt}^I(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return, $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise, and I_{nt}^I is a dummy variable equal to one if at least one firm in country n issued ADR in month $t+j$ and month $t+j$ is past the endogenous financial integration date for country n estimated by Bekaert, Harvey, and Lumsdaine (2002, Table 3) and zero otherwise. Time 0 is the ADR issuance month.

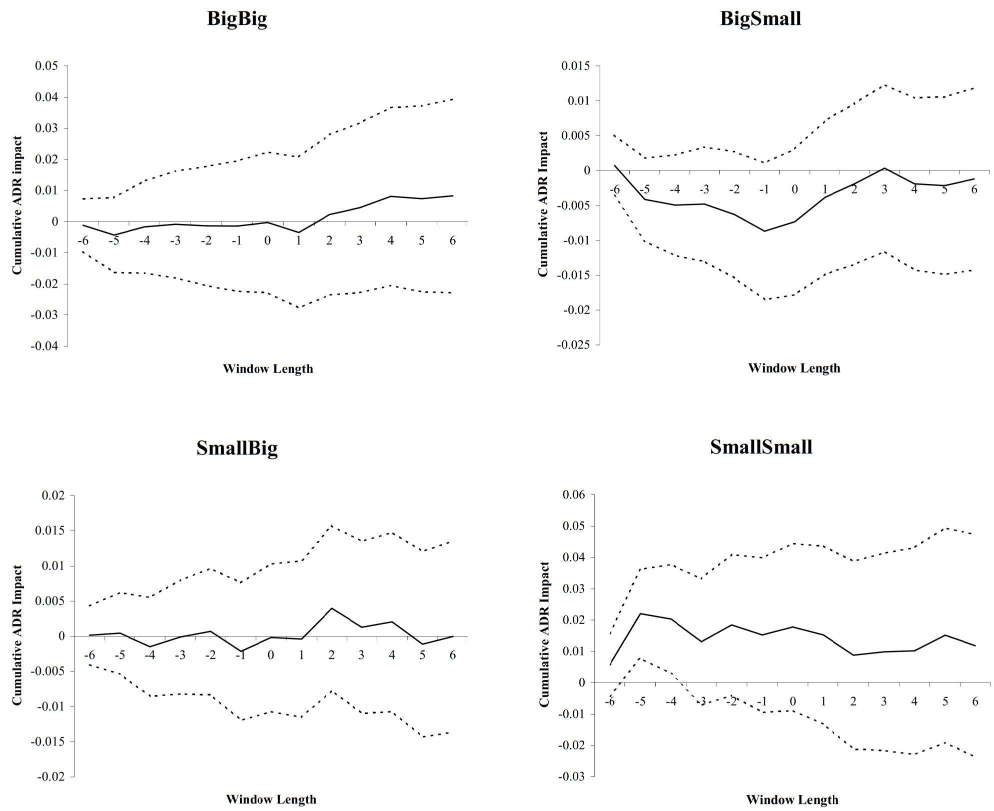

Figure 3.5: Cumulative Abnormal Exchange Rate Returns around ADR Issues: Firm and Issue Size
This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue dates across four subsets of all firms issuing ADR in our sample: "BigBig" includes all large ADR issues (i.e., above the median relative ADR issue size) from large firms (i.e., above the median issuing firm size) in a country; "BigSmall" include all large ADR issues (i.e., above the median relative ADR issue size) from small firms (i.e., below the median issuing firm size) in a country; "SmallBig" includes all small ADR issues (i.e., below the median relative ADR issue size) from large firms (i.e., above the median issuing firm size) in a country; and "SmallSmall" include all small ADR issues (i.e., below the median relative ADR issue size) from small firms (i.e., below the median issuing firm size) in a country. The cumulative abnormal impact of ADR issues is measured as the cumulative sums of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$ in the following regression:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return and $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise. Time 0 is the ADR issuance month. The estimated cumulative abnormal impact of ADR issues are plotted in solid lines. Their 95% confidence intervals are plotted in dotted lines.

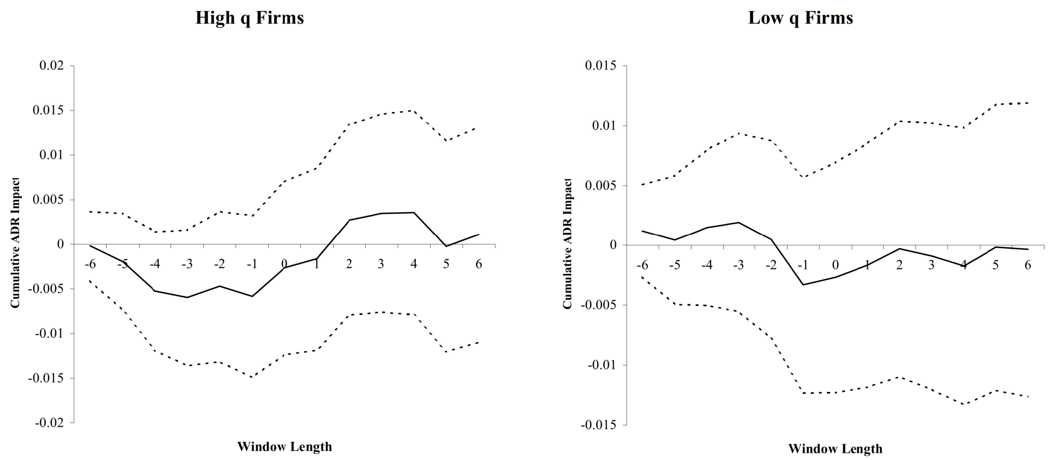

Figure 3.6: Cumulative Abnormal Exchange Rate Returns around ADR Issues: Tobin's q

This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue dates for two subsets of all firm issuing ADR in our sample: "High q Firms" includes all ADR issues from firms with above median Tobin's q in a country; and "Low q Firms" includes all ADR issues from firms with below median Tobin's q in a country. The cumulative abnormal impact of ADR issues is measured as the cumulative sum of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$ in the following regression:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return and $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise. Time 0 is the ADR issuance month. The estimated cumulative abnormal impact of ADR issues are plotted in solid lines. Their 95% confidence intervals are plotted in dotted lines.

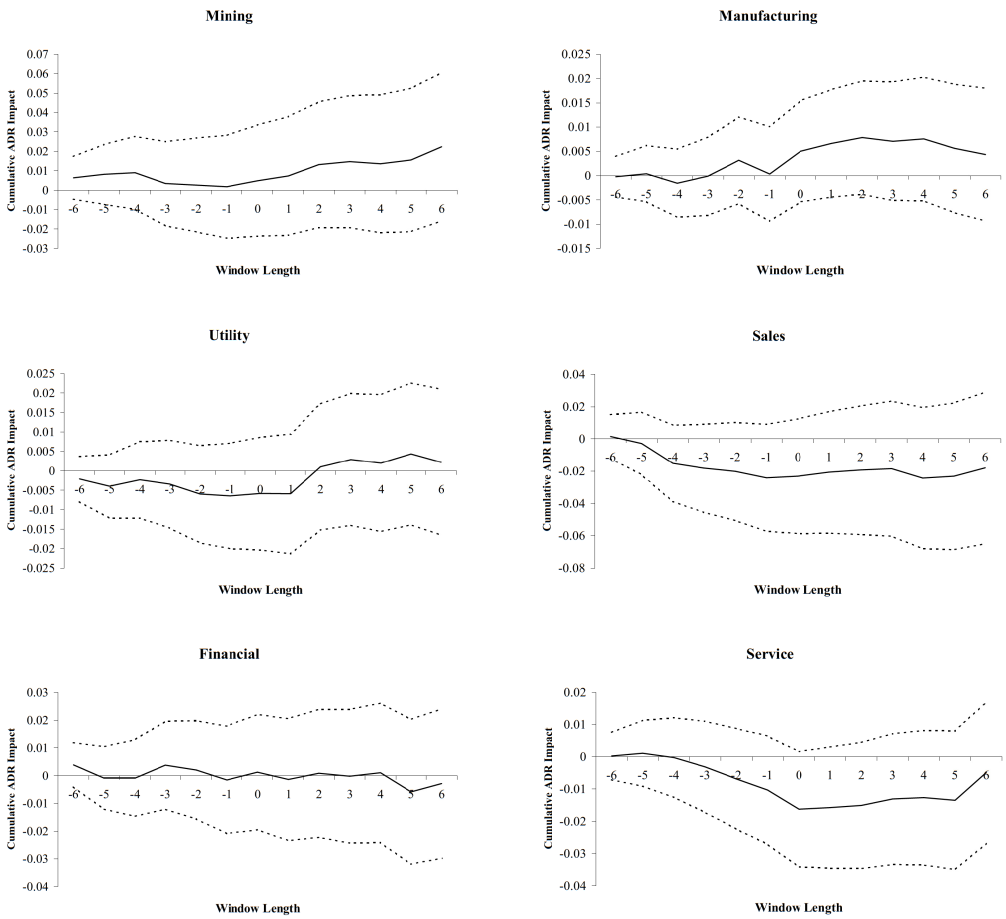

Figure 3.7: Cumulative Abnormal Exchange Rate Returns Around ADR Issues: Industries

This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue dates for six major industry groups across all ADR issuing firms in our sample. We use SIC codes to classify firms into eight industries: Agriculture, Mining, Manufacturing, Utility, Sales, Financial, Construction, and Service. The cumulative abnormal impact of ADR issues is measured as the cumulative sum of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$ in the following regression:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return and $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise. Time 0 is the ADR issuance month. The model could not be estimated for Agriculture and Construction industries since less than five ADR issues were available for each of them. The estimated cumulative abnormal impact of ADR issues are plotted in solid lines. Their 95% confidence intervals are plotted in dotted lines.

