THREE ESSAYS IN CROSS-BORDER FINANCE

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This Ph.D. dissertation investigates various areas in financial economics: market microstructure, corporate finance, asset pricing, and financial econometrics. The three comprising essays have a common ground: cross-border finance.

Chapter One documents the impact of differential private information on relative asset pricing across borders by studying the probability of informed trading (PIN) for Canadian shares traded on exchanges separated by Niagara Falls. Relative to the New York Stock Exchange (NYSE), the Toronto Stock Exchange (TSX) has more informed trades and accounts for a larger information share, indicating that informed traders contribute to cross-border price discovery. The information imbalance across the two markets is associated with small but positive price premiums for New York trades. The dynamics of these premiums depends on trade informedness. Lastly, the PIN of a TSX-listed share typically rises upon cross-listing on the NYSE, which is consistent with negative abnormal returns of the original listing.

The theory of corporate governance suggests that managers of poorly governed firms are more likely to make poor investment decisions, and the evidence on high antitakeover provision (ATP) firms is consistent. In Chapter Two, I study the effect of domestic and foreign takeovers by U.S. firms and find that high-ATP bidders tend to pay relatively high premiums for either targets. While this suggests that these firms make poor decisions, high-ATP bidders also experience relatively high event study returns at times of foreign takeover news.

This contradicts the findings of Masulis *et al.* (2007) for domestic takeovers.

Finally, Chapter Three explores the convergence between the prices of American Depositary Receipts (ADRs) listed by Asia-Pacific firms and their original shares listed on home exchanges. Instead of relying on conventional parametric approaches that carry embedded model-specification errors, I contribute to the literature by introducing a nonparametric technique to estimate the convergence speed parameter. I present the time-varying characteristics of both firm and country-level convergence speed parameters. Furthermore, I empirically verify and visually corroborate the comparative dynamics of convergence with respect to short sales restrictions, trading time differences, and market-tier measures proxied by the Morgan Stanley Capital International indices. I conclude that enhancement in market efficiency accelerates the reversion to the parity of ADR-pairs.

BIOGRAPHICAL SKETCH

I received a Bachelor of Arts (B.A.) in economics from Yonsei University in 1997 and a Master of Arts (A.M. *Cantab.*) in statistics from Harvard University in 2005. I have served on the faculty of the Ewha School of Business as a tenure-track professor of finance since March 2010.

This dissertation is dedicated to my parents, Chun Sik Choi and Sam Jo Kim, loving wife, Hye Yeon Nam, and dearest son, Daniel Minsoo Nam Choi.

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CHAPTER 1

A TRANS-NIAGARA TALE OF INFORMED TRADERS

1.1 Introduction

Canada and the United States are among the most integrated economies in the world and share comparable accounting standards and institutions. Can information asymmetry explain cross-border pricing effects for Canadian shares listed in both Canadian and U.S. equity markets? This research begins by showing how dominance in private information in one market can yield a positive premium[1] in the other market. Empirical tests relate information asymmetry to the level and dynamics of these premiums, and to cross-listing announcement effects. The probability of informed trading (PIN) proves itself to be an effective tool for revealing "how information is priced" in stock trading dispersed across the border, across time, and beyond the initial cross-listing event.

Over the past several decades, many firms have listed their common shares on exchanges outside their home country. According to the World Federation of Exchanges, as of 2005, the global market capitalization of stocks listed outside their home country by 2,636 foreign companies amounted to U.S.$5.76 trillion, an increase of 16.3% from 2004. In the U.S. alone, almost 2,000 cross-listings[2]

[1]It is the relative premium of a cross-listed stock traded on a foreign exchange against the home market share, adjusted by the exchange rate. January 1998 through December 2000, the ten-minute frequency relative premium for 56 pairs of Canadian stocks traded on the New York Stock Exchange (NYSE) and the Toronto Stock Exchange (TSX) has arithmetic mean, median, and standard deviation of 0.00306, 0.00004, and 0.03031 respectively (Panel A of Table A.2). The term "cross-listing premium" defined by Doidge, Karolyi, and Stulz (2004) is the excess value of foreign firms cross-listed in the U.S. relative to those not in terms of Tobin's (1969) q ratio.

[2]This includes Levels I *&* II Depositary Receipts (DRs), Level I over-the-Counter (OTC) DRs, Rule 144a private placement DRs, ordinary shares, and Global Registered Shares (GRSs). See Bank of New York's (2006) *The Depositary Receipt Markets*.

were recorded. By September 2005, the total value of American Depositary Receipts (ADRs) reached U.S.$657 billion, an increase of 36% over the preceding twelve months. The popularity of international cross-listings has prompted many publications on this subject, most of which focus on the benefits of cross-border listings. See Karolyi (2006) for an excellent survey.

Cross-listing is a cross-border version of "fragmentation."[3] Consequently, the same questions asked of domestically fragmented trading also arise with international cross-listing.[4] If a stock lists on both home and foreign exchanges, where does price information originate and where does price discovery take place? What is the dynamic relationship between the two? Do both markets reflect the same fundamental values? Does the trading of identical stocks in two distinct markets reveal the same information on the company?

Hasbrouck (1995) confirms that the New York Stock Exchange (NYSE) dominates other regional exchanges in contributing to price discovery: order purchase agreements may seek to divert small retail trades to regional locations but leave the larger and potentially more information-based trades to the NYSE. When a non-U.S. stock lists on the NYSE, the host exchange may no longer be the overwhelming source of new information being collected about the cross-listed pair. On the other hand, trades on the non-U.S. home exchange can be more influential if more information (either private or public) is traded in the home market.

In this paper, I study the trading of Canadian shares listed on the NYSE, along with their original listings on the Toronto Stock Exchange (TSX). The Canadian

[3] In the market microstructure literature, fragmentation refers to domestic multi-marketing trading of a stock listing.

[4] Previous studies on "intra-border" fragmentation include Hasbrouck (1995) and Easley, Kiefer, and O'Hara (1996).

shares traded in the U.S. are identical to those traded at home in terms of dividends, voting rights, and other characteristics, and can be bought and sold on either market. Furthermore, the U.S. and Canadian economies are highly integrated, implying identical costs of capital and identical stock prices in both markets. While a positive but small relative premium in New York trading (Panel A of Table A.2) is not likely to yield consistent arbitrage profits after considering bid-ask spreads and other trading costs, it may, as we shall see, reveal the impact of private information in interesting and useful ways.

Eun and Sabherwal (2003) and Gagnon and Karolyi (2004, 2009a) document price differentials between Canadian cross-listings in the U.S. and their original listings in Canada. By extension, my theoretical arguments and empirical results show that information asymmetry that varies across the border, firms, and time manifests itself in relative pricing of Canadian cross-listed pairs on the NYSE and the TSX. The PIN on a stock proxies for the proportion of informed transactions among all trades in a particular market. Following Easley, Kiefer, O'Hara, and Paperman (1996), I individually estimate PIN for both the TSX and NYSE trading of each cross-listed pair.

Easley, Hvidkjaer, and O'Hara (2002) note that, in equilibrium, a high-PIN stock carries an adverse-selection discount. Similarly, I reason that a non-zero price gap arises between New York and Toronto trades if one market features relatively more private information. Building on the noisy rational expectations model of Grossman and Stiglitz (1980), I show that a higher-PIN TSX-listed stock must trade at a lower price than on the NYSE in a no-arbitrage equilibrium given a sufficient condition of "home market liquidity dominance." Put another way, a price discount is needed to induce buyers to trade in the market which is more likely to be plagued by informed traders.

Hasbrouck's (1995) "information share" is a relative measure of the contribution made by a particular stock exchange to price discovery when trade in an asset is dispersed across multiple domestic sites. This idea is also valid beyond the border. The exchange with a higher proportion of informed traders (PIN) is expected to lead the other market in cross-border price discovery, reflected in a higher information share. Given a "Trans-Niagara" imbalance in asymmetric information, a slightly higher NYSE price is sensible. The volatility of the relative premiums in New York (Panel A of Table A.2) can attract arbitrageurs. In turn, the degree to which arbitrage pushes NYSE and TSX prices to converge to parity can be measured by the convergence speed parameter of Gagnon and Karolyi (2004). I relate the estimated convergence speed to trade informedness (PIN) in order to explain the dynamics of cross-border arbitrage returns.

Cross-listing appears to affect the home exchange in a number of dimensions. Foerster and Karolyi (1998) report that, on average, the bid-ask spread narrows on the TSX upon a cross-listing in New York. The original listings also experience negative abnormal returns upon cross-listing (Foerster and Karolyi (1999)). Given that fewer noise trades occur in the market with lower trading costs (Eun and Sabherwal (2003)), a higher proportion of informed traders on the TSX is likely after a cross-listing on the NYSE. This prediction deserves due attention because the PIN is a risk factor in determining asset returns.

