Strategic Disclosure and Stock Returns: Theory and Evidence from US Cross-Listing

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When a firm exercises discretion to disclose or withhold information (strategic disclosure), risk-averse investors command higher expected returns when expected cash flows decrease, producing a negative correlation between these expectations. 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 American Depositary Receipts (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 less-developed/emerging markets and code-law countries. (*JEL* G14, G15, F30)

The integrity of corporate disclosure sustains investor 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 the strategic decisions of self-interested managers, as it involves a number of estimates, judgments, and assumptions. Thus, managers may have an incentive to obfuscate the firm's true performance, within the allowances of investor-protection regulations.

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This paper provides a theoretical framework to analyze the effects of firm disclosure strategies on stock returns. Its implications are then tested with a sample of foreign firms that have listed American Depositary Receipts (ADRs).[1] 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, a firm has multiple projects, each with two possible outcomes: success or failure. The firm's manager observes some of these outcomes, while investors observe only his disclosure. The manager has discretion to disclose or withhold what he observes, as what information he possesses 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.

Providing a complete treatment of full disclosure under risk aversion, we extend Shin's (2003) analysis to investigate the effects of strategic disclosure on the time-series behavior of stock returns in comparison to those of full disclosure. We propose that, when risk-averse investors are uncertain about the manager's information and the manager does not credibly make full disclosure, investors require higher expected returns as expected cash flows decrease. As a result, changes in cash-flow expectations ("cash-flow news") and changes in expected returns ("expected-return news") are negatively correlated. Intuitively, as the manager reports fewer successes, investors become increasingly skeptical regarding the informativeness of his disclosure, because they do not know whether he is withholding more failures or if he simply does not observe many outcomes. This skepticism compels them to 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. Moreover, the skepticism also produces a stronger, negative return (return reversal) under strategic disclosure, as a poorer reported performance makes the investors demand a greater price discount, which tends to be reversed subsequently.

While we show these results using a simple two-period model, we illustrate how they carry over to a multiperiod setting. We also emphasize that the intuition behind our theory does not depend sensitively on the specific model we employ. In fact, the compelling prediction of the discretionary disclosure literature is the existence of a threshold information level above which a manager discloses and below which he withholds his information.[2] The common implication is that investors become more skeptical about the manager's

[1] ADRs are created by US-based depositary banks as claims on non-US firms' equity shares in their home markets, and are traded just like US stocks.

[2] See, for example, Verrecchia (1983, 1990); Dye (1985); and Jung and Kwon (1988). Verrecchia's (2001) Section 3 provides an extensive review of the literature.

disclosure when he discloses less good news. Therefore, a negative relation between the reported firm performance and information risk premium would naturally follow as a robust outcome of the uncertainty regarding the manager's information.[3] To our knowledge, no other existing studies analyze the link between cash-flow expectations and expected returns along this line.

We test our predictions using a panel of foreign firms that listed ADRs between 1983 and 2003. 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.[4] Furthermore, expected litigation costs for withholding unfavorable information are much higher in the United States than in other countries (see, e.g., Skinner, 1994, 1997). To the extent that these conditions are true, negative news correlation and return reversal should characterize stock returns before the ADRs, but these return properties should be significantly attenuated afterward.

We employ the firm-level panel vector-autoregression (VAR) framework of Vuolteenaho (2002) to examine potential shifts in the time-series properties of foreign firms' stock returns before and after the ADRs. We identify cash-flow news and expected-return news from the VAR (also see Campbell, 1991). We test our predictions with a generalized method of moments (GMM) estimator (see, e.g., Arellano and Bond, 1991; Arellano and Bover, 1995; and Blundell and Bond, 1998) and make statistical inferences using conservative Jackknife standard errors.

The results provide compelling evidence for our theoretical predictions. We find that the negative correlation between cash-flow news and expected-return news indeed diminishes substantially upon cross-listing; the correlation changes from −0.73 to −0.26, and the difference is statistically significant at the 1% level. We also find a significant reduction in return reversal upon cross-listing. Both the negative news correlation and the return reversal are further reduced for firms that experience large shifts in accounting standards; that is, more for (i) firms that cross-list on NYSE, AMEX, or NASDAQ (US GAAP compliance required) than firms that cross-list on over-the-counter (OTC) Pink Sheets or PORTAL (the market for firms issuing equity under SEC Rule 144a; US GAAP compliance not required); (ii) firms from less-developed/emerging capital markets than those from advanced capital markets; and (iii) firms from countries with code-law traditions than those with common-law traditions, which allow less room for managers to withhold unfavorable information (Ball,

[3] We thank an anonymous referee for this insight.

[4] Non-US firms listing on major US trading venues (i.e., NYSE, AMEX, or 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 with the SEC, where the Form 20-F provides a reconciliation with US GAAP. On the other hand, firms that cross-list on over-the-counter (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 US GAAP. Nevertheless, they usually need to disclose more than the level mandated by the home market regulators.

Kothari, and Robin, 2000). Our results are robust to possible changes in illiquidity levels associated with cross-listing and to general market developments that may affect expected returns.

Using the US data, Rogers, Schrand, and Verrecchia (2006) test Shin's (2003) prediction for the negative return–volatility relation (the leverage effect) under strategic disclosure. Their choice of data takes advantage of the observation that the Private Securities Litigation Reform Act of 1995 is supposed to have increased managers' incentives to engage in strategic disclosure. Our work differs from theirs in two fundamental ways. First, we derive and test a novel implication of Shin's (2003) model, rather than examine some of his original predictions. Second, we look at the sample of firms that we believe experience maximal changes in Shin's (2003) key assumptions, namely, foreign firms that cross-list in the United States rather than US firms that undergo a legal reform. In the international cross-listing literature, Bailey, Karolyi, and Salva (2006) show that changes in the cross-listing firm's disclosure environment significantly affect the way that its stock return and trading volume respond to its announced earnings. We add to their contributions by studying another channel through which changes in disclosure environment affect stock returns. 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 firm's expected return. These papers attribute their findings to Merton's (1987) investor recognition hypothesis.[5] 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. However, none of the studies cited in this paragraph analyzes potential shifts in the time-series properties of stock returns when these firms subject themselves to the more stringent disclosure environment in US markets.

The rest of the paper is organized as follows. In Section 1, we present our theory and derive its testable implications. Section 2 explains our data, provides empirical evidence, and conducts robustness tests. Section 3 concludes. The appendices provide 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 prices and returns. 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 show, the risk aversion of investors produces

[5] See also Errunza and Miller (2000); and Sarkissian and Schill (2008) for alternative explanations.

rich dynamics in stock returns when the manager has discretion to disclose or withhold information.

1.1 Setup

The notation used here generally follows 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. 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 θ 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 as 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) where $0 \leq s_1 \leq s$ and $0 \leq f_1 \leq f$, because it will be consistent with any terminal outcome 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 of the unobserved project outcomes turn out to be failures at date 2.

The implicit understanding is that the manager's disclosure must be accompanied by verifiable evidence. If the manager knows that a project has failed, he cannot claim that it has succeeded, since antifraud 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 the time of disclosure is not verifiable even at a later date. This assumption is at the heart of Shin's (2003) model, and implies that disclosure is discretionary and endogenous rather than mandatory and exogenous.[6]

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_t^a, $t = 0, 1, 2$. Throughout this paper, superscript a represents

[6] The vast literature on strategic or discretionary disclosure relies on the verifiability assumption as it can be justified by appealing to the potential litigation and human capital erosion costs associated with dissembling (Verrecchia, 2001). Admati and Pfleiderer (2000) also note that 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.

a risk-averse quantity and distinguishes it from its risk-neutral counterpart. The stock pays in units of a consumption good, which serves as numeraire of the economy. Without loss of generality, we normalize the per capita 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, $\mathcal{F}_t$:

$$\mathbf{E}_t[U(W_2)|\mathcal{F}_t], \tag{1}$$

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.[7]

1.2 Firm value

We examine two manager strategies: one in which the manager discloses only the number of successful projects, and another in which he additionally discloses the number of failures. Shin (2003) calls the former "the sanitization strategy" and the latter "full disclosure."[8] Under the sanitization strategy, the manager chooses not to disclose negative information that reduces firm value.[9] Given the manager's 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 using the standard solution technique. The assumption of a complete market implies the existence of state prices,

$$\varphi_t^a = \frac{h(k|\mathcal{F}_t)U'(u^k d^{N-k})}{\sum_k h(k|\mathcal{F}_t)U'(u^k d^{N-k})}, \tag{2}$$

where $h(k|\mathcal{F}_t)$ is the probability of k terminal successes assessed at date $t = 0, 1$. The firm's value is then given by

$$V_t^a = \sum_k \varphi_t^a u^k d^{N-k}, \tag{3}$$

7 Recently, Shin (2006) extends his own (2003) framework to make both the number of projects (N) and the observation probability (θ) random. He also considers the case with risk-averse investors. His objective is to explain short-run return continuation (drift) and long-run reversal following earnings announcement by way of heightened risk during the disclosure period. In contrast, we show in a simpler model that investors' risk aversion and skepticism alone generate novel and interesting implications for the time series of stock returns.

8 We will use the term "strategic disclosure" to mean the sanitization strategy. See Rogers, Schrand, and Verrecchia (2006) for a similar treatment.

9 Consistent with this, more than 90% 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.

where $t = 0, 1$, 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 no matter what disclosure strategy the manager will follow at date 1. 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. *(Firm value) (i) The firm value at date* 0 *is given by*

$$V_0^a = [\psi u + (1 - \psi)d]^N, \tag{4}$$

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

(ii) (Sanitization strategy, Shin, 2003) 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*

$$V_1^a(s) = u^s[\pi u + (1 - \pi)d]^{N-s}, \tag{5}$$

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

(iii) (Full disclosure) The firm value at date 1 *when the manager discloses the observed number of both successes, s, and failures, f, is given by*

$$V_1^a(s, f) = u^s d^f[\psi u + (1 - \psi)d]^{N-s-f}. \tag{6}$$

Proof. All proofs are contained in Appendix A. ■

The price function in Equation (4) is intuitive. It is the product of the expected values of N independent projects, where each project's expected value is computed by a modified success probability, ψ. 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 in the "up" and "down" states.

The date 1 price in Equation (5) 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 π, which is a combination of the effects of risk aversion and such skepticism. These two effects are roughly captured by the differences in success probabilities, $r - \psi$ and $\psi - \pi$, respectively.

Finally, Equation (6), under full disclosure, prices the $N - s - f$ undisclosed projects with the original modified success probability, ψ. Note that, as Shin (2003) shows for the risk-neutral case, this will not constitute an equilibrium price unless full disclosure is enforced exogenously, say by regulators. To see this point, suppose the price is given by Equation (6). However, the manager then has an incentive to deviate from his strategy by hiding all 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 ψ to the success of undisclosed projects. Therefore, the manager cannot credibly achieve full disclosure without a mandatory stipulation.

We emphasize that this result does not sensitively depend on the specific assumptions of our model. In fact, the seminal conclusion of the vast literature on strategic or discretionary disclosure is that, in the presence of uncertainty or disclosure costs (broadly defined), a partial disclosure equilibrium can be sustained in which a manager strategically withholds bad news and discloses good news (e.g., see Verrecchia, 2001 for a review). We build on Shin's (2003) setup because it allows us to draw clear predictions about the time-series properties of stock returns. Specifically, it allows a simple and intuitive decomposition of stock returns into changes in expected payoff (expected cash flows) and changes in expected return (information risk premium), which is the subject of the next subsection.

1.3 Disclosure and time-series properties of stock returns

Using the price functions derived above, we can examine the properties of returns in each of the two disclosure regimes. In the following, $R_1^a \equiv V_1^a / V_0^a$ and $R_2^a \equiv V_2^a / V_1^a$ denote the gross return on the firm's equity in the first and second periods, respectively. When necessary, we distinguish returns under the two disclosure strategies by their arguments. For example, $R_1^a(s)$ and $R_1^a(s, f)$ denote the first-period returns under the sanitization strategy and full disclosure, respectively.

1.3.1 Theoretical predictions. The manager's disclosure affects both investors' expectations about the firm's future payoff (cash flow) and expected return. 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 number of disclosed failures, f, as well as the number of disclosed successes, s.[10]

By definition, stock returns are driven by cash-flow news and expected-return news (Campbell, 1991). Let us consider the following decomposition of the log first-period return:

$$\ln \tilde{R}_1^a = N_{cf} - N_{er}, \tag{7}$$

where $\ln \tilde{R}_1^a \equiv \ln R_1^a - E_0[\ln R_1^a]$. N_{cf} and N_{er} denote cash-flow news and expected-return news, respectively. Notice that if investors were risk-neutral,

[10] To see this formally, compute the conditional second-period returns $E(R_2^a|s) = (\frac{qu+(1-q)d}{\pi u+(1-\pi)d})^{N-s}$ under the sanitization strategy (Shin, 2003, p. 121) and $E(R_2^a|s, f) = (\frac{ru+(1-r)d}{\psi u+(1-\psi)d})^{N-s-f}$ under full disclosure.

there should be no time variation in expected returns, and hence $N_{er} = 0$. Thus, we define cash-flow news as the first-period return under risk-neutrality, i.e., $N_{cf} = \ln \tilde{R}_1 \equiv \ln R_1 - E_0[\ln R_1]$. Then, expected-return news is given by $N_{er} = \ln \tilde{R}_1 - \ln \tilde{R}_1^a$.

Based on decomposition (7), our objective is to show how expected cash flows interact with expected returns (information risk premium). The following proposition establishes that cash-flow news and expected-return news covary differently under the two disclosure policies.

Proposition 1. *(News covariance) (i) Under the sanitization strategy, cash-flow news and expected-return news are negatively correlated:* $Cov(N_{cf}, N_{er}) < 0$. *(ii) Under full disclosure, cash-flow news and expected-return news are positively correlated:* $Cov(N_{cf}, N_{er}) > 0$.

The 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 as to 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 changes in cash-flow expectations and changes in 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, which removes the price discount due to skepticism. Second, an increase in the reported number of failures will decrease not only the expected cash flow, but also the price discount due to risk aversion because it partially resolves uncertainty about the future payoff. The result in part (ii) of the proposition indicates that these two effects collectively overturn the negative correlation to a positive one.

Proposition 1 highlights the effect of investor skepticism on stock returns; a poorer disclosed performance exacerbates the investors' uncertainty about the manager's hidden unfavorable information, and hence investors require higher expected returns. This effect tends to be subsequently reversed as more information becomes available to the investors. It is then natural to conjecture that stock returns will exhibit a stronger reversal (negative autocovariance) under strategic disclosure than under full disclosure. The following proposition establishes that this conjecture is indeed true.

Proposition 2. *(Return reversal) The stock return exhibits a stronger negative autocovariance under the sanitization strategy than under full disclosure;* $cov(R_1^a(s), R_2^a(s)) < cov(R_1^a(s, f), R_2^a(s, f))$.

1.3.2 A multiperiod extension. While the two-period model in the previous section offers sharp predictions with a minimal setup, we emphasize that similar results hold in a multiperiod extension. The most direct extension would be an overlapping-generations model in which the events of the two-period model repeat indefinitely. The primary difference from the two-period model is that the investors' wealth includes the selling price of the stock, as well as the dividend. Despite this difference, we can show that the stock price is proportional to the price from the two-period model. This implies that Propositions 1 and 2 remain valid within a generation. At the turn of the generation, the news covariance is zero under both disclosure policies simply because there is no disclosure; there is nothing that changes expected returns. The return autocovariance is also zero as long as disclosures in two successive generations are independent. To summarize, the two propositions continue to hold in an overlapping-generations model if we relax the strict inequalities to weak inequalities.[11] In this way, the intuition from our two-period model also applies in a dynamic setting.

1.4 Empirical hypotheses

A large body of literature on ADR or US cross-listing documents empirical evidence that foreign firms' expected returns decrease substantially when they cross-list on US markets (e.g., Foerster and Karolyi, 1999; Miller, 1999; Errunza and Miller, 2000; Hail and Leuz, 2006a, b; and Sarkissian and Schill, 2008). In addition, the reduction is larger for firms that tend to undergo more substantial shifts in disclosure policies around cross-listing, i.e., firms that cross-list on NYSE, AMEX, or NASDAQ and firms from countries with high investment barriers. We add to this literature by proposing and testing new theoretical implications for the effects of strategic disclosure formalized in Propositions 1 and 2.

We examine these predictions in a panel of foreign firms that cross-list on US markets via the ADR programs. Specifically, we compare the time-series behavior of stock returns before and after cross-listing. Since disclosure requirements and investor-protection regulations are much stricter in the US markets than in other markets, US cross-listing by a foreign firm is generally regarded as a canonical event that increases its commitment to fuller disclosure.

The main empirical 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 cross-listing.

[11] A formal treatment of the multiperiod extension is available in the technical appendix.

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 than for firms that cross-list on OTC Pink Sheets or PORTAL, and diminish more for firms from less-developed/emerging capital markets than for firms from well-developed capital markets. We also examine if the effects differ between firms from code-law countries and firms from common-law countries.

1.5 Identifying cash-flow news and expected-return news

The primary objective of our empirical analysis is to estimate the effects of changes in disclosure regimes on the time-series properties of stock returns. The key of this exercise is to decompose stock returns into cash-flow news and expected-return news (Campbell, 1991). In our model, the former corresponds to changes in the investors' expectation of the firm's payoff (cash flows), while the latter corresponds to the implicit discount the investors require due to the uncertainty of the manager's information. To empirically identify these two news components, we rely on Vuolteenaho's (2002) accounting-based dynamic present-value relation, which implies the following return decomposition as an empirical counterpart of Equation (7): $r_t - E_{t-1}r_t = N_{cf,t} - N_{er,t}$, where

$$N_{cf,t} = (E_t - E_{t-1})\left[\sum_{j=0}^{\infty} \rho^j \times roe_{t+j} + \xi_t\right], \tag{8}$$

$$N_{er,t} = (E_t - E_{t-1})\left[\sum_{j=1}^{\infty} \rho^j \times ret_{t+j}\right]. \tag{9}$$

In Equations (8) and (9), roe_t is the logarithm of one plus return on equity (ROE), i.e., earnings per book value of equity, and ret_t is the local log stock return in excess of the local interest rate. Henceforth, we use the terms ROE and stock return to mean excess ROE and excess return, respectively. The term ξ_t reflects possible effects of extraordinary items on ROE, the time variation in the dividend payout ratio, and the approximation error due to log linearization. ρ is a constant close to but less than 1. We set $\rho = 0.967$ as in Vuolteenaho (2002), but a small variation in ρ has no material effects on our conclusions. $N_{cf,t}$ and $N_{er,t}$, as defined in Equations (8) and (9), are respectively the empirical counterparts of cash-flow news (N_{cf}) and expected-return news (N_{er}) discussed in Section 1.3.1.

As in Cohen, Gompers, and Vuolteenaho (2002); and Vuolteenaho (2002), we use a firm-level panel VAR to model the conditional expectation E_t. Let us express the VAR for each firm as

$$z_t = \Gamma z_{t-1} + \eta + w_t;\ \Sigma = E[w_t w_t'], \tag{10}$$

where Γ and Σ are assumed to be constant, both across time and across firms in each subperiod, and η is a vector of individual firm-specific effects. Following Vuolteenaho (2002), we set our state vector as $z_t = (ret_t, roe_t, (p\text{-}b)_t)'$, where $(p\text{-}b)_t$ is the log market-to-book (ME/BE) ratio. Vuolteenaho's model implies that the ME/BE ratio contains useful information about future ROEs and future stock returns.

With this specification, we can identify expected-return news and cash-flow news as $N_{er,t} = e1'\Upsilon w_t$ and $N_{cf,t} = e1'(I + \Upsilon)w_t$, 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., $e1'z_t = ret_t$. It then follows that

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

To put the implications of our strategic disclosure model into practice, we first estimate a firm-level panel VAR (10) for the pre- and post-ADR periods separately. Since traditional least-squares estimators are biased in panel autoregressions, especially in a short panel like ours, we implement the so-called GMM system estimator (Arellano and Bover, 1995; and Blundell and Bond, 1998) to estimate our panel VAR. We then inspect the significance of the shifts in the covariance (11) associated with ADRs. We test our hypotheses based on the Jackknife standard errors that are known to be conservative (Efron and Stein, 1981; and Efron, 1982) and robust to outliers. Appendix B explains our econometric methodology in detail.

2. Evidence

2.1 Data and descriptives

We obtain information about ADR listings from the Bank of New York. Our firm-level panel is from Datastream/Worldscope, and covers 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 the effective date to the last available year. Since our focus is on the time-series behavior of stock returns and cash flows around cross-listing, we only retain firms that have three or more consecutive years of data in each of the pre- and post-ADR subperiods. 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 obtain local annual interest rates at the end of the previous year, yearly local total stock returns, annual accounting ROEs, and the market-to-book ratios from Datastream/Worldscope.[12] 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 well-developed markets and firms from less-developed/emerging markets.[13] We exclude UK firms from our analysis since UK markets have a distinctly superior information environment for US investors than other markets, and the London Stock Exchange offers a popular alternative venue for international cross-listing.[14] We also exclude three countries, Argentina, Brazil, and Peru, because they had years of triple-digit annual inflation rates during our sample period. Past ROE and ME/BE become meaningless under high inflation, because both depend on the book value of equity that reflects accumulated earnings, and past earnings are accumulated at historical costs without adjusting for inflation over time.

Table 1 summarizes the number of firms in the sample by country and effective year. Our final sample consists of 2,636 firm-year observations for 227 firms from 28 markets that cross-listed on US markets between 1983 and 2003;[15] 170 firms are from 16 well-developed markets and 57 firms are from 12 less-developed/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.[16] The number of ADR listings generally grew steadily until 2001 and has decreased since 2002, the year in which the Sarbanes-Oxley Act took effect. We also classify firms into two categories by trading venues for their ADRs: firms that list ADRs on NYSE, AMEX, or NASDAQ (85 firms), and firms that list ADRs on OTC Pink Sheets or PORTAL (142 firms). This classification is important for our study since NYSE, AMEX, and NASDAQ require compliance

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

[13] See Table 1 for our classification. To qualify as well developed, a country must have a per capita GNP of at least \$12,600 (New Zealand; in 1994 US constant dollars) and its accounting standard index (La Porta et al., 1998) must be at least 60. Note that Austria is categorized as less developed, since its accounting standard index (54) is among the lowest in our sample, although it has a high per capita GDP (\$23,510). Leuz, Nanda, and Wysocki (2003) also report that, in a sample of 31 countries, Austria allows the highest degree of flexibility for managers to engage in earnings management. However, the grouping of Austria has minimal effects on our estimation results for both well-developed and less-developed markets.

[14] Our sample does not include Canadian firms since they can use Canadian GAAP to directly list on US markets. In fact, Canadian GAAP is very similar to US GAAP (Bailey, Karolyi, and Salva, 2006, footnote 11).

[15] Among the 227 firms, 25 firms (11% of the sample) are banks. Besides this, we do not see any particular industry concentration in our sample. Excluding banks from the sample does not change our conclusions.

[16] Japan liberalized its financial market in 1983 (Bekaert and Harvey, 2000). Excluding this or another year does not change our conclusion.

Table 1
Number of sample firms by country of origin, trading location, and ADR effective year

Well-developed markets		ADR effective year																					Total
Country of origin	Trading location	1983	84	85	86	87	88	89	90	91	92	93	94	95	96	97	98	99	2000	01	02	03	1983–2003
Australia	NYSE/AMEX/NASDAQ															1							1
(Common-law)	Pink Sheets/Portal								1	1	2	1				1			1		1		8
Belgium	NYSE/AMEX/NASDAQ																			1			1
(Code-law)	Pink Sheets/Portal												1										1
Denmark	NYSE/AMEX/NASDAQ									1											1		2
(Code-law)	Pink Sheets/Portal																						0
Finland	NYSE/AMEX/NASDAQ											1											1
(Code-law)	Pink Sheets/Portal											1											1
France	NYSE/AMEX/NASDAQ									1					2	2			3	3	1		12
(Code-law)	Pink Sheets/Portal	1							1		1	2	3							1	1		10
Germany	NYSE/AMEX/NASDAQ															1			4	2	3		10
(Code-law)	Pink Sheets/Portal	1					1							1				2	1	3	1	1	11
Hong Kong	NYSE/AMEX/NASDAQ																			1			1
(Common-law)	Pink Sheets/Portal										1	7	4	5	2	2		2	4	2	3	1	33
Ireland	NYSE/AMEX/NASDAQ																						0
(Common-law)	Pink Sheets/Portal																	1			1		2
Italy	NYSE/AMEX/NASDAQ							1									1						2
(Code-law)	Pink Sheets/Portal																		1				1
Japan	NYSE/AMEX/NASDAQ	1								1	1		1			1	1	1		3	4		14
(Code-law)	Pink Sheets/Portal	9					1			1	2	2	1	1		3	1		1	2		2	26
Netherlands	NYSE/AMEX/NASDAQ		1			1										3	1			2			8
(Code-law)	Pink Sheets/Portal							1												1		1	3
New Zealand	NYSE/AMEX/NASDAQ																				1		1
(Common-law)	Pink Sheets/Portal																						0
Singapore	NYSE/AMEX/NASDAQ																						0
(Common-law)	Pink Sheets/Portal									1	1			1		1							4
Spain	NYSE/AMEX/NASDAQ														2				2				4
(Code-law)	Pink Sheets/Portal																1		1				2
Sweden	NYSE/AMEX/NASDAQ								1														1
(Code-law)	Pink Sheets/Portal								1					1									2
Switzerland	NYSE/AMEX/NASDAQ													1		1			2	1			5
(Code-law)	Pink Sheets/Portal										1			1	1								3
	Well-developed markets total	12	1	0	0	1	2	2	4	6	9	14	10	11	7	16	5	6	20	22	17	5	170

Less-developed/Emerging markets		ADR effective year																					Total
Country of origin	Trading location	1983	84	85	86	87	88	89	90	91	92	93	94	95	96	97	98	99	2000	01	02	03	1983–2003
Austria	NYSE/AMEX/NASDAQ																						0
(Code-law)	Pink Sheets/Portal														2	2	1			1	1		7
Greece	NYSE/AMEX/NASDAQ																	1			1		2
(Code-law)	Pink Sheets/Portal																1						1
India	NYSE/AMEX/NASDAQ																		1	4			5
(Common-law)	Pink Sheets/Portal																						0
Korea	NYSE/AMEX/NASDAQ														1								1
(Code-law)	Pink Sheets/Portal																				1		1
Malaysia	NYSE/AMEX/NASDAQ																						0
(Common-law)	Pink Sheets/Portal																1	1	1		1		4
Mexico	NYSE/AMEX/NASDAQ																1	1			2		4
(Code-law)	Pink Sheets/Portal													1		1			1				3
Portugal	NYSE/AMEX/NASDAQ																						0
(Code-law)	Pink Sheets/Portal																1						1
Russia	NYSE/AMEX/NASDAQ																				1		1
(Code-law)	Pink Sheets/Portal																		1	2			3
South Africa	NYSE/AMEX/NASDAQ												1		1		2				1		5
(Common-law)	Pink Sheets/Portal												4			1		1		3	1	2	12
Taiwan	NYSE/AMEX/NASDAQ																		3				3
(Code-law)	Pink Sheets/Portal																						0
Turkey	NYSE/AMEX/NASDAQ																						0
(Code-law)	Pink Sheets/Portal												1				1						2
Venezuela	NYSE/AMEX/NASDAQ																		1				1
(Code-law)	Pink Sheets/Portal																1						1
	Less-developed markets total	0	0	0	0	0	0	0	0	0	0	0	6	1	4	4	9	4	8	10	9	2	57

Number of sample firms by country of origin, trading location, and ADR effective year. This table 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. Each country's legal origin (common-law or code-law tradition) is also shown in parentheses.

Table 2
Descriptives

Variable	Group (*N* of firms)	Pre-ADR Mean	Pre-ADR St. dev.	Post-ADR Mean	Post-ADR St. dev.	Post−Pre Δ*mean*
Excess return	All (227 firms)	7.6	42.5	0.3	46.5	−7.9
(log return	NYSE/AMEX/NASDAQ (85)	9.1	40.4	0.1	47.3	−9.0
in annual %)	Pink Sheets/PORTAL (142)	6.9	43.8	0.4	47.8	−7.3
	Well-developed markets (170)	8.1	37.7	0.3	44.6	−8.4
	Less-developed markets (57)	6.7	55.2	0.7	51.5	−6.0
Excess ROE	All (227 firms)	5.5	22.3	4.8	13.7	−0.7
(log return	NYSE/AMEX/NASDAQ (85)	6.3	20.4	5.7	15.4	−0.5
in annual %)	Pink Sheets/PORTAL (142)	4.9	23.4	4.2	13.6	−0.7
	Well-developed markets (170)	5.6	21.7	4.8	13.5	−0.8
	Less-developed markets (57)	4.9	24.1	4.2	14.3	−0.7
ME/BE	All (227 firms)	0.66	0.78	0.58	0.71	−0.09
(in log)	NYSE/AMEX/NASDAQ (85)	0.70	0.79	0.70	0.67	+0.00
	Pink Sheets/PORTAL (142)	0.64	0.76	0.51	0.72	−0.13
	Well-developed markets (170)	0.67	0.73	0.59	0.73	−0.07
	Less-developed markets (57)	0.66	0.92	0.53	0.62	−0.13

This table reports means and standard deviations of the constructed variables for the pooled firm-years. The last column reports changes in the pooled sample means.

with US GAAP for cross-listing firms, while OTC and PORTAL do not.[17] We also control for firms switching from one ADR type to another.[18]

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 annual excess returns in local currencies are about 8% lower in the post-ADR period than in the pre-ADR period for the whole sample. This observation is consistent with the argument that ADR listing decreases firms' expected returns significantly (Hail and Leuz, 2006a, b). It is also consistent with the observation of prelisting overperformance and significant reduction in the cost of capital after cross-listing (e.g., Foerster and Karolyi, 1999; Miller, 1999; Errunza and Miller, 2000; and Sarkissian and Schill, 2008). On the other hand, average excess ROEs for the whole sample decline only slightly (0.7%) after cross-listing. Meanwhile, the standard deviation of ROEs drops from 22.3% to 13.7% after cross-listing in the whole sample.[19]

17 Since 1999, the SEC mandates firms to comply with US GAAP to be quoted on the OTC Bulletin Board. This does not affect our analysis since none of our sample ADRs are traded on the OTC Bulletin Board.

18 Two firms in our sample, National Australia Bank (Australia) and Anglogold Ashanti (South Africa), listed their ADRs on the NYSE first and subsequently used Rule 144a private placements (RADRs). We categorize them as NYSE/AMEX/NASDAQ firms. Ten other firms in our sample switched their trading locations between PORTAL (for RADRs) and OTC Pink Sheets, which do not affect their OTC/PORTAL classification. None of our sample firms cross-listed on the OTC/PORTAL first and switched to the NYSE/AMEX/NASDAQ later.

19 Consistent with the notion that the quality of disclosure improves when foreign firms cross-list on US markets, we find a significantly stronger association between ROEs and lagged stock returns after cross-listing than before. In other words, the reported earnings become more value-relevant after cross-listing. This result is consistent with the findings of Lang, Raedy, and Yetman (2003); and Lang, Raedy, and Wilson (2006).

2.2 Test results

This section first states our empirical predictions (see Section 1.4) as statistical hypotheses. We then test the hypotheses and examine whether the effects are stronger for firms that cross-list on the NYSE, AMEX, or NASDAQ than for firms that cross-list on other venues that do not require US GAAP compliance. We also test whether the effects are stronger for firms from less-developed/emerging markets than for firms from developed markets.

2.2.1 Relation between cash-flow news and expected-return news. Our theory generates a set of implications for the covariance between cash-flow news (N_{cf}) and expected-return news (N_{er}). Relying on Vuolteenaho (2002), we identify the two news components using a firm-level panel VAR and estimate their covariances and correlations before and after cross-listing.[20] 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-1: $\quad Cov^*(N_{er,t}, N_{cf,t})^{pre} < 0,$
Hypothesis A-1': $\quad Corr(N_{er,t}, N_{cf,t})^{pre} < 0,$

where the superscript "*pre*" (and "*post*" below) represents the periods before (after) cross-listing. 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 the following hypotheses:

Hypothesis B-1: $\quad Cov^*(N_{er,t}, N_{cf,t})^{pre} < Cov^*(N_{er,t}, N_{cf,t})^{post},$
Hypothesis B-1': $\quad Corr(N_{er,t}, N_{cf,t})^{pre} < Corr(N_{er,t}, N_{cf,t})^{post},$

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

Table 3 presents our results with Jackknife standard errors. The cash-flow news and expected-return news are significantly negatively correlated at about -0.2 for all groups of ADR firms before cross-listing. 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 less-developed/emerging market firms.

After cross-listing, 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. The increase in the correlation coefficient is stronger and statistically significant for firms that cross-list on the NYSE, AMEX, or NASDAQ ($+0.55$, $t = 2.5$) than those that cross-list on the OTC or PORTAL ($+0.30$, $t = 1.6$), and for firms from

[20] Note that we calculate the news covariance according to Equation (11) and do not extract the news time series.

Table 3
Covariance and correlation tests

	$Cov^*(N_{cf,t}, N_{er,t})$			$Corr(N_{cf,t}, N_{er,t})$		
Group	Pre-	Post-	Change	Pre-	Post-	Change
All	−0.20	−0.10	+0.10	−0.73	−0.26	+0.47
(227 firms)	−11.8***	−3.40***	2.76***	−5.80***	−2.72***	2.73***
NYSE/AMEX/NASDAQ	−0.20	−0.06	+0.14	−0.68	−0.13	+0.55
(85 firms)	−12.2***	−0.74	1.89*	−5.57***	0.64	2.49**
Pink Sheets/PORTAL	−0.19	−0.12	+0.08	−0.62	−0.33	+0.30
(142 firms)	−6.47***	−3.79***	1.78*	−3.97***	−3.11***	1.59
Well-developed markets	−0.17	−0.11	+0.06	−0.54	−0.30	+0.24
(170 firms)	−5.94***	−3.77***	1.31	−3.64***	−2.87***	1.22
Less-developed markets	−0.24	−0.11	+0.13	−0.96	−0.27	+0.69
(57 firms)	−14.6***	−2.03**	2.21**	−12.8***	−1.59	3.74***

This table reports $Cov^*(N_{cf,t}, N_{er,t}) \equiv Cov(N_{cf,t}, N_{er,t})/Var(\tilde{r}_t)$, where $\tilde{r}_t \equiv ret_t - E_{t-1}ret_t$, and $Corr(N_{cf,t}, N_{er,t})$ 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 1%, 5%, and 10%, respectively.

less-developed/emerging markets $(+0.69, t = 3.7)$ than those from well-developed markets $(+0.24, t = 1.2)$.

Recall that Proposition 1 implies that the covariance between cash-flow news and expected-return news is positive under full disclosure. In Table 3, 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 the NYSE, AMEX, or NASDAQ. This result suggests that the time-series properties of post-ADR stock returns are not fully consistent with full disclosure.[21] Nevertheless, our evidence for Hypotheses B-1 and B-1' lends support to our argument that the effects of strategic disclosure significantly decrease after cross-listing. Interestingly, the evidence of the negative news covariance contrasts to the positive news covariance that Cohen, Gompers, and Vuolteenaho (2002) document for US small stocks.[22]

2.2.2 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,

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

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

[21] This may not be surprising given that fraudulent disclosures are possible even under US GAAP, as illustrated by the recent corporate accounting scandals by major corporations. Meanwhile, La Porta et al. (2000); Siegel (2005); and Lang, Raedy, and Wilson (2006) discuss some limitations of SEC regulatory enforcement for cross-listed foreign firms.

[22] Cohen, Gompers, and Vuolteenaho (2002) attribute the positive covariance between cash-flow news and expected-return news of the US small firms to "under-reaction," which institutional investors appear to exploit only partially. On the other hand, US large firms do not exhibit the "under-reaction" phenomenon. Most of the foreign firms that cross-list in US markets are among the largest firms in their respective home markets to which the under-reaction story is unlikely to apply.

Table 4
Return reversal

Group	Pre-ADR	Post-ADR	Change
All	−0.26	−0.14	+0.12
(227 firms)	−3.30***	−3.00***	1.18
NYSE/AMEX/NASDAQ	−0.25	−0.08	+0.17
(85 firms)	−3.71***	−1.23	2.06**
Pink Sheets/PORTAL	−0.29	−0.18	+0.12
(142 firms)	−2.63***	−3.09***	0.77
Well-developed markets	−0.10	−0.17	−0.07
(170 firms)	−2.04**	−3.41***	−1.24
Less-developed markets	−0.55	−0.15	+0.40
(57 firms)	−5.57***	−1.65*	2.20**

This table 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. ***, **, and * indicate significance at 1%, 5%, and 10%, respectively.

and B are

$$\text{Hypothesis A-2:} \qquad \phi_{ret}^{pre} < 0.$$

$$\text{Hypothesis B-2:} \qquad \phi_{ret}^{post} - \phi_{ret}^{pre} > 0.$$

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

Table 4 presents strong evidence for return reversal, i.e., $\phi_{ret}^{pre} < 0$, for all groups of firms in the pre-ADR period. For the whole sample before cross-listing, the autoregressive coefficient ϕ_{ret} is significantly negative at −0.26 ($t = -3.3$). The magnitude of negative autocorrelation is much stronger for firms from countries with less-developed/emerging markets ($\phi_{ret} = -0.55$, $t = -5.6$) than for those from well-developed markets ($\phi_{ret} = -0.10$, $t = -2.0$). This result is not unusual since emerging markets tend to have weaker disclosure standards.

The return reversal persists, but diminishes significantly after cross-listing. We are able to reject the null hypothesis of $\phi_{ret}^{post} = \phi_{ret}^{pre}$ for firms that cross-list on the NYSE, AMEX, or NASDAQ and for firms from less-developed/emerging markets at a 5% significance level. Cross-listings of these firms are likely to convey stronger signals about their commitments to disclosing information more truthfully. Thus, this evidence supports our hypothesis that return reversal is stronger in weaker regulatory environments, where investors presume firms' strategic disclosure behavior, but the reversal tends to diminish as firms credibly convince investors of their commitment to increased disclosure.

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 et al., 1998). In particular, it has been documented that countries with common-law traditions provide stronger protection for shareholders than countries with code-law traditions.

Table 5
Common-law countries versus code-law countries

Panel A: Covariance and correlation tests (cf. Table 4)

	$Cov^*(N_{cf,t}, N_{er,t})$			$Corr(N_{cf,t}, N_{er,t})$		
Legal origin	Pre-	Post-	Change	Pre-	Post-	Change
Common-law	−0.18	−0.14	+0.04	−0.63	−0.39	+0.23
(76 firms)	−5.71***	−4.35***	0.93	−2.83***	−3.27***	0.91
Code-law	−0.21	−0.06	+0.15	−0.71	−0.14	+0.57
(151 firms)	−8.26***	−1.30	2.62***	−5.68***	−1.12	3.01***

Panel B: Return reversal (cf. Table 3)

Legal origin	Pre-ADR	Post-ADR	Change
Common-law	−0.23	−0.25	−0.02
(76 firms)	−3.14***	−4.44***	−0.18
Code-law	−0.28	−0.07	+0.21
(151 firms)	−2.62***	−1.36	1.36

Panel A reports $Cov^*(N_{cf,t}, N_{er,t}) \equiv Cov(N_{cf,t}, N_{er,t})/Var(\tilde{r}_t)$, where $\tilde{r}_t \equiv ret_t - E_{t-1}ret_t$, and $Corr(N_{cf,t}, N_{er,t})$ for the pre- and post-ADR periods, as well as their changes. The results are based on GMM estimates of the firm-level panel VARs. t-statistics based on Jackknife standard errors are shown below the point estimates. Panel B 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. ***,**, and * indicate significance at 1%, 5%, and 10%, respectively. Common-law countries in our sample are Australia, Hong Kong, 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.

In the accounting literature, Ball, Kothari, and Robin (2000) present evidence that firms in countries with code-law traditions are more likely to withhold bad news than firms in countries with common-law traditions. 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. If this is true, 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 using firms from countries with different legal origins, following the classification of La Porta et al. (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 standardized covariances and correlation coefficients between cash-flow news and expected-return news for the two groups before and after cross-listing. Again, the increase in the correlation coefficient, $Corr(N_{er,t}, N_{cf,t})^{post} - Corr(N_{er,t}, N_{cf,t})^{pre}$, is more significant for firms from code-law countries ($+0.57$, $t = 3.0$) than those from common-law countries ($+0.23, t = 0.91$). Panel B of Table 5, comparable to Table 4, reports the estimated autoregressive coefficient from regression (12). The reduction in the magnitude of return reversal, $\phi_{ret}^{post} - \phi_{ret}^{pre}$, is -0.02 ($t = -0.18$) for firms from common-law countries, and 0.21 ($t = 1.4$) for firms from code-law countries. Generally, these results are consistent with our prediction that US 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 substantial shifts in the time-series properties of stock returns around 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, such as developments in local capital markets. In order to inspect this possibility, we focus on less-developed/emerging markets and repeat our tests using a sample of noncross-listing firms in the same industries and countries as our ADR firms. Specifically, we match the ADR firms with their noncross-listing local market competitors in each of the 12 less-developed/emerging markets by carefully matching the four-digit industry code used in Datastream/Worldscope.[23] If no matching firm is available, we match the first three digits. Using 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 pseudo pre- and post-ADR periods. For these firms, we apply the same data requirements (e.g., at least three years of data in each of the 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 noncross-listing local market competitors of our ADR firms, with 620 and 692 firm-year observations in pseudo pre- and post-ADR periods, respectively.

Table 6 summarizes the results from our matched sample analysis. The magnitude of return reversal diminishes after cross-listing only for the ADR firms. We do not observe a similar decline in return reversal for non-ADR matched firms. The autocorrelation exhibits a decline, albeit insignificant, for the non-ADR firms. Similarly, we can detect a statistically significant increase in the correlation between cash-flow news and expected-return news for the ADR firms, but not for the non-ADR firms.

In summary, while we report significant evidence for shifts in the time-series properties of stock returns around cross-listing, we do not find any reliable evidence that such shifts also occur in the matched sample of non-ADR firms. Therefore, it is unlikely that other factors, such as financial market development, can account for our results.

2.3.3 Interaction with liquidity effects. Although our theory offers no implications on trading or liquidity, we are aware that changes in market

[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 subindustry groups.

Table 6
Matched sample analysis: ADR firms versus non-ADR firms from less-developed/emerging capital markets

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

This table summarizes the results from the matched sample analysis. We match the ADR firms with their noncross-listing local market competitors in each of the 12 less-developed/emerging markets, 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 the estimation and hypothesis testing for the non-ADR matched firms. The table reports changes in $Cov^*(N_{cf,t}, N_{er,t})$ and $Corr(N_{cf,t}, N_{er,t})$, as well as changes in the AR(1) coefficients for stock returns 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 1%, 5%, and 10%, respectively. The last column summarizes the corresponding test results for the ADR firms reported in Tables 3 and 4.

microstructures associated with cross-listing 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) also argue that an increase in a firm's commitment to more disclosure should lower the price impact of a trade.[24] Leuz and Verrecchia (2000) support this argument by showing that German firms that list their stocks on the Neuer Market, which requires higher disclosure standards 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 the liquidity of the firm's securities.

US cross-listing should also improve the liquidity of a firm's stock, since it represents the firm's 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 may try to infer it from current prices, which tends to amplify the price changes caused by supply shocks or noise trading and produce a negative-return autocorrelation (Brown and Jennings, 1989; and 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; and Llorente et al., 2002).[25] Yet another complication, separate from the effect of

[24] Bailey, Karolyi, and Salva (2006) find stronger volatility and volume reactions to earnings announcements after cross-listing than before, especially among firms with higher disclosure standards. Kim and Verrecchia (2001) suggest that the slope coefficient in the regression of absolute return on trading volume provides a useful measure of the level of disclosure.

[25] Using foreign stocks cross-listed on US markets, Gagnon and Karolyi (2007) provide evidence of return spillover across international stock markets along the lines of Llorente et al. (2002).

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; and 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 liquidity.

As an attempt to control for these liquidity effects parsimoniously, we include Amihud's (2002) illiquidity ratio in our VAR. The ratio is a price-impact measure that is closely related to the asymmetric information discussed above. We hope that its inclusion in the VAR system allows us to at least partially isolate the effect of illiquidity due to asymmetric information from the effect of investor skepticism. Specifically, we augment the state vector as $z_t = (ret_t, roe_t, (p\text{-}b)_t, ilq_t)'$ with $ilq_t = \ln(ILLIQ_t)$. Here, $ILLIQ_t$ is Amihud's (2002) illiquidity ratio constructed as

$$ILLIQ_t = \frac{1}{days_t} \sum_{d=1}^{days_t} \frac{|r_{t,d}|}{VOLD_{t,d}}, \tag{13}$$

where $days_t$ is the number of days in year t, and $r_{t,d}$ and $VOLD_{t,d}$ are the daily stock return and daily trading volume (in millions of units in local currencies), respectively, on day d of year t. Unfortunately, the availability of daily stock returns, prices, and volume reduces the sample size from 227 firms to 170 firms that cross-list between 1989 and 2003. However, the VAR analysis with an additional variable (ilq) and a 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 US markets.

Our VAR analysis allows us to examine the liquidity effects on foreign firms' stock returns before and after cross-listing. For convenience, we call shocks representing the illiquidity effects "illiquidity news," which is measured as

$$N_{ilq,t} = (E_t - E_{t-1}) \left[\sum_{j=1}^{\infty} \rho^j \times ilq_{t+j} \right] = e4'\rho\Gamma(I - \rho\Gamma)^{-1} w_t, \tag{14}$$

where $e4$ is a choice vector with 1 in its fourth element and 0 elsewhere.

Panel B of Table 7 reports the covariances of $N_{ilq,t}$ with unexpected stock return, cash-flow news, and expected-return news. Illiquidity news is negatively correlated with cash-flow news in all groups, implying 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

$$N_{er,t} = \beta_{er|ilq} N_{ilq,t} + \tilde{N}_{er,t}, \tag{15}$$

Table 7
Liquidity effects

Panel A: Descriptives for ilq

Group (N of firms)	Pre-ADR Mean	Pre-ADR St. dev.	Post-ADR Mean	Post-ADR St. dev.	Post − Pre Δ mean
All (170 firms)	−5.9	3.1	−6.9	3.0	−1.0
NYSE/AMEX/NASDAQ (67)	−6.5	3.2	−8.1	2.7	−1.6
Pink Sheets/PORTAL (103)	−5.6	2.9	−6.2	2.9	−0.6
Well-developed markets (127)	−6.3	3.0	−7.2	3.0	−0.9
Less-developed markets (43)	−4.8	3.1	−6.2	2.9	−1.4

Panel B: Relation between illiquidity news (N_{ilq}) and stock return components

		Covariances of $N_{ilq,t}$ with				Regression	
Groups	Period	$N_{ilq,t}$	$\tilde{r}_t$	$N_{cf,t}$	$N_{er,t}$	$\beta_{er\|ilq}$	R^2
All	Pre-ADR	0.45	−0.11	−0.04	+0.06	0.14	0.35
(170 firms)	Post-ADR	0.33	−0.04	−0.04	+0.07	0.22	0.23
NYSE/AMEX/NASDAQ	Pre-ADR	0.28	−0.11	−0.08	+0.04	0.13	0.29
(67 firms)	Post-ADR	0.42	−0.17	−0.08	+0.10	0.23	0.28
Pink Sheets/PORTAL	Pre-ADR	0.39	−0.10	−0.02	+0.08	0.21	0.40
(103 firms)	Post-ADR	0.26	−0.09	−0.02	+0.07	0.26	0.25
Well-developed markets	Pre-ADR	0.54	−0.12	−0.02	+0.10	0.18	0.59
(127 firms)	Post-ADR	0.35	−0.12	−0.04	+0.07	0.21	0.21
Less-developed markets	Pre-ADR	0.23	−0.13	−0.09	+0.04	0.17	0.32
(43 firms)	Post-ADR	0.59	−0.13	−0.02	+0.11	0.19	0.38

Panel A reports means and standard deviations of $ilq \equiv \ln(ILLIQ)$ for the pooled firm-years. The last 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 $z_{i,t} = (ret_{i,t}, roe_{i,t}, (p\text{-}b)_{i,t}, ilq_{i,t})'$. All variances and covariances are deflated by $Var(\tilde{r})$. The last two columns report coefficients and R^2s for the regression, $N_{er} = \beta_{er|ilq} N_{ilq} + \tilde{N}_{er}$, $N_{ilq} \perp \tilde{N}_{er}$. The figures are calculated from the GMM estimates of the firm-level panel VAR.

where $\tilde{N}_{er,t}$ represents the component of $N_{er,t}$ orthogonal to $N_{ilq,t}$, 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 in expected returns cannot be accounted for by the variation in illiquidity news.

Table 8 repeats the covariance and correlation tests in Table 3 using the illiquidity-augmented VAR. 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 the additional data requirements for daily stock returns and trading volume. Nevertheless, our conclusions from Table 3 remain intact in Table 8.

We measure the liquidity effects by the term $\beta_{er|ilq} N_{ilq,t}$ in Equation (15). We then ask to what extent it can account for the negative covariance between cash-flow news and expected-return news, as well as the shift in the covariance around cross-listing. Specifically, we consider the following decomposition:

$$Cov(N_{er,t}, N_{cf,t}) = \beta_{er|ilq} Cov(N_{ilq,t}, N_{cf,t}) + Cov(\tilde{N}_{er,t}, N_{cf,t}), \qquad (16)$$

Table 8
Covariance and correlation tests (VAR with ilq)

	$Cov^*(N_{cf,t}, N_{er,t})$			$Corr(N_{cf,t}, N_{er,t})$		
Group	Pre-	Post-	Change	Pre-	Post-	Change
All	−0.20	−0.12	+0.09	−0.70	−0.32	+0.38
(170 firms)	−10.2***	−4.02***	2.31**	−5.38***	−3.13***	2.19**
NYSE/AMEX/NASDAQ	−0.19	−0.07	+0.12	−0.74	−0.15	+0.59
(67 firms)	−5.66***	−0.87	1.60	−3.60***	−0.74	2.09**
Pink Sheets/PORTAL	−0.17	−0.13	+0.04	−0.52	−0.37	+0.15
(103 firms)	−3.44***	−3.68***	0.65	−2.58***	−2.96***	0.83
Well-developed markets	−0.15	−0.12	+0.03	−0.42	−0.33	+0.10
(127 firms)	−3.51***	−3.45***	0.50	−2.45***	−2.67***	0.43
Less-developed markets	−0.23	−0.10	+0.13	−0.95	−0.25	+0.70
(43 firms)	−6.26***	−1.32	1.59	−7.53***	−1.05	2.75***

This table reports $Cov^*(N_{cf,t}, N_{er,t}) \equiv Cov(N_{cf,t}, N_{er,t})/Var(\tilde{r}_t)$, where $\tilde{r}_t \equiv ret_t - E_{t-1}ret_t$, and $Corr(N_{cf,t}, N_{er,t})$ 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 $z_{i,t} = (ret_{i,t}, roe_{i,t}, (p\text{-}b)_{i,t}, ilq_{i,t})'$. ***,**, and * indicate significance at 1%, 5%, and 10%, respectively.

Table 9
Significance of liquidity effects on news covariance

Group	Periods	Liquidity effects	Residual effects
All	Pre-ADR	−0.09	−0.11
(170 firms)	Post-ADR	−0.07	−0.05
	Post − Pre	+0.02	+0.07
NYSE/AMEX/NASDAQ	Pre-ADR	−0.08	−0.11
(67 firms)	Post-ADR	−0.02	−0.05
	Post − Pre	+0.06	+0.06
Pink Sheets/PORTAL	Pre-ADR	−0.04	−0.13
(103 firms)	Post-ADR	−0.04	−0.09
	Post − Pre	+0.00	+0.04
Well-developed markets	Pre-ADR	−0.04	−0.11
(127 firms)	Post-ADR	−0.06	−0.06
	Post − Pre	+0.02	+0.01
Less-developed markets	Pre-ADR	−0.09	−0.14
(43 firms)	Post-ADR	−0.04	−0.06
	Post − Pre	+0.05	+0.08

This table reports the liquidity effects, $\beta_{er|ilq}Cov(N_{ilq,t}, N_{cf,t})/Var(\tilde{r}_t)$, where $\tilde{r}_t \equiv ret_t - E_{t-1}ret_t$, and the residual effects, $Cov(N_{er,t} - \beta_{er|ilq}N_{ilq,t}, N_{cf,t})/Var(\tilde{r}_t)$, estimated for both the 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 $z_{i,t} = (ret_{i,t}, roe_{i,t}, (p\text{-}b)_{i,t}, ilq_{i,t})'$.

where the term $\beta_{er|ilq}Cov(N_{ilq,t}, N_{cf,t})$ represents the liquidity effects on the covariance between cash-flow news and expected-return news.

Table 9 reports $Cov(N_{er,t}, N_{cf,t})$ and $\beta_{er|ilq}Cov(N_{ilq,t}, N_{cf,t})$, both deflated by $Var(\tilde{r})$ as in Table 3. The liquidity 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 the NYSE, AMEX, or NASDAQ and for firms from less-developed/emerging markets. Nevertheless, the majority of $Cov(N_{er,t}, N_{cf,t})$ and its change are left unexplained. This result suggests that enforced disclosure can affect expected stock returns well beyond its liquidity effects.

3. Conclusion

In the presence of information asymmetry between a firm's manager and investors, the manager has discretion to disclose or withhold his private information about the firm's performance. Knowing this, investors try to infer the unfavorable information potentially withheld by the manager. They become more skeptical when the manager reports little favorable information, because they do not know whether he is withholding unfavorable information or if he simply does not observe much information. We build on Shin's (2003) model and show that such skepticism produces a stronger return reversal because investors require an additional price discount on top of the usual one resulting from risk aversion. The skepticism and the absence of negative reported performance also lead to a negative correlation between cash-flow news and expected-return news, as opposed to a positive correlation resulting from truthful disclosure.

We find compelling evidence for our theoretical predictions. Applying a GMM estimator for Vuolteenaho's (2002) panel VAR to foreign firms that cross-list ADRs, we find a significant negative correlation between the two news components of stock returns, and a significant reversal before the ADRs. Consistent with the hypothesis that US cross-listing commits foreign firms' managers to fuller disclosure, both the news correlation and the return reversal become significantly weaker after cross-listing. The changes are stronger for: firms that raise disclosure standards substantially, namely, firms that cross-list on the NYSE, AMEX, or NASDAQ, than firms that cross-list on OTC Pink Sheets or PORTAL; firms from less-developed/emerging capital markets than those from more advanced capital markets; and firms from code-law countries than those from common-law countries. Our findings are robust to possible changes in the liquidity level associated with cross-listing and to potential effects of local capital market developments.

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 asymmetry between firm managers and investors, as well as the litigation risks for withholding unfavorable information, varies across industries and firms. Therefore, extending our model to multiple firms with different levels of information asymmetry should help us understand how strategic disclosure produces a cross-sectional difference in information risk premiums. Second, although we empirically control for the possible effects of liquidity on stock returns, our theoretical model assumes homogeneous investors and does not allow us to analyze the effect of strategic disclosure on trading or liquidity. Incorporating heterogeneity among investors is a natural and interesting theoretical extension for further study.

Finally, this paper presents significant and robust evidence consistent with the predictions of our theory. Nevertheless, our results should be interpreted cautiously as the effects of disclosure strategies on expected returns are difficult to isolate from the potential effects of other factors, such as the composition of investor base, the 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 relation 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, $h(k|\mathcal{F}_0) = \binom{N}{k} r^k (1-r)^{N-k}$ in Equation (2). Then Equation (3) becomes

$$V_0^a = \frac{\sum_{k=0}^{N} \binom{N}{k} r^k (1-r)^{N-k} (u^k d^{N-k})^{1-\alpha}}{\sum_{k=0}^{N} \binom{N}{k} r^k (1-r)^{N-k} (u^k d^{N-k})^{-\alpha}}. \tag{A1}$$

Here, the denominator can be computed as

$$\begin{aligned}
&\sum_{k=0}^{N} \binom{N}{k} (r u^{-\alpha})^k [(1-r) d^{-\alpha}]^{N-k} \\
&= \underbrace{\sum_{k=0}^{N} \binom{N}{k} \left(\frac{r u^{-\alpha}}{r u^{-\alpha} + (1-r) d^{-\alpha}} \right)^k \left(\frac{(1-r) d^{-\alpha}}{r u^{-\alpha} + (1-r) d^{-\alpha}} \right)^{N-k}}_{=1} \\
&\quad \times [r u^{-\alpha} + (1-r) d^{-\alpha}]^N \\
&= [r u^{-\alpha} + (1-r) d^{-\alpha}]^N.
\end{aligned} \tag{A2}$$

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

$$V_0^a = \left[\frac{r u^{1-\alpha} + (1-r) d^{1-\alpha}}{r u^{-\alpha} + (1-r) d^{-\alpha}} \right]^N \equiv [\psi u + (1-\psi) d]^N. \tag{A3}$$

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 π in his Equation (16) and q in his Lemma 1. We note that $h(k|\mathcal{F}_1) = h(k|s)$ is also 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 zero failures, $f = 0$, because any other strategy will result in a lower firm value. 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 the observed number of both successes and failures, i.e., $\mathcal{F}_1 = \{s, f\}$.

Lemma 1. *When the manager makes full disclosure,*

$$h(k|s, f) = \begin{cases} \binom{N-f-s}{k-s} r^{k-s}(1-r)^{N-f-k} & \text{if } s \le k \le N-f, \\ 0 & \text{otherwise}. \end{cases} \tag{A4}$$

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. A mathematical proof can be found in the technical appendix available on the journal's and the authors' Web sites.

Then, using the pricing Equation (3) with state prices in Equation (2),

$$\begin{aligned} V_1^a(s, f) &= \frac{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} r^{k-s}(1-r)^{N-f-k}(u^k d^{N-k})^{1-\alpha}}{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} r^{k-s}(1-r)^{N-f-k}(u^k d^{N-k})^{-\alpha}} \\ &= \frac{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} (ru^{1-\alpha})^{k-s}[(1-r)d^{1-\alpha}]^{N-f-k}}{\sum_{k=s}^{N-f} \binom{N-f-s}{k-s} (ru^{-\alpha})^{k-s}[(1-r)d^{-\alpha}]^{N-f-k}} \cdot \frac{u^{(1-\alpha)s} d^{(1-\alpha)f}}{u^{-\alpha s} d^{-\alpha f}} \\ &= \frac{[ru^{1-\alpha} + (1-r)d^{1-\alpha}]^{N-s-f}}{[ru^{-\alpha} + (1-r)d^{-\alpha}]^{N-s-f}} \cdot \frac{u^{(1-\alpha)s} d^{(1-\alpha)f}}{u^{-\alpha s} d^{-\alpha f}} \\ &= [\psi u + (1-\psi)d]^{N-s-f} \cdot u^s d^f. \quad \blacksquare \end{aligned} \tag{A5}$$

A.2 Proof of Proposition 1

Part (i): Sanitization strategy. Using the result of Theorem 1, compute the gross first-period return under the sanitization strategy,

$$R_1^a(s) = \frac{V_1^a(s)}{V_0^a} = \frac{u^s[\pi u + (1-\pi)d]^{N-s}}{[\psi u + (1-\psi)d]^N}. \tag{A6}$$

Under risk neutrality ($\alpha = 0$), we have $\psi = r$ and $\pi = q$. Substitute them into the above equation to calculate the cash-flow news,

$$N_{cf} = \ln R_1 - E_0[\ln R_1] = s \ln\left(\frac{u}{qu + (1-q)d}\right) + const., \tag{A7}$$

where *const.* generally represents a constant up to a relevant information set. Then, the expected-return news is given by

$$\begin{aligned} N_{er} &= \ln \tilde{R}_1 - \ln \tilde{R}_1^a = \ln R_1 - \ln R_1^a + const. \\ &= s \ln\left(\frac{\pi u + (1-\pi)d}{qu + (1-q)d}\right) + const. \end{aligned} \tag{A8}$$

Since $0 < \pi < q < r < 1$, we have

$$Cov(N_{cf}, N_{er}) = Var(s) \underbrace{\ln\left(\frac{u}{qu + (1-q)d}\right)}_{>0} \times \underbrace{\ln\left(\frac{\pi u + (1-\pi)d}{qu + (1-q)d}\right)}_{<0} < 0. \tag{A9}$$

Part (ii): Full disclosure. Again, using the result of Theorem 1, the gross first-period return under full disclosure is

$$R_1^a(s, f) = \frac{V_1^a(s, f)}{V_0^a} = \frac{u^s d^f}{[\psi u + (1-\psi)d]^{s+f}}. \tag{A10}$$

Following the procedure similar to the sanitization strategy, we can calculate the two news components as

$$N_{cf} = s \ln\left(\frac{u}{ru + (1-r)d}\right) + f \ln\left(\frac{d}{ru + (1-r)d}\right) + const., \tag{A11}$$

$$N_{er} = (s + f) \ln\left(\frac{\psi u + (1-\psi)d}{ru + (1-r)d}\right) + const. \tag{A12}$$

Then, with $n \equiv s + f$, we have

$$\begin{aligned} Cov(N_{cf}, N_{er}) = \Bigg[& Cov(s, n) \times \ln\left(\frac{u}{ru + (1-r)d}\right) + Cov(f, n) \\ & \times \ln\left(\frac{d}{ru + (1-r)d}\right)\Bigg] \times \ln\left(\frac{\psi u + (1-\psi)d}{ru + (1-r)d}\right). \end{aligned} \tag{A13}$$

It is straightforward to show that, given n, s is distributed binomially with the probability mass function $h(s|n) = \binom{n}{s} r^s (1-r)^{n-s}$ with mean $E[s|n] = nr$. Similarly, $E[f|n] = n(1-r)$. In addition, n is distributed binomially with the probability mass function $h(n) = \binom{N}{n} \theta^n (1-\theta)^{N-n}$ with variance $N\theta(1-\theta)$. Using the Law of Iterated Expectations,

$$Cov(s, n) = Cov(E[s|n], n) = Cov(nr, n) = Var(n)r = N\theta(1-\theta)r, \tag{A14}$$

$$\begin{aligned} Cov(f, n) &= Cov(E[f|n], n) = Cov(n(1-r), n) = Var(n)(1-r) \\ &= N\theta(1-\theta)(1-r). \end{aligned} \tag{A15}$$

It then follows that, using Jensen's inequality to sign the first logarithm below,

$$Cov(N_{cf}, N_{er}) = N\theta(1-\theta)\underbrace{\ln\left(\frac{u^r d^{1-r}}{ru+(1-r)d}\right)}_{<0} \times \underbrace{\ln\left(\frac{\psi u+(1-\psi)d}{ru+(1-r)d}\right)}_{<0} > 0. \quad \blacksquare \tag{A16}$$

A.3 Proof of Proposition 2

To save space, we only describe the synopsis of the proof. A complete proof is available in the technical appendix. We first derive the necessary moment expressions. 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, using Equation (A6), the expected first-period return under the sanitization strategy is given by

$$E\left[R_1^a(s)\right] = \sum_{s=0}^{N} h(s)R_1^a(s) = [r\theta\gamma_0 + (1-r\theta)\gamma_1]^N, \tag{A17}$$

where we have defined $\gamma_0 \equiv \frac{u}{\psi u+(1-\psi)d} > 1$ and $\gamma_1 \equiv \frac{\pi u+(1-\pi)d}{\psi u+(1-\psi)d} < 1$. Similarly, we can calculate

$$R_2^a(s) = \frac{V_2^a}{V_1^a(s)} = \frac{u^k d^{N-k}}{u^s[\pi u+(1-\pi)d]^{N-s}}, \tag{A18}$$

$$E\left[E\left(R_2^a(s)\middle|s\right)\right] = [r\theta + (1-r\theta)\gamma_2]^N > 1, \tag{A19}$$

$$E\left[R_1^a(s)\cdot E\left(R_2^a(s)\middle|s\right)\right] = [r\theta\gamma_0 + (1-r\theta)\gamma_1\gamma_2]^N, \tag{A20}$$

where $\gamma_2 \equiv \frac{qu+(1-q)d}{\pi u+(1-\pi)d} > 1$. From these moments, we can compute the return autocovariance under the sanitization strategy using the Law of Iterated Expectations,

$$\begin{aligned} cov\left(R_1^a(s), R_2^a(s)\right) &= E\left[R_1^a(s)\cdot E\left(R_2^a(s)\middle|s\right)\right] - E\left[R_1^a(s)\right]E\left[E\left(R_2^a(s)\middle|s\right)\right] \\ &= [r\theta\gamma_0 + (1-r\theta)\gamma_1\gamma_2]^N \\ &\quad - [\{r\theta\gamma_0 + (1-r\theta)\gamma_1\}\{r\theta + (1-r\theta)\gamma_2\}]^N. \end{aligned} \tag{A21}$$

Comparing the terms inside the square brackets,

$$\begin{aligned} &r\theta\gamma_0 + (1-r\theta)\gamma_1\gamma_2 - \{r\theta\gamma_0 + (1-r\theta)\gamma_1\}\{r\theta + (1-r\theta)\gamma_2\} \\ &\quad = -r\theta(1-r\theta)(\gamma_2-1)(\gamma_0-\gamma_1) < 0, \end{aligned} \tag{A22}$$

which implies that $cov(R_1^a(s), R_2^a(s)) < 0$ for all N.

Next, consider full disclosure. 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. Using this and Lemma 1, we can similarly calculate

$$E\left[R_1^a(s,f)\right] = [\theta\gamma_3 + 1 - \theta]^N > 1,$$
$$R_2^a(s,f) = \frac{V_2^a}{V_1^a(s,f)} = \frac{u^k d^{N-k}}{u^s d^f[\psi u + (1-\psi)d]^{N-s-f}}, \quad \text{(A23)}$$
$$E\left[E\left(R_2^a(s,f)\middle|s,f\right)\right] = [\theta + (1-\theta)\gamma_3]^N > 1, \quad \text{(A24)}$$
$$E\left[R_1^a(s,f)\cdot E\left(R_2^a(s,f)\middle|s,f\right)\right] = \gamma_3^N > 1, \quad \text{(A25)}$$
$$cov\left(R_1^a(s,f), R_2^a(s,f)\right) = \gamma_3^N - [\{\theta\gamma_3 + 1 - \theta\}\{\theta + (1-\theta)\gamma_3\}]^N, \quad \text{(A26)}$$

where $\gamma_3 \equiv \frac{ru+(1-r)d}{\psi u+(1-\psi)d} > 1$. Comparing the terms in the last line ignoring the exponent,

$$\gamma_3 - \{\theta\gamma_3 + 1 - \theta\}\{\theta + (1-\theta)\gamma_3\} = -\theta(1-\theta)(\gamma_3 - 1)^2 < 0, \quad \text{(A27)}$$

which implies that $cov(R_1^a(s,f), R_2^a(s,f)) < 0$ for all N.

Further, by the Law of Iterated Expectations, $E[R_1^a(s,f)\cdot E(R_2^a(s,f)|s,f)] = E[R_1^a(s)\cdot E(R_2^a(s)|s)]$. Thus,

$$cov\left(R_1^a(s), R_2^a(s)\right) - cov\left(R_1^a(s,f), R_2^a(s,f)\right) = -E[R_1^a(s)]\cdot E[E(R_2^a(s)|s)] + E[R_1^a(s,f)]\cdot E[E(R_2^a(s,f)|s,f)]. \quad \text{(A28)}$$

Inspecting the above moment expressions, we see that this expression is negative for all N if it is so for $N = 1$. Therefore, it suffices to show that, for $N = 1$,

$$cov\left(R_1^a(s), R_2^a(s)\right) - cov\left(R_1^a(s,f), R_2^a(s,f)\right) = \theta(1-\theta)(\gamma_3 - 1)^2 - r\theta(1-r\theta)(\gamma_2 - 1)(\gamma_0 - \gamma_1) \quad \text{(A29)}$$

is negative, where we have used Equations (A27) and (A22). Substituting the definitions of γ_0, γ_1 , γ_2, and γ_3 and further those of π, ψ, and q, we can show that the above expression is proportional to

$$\frac{d^{-\alpha} - u^{-\alpha}}{[\psi u + (1-\psi)d][ru^{-\alpha} + (1-r)d^{-\alpha}]^2} - \frac{d^{-\alpha}}{[\pi u + (1-\pi)d][(1-\theta)ru^{-\alpha} + (1-r)d^{-\alpha}]^2} < 0, \quad \text{(A30)}$$

with a positive proportionality constant after tedious but straightforward algebra. The inequality holds because $d^{-\alpha} - u^{-\alpha} < d^{-\alpha}$, $\psi u + (1-\psi)d > \pi u + (1-\pi)d$, and $ru^{-\alpha} + (1-r)d^{-\alpha} > (1-\theta)ru^{-\alpha} + (1-r)d^{-\alpha}$ for $\alpha > 0$. ■

Appendix B: Econometric Methodology

B.1 Dynamic panel data estimation

Recall that our firm-specific state vector for firm i, denoted by $z_{i,t}$, consists of three state variables: $z_{i,t} = (ret_{i,t}, roe_{i,t}, (p\text{-}b)_{i,t})'$, 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 us take the first state variable $ret_{i,t}$ as an example. The single-equation panel autoregressive model with individual effects can be written as

$$ret_{i,t} = \gamma_{11} ret_{i,t-1} + \gamma_{12} roe_{i,t-1} + \gamma_{13}(p\text{-}b)_{i,t-1} + \eta_i + \tilde{v}_{i,t}, \tag{A31}$$

where η_i is the individual firm-specific effect. γ_{11} denotes the (1, 1) element of the VAR coefficient matrix Γ.[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 $\tilde{v}_{i,t}$ as $\overline{v}_i = \frac{1}{T}\sum_{t=1}^{T} \tilde{v}_{i,t}$. A simple within-group transformation would show that the strict exogeneity condition is violated when regressors include lagged dependent variables,

$$\sum_{t=1}^{T} E[ret_{i,t-1}(\tilde{v}_{i,t} - \overline{v}_i)] \neq 0. \tag{A32}$$

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; and 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 transformed variables and levels. We use the past five available lagged endogenous variables as instruments in "transformed regressions" and the most recent lagged transformations of endogenous variables in "level regressions."

Specifically, our moment conditions for the "transformed regressions" are

$$E\begin{bmatrix} ret_{i,t-\tau} \\ roe_{i,t-\tau} & \otimes \tilde{v}^*_{i,t} \\ (p\text{-}b)_{i,t-\tau} \end{bmatrix} = 0 \text{ for } \tau = 2, \ldots, 6; t = 3, \ldots T, \tag{A33}$$

[26] A univariate autoregression for $ret_{i,t}$ is a special case in which γ_{12} and γ_{13} 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 inferences based on the two-step estimators and asymptotic standard errors are inappropriate in typical sample sizes.

where $\otimes$ denotes the Kronecker product and $\tilde{\upsilon}^*_{i,t}$ is the residual from the regressions on variables after taking orthogonal deviations,

$$z^*_{i,t} \equiv \left(z_{i,t} - \frac{z_{i,t+1} + \cdots + z_{i,T}}{T-t}\right)\left(\frac{T-t}{T-t+1}\right)^{1/2} \text{ for } t = 1, \ldots, T-1. \tag{A34}$$

Our moment conditions for the "level regressions" are

$$E\left[\tilde{\nu}_{i,t} \otimes \begin{array}{c} ret^*_{i,t-1} \\ roe^*_{i,t-1} \\ (p\text{-}b)^*_{i,t-1} \end{array}\right] = 0 \text{ for } t = 3, \ldots T, \tag{A35}$$

where $\tilde{\nu}_{i,t} = \eta_i + \tilde{\upsilon}_{i,t}$.

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 uses 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 cross-listing, we compute the differences in these parameters (denoted by K) using all the firms (for example, $\gamma^{post}_{11} - \gamma^{pre}_{11}$). Then we recalculate the differences in these parameters, deleting one firm at a time. Denote by $K_{(-j)}$ this estimator computed with the jth firm removed from the total L firms. Then, we can calculate the Jackknife standard errors as

$$\sqrt{\widehat{Var(K)}_J} = \sqrt{\frac{\sum_{j=1}^{L}(K^*_j - J(K))^2}{L(L-1)}}, \tag{A36}$$

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 $K/\sqrt{\widehat{Var(K)}_J}$. Monte Carlo studies indicate that $\widehat{Var(K)}_J$ often overestimates the variance (Efron and Stein, 1981; and Efron 1982). Consequently, our t-tests provide very conservative evidence against the null hypothesis.

Supplementary Data

Supplementary data are available online at http://rfs.oxfordjournals.org/.

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