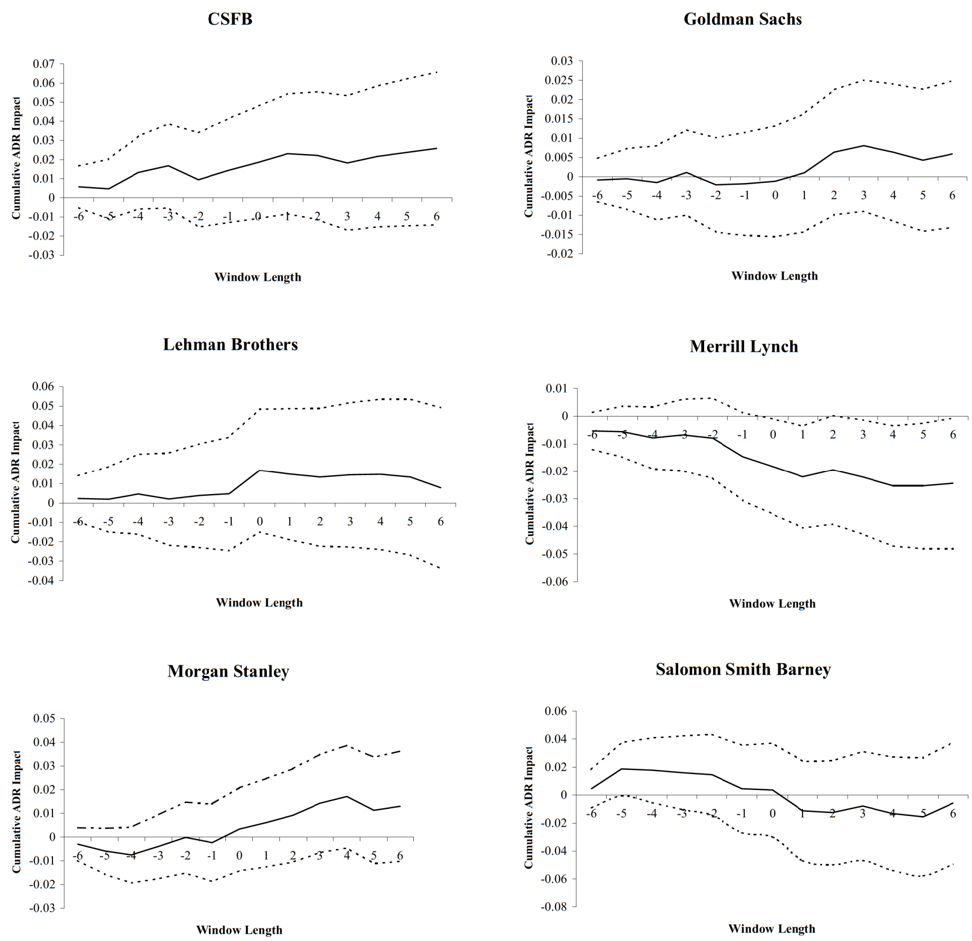

Figure 3.8: Cumulative Abnormal Exchange Rate Returns around ADR Issues: Underwriters

This figure plots the estimated cumulative abnormal impact of ADR issues on exchange rate returns around ADR issue dates for the subsets of ADRs underwritten by each of the top six major ADR underwriting investment banks during our sample period: Credit Suisse First Boston (CFSB), Goldman Sachs, Lehman Brothers, Merrill Lynch, Morgan Stanley, and Salomon Smith Barney. The cumulative abnormal impact of ADR issues is measured as the cumulative sum of estimated coefficients $\sum_{j\in[-6,6]}^{6} \delta_j$ in the following regression:

$$\epsilon_{nt} = \alpha + \sum_{j=-6}^{6} \delta_j I_{nt}(j) + \eta_{nt},$$

where ϵ_{nt} is the detrended exchange rate return and $I_{nt}(j)$ is a dummy variable equal to one if there is at least one ADR issue in country n in month $t+j$ and zero otherwise. Time 0 is the ADR issuance month. The estimated cumulative abnormal impact of ADR issues are plotted in solid lines. Their 95% confidence intervals are plotted in dotted lines.

CHAPTER IV

Are Investors Moonstruck? Lunar Phases and Stock Returns

"It is the very error of the moon. She comes more near the earth than she was wont. And makes men mad."
(Othello, Act V, Scene ii)

4.1 Introduction

The belief that phases of the moon affect mood and behavior dates back to ancient times. The lunar effect on the human body and mind is suggested anecdotally as well as empirically in the psychological and biological literature. Do lunar phases also affect the securities markets?

If investors make decisions strictly through rational maximization, then the answer is no. However, research evidence suggests that investors are subject to various psychological and behavioral biases when making investment decisions, such as loss-aversion, overconfidence, and mood fluctuation (for example, Harlow and Brown (1990); Odean, (1998), (1999)). On a general level, numerous psychological studies suggest that mood can affect human judgment and behavior (for example, Schwarz and Bless (1991) Frijda (1988)). The behavioral finance literature documents evidence on the effects of mood on asset prices (for example, Avery and Chevalier (1999); Kamstra et. al (2000), (2003); Hirshleifer and Shumway (2003); Coval and Shumway (2005)). If lunar phases affect mood, by extension, these phases may affect investor behavior and thus asset prices. If so, asset returns during full moon phases may be different from those during new moon phases. More specifically, since psychological studies associate full moon phases with depressed mood, this study hypothesizes that stocks are valued less and thus returns are lower during full moon periods.

This study is motivated by a psychological hypothesis. In modern societies the lunar cycle has little tangible impact on people's economic and social activities. Consequently, it would be difficult to find rational explanations for any correlation between lunar phases and stock returns. The causality would be obvious if there is such an effect. Therefore, investigating the lunar effect on stock returns is a strong test of whether investor behavior affects asset prices. Nevertheless, it is also important to recognize the possibility that the relation between lunar phases and stock returns could be spurious. As many researchers study the patterns of historical stock returns, some will find significant results simply due to chance.[1]

To investigate the relation between lunar phases and stock returns, we first examine the association of lunar phases with the returns of an equal-weighted and a value-weighted global portfolio of 48 country stock indices. The findings indicate that

[1]For example, Sullivan, Timmermann, and White (1999) argue that data snooping biases occur when a given set of data is used more than once for the purpose of model selection or inference. When such data reuse happens, there is always a possibility that results are due to chance rather than any merits inherit in the method. They quantify the data-snooping bias and adjust for its effect in the context of technical trading rules.

global stock returns are significantly lower during full moon periods than new moon periods. For the equal-weighted global portfolio, the cumulative return difference between the new moon periods and the full moon periods is 40.26 bps per lunar cycle for the 15-day window specification and 27.48 bps per lunar cycle for the 7-day window specification; both are significant at the 5% level. For the value-weighted global portfolio, the corresponding return difference is 30.44 bps for the 15-day window specification and 25.87 bps for the 7-day window specification, which are significant at the 10% and the 5% levels respectively. These numbers translate into annual return differences of 3% to 5%. The differences in the average daily logarithmic returns between the new and the full moon periods are consistent with the above findings.

A sinusoidal model is also estimated to test for the cyclical pattern of the lunar effect. According to this model, the lunar effect reaches its peak at the time of full moon and declines to a trough at the time of new moon, following a cosine curve with a period of 29.53 days (the mean length of a lunar cycle). The results indicate a significant cyclical lunar pattern in stock returns.

To fully utilize the panel data, a pooled regression was estimated with panel-corrected standard errors (PCSE) for all 48 countries and for the following subgroups of countries: the G-7 countries, the other developed countries, and the emerging-market countries. The PCSE specification adjusts for the contemporaneous correlation and heteroskedasticity among country index returns, as well as for the autocorrelation within each country's stock index returns. When all countries are included in the analysis, a statistically significant relation is found between moon phases and stock returns for both the 15-day and 7-day window specifications. Stock returns are, on average, 4 bps lower daily (about 5% annually) for the 15 days around the full moon than for the 15 days around the new moon. Using a 7-day window, stock returns are, on average, 6 bps lower daily (about 4% annually) on the full moon days than on the new moon days. The estimated effect remains similar when country group fixed effects are included. When country fixed effects are included, the estimated lunar effect becomes stronger. Another interesting observation is that the magnitude of this lunar effect is larger in the emerging market countries than in the developed countries.

To study the relation between the lunar effect and investor sentiment, we examine whether the lunar effect on stock returns is related to stock size, and thus individual versus institutional decision-making, since institutional ownership is higher for large cap stocks. Indeed, for U.S. stocks, we find evidence that the lunar effect is more pronounced for NASDAQ and small cap stocks than for NYSE-AMEX and large cap stocks.[2] Thus, the evidence suggests that the lunar effect is stronger for stocks that are held mostly by individuals. This finding is consistent with the notion that lunar phases affect individual moods, which in turn affect investment behavior.

To better understand the relation between lunar phases and stock markets, we investigate whether lunar phases relate to stock trading volumes and return volatility. No evidence is found that the lunar effect observed in stock returns is associated with trading volumes or risk differentials during the full moon and the new moon periods. We then examine whether the lunar effect can be explained by macro-economic events

[2]The exception is for the smallest size decile in NASDAQ stocks. Market microstructure and liquidity-related issues are more likely to have a significant impact on the pricing of extremely small stocks.

and other documented calendar anomalies. The findings indicate that the lunar effect is not due to the average effect of macro-economic announcements and the changes in short-term interest rates. Nor can the lunar effect be fully explained by global shocks. The lunar effect remains similar after we control for other calendar-related anomalies, such as the January effect, the day-of-week effect, the calendar month effect, and the holiday effects (including lunar holidays). Thus, we conclude that the lunar effect is unlikely a manifestation of these calendar anomalies. We further check the robustness of the lunar effect using various lunar window lengths, alternative ARIMA specifications, and a test of random 30-day cycles.

The remainder of the chapter is organized as follows. Section 2 discusses the literature on how lunar phases affect human mood and behavior. Section 3 describes the data. Section 4 discusses the methodology and results. Section 5 concludes.

4.2 Literature Review

One difficulty in testing whether psychological biases and sentiments affect investor trading behavior and asset prices is to find a proxy variable for sentiment or mood that is observable and exogenous to economic variables. Nonetheless, there have been several creative attempts. For example, Avery and Chevalier (1999) show that sentimental bettors can affect the path of prices in football betting. Saunders (1993) and Hirshleifer and Shumway (2003), drawing on psychological evidence that sunny weather is associated with an upbeat mood, find that sunshine is strongly correlated with stock returns. In a study of the seasonal-variation of risk premia in stock market returns, Kamstra et. al (2003) draw on a documented medical phenomenon, Seasonal Affective Disorder (SAD), to proxy investor mood and find a statistically significant relation between SAD and stock market returns. In another study, Kamstra, Kramer, and Levi (2000) relate yearly daylight fluctuations to mood changes and in turn to stock market returns.

This chapter exploits the popular perception that lunar phases affect mood and behavior, and analyzes the relation between lunar phases and stock returns. The hypothesis is that the lunar effect is an exogenous proxy for mood since lunar phases do not have tangible effects on economic and social activities. Furthermore, while the level of sunshine studied in Hirshleifer and Shumway (2003) is specific to geographical locations, lunar cycles are the same around the globe. Thus the lunar effect does not depend on the geographical locations of investors. Lunar cycles are also predictable. A relation between lunar cycles and stock returns will indicate that stock prices are predictable in a way uncorrelated with economic fundamentals, which is a strong violation of the efficient market hypothesis.

The idea that the moon affects individual moods has ancient roots. The moon has been associated with mental disorder since ancient time, as reflected by the word "lunacy," which derives from Luna, the Roman goddess of the moon. Popular belief has linked the full moon to such disparate events as epilepsy, somnambulism, crime, suicide, mental illness, disasters, accidents, birthrates, and fertility.

Biological evidence suggests that lunar phases have an impact on the human body and behavior. Research on biological rhythms documents a circatrigintan cycle, which is a moon-related human cycle. The most common monthly cycle is menstruation. A woman's menstrual cycle is about the same length as a lunar cycle, which suggests the influence of the moon. Law (1986) finds a synchronous relationship between the menstrual cycle and the lunar phases. Studies also find a lunar effect on fertility; for example, Criss and Marcum (1981) document that births vary systematically over lunar cycles with peak fertility during the third lunar quarter. In addition, lunar phases affect human nutrient intake: De Castro and Pearcey (1995) document an 8% increase in meal size and a 26% decrease in alcohol intake at the time of full moon.

Much attention has been paid to the lunar effect on human mood and behavior in the psychology literature. A recent study by Neal and Colledge (2000) documents an increase in general practice consultations during the full moon phase. Lieber (1978) and Tasso and Miller (1976) indicate a disproportionately higher number of criminal offenses occur during the full moon phase. Weiskott (1974) reports evidence that the number of crisis calls is higher during full moon and waning phases. Hicks-Caskey and Potter (1991) suggest an effect of the day of a full moon on the acting-out behavior of developmentally delayed, institutionalized women. Sands and Miller (1991) document that a full moon is associated with a significant but slight decrease in absenteeism after controlling for the effects of the day of the week, month, and proximity to a holiday.

Overall, the effect of the moon has been studied informally and formally for years. However, despite the attention this effect has received, psychological evidence for the lunar hypothesis in general is not conclusive even though biological evidence is strong. For example, in a review of empirical studies, Campbell and Beets (1978) conclude that lunar phases have little effect on psychiatric hospital admissions, suicides, or homicides. On the other hand, researchers argue that this lack of relation does not preclude a lunar effect. It may simply mean that the effect has not been adequately tested due to small sample sizes and short sample time periods (Cyr and Kaplan (1987); Garzino (1982)). Moreover, the psychology literature has mainly focused on trying to link the moon's phases to extreme behavioral problems in a few disturbed people, rather than a less drastic effect on human beings in general. By studying the relation between lunar phases and asset prices, this chapter extends the literature of the lunar effect on behavior.

In addition, survey evidence suggests a wide belief in the lunar effect. Rotton and Kelly (1985b) find that 49.4% of the respondents to their survey believe in lunar phenomena. Interestingly, among psychiatric nurses, this percentage rises to 74% (Agus (1973)). Vance (1995) reports a similar result. Danzl (1987) finds survey evidence that 80% of the nurses and 64% of the physicians in the emergency department believe that the moon cycle affects patients. Scientific explanations have been proposed to account for the moon's effect on the brain: sleep deprivation, heavy nocturnal dew, tidal effect, weather patterns, magnetism and polarization of the moon's light (Kelley (1942); Katzeff (1981); Szpir (1996); Raison, Klein, and Steckler (1999)).

Given the extensive documentation of the correlation between lunar phases and human feelings, thoughts, and behaviors, more specifically, the correlation between full moon periods and sleep deprivation, depressed mood, and suicidal events, the hypothesis in this study is that investors may value financial assets less during full moon periods than during new moon periods due to the changes in mood associated with lunar phases.[3]

This chapter is not the first attempt to link lunar phases to stock returns. Rotton and Kelly (1985a) cite a working paper by Rotton and Rosenberg (1984) that investigates the relation between lunar phases and Dow-Jones average closing prices. They find no significant relation between lunar phases and the Dow Jones Index prices.[4] The current study differs in that it examines returns rather than prices. In addition, heteroscedasticity and autocorrelations are corrected for in the return series, thus providing a more precise test for the relation. Most importantly, a broad sample of 48 countries is examined, which constitutes a more comprehensive and powerful test. Dichev and Janes (2003) also report a significant lunar effect on stock returns. Their study is concurrent with and independent of this study. Their findings and the findings of this chapter complement each other. Dichev and Janes (2003) focus more on the U.S. market, while this chapter provides global evidence on 48 countries with different levels of market development.

4.3 Data

A lunar calendar was obtained from the United States Naval Observatory (USNO) website.[5] This site provides the date and time (Greenwich Mean Time) of four phases of the moon for the time period of 1700 to 2015. The four phases are: new moon,

[3] We follow the evidence and argument in Hirshleifer and Shumway (2001) that good mood is associated with high asset returns. Since we assume that investors' mood follows a sinusoidal model and positive mood is associated with high asset returns, the hypothesis corresponds to a cycle in returns that meets its peak at new moon and its trough at the full moon.

[4] We were unable to obtain the working paper by Rotton and Rosenberg (1984). Our comments are based on the discussion provided in Rotton and Kelly (1985a).

[5] http://aa.usno.navy.mil/AA/

first quarter, full moon and last quarter. For the year 2000, the length of the mean synodic month (new moon to new moon) is 29.53059 days.

Stock market information on returns and trading volumes was obtained through Datastream. The sample period is from January 1973 to July 2001. The return sample consists of 48 countries listed in Morgan Stanley Capital International (MSCI) as developed markets or emerging markets. The country index calculated by Datastream (Datastream Total Market Index) was used unless a country did not have this series for at least five years; in these cases, the country index from other sources in Datastream was used. All returns were measured as nominal returns in local currencies. Trading volume data was collected for 40 of the corresponding 48 stock indices. Eight of these 48 indices did not have trading volume data in Datastream. Summary statistics appear in Table 4.1.

4.4 Empirical Findings

This section describes the empirical results of testing the hypothesis that stock returns are associated with lunar phases. We first report findings using an equal-weighted and a value-weighted global portfolio of the 48 country stock indices. This set of results indicates the significance of lunar effect on global stock returns.

Next, a panel regression was estimated using the entire panel of countries as well as panels of the following country categorizations: the G-7 countries, the other developed countries, and the emerging market countries.

To better understand the lunar effect on stock returns, we examine whether such an effect is related to stock capitalizations, patterns in trading volumes and stock market volatility. We also investigate whether the lunar effect is driven by macroeconomic announcements, global shocks, and calendar-related anomalies, such as the January effect, the day-of-week effect, the calendar month effect, the holiday effect, and the lunar holiday effect. Finally we check the robustness of the lunar effect to various lunar window lengths, several ARIMA specifications, and a test of random 30-day cycles.

4.4.1 Lunar Effect on the Global Portfolios

Since lunar cycles are common across the globe, we examine the lunar effect on an equal-weighted and a value-weighted global portfolio of 48 countries.[6] Specifically, we compare the returns of the full moon periods to the returns of the new moon periods for the global portfolios. Table 2 reports the test results. Panel A presents average cumulative returns; Panel B presents average daily logarithmic returns. The results indicate that the returns during the new moon periods are significantly higher than those during the full moon periods.[7] The effect is stronger for the 7-day window specification than for the 15-day window specification. The effect is also stronger for the equal-weighted portfolio than for the value-weighted portfolio. The findings are consistent regardless of using cumulative returns or daily logarithmic returns.

[6] At each point of time, we form the global portfolio using countries for which the return information is available.

[7] A full moon period is defined as N days before the full moon day + the full moon day + N days after the full moon day (N = 3 or 7). Similarly, a new moon period is defined as N days before the new moon day + the new moon day + N days after the new moon day (N = 3 or 7). In the case of the 15-day window, a new moon period can be less than 15 days since a lunar month may be less than 30 days. In these cases, the new moon period is defined as the remaining days of the lunar month.

In Table 4.2, Panel A, the cumulative return difference is -40.26 bps per lunar cycle for the 15-day window specification and -27.48 bps per lunar cycle for the 7-day window specification. A trading strategy with a long position in the portfolio during the new moon periods and a short position during the full moon periods on average yields a return of 40.26 bps for a lunar month using the 15-day window specification. A similar strategy using the 7-day window specification yields a return of 27.48 bps. These numbers translate into annual returns of 4.8% (40.26 bps *12) and 3.3% (27.48 bps *12) for the trading strategies, both significant at the 5% level. For the value-weighed portfolio, the corresponding return differences are -30.44 bps for the 15-day window and -25.87 bps for the 7-day window. These numbers translate into annual returns of 3.7% (30.44 bps *12) and 3.1% (25.87 bps *12) for the corresponding trading strategies using the value-weighted portfolio. To further gauge the economic significance, the transaction cost of implementing this trading strategy was estimated using exchange-traded-funds. The bid-ask spread for emerging market Exchange-Traded Funds (ETFs), such as iShares MSCI Emerging Markets ETF,[8] typically is around 0.10% of the traded value. Since the trading strategy used here involves 12 round-trip transactions (i.e., 24 transactions), a rough estimate of the transaction cost is 1.2%. Hence, the annual returns net of transaction costs for the trading strategies range from 1.9% to 3.6%.

Consistent with the evidence in Panel A, Panel B shows that the mean daily logarithmic returns are lower for the full moon periods than for the new moon periods.

[8] The ticker symbol of ishares MSCI Emerging Markets ETF is EEM and was incepted on April 7, 2003.

The average daily return difference for the equal-weighted portfolio is -3.53 bps for the 15-day window specification and -4.94 bps for the 7-day window specification. For the value-weighted portfolio, the average daily return difference is -2.80 bps for the 15-day window specification and -4.82 bps for the 7-day window specification. These numbers translate into annual returns of 4.4% (=3.53 bps *125) and 3.1% (=4.94 bps *62) for the corresponding trading strategies using the equal-weighted portfolio and annual returns of 3.5% (=2.80 bps *125) and 3.0% (=4.82 bps *62) for the trading strategies using the value-weighted portfolio respectively. Again, the lunar effect is stronger for the 7-day window specification and for the equal-weighted portfolio. Figure 1 plots the corresponding average daily logarithmic returns of the equal-weighted global portfolio for the full moon periods versus the new moon periods.

The documented return differences in both panels in Table 2 are statistically significant. The results are similar when alternative AR specifications of the Newey-West estimates are used. The bootstrapped p-values further confirm that the return differences are unlikely driven by the non-normality of the return distributions and by pure chance. The p-values of the nonparametric signed-rank test are all less than 5Next, a sinusoidal model of continuous lunar impact is used to test for the cyclical pattern of the lunar effect. According to the model, the lunar effect reaches its peak at the time of the full moon and declines to the trough at the time of the new moon, following a cosine curve with a period of 29.53 days (the mean length of a lunar cycle). The following regression is estimated for the portfolios:

$$R_t = \alpha + \beta \cos(2\pi d_t/29.53) + e_t, \tag{4.1}$$

where d is the number of days since the last full moon day and the β coefficient indicates the association between stock returns and lunar cycles. There is a negative relation (β = -2.88) between the global stock returns and lunar cycles. The test result is statistically significant at the 1% level (Figure 2). Overall, the sinusoidal model suggests that the lunar effect is cyclical.

In summary, we find global evidence of a significant correlation between stock returns and lunar phases. We document that on average returns are higher during the new moon periods than during the full moon periods.

4.4.2 Panel Analysis

A panel of country level average daily logarithmic returns for each lunar period is set up to fully utilize the cross-sectional and time series data. We estimate a pooled regression with panel corrected standard errors (PCSE) as the following:

$$R_{it} = \alpha_i + \beta Lunardummy_t + e_t, \tag{4.2}$$

R_it is the average daily logarithmic return during a full moon or a new moon period for country i at time t. *Lunardummy* is a dummy variable indicating a full moon or a new moon period; it takes on a value of one for a full moon period and zero for a new moon period.[9] The coefficient on this variable indicates the difference in the mean daily logarithmic returns between the lunar periods. The PCSE specification adjusts for the contemporaneous correlation and heteroskedasticity among country

[9]For the 7-day window specification, we only include days of a full moon period and a new moon period. Other days of a lunar month are excluded from the regression.

index returns, as well as for the autocorrelation within each country's stock index returns (Beck and Katz, 1995).

Table 4.3 presents estimation results of the pooled regression for both the 15-day and the 7-day window specifications. The results indicate that stock returns of the new moon periods are significantly higher than those of the full moon periods. Regardless of model specifications, the coefficients on *Lunardummy* are negative. When all countries are included in the analysis, the returns of the new moon periods are on average 3.95 bps and 5.93 bps higher than returns of the full moon periods for the 15-day and 7-day windows respectively. Both estimates are significant at the 5% level. The estimated coefficients on the *Lunardummy* remain similar when country group dummy variables are included. When country fixed effects are included, the estimated coefficients become larger in magnitude and higher in statistical significance. Interestingly, the extent of the lunar effect seems to vary across different levels of market maturity. The strongest lunar effect is found among the emerging market countries: a 5.60 bps daily difference for the 15-day window and an 11.27 bps daily difference for the 7-day window; both are significant at the 5% level. The lunar effect in the other developed markets and the G-7 countries is less strong: a 3.21 bps daily difference for the 15-day window and a 3.19 bps daily difference for the 7-day window for the other developed markets; and a 2.43 bps daily difference for the 15-day window and a 2.45 bps daily difference for the 7-day window for the G-7 countries[10]

In summary, the panel analysis confirms the earlier findings using the global port-

[10] R^2 statistics for these regressions are between 0.1% and 0.5%. The values are low but not very surprising since we do not expect that the lunar effect would explain a large proportion of variations in stock returns.

folio: Stock returns of the new moon periods are significantly higher than those of the full moon periods, more so for the emerging market countries. Maturity of the stock markets and the percentage of institutional investors might explain the differences in the magnitude of lunar impact in these markets.

4.4.3 Large vs. Small Capitalization Stocks

In this section, we examine whether the lunar effect is related to stock capitalization. This test is motivated by the empirical finding that institutional ownership is positively correlated with stock capitalization.[11] Specifically, large capitalization stocks have a higher percentage of institutional ownership than small capitalization stocks. Since investment decisions of individual investors are more likely to be affected by sentiments and mood than those of institutional investors, we expect the lunar effect to be more pronounced in the pricing of small capitalization stocks. To assess the relation between lunar phases and stock capitalization, 10 stock portfolios were formed based on market capitalization for stocks traded on the NYSE, AMEX and NASDAQ. Returns and market capitalization for the NYSE, AMEX, and NASDAQ stocks were obtained from CRSP.

Table 4.4 reports results of a regression of daily returns of market capitalization ranked portfolios on lunar phases. The estimated lunar effect is stronger for NASDAQ stocks than for NYSE and AMEX stocks. Moreover, the lunar effect is stronger for smaller size deciles with the exception of the smallest decile in NASDAQ.[12] The

[11] For example, see Sias and Starks (1997).

[12] Liquidity and market microstructure related issues are likely to have a first-order effect in pricing extreme small stocks rather than mood; hence, a weaker lunar effect for stocks that are extremely small in capitalization is not entirely surprising.

Spearman rank correlation is -0.81 for the NYSE and AMEX deciles and significant at the 1% level. The correlation is -0.65 for the NASDAQ deciles and significant at the 10% level (excluding the smallest decile).

Overall, these findings are consistent with the hypothesis that stocks with more individual investor ownership display a stronger lunar effect and thus provide further evidence that mood or sentiment may affect asset prices.

4.4.4 Trading Volume and Market Volatility

In order to determine whether the observed lunar effect is related to trading volumes and return volatility, we estimate the following regressions for an equal-weighted portfolio and a panel of 48 countries for the 15-day full moon window:

$$normvolume_{jt} = \alpha_j + \lambda_j Lunardummy_t + e_{jt}, \tag{4.3}$$

$$volatility_{jT} = \alpha_j + \lambda_j Lunardummy_T + e_{jT}, \tag{4.4}$$

where the variable, *normvolume*, is the daily trading volume normalized by average daily volume in the month and t is the time index for each day; the variable, *volatility*, is the standard deviation of daily logarithmic stock returns in a lunar period, and T is the time index for a lunar period.

Test results for Equation 4.3 are reported in Table 4.5, Panel A. The coefficient on *Lunardummy* is not significant for the portfolio, nor is it significant for the pooled regression of 48 countries, which indicates that there is little evidence that trading volumes are related to lunar phases in a systematic manner. Therefore, the observed lunar effect is not due to patterns in trading volume that are related to lunar phases. Test results for Equation 4.4 are reported in Table 4.5, Panel B. The coefficients on *Lunardummy* of the portfolios and the panel regression are of different signs and are both insignificant, which indicates that stock market volatilities are not related to lunar phases in a systematic manner. Hence, the observed lunar effect in stock returns cannot be explained by the risk differentials between the full moon and the new moon periods.

4.4.5 Macroeconomic Events

It is possible that the return differential between the full moon and the new moon periods reflects the average effect of macroeconomic events or common market shocks. In this section, we examine to what extent the estimated lunar effect is explained by macroeconomic events. Specifically, we investigate the lunar effect on the global portfolio, controlling for the following three types of events: macroeconomic announcements, major global shocks, and movements in short-term interest rates. We estimate the lunar effect on the daily logarithmic returns of the global portfolios, controlling for macro-economic events. Initially, a base case of lunar effect is identified by estimating the following regression for the portfolios at the daily frequency:

$$R_t = \alpha + \beta Lunardummy_t + e_t, \tag{4.5}$$

where R_t is the daily logarithmic return, and t is the time index of daily frequency. The Newey-West adjusted standard errors are used, assuming an AR1 process. The indicator variable, $Lunardummy_t$, is set equal to one if day t falls in a full moon period and zero if day t falls in a new moon period. Equation 4.5 is then re-estimated, controlling for macroeconomic announcements, major global shocks, and changes in short-term interest rates.

The test results are reported in Table 4.6. Model 1 reports the base case. The coefficient on *Lunardummy* is -4.26 (-5.40) bps for the equal-weighted portfolio and -3.47 (-4.77) bps for the value-weighted portfolio, for the 15-day (the 7-day) window specification. All estimates are significant at the 5% level. These results confirm the earlier finding that stock returns are higher during new moon periods than full moon periods.

To examine whether the return differences between the new moon and the full moon periods are due to macroeconomic announcements, two tests are performed. In the first test, Equation (5) is re-estimated by excluding the days with specific macroeconomic announcements: Consumer Price Index, Federal Reserve Open Market Committee announcements, Gross National Product, Retail Sales, Employment Report, Employment Cost Index, Trade Deficit, and National Association of Purchasing Managers Survey Index (Gerlach (2004)). The resulting estimates, reported in Table 6, Model 2, are similar to those of the base case for the 15-day window specification and larger in magnitude for the 7-day window specification, indicating that the lunar effect cannot be explained by the average effect of macroeconomic announcements. In the second test, the number of days with macroeconomic announcement during the new moon and the full moon phases is plotted to compare the distribution of announcements across the lunar periods. Figure 3 shows that macro-announcements occur quite evenly during the two periods for both the 15-day and the 7-day window specifications. Overall, the evidence indicates that the lunar effect is unlikely due to macroeconomic announcements.

Next, the relevance of global shocks to the lunar effect is examined. Equation 4.5 is re-estimated excluding various global shocks. As reported in Table 6, Model 3, the estimates are still negative, albeit smaller in magnitude and lower in statistical significance than the base case. The lunar effect remains negative and significant for the equal-weighted portfolio; the coefficients on *Lunardummy* are negative but not statistically significant for the value-weighted portfolio. Thus, excluding the periods of global shocks from the analysis weakens the lunar effect to some extent; however, these shocks cannot fully explain the documented lunar effect.

Finally, the lunar effect is examined by controlling for changes in the short-term interest rates. Equation 4.5 is modified by including short-term interest rates as an explanatory variable in the following regression:

$$R_t = \alpha + \beta Lunardummy_t + \delta shortterminterestrate_t + e_t, \tag{4.6}$$

where the *shortterminterestrate* is the three-month Treasury bill rate. The results are presented in Table 6, Model 4. The coefficients on the short-term interest rate are negative for all specifications, indicating that higher interest rates are correlated with lower stock prices. The estimated lunar effect remains similar although slightly weaker, which indicates that changes in short-term interest rates do not explain the observed lunar effect. Overall, the evidence indicates that the lunar effect cannot be explained away by macro-economic announcements, common shocks in the stock markets, and changes in short-term interest rates.

4.4.6 The Lunar Effect and Other Calendar Anomalies

This section examines whether the lunar effect can be explained by other calendar anomalies.

4.4.6.1 The January Effect

The lunar effect is unlikely a manifestation of the January effect,[13] since lunar months do not correspond to calendar months. Nevertheless, to test for the relation between the lunar effect and the January effect, a January dummy variable was added to the following regression:

$$R_t = \alpha + \beta Lunardummy_t + \delta januarydummy_t + e_t, \tag{4.7}$$

$Januarydummy$ is a dummy variable equal to one in the month of January and zero otherwise.

Table 7, Model 1, shows both a significant January effect and a significant lunar effect. Compared with the base case findings in Table 6, where the January effect is not controlled, the magnitude and significance of the lunar effect become only slightly

[13] The January effect has been documented by, for example, Rozeff and Jr. (1976) and Reinganum (1983)

smaller; thus, the January anomaly is not a driving force behind the observed lunar effect.

4.4.6.2 Day-of-Week Effect

If most full moon days fall on Mondays, it is possible that the Monday effect may explain the observed lunar effect. Figure 4 shows that the full moon days fall evenly on each day of the week in the sample. In an unreported panel regression of all countries, the estimated lunar effect becomes stronger when the day-of-week fixed effects are included; thus, the lunar effect on stock returns is not related to the day-of-week effect.

4.4.6.3 Calendar Month Effect

Ariel (1987) shows that the mean U.S. stock return for days during the first half of a calendar month is higher than the mean stock return during the second half of the month. Thus, it is conceivable that the lunar effect documented in this chapter may be a manifestation of the calendar month effect. To test for this possibility, a calendar dummy is added in the regression and Equation 4.5 is re-estimated, as follows:

$$R_t = \alpha + \beta Lunardummy_t + \delta Calendardummy_t + e_t, \tag{4.8}$$

$Calendardummy$ is a dummy variable equal to one for the first half of a calendar month and zero otherwise. As shown in Model 2 of Table reftable:lunar7, the calendar month effect is not significant for the portfolios; nevertheless, the magnitude and significance of the *Lunardummy* is consistent with the earlier results. Thus, the test statistics suggest that the calendar month effect cannot explain the observed lunar effect.

4.4.6.4 Holiday Effect

Ariel (1990) documents that, on the trading day prior to holidays, stocks advance with disproportionate frequency and show high mean returns averaging nine to 14 times the mean returns for the other days. To examine this effect, the day before a holiday is excluded for each country. Equation (5) is then re-estimated using the holiday-adjusted global index returns. As reported in Model 3 of Table 7, the lunar effect is significant at the 5% level for both portfolios and for both the 15-day and 7-day window specifications. Thus, the lunar effect does not appear to be driven by the holiday effect.

4.4.6.5 Lunar Holidays

Frieder and Subrahmanyam (2004) show that Jewish holidays have a significant impact on the U.S. equity market. They find that returns are significantly positive around Rosh Hashanah and significantly negative around Yom Kippur. Two tests are used to check the robustness of the lunar cycle effect: 1) lunar holiday dummy variables are added to Equation 4.5 because many Jewish, Islamic, Hindu, Chinese, Japanese, and Korean holidays fall on the fixed days of a lunar-based calendar (Table 7); and 2) lunar holiday dummy variables are added to the estimation of the lunar effect at the country level for relevant countries (Table 8).

Table 4.7 reports the estimates on the lunar dummy and the lunar holiday dummies for two Jewish holidays: Yom Kippur and Rosh Hashanah. The coefficients for these two holidays are not significant for the portfolios. The coefficient on the lunar dummy for the equal-weighted portfolio is -3.76 for the 15-day window specification and -4.33 for the 7-day window specification, which are significant at the 5% and 10% levels respectively. The coefficient on the lunar dummy for the value-weighted portfolio is -2.57 for the 15-day window specification and -3.54 for the 7-day window specification, both statistically insignificant.

Interestingly, in the country level analysis, the Jewish holiday dummies are statistically significant for the U.S. and the Israeli markets while the lunar holiday dummies for other countries are not significantly different from zero (except in Korea). These results are consistent with the findings for the U.S. stock market in Frieder and Subrahmanyam (2004). For both the U.S. and Israeli market, the returns are lower around Yom Kippur (a somber holiday) and higher around Rosh Hashanah (a cheerful holiday). Nevertheless, the coefficients on the lunar dummies do not change much when the lunar holiday dummies are included, which indicates that the lunar holiday effect is probably independent of the lunar cycle effect. Thus, the observed lunar effect is not likely just a manifestation of other documented calendar anomalies.

4.4.7 Additional Robustness Checks

The robustness of the lunar effect was further checked by examining various lunar window lengths, alternative ARIMA specifications, and a test of random 30-day cycles.

4.4.7.1 Lunar Window Length

To address the concern that the estimated lunar effect may be due to the choice of window length, Equation 4.5 is re-estimated for window lengths of one to 15 (Table 4.9). The stock returns are higher during the new moon phases than the full moon phases for all window lengths for both portfolios. Except for the one-day window, the p-values of all estimates are lower than the 10% level of significance. Consistent with the earlier findings, except for the three-day window, the lunar effect is stronger and more significant for the equal-weighted portfolio than for the value-weighted portfolio. Since the return differences are quite consistent across the window lengths, it is unlikely that the lunar effect is due to the choice of window length.

4.4.7.2 ARIMA

Different ARIMA specifications are used to adjust the returns of the portfolios. Equation 4.5 is then re-estimated (Table 4.10). Both the magnitude and the statistical significance of the estimates are consistent across the different specifications, indicating that the documented lunar effect is not due to the time-series properties of stock returns.

4.4.8 30-Day Cycle Effect

To test whether the observed lunar effect reflects a general pattern in stock returns, rather than a lunar-driven cycle, the lunar phase is shifted by one to 29 days. That is, a 30-day cycle is started on day one to 29 after the first full moon day, and the 30-day cycle effect is estimated for each specification, using the following pooled regression with PCSE:

$$R_{it} = \alpha_i + \beta 30daydummy_t + e_{it}, \tag{4.9}$$

where R_{it} is the daily logarithmic return for country i and date t, and $30daydummy$ is a dummy variable that indicates the phase of a 30-day cycle. The $30daydummy$ takes on a value of one for 7 days before the starting day + the starting day + 7 days after the starting day, and a value of zero otherwise.

Table 4.11 shows that the 30-day cycle effects for the cycles starting one to eight days and 24 to 29 days after the full moon display negative signs. Moreover, the statistical significance of the estimated 30-day cycle effect declines as these 30-day cycles deviate more from the lunar cycle. In fact, for the cycles starting 10 to 23 days after the full moon, the pattern is reversed. Figure 5 graphs the estimates of the 30-day cycle effect and shows that the documented lunar effect cannot arise from any 30-day cycle except for those that closely track the lunar cycle. Overall, the findings indicate that the lunar effect on stock returns is robust to various lunar window lengths, alternative ARIMA specifications of the stock returns, and a test of random 30-day cycles.

4.5 Conclusion

This chapter investigates the relation between lunar phases and stock returns for a sample of 48 countries. Strong global evidence indicates that stock returns are lower on days around a full moon than on days around a new moon. The return differences are statistically and economically significant during the sample period. Since lunar phases are likely to be related to investor mood and are not related to economic activities, the findings are thus not consistent with the predictions of traditional asset pricing theories that assume fully rational investors. The positive association identified between lunar phases and stock returns suggests that it might be valuable to go beyond a rational asset pricing framework to explore investor behavior.

The psychology literature has provided numerous theories on how mood affects perceptions and preferences. One theory is that mood affects perception through misattribution: attributing feelings to wrong sources leads to incorrect judgments (Schwarz and Clore (1983); Frijda (1988)). Alternatively, mood may affect people's ability to process information. In particular, investors may react to salient or irrelevant information when feeling good (Schwarz (1990); Schwarz and Bless (1991)). Finally, mood may affect preferences (Loewenstein (2000); Mehra and Sah (2000)). This chapter is a first step toward documenting the possible effect of mood on asset prices. It would be interesting to better understand how mood affect asset prices. In a survey paper, Hirshleifer (2001) pointed out that one area of future research is to conduct experimental testing of behavioral hypotheses. In a related vein, future work could study the effect of mood on asset prices in an experimental setting. For example, does investment behavior in experimental settings differ during different phases of a lunar cycle?

Table 4.1: Summary Statistics

This table reports the summary statistics for the 48 country stock indices. All sample periods end on July 31, 2001. All returns are in basis points.

Country	Code	Starting Date	Number of Observations	Mean Daily Log Return	Std Dev of Daily Log Return
Argentina	TOTMKAR	Jan-88	3510	28.42	360.6
Australia	TOTMKAU	Jan-73	7213	3.35	111.79
Austria	TOTMKOE	Jan-74	6355	2.55	86
Belgium	TOTMKBG	Jan-73	7124	2.96	82.26
Brazil	BRBOVES	Jan-72	2475	57.11	646.73
Canada	TOTMKCN	Jan-73	7226	2.97	84.13
Chile	TOTMKCL	Jul-89	3013	8.14	103.21
China	TOTMKCH	Jan-91	2443	11.36	291.94
Czech	CZPX50I	Apr-94	1750	-5.5	127.19
Denmark	TOTMKDK	Jan-74	6377	5.34	108.49
Finland	TOTMKFN	Jan-88	3339	5.46	183.89
France	TOTMKFR	Jan-73	7264	4.17	111.26
Germany	TOTMKBD	Jan-73	7192	2.72	95.33
Greece	TOTMKGR	Jan-88	3385	7.86	191.51
Hong Kong	TOTMKHK	Jan-73	7103	3.97	192.03
Hungary	BUXINDX	Feb-91	2629	7.17	177.12
India	IBOMBSE	Apr-84	2903	6.29	188.61
Indonesia	TOTMKID	Apr-84	2761	-1.18	251.78
Ireland	TOTMKIR	Jan-73	7103	4.69	108.82
Israel	ISTGNRL	Jan-84	4179	14.25	143.62
Italy	TOTMKIT	Jan-73	7445	4.33	134.2
Japan	TOTMKJP	Jan-73	7145	1.81	101.45
Jordan	AMMANFM	Nov-88	2176	2.68	86.07
Korea	TOTMKKO	Jan-75	3322	1.04	207.68
Luxembourg	TOTMKLX	Jan-92	2370	5.65	100.19
Malaysia	TOTMKMY	Jan-88	3349	3.52	164.16
Mexico	TOTMKMX	Jan-88	3436	11.71	170.9
Morocco	MDCFG25	Dec-87	1820	11.99	91.44
Netherlands	TOTMKNL	Jan-73	7219	3.51	95.83
New Zealand	TOTMKNZ	Jan-88	3409	1.71	114.76
Norway	TOTMKNW	Jan-80	5419	3.99	142.43
Pakistan	PKSE100	Dec-88	2795	2.63	162.94
Peru	PEGENRL	Jan-91	2597	15.25	158.34
Philippines	TOTMKPH	Sep-87	3464	4.86	154.32
Poland	TOTMKPO	Jan-94	1803	-2.07	231.97
Portugal	TOTMKPT	Jan-90	2858	1.76	93.31
Russia	RSMTIND	Sep-94	1676	18.85	369.42
Singapore	TOTMKSG	Jan-73	7128	1.2	144.94
South Africa	TOTMKSA	Jan-73	7170	5.53	135.84
Spain	TOTMKES	Jan-88	3623	3.34	116.1
Sweden	TOTMKSD	Jan-82	4903	6.07	134.73
Switzerland	TOTMKSW	Jan-73	7174	2.87	85.17
Taiwan	TOTMKTA	Sep-87	3371	1.89	223.19
Thailand	TOTMKTH	Jan-88	3349	2.09	200.12
Turkey	TOTMKTK	Jan-88	3467	21.28	298.62
United Kingdom	TOTMKUK	Jan-73	7258	3.78	103.42
United States	TOTMKUS	Jan-73	7216	3.26	98.8
Venezuela	TOTMKVE	Jan-90	2829	12.72	249.88
Global Portfolio	Equal-weighted	Jan-73	7456	5.38	58.93
Global Portfolio	Value-weighted	Jan-73	7456	3.07	67.4

Table 4.2: Lunar phases and stock returns: the global portfolio

This table compares the returns of the equal-weighted and value-weighted global portfolio between the full moon and the new moon periods. We define the full moon and the new moon periods using the 15-day and the 7-day windows. In the 15-day window analysis, the full moon period is 7 days before and after the full moon day plus the full moon day; the new moon period is defined as the rest of the lunar month. In the 7-day window analysis, the full (new) moon period is 3 days before and after the full (new) moon day plus the full (new) moon day. Panel A examines the average cumulative returns and Panel B examines the average daily logarithmic returns. We report Newey-West adjusted T-statistics, p-values t2.20 from bootstrap analysis and signed-rank test. The returns are in basis points.

	15-Day Window		7-Day Window	
The Global Portfolio	Equal-Weighted	Value-Weighted	Equal-Weighted	Value-Weighted
			Panel A: Average Cumulative Returns	
Full Moon Return C New Moon Return	-40.26	-30.44	-27.48	-25.87
Newey-West adjusted T-statistic (AR1)	(-2.40)	(-1.87)	(-3.46)	(-2.40)
Newey-West adjusted T-statistic (AR2)	(-2.82)	(-1.94)	(-3.46)	(-2.39)
Newey-West adjusted T-statistics (AR3)	(-2.99)	(-1.99)	(-3.54)	(-2.37)
Bootstrapped p-value	-0.01	-0.07	-0.01	0
Signed-Rank Test (p-value)	0	-0.02	0	-0.01
			Panel B: Average Daily Logarithmic Returns	
Full Moon Return C New Moon Return	-3.53	-2.8	-4.94	-4.82
Newey-West adjusted T-statistic (AR1)	(-2.61)	(-1.75)	(-2.68)	(-2.02)
Newey-West adjusted T-statistic (AR2)	(-2.82)	(-1.82)	(-2.74)	(-2.03)
Newey-West adjusted T-statistics (AR3)	(-2.95)	(-1.86)	(-2.81)	(-2.04)
Bootstrapped p-value	-0.02	-0.05	-0.02	-0.03
Signed-Rank Test (p-value)	0	-0.02	0	-0.02

Table 4.3: Lunar Phases and Stock Returns: Joint Tests

This table reports the estimates of a pooled regression with panel corrected standard errors (PCSE): $R_{it} = \alpha_i + \beta Lunardummy_t + e_t$ for the 15-day window and 7-day window, respectively, where $R_i t$ is average daily logarithmic returns for country i in lunar month t for each full moon and new moon period. Lunardummy is a dummy variable equal to one if it is a full moon period and zero if it is a new moon period. We define the full moon and the new moon periods using the 15-day and the 7-day windows. In the 15-day window specification, we define the full moon period as 7 days before and after the full moon day plus the full moon day, and define the new moon period as the rest of the lunar month. In the 7-day window specification, we define the full (new) moon period as 3 days before and after the full (new) moon day plus the full (new) moon day. The PCSE specification adjusts for the contemporaneous correlation and heteroscedasticity among country indices and for the autocorrelation within each countrys stock index. T-statistics are reported in the parentheses. The daily returns are in basis points. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

	Lunar Dummy (β)	
	15-Day Window	7-Day Window
G7	-2.43	-2.45
	(-1.53)	(-1.09)
Other Developed Markets	-3.21**	-3.19
	(-2.08)	(-1.47)
Emerging Markets	-5.60**	-11.27***
	(-2.11)	(-3.12)
All Markets	-3.95**	-5.93**
	(-2.25)	(-2.44)
All Markets with Country Group Dummies	-3.94**	-5.92**
	(-2.24)	(-2.43)
All Markets with Country Fixed Effects	-4.63***	-8.19***
	(-3.54)	(-3.59)

Table 4.4: Lunar Effect and Stock Sizes

This table reports results from estimating a regression of daily returns of market capitalization ranked portfolios on lunar phases. The portfolios are constructed using stocks traded in the U.S. markets: NYSE and AMEX, and NASDAQ, respectively. Decile 1 corresponds to the largest market capitalization stocks. The following regression is used for each portfolio: $R_{it} = \alpha_i + \beta Lunardummy_t + e_{it}$, Lunardummy is a dummy variable indicating the phase of a lunar cycle, equal to one during a full moon period and zero during a new moon period. The full moon period is 7 days before and after the full moon day plus the full moon day, and the new moon period is the rest of the lunar month. T-statistics with the Newey-West robust standard errors are in the parentheses. The daily returns are in basis points. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

	Lunar Dummy (β)		
Decile	NYSE and AMEX	NASDAQ	
1	-0.66	-3.20*	
	(-0.19)	(-1.66)	
2	-2.73	-3.44*	
	(-1.07)	(-1.86)	
3	-2.05	-3.93**	
	(-0.87)	(-2.01)	
4	-2.94	-4.08**	
	(-1.34)	(-2.00)	
5	-2.71	-3.33	
	(-1.26)	(-1.54)	
6	-3.01	-4.08*	
	(-1.41)	(-1.80)	
7	-2.84	-3.84	
	(-1.32)	(-1.57)	
8	-2.91	-3.97	
	(-1.35)	(-1.58)	
9	-3.35	-5.52*	
	(-1.56)	(-1.94)	
10	-3.03	-2.16	
	(-1.30)	(-0.65)	
	Deciles 1-10	Deciles 1-10	Deciles 1-9
Spearman Rank Correlation (p-value)	-0.81***	-0.2	-0.65*
	-0.005	-0.578	-0.056

Table 4.5: Lunar Phases, Trading Volumes, and Volatility

Panel A reports test results from estimating the following regressions of daily trading volume on lunar phases for the global portfolio and a pooled sample of 48 countries: $normvolume_{jt} = \alpha_j + \lambda_j Lunardummy_t + e_{jt}$. *normvolume* is daily trading volume normalized by average daily volume in that month. Panel B reports the following regression estimates for the global portfolio and a pooled sample of 48 countries: $volatility_{jT} = \alpha_j + \lambda_j Lunardummy_T + e_{jT}$. Volatility is the standard deviation of daily logarithmic stock returns in the full moon and the new moon period for each lunar month. Lunardummy is a dummy variable equal to one during a full moon period and zero during a new moon period. The full moon period is 7 days before and after the full moon day plus the full moon day, and the new moon period is the rest of the lunar month. The estimates of a pooled regression use panel corrected standard errors. The estimates of an OLS regression use the Newey-West robust standard errors with one lag. T-statistics are reported in the parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

	Panel A: Trading Volumes
	Lunar Dummy (λ)
Global Portfolio (equal-weighted)	36.27
	-0.64
Pooled Regression of 48 countries	48.802
	-1.01
	Panel B: Return Volatility
Global Portfolio (equal-weighted)	0.1
	-0.05
Pooled Regression of 48 countries	-0.11
	(-0.10)

Table 4.6: Lunar Phases, Stock Returns, and Macro-Variables

This table compares the returns of global portfolios between the full moon and the new moon periods, excluding macroannouncement dates, periods of major global shocks, or controlling for short-term interest rates. We define the full moon and the new moon periods using the 15-day and the 7-day windows. In the 15-day window specification, we define the full moon period as 7 days before and after the full moon day plus the full moon day, and define the new moon period as the rest of the lunar month. In the 7-day window specification, we define the full (new) moon period as 3 days before and after the full (new) moon day plus the full (new) moon day. Model 1 estimates the following regressions: $R_t = \alpha + \beta Lunardummy_t + e_t$, where R_t is daily logarithmic returns. Lunardummy is a dummy variable equal to one if it is a full moon period and zero if it is a new moon period. Model 2 estimates the model excluding dates for eight macroannouncements: Consumer Price Index, Federal Reserve Open Market Committee announcements, Gross National Product, Retail Sales, Employment Report, Employment Cost Index, Trade Deficit, and National Association of Purchasing Managers Survey Index. Model 3 excludes periods of global shocks: the 1987 U.S. stock market crash (October 19, 1987), the Gulf War (January 17, 1991 to February 17, 1991), the Mexican Peso crisis (December 20, 1994 to January 31, 1995), the Asian financial crisis (July 2, 1997 to December 3, 1997), and the Russian crisis (August 11, 1998 to January 15, 1999). Model 4 estimates the following regression: $R_t = \alpha + \beta Lunardummy_t + \delta shortterminterestrate_t + e_t$, where $shortterminterestrate$ is 3-month Treasury bill rate. Newey-West adjusted T-statistics (with one lag) are reported in the parentheses. The daily returns are in basis points. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

	15-Day Window		7-Day Window	
	Equal-Weighted Global Portfolio	Value-Weighted Global Portfolio	Equal-Weighted Global Portfolio	Value-Weighted Global Portfolio
Model 1: Global Portfolio (Base Case)				
Lunar Dummy	-4.26***	-3.47**	-5.40**	-4.77*
	(-2.81)	(-2.03)	(-2.40)	(-1.89)
Model 2: Excluding Macro-announcement Dates				
Lunar Dummy	-4.30**	-3.65*	-6.21**	-5.37*
	(-2.40)	(-1.84)	(-2.40)	(-1.82)
Model 3: Excluding Major Global Shocks				
Lunar Dummy	-3.65**	-2.57	-4.07*	-2.87
	(-2.51)	(-1.50)	(-1.90)	(-1.15)
Model 4: Controlling for Short-term Interest Rates				
Lunar Dummy	-4.08**	-3.02*	-4.92**	-4.02
	(-2.66)	(-1.72)	(-2.16)	(-1.56)
Short-Term Interest Rate	-0.94***	-0.69**	-1.12***	-1.30***
	(-3.51)	(-2.32)	(-3.05)	(-3.21)

Table 4.7: Lunar Phases, Stock Returns, and Other Calendar Anomalies

This table reports regression results of daily logarithmic stock returns on lunar phases (as in Table 4.6) with controls for other calendar anomalies. Model 1 controls for the January effect. Model 2 controls for the calendar month effect. Model 3 controls for the holiday effect. Model 4 controls for lunar holiday effects. Yomcum dummy is equal to 1 for the day of and the day following Yom Kippur. Roshcum dummy is equal to 1 for the first day of Rosh Hashanah and the following day. Islamic, Hindu, Chinese, and Korean lunar holidays are also controlled for (the coefficient estimates for these lunar holidays are not reported). Newey-West adjusted T-statistics (with one lag) are reported in the parentheses. The daily returns are in basis points. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

	15-Day Window		7-Day Window	
	Equal-Weighted Global Portfolio	Value-Weighted Global Portfolio	Equal-Weighted Global Portfolio	Value-Weighted Global Portfolio
				Model 1: January Effect
Lunar Dummy	-3.98***	-3.00*	-4.77**	-3.96
	(-2.58)	(-1.70)	(-2.09)	(-1.52)
January Dummy	13.81***	8.20**	16.67***	7.96
	-4.4	-2.54	-3.47	-1.63
				Model 2: Calendar Effect
Lunar Dummy	-3.99***	-3.01*	-4.73**	-3.92
	(-2.59)	(-1.71)	(-2.07)	(-1.51)
Calendar Month Dummy	0.66	0.56	-1.34	-1.19
	-0.43	-0.32	(-0.60)	(-0.47)
				Model 3: Holiday Effect
Lunar Dummy	-4.24***	-4.78***	-4.82**	-7.74***
	(-2.79)	(-2.59)	(-2.13)	(-2.78)
				Model 4: Lunar Holiday Effect
Lunar Dummy	-3.76**	-2.57	-4.33*	-3.53
	(-2.42)	(-1.44)	(-1.84)	(-1.31)
Yomcum Dummy	-17.85	-34.84**	5.43	-11.83
	(-1.18)	(-2.23)	-0.69	(-0.81)
Roshcum Dummy	0.46	7.21	4.49	11.93
	-0.06	-0.86	-0.46	-1.09

Table 4.8: Lunar Holidays

This table reports the 15-day window regression results of daily logarithmic stock returns on lunar phases controlling for the January effect and the lunar holiday effect. Yomcum dummy is equal to one for the day of and the day following Yom Kippur. Roshcum dummy is equal to one for the first day of Rosh Hashanah and the day following. Other lunar holiday dummies are country/religion specific. Newey-West adjusted T-statistics (with one lag) are reported in the parentheses. The daily returns are in basis points. ***, **, * indicate 1%, 5%, 10% significance levels respectively using a two-tailed test.

				Independent Variables		
Dependent Variables	Intercept	Lunar Dummy	January Dummy	Yomcum Dummy	Roshcum Dummy	Other Lunar Holiday Dummy
U.S.	3.67**	-1.89	6.81	-39.39**	17.44*	
	-2	(-0.78)	-1.56	(-2.27)	-1.67	
Israel	19.01***	-11.17**	8.49	-54.39	71.00**	
	-5.74	(-2.41)	-0.91	(-0.80)	-2.15	
China	14.71	-8.45	7.34			34.36
	-1.51	(-0.70)	-0.48			-0.93
Japan	3.41*	-4.57*	8.71*			0.23
	-1.82	(-1.81)	-1.79			-0.03
Korea	-2.49	1.96	27.62*			94.32*
	(-0.74)	-0.27	-1.74			-1.77
India	9.96*	-8.15	7.98			-11.87
	-1.84	(-1.11)	-0.62			(-0.48)
Indonesia	7.77	-19.23**	25.66			-48.29
	-1.28	(-2.23)	-1.48			(-0.52)
Jordan	2.54	-1.23	8.9			-1.26
	-0.91	(-0.32)	-1.29			(-0.10)
Malaysia	7.16*	-8.28	0.26			23.58
	-1.73	(-1.43)	-0.02			-1.6
Morocco	12.15***	-1.39	9.03			-9.37
	-3.84	-0.31	-0.91			(-1.13)
Pakistan	3.46	-1.17	-2.3			-3.1
	-0.74	(-0.18)	(-0.19)			(-0.14)
Turkey	22.98**	-12.62	52.28**			3.08
	-2.95	(-1.20)	-2.4			-0.1

Table 4.9: Lunar Phases, Stock Returns, and Varying Lunar Window Length

TThe following regressions are estimated: $R_t = \alpha + \beta Lunardummy_t + e_t$, where R_t is daily logarithmic returns. Lunardummy is a dummy variable indicating the phase of a lunar cycle. We define a full moon period as N days before the full moon day+the full moon day+N days after the full moon day (N =0 to 7). Similarly, we define a new moon period as N days before the new moon day+the new moon day+N days after the new moon day (N =0 to 7). Lunardummy is equal to one during a full moon period and zero during a new moon period. Window length is 2*N +1. Newey-West adjusted T-statistics (with one lag) are reported in the parentheses. The daily returns are in basis points. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

			Independent Variables			
Dependent Variables	Intercept	Lunar Dummy	January Dummy	Yomcum Dummy	Roshcum Dummy	Other Lunar Holiday Dummy
U.S.	3.67**	-1.89	6.81	-39.39**	17.44*	
	-2	(-0.78)	-1.56	(-2.27)	-1.67	
Israel	19.01***	-11.17**	8.49	-54.39	71.00**	
	-5.74	(-2.41)	-0.91	(-0.80)	-2.15	
China	14.71	-8.45	7.34			34.36
	-1.51	(-0.70)	-0.48			-0.93
Japan	3.41*	-4.57*	8.71*			0.23
	-1.82	(-1.81)	-1.79			-0.03
Korea	-2.49	1.96	27.62*			94.32*
	(-0.74)	-0.27	-1.74			-1.77
India	9.96*	-8.15	7.98			-11.87
	-1.84	(-1.11)	-0.62			(-0.48)
Indonesia	7.77	-19.23**	25.66			-48.29
	-1.28	(-2.23)	-1.48			(-0.52)
Jordan	2.54	-1.23	8.9			-1.26
	-0.91	(-0.32)	-1.29			(-0.10)
Malaysia	7.16*	-8.28	0.26			23.58
	-1.73	(-1.43)	-0.02			-1.6
Morocco	12.15***	-1.39	9.03			-9.37
	-3.84	-0.31	-0.91			(-1.13)
Pakistan	3.46	-1.17	-2.3			-3.1
	-0.74	(-0.18)	(-0.19)			(-0.14)
Turkey	22.98**	-12.62	52.28**			3.08
	-2.95	(-1.20)	-2.4			-0.1

Table 4.10: Lunar Phases, Stock Returns, and ARIMA Specifications

ARIMA specifications are estimated for the daily logarithmic returns of the portfolios and then the following regression is estimated: $R_t = \alpha + \beta Lunardummy_t + e_t$, where R_t is the residual of the ARIMA model. Lunardummy is a dummy variable indicating the phase of a lunar cycle, equal to one during a full moon period and zero during a new moon period. The full moon period is 7 days before and after the full moon day plus the full moon day, and the new moon period is the rest of the lunar month. T-statistics are reported in the parentheses. The daily returns are in basis points. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

	Lunar Dummy (β)	
	Equal-Weighted Global Portfolio	Value-Weighted Global Portfolio
ARIMA (1, 1)	-4.12**	-3.71**
	(-2.66)	(-2.05)
ARIMA (1, 0)	-4.08***	-3.60*
	(-2.60)	(-1.93)
ARIMA (0, 1)	-4.14***	-3.71**
	(-2.76)	(-2.04)
ARIMA (1, 2)	-4.11***	-3.73**
	(-2.82)	(-2.07)
ARIMA (2, 1)	-4.10**	-3.74**
	(-2.64)	(-2.09)

Table 4.11: 30-Day Cycles and Stock Returns

This table reports the estimates of a pooled regression with panel corrected standard errors (PCSE): $R_{it} = \alpha_i + \beta 30daydummy_t + e_{it}$ for a 15-day window when lunar phases are shifted by N calendar days. A 30-day cycle is started N days after the first full moon (N =1 to 29), and then the 30-day cycle effect is estimated. 30daydummy takes on a value of one for 7 days before the starting day+the starting day+7 days after the starting day, and a value of zero otherwise. The lunar cycle is represented by N =0. T-statistics are in parentheses. The daily logarithmic returns are in basis points. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using a two-tailed test.

N	β	N	β
1	-3.37*** (-4.11)	16	2.73*** (3.33)
2	-3.28*** (-4.00)	17	3.40*** (4.14)
3	-2.69*** (-3.28)	18	2.53*** (3.08)
4	-3.44*** (-4.19)	19	2.54*** (3.08)
5	-3.07*** (-3.74)	20	3.17*** (3.85)
6	-3.19*** (-3.88)	21	2.39*** (2.90)
7	-0.85 (-0.03)	22	0.10 (0.72)
8	-0.27 (-0.33)	23	0.75 (0.96)
9	0.22 (0.27)	24	-0.80 (-1.03)
10	1.32 (1.60)	25	-1.49* (-1.92)
11	3.44*** (4.19)	26	-3.63*** (-4.67)
12	3.89*** (4.74)	27	-4.44*** (-5.71)
13	4.26*** (5.19)	28	-4.10*** (-5.27)
14	4.16*** (5.07)	29	-3.85*** (-4.95)
15	4.48*** (5.45)	30	-4.55*** (-5.55)

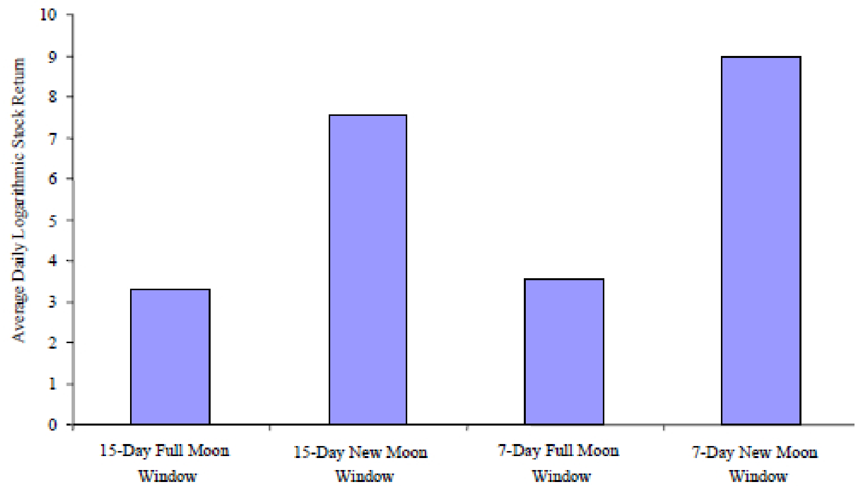

Figure 4.1: Average daily logarithmic stock returns of the global portfolio by lunar phases. This figure plots the average daily stock returns of an equal-weighted global portfolio of the 48 country stock indices in a full moon period and a new moon period. The two bars on the left are average returns of a 15-day window; the two bars on the right are average returns of a 7-day window. All returns are in basis points.

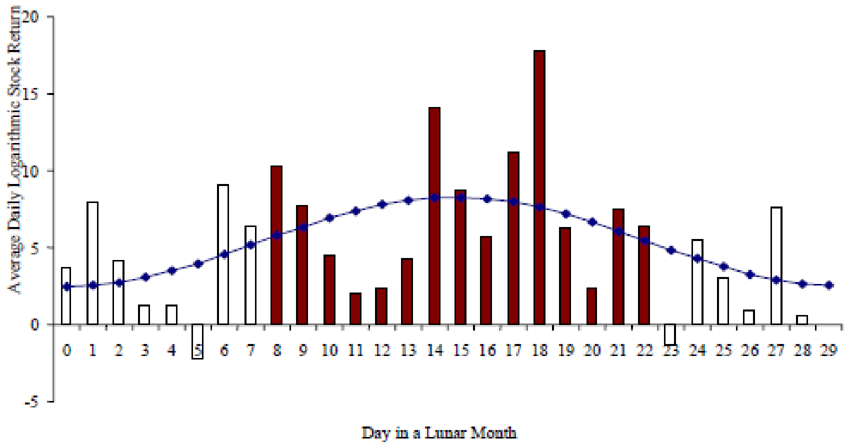

Figure 4.2: Average daily logarithmic return of the global portfolio by lunar dates. This figure graphs, for each day of a lunar month, the average daily logarithmic stock returns of an equal-weighted global portfolio of the 48 country stock indices. Day 0 is a full moon day and day 15 is around a new moon day (day 15 is around new moon day since the length of a lunar month varies). The curved line is the estimated sinusoidal model of the lunar effect on stock returns from the following estimated equation: $R_t = \alpha + \beta \cos(2\pi d_t/29.53) + e_t$, where d is the number of days since the last full moon.

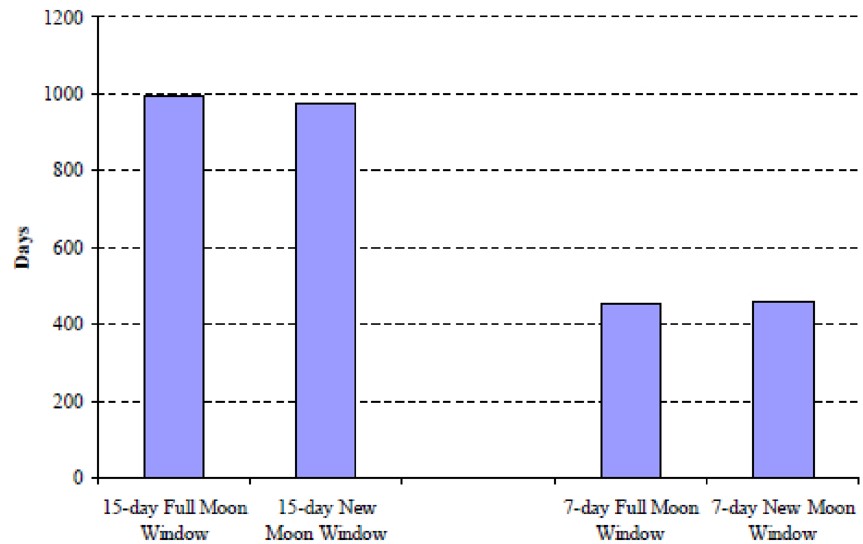

Figure 4.3: Distribution of macro-announcement days on full moon and new moon phases. This figure plots the number of announcement dates in full moon and new moon phases in the sample.

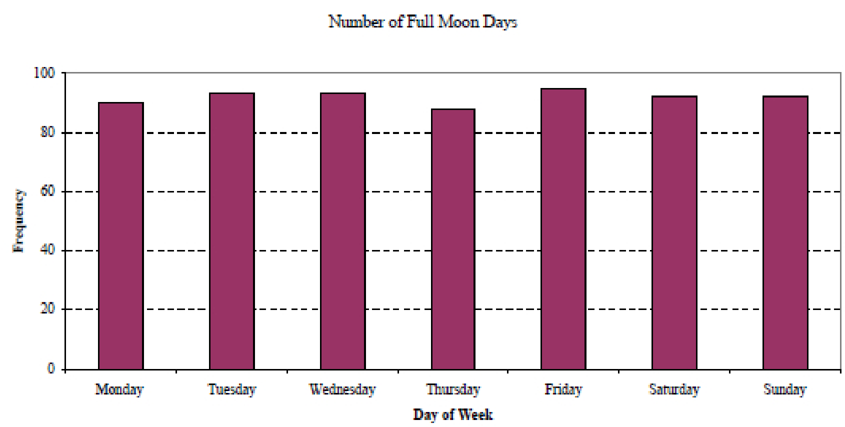

Figure 4.4: Distribution of full moon days on days of a week. This figure plots the number of full moon days falling on each weekday during the sample period.

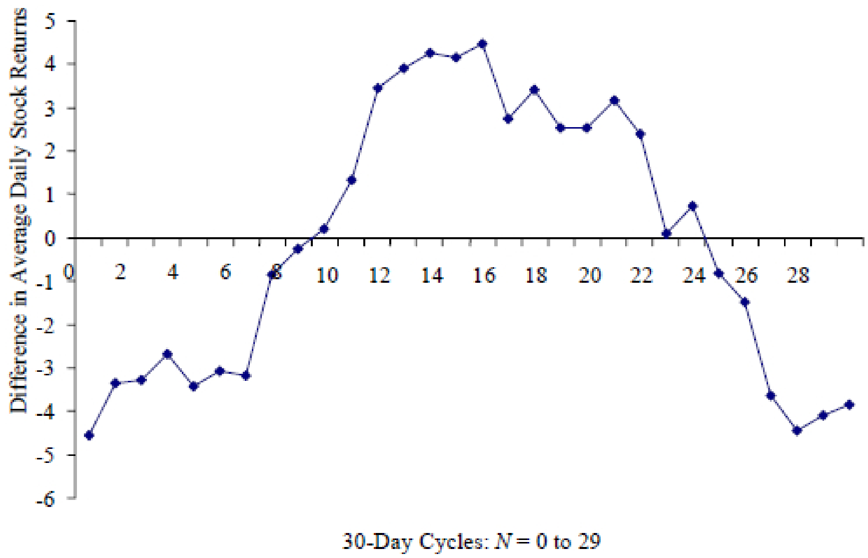

Figure 4.5: 30-Day cycles and stock returns. This figure graphs the estimates of pooled regressions with panel corrected standard errors (PCSE): $R_{it} = \alpha_i + \beta 30daydummy_t + e_{it}$ for a 15-day window when lunar phases are shifted by N calendar days. More specifically, a 30-day cycle of N days is started after the first full moon (N =1 to 29), and then estimate the 30-day cycle effect for each specification. 30daydummy takes on a value of one for 7 days before the starting day+the starting day+7 days after the starting day, and a value of zero otherwise. The lunar cycle is represented by N =0. The X-axis indicates 30-day cycles ordered by N. The Y-axis marks b estimates. The daily returns are in basis points.

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