Following Eun and Sabherwal (2003), I choose to study Canadian stocks listed in the U.S. for several reasons. First, Canadian equities are the largest group of stocks cross-listed in the U.S. from a single country. Thus, a large cross-section that holds the nationality of the shares constant is available for study. Second, many of these Canadian stocks trade actively on both the NYSE and the TSX which is essential for conducting intraday tests. Third, the trading hours of the TSX coincide with that of the NYSE (9:30AM—4:00PM, EST), a distinct advantage for studying Canadian stocks relative to those from Europe and Asia with little or no overlap in trading times between home and U.S. markets. Since the potential noise and bias from trading-time differences are eliminated, analyses based on information asymmetry are more reliable. Finally, Canadian stocks trade in the U.S. as ordinary shares due to compatible accounting standards, whereas most other cross-listed shares are ADRs issued by U.S. custodian banks. This implies that arbitrage between the U.S. and Canada is particularly simple as it is not necessary to create or destroy depositary receipts (DRs).

The main empirical findings of my study are as follows. First, relative to the NYSE, the TSX has denser population of informed traders (higher in PIN) and typically accounts for more of the measured information share. This is explicit evidence of the informed traders' contribution to cross-border price discovery, confirming Eun and Sabherwal's (2003) finding. The reason for using the PIN to proxy for information asymmetry is that its cross-border difference is central in relative pricing of cross-listings. Per the extended version of Grossman and Stiglitz's (1980) noisy rational expectations model presented in Section 2, having relatively better liquidity (lower in relative quoted spread) on the TSX provides an empirical support for the sufficient condition to give rise to an, overall, slightly positive premium on the NYSE-listing against its original TSX-listing. A higher PIN on one exchange reflects a larger proportion of informed traders who have a better understanding of the firm. However, this is likely to be the result of institutional background of the TSX where insider trading was more feasible due to delayed prosecution by the authority (King and Segal (2004)).

Nonetheless, the exchange with relatively more informed traders is more likely to generate relevant information that stokes price discovery in both markets.

Second, the tendency of pairs of prices to converge appears to be fostered by discretionary liquidity traders. This novel finding is among a few articles in the literature relating the dynamics of premiums and discounts on pairs of cross-listed shares to information asymmetry.[5] It turns out that lower-PIN pairs converge more rapidly to parity, perhaps because arbitrageurs avoid informed traders, trading with "non-discretionary" liquidity traders instead. Thus, a low PIN on a pair with a quickly vanishing premium reflects active participation of discretionary liquidity traders. Pairs trades can be done without private information on the issuers of diverged stocks as timely execution and unwinding of positions suffice.[6]

Finally, the PIN on a TSX-listed stock, on average, rises upon cross-listing on the NYSE. In other words, the information asymmetry surrounding the issuer on its home exchange intensifies once it cross-lists away from home.[7] This increase in adverse selection is consistent with a finding of negative abnormal returns on the TSX upon cross-listing (Foerster and Karolyi (1999)). The managers of Canadian firms may have been led to trade on inside information upon cross-listings that resulted in undermining their existing shareholder values. This is a mild contradiction to Coffee's (1999) bonding hypothesis which posits that managerial incentives of cross-listing firms are aligned with shareholders' interests.

[5]Gagnon and Karolyi (2009a) find that, controlling for various proxies of information asymmetry, holding costs of long-short portfolios of cross-listed pairs significantly explain the cross-sectional and time-series variation in price parity deviations.

[6]Statistical arbitrage, or pairs trade, is a risk-taking trading strategy on a pair of assets whose price difference is expected to diminish over a relatively short holding period. It contrasts with a true risk-free, pure arbitrage in which one simultaneously submits and settles buy and sell orders on both exchanges.

[7]This finding provides an answer to the open question mentioned in the conclusion of Halling *et al.* (2008).

These three key results effectively address "how information asymmetry is priced" in stock trading that is fragmented across a border, over time, and around cross-listing events. The remainder of this paper is organized as follows. First, Section 2 shows the existence of a positive relative premium with an extended version of Grossman and Stiglitz's (1980) model. Section 3 presents key hypotheses based on the existing literature. Section 4 describes the data and exhibits preliminary results. Section 5 provides my main empirical results. I conclude in Section 6.

1.2 Extended Grossman and Stiglitz (1980) model

Easley, Hvidkjaer, and O'Hara (2002) note that, in equilibrium, a high-probability of informed trading (PIN) stock carries an adverse-selection discount since it requires an additional return.[8] Similarly, I reason that, for a cross-listed pair, the cross-listing yields a positive relative premium[9] if the original home-listing carries heavier private information. For a Canadian company that trades it at $p_{\mathrm{T}} > 0$ on the Toronto Stock Exchange (TSX), its cross-listing decision on the New York Stock Exchange (NYSE) creates a replica that trades at $p_{\mathrm{N}} > 0$ with the same underlying fair value, adjusted for the exchange rate.

Formally, a cross-listing event gives rise to a relative premium in the cross-listing, $\kappa \equiv p_{\mathrm{N}}/p_{\mathrm{T}} - 1 \gtrless 0$, then $p_{\mathrm{N}} = p_{\mathrm{T}} + \alpha\,(\pi_{\mathrm{T}} - \pi_{\mathrm{N}})$ for some $\alpha > 0$, where π_{T} and π_{N} are the respective proportions of informed traders on the TSX and the NYSE whose empirical proxies are the exchange-specific PINs. Thus, the relative

[8] See Appendix A.3 for derivation of the PIN.

[9] I use the term "relative premium" as the relative premium of a cross-listed stock on a foreign exchange against its home market share, adjusted by the exchange rate.

premium is determined as follows: $\kappa = \left(\frac{\alpha}{p_{\mathrm{T}}}\right)(\pi_{\mathrm{T}} - \pi_{\mathrm{N}}) \gtrless 0$ for $\pi_{\mathrm{T}} \gtrless \pi_{\mathrm{N}}$.

Following the noisy rational expectations model introduced by Grossman and Stiglitz (1980), informed traders and uninformed traders have respective proportions on their own exchanges of π_i and $(1 - \pi_i)$, where $i = \mathrm{T(SX), N(YSE)}$. Arbitrageurs do not have an *a priori* proportion on either exchange in a "no-arbitrage" equilibrium. Informed traders and uninformed traders share the same constant relative risk aversion (CARA) utility function with a risk aversion coefficient (ρ) or a risk tolerance parameter ($\eta \equiv 1/\rho$). Arbitrageurs are risk-neutral.

The future earnings (υ) of the firm is uncertain, $\upsilon \sim N\left(\overline{\upsilon}, \sigma_{\upsilon}^2\right)$. Informed traders recognize a signal S about υ with random noise $\epsilon_s \sim N\left(0, \sigma_s^2\right)$, such that $S = \upsilon + \epsilon_s$. The exchange-specific aggregate supply of shares is $Y_i \sim N\left(\overline{y}_i, \sigma_i^2\right)$ and is proportionately driven by uninformed (noise) traders. For convenience, all variances are expressed in precision terms in the following discussion: $\tau_{\upsilon} \equiv 1/\sigma_{\upsilon}^2, \tau_i \equiv 1/\sigma_i^2$, and $\tau_s \equiv 1/\sigma_s^2$.

Neither informed nor uninformed traders cross the Niagara Falls, and they trade on their own exchanges. Informed traders on both exchanges receive the same earnings signal, and they trade based on their updated expectations of future earnings of the firm. Uninformed traders extract information from historical price data only from their respective exchange. Their bias is reasonable since uninformed investors cannot tell informativeness of prices so they only refer to familiar listings. The two markets share the same risk-free asset with a guaranteed net return of r which serves as the common opportunity cost of capital.

Arbitrageurs can buy and sell in both markets, and their demand only depends on the relative premium, or discount. Specifically, their demand for one side of the cross-listed pair (in order to shortsell) is given by x_i^A on each exchange, and it satisfies $x_{\mathrm{T}}^A + x_{\mathrm{N}}^A = 0$ since "pure" arbitrageurs use a perfect hedged strategy. Thus, their short position on the TSX equals their long position on the NYSE, $\mu \equiv x_{\mathrm{T}}^A = -x_{\mathrm{N}}^A$.

Denote the surprises in the earnings signal and the exchange-specific supply of shares as $\Delta S \equiv S - \overline{S}$, and $\Delta Y_i \equiv Y_i - \overline{y}_i$, respectively. The prices of the cross-listed pair are bullish on a positive earnings shock ($\Delta S > 0$), and bearish on positive liquidity excesses ($\Delta Y_i > 0$) and shortsells ($x_i^A > 0$) on respective exchanges. Thus, the prices on the TSX and the NYSE are conjectured to be:

$$\begin{aligned} p_{\mathrm{T}} &= \beta_{\mathrm{T}}^0 + \beta_{\mathrm{T}}^S \Delta S - \beta_{\mathrm{T}}^Y \Delta Y_{\mathrm{T}} - \beta_{\mathrm{T}}^A x_{\mathrm{T}}^A, \\ p_{\mathrm{N}} &= \beta_{\mathrm{N}}^0 + \beta_{\mathrm{N}}^S \Delta S - \beta_{\mathrm{N}}^Y \Delta Y_{\mathrm{N}} - \beta_{\mathrm{N}}^A x_{\mathrm{N}}^A. \end{aligned}$$

Informed traders in the two markets observe the same private signal S and use it to update their beliefs. Upon receiving a new earnings signal, their updated (posterior) earnings forecast ($\mathbb{E}(\upsilon|S)$) and updated earnings forecast precision ($\tau(\upsilon|S)$) are given by

$$\begin{aligned} \mathbb{E}(\upsilon|S) &= \overline{\upsilon} + \left(\frac{\tau_s}{\tau_s + \tau_\upsilon}\right)\Delta S, \\ \tau(\upsilon|S) &\equiv \frac{1}{\mathrm{Var}(\upsilon|S)} = \tau_s + \tau_\upsilon. \end{aligned}$$

Under the CARA utility function assumption, exchange-specific informed traders' demand for shares is

$$\begin{aligned} x_i^I(p_i, S) &= \frac{\mathbb{E}(\upsilon|S) - p_i(1+r)}{\rho \mathrm{Var}(\upsilon|S)} \\ &= \eta(\tau_s + \tau_\upsilon)\left\{\overline{\upsilon} + \left(\frac{\tau_s}{\tau_s + \tau_\upsilon}\right)\Delta S - p_i(1+r)\right\}. \end{aligned}$$

Uninformed traders observe prices on their respective exchanges and form their expectations of future earnings. Their price-contingent updated (posterior) earnings forecast ($\mathbb{E}(\upsilon|p_i)$), updated earnings precision ($\tau(\upsilon|p_i)$) and demand function are, respectively, given by

$$\mathbb{E}(\upsilon|p_i) = \bar{\upsilon} + \left(\frac{1}{\beta_i^S}\right)\left(\frac{\phi_i\tau_s}{\phi_i\tau_s + \tau_\upsilon}\right)\Delta p_i,$$

$$\tau(\upsilon|p_i) \equiv \frac{1}{\mathrm{Var}(\upsilon|p_i)} = \left(\frac{\tau_i}{\tau_i + h_i^2\tau_s}\right)\tau_s + \tau_\upsilon,$$

$$\begin{aligned} x_i^U(p_i) &= \frac{\mathbb{E}(\upsilon|p_i) - p_i(1+r)}{\rho\mathrm{Var}(\upsilon|p_i)} \\ &= \eta(\phi_i\tau_s + \tau_\upsilon)\left\{\bar{\upsilon} + \left(\frac{1}{\beta_i^S}\right)\left(\frac{\tau_s}{\tau_s + \tau_\upsilon}\right)\Delta p_i - p_i(1+r)\right\}, \end{aligned}$$

where $h_i \equiv \beta_i^Y/\beta_i^S$ and $\phi_i \equiv \tau_i/\left(\tau_i + h_i^2\tau_s\right)$.

The market clearing condition on each exchange prescribes

$$\pi_i x_i^I(p_i, S) + (1-\pi_i)\, x_i^U(p_i, S) = Y_i - x_i^A.$$

Consequently, for a given arbitrageurs' position (μ), solving the market-clearing condition for the coefficients ($\beta_i^0, \beta_i^S, \beta_i^Y$, and β_i^A) of conjectured prices yields

$$\beta_i^0 = \frac{\bar{\upsilon}}{1+r} - \frac{\bar{y}_i}{(1+r)\left(\omega_i^I + \omega_i^U\right)},$$

$$\beta_i^S = \frac{1}{(1+r)\left(\omega_i^I + \omega_i^U\right)}\left\{\omega_i^I\left(\frac{\tau_s}{\tau_s + \tau_\upsilon}\right) + \omega_i^U\left(\frac{\phi_i\tau_s}{\phi_i\tau_s + \tau_\upsilon}\right)\right\},$$

$$\beta_i^Y = \frac{\omega_i^I\{\tau_s/(\tau_\epsilon + \tau_\upsilon)\}}{(1+r)\left(\omega_i^I + \omega_i^U\right)}\left\{\omega_i^I\left(\frac{\tau_s}{\tau_s + \tau_\upsilon}\right) + \omega_i^U\left(\frac{\phi_i\tau_s}{\phi_i\tau_s + \tau_\upsilon}\right)\right\},$$

$$\beta_i^A = \frac{1}{(1+r)\left(\omega_i^I + \omega_i^U\right)},$$

where $\phi_i \equiv \frac{\pi_i^2 \eta^2 \tau_s \tau_i}{1+\pi_i^2 \eta^2 \tau_s \tau_i}$, $\omega_i^I \equiv \pi_i \eta (\tau_s + \tau_v)$, and $\omega_i^U \equiv (1-\pi_i)\eta(\phi_i \tau_s + \tau_v)$.[10]

In a no-arbitrage equilibrium ($\mu = \Delta S = \Delta Y_{\mathrm{T}} = \Delta Y_{\mathrm{N}} = 0$), the dollar premium on the cross-listing is as follows.

$$
\begin{aligned}
p_{\mathrm{N}} - p_{\mathrm{T}} &= \beta_{\mathrm{N}}^0 - \beta_{\mathrm{T}}^0 \\
&= \frac{\overline{y}_{\mathrm{T}}}{(1+r)(\omega_{\mathrm{T}}^I + \omega_{\mathrm{T}}^U)} - \frac{\overline{y}_{\mathrm{N}}}{(1+r)(\omega_{\mathrm{N}}^I + \omega_{\mathrm{N}}^U)} \\
&= \frac{\overline{y}_{\mathrm{T}}}{(1+r)\,\omega_{\mathrm{T}}} - \frac{\overline{y}_{\mathrm{N}}}{(1+r)\,\omega_{\mathrm{N}}} = \frac{\omega_{\mathrm{N}}\,\overline{y}_{\mathrm{T}} - \omega_{\mathrm{T}}\,\overline{y}_{\mathrm{N}}}{(1+r)\,\omega_{\mathrm{T}}\,\omega_{\mathrm{N}}} \\
&= \frac{\omega_{\mathrm{N}}\,\{\overline{y}_{\mathrm{T}} - (\omega_{\mathrm{T}}/\omega_{\mathrm{N}})\overline{y}_{\mathrm{N}}\}}{(1+r)\,\omega_{\mathrm{T}}\,\omega_{\mathrm{N}}}.
\end{aligned}
$$

If $\pi_{\mathrm{T}} > \pi_{\mathrm{N}}$, then $\omega_{\mathrm{T}} \equiv \left(\omega_{\mathrm{T}}^I + \omega_{\mathrm{T}}^U\right) > \omega_{\mathrm{N}} \equiv \left(\omega_{\mathrm{N}}^I + \omega_{\mathrm{N}}^U\right)$, thus $\omega_{\mathrm{T}}/\omega_{\mathrm{N}} > 1$. With a sufficient "home market liquidity dominance" condition that $\overline{y}_{\mathrm{T}}/\overline{y}_{\mathrm{N}} > \omega_{\mathrm{T}}/\omega_{\mathrm{N}} > 1$,[11] the stock is dearer on the NYSE than on the TSX such that $p_{\mathrm{N}} > p_{\mathrm{T}}$. In other words, as long as liquidity on the home exchange is relatively "better" than on the host exchange, a higher proportion of informed traders on the home-listed stock must give rise to a strictly positive relative premium in the cross-listed stock. A price discount on the original listing is needed to induce buyers to trade in the market which is more likely to be plagued by informed traders. This premium on the cross-listing does not attract arbitrageurs and, thus, neither side of the pair is mispriced.

If $\overline{y}_{\mathrm{T}} = \overline{y}_{\mathrm{N}}$, $\pi_{\mathrm{T}} > \pi_{\mathrm{N}}$ implies $p_{\mathrm{T}} > p_{\mathrm{N}}$, which is consistent with Chan, Menkveld, and Yang's (2008) application to the Chinese A and B share markets. The no-arbitrage condition in an equilibrium ($\Delta S = \Delta Y_{\mathrm{T}} = \Delta Y_{\mathrm{N}} = 0$) is $p_{\mathrm{N}} - p_{\mathrm{T}} = \beta_{\mathrm{N}}^0 - \beta_{\mathrm{T}}^0$

[10] It can be shown that *1.* $\partial\beta_i^0(\pi_i)/\partial\pi_i > 0$ for all $\pi_i \in [0,1]$; *2.* $\partial\beta_i^S(\pi_i)/\partial\pi_i > 0$ for all $\pi_i \in [0,1]$; *3.* $\partial\beta_i^Y(\pi_i)/\partial\pi_i < 0$ for some large π_i; and *4.* $\partial\beta_i^A(\pi_i)/\partial\pi_i < 0$ for all $\pi_i \in [0,1]$. $\partial\beta_i^S(\pi_i)/\partial\pi_i > 0$ is an intuitive result since the price is expected to reflect more information shocks with an increase in the proportion of informed traders. See proofs in Appendix A.5.

[11] This sufficient condition is reasonable since the higher adverse-selection risk side of a cross-listed pair is offering better liquidity, or facilitating easier exit, in addition to a commensurate discount to attract investors.

(see proof in Appendix A.5).

1.3 Hypotheses

Theoretically speaking, cross-border differential in private information can explain relative pricing of Canadian shares concurrently traded on the TSX and the NYSE. I subsequently raise testable hypotheses of empirical support for the institutional background of information asymmetry, the dynamics of relative premiums, and the informational and economic consequences of cross-listings on the home exchange.

1.3.1 Informed trading and cross-border price discovery

Unlike articles that focus on the joint distribution of trades and prices,[12] Easley, Kiefer, and O'Hara (1997a, 1997b) and Easley, Kiefer, O'Hara, and Paperman (1996) make parametric assumptions to estimate a relative measure of adverse selection using buy and sell order indicators instead of price data. In their theoretical setting, there are risk-averse and competitive market makers, informed traders, and uninformed (liquidity) traders.

The four parameters of the maximum likelihood model are: the probability that an information event occurs on a given day (α); the probability that the information event is pessimistic (δ); and the respective (Poisson) order arrival rates of informed and uninformed traders (μ and η). As a result, the probability

[12] Bagehot (1971), Grossman and Stiglitz (1980), Kyle (1985), and Glosten and Milgrom (1985).

of informed trading[13] (PIN) measures the relative degree of private information-based trades among all trades. Easley, Kiefer and O'Hara (1997b) argue that, as informed traders gain weight in the market, adverse selection is aggravated and the trading volume increases.

Fragmentation is the dispersal of trading in a security to multiple exchanges or markets. As an early bridge between fragmentation and informed trading, Chowdhry and Nanda (1991) note that information lags between distinct trading locations yield transitory disparities in the prices of an identical security. Blume and Goldstein (1991) and Lee (1993) report that price discovery (convergence towards an equilibrium price) on U.S. exchanges occurs primarily on the NYSE. Similar results are drawn by Harris, Mclnish, Shoesmith, and Wood (1995) and Gardner and Subrahmanyam (1994).[14]

When a NYSE-listed stock trades not only on the NYSE but also on the regional exchanges, the fragmented security prices may not be identical but they also cannot differ too much in the long run either. Hasbrouck's (1995) "information share"[15] is a relative measure of contribution made by a stock exchange to price discovery of shares fragmented on multiple exchanges. Hasbrouck (1995) finds that price discovery of fragmented stocks appears to be concentrated on the NYSE whose information share is shown to be the highest.

[13] PIN $\equiv \frac{\alpha\mu}{\alpha\mu+2\eta}$. See Appendix A.3.

[14] Extending the works of Hasbrouck (1991, 1995), Gardner and Subrahmanyam (1994) conclude that fewer informed trades are executed on the regional exchanges than on the NYSE.

[15] Information shares are estimated by the vector error correction model (ECM) provided that the dispersed security prices are "cointegrated." Security prices are cointegrated if there exists a linear combination of the non-stationary prices that can be toned stationary. A time series is strongly stationary if its probability distribution is time-invariant, and weakly stationary up to its second moments: mean, variance, and covariance. This property renders Sims's (1980) original vector autoregressive (VAR) model unwieldy. That is why Hasbrouck (1995) takes an ECM (Engle and Granger (1987), and Engle and Yoo (1987)) approach to propose "information shares." See Appendix A.4.

Easley, Kiefer, and O'Hara (1996) show that there is a significant difference in the information content of orders executed in New York and in Cincinnati, and that this difference is consistent with the "cream-skimming" hypothesis, instead of the competition hypothesis. The notion that trades in distinct U.S. locations carry different levels of information is also relevant to cross-border fragmentation.

Extending the fragmentation idea to the international finance literature, based on U.S.-listed Canadian stocks, Eun and Sabherwal (2003) find that prices on the Toronto Stock Exchange (TSX) and U.S. exchanges are mutually convergent, following Harris, McInish, Shoesmith, and Wood (1995). They report that the U.S. share of price discovery ranges from 0.2 percent to 98.2 percent, with an average of 38.1 percent.

Across the global equity markets, Bailey, Mao, and Sirodom (2006) and Chan, Menkveld, and Yang (2008) describe intriguing multi-board trading structures in Thailand and China, respectively, and explain how information asymmetry affects fragmented trading. Also, foreigners are disadvantaged in Korea (Choe, Kho, and Stulz (2005)) while they wield superior information processing capability in Thailand and Singapore (Bailey, Mao, and Sirodom (2007)).

If a stock listed on an exchange has a higher PIN than its cross-listed stock traded on the other cross-border exchange, this reflects a greater proportion of informed traders who have private information of the issuer. Since informed traders are believed to contribute to price discovery, it is also likely that the exchange with heavier intensity of informed trades generates more relevant information which fosters price discovery.[16]

[16]Hasbrouck (2007) notes that a vector ECM analysis assigns quote changes to the influx of trades. Asymmetric information is then reflected in a wide price change. In this sense, the

By definition, an exchange is said to *lead* the other exchange if it accounts for more price discovery (reflected in its higher information share). However, unlike domestically dispersed stocks, trades in TSX-NYSE cross-listed pairs are exposed to aggregate shocks hitting the two exchanges and the foreign exchange market. In other words, cross-border fragmentation is a more intricate mechanism of price discovery than the domestic case. My first hypothesis attempts to verify the role of informed traders in determining cross-border price discovery. Specifically,

> *H1: compared to the other exchange, the lead market (with a higher average information share) has relatively more informed trades (with a higher average* PIN*), and vice versa.*

1.3.2 Dynamics of relative premiums

Easley, Hvidkjaer, and O'Hara (2002) note that, in equilibrium, a high-PIN stock carries an adverse-selection discount since it requires an additional return. Similarly, as discussed in Section 2, I reason that a cross-listed pair yields either a positive or negative relative premium[17] if one side carries relatively more private information.[18] Unless that relative price spread is believed to persist due to severe liquidity constraints, shortsale restrictions, or other frictions, an arbitrageur will buy the discounted stock and short the other side with favorable

information share is expected to be directionally equivalent to the PIN.

[17]This as defined as the relative premium of a cross-listed stock on a U.S. exchange against its home market basis share, adjusted by the exchange rate.

[18]In Section 2, based on an extended version of the noisy rational expectations model (Grossman and Stiglitz (1980)), I provide a sufficient condition ("home market liquidity dominance") under which a higher-PIN TSX-listed stock must be priced lower than its NYSE-listed replica in a "no-arbitrage" equilibrium.

assumptions on the exchange rate.

The international finance literature has accumulated articles on arbitrage opportunities created by cross-listed shares. The early studies (Maldonado and Saunders (1983), Kato, Linn, and Schallheim (1991), Park and Tavakkol (1994), Miller and Morey (1996), and Karolyi and Stulz (1996)) conclude that arbitrage profits for cross-listed shares do not exist and thus they are priced at parity. Wahab, Lashgari, and Cohn (1992) show that there are arbitrage opportunities in cross-listed pairs. Froot and Dabora (1999) study pricing of a couple of dual-listed corporations (Royal Dutch and Shell, and Unilever N.V. and Unilever PLC) and find a sizable and significant price deviation from parity.[19]

Gagnon and Karolyi (2009a) record significant price deviations in 506 cross-listed pairs under their study: they report discounts of up to 40.4% and premiums of up to 127.4%. The speed at which a relative premium converges to parity is measured by a parameter proposed by Gagnon and Karolyi (2004). According to their empirical model each firm's relative premium can be explained by its first-lag term, and its time-distributed risk exposure to the respective returns on the home and host market indices and the foreign exchange rate.[20]

[19]See Kim, Szakmary, and Mathur (2000) for vector autoregressive (VAR) and seemingly unrelated estimation (SURE) methods that analyze adjustments in ADR-implied prices.

[20]For each cross-listed pair (i), $\text{SPEED}_{\text{CONV}}$ ($\equiv \theta_i$) measures the reciprocal speed of the parity-convergence of relative premium, following Gagnon and Karolyi's (2004) empirical model:

$$\text{DR}_i(t) = \alpha_i + \theta_i \, \text{DR}_i(t-1) + \sum_{j=-1}^{1} \beta_j^{US} R_M^{US}(t+j) + \sum_{j=-1}^{1} \beta_j^{C} R_M^{C}(t+j) + \sum_{j=-1}^{1} \beta_j^{FX} R_{FX}(t+j) + \varepsilon_i(t).$$

The daily relative premium $\left(\text{DR}_i(t) \equiv \left(P_i^{US}(t) - P_i^{C}(t)\right)/P_i^{C}(t)\right)$ can be explained by *1.* its own lag ($\text{DR}_i(t-1)$) associated with 2. the *convergence speed parameter* (θ_i): the closer the absolute value to zero, the faster the convergence to parity; and lag-distributed (yesterday ($j=-1$), today ($j=0$), and tomorrow ($j=+1$)) returns on *3.* the S&P 500 Index $\left(R_M^{US}(t+j)\right)$, *4.* the S&P TSX Composite Index $\left(R_M^{C}(t+j)\right)$, and *5.* the Canada-U.S. exchange rate return ($R_{FX}(t+j)$), a positive R_{FX} implies a depreciation in the Canadian dollar. The forward-lag is due to information leakages and market impact.

In a rational expectations equilibrium, informed investors impound information in prices (Grossman and Stiglitz (1980)) and, thus, catalyze price discovery. In the cross-section, a higher PIN implies enhanced price discovery. Hence, for the "synchronous" relative premium of a Canada-U.S. cross-listed pair, its dynamics (convergence speed) is expected to depend on the informedness of trades, after controlling for market friction, liquidity constraint, and firm characteristics. Parity-convergence can, therefore, be accelerated by the degree of private information on the cross-lister. In this regard, my second conjecture states that

> *H2: the higher the* PIN *on a cross-listed pair, the faster the parity-convergence of relative premiums.*[21]

1.3.3 Cross-listing effects on the home exchange

De Long, Shleifer, Summers, and Waldman (1990) argue that, since noise traders do not reflect information on the fundamentals their trades dislocate prices from their intrinsic values, reducing price informativeness while increasing volatility (noise trader risk). Eun and Sabherwal (2003), Fleming, Ostdiek, and Whaley (1996), and Jones and Seguin (1997) suggest that less noise trades occur in the markets with lower trading costs.

Foerster and Karolyi (1998) document that post-cross-listing spreads in Canada decrease. The augmented liquidity gives rise to TSX market makers' competitive reaction by setting bid-ask spreads lower.[22] The bid-ask spread

[21] By specification, a lower absolute value of parameter below one is equivalent to a higher convergence speed.

[22] The decrease in spreads on the TSX is heavily weighed on the stocks whose trading volume

represents a significant portion in transaction costs, thus cross-listings can reduce noise trader risk on the home exchange. This, in turn, may enhance price discovery, since less noisy fluctuation contributes to setting a more precise and stable process towards the fair price of a security.

A subsequent question will be: "whether less *volatility* entails a higher *proportion* of informed trades?" Further, "does cross-listing exacerbate the home market information environment with relatively more grave adverse selection?" The reason for focusing a cross-listing effect on the PIN, as a proxy information asymmetry, is that it serves as a risk factor in relative pricing of cross-listed pairs. My last hypothesis is that

> *H3: after cross-listing on the NYSE, on average, information asymmetry on a TSX-listed stock intensifies (the* PIN *rises).*

Cross-listings can be a good source of additional liquidity to the existing home-listed stocks. Intensifying adverse selection captured by the PIN and increasing trading volume are positively correlated (Easley, Kiefer, and O'Hara (1997b)) and this further leverages my hypothesis. The additional liquidity on the TSX forces market makers to set spreads narrower. See Admati and Pfleiderer (1988) for a similar discussion.

contribution by the U.S. exchanges is relatively large.

1.4 Data and preliminary results

1.4.1 Data

56 TSX-NYSE pairs are identified through the sample period: January 1, 1998, through December 31, 2000.[23] In order to conduct microstructure analyses, high-frequency data are required for the shares co-listed on the TSX and the NYSE, and the U.S.-Canada exchange rate. Accordingly, the tick-by-tick trade and quote data for the TSX-listed Canadian stocks and the Trade-And-Quote (TAQ) data of their cross-listings on the NYSE through the period are used. The exchange rate intraday data is purchased from Olson & Associates.

Unlike a specialist-based auction exchange NYSE, electronic exchange TSX uses a Central Limit Order Book (CLOB) system, thus orders are required to be in the book to have standing.[24] By studying decrements in the inside depth on one side of the quote that correspond to uncommon trade sizes (like a trade of 1,300 shares), matching trades with prevailing quotes of five-second lead (Lee and Ready (1991)) is reasonable: a trade is considered buyer-initiated if it is higher than the five-second earlier mid-quote, and seller-initiated if lower.[25]

I construct the preliminary datasets for estimation of the PIN following Easley, Kiefer, O'Hara, and Paperman (1996), and Easley, Hvidjkaer, and O'Hara

[23]Following Eun and Sabherwal (2003), the augmented Dickey-Fuller (1981) unit root test is conducted for each pair of daily closing price time series with appropriate lag lengths, per Akaike (1974), to verify first-order integration ($I(1)$). Applying Johansen's (1991) either the trace or eigen-value tests yielded one co-integrating equation for each TSX-NYSE co-listed pair. These results provide justification for constructing error correction models (ECMs) to estimate the information shares of each co-listed pair's exchanges.

[24]I owe this comment to Daniel Weaver. See Eun and Sabherwal (2003) for a detailed institutional comparison between the TSX and the NYSE.

[25]See Schultz and Shive (2008) for trade misclassification of the TAQ on the NYSE which becomes severe after 2000.

(2002). The NYSE-resident specialists are central to the theory of the PIN (Easley, O'Hara, and Saar (2001), and Duarte and Young (2008)). There are official market makers, known as registered traders, on the TSX whose function is akin to that of NYSE specialists. Thus, a comparison of trade informedness on the two exchanges by the PIN is deemed appropriate.[26]

1.4.2 Preliminary results

The PINs for TSX- and NYSE-listed Canadian stocks are estimated following Easley, Kiefer, O'Hara, and Paperman (1996) and Easley, Kiefer and O'Hara (1997a, 1997b).[27] The arithmetic means of monthly PIN estimates of 56 Canadian cross-listers on the TSX and the NYSE are plotted in Figure A.1. It appears that the TSX, on average, dominates the NYSE in terms of the PIN in annual estimates for the cross-listed pairs through the sample period.[28]

The bid-ask spreads[29] are adjusted by the mid-quotes and, thus, measure the relative discrepancy between bid and ask quotes free from the exchange rate. Following Eun and Sabherwal (2003), the mid-points of U.S.-Canada exchange rate bid and ask quotes are updated every minute. The bid and ask quotes of the NYSE-listed Canadian stocks are matched with their concurrent minutes' exchange rate quote mid-points. Based on mutual interaction (orthogonalized

[26]I owe this comment to Lawrence Kryzanowski. See Fuller, Van Ness, and Van Ness (2008) for difficulties in estimation of the PIN for NASDAQ trades.

[27]I adopt Easley, Engle, O'Hara, and Wu's (2008) log-likelihood function specification for improved numerical stability in computing the the PIN. See Appendix A.3.

[28]The annual estimates for the PIN on the TSX are {0.242, 0.213, 0.206} in 1998, 1999, and 2000, respectively, while the corresponding estimates for the NYSE are {0.204, 0.212, 0.196}, over the same period. The spikes in PIN are seen in the post-decimalization period between November and December 1999, a finding consistent with Zhao and Chung (2006).

[29]$\text{SPREAD}_{\text{NYSE}} \equiv \frac{\text{ask}_{\text{NYSE}} - \text{bid}_{\text{NYSE}}}{(\text{ask}_{\text{NYSE}} + \text{bid}_{\text{NYSE}})/2}$; and $\text{SPREAD}_{\text{TSX}} \equiv \frac{\text{ask}_{\text{TSX}} - \text{bid}_{\text{TSX}}}{(\text{ask}_{\text{TSX}} + \text{bid}_{\text{TSX}})/2}$.

impulse responses) of bid and ask quotes on the TSX and the NYSE, the information shares[30] of the TSX and the NYSE for each cross-listed pair are estimated per Hasbrouck (1995, 2007).

The averages across monthly estimates of PINs, spreads, and information shares of each pair over the entire sample period are listed in Table A.1.[31] About twenty firms in the sample exhibit higher PINs on the NYSE than on the TSX. For some cross-listers, like Manulife Financial Corp. and Suncor Energy Inc., there is no significant difference between the PINs on the two exchanges. Only nine firms in the sample show higher spreads on the TSX, and only two firms have higher information shares on the NYSE.

First, on average, the PIN on the TSX (0.242) exceeds that on the NYSE (0.214). Second, the relative quoted spread on the TSX (0.015) is narrower than that on the NYSE (0.022). Third, the information share of the TSX (0.544) is higher than that of the NYSE (0.455). For a Canadian cross-lister, on average, it appears that more price discovery takes place on the TSX (the lead market) where the intensity of informed trades tends to be heavier (a higher PIN) and yet with lower spreads (competitive market making).

The impulse response function plots of bid and ask quotes for Abitibi Consolidated, Inc. are shown in Figure A.2. Each of the four consecutive charts specifies the source of innovation by two standard deviations. The quotes on the NYSE rarely affect the quotes on the TSX. To the contrary, positive increases

[30]Since there are four quote prices (bid_{TSX}, ask_{TSX}, bid_{NYSE}, and ask_{NYSE}), there are 24 (= 4!) orderings in terms of Cholesky exogeneity. For each TSX-NYSE co-listed pair, thus, there are 24 pairs of TSX-NYSE information shares. Averaging across varying exogeneity reduces them to a single pair of information shares for each cross-listed pair. See Appendix A.4.

[31]For brevity, in Table A.1, I do not present the monthly estimates (January 1998 through December 2000) of the PINs , spreads, and information shares for the cross-listed pairs. They are, however, available upon request.

in ask and bid prices on the TSX are followed by changes in ask and bid prices on the NYSE, respectively. This pattern does not hold for all cross-listed stocks, and the degree to which an exchange responds to the other side is reflected in the magnitude of information share.

Based on the ten-minute frequency relative premiums of 56 cross-listed pairs traded through the sample period, the arithmetic mean, the median, and the standard deviation are 0.00306, 0.00004, and 0.03031, respectively (Panel A of Table A.2). The average relative premium of 30.1 basis points with a 3.03 percent volatility is a statistically insignificant deviation from parity. This suggests the extent to which Toronto and New York are integrated.[32] A regression analysis of relative premiums against cross-border differences in the proportions of informed traders is conducted in Panel B of Table A.2. It shows that a higher PIN on a stock listed on the TSX, on average, is associated with a positive premium on the cross-listed stock traded on the NYSE. This strongly supports the extended Grossman and Stiglitz (1980) model presented in Section 2. The seemingly unarbitrageable and negligibly positive average daily relative premium is a result of cross-border imbalance in private information.

[32]See Kryzanowski and Zhang (2002) for further intraday analyses of price differences of Canadian cross-listed pairs traded in Toronto and New York.

1.5 Results

1.5.1 Informed trading and cross-border price discovery

Based on monthly estimates, the statistical significance of the TSX's dominance over the NYSE in terms of the PIN can be verified by the Wilcoxon signed-rank test.[33] In the first column of Table A.3, the Wilcoxon-test statistic, under the null hypothesis is very strongly rejected at a 1% right-tail significance level. Thus, the traders on the TSX posses relatively more private information on Canadian cross-listed stocks than their counterparts on the NYSE. However, this is likely to be the result of institutional background of the TSX where insider trading was more feasible due to delayed prosecution by the authority (King and Segal (2004)).[34]

Harris, McInish, and Wood (2002) report that the influence of the NYSE on price discovery against its regional counterparts increases as its spreads compared to those of the regionals' decrease. In the cross-border context, competitive market making by the TSX versus the NYSE can be inferred from, similarly, comparing the bid-ask spreads on the TSX and on the NYSE.[35] The test result overwhelmingly agrees with the alternative hypothesis as seen in the second column of Table A.3. As a result, the market makers on the TSX are more competitive in setting quote spreads than their competitors on the NYSE are.

Relative dominance of the TSX over the NYSE in terms of information share

[33] H_0 : $\text{PIN}_{\text{TSX}} = \text{PIN}_{\text{NYSE}}$ versus H_1 : $\text{PIN}_{\text{TSX}} > \text{PIN}_{\text{NYSE}}$.

[34] Canadian insider trading was no less egregious than that of the U.S. until 2003 when the anti-white collar crime act was legislated under the Criminal Code. See King and Segal (2004) for an excellent survey on this issue.

[35] H_0 : $\text{SPREAD}_{\text{TSX}} = \text{SPREAD}_{\text{NYSE}}$ versus H_1 : $\text{SPREAD}_{\text{TSX}} < \text{SPREAD}_{\text{NYSE}}$.

can be empirically checked[36] and the test confirms that the information share of the TSX, on average, far exceeds that of the NYSE as seen in the third column of Table A.3. Thus, the TSX contributes more to price discovery than the NYSE does.

In order to check for robustness of the Wilcoxon test results shown in Table A.3, I construct a monthly panel dataset of the PIN, spread, the information share, volume,[37] and the TSX indicator.[38] In Panels A, B, and C of Table A.4, the PIN, spread, and the information share are, respectively, regressed against the others controlling for volume and the TSX dummy variable. The signs of the binary TSX variable in Models 2 and 3 confirm the results shown in Table A.3. Trade informedness (PIN) is graver on the exchange with a higher information share (vice versa) as shown by Models 1 and 2 in Panel A (Panel C) of Table A.4.[39]

In summary, I find that for the cross-listed of pairs the TSX assumes leadership in price discovery and also shows an, overall, higher PIN than the NYSE. In other words, the trading venue with heavier intensity of informed trades contributes more to the price discovery of cross-listed pairs. This is explicit empirical evidence that informed traders catalyze cross-border price discovery.

Eun and Sabherwal (2003) conclude that informed traders prefer to trade in a market where more original information can be found. By extension, I use direct relative measures of informed trades (PIN) and contribution to price discovery (information share). The trades executed on the lead exchange, TSX, are more

[36] $H_0 : \mathrm{IS}_{\mathrm{TSX}} = \mathrm{IS}_{\mathrm{NYSE}}$ versus $H_1 : \mathrm{IS}_{\mathrm{TSX}} > \mathrm{IS}_{\mathrm{NYSE}}$.

[37] The monthly average of the logs of total daily trading volumes.

[38] Equals one if the estimated numerical value is of the TSX, or zero if the NYSE.

[39] Further, the panel regression results in Table A.4 are robust to Fama and MacBeth's (1973) test.

likely to be information-based than the trades executed on the lag exchange, NYSE. The PINs of a cross-listed pair represent the proportions of exchange-specific informed traders.

1.5.2 Dynamics of relative premiums

The Canadian listings on the NYSE, on average, carry slightly positive and highly volatile relative premiums relative to their home listing on the TSX through the sample period (Panel A of Table A.2). Throughout trading hours, the pairs appear to be fairly priced and the small premium on an average NYSE-cross-listing carries against its original TSX-listing is not surprising given the implication of the extended Grossman and Stiglitz (1980) model presented in Section 2.[40]

As the relatively high standard deviation suggests, there evidently are more-than-profitable, but short-lived, relative premiums which subsequently attract pairs traders. It is natural to ask how quickly *and* by whom a temporarily profitable relative premium is pushed back towards parity. Following Gagnon and Karolyi (2004), I estimate the convergence speed parameter in a daily frequency for each firm. The PIN effect on the convergence speed can be inferred from regressing the convergence speed parameter ($\text{SPEED}_{\text{CONV}}$) onto the average PIN on both exchanges, since convergence speed is a mutual concept, and average spread (on both exchanges), controlling for firm size,[41] industry dummy,[42] vol-

[40]In Section 2, based on an extended version of the noisy rational expectations model of Grossman and Stiglitz (1980), I derive an implication of "home market liquidity dominance": higher-PIN TSX-listed stock must be priced lower than its NYSE-listed share in a "no-arbitrage" equilibrium.

[41]The average log market capitalization on the TSX and the NYSE.

[42]Equals one if the cross-lister is a manufacturing firm, and zero otherwise.

ume,[43] and governance index[44] as follows

$$\text{SPEED}_{\text{CONV}} = \gamma_1 \text{PIN}_{\text{AVG}} + \gamma_2 \text{SPREAD}_{\text{AVG}} + \gamma_3 \text{SIZE} + \gamma_4 \text{INDUSTRY} + \gamma_5 \text{VOLUME} + \gamma_6 \text{GOVERNANCE} + \eta.$$

According to the regression model, the dynamics of synchronous relative premiums is explained by the asymmetric information component (PIN) and market friction (spread) while holding liquidity constraint (volume) and firm characteristics (size, industry, and the level of corporate governance) constant.

It turns out that, in Panel A of Table A.5, a higher PIN on either exchange very significantly impedes the convergence to parity in all specifications, since the convergence speed parameter is reciprocal to actual speed. This is against the second hypothesis raised in Section 3. The uninformed traders appear to deplete relative premiums faster than their informed cohort. The PIN effect appears robust controlling for liquidity of cross-listed pairs in Models 2, 3, and 4. The higher the spread on either exchange (the higher the average spread as a result) the slower the convergence speed in Models 1, 2, and 4.

Practitioners executing statistical arbitrages (pairs trades) and profiting from relative premiums need not be informed of the issuer's fundamental value. Timely execution and unwinding of their positions will suffice. Thus, statistical arbitrageurs are believed to be discretionary liquidity traders who are responsible for quickly converging *and* low-PIN cross-listed pairs.[45] This novel finding is among a few articles in the literature relating the dynamics of premiums and

[43]The log of total daily trading volume on the TSX and the NYSE, respectively.

[44]The Report on Business governance index of Canadian firms is published by *Globe and Mail* (McFarland (2002)). Full scores in the four following criteria total up to 100 points: board composition (40), compensation (23), shareholder rights (22), and disclosure (15). The higher the index score, the better the firm is governed. I appreciate Stephen Foerster and Michael King for sharing this information.

[45]Admati and Pfleiderer (1988) distinguish discretionary liquidity traders who can skillfully and strategically time their executions, in contrast to non-discretionary liquidity (or noise) traders.

discounts on pairs of cross-listed shares to information asymmetry. Similarly, Gagnon and Karolyi (2009a) find that, controlling for various proxies of information asymmetry, holding costs of long-short portfolios of cross-listed pairs significantly explain the cross-sectional and time-series variation in price parity deviations.

I further explore the cross-sectional relationship between the average spread across the exchanges against the average PIN on both exchanges, and convergence speed, controlled for firm size and industry dummy. In Panel B of Table A.5, the average PIN is very significantly positively associated with the average spread which is consistent with the finding of Easley, Kiefer, O'Hara, and Paperman (1996).

$$\text{SPREAD}_{\text{AVG}} = \delta_1\,\text{PIN}_{\text{AVG}} \;+\; \delta_2\,\text{SPEED}_{\text{CONV}} \;+\; \delta_3\,\text{SIZE} \;+\; \delta_4\,\text{INDUSTRY} \;+\; \epsilon.$$

Panel B of Table A.2 shows that a higher PIN of a stock listed on the TSX gives rise to a positive premium in the stock cross-listed on the NYSE. This relation is robust to controlling for convergence speed and governance index as shown in Panel C of Table A.5. This provides further support for the extended Grossman and Stiglitz (1980) model presented in Section 2.

$$(p_{\text{NYSE}} - p_{\text{TSX}})\,/\,p_{\text{TSX}} \;=\; \beta_1\,(\text{PIN}_{\text{TSX}} \;-\; \text{PIN}_{\text{NYSE}}) + \beta_2\,\text{SPEED}_{\text{CONV}} + \beta_3\,\text{GOVERNANCE} + \epsilon.$$

One arbitrageur may prefer to short-sell on the NYSE and to long on the TSX, while another to short-sell on the TSX and to long on the NYSE for liquidity reasons. This may render using the quote mid-points of U.S.-Canada exchange rate problematic.[46] For example, it may be easier to short-sell on the TSX than

[46] I owe this point to Bhagwan Chowdhry. In other words, dynamics in the foreign exchange market are another source of innovation to the cointegrated system of cross-listed pairs (Grammig *et al.* (2005)).

on the NYSE. The *cross-border relative quoted spreads* are defined as follows

$$\widetilde{\text{SPREAD}}_{\text{NT}} \equiv \{\text{ask}_{\text{NYSE}} - \text{bid}_{\text{TSX}} \cdot (\text{US\$/CAN\$})_{\text{ask}}\} / \{\text{bid}_{\text{TSX}} \cdot (\text{US\$/CAN\$})_{\text{ask}}\},$$

$$\widetilde{\text{SPREAD}}_{\text{TN}} \equiv \{\text{ask}_{\text{TSX}} \cdot (\text{US\$/CAN\$})_{\text{bid}} - \text{bid}_{\text{NYSE}}\} / \text{bid}_{\text{NYSE}}.$$

$\widetilde{\text{SPREAD}}_{\text{NT}}$ is the percentage cross-border arbitrage profit from buying on the TSX and selling on the NYSE, and $\widetilde{\text{SPREAD}}_{\text{TN}}$ is from buying on the NYSE and selling on the TSX. The first strategy narrows down $\widetilde{\text{SPREAD}}_{\text{NT}}$, while the second pairs trade squeezes $\widetilde{\text{SPREAD}}_{\text{TN}}$. Either strategy may turn out more lucrative than the other due to the existence of bid-ask spread in the exchange rate.

In Table A.6, monthly averages of cross-border relative quoted spreads (updated every minute) of 56 cross-listed pairs are tested for differences using the Wilcoxon test. It turns out that the two spread measures are empirically equivalent. In other words, arbitrageurs' positions are not skewed towards either trans-Niagara trading venue due to exchange rate market friction. Thus, using exchange rate mid-quotes appears reasonable.

1.5.3 Cross-listing effects on the home exchange

Table A.7 shows fifteen Canadian firms that cross-listed on the NYSE during the sample period. Twelve firms had been listed on the TSX before they cross-listed on the NYSE. The firms without the PIN either have cross-listing dates too near the end of the sample period or are insufficiently liquid. For the PIN estimates before and after cross-listing events, there are eight pairs with a six-month window, six pairs with a twelve-month window, and nine pairs with an exhaustive window.

The arithmetic means of the columns of the PIN show that they rise around the cross-listing events. The pre- versus post-cross-listing scatter plots are provided for respective event windows in Figures A.3, A.4, and A.5. The PIN on the TSX, on average, rises upon cross-listing on the NYSE within all event windows. The significance of the PIN increase (rise in the relative degree of adverse selection) around cross-listings can be verified by the Wilcoxon test with the difference in PINs before and after cross-listings.[47]

In Panel A of Table A.8, each of the null hypotheses against the alternative hypotheses are rejected at a 10% right-tail significance level. This result that the PIN rises (or that the intensity of private information increases) on the home exchange upon cross-listing unifies and extends the existing claims in the cross-border finance literature.

Cross-listing lowers transaction costs and narrows the spreads on the TSX and, resultantly, reduces noise trader risk (Eun and Sabherwal (2003), Fleming, Ostdiek, Whaley (1996), and Jones and Seguin (1997)), or subdues excessive volatility borne by liquidity trades. The more grave degree of adverse selection in the home market shown in Panel A of Table A.8 is the first documentation of *relative* cross-listing effects on the home exchange information environment.[48] The aforementioned articles only mention the decrease in *absolute* magnitude of noise trades.

The TSX-listed firms, on average, post negative cumulative abnormal returns (CARs) within all event windows around cross-listings in Table A.7. This result

[47] $H_0 : \text{PIN}_{+3\text{M}} = \text{PIN}_{-3\text{M}}$ versus $H_1 : \text{PIN}_{+3\text{M}} > \text{PIN}_{-3\text{M}}$,
$H_0 : \text{PIN}_{+6\text{M}} = \text{PIN}_{-6\text{M}}$ versus $H_1 : \text{PIN}_{+6\text{M}} > \text{PIN}_{-6\text{M}}$,
$H_0 : \text{PIN}_{\text{after}} = \text{PIN}_{\text{before}}$ versus $H_1 : \text{PIN}_{\text{after}} > \text{PIN}_{\text{before}}$.

[48] In a comparable case, Chan, Menkveld, and Yang (2008) report that the PIN on B shares in China (that had only been legally traded by foreign investors) rises on opening access to locals.

bears resemblance to that of Foerster and Karolyi (1999). It is reasonable that Canadian firms who cross-list in the U.S. do not benefit from lower costs of capital. Unlike those in the emerging market economies, Canadian managers can easily diversify their financing risk across the border.

There appears to be no discernable relative premium due to diminished market incompleteness (Merton (1987)) for Canadian cross-listers in the U.S. The higher post-cross-listing PIN intuitively explains the negative event study returns on the home-listed stocks. As the original TSX listings become more concentrated with private information, they must reflect relative discounts in equilibrium, as in Easley, Hvidkjaer, and O'Hara (2002).

$$\begin{aligned}\text{RETURN}_{\text{AB}} &= \beta_0 + \beta_1\,\text{PIN} + \beta_2\,\text{CROSS-LIST} + \beta_3\,\text{SPREAD} + \beta_4\,\text{VOLUME} + \beta_5\,\text{VOLATILITY}\\ &\quad + \beta_6\,\text{PIN}\times\text{CROSS-LIST} + \beta_7\,\text{SPREAD}\times\text{CROSS-LIST}\\ &\quad + \beta_8\,\text{VOLUME}\times\text{CROSS-LIST} + \beta_9\,\text{VOLATILITY}\times\text{CROSS-LIST} + \epsilon.\end{aligned}$$

Accordingly, the negative abnormal returns on the TSX-listed stocks upon cross-listing on the NYSE are associated with heavier trade informedness in Table A.9. In the fixed-effect panel regression analyses, the abnormal returns[49] ($\text{RETURN}_{\text{AB}}$) on the original listings on the TSX are regressed, on a monthly basis, onto the PIN, cross-listing dummy,[50] spread,[51] volume,[52] return volatility,[53] and cross-listing interaction terms. Once cross-listed, the home-listed stocks' underperformance against the market typically magnifies as the intensity of informed trades (PIN×CROSS-LIST) increases. This relation is robust to controlling for spread, volume, and volatility measures which do not appear as economically and statistically significant as the PIN after cross-listings on the NYSE.

[49] The monthly abnormal return, following Binder (1998), using the S&P TSX Composite Index.
[50] A dummy variable which equals one in the month of cross-listing event, or zero otherwise.
[51] The monthly average relative quoted spread.
[52] The monthly average of the logs of total daily trading volumes.
[53] The standard deviation of daily returns multiplied by 250/12.

Panel B of Table A.8 shows that the bid-ask spreads evidently narrow after cross-listing events over the exhaustive threshold window (before and after cross-listing through the sample period), a finding consistent with Foerster and Karolyi (1998). Whether Canadian firms' cross-listings on the NYSE facilitate enhanced volume[54] on the home exchange is shown in Panel C of Table A.9. Statistically, the incremental effect of cross-listing on home market liquidity is not strong, perhaps due to the limited sample size. This may also reflect Karolyi's (2006) summarizing remark that "... Price discovery does not necessarily originate in the markets with the highest relative turnover, but rather where the informed traders are going with limited market impact."

The above findings suggest that, at least within integrated economies, cross-listings boost the intensity of private information-based trades in home-listed stocks. A higher proportion of informed traders is a double-edged sword: it fosters price discovery *and* exacerbates adverse selection. This shift in information environment lends support to the claim of Bailey, Karolyi, and Salva (2006) that cross-listings may not reduce information asymmetry. The managers of Canadian firms may have been led to trading on inside information upon cross-listings that resulted in undermining their existing shareholder values given the comparatively lax insider trading environment on the TSX (King and Segal (2004)) during the sample period. The result herein may contradict the bonding hypothesis (Coffee (1999)) which states that insiders have "less" incentive to trade after cross-listings.

[54]The log of daily total trading volume of TSX-listed shares of NYSE-cross-listed Canadian firms.

1.6 Conclusion

In this paper, I address how information asymmetry determines relative pricing of Canadian stocks that trade across the Niagara Falls. The theoretical prediction is empirically supported with evidence on Canadian shares listed on both the Toronto and New York stock exchanges, from January 1998 through December 2000. The three key results reveal "information asymmetry is priced" across the border, across time, and around cross-listing events. Overall, the PIN proves to be a useful for understanding the effect of asymmetric information on stock trading fragmented across an international border.

My first empirical finding reveals that, on average, the TSX leads the NYSE in price discovery (measured by information share) and shows a higher PIN. In other words, the exchange with greater intensity of informed trading contributes more to price discovery. This is explicit *cross-border* evidence that informed traders stoke price discovery. However, the higher proportion of informed traders on the TSX is, likely to be, due to the comparatively lax regulatory environment therein by then. Second, I find that New York and Toronto prices of lower-PIN stocks converge more rapidly. Specifically, a preponderance of discretionary liquidity traders yields a low PIN, and some of them attempt to arbitrage the relative premium when there are fewer informed traders around. This novel finding is among a few notable documentations that relate the dynamics of premiums and discounts on home versus foreign listings to asymmetric information. Finally, on average, the PIN on a TSX-listed stock rises upon cross-listing on the NYSE. This finding of *relative* cross-listing effects on the home market information environment not only explains negative cross-listing announcement event study returns but also unifies and extends existing findings in the literature. Previous articles mention a reduced noise trader risk as a result of decreased transaction costs on the home exchange following cross-listings. That fragmentation due to cross-listing aggravates adverse selection, or exacerbates transparency in order flow, on the home exchange may offer useful insight into recent research direction in market microstructure.

There are numerous unresolved issues for cross-listings between integrated markets. The consequences of cross-listings by Canadian firms I have shown imply that insiders may trade more on hidden corporate information in their home market as their companies cross-list overseas. This is likely to be a downside of cross-listing. As this contradicts the bonding hypothesis, I leave a testable hypothesis for future research. Cross-listing emerging market firms may warrant higher event study returns on their home exchanges than for developed country firms. This is possible since the former group's bonding effect is dominant while the latter group's adverse selection aggravates like I have shown in the paper.

Lastly, I would like to mention limitations of this study. Unlike in Gagnon and Karolyi (2009a), the herein defined relative premiums may overstate the actual relative arbitrage returns since I do not account for arbitrage costs. The cross-listing effect on the TSX can be assessed in a finer detail by estimating effective transaction costs per Bessembinder and Kaufman (1997), Hasbrouck (2009), and Tiwari (2004). By incorporating random shocks from the foreign exchange market (Grammig *et al.* (2005)) into quote analysis of the cross-listed pairs on the TSX and the NYSE, there can be more implications to shed light on unanswered questions in the literature.

CHAPTER 2

CORPORATE GOVERNANCE & CROSS-BORDER ACQUIREE RETURNS

2.1 Introduction

Mergers and acquisitions are the most frequent means of corporate control transfer, and each deal potentially creates value to be transferred among parties. For example, Jensen (1993) finds that the market for corporate control is a key conduit for the efficient handling of excess capacity and further points out that, in the absence of such a market, weak internal control generally delays value-creating restructuring. Even when no explicitly measurable wealth is created, existing value may still be transferred among parties, so it is reasonable to investigate the role of corporate governance in mergers and acquisitions. Specifically, we explore here whether mergers and acquisitions act as a tool to enforce corporate governance discipline, as firms seeking to benefit from the corporate control market must first convince participants that they are well governed. Following a well-developed corporate governance literature, we study anti-takeover provisions (ATPs) as a proxy for corporate governance under the observation that firms with more ATPs are more strongly protected and hence more difficult for the control market to punish should they behave unscrupulously, i.e. more ATPs signifies weaker governance.

We find that, in domestic U.S. deals, targets exhibit higher abnormal returns upon merger announcement when their bidders are more poorly governed, consistent with the notion that the market expects targets of poorly governed firms to be relatively overpaid. Our study is the first to document this relation in the context of ATPs, though other studies find consistent qualitative conclusions using different measures. We also investigate cross-border deals, as the multiple market nature of these deals presents additional unique and non-negligible take-over challenges, and find that overseas targets also benefit more from poorly governed U.S. acquirers than well governed ones. However, surprisingly, in cross-border deals, poorly governed acquirers also enjoy greater merger announcement premiums than their well-governed counterparts. This finding is contrary to both the aforementioned intuition and the U.S. domestic market findings of Masulis *et al.* (2007) and others. While we conjecture possible explanations, we leave robust resolution of this puzzle as a topic of future research.

It is conventional wisdom that, in the market for control, bidders lose in the stock market when they overpay for targets. Masulis *et al.* (2007) finds that the acquirers with more ATPs perform worse than their peers at merger announcement, ostensibly because firms with more ATPs are more poorly governed and hence tend to overpay on overage. As suggested by Gompers *et al.* (2003), these so-called "dictator" acquirers, who are more difficult to take over and who are more entrenched, generate premiums for targets because their corporate structures offer weak or no disincentives for reckless behