Strategic Disclosure and Stock Returns: Theory and Evidence from U.S. Cross-listing*

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Abstract

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Abstract

When a firm exercises discretion to disclose or withhold information (strategic disclosure), risk-averse investors command higher expected returns when expected cash flows decrease, leading to a negative correlation between expected cash flows and expected returns. Moreover, stock returns exhibit stronger reversal than they do when full disclosure is enforced. We propose a model that makes these predictions and provide consistent evidence using a panel of foreign firms that list ADRs. We find significant shifts in the time-series properties of stock returns for firms that undergo large changes in disclosure environments, such as those cross-listing on the NYSE/AMEX/NASDAQ and those from emerging markets.
Introduction

The integrity of corporate disclosure sustains investors’ confidence in trading securities at fair prices, and hence is at the heart of well-functioning capital markets. In practice, however, a firm’s disclosure may reflect a strategic decision of its self-interested manager, as it involves a number of estimates, judgments, and assumptions. In reality, managers may obfuscate the true firm performance within the allowances of investor protection regulations.

This paper provides a theoretical framework to analyze the effects of firms’ disclosure strategies on stock returns, which we test with a sample of foreign firms that have listed American Depositary Receipts (ADRs). Our theory builds on Shin’s (2003) ‘strategic disclosure’ model, which formalizes the notion that “although the manager has to tell the truth, he cannot be forced to tell the whole truth” (p.108). In this model, the firm has multiple projects, each of which has two possible outcomes, a success and a failure. The manager observes some of these projects’ outcomes, while investors observe only the manager’s disclosure. The manager has discretion to disclose or withhold what he observes, as the information he has at the time of disclosure is unverifiable even at a later date. In this setting, Shin shows that full disclosure is not supported in equilibrium, but a strategy that discloses all observed successes and withholds all failures is.

The motivation for this study stems from a desire to understand how strategic disclosure affects the joint behavior of reported earnings and stock returns, especially in comparison to full disclosure. Therefore, we extend Shin’s (2003) analysis by providing a complete treatment of full disclosure under risk aversion, and analyze the effects of strategic disclosure on expected stock returns in

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1ADRs are created by U.S.-based depositary banks as claims on non-U.S. firm’s equity shares in their home markets, and are traded just like U.S. stocks. Since most U.S. cross-listings utilize ADRs, this paper uses the terms ‘ADR’ and ‘U.S. cross-listing’ interchangeably.
comparison to those of full disclosure. One direct implication of our model is that a firm’s expected stock return is lower under full disclosure than under strategic disclosure due to reduced information risk, which is consistent with the theoretical predictions of Diamond and Verrecchia (1991) and Easley and O’Hara (2004). Moreover, stock returns will exhibit stronger reversal (negative autocorrelation) under strategic disclosure than under full disclosure. This is because risk-averse investors require a large price discount (information risk premium) in receipt of disclosure when they are uncertain as to whether the manager is withholding information or he is simply uninformed. The price discount tends to be reversed subsequently as more information becomes available.

We further propose that, when risk-averse investors are uncertain about the manager’s information and the manager does not credibly make full disclosure, investors command higher expected returns as expected cash flows decrease. As a result, changes in cash flow expectations (the “cash flow news”) and changes in expected returns (the “expected return news”) are negatively correlated. Intuitively, as the manager reports fewer successes, investors become more skeptical as to whether he is withholding more failures. This makes them require higher expected returns and produces a negative relation between cash flow expectations and expected returns. On the other hand, the negative correlation will be attenuated by an event that unambiguously increases the firm’s commitment to more truthful disclosure, such as listing its shares on markets with higher disclosure requirements. To our knowledge, no other existing studies provide similar implications for the effects of information on stock returns, especially on the relation between cash flow expectations and expected returns.

We test our predictions using a panel of foreign firms that listed ADRs between 1983 and
2003 at annual frequency.\textsuperscript{2} ADRs substantially limit the discretion of foreign managers to disclose strategically, as the US Securities and Exchange Commission (SEC) mandates a much higher disclosure level than most foreign market regulators do.\textsuperscript{3} Furthermore, expected litigation costs for withholding unfavorable information are much higher in the U.S. than in other countries (e.g., Skinner 1994, 1997). Since our sample is a short panel, we implement our econometric analysis using the panel vector-autoregression (VAR) framework of Vuolteenaho (2002). We then employ the return decomposition of Campbell (1991) and Vuolteenaho (2002) to identify the cash flow news and expected return news. Since standard panel data estimators are known to be biased when the regressors include lagged dependent variables, we implement the generalized method of moments (GMM) estimator of Arellano and Bond (1991) and Arellano and Bover (1995). We test our hypotheses based on the Jackknife standard errors that are conservative and robust to outliers.

Our empirical results provide compelling evidence for our predictions. Before the cross-list we find significant return reversal, which diminishes remarkably afterward. The decrease in return reversal is significant for firms that experience large shifts in accounting standards; that is, for firms that comply with U.S. GAAP to cross-list on NYSE, AMEX, and NASDAQ, as opposed to OTC Pink Sheets and PORTAL that do not require such compliance. The decrease is also significant for firms from countries with less-developed or emerging capital markets than those

\textsuperscript{2}‘Disclosure’ in our empirical analysis thus refers to annual corporate financial reports. The effects of ‘voluntary disclosure’ are outside the scope of this study.

\textsuperscript{3}Non-U.S. firms listing on major U.S. exchanges (i.e., NYSE, AMEX, and NASDAQ) are subject to the Securities Act of 1933 and the Securities Exchange Act of 1934. They must file annual reports and Form 20-F to the SEC, where the Form 20-F provides a reconciliation with U.S. GAAP. On the other hand, firms that cross-list on OTC “Pink Sheets” or PORTAL (the market for firms issuing equity under SEC Rule 144a) are exempt from Form 20-F, i.e., they are not required to comply with U.S. GAAP. Nevertheless, they usually need to disclose more than the level mandated by the home market regulators.
from developed countries. We also find that the negative correlation between the cash flow news and expected return news diminishes substantially upon cross-listing; the correlation changes from −0.73 to −0.26, and the difference is statistically significant at the 1% level. The post-listing correlation is again much weaker for firms that cross-list on NYSE, AMEX, or NASDAQ, and for firms from emerging markets.

Recently, a substantial body of research has shown that legal origin matters for the quality of financial disclosure. For example, Ball, Kothari, and Robin (2000) show that managers in countries with code-law (civil-law) tradition are more likely to withhold unfavorable information than managers in countries with common-law tradition. This suggests that strategic disclosure plays a larger role in code-law countries than in common-law countries. Then it is natural to ask if firms from code-law countries experience larger shifts in return properties after the cross-list. Consistent with our prediction, we find that the magnitude of return reversal and the negative correlation between the cash flow news and expected return news diminish more significantly among firms from code-law countries than those from common-law countries.

Furthermore, our time-series evidence remains robust even after controlling for the potential microstructural effects of cross-listing. We also show that these effects are not likely to be driven by other factors, as we do not find similar evidence for a matched sample of otherwise similar firms that do not cross-list on U.S. markets. The auxiliary evidence lends further support to our argument that changes in the disclosure environment around ADR listing affect the time-variation of expected stock returns.

Rogers, Schrand, and Verrecchia (2006) recently test Shin’s (2003) prediction for the negative return-volatility relation (the leverage effect) using an EGARCH specification. Their modest evi-
dence for the prediction could be attributable to the fact that variation in the degree of strategic disclosure among U.S. firms is already small under the strict requirements of the U.S. GAAP. Our work differs from theirs in two ways. First, we derive and test a novel implication of Shin’s (2003) model, rather than empirically examine some of his original predictions. Second, we look at the sample of ADR firms that we believe experience maximal changes in Shin’s (2003) key assumptions; namely, those that shift from an environment that leaves much room for strategic disclosure to one that enforces fuller disclosure.

Bailey, Karolyi, and Salva (2006) show that changes in the cross-listing firms’ disclosure environment significantly affect the way that stock returns and trading volumes respond to announced earnings. Rather than the short-run effects of information release in different disclosure environments that they examine, our primary focus is on the effects of disclosure environments on the time-series behavior of expected returns and expected cash flows. Consequently, we investigate the low frequency variation in stock returns before and after ADRs.

In the international cross-listing literature, Foerster and Karolyi (1999), Miller (1999), Baker, Nofsinger, and Weaver (2002), and Lang, Lins, and Miller (2003), among others, show that various proxies for improved information flows are associated with declines in the cross-listing firms’ expected returns. These papers attribute their findings to Merton’s (1987) investor recognition hypothesis. Doidge, Karolyi, and Stulz (2004) provide a detailed assessment of other explanations for the valuation gains of cross-listing firms, and emphasize the role of cross-listing as a commitment to protect minority shareholders against potential cash flow expropriations by controlling shareholders. Our study contributes to this strand of literature by showing that the time-series property of foreign firms’ stock returns and earnings change significantly when they subject themselves to a
more stringent disclosure environment in U.S. markets.

The rest of the paper is organized as follows. In the next section, we present our theory and derive its testable implications. Section 2 explains our data and provides empirical evidence. It also conducts robustness tests. Section 3 concludes. The Appendix provides proofs and detailed descriptions of our econometric methodology.

1 Strategic Disclosure and Stock Returns

This section analyzes the effect of strategic disclosure on the formation of price and return. Following Shin (2003), we set up our model as a “disclosure game” between a firm manager and investors. The objective is to draw new theoretical predictions about the effects of different disclosure strategies on the time-series properties of stock returns. As we will see, risk aversion of investors produces rich dynamics in stock returns when the manager has discretion to disclose or withhold information.

1.1 Set-up

The notation generally conforms to Shin (2003). There are three dates, 0, 1, and 2. At date 0, a firm undertakes \( N \) independent and identical projects that will be completed over time. Each project succeeds with probability \( r \) and fails with probability \( 1 - r \). The firm’s liquidation value evolves through a binomial tree. If a project succeeds, the firm value ‘climbs up’ one subtree and changes by a factor of \( u \), while if it fails, the firm value ‘climbs down’ by a factor of \( d \) (\( 0 < d < u \)). Thus, the firm’s liquidation value is given by \( u^k d^{N-k} \) when \( k \) projects are completed successfully.

Each project outcome is realized at date 1 with probability \( \theta \) and observed by the firm’s manager. This probability is identical across projects, and whether the outcome is realized and observed is
independent across projects. We denote the number of successes and failures observed by the manager by $s$ and $f$, respectively, compactly represented by $(s, f)$. However, investors observe only the manager’s disclosure. Given this information asymmetry, the manager chooses his disclosure policy to maximize firm value. Following Shin (2003), we assume that the manager is free to disclose some or all of his information at date 1. For example, if the manager observes $(s, f)$ at date 1, he can disclose $(s_1, f_1), 0 \leq s_1 \leq s, 0 \leq f_1 \leq f$, because it will be consistent with any terminal outcome $(s_2, f_2), s_2 \geq s, f_2 \geq f$, that may be realized at date 2. In particular, the set of admissible disclosures includes $(s, f)$ and $(s, 0)$. However, he cannot disclose for example $(s + 1, 0)$, because it will be inconsistent when all unobserved project outcomes turn out to be failures at date 2.

The implicit understanding is that the manager’s disclosures must be accompanied by verifiable evidence. If the manager knows that project $j$ has failed, he cannot claim that it has succeeded, since anti-fraud laws can impose a very large penalty if the disclosure is inconsistent with the evidence. Nevertheless, he can choose not to reveal the whole truth because the amount of information he had at date 1 is not verifiable even at a later date. This assumption is at the heart of Shin’s (2003) model and makes disclosure discretionary and endogenous rather than mandatory and exogenous.\footnote{The verifiability assumption offers a reasonable characterization of financial disclosures, since disclosures are audited by third parties, such as accounting firms, who are not directly affected by the content of the disclosure and for whom the reputation for truthfulness is valuable (Admati and Pfleiderer 2000).}

By date 2, all the projects are completed and their outcomes are announced to all market participants. The firm is liquidated and consumption takes place. For simplicity, we assume no interim consumption at date 1.

The firm’s stock is traded in the financial market. The price of the stock at date $t$ is denoted by $V^a_t, t = 0, 1, 2$. Throughout this paper, superscript $a$ represents a risk-averse quantity and
distinguishes it from a risk-neutral quantity. The stock pays in units of a consumption good, which serves as numeraire of the economy. Without loss of generality, we normalize the supply of the stock at one and assume a zero interest rate. We further assume that the market is complete. At each of the two trading dates \( t = 0, 1 \), investors maximize their expected terminal utility given their information set, \( \mathcal{F}_t \),

\[
E_t[U(W_2)|\mathcal{F}_1],
\]

where \( W_2 \) is their wealth at date 2, and \( U(W) = (W^{1-\alpha} - 1)/(1 - \alpha) \) is a power utility function with a coefficient of relative risk aversion \( \alpha > 0 \). Note that investors’ information set at date 1, \( \mathcal{F}_1 \), depends on the manager’s disclosure strategy. Prior to trading at date 0, investors are endowed with one share of the firm’s stock and no consumption good.\(^5\)

1.2 Equilibrium

We examine two manager strategies, one in which he discloses only the number of successful projects and the other in which he additionally discloses the number of failures. Shin (2003) calls the former “the sanitization strategy” and the latter “full disclosure.”\(^6\) Under the sanitization strategy, the

\(^5\)Recently Shin (2006) extends his own (2003) framework to make both the number of projects (\( N \)) and the probability parameter \( \theta \) uncertain in order to explain short-run momentum (drift) and long-run reversal of stock returns following earnings announcements. He also considers the case of risk-averse investors. Shin (2006, Theorem 6) is able to show that short-run expected return is increasing in \( \theta \) but long-run expected return is decreasing in \( \theta \). By treating \( \theta \) as random, however, Shin’s (2006) analysis introduces a subtle issue: the sanitization strategy may not be an equilibrium strategy, as Shin himself notes. This issue does not concern our analysis since we maintain Shin’s (2003) original framework in which \( N \) and \( \theta \) are known constants. Rather than explaining existing empirical regularities, we aim to draw new theoretical predictions for the time-series properties of stock returns. Thus our model predictions do not overlap with Shin’s (2006) results. We will show that risk-aversion of investors alone generates novel and interesting implications for the time-series of expected returns.

\(^6\)The basic idea of this model goes back to Jung and Kwon (1988), who show that a partial disclosure equilibrium in which a manager strategically withholds bad news but discloses good news can be sustained if investors have
manager chooses not to disclose negative information that reduces firm value.\(^7\) Given a manager strategy, investors maximize their expected utility and set the price, or the firm value (these two terms will be used interchangeably). The manager’s optimal strategy is then to follow the disclosure policy that elicits the maximal firm value.

The firm’s share can be priced by the standard solution technique. The assumption of a complete market implies the existence of state prices,

\[
\phi_t^a = \frac{h(k|F_t)U'(u^k d^{N-k})}{\sum_{k=0}^{N} h(k|F_t)U'(u^k d^{N-k})},
\]

where \(h(k|F_t)\) is the probability of \(k\) terminal successes assessed at date \(t = 0, 1\). Then the firm’s value is given by

\[
V_t^a = \sum_{k=0}^{N} \phi_t^a u^k d^{N-k},
\]

\(t = 0, 1\), and by convention its terminal value is \(V_2^a = u^k d^{N-k}\). Since investors always start with no information about the project outcomes at date 0, we immediately see that the initial firm value is identical between the two manager strategies. However, the share price at date 1 differs between the two strategies due to the difference in investors’ information set. The following theorem calculates these prices and establishes the existence of an equilibrium under the sanitization strategy.

**Theorem 1** (i) The firm value at date 0 is given by

\[
V_0^a = [\psi u + (1 - \psi)d]^N,
\]

where \(\psi \equiv ru^-a/[ru^-a + (1 - r)d^{-a}]\).

\(^7\)Consistent with this, more than 90 percent of the earnings restatements are downward revisions, meaning that firms are more likely to overstate earnings than understate them (e.g., Lev 2003. See also Panel of Audit Effectiveness, available at www.pobauditpanel.org.) In our model, firms can overstate earnings by withholding failures.
There exists an equilibrium in which the manager reports the observed number of successes, $s$, and zero failures. The firm value at date 1 is given by

$$V_1^a(s) = u^s [\pi u + (1 - \pi) d]^{N-s}, \quad (4)$$

where $\pi \equiv qu^{-\alpha}/[qu^{-\alpha} + (1 - q)d^{-\alpha}]$ and $q \equiv (r - r\theta)/(1 - \theta r)$.

The firm value at date 1 when the manager discloses both the observed numbers of successes, $s$, and failures, $f$, is given by

$$V_1^a(s, f) = u^s[df[\psi u + (1 - \psi)d]^{N-s-f}. \quad (5)$$

**Proof.** See the Appendix.

The price function in Equation (3) is intuitive. It is the product of expected values of $N$ independent projects, where each project’s expected value is computed by a modified success probability, $\psi$. Because investors are risk-averse, the price is lower than the risk-neutral price, $[ru + (1 - r)d]^{N}$, which corresponds to the limiting case as $\alpha \to 0$. The price discount due to risk aversion ($\alpha > 0$) is incorporated into $\psi < r$; it is the original success probability weighted by the marginal utility.

The date 1 price in Equation (4) under the sanitization strategy uses another modified success probability, $\pi < \psi$. In this case, anticipating that the manager is hiding possible failures, investors do not take the disclosure at its face value. The discount for the value of the $N - s$ undisclosed projects is represented by $\pi$, which is a combination of the effects of such skepticism and the risk aversion. These two effects are roughly captured by the differences in success probabilities, $\psi - \pi$ and $r - \psi$, respectively.

Finally, Equation (5) under full disclosure prices the $N - s - f$ undisclosed projects with the original modified success probability, $\psi$. Note that, as Shin (2003) shows for the risk-neutral case, this will not constitute an equilibrium price unless full disclosure is enforced exogenously. To see this, suppose the price is given by Equation (5). But then the manager has an incentive to deviate
from his strategy by hiding the failures, which will elicit a higher price $u^s[\psi u + (1 - \psi)d]^{N-s}$.

Knowing this, the investors will not attach probability $\psi$ to the success of undisclosed projects.

Therefore, the manager cannot credibly make full disclosure without a mandatory stipulation.

1.3 Implications for Returns

Using the price functions derived above, we can examine the properties of returns given each of the two disclosure policies. We first examine how a firm’s disclosure policy affects its cost of capital. The following proposition shows that an event that unambiguously increases the manager’s commitment to full disclosure will decrease the firm’s expected return and increase its market value.

**Proposition 2** In receipt of the manager’s disclosure, investors command higher expected return under the sanitization strategy than under full disclosure; $E[E(R^2_s|s)] > E[E(R^2_f|s, f)]$.

**Proof.** See the Appendix. ■

An important consideration in the examination of expected return is the effect of diversification. When investors hold a diversified portfolio, state prices (and hence the pricing kernel) will generally depend on the aggregate wealth rather than on the firm’s payoff. However, as long as the firm’s payoff is strictly positively correlated with the aggregate wealth, it is straightforward to obtain the above proposition and all others to follow.

Proposition 2 complements extant theories that suggest a few channels through which increased disclosure can reduce firms’ expected returns. For example, Barry and Brown (1985) and Coles and Lowenstein (1988), among others, ask if uncertainty about firms’ future payoffs affects expected returns. They show that investors command higher expected returns to hold securities for which they are less confident about the return distributions. Thus, by reducing investors’ estimation risk, increased disclosure may lower the firms’ expected returns. Second, Easley, Hvidkjaer and
O’Hara (2002), Easley and O’Hara (2004) and O’Hara (2003) argue that stocks with more private information must yield higher expected returns, and increased disclosure should lower the firm’s expected returns by suppressing the effects of private information. Investors in our model also require information risk premium when they are uncertain about the manager’s information, unless the manager convinces investors that he is fully disclosing his information. Our model further provides new and refutable implications for the time-series variation of expected stock returns as we will show in the next section.

1.4 Disclosure and Time-Series Properties of Stock Returns

1.4.1 Return Reversal

When investors are risk-neutral, the stock return is serially uncorrelated because the conditional second-period return is always constant at 1. Given this, we may conjecture that return will be negatively serially correlated when investors are risk-averse, because they will require price concession (risk premium) at date 1. Moreover, the magnitude of the negative autocovariance should be larger under the sanitization strategy than under full disclosure because investors will require more price concession in the former case. Formally, define the gross first-period returns under the sanitization strategy and full disclosure,

\[
R_1^a(s) = \frac{V_1^a(s)}{V_0^a} = \frac{u^a[\pi u + (1 - \pi)d]^N s}{[\psi u + (1 - \psi)d]^N},
\]

\[
R_1^f(s, f) = \frac{V_1^a(s, f)}{V_0^a} = \frac{u^a d}{[\psi u + (1 - \psi)d]^{s+f}},
\]

respectively. The following proposition establishes that the above conjectures are indeed true.

Diamond and Verrecchia (1991) argue that increased disclosure reduces future price impact of a trade by mitigating information asymmetries. This attracts large investors’ demand for the firm’s securities, which lowers the firm’s cost of capital. Our representative investor model is silent about trading or liquidity.
**Proposition 3** \(\text{Returns are serially negatively correlated under both the sanitization strategy and full disclosure. Moreover, the magnitude of the negative return autocovariance is larger under the former than the latter;}\) \(\text{cov}(R_1^a(s), E[R_2^a|s]) < \text{cov}(R_1^a(s), E[R_2^a|s, f]) < 0.\)

**Proof.** See the Appendix. ■

### 1.4.2 News to Future Cash Flows and Expected Returns

The manager’s disclosure affects both investors’ expectations about the firm’s future payoffs (cash flows) and expected returns. One immediate implication of our model is that the firm’s expected return decreases in the number of disclosed successes under the sanitization strategy. Under full disclosure, however, the expected return decreases in the total number of disclosed outcomes \((s + f)\) rather than the number of successes \((s)\) alone.\(^9\)

By definition, stock returns can be attributed to changes in expectations of future cash flows (“cash flow news”) and/or changes in expectations of future stock returns (“expected return news”). Following the literature (e.g., Campbell 1991, Vuolteenaho 2002), let us consider the following decomposition of the log first-period return:

\[
\ln \tilde{R}_1^a = N_{cf} - N_{er},
\]

where \(\ln \tilde{R}_1^a = \ln R_1^a - E_0[\ln R_1^a]\). \(N_{cf}\) and \(N_{er}\) denote the cash flow news and expected return news, respectively. Notice that, if investors were risk-neutral, there would be no time-variation in expected returns and hence \(N_{er} = 0\) in our model. Thus we define the cash flow news as the first-period return under risk-neutrality, i.e., \(N_{cf} = \ln \tilde{R}_1 = \ln R_1 - E_0[\ln R_1]\). It then follows that

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\(^9\)The conditional second-period returns are: \(E(R_2^a|s) = \left(\frac{su+(1-q)d}{sw+(1-q)d}\right)^{N-s}\) under the sanitization strategy (Shin 2003, p.121), and \(E(R_2^a|s, f) = \left(\frac{su+(1-r)d}{sw+(1-r)d}\right)^{N-s-f}\) under full disclosure.

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the expected return news is given by

\[ N_{er} = \ln \tilde{R}_1 - \ln \tilde{R}^2_1. \]

Based on the decomposition (6), our objective is to show how the expected cash flows interact with the expected returns (information risk premium). The following proposition establishes that the cash flow news and expected return news covary differently under the two disclosure policies.

**Proposition 4**  
(i) Under the sanitization strategy, the cash flow news and the expected return news are negatively correlated:

\[ \text{Cov}(N_{cf}, N_{er}) < 0. \]

(ii) Under full disclosure, the cash flow news and the expected return news are positively correlated:

\[ \text{Cov}(N_{cf}, N_{er}) > 0. \]

**Proof.** See the Appendix.

Economic intuition behind this proposition is as follows: when investors are unsure about the manager’s true information under the sanitization strategy, they try to infer his withheld information from the disclosure. Since the manager reports only the number of successes, when the reported performance is good (high \( s \)), the investors are more certain that he is not withholding unfavorable information (\( f \)). On the other hand, when the manager reports a poor performance (low \( s \)), the investors are less certain whether he is withholding unfavorable information or he simply does not observe much information. Having to bear such information risk, the investors require higher expected return to hold the stock. This monotonic negative relation between the reported performance and investors’ uncertainty leads to a negative correlation between cash flow expectations and expected returns.
There are two reasons why the negative correlation will be mitigated under full disclosure. First, there is no uncertainty about the manager’s information. Second, an increase in the reported number of failures will decrease not only the value of the cash flow component but also the discount due to risk aversion because it provides more information about the future payoff. The result in Part (ii) of the proposition indicates that these two effects collectively overturn the negative correlation to positive.

Finally, we note that Proposition 4 holds in a multi-period model. Specifically, it is straightforward to obtain the same result in a model with multiple disclosure dates as set up in Shin’s (2003) Section 3.3. Therefore, the result is not specific to the fact that all uncertainty will be resolved in the period just following the manager’s announcement in our three-date model.

1.5 Empirical Predictions

A large body of literature on ADR or U.S. cross-listing documents empirical evidence that is consistent with Proposition 2. Foreign firms’ expected returns decrease substantially when they cross-list on U.S. markets (e.g., Foerster and Karolyi 1999, Miller 1999, Errunza and Miller 2000, Sarkissian and Schill 2005), and the reduction is larger for firms that tend to undergo more substantial shifts in disclosure policies around the cross-list, i.e., firms that cross-list on NYSE, AMEX, and NASDAQ and firms from countries with high investment barriers.

Theoretical predictions of Propositions 3 and 4 are new to the literature. We will examine these predictions in a panel of foreign firms that cross-list on U.S. markets via the ADR programs. Specifically, we compare the time-series behavior of stock returns and earnings before and after the cross-listing. Since disclosure requirements and investor protection regulations are much stricter in the U.S. markets than in other markets, U.S. cross-listing by foreign firms represents a canonical
event that unambiguously increases a firm’s commitment to more disclosure.

The main hypotheses of this paper are summarized as follows:

**Hypothesis A**: The effects of the sanitization strategy prevail in the time-series of stock returns and cash flows in the pre-ADR period.

**Hypothesis B**: The time-series effects of the sanitization strategy diminish after the cross-list. That is, the time-series effects of the sanitization strategy are stronger in the pre-ADR period than in the post-ADR period.

As an auxiliary hypothesis to Hypothesis B, we posit that the decline in the effects of the sanitization strategy is stronger for firms that commit to stricter disclosures. That is, the effects of the sanitization strategy diminish more for firms that cross-list on NYSE, AMEX, or NASDAQ as opposed to OTC Pink Sheets or PORTAL and for firms from emerging market countries rather than from countries with well-developed markets. We also examine if the effects differ between firms from code-law countries and those from common-law countries.

1.6 Identifying the Cash Flow News and the Expected Return News

In order to test the theoretical implications of our model, we need to empirically identify the cash flow news and the expected return news from the time-series of stock returns and earnings. To this end, we rely on Vuolteenaho’s (2002) accounting-based dynamic present value relation, which implies the following return decomposition:

\[ r_t - E_{t-1}r_t = N_{cf,t} - N_{er,t}, \]
where

\[
N_{cf,t} = (E_{t} - E_{t-1}) \left[ \sum_{j=0}^{\infty} \rho^j \times \text{roe}_{t+j} + k_{t} \right],
\]

(7)

\[
N_{er,t} = (E_{t} - E_{t-1}) \left[ \sum_{j=1}^{\infty} \rho^j \times \text{ret}_{t+j} \right].
\]

(8)

In (7) and (8), \text{roe}_t is the log ROE (log of one plus earnings per book value of equity) in excess of the local interest rate (the excess ROE), and \text{ret}_t is the local log stock return in excess of the local interest rate (the excess return). Henceforth, we use the term ROEs and stock returns to mean “excess ROEs” and “excess returns”. The term \(k_t\) reflects the difference between the accounting ROE and the clean-surplus ROE, the time-variation in the dividend payout ratio, and the approximation error due to log linearization. \(\rho\) is a constant close to but less than one.\(^{10}\) We set \(\rho = 0.967\) as in Vuolteenaho (2002), but the small variation in \(\rho\) has no material effects on our conclusions. Following the literature (e.g., Campbell 1991, Vuolteenaho 2002), we call \(N_{cf,t}\) and \(N_{er,t}\) the cash flow news and expected return news, respectively, in interpreting our empirical results.

As in Vuolteenaho (2002) and Cohen, Gompers, and Vuolteenaho (2002), we use a firm-level panel VAR to model the conditional expectation \(E_t\). Let us express the VAR (suppressing the firm-specific constant term) as

\[
z_t = \Gamma z_{t-1} + w_t; \quad \Sigma = E \left[ w_t w_t' \right],
\]

(9)

where \(\Gamma\) and \(\Sigma\) are assumed constant, both across time and across firms, in each subperiod. Fol-

\(^{10}\)Vuolteenaho’s (2002) accounting-based present value model relies on a clean-surplus relation and constant payout ratio.
Following Vuolteenaho (2002), we set our state vector as

\[ z_t = (\text{ret}_t, \text{roe}_t, (p-b)_t)' , \]

where \((p-b)_t\) is the log market-to-book (ME/BE) ratio. Vuolteenaho’s (2002) model implies that the ME/BE ratio contains useful information about future ROEs and future stock returns.

With this specification, we can identify the cash-flow news and the expected return news as

\[
N_{er,t} = e1' \Upsilon w_t, \\
N_{cf,t} = e1' (I + \Upsilon) w_t, \tag{10}
\]

where \(\Upsilon = \rho \Gamma (I - \rho \Gamma)^{-1}\) and \(e1\) is a choice vector with 1 in its first element and 0 elsewhere, e.g., \(e1' z_t = \text{ret}_t\). It then follows that

\[
\text{Cov} (N_{er,t}, N_{cf,t}) = e1' \Sigma (I + \Upsilon') e1. \tag{11}
\]

2 Evidence

2.1 Data and Descriptives

We obtain data on ADR listings from the Bank of New York. Our firm-level panel is from Datastream/Worldscope, covering the period from 1980 through 2005 (26 years). We then match the ADR effective dates (the dates when the ADRs began trading) with foreign firms and divide the data into pre- and post-ADR subperiods. The pre-ADR subperiod is from the earliest available year to one year before the ADR effective date. The post-ADR subperiod is from the year of effective date to last available year. Since our focus is on the time-series behavior of stock returns and cash flows around the cross-list, and due to the nature of our dynamic panel data, we retain
only firms that have three or longer consecutive yearly data in each of the pre- and post-ADR sub-periods. Meanwhile, in order to focus on the effects of ADR listing, we keep the time-dimension of the panel relatively short. Specifically, we work with unbalanced panels with at most seven yearly observations in each of the pre- and post-ADR subperiods.

For each firm year, we collect data for local annual interest rates at the end of the previous year, yearly local stock total returns, annual accounting ROEs (earnings per book value of equity), and the market value of equity per book value of equity (ME/BE or the market-to-book ratio). To ensure that our results are not driven by outliers, we exclude firms with ME/BE ratios either lower than 0.01 or higher than 100, as in Vuolteenaho (2002).

We classify the firms into two categories: firms from developed markets versus firms from emerging markets. We exclude U.K. firms from our analysis since U.K. markets have a distinctly superior information environment for U.S. investors than other markets, and the London Stock Exchange offers a popular alternative venue for international cross-listing. We also exclude three countries, Argentina, Brazil, and Peru, because they had years of triple-digit annual inflation rates during the sample period. Under high inflation, past ROE and ME/BE become meaningless, because both depend on the book value of equity that reflects accumulated earnings, and past

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11 Worldscope mnemonics for stock returns, ROE, and ME/BE are 08801, 08371, and 09704, respectively. For interest rate data, we primarily use 1-year off-shore money market rates if available. Otherwise, we use domestic money market rates whose maturities are close to 1 year and have long historical data on Datastream.

12 See Table 1 for our classification. We classify Austria as an emerging market because of its low accounting standards (e.g., La Porta, Lopez-de-Silanes, Shleifer, and Vishny 1998). Leuz, Nanda, and Wysocki (2003) also report that, in a sample of 31 countries, Austria allows the highest degree of flexibility for managers to exercise their discretions to manage reported earnings. However, the grouping of Austria has minimal effects on our estimation results for both developed markets and emerging markets.

13 Our sample does not include Canadian firms since they can use Canadian GAAP to directly list on U.S. markets. In fact, Canadian GAAP is very similar to U.S. GAAP (Bailey, Karolyi, and Salva 2006, footnote 11).
earnings are accumulated at historical costs without adjusting for inflation over time.

Panel A of Table 1 summarizes the number of firms in the sample by country and by effective year. Our final sample consists of 2,636 firm-year observations for 227 firms from 28 markets that cross-listed on U.S. markets between 1983 and 2003. 170 firms are from 16 developed markets and 57 firms are from 12 emerging markets. ADR listings were sparse in the 1980s with the exception of 1983 in which there are 12 listings, 10 of which are from Japan, in our sample.\textsuperscript{14} The number of ADR listings grew steadily until the Sarbanes-Oxley Act of 2002 slowed it down. We also classify firms into two categories by the trading locations for their ADRs: firms that list ADRs on NYSE, AMEX, or NASDAQ (85 firms) \textit{versus} firms that list ADRs on OTC “Pink Sheets” or PORTAL (for Rule 144a private placements) (142 firms). This classification is important for our study since NYSE, AMEX, and NASDAQ require compliance with U.S. GAAP for cross-listing firms, while OTC and PORTAL do not.\textsuperscript{15} Panel B of Table 1 reports the number of firms by industry and by effective year. 25 firms (11\% of the sample) are banks. Besides this, we do not see any particular industry concentration in our sample. We also control for firms switching from one ADR type to another.\textsuperscript{16}

\textsuperscript{14}Japan liberalized its financial market in 1983 (Bekaert and Harvey, 2000). Excluding this year or other years does not change our conclusion.

\textsuperscript{15}Since 1999, the SEC mandates firms to comply with the U.S. GAAP to be quoted on the OTC Bulletin Board which is an electronic quoting medium operated by the National Association of Securities Dealers. This does not affect our analysis since none of our sample ADRs are traded on the OTC Bulletin Board.

\textsuperscript{16}Two firms in our sample, National Australia Bank (Australia) and Anglogold Ashanti (South Africa) listed their ADRs on NYSE first and subsequently used Rule 144a private placements (RADRs). We categorize them as the NYSE/AMEX/NASDAQ firms. Ten other firms in our sample switched their trading locations between the PORTAL (for RADRs) and the OTC Pink Sheets, which do not affect their OTC/PORTAL classification. None of our sample
Table 2 reports the means and standard deviations of the three state variables for each of the five groups. These figures need to be interpreted with caution since they are estimated for the pooled sample of firm-years. Nevertheless, stock returns display stark contrasts between the pre- and post-ADR periods. For example, average excess stock returns per year are about 8% lower in the post-ADR period than in the pre-ADR period for the whole sample.\footnote{These figures are local stock returns.} This observation is consistent with Proposition 2 and more broadly the argument that ADR listing decreases the firm’s expected returns significantly (e.g., Foerster and Karolyi 1999, Miller 1999, Errunza and Miller 2000). It is also consistent with Sarkissian and Schill’s (2005) observation of pre-listing over-performance and significant cost of capital gains after the cross-listing. On the other hand, average excess ROEs for the whole sample decline only slightly (0.7%) after the cross-list. Meanwhile, standard deviation of ROEs drops from 22.3% to 13.7% after the cross-list in the whole sample.\footnote{Consistent with the notion that the quality of disclosure improves when foreign firms cross-list on U.S. markets, we find a significantly stronger association between ROEs and lagged stock returns after the cross-list than before. In other words, the reported earnings become more value-relevant after the cross-list. This result is consistent with the findings of Lang, Raedy, and Yetman (2003) and Lang, Raedy, and Wilson (2006).}

### 2.2 Test Results

Our theory generates two central predictions. First, we predict that stock returns of foreign firms are serially negatively correlated in their local markets, and the magnitude of this “return reversal” diminishes after they cross-list their shares on U.S. markets (Hypothesis A). Second, we predict that changes in expected stock returns are negatively correlated with changes in cash flow expectations firms cross-listed on OTC/PORTAL first and listed on NYSE/AMEX/NASDAQ later.
for foreign in their local markets, and the magnitude of the covariance and correlation between the expected return news and cash flow news will decline after the cross-list (*Hypothesis B*). As we discussed, poorer disclosed earnings exacerbate the investors’ uncertainty about the manager’s hidden unfavorable information and hence investors require higher expected returns. However, as the manager credibly communicates his commitments to more truthful disclosure by cross-listing on U.S. markets, the negative covariance between cash flow news and expected return news should diminish.

We also predict that the reduction in the effects of strategic disclosure is stronger for firms that list ADRs on NYSE, AMEX, or NASDAQ than firms that choose OTC Pink Sheets or PORTAL as trading locations for their ADRs. Cross-listing on NYSE, AMEX, or NASDAQ requires the firm to go through the substantial challenge of reconciling its financial statements with the U.S. GAAP. Meanwhile, this allows the firm to signal its commitment to more truthful disclosure. We also predict that the magnitude of reversals are stronger for firms from countries with emerging capital markets than for firms from countries with developed markets, since the disclosure standards and investor protection regulation tend to be less strict in less-developed capital markets.

In the following, we first state our theoretical predictions as statistical hypotheses. We then test them and examine whether the effects are stronger for firms that cross-list on NYSE, AMEX, or NASDAQ and for firms from emerging markets, and for firms from code-law countries.
2.2.1 Return Reversal

In order to test for the return reversal in our data, we first consider a panel univariate autoregression
for annual excess stock returns

$$ret_{i,t} = \phi_{ret} ret_{i,t-1} + \eta_{ret,i} + \tilde{\epsilon}_{ret,i,t-1}$$ (12)

for both pre- and post-ADR periods. In (12), \(\eta_{ret,i}\) represents the firm-specific effect and the
autoregressive coefficient \(\phi_{ret}\) provides a measure of autocorrelation. Our statistical hypotheses
corresponding to Hypotheses A and B are:

Hypothesis A-1: \(\phi_{ret}^{pre} < 0\).

Hypothesis B-1: \(\phi_{ret}^{post} - \phi_{ret}^{pre} > 0\).

The null hypotheses are \(\phi_{ret}^{pre} = 0\) and \(\phi_{ret}^{post} - \phi_{ret}^{pre} = 0\), respectively.

Table 3 presents strong evidence for a return reversal, i.e., \(\phi_{ret}^{pre} < 0\), for all groups of firms in
the pre-ADR period. For the whole sample before the cross-list, the autoregressive coefficient \(\phi_{ret}\)
is significantly negative at \(-0.26\) (\(t = -3.3\)). The magnitude of negative autocorrelation is much
stronger for firms from countries with emerging markets than for those from developed markets.
In the pre-ADR period, the \(\phi_{ret}\) coefficients are \(-0.55\) (\(t = -5.6\)) for firms from emerging markets
and \(-0.10\) (\(t = -2.0\)) for firms from developed markets. This result is not unusual since emerging
markets tend to have weaker disclosure standards.

The significant negative correlations persist even after the cross-list, but the degree of “return
reversal” diminishes significantly afterward. We are able to reject the null hypothesis of \(\phi_{ret}^{post} = \phi_{ret}^{pre}\).
for firms that cross-list on NYSE, AMEX, and NASDAQ and for firms from emerging markets at a 5% significance level. Those firms that cross-list on NYSE, AMEX, or NASDAQ and those from emerging markets are likely to convey stronger signals about their commitments to disclose information more truthfully, which in turn suppresses the reversal of their returns more significantly than other firms. Thus, this evidence supports our hypothesis that “return reversals” tend to arise in weak regulatory environments in which investors presume firms’ strategic disclosure behavior, but they tend to diminish as firms credibly convince investors of their commitment to increased disclosure.

2.2.2 Relation between Cash Flow News and Expected Return News

Our theory generates a set of implications for the covariance between the cash flow news \( (N_{cf}) \) and the expected return news \( (N_{er}) \). Relying on Vuolteenaho (2002), we identify the two news components as in equation (10) using a firm-level panel VAR and estimate their covariances and correlations before and after the cross-list. In order to control for the effects of heteroscedasticity, we deflate the covariance by the variance of unexpected stock return, \( \tilde{r}_t \equiv ret_t - E_{t-1}ret_t \), and define \( Cov^* (N_{er,t}, N_{cf,t}) \equiv Cov (N_{er,t}, N_{cf,t}) / Var (\tilde{r}_t) \). Then, we cast our theoretical predictions in the following statistical hypotheses:

\[ Hypothesis \ A-3: \] \[ Cov^* (N_{er,t}, N_{cf,t})^{pre} < 0. \]

\[ Hypothesis \ A-3': \] \[ Corr (N_{er,t}, N_{cf,t})^{pre} < 0. \]

The null hypotheses are \( Cov^* (N_{er,t}, N_{cf,t})^{pre} = 0 \) and \( Corr (N_{er,t}, N_{cf,t})^{pre} = 0 \), respectively. We also have

\(^{19}\)Our inference is based on jackknife standard errors, that are known to overestimate the population standard errors (e.g., Efron and Stein 1981, Efron 1982). In other words, we are conducting conservative tests against the null.
Hypothesis B-3: \( \text{Cov}^* (N_{er,t}, N_{cf,t})^{\text{pre}} < \text{Cov}^* (N_{er,t}, N_{cf,t})^{\text{post}}, \)

Hypothesis B-3': \( \text{Corr} (N_{er,t}, N_{cf,t})^{\text{pre}} < \text{Corr} (N_{er,t}, N_{cf,t})^{\text{post}}, \)

for which the respective null hypotheses are \( \text{Cov}^* (N_{er,t}, N_{cf,t})^{\text{pre}} = \text{Cov}^* (N_{er,t}, N_{cf,t})^{\text{post}} \) and \( \text{Corr} (N_{er,t}, N_{cf,t})^{\text{pre}} = \text{Corr} (N_{er,t}, N_{cf,t})^{\text{post}}, \) respectively.

Table 4 presents our results with jackknife standard errors. The cash flow news and the expected return news are significantly negatively correlated at about \(-0.2\) for all groups of ADR firms before the cross-list. The correlation coefficients are also significantly negative at \(-0.73\) \((t = -5.8)\) for the whole sample and in particular \(-0.96\) \((t = -12.8)\) for emerging market firms.

After the cross-list, however, the magnitude of the negative covariance and correlation diminishes. For example, the correlation changes by 0.47 \((t = 2.7)\) from \(-0.73\) to \(-0.26\) for the whole sample across the cross-list. The increase in the correlation coefficient is stronger and statistically significant for firms that cross-list on NYSE, AMEX, or NASDAQ \((+0.55; t = 2.5)\) than those that cross-list on OTC or PORTAL \((+0.30; t = 1.6)\) and for firms from emerging markets \((+0.69; t = 3.7)\) than those from developed markets \((+0.24; t = 1.2)\).

Recall that our Proposition 4 implies that the covariance between the cash flow news and expected return news is positive under full disclosure. In Table 4, the estimated covariances are negative across all groups of firms even in the post-ADR period, though the covariance is not significantly different from zero for firms that cross-list on NYSE, AMEX, or NASDAQ. This suggests that the time-series properties of post-ADR stock returns are not fully consistent with full
disclosure. Nevertheless, our evidence for Hypotheses B-3 and B-3 lends support to our argument that the effects of strategic disclosure significantly decrease after the cross-list.

The negative correlation implies that expected returns tend to decrease with an increase in cash flow expectations. Consequently, stock returns appear to “over-react” to cash flow news. For example, in a regression \( \tilde{r}_t = b_{er|cf}N_{cf,t} + \tilde{\varepsilon}_t \) with \( \tilde{r}_t \equiv ret_t - E_{t-1}ret_t \), the coefficient is \( b_{er|cf} = 1 - \text{Cov}(N_{cf,t}, N_{er,t}) / \text{Var}(N_{er,t}) \), which exceeds one when \( \text{Cov}(N_{cf,t}, N_{er,t}) \) is negative. This “over-reaction” is consistent with our theory in which, when rational investors are uncertain about the manager’s information and the manager does not communicate all dimensions of his information, rational investors tend to command a higher risk premium when the firm discloses poorer performance. Interestingly, the evidence of negative \( \text{Cov}(N_{cf,t}, N_{er,t}) \) (over-reaction) contrasts to the under-reaction phenomenon that Cohen, Gompers, and Vuolteenaho (2002) document for U.S. small stocks.

In summary, consistent with the time-series implications of our theory, we find compelling evidence for return reversals and negative correlations between cash flow news and expected returns news in foreign firms’ stock returns in their home markets. These effects diminish significantly when

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20 This may not be surprising given that fraudulent disclosures are possible even under the U.S. GAAP as illustrated by the recent corporate accounting scandals by major corporations such as AIG, Enron, Freddie Mac, WorldCom, Xerox, etcetera. Meanwhile, La Porta, Lopez-de-Silanes, Shleifer, and Vishny (2000), Siegel (2005), and Lang, Rady, and Wilson (2006) discuss some limitations of SEC regulatory enforcement for cross-listed foreign firms.

21 Cohen, Gompers and Vuolteenaho (2002) use this \( b_{er|cf} \) to measure the degree of under-/over-reaction.

22 The positive covariance between the cash flow news and expected return news of the U.S. small firms has been attributed to “under-reaction,” which institutional investors appear to exploit only partially (Cohen, Gompers, and Vuolteenaho 2002). On the other hand, U.S. large firms do not exhibit the “under-reaction” phenomenon. Most of the foreign firms that have cross-listed are among the largest firms in their respective home markets to which the under-reaction story is unlikely to apply. Furthermore, our theory implies that the covariance between the cash flow news and expected return news reflects time-varying information risk premium rather than investors’ “over-reaction”.
the firms cross-list their shares on U.S. markets. As we predicted, the reduction in the time series effects is stronger for firms that comply with U.S. GAAP (i.e., cross-list on NYSE, AMEX, or NASDAQ) and for firms from emerging markets.

2.3 Additional Tests

2.3.1 Effects of Legal Origin

A substantial body of literature has suggested that legal origin matters for the quality of corporate financial disclosure, corporate governance structure, and other institutional factors (e.g., La Porta, Lopez-de-Silanes, Shleifer, and Vishny 1998). In particular, it has been documented that countries with a common-law tradition provide stronger protection for shareholders than countries with a code-law tradition. In a recent empirical study, Reese and Weisbach (2002) find that legal origins of home countries explain differences in the choice of trading venues for ADRs. Leuz, Nanda, and Wysocki (2003) and Doidge, Karolyi, and Stulz (2004), among others, also explicitly account for the effects of legal origin in examining the role of U.S. cross-listing in improving minority shareholder protection.

In the accounting literature, Ball, Kothari, and Robin (2000) present evidence that firms in countries with code-law tradition are more likely to withhold bad news than firms in countries with common-law tradition. In code-law countries, private communication among “stakeholders” such as shareholders, managers, employees, the government, and banks plays a relatively more important role than public disclosure. By contrast, the demand for more transparent public disclosure is much higher in common-law countries to mitigate information asymmetry between managers and investors. Then firms from code-law countries should exhibit more pronounced shifts in return properties upon cross-listing, because they are likely to experience
larger changes in accounting standards than firms from common-law countries.

In order to examine this point, we repeat our analysis by firms from countries with different legal origins, following the classification of La Porta, Lopez-de-Silanes, Shleifer, and Vishny (1998). In our sample, 76 firms are from 8 common-law countries and 151 firms are from 20 code-law countries.

Panel A of Table 5, comparable to Table 3, reports the estimated autoregressive coefficient from the regression (12). The reductions in the magnitude of return reversal, \( \phi_{ret}^{\text{post}} - \phi_{ret}^{\text{pre}} \), are 0.12 \( (t = 0.97) \) for firms from common-law countries and 0.29 \( (t = 2.4) \) for firms from code-law countries. Panel B of Table 5, comparable to Table 4, reports standardized covariances and correlation coefficients between the cash flow news and the expected return news for the two groups before and after the cross-list. Again, the increase in the correlation coefficient, \( \text{Corr} \left( N_{er,t}, N_{cf,t} \right)^{\text{post}} - \text{Corr} \left( N_{er,t}, N_{cf,t} \right)^{\text{pre}} \), is more significant for firms from code-law countries \( (+0.57; \ t = 3.0) \) than those from common-law countries \( (+0.23; \ t = 0.91) \). Indeed, these results are consistent with our prediction that U.S. cross-listing should affect the return properties of firms from code-law countries more strongly than those from common-law countries.

2.3.2 Comparison with Non-ADR firms – A Matched Sample Analysis

We have documented a substantial shift in the time-series property of stock returns and reported earnings around the cross-listing, especially among firms from countries with less-developed capital markets. This subsection checks against the possibility that our findings merely reflect shifts in factors other than disclosure environments. For example, developments in local capital markets
could cause some time trend in the quantities that we examine. In order to inspect this, we repeat our tests using a sample of otherwise-comparable emerging market firms that have not cross-listed on U.S. markets. Specifically, we match the ADR firms with their non-cross-listing local market competitors in each of the 12 emerging countries, by carefully matching the four-digit industry code on the Datastream/Worldscope.\textsuperscript{23} If no matching firm is available, then we match the first three digits. From this procedure, we are able to screen out 574 non-ADR comparable firms. Using the ADR effective dates of their matches, we assign the pseudo-ADR years to the non-ADR firms to split the samples into the pseudo pre- and post-ADR periods. For these firms, we apply the same data requirements (e.g., at least 3-years of data in both pseudo pre- and post-ADR periods) to the non-ADR matched firms as we do to the ADR firms. Through this procedure, we are able to identify 120 comparable non-ADR firms, with 620 and 692 firm-year observations in pseudo pre- and post-ADR periods.

Table 6 about here

Table 6 summarizes the results from our matched sample analysis. The magnitude of return reversal (negative return autocorrelation) diminishes after the cross-list only for the ADR firms. By contrast, we do not observe a similar decline in return reversal for non-ADR matched firms. The return autocorrelation exhibits a statistically significant increase of +0.40 for the ADR firms with a Jackknife t-statistic of 2.2. The autocorrelation exhibits a decline, albeit insignificant, for the non-ADR firms. Similarly, we can detect statistically significant increase in the correlation between the cash flow news and expected return news only for the ADR firms, but we do not find

\textsuperscript{23}We look up the four-digit industry code in Field 06011. The Datastream/Worldscope reports 27 major industry groups (reflected in the first two digits), each of which is further broken down into sub-industry groups.
a statistically reliable shift in the correlation for the non-ADR firms.

In summary, while we report significant evidence for shifts in the time-series of stock returns across the cross-list, we do not find any reliable evidence that such shifts also occur in the matched sample of non-ADR firms. Therefore, it is very unlikely that other factors such as financial market development explain our results. This auxiliary analysis further strengthens our argument that changes in disclosure environments exert significant effects on the time-series property of stock returns and earnings upon the cross-list.

### 2.3.3 Interaction with Microstructure Effects

Although our theory offers no implications on trading or liquidity, we are aware that changes in market microstructure might explain our empirical results. For example, Diamond and Verrecchia (1991) argue that increased disclosure reduces the future price impact of a trade, which attracts additional demand from large investors and reduces expected returns. Kim and Verrecchia (2001) further argue that an increase in a firm’s commitment to more disclosure should lower the price impact of a trade. Leuz and Verrecchia (2001) support this argument by showing that German firms that list their stocks on the Neuer Market by satisfying higher disclosure requirements than the Frankfurt Exchange enjoy higher liquidity (e.g., lower bid-ask spreads) for their shares. Moerman’s (2006) recent evidence suggests that timely incorporation of losses reduces information costs and hence increases liquidity of the firm’s securities.

U.S. cross-listing should also improve the liquidity of the firm’s stock since it represents a firm’s

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24Bailey, Karolyi, and Salva (2006) find stronger absolute return reactions and volatility reactions to earnings announcements after the cross-list than before, especially among firms with higher disclosure standards. Kim and Verrecchia (2001) suggests that the slope coefficient of absolute return on trading volume provides a useful measure for the level of disclosure.
commitment to stricter disclosure, which mitigates asymmetric information between firm managers and investors. However, the effect on return autocorrelation through the microstructure channel is subtle. Investors who lack complete information about the future security payoff would try to infer it from current returns, which tends to strengthen the price changes caused by supply shocks or noise trading and produce a negative return autocorrelation (Brown and Jennings 1989, Wang 1993). However, when private information is long-lived, it can produce a positive return autocorrelation because information gets impounded into prices gradually over time (Wang 1994, Llorente, Michaely, Saar, and Wang 2002). Yet another complication, separate from the effect of asymmetric information, is that when risk-averse investors accommodate non-informational trades, the price pressure tends to be reversed over time, inducing a negative return autocorrelation (Campbell, Grossman, and Wang 1993, Pastor and Stambaugh 2003). Which of these effects ultimately prevails is an empirical question. It is therefore important to distinguish the effects of changes in disclosure policies from the effects of changes in market microstructure.

To control for these microstructure effects parsimoniously, we include Amihud’s (2002) illiquidity ratio in our VAR. The ratio is a price impact measure that is closely related to asymmetric information discussed above. Specifically, we augment our state vector as

\[ z_t = (\text{ret}_t, \text{roe}_t, (p-b)_t, ilq_t)' \]

with \( ilq_t = \ln(ILLIQ_t) \). Here, \( ILLIQ_t \) is Amihud’s (2002) illiquidity ratio constructed as

\[ ILLIQ_t = \frac{1}{\text{days}_t} \sum_{d=1}^{\text{days}_t} \frac{|r_{t,d}|}{VOLD_{t,d}} \]

where \( \text{days}_t \) is the number of days in year \( t \), \( r_{t,d} \) and \( VOLD_{t,d} \) are the daily stock return and daily trading volume (in millions of units in local currencies) on day \( d \) of year \( t \). Unfortunately,
the requirements for daily stock returns reduces the sample size from 227 firms to 170 firms that cross-list between 1989 and 2003. However, a VAR analysis with additional variable (ilq) and the smaller sample should provide a useful robustness check for our empirical results. As we can see in Panel A of Table 7, illiquidity of foreign firms declines (i.e., liquidity improves) on average after they cross-list on U.S. markets (see also Miller 1999, Foerster and Karolyi 1999).

Table 7 about here

Our VAR analysis allows us to examine the microstructure effects on foreign firms’ stock returns before and after the cross-list. For convenience, we call shocks representing the microstructure effects “illiquidity news,” which is measured as

\[
N_{itq,t} = (E_t - E_{t-1}) \left[ \sum_{j=1}^{\infty} \rho^j \times ilq_{t+j} \right] = e^{4t} \rho \Gamma (I - \rho \Gamma)^{-1} w_t.
\]

Panel B of Table 7 reports the covariances of \( N_{itq,t} \) with unexpected stock return, cash flow news, and expected return news. Illiquidity news is negatively correlated with cash flow news in all groups, meaning that liquidity expectations and cash flow expectations tend to move together. Meanwhile, illiquidity news is positively correlated with expected return news in all groups. This result is consistent with the notion that an increase in expected illiquidity is compensated by higher expected returns (Amihud 2002). In a projection

\[
N_{er,t} = \beta_{er|ilq} N_{itq,t} + \tilde{N}_{er,t}; \; N_{itq,t} \perp \tilde{N}_{er,t}, \tag{13}
\]

In the VAR, \( ilq \) does not show any significant lead-lag relation with \( ret \) or \( roe \). On the other hand, innovations to \( ilq \) exhibit strong contemporaneous negative correlation with innovations to \( ret \) and \( roe \). This is consistent with the notion that illiquidity effects arise at a higher frequency than our annual sample frequencies.
the coefficient $\beta_{er|ilq}$ is around 0.2 for all sample groups. For the whole sample, the term $\beta_{er|ilq} N_{ilq,t}$ explains 35% and 23% of the time-series variation of expected returns in the pre- and post-ADR periods respectively. However, more than 60% of the time-series variation of expected returns cannot be accounted for by the variation of illiquidity news.

Table 8 reproduces the covariance and correlation test results (Table 4) using the augmented VAR with $z_t = (\text{ret}_t, \text{roe}_t, (p-b)_t, \text{ilq}_t)'$. The sample period is shorter (1989-2003 as opposed to 1983-2003) and the sample size is smaller (170 total firms as opposed to 227 total firms) due to additional data requirements for daily stock returns and trading volume. Nevertheless, our conclusions from Table 4 remain intact in Table 8, indicating robustness of our results to changes in samples and the inclusion of the microstructure measure.

We measure microstructure effects by the term $\beta_{er|ilq} N_{ilq,t}$ in equation (13). We then ask to what extent it can account for the negative covariance between cash flow news and expected return news, and the shift in the covariance around the cross-list. Specifically, we consider the following decomposition,

$$\text{Cov}(\text{Ne}_{r,t}, \text{Ne}_{f,t}) = \beta_{er|ilq} \text{Cov}(N_{ilq,t}, \text{Ne}_{f,t}) + \text{Cov}(\tilde{N}_{er,t}, \text{Ne}_{f,t}),$$

where the term $\beta_{Ner|N_{ilq}} \text{Cov}(N_{ilq,t}, \text{Ne}_{f,t})$ represents the microstructure effects on the covariance between cash flow news and expected return news.

Table 9 reports $\text{Cov}(\text{Ne}_{r,t}, \text{Ne}_{f,t})$ and $\beta_{Ner|N_{ilq}} \text{Cov}(N_{ilq,t}, \text{Ne}_{f,t})$, both deflated by $\text{Var}(\tilde{r})$ as in Table 4. The microstructure effects account for a measurable fraction of the negative covariance,
and contribute more to the increase in $Cov(N_{er,t}, N_{cf,t})$ for firms that cross-list on NYSE, AMEX, or NASDAQ and for firms from emerging markets. Nevertheless, the lion’s share of $Cov(N_{er,t}, N_{cf,t})$ and its change is left unexplained. This result suggests that enforced disclosure can affect expected stock returns well beyond its microstructural effects.

3 Conclusion

In the presence of information asymmetry between a firm manager and investors, the manager has discretion to disclose or withhold his private information. In such a situation, the manager cannot credibly make full disclosure unless a mandatory regulation enforces it. Shin (2003) and other models of strategic/discretionary disclosure have shown that the manager typically discloses only favorable news and withholds unfavorable news in equilibrium.

Knowing this, investors try to infer the unfavorable information withheld by the manager. The degree of uncertainty about the hidden unfavorable information increases when the manager reports little favorable information, since investors do not know whether the manager is withholding a large amount of unfavorable information or he simply does not observe much. Because of this adverse selection problem, risk-averse investors require higher expected returns (information risk premium) when the firm reports poor performance (such as earnings and cash flows). This monotonic negative relation between the reported performance and investors’ uncertainty leads to a negative correlation between expected returns and expected cash flows. Consequently, the firm’s expected return will be more negatively correlated with its realized return than under full disclosure, implying a stronger return reversal.

We find compelling evidence for our model’s predictions. Applying a GMM version of Vuolteenaho’s
(2002) panel VAR to foreign firms that cross-list ADRs, we find a significant negative return autocorrelation (reversal) and a significant negative correlation between cash flow news and expected return news before the ADRs. Consistent with the predicted effects of the commitment to increased disclosure, U.S. cross-listing significantly reduces the return reversal and the negative correlation between the two news. This is stronger among the firms that raise disclosure standards substantially, namely, among the firms that comply with U.S. GAAP to cross-list on NYSE, AMEX, or NASDAQ (as opposed to OTC Pink Sheets or PORTAL). It is also more pronounced for firms from countries with emerging capital markets than for firms from more advanced economies. Furthermore, the reductions are more significant for firms from code-law countries than those from common-law countries, consistent with the notion that firm managers in code-law countries have more discretion in disclosure.

We also repeat our tests for the comparable firms from emerging markets that have not cross-listed in U.S. markets. We do not detect any statistically reliable shifts in the time-series of stock returns among the non-ADR comparable firms. Consequently, it is unlikely that other factors such as financial market developments explain our evidence. This additional exercise lends further support to our argument that changes in disclosure environments have significant effects on the time-series of stock returns and earnings upon cross-listing. Moreover, our evidence remains robust even after controlling for possible microstructure effects associated with disclosure changes.

Our analysis leaves some unresolved issues. First, while we show that disclosures matter for the time-series variation in expected returns, our single-firm model does not allow us to draw cross-sectional implications for expected returns. In practice, the degree of information asymmetries between firm managers and investors, as well as the litigation risks for withholding unfavorable
information, vary across industries and firms. Therefore, extending our model to multiple firms with heterogeneous information should help us achieve a better understanding of how strategic disclosure produces a cross-sectional difference in information risk premia. Second, although we discuss the potential effects that market microstructure has on our findings, our model assumes homogeneous investors and does not offer implications for trading or liquidity. Incorporating heterogeneity among investors is a natural and interesting extension for further study.

Finally, this paper presents significant and robust evidence consistent with the prediction of our theory. Nevertheless, our result should be interpreted cautiously as the effects of disclosure strategies on expected returns are difficult to isolate from other effects such as corporate governance structure and the quality of law enforcement to protect investors. Although this paper provides a new insight and evidence, there remains more to learn about the theoretical relations between disclosure strategies and expected returns.
Appendix

A Proofs

A.1 Proof of Theorem 1

**Date 0.** Since investors have no information at date 0, \( h(k|\mathcal{F}_0) = \binom{N}{k} r^k (1-r)^{N-k} \). Equation (2) becomes

\[
V_0^a = \frac{\sum_{k=0}^{N} \binom{N}{k} r^k (1-r)^{N-k} (u^k d^{N-k})^{1-\alpha}}{\sum_{k=0}^{N} \binom{N}{k} r^k (1-r)^{N-k} (u^k d^{N-k})^{-\alpha}}.
\]

Here, the denominator can be computed as

\[
\sum_{k=0}^{N} \binom{N}{k} (ru^{-\alpha})^k [(1-r)d^{-\alpha}]^{N-k} = [ru^{-\alpha} + (1-r)d^{-\alpha}]^N.
\]

Similarly, the numerator is given by \([ru^{1-\alpha} + (1-r)d^{1-\alpha}]^N\). Therefore,

\[
V_0^a = \left[ru^{1-\alpha} + (1-r)d^{1-\alpha}\right]^N / \left[ru^{-\alpha} + (1-r)d^{-\alpha}\right]^N \\
\equiv [\psi u + (1 - \psi)d]^N,
\]

where we have defined \( \psi \equiv ru^{-\alpha} / [ru^{-\alpha} + (1-r)d^{-\alpha}] \).

**Sanitization strategy.** When the manager follows the sanitization strategy, the investors’ information set at date 1 is given by \( \mathcal{F}_1 = \{s\} \). \( V_1^a(s) \) in Part (ii) is given in Shin’s (2003) Equation (17), with \( \pi \) in Equation (16) and \( q \) in Lemma 1. We note that \( h(k|\mathcal{F}_1) = h(k|s) \) is given in his Lemma 1. Given \( V_1^a(s) \), the manager’s optimal disclosure policy is to announce the observed
number of successes, $s$, and a zero failure, $f = 0$, because any other strategy will result in a lower price. For example, reducing $s$ by 1 would result in $u^{s-1}[\pi u+(1-\pi)d]^{N-s+1} < u^s[\pi u+(1-\pi)d]^{N-s}$.

**Full disclosure.** The following lemma gives the probability of $k$ terminal successes when the manager announces both the observed numbers of successes and failures, i.e., $\mathcal{F}_1 = \{s, f\}$.

**Lemma 5** When the manager makes full disclosure,

$$h(k|s, f) = \begin{cases} \binom{N-f-s}{k-s} r^{k-s} (1-r)^{N-f-k} & \text{if } s \leq k \leq N - f, \\ 0 & \text{otherwise.} \end{cases}$$

**Proof.** By the total probability theorem,

$$h(s, f, k) = h(k)h(s, f | k)$$

$$= \frac{\binom{N}{k} r^k (1-r)^{N-k} \theta^s (1-\theta)^{k-s} \theta^f (1-\theta)^{N-k-f}}{\binom{N}{k} r^k (1-r)^{N-k} \theta^s (1-\theta)^{k-s} \theta^f (1-\theta)^{N-k-f}}.$$

Considering the relation

$$\frac{h(k|s, f)}{h(k-1|s, f)} = \frac{h(s, f, k)/h(s, f)}{h(s, f, k-1)/h(s, f)},$$

we have

$$h(k|s, f) \propto \frac{1}{(k-s)!(N-k-f)!} r^k (1-r)^{N-k}.$$

Normalizing, we obtain the desired expression. ■

Intuitively, under full disclosure there is no skepticism about the manager hiding negative information, and therefore investors’ subjective probability in drawing successes is the original success probability, $r$, rather than the reduced success probability, $q$. 

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Then using the pricing equation (2),

\[ V_a^1(s, f) = \frac{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} r^{k-s}(1-r)^{N-f-k}(u^k d^{N-k})^{1-\alpha}}{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} r^{k-s}(1-r)^{N-f-k}(u^k d^{N-k})^{-\alpha}} \]

\[ = \frac{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} (ru^{1-\alpha})^{k-s}(1-r)^{d^{-\alpha}N-f-k}}{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} (ru^{-\alpha})^{k-s}(1-r)^{d^{-\alpha}N-f-k}} \cdot \frac{u^{(1-\alpha)s d^{(1-\alpha)f}}}{u^{-\alpha s d^{-\alpha f}}} \]

\[ = \frac{[ru^{1-\alpha} + (1-r)d^{1-\alpha}]^{N-s-f} \cdot u^{(1-\alpha)s d^{(1-\alpha)f}}}{[ru^{-\alpha} + (1-r)d^{-\alpha}]^{N-s-f}} \cdot u^{-\alpha s d^{-\alpha f}} \]

\[ = [\psi u + (1-\psi)d]^{N-s-f} \cdot u^s d^f. \]

\[ \blacksquare \]

A.2 Proof of Propositions 2, 3, and 4

We first derive the necessary moment expressions and then prove Proposition 3 followed by Propositions 2 and 4. Let \( h(s) = \binom{N}{s} (r\theta)^s (1-r\theta)^{N-s} \) be the unconditional probability of the manager announcing \( s \) successes at date 1. Then the expected first-period return under the sanitization strategy is given by

\[ E[R_1^a(s)] = \sum_{s=0}^{N} h(s) R_1^a(s) \]

\[ = \sum_{s=0}^{N} \binom{N}{s} (r\theta)^s (1-r\theta)^{N-s} u^s [\psi u + (1-\pi)d]^{N-s} \frac{[\psi u + (1-\psi)d]^N}{[\psi u + (1-\psi)d]^N} \]

\[ = \left[ r\theta \cdot u + (1-r\theta) \{\pi u + (1-\pi)d\} \right]^N \psi u + (1-\psi)d \]

\[ \equiv [r\theta \gamma_0 + (1-r\theta)\gamma_1]^N, \quad (14) \]

where we have defined

\[ \gamma_0 \equiv \frac{u}{\psi u + (1-\psi)d} > 1, \]

\[ \gamma_1 \equiv \frac{\pi u + (1-\pi)d}{\psi u + (1-\psi)d} < 1. \]
Similarly, we can calculate

$$E[E(R_2^2|s)] = \sum_{s=0}^{N} h(s) E(R_2^2|s)$$

$$= \left[ r\theta + (1-r\theta) \left( \frac{q u + (1-q)d}{\pi u + (1-\pi)d} \right) \right]^N$$

$$\equiv [r\theta + (1-r\theta)\gamma_2]^N > 1,$$

$$E[R_1^s(s) \cdot E(R_2^2|s)] = \sum_{s=0}^{N} h(s) \cdot R_1^s(s) E(R_2^2|s)$$

$$= \left[ \frac{r\theta \cdot u + (1-r\theta)\{q u + (1-q)d\}}{\psi u + (1-\psi)d} \right]^N$$

$$\equiv [r\theta\gamma_0 + (1-r\theta)\gamma_1\gamma_2]^N,$$

where we have defined

$$\gamma_2 \equiv \frac{qu + (1-q)d}{\pi u + (1-\pi)d} > 1.$$  

Next, let $$h(s,f) = \binom{N}{s} \binom{N-s}{f} (r\theta)^s [(1-r\theta)^f (1-\theta)^{N-s-f}$$ be the unconditional probability of the manager announcing $$s$$ successes and $$f$$ failures at date 1. We can also compute the corresponding moments under full disclosure,

$$E[R_1^s(s,f)] = \sum_{s=0}^{N} \sum_{f=0}^{N-s} h(s,f) \cdot R_1^s(s,f)$$

$$= \left[ \theta \cdot \frac{ru + (1-r)d}{\psi u + (1-\psi)d} + 1 - \theta \right]^N$$

$$\equiv [\theta\gamma_4 + 1 - \theta]^N > 1, \quad (15)$$

$$E[E(R_2^2|s,f)] = \sum_{s=0}^{N} \sum_{f=0}^{N-s} h(s,f) \cdot E(R_2^2|s,f)$$

$$= \left[ \theta + (1-\theta) \left( \frac{ru + (1-r)d}{\psi u + (1-\psi)d} \right) \right]^N$$

$$\equiv [\theta + (1-\theta)\gamma_4]^N > 1,$$

\[42\]
\[ E[R_1^a(s,f) \cdot E(R_2^a|s,f)] = \sum_{s=0}^{N} \sum_{f=0}^{N-s} h(s,f) \cdot R_1^a(s,f)E(R_2^a|s,f) \]
\[ = \left[ \frac{ru + (1-r)d}{\psi u + (1-\psi)d} \right]^N \]
\[ \equiv \gamma_4^N > 1, \]

where we have defined
\[ \gamma_4 \equiv \frac{ru + (1-r)d}{\psi u + (1-\psi)d} > 1. \]

\[ \text{A.2.1 Proof of Proposition 3} \]

We first show that \( \text{cov}(R_1^a(s,f), E[R_2^a|s,f]) < 0 \). Using the above result,
\[ \text{cov}(R_1^a(s,f), E(R_2^a|s,f)) = E[R_1^a(s,f) \cdot E(R_2^a|s,f)] - E[R_1^a(s,f)]E(R_2^a|s,f) \]
\[ = \gamma_4^N - [\theta \gamma_4 + 1 - \theta]^N[\theta + (1 - \theta)\gamma_4]^N. \]

Clearly, it suffices to show the case with \( N = 1 \),
\[ \gamma_4 - [\theta \gamma_4 + 1 - \theta][\theta + (1 - \theta)\gamma_4] = -\theta(1 - \theta)(\gamma_4 - 1)^2 < 0, \] (16)

which is true.

Next, using the earlier result, we can calculate the return autocovariance under full disclosure,
\[ \text{cov}(R_1^a(s), E(R_2^a|s)) \]
\[ = E[R_1^a(s) \cdot E(R_2^a|s)] - E[R_1^a(s)] \cdot E(E(R_2^a|s)) \]
\[ = [r\theta \gamma_0 + (1 - r\theta)\gamma_1 \gamma_2]^N - [r\theta \gamma_0 + (1 - r\theta)\gamma_1]^N \cdot [r\theta + (1 - r\theta)\gamma_2]^N. \] (17)

Again, when \( N = 1 \),
\[ [r\theta \gamma_0 + (1 - r\theta)\gamma_1 \gamma_2] - [r\theta \gamma_0 + (1 - r\theta)\gamma_1] \cdot [r\theta + (1 - r\theta)\gamma_2] \]
\[ = -r\theta(1 - r\theta)(\gamma_2 - 1)(\gamma_0 - \gamma_1) < 0, \] (18)
which also implies that $\text{cov}(R_1^a(s), E(R_2^a|s)) < 0$ for all $N$.

To show that $\text{cov}(R_1^a(s), E[R_2^a|s]) < \text{cov}(R_1^a(s, f), E[R_2^a|s, f])$, we first claim that it suffices to show the case with $N = 1$. To see this, note that by the law of iterated expectations,

$$E[R_1^a(s, f) \cdot E(R_2^a|s, f)] = E\left[ \frac{V^a_1(s, f)}{V_0^a} \cdot E\left( \frac{V^a_2}{V^a_1(s)|s, f} \right) \right]$$

$$= \frac{E[V^a_2]}{V_0^a}$$

$$= E\left[ \frac{V^a_1(s)}{V_0^a} \cdot E\left( \frac{V^a_2}{V^a_1(s)|s} \right) \right]$$

$$= E[R_1^a(s) \cdot E(R_2^a|s)],$$

or equivalently $r \theta \gamma_0 + (1 - r \theta) \gamma_1 \gamma_2 = \gamma_4$ (this can also be shown by a Brute-force calculation).

Thus, $\text{cov}(R_1^a(s), E[R_2^a|s]) < \text{cov}(R_1^a(s, f), E[R_2^a|s, f])$ if and only if

$$-E[R_1^a(s)] \cdot E[E(R_2^a|s)] + E[R_1^a(s, f)] \cdot E[E(R_2^a|s, f)] < 0. \quad (19)$$

But for all $N$, the sign of the left hand side of this inequality is identical to that with $N = 1$, which in turn equals the sign of the difference between the two covariances with $N = 1$ derived above.

From Equations (16) and (18), this difference is

$$-\theta(1 - \theta)(\gamma_4 - 1)^2 - [-r \theta(1 - r \theta)(\gamma_2 - 1)(\gamma_0 - \gamma_1)]$$

$$= \frac{\theta(u - d)^2}{\psi u + (1 - \pi)d} \cdot \left[ \frac{(1 - \theta)(r - \psi)^2}{\psi u + (1 - \psi)d} - \frac{r(1 - r \theta)(q - \pi)(1 - \pi)}{\pi u + (1 - \pi)d} \right],$$

where we have substituted the expressions for $\gamma_0$, $\gamma_1$, $\gamma_2$, and $\gamma_4$. Further substituting for the
definitions of $\pi$, $\psi$, and $q$, the square bracket in the last line can be rewritten as

$$(1 - \theta)\psi^2(1 - r)^2(\alpha - u^{-}\alpha).$$

$$
\left[ \frac{d^{-\alpha} - u^{-\alpha}}{[\psi u + (1 - \psi)d][ru^{-\alpha} + (1 - r)d^{-\alpha}]^2} - \frac{d^{-\alpha}}{[\pi u + (1 - \pi)d][(1 - \theta)ru^{-\alpha} + (1 - r)d^{-\alpha}]^2} \right]
< 0,
$$
after tedious but straightforward algebra. This is negative for $\alpha > 0$ because $d^{-\alpha} - u^{-\alpha} < d^{-\alpha}$, $\psi u + (1 - \psi)d > \pi u + (1 - \pi)d$, and $ru^{-\alpha} + (1 - r)d^{-\alpha} > (1 - \theta)ru^{-\alpha} + (1 - r)d^{-\alpha}$. ■

### A.2.2 Proof of Proposition 2

We show the following corollary first.

**Corollary 6** The expected first period return is lower under the sanitization strategy than under full disclosure; $E[R_1^0(s)] < E[R_1^0(s, f)]$.

**Proof.** Inspecting equations (14) and (15), we again see that it suffices to work with $N = 1$. In this case, $E[R_1^0(s)] - E[R_1^0(s, f)]$ can be written as

$$[r\theta\gamma_0 + (1 - r\theta)\gamma_1] - [\theta\gamma_4 + 1 - \theta] = \frac{[(1 - r\theta)\pi - (1 - \theta)\psi](u - d)}{\psi u + (1 - \psi)d} < 0$$

after straightforward manipulations, where we have used

$$(1 - r\theta)\pi = \frac{r(1 - \theta)u^{-\alpha}}{qu^{-\alpha} + (1 - q)d^{-\alpha}} < (1 - \theta)\psi$$

to sign the expression. ■

This corollary and inequality (19) imply that $E[E(R_2^0|s)] > E[E(R_2^0|s, f)]$. ■
A.2.3 Proof of Proposition 4

Part (i) immediately follows from straightforward algebra. To prove part (ii), note that the cash flow news and the expected return news are, respectively,

\[ N_{\text{cf}} = s \ln \left( \frac{u}{ru + (1-r)d} \right) + f \ln \left( \frac{d}{ru + (1-r)d} \right) + \text{const.} \]

\[ N_{\text{er}} = (s + f) \ln \left( \frac{\psi u + (1-\psi) d}{ru + (1-r)d} \right) + \text{const.} \]

Then, with \( n = s + f \), we have

\[ \text{Cov}(N_{\text{cf}}, N_{\text{er}}) = \text{Cov}(s, n) \times \ln \left( \frac{u}{ru + (1-r)d} \right) \times \ln \left( \frac{\psi u + (1-\psi) d}{ru + (1-r)d} \right) \]

\[ + \text{Cov}(f, n) \times \ln \left( \frac{d}{ru + (1-r)d} \right) \times \ln \left( \frac{\psi u + (1-\psi) d}{ru + (1-r)d} \right) . \]

Intuitively, \( s \) and \( f \) are drawn from binomial distributions conditional on \( n \). The following lemma proves that this intuition is true.

**Lemma 7** The probabilistic assumptions \( k \sim \text{Bin}(k|N, r) \), \( s|k \sim \text{Bin}(s|k, \theta) \) and \( f|k \sim \text{Bin}(f|N-k, \theta) \) imply, with \( n = s + f \),

\[ s|n \sim \text{Bin}(s|n, r), \ f|n \sim \text{Bin}(f|n, 1-r) . \]

**Proof.** The joint probability mass function for \( s, n, k \) (\( 0 \leq s \leq n \leq N, \ 0 \leq s \leq k \leq N \)) is

\[ \Pr(s, n, k) = \frac{\binom{k}{s} \binom{N-k}{n-s}}{\binom{N}{n}} \times \binom{N}{n} (\theta)^n (1-\theta)^{N-n} \times \binom{N}{k} r^k (1-r)^{N-k} . \]


Marginalizing over $k$ yields

$$Pr(s,n) = \binom{N}{n} \binom{n}{s} \theta^n \left(1 - \theta\right)^{N-n} \sum_{k=0}^{N} \left\{ \binom{k}{s} \binom{N-k}{n-s} \binom{r^k}{k} (1 - r)^{N-k} \right\}$$

$$= \binom{N}{n} \binom{n}{s} \theta^n (1 - \theta)^{N-n} r^s (1 - r)^{n-s} \sum_{k=0}^{N} \left\{ \binom{N-n}{k-s} r^{k-s} (1 - r)^{N-n-(k-s)} \right\}$$

$$= \binom{N}{n} \theta^n (1 - \theta)^{N-n} \binom{n}{s} r^s (1 - r)^{n-s}$$

binomial distribution $\text{Bin}(n;N,\theta)$ \hspace{1cm} binomial distribution $\text{Bin}(s;n,r)$

It follows that $Pr(s|n) = \binom{n}{s} r^s (1 - r)^{n-s}$ and $Pr(f|n) = \binom{n}{f} r^{n-f} (1 - r)^f$. 

>From Lemma 7 we can calculate $Cov(s,n)$ and $Cov(f,n)$ as

$$Cov(s,n) = E[Cov(s,n|n)] + Cov(E[s|n],n) = Cov(E[s|n],n)$$

$$= Cov(nr,n) = Var(n) r = N \theta (1 - \theta) r.$$

$$Cov(f,n) = Var(n) (1 - r) = N \theta (1 - \theta) (1 - r).$$

It then follows that

$$Cov(N_{ef},N_{er}) = N \theta (1 - \theta) \times \ln \left( \frac{\psi u + (1 - \psi) d}{ru + (1 - r) d} \right) \times \ln \left( \frac{u^d d^{1-r}}{ru + (1 - r) d} \right) > 0. \quad \blacksquare$$

**B Econometric Methodology**

**B.1 Dynamic Panel Data Estimation**

Recall that our firm-specific state vector for firm $t$, denoted by $z_{i,t}$, consists of three state variables:

$$z_{i,t} = (ret_{i,t},roe_{i,t},(p-b)_{i,t})',$$

where $ret_{i,t}$ and $roe_{i,t}$ are in excess of the local money market rate.

In order to explain the estimation of our panel VAR, let’s take the first state variable $ret_{i,t}$ as an
example. The single equation panel autoregressive model with individual effects can be written as

\[ ret_{i,t} = \gamma_{11} ret_{i,t-1} + \gamma_{12} \text{roe}_{i,t-1} + \gamma_{13} (p-b)_{i,t-1} + \eta_i + \tilde{v}_{i,t} \]

where \( \eta_i \) is individual firm-specific effect. \( \gamma_{11} \) denotes the \((1,1)\) element of the VAR coefficient matrix \( \Gamma \).\(^{26}\) We apply the following procedure to all the other equations in our panel VAR system.\(^{27}\)

It is well known that the traditional LSDV (least squares dummy variable) method is biased in the above panel autoregressive model with individual effects. To see this, denote the time mean of \( \bar{\tilde{v}}_{i,t} \) as \( \bar{v}_i = \sum_{t=1}^{T} \tilde{v}_{i,t} \). Simple within-group transformation would show that the strict exogeneity condition is violated when regressors include lagged dependent variables:

\[ \sum_{t=1}^{T} E[ret_{i,t-1}(\tilde{v}_{i,t} - \bar{v}_i)] \neq 0 \]

When the time dimension of the panel data \( T \) is small the biases will be very large regardless of the number of cross-sections.

To address this issue, we use the one-step GMM system estimator (Arellano and Bover, 1995, Blundell and Bond, 1998) to estimate the panel VAR system.\(^{28}\) The GMM system estimator employs two moment conditions to jointly estimate the regressions in transforms of the variables and regressions in levels. We use the past five available lagged endogenous variables as instruments in “transformed regressions” and the most recent lagged differences of endogenous variables in “level regressions”.

\(^{26}\) A univariate autoregression for \( ret_{i,t} \) is a special case in which \( \gamma_{12}, \gamma_{13}, \) and \( \gamma_{14} \) are set to zero.

\(^{27}\) We benefit greatly from Arellano’s DPD package written in OX (Doornik, 2002) in implementing our methodology.

\(^{28}\) Arellano and Bond (1991) and Blundell and Bond (1998) show that any inference based on the two-step estimators will not be reliable, since the asymptotic standard errors for the two-step estimators are not appropriate in typical sample sizes.
Specifically, our moment conditions for the “transformed regressions” are

\[ E \begin{bmatrix} \text{ret}_{i,t-\tau} \\ \text{roe}_{i,t-\tau} \otimes \tilde{\nu}_{i,t}^* \\ (p-b)_{i,t-\tau} \end{bmatrix} = 0 \text{ for } \tau = 2, \ldots, 6; \ t = 3, \ldots, T, \]

where \( \otimes \) denotes the Kronecker product and \( \tilde{\nu}_{i,t}^* \) is the residuals from the regressions on variables after taking orthogonal deviations,

\[ \tilde{\nu}_{i,t}^* = (\tilde{\nu}_{i,t} - \tilde{\nu}_{i,t+1} + \ldots + \tilde{\nu}_{i,T}) (\frac{T - t}{T - t + 1})^{1/2} \text{ for } t = 1, \ldots, T - 1. \]

Our moment conditions for the “level regressions” are

\[ E \begin{bmatrix} \text{ret}_{i,t-1} \\ \text{roe}_{i,t-1} \\ (p-b)_{i,t-1} \end{bmatrix} = 0 \text{ for } t = 3, \ldots, T, \text{ where } \tilde{\nu}_{i,t} = \eta_i + \tilde{\nu}_{i,t}. \]

**B.2 Jackknife**

In order to inspect the significance of shifts in disclosure policies with ADRs, we conduct our statistical inference using the jackknife method. The jackknife method makes use of systematic partitions of a dataset and is especially appropriate in a panel-data setting. As such, the sample is naturally partitioned into \( L \) firms. When we compare the shifts of parameters before and after the cross-list, we compute the differences in these parameters (denoted by \( K \)) using all the firms (for example, \( \gamma_{11}^{\text{post}} - \gamma_{11}^{\text{pre}} \)). Then we recalculate the differences in these parameters deleting one firm at a time. Denote by \( K_{(-j)} \) this estimator computed with the \( j^{th} \) firm removed from the total \( L \) firms. Then, we can calculate the jackknife standard errors as

\[
\sqrt{\text{Var}(K)}_j = \sqrt{\frac{\sum_{j=1}^{L} (K_j^* - J(K))^2}{L(L-1)}},
\]

49
where $K_j^* = LK - (L-1)K_{(-j)}$ (often called the “pseudo-values”) and $J(K) = \frac{1}{L} \sum_{j=1}^{L} K_j^*$. Using this, we compute the t-statistic as

$$\frac{K}{\sqrt{\text{Var}(K)_J}}.$$

Monte Carlo studies indicate that $\text{Var}(K)_J$ often overestimates the variance (Efron and Stein 1981, Efron 1982). Consequently, our t-test provides very conservative evidence against the null hypothesis.
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Table 1

Panel A: Number of sample firms by country of origin, trading location, and ADR effective year

<table>
<thead>
<tr>
<th>Developed Markets</th>
<th>ADR Effective Year</th>
<th>Total 1983-2003</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td></td>
</tr>
<tr>
<td>Belgium</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>11</td>
</tr>
<tr>
<td>Denmark</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>0</td>
</tr>
<tr>
<td>Finland</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>57</td>
</tr>
<tr>
<td>France</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>12</td>
</tr>
<tr>
<td>Germany</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>10</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>33</td>
</tr>
<tr>
<td>Ireland</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>2</td>
</tr>
<tr>
<td>Italy</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>2</td>
</tr>
<tr>
<td>Japan</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>14</td>
</tr>
<tr>
<td>Netherlands</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>8</td>
</tr>
<tr>
<td>New Zealand</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>3</td>
</tr>
<tr>
<td>Singapore</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>4</td>
</tr>
<tr>
<td>Spain</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>4</td>
</tr>
<tr>
<td>Sweden</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>2</td>
</tr>
<tr>
<td>Switzerland</td>
<td>Nyse/Amex/Nasdaq</td>
<td>1</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets / Portal</td>
<td>3</td>
</tr>
<tr>
<td>Developed Markets</td>
<td>Total</td>
<td>12</td>
</tr>
</tbody>
</table>
Table 1
Panel A (continued): Number of sample firms by country of origin, trading location, and ADR effective year

<table>
<thead>
<tr>
<th>Less-Developed/Emerging Markets</th>
<th>ADR Effective Year</th>
<th>Total</th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Greece</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>India</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Korea</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Malaysia</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Mexico</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Portugal</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Russia</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>South Africa</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Taiwan</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Turkey</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Venezuela</td>
<td>Nyse/Amex/Nasdaq</td>
<td>Pink Sheets / Portal</td>
</tr>
<tr>
<td>Emerging Markets Total</td>
<td></td>
<td>0</td>
</tr>
</tbody>
</table>

Notes: Panel A presents the distribution of our sample firms according to the effective dates reported by Bank of New York, the ADR trading locations, and the original countries they are from.
Table 2: Descriptives

<table>
<thead>
<tr>
<th>Variable</th>
<th>Group (N of firms)</th>
<th>Pre-ADR</th>
<th>Post-ADR</th>
<th>Post–Pre</th>
<th>Δmean</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>mean</td>
<td>st.dev.</td>
<td>mean</td>
<td>st.dev.</td>
</tr>
<tr>
<td>Excess Return</td>
<td>All (227 firms)</td>
<td>7.6</td>
<td>42.5</td>
<td>0.3</td>
<td>46.5</td>
</tr>
<tr>
<td></td>
<td>Nyse/Amex/Nasdaq (85)</td>
<td>9.1</td>
<td>40.4</td>
<td>0.1</td>
<td>47.3</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets/Portal (142)</td>
<td>6.9</td>
<td>43.8</td>
<td>0.4</td>
<td>47.8</td>
</tr>
<tr>
<td></td>
<td>Developed markets (170)</td>
<td>8.1</td>
<td>37.7</td>
<td>0.3</td>
<td>44.6</td>
</tr>
<tr>
<td></td>
<td>Emerging markets (57)</td>
<td>6.7</td>
<td>55.2</td>
<td>0.7</td>
<td>51.5</td>
</tr>
<tr>
<td>Excess ROE</td>
<td>All (227 firms)</td>
<td>5.5</td>
<td>22.3</td>
<td>4.8</td>
<td>13.7</td>
</tr>
<tr>
<td></td>
<td>Nyse/Amex/Nasdaq (85)</td>
<td>6.3</td>
<td>20.4</td>
<td>5.7</td>
<td>15.4</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets/Portal (142)</td>
<td>4.9</td>
<td>23.4</td>
<td>4.2</td>
<td>13.6</td>
</tr>
<tr>
<td></td>
<td>Developed markets (170)</td>
<td>5.6</td>
<td>21.7</td>
<td>4.8</td>
<td>13.5</td>
</tr>
<tr>
<td></td>
<td>Emerging markets (57)</td>
<td>4.9</td>
<td>24.1</td>
<td>4.2</td>
<td>14.3</td>
</tr>
<tr>
<td>ME/BE</td>
<td>All (227 firms)</td>
<td>.66</td>
<td>.78</td>
<td>.58</td>
<td>.71</td>
</tr>
<tr>
<td></td>
<td>Nyse/Amex/Nasdaq (85)</td>
<td>.70</td>
<td>.79</td>
<td>.70</td>
<td>.67</td>
</tr>
<tr>
<td></td>
<td>Pink Sheets/Portal (142)</td>
<td>.64</td>
<td>.76</td>
<td>.51</td>
<td>.72</td>
</tr>
<tr>
<td></td>
<td>Developed markets (170)</td>
<td>.67</td>
<td>.73</td>
<td>.59</td>
<td>.73</td>
</tr>
<tr>
<td></td>
<td>Emerging markets (57)</td>
<td>.66</td>
<td>.92</td>
<td>.53</td>
<td>.62</td>
</tr>
</tbody>
</table>

Notes: This table reports means and standard deviations for the variables of the available (pooled) firm-years. Δmean column reports changes in the pooled sample means.
Table 3: Return Reversal

<table>
<thead>
<tr>
<th>Group</th>
<th>Pre-ADR</th>
<th>Post-ADR</th>
<th>Change</th>
</tr>
</thead>
<tbody>
<tr>
<td>All</td>
<td>-.26</td>
<td>-.14</td>
<td>+.12</td>
</tr>
<tr>
<td>(227 firms)</td>
<td>-3.30***</td>
<td>-3.00***</td>
<td>1.18</td>
</tr>
<tr>
<td>Nyse/Amex/Nasdaq</td>
<td>-.25</td>
<td>-.08</td>
<td>+.17</td>
</tr>
<tr>
<td>(85 firms)</td>
<td>-3.71***</td>
<td>-1.23</td>
<td>2.06**</td>
</tr>
<tr>
<td>Pink Sheets/Portal</td>
<td>-.29</td>
<td>-.18</td>
<td>+.12</td>
</tr>
<tr>
<td>(142 firms)</td>
<td>-2.63***</td>
<td>-3.09***</td>
<td>0.77</td>
</tr>
<tr>
<td>Developed Markets</td>
<td>-.10</td>
<td>-.17</td>
<td>-.07</td>
</tr>
<tr>
<td>(170 firms)</td>
<td>-2.04**</td>
<td>-3.41***</td>
<td>-1.24</td>
</tr>
<tr>
<td>Emerging Markets</td>
<td>-.55</td>
<td>-.15</td>
<td>+.40</td>
</tr>
<tr>
<td>(57 firms)</td>
<td>-5.57***</td>
<td>-.165*</td>
<td>2.20**</td>
</tr>
</tbody>
</table>

Notes: This table reports AR(1) coefficients of a panel first-order autoregressive model for $r_{it}$ that controls for firm-specific effects, estimated for the pre- and post-ADR periods. t-statistics based on GMM standard errors are shown below the coefficient estimates. The last column reports changes in the AR(1) coefficients from the pre-ADR period to the post-ADR period. t-statistics based on jackknife standard errors are shown below the point estimates. ***, **, and * indicate significance at the 1%, 5% and 10% level.
Table 4: Covariance and Correlation Tests

<table>
<thead>
<tr>
<th>Group</th>
<th>Pre-</th>
<th>Post-</th>
<th>Change</th>
<th>Pre-</th>
<th>Post-</th>
<th>Change</th>
</tr>
</thead>
<tbody>
<tr>
<td>All (227 firms)</td>
<td>-.20</td>
<td>-.10</td>
<td>+.10</td>
<td>-.73</td>
<td>-.26</td>
<td>+.47</td>
</tr>
<tr>
<td>Nyse/Amex/Nasdaq (85 firms)</td>
<td>-.20</td>
<td>-.06</td>
<td>+.14</td>
<td>-.68</td>
<td>-.13</td>
<td>+.55</td>
</tr>
<tr>
<td>Pink Sheets/Portal (142 firms)</td>
<td>-.19</td>
<td>-.12</td>
<td>+.08</td>
<td>-.62</td>
<td>-.33</td>
<td>+.30</td>
</tr>
<tr>
<td>Developed Markets (170 firms)</td>
<td>-.17</td>
<td>-.11</td>
<td>+.06</td>
<td>-.54</td>
<td>-.30</td>
<td>+.24</td>
</tr>
<tr>
<td>Emerging Markets (57 firms)</td>
<td>-.24</td>
<td>-.11</td>
<td>+.13</td>
<td>-.96</td>
<td>-.27</td>
<td>+.69</td>
</tr>
</tbody>
</table>

Notes: This table reports \( \text{Cov}^* (N_{cf,t}, N_{er,t}) = \text{Cov} (N_{cf,t}, N_{er,t}) / \text{Var} (\tilde{r}_t) \) and \( \text{Corr} (N_{cf,t}, N_{er,t}) \) for the pre- and post-ADR periods as well as their changes. The results are based on GMM estimates of our firm-level panel VARs. t-statistics based on jackknife standard errors are shown below the point estimates. ***, **, and * indicate significance at the 1%, 5% and 10% level.
Table 5: Common Law Countries versus Code Law Countries

Panel A: Return Reversal (cf. Table 3)

<table>
<thead>
<tr>
<th>Legal origin</th>
<th>Pre-ADR</th>
<th>Post-ADR</th>
<th>Change</th>
</tr>
</thead>
<tbody>
<tr>
<td>Common Law</td>
<td>-.14</td>
<td>-.02</td>
<td>+.12</td>
</tr>
<tr>
<td>(76 firms)</td>
<td>-1.72*</td>
<td>-.42</td>
<td>.97</td>
</tr>
<tr>
<td>Code Law</td>
<td>-.20</td>
<td>+.10</td>
<td>+.29</td>
</tr>
<tr>
<td>(151 firms)</td>
<td>-2.13**</td>
<td>1.86*</td>
<td>2.40***</td>
</tr>
</tbody>
</table>

Panel B: Covariance and Correlation Tests (cf. Table 4)

<table>
<thead>
<tr>
<th>Legal origin</th>
<th>Cov*(N_{c,t}, N_{er,t})</th>
<th>Corr(N_{c,t}, N_{er,t})</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Pre-</td>
<td>Post-</td>
</tr>
<tr>
<td>Common Law</td>
<td>-.18</td>
<td>-.14</td>
</tr>
<tr>
<td>(76 firms)</td>
<td>-5.71***</td>
<td>-4.35***</td>
</tr>
<tr>
<td>Code Law</td>
<td>-.21</td>
<td>-.06</td>
</tr>
<tr>
<td>(151 firms)</td>
<td>-8.26***</td>
<td>-1.30</td>
</tr>
</tbody>
</table>

Notes: Panel A reports AR(1) coefficients of a panel first-order autoregressive model for ret_t that controls for firm-specific effects, estimated for the pre- and post-ADR periods. t-statistics based on GMM standard errors are shown below the coefficient estimates. The last column reports changes in the AR(1) coefficients from the pre-ADR period to the post-ADR period. t-statistics based on jackknife standard errors are shown below the point estimates. Panel B reports Cov*(N_{c,t}, N_{er,t}) = Cov(N_{c,t}, N_{er,t}) / Var(\tilde{r}_t) (\tilde{r}_t = ret_t - E_{t-1}ret_t) and Corr(N_{c,t}, N_{er,t}) for the pre- and post-ADR periods as well as their changes. The results are based on GMM estimates of our firm-level panel VARs. t-statistics based on jackknife standard errors are shown below the point estimates. ***, **, and * indicate significance at the 1%, 5% and 10% level. Common Law countries in our sample are: Australia, HK, India, Ireland, Malaysia, New Zealand, Singapore, and South Africa. Code Law countries in our sample are: Belgium, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Spain, Sweden, Switzerland, Austria, Greece, Korea, Mexico, Portugal, Russia, Taiwan, Turkey and Venezuela.
Table 6: Matching Sample Analysis: ADR firms versus Non-ADR firms from Emerging Markets

<table>
<thead>
<tr>
<th>Post ADR – Pre ADR differences</th>
<th>Non-ADR firms (120 firms)</th>
<th>ADR firms (57 firms)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Return auto-correlation</td>
<td>−0.10</td>
<td>+0.40</td>
</tr>
<tr>
<td>(cf. Table 3)</td>
<td>1.12</td>
<td>2.20**</td>
</tr>
<tr>
<td>$Cov^* (N_{cf,t}, N_{er,t})$</td>
<td>+0.07</td>
<td>+0.13</td>
</tr>
<tr>
<td>(cf. Table 4)</td>
<td>1.11</td>
<td>2.21**</td>
</tr>
<tr>
<td>$Corr (N_{cf,t}, N_{er,t})$</td>
<td>+0.30</td>
<td>+0.69</td>
</tr>
<tr>
<td>(cf. Table 4)</td>
<td>1.24</td>
<td>3.74***</td>
</tr>
</tbody>
</table>

Notes: This table summarizes the results from our matched sample analysis. Specifically, we match the ADR firms with their non-cross-listed local market competitors in each of the 12 emerging countries, by carefully matching the four-digit industry code on the Datastream/Worldscope (Field 06011). If no matching firm is available, then we match the first three digits. The ADR effective dates of the matching ADR firms are used as the pseudo-ADR effective dates for the non-ADR matched firms. We then repeat our estimation and hypothesis testing procedures for the non-ADR matched firms. The table reports changes in the AR(1) coefficients for stock returns, as well as changes in $Cov^* (N_{cf,t}, N_{er,t})$ and $Corr (N_{cf,t}, N_{er,t})$ from the (pseudo-)pre-ADR sample period to the (pseudo-)post-ADR sample period. t-statistics based on jackknife standard errors are shown below the point estimates. ***, **, and * indicate significance at the 1%, 5% and 10% level. The last column summarizes the corresponding test results for the ADR firms, that are reported in Tables 3, 4, and 5.
### Table 7: Microstructure Effects

#### Panel A: Descriptives for \( ilq \)

<table>
<thead>
<tr>
<th>Group (N of firms)</th>
<th>Pre-ADR</th>
<th>Post-ADR</th>
<th>Post – Pre</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>mean</td>
<td>st.dev.</td>
<td>mean</td>
</tr>
<tr>
<td>All (170 firms)</td>
<td>-5.9</td>
<td>3.1</td>
<td>-6.9</td>
</tr>
<tr>
<td>Nyse/Amex/Nasdaq (67)</td>
<td>-6.5</td>
<td>3.2</td>
<td>-8.1</td>
</tr>
<tr>
<td>Pink Sheets/Portal (103)</td>
<td>-5.6</td>
<td>2.9</td>
<td>-6.2</td>
</tr>
<tr>
<td>Developed markets (127)</td>
<td>-6.3</td>
<td>3.0</td>
<td>-7.2</td>
</tr>
<tr>
<td>Emerging markets (43)</td>
<td>-4.8</td>
<td>3.1</td>
<td>-6.2</td>
</tr>
</tbody>
</table>

#### Panel B: Relation between Illiquidity News (\( N_{ilq} \)) and Stock Return Components

<table>
<thead>
<tr>
<th>Groups</th>
<th>Period</th>
<th>Covariances of ( N_{ilq,t} ) with ( \tilde{r}<em>t ), ( N</em>{cf,t} ), ( N_{er,t} )</th>
<th>Regression</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>( N_{ilq,t} )</td>
<td>( \tilde{r}_t )</td>
</tr>
<tr>
<td>All (170 firms)</td>
<td>Pre-ADR</td>
<td>.45</td>
<td>-.11</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>.33</td>
<td>-.04</td>
</tr>
<tr>
<td>Nyse/Amex/Nasdaq (67)</td>
<td>Pre-ADR</td>
<td>.28</td>
<td>-.11</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>.42</td>
<td>-.17</td>
</tr>
<tr>
<td>Pink Sheets/Portal (103)</td>
<td>Pre-ADR</td>
<td>.39</td>
<td>-.10</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>.26</td>
<td>-.09</td>
</tr>
<tr>
<td>Developed Mkts (127)</td>
<td>Pre-ADR</td>
<td>.54</td>
<td>-.12</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>.35</td>
<td>-.12</td>
</tr>
<tr>
<td>Emerging Mkts (43)</td>
<td>Pre-ADR</td>
<td>.23</td>
<td>-.13</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>.59</td>
<td>-.13</td>
</tr>
</tbody>
</table>

Notes: Panel A reports means and standard deviations for \( ilq = \ln (ILLIQ) \) of the available (pooled) firm-years. \( \Delta \text{mean} \) column reports changes in the pooled sample means. Panel B reports variances and covariances of the illiquidity news (\( N_{ilq} \)) with unexpected stock returns (\( \tilde{r} \)), cash flow news (\( N_{cf} \)), and expected return news (\( N_{er} \)) that are identified in a firm-level panel VAR with \( z_{i,t} = (\text{ret}_{i,t}, \text{roe}_{i,t}, (p-b)_{i,t}, \text{ilq}_{i,t})' \). All variances and covariances are deflated by \( V\text{ar}(\tilde{r}) \). The last two columns report coefficients and \( R^2 \)'s for the regression: \( N_{er} = \beta_{er|ilq} N_{ilq} + \tilde{N}_{er} \). The figures are calculated from the GMM estimates of our firm-level panel VAR.
Table 8: Covariance and Correlation Tests (VAR with \(ilq\))

<table>
<thead>
<tr>
<th>Group</th>
<th>(\text{Cov}^* (N_{cf,t}, N_{er,t}))</th>
<th>(\text{Corr} (N_{cf,t}, N_{er,t}))</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Pre-</td>
<td>Post-</td>
</tr>
<tr>
<td>All</td>
<td>-0.20</td>
<td>-0.12</td>
</tr>
<tr>
<td>(170 firms)</td>
<td>-10.2***</td>
<td>-4.02***</td>
</tr>
<tr>
<td>Nyse/Amex/Nasdaq (67 firms)</td>
<td>-0.19</td>
<td>-0.07</td>
</tr>
<tr>
<td>Pink Sheets/Portal (103 firms)</td>
<td>-3.44***</td>
<td>-3.68***</td>
</tr>
<tr>
<td>Developed Markets (127 firms)</td>
<td>-0.15</td>
<td>-0.12</td>
</tr>
<tr>
<td>Emerging Markets (43 firms)</td>
<td>-0.23</td>
<td>-0.10</td>
</tr>
</tbody>
</table>

Notes: This table reports \(\text{Cov}^* (N_{cf,t}, N_{er,t}) = \text{Cov} (N_{cf,t}, N_{er,t}) / \text{Var} (\hat{r}_t) (\hat{r}_t = \text{ret}_t - E_{t-1}\text{ret}_t)\) and \(\text{Corr} (N_{cf,t}, N_{er,t})\) for the pre- and post-ADR periods as well as their changes. t-statistics based on jackknife standard errors are shown below the point estimates. The figures are based on GMM estimates of the augmented firm-level panel VAR with \(z_{i,t} = (\text{ret}_{i,t}, \text{roe}_{i,t}, (p-b)_{i,t}, ilq_{i,t})'\). ***, **, and * indicate significance at the 1%, 5% and 10% level.
Table 9: Significance of Microstructure Effects for $\text{Cov}^* (N_{cf,t}, N_{er,t})$

<table>
<thead>
<tr>
<th>Group</th>
<th>Periods</th>
<th>Microstructure Effects</th>
<th>Other Effects</th>
</tr>
</thead>
<tbody>
<tr>
<td>All (170 firms)</td>
<td>Pre-ADR</td>
<td>-.09</td>
<td>-.11</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>-.07</td>
<td>-.05</td>
</tr>
<tr>
<td></td>
<td>Post – Pre</td>
<td>+.02</td>
<td>+.07</td>
</tr>
<tr>
<td>Nyse/Amex/Nasdaq (67 firms)</td>
<td>Pre-ADR</td>
<td>-.08</td>
<td>-.11</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>-.02</td>
<td>-.05</td>
</tr>
<tr>
<td></td>
<td>Post – Pre</td>
<td>+.06</td>
<td>+.06</td>
</tr>
<tr>
<td>Pink Sheets/Portal (103 firms)</td>
<td>Pre-ADR</td>
<td>-.04</td>
<td>-.13</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>-.04</td>
<td>-.09</td>
</tr>
<tr>
<td></td>
<td>Post – Pre</td>
<td>+.00</td>
<td>+.04</td>
</tr>
<tr>
<td>Developed Markets (127 firms)</td>
<td>Pre-ADR</td>
<td>-.04</td>
<td>-.11</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>-.06</td>
<td>-.06</td>
</tr>
<tr>
<td></td>
<td>Post – Pre</td>
<td>+.02</td>
<td>+.01</td>
</tr>
<tr>
<td>Emerging Markets (43 firms)</td>
<td>Pre-ADR</td>
<td>-.09</td>
<td>-.14</td>
</tr>
<tr>
<td></td>
<td>Post-ADR</td>
<td>-.04</td>
<td>-.06</td>
</tr>
<tr>
<td></td>
<td>Post – Pre</td>
<td>+.05</td>
<td>+.08</td>
</tr>
</tbody>
</table>

Notes: This table reports the microstructure effects, $b_{N_{er,t}} N_{ilq,t} \text{Cov} (N_{ilq,t}, N_{cf,t}) / \text{Var} (\tilde{r}_t)$; $\tilde{r}_t = ret_t - E_{t-1} ret_t$, and other effects $\text{Cov} (N_{er,t} - b_{N_{er,t}} N_{ilq,t}, N_{cf,t}) / \text{Var} (\tilde{r}_t)$, estimated for both pre- and post-ADR periods as well as their changes. The figures are calculated from the GMM point estimates of the augmented firm-level panel VAR with $z_{i,t} = (ret_{i,t}, roe_{i,t}, (p-b)_{i,t}, ilq_{i,t})'$. 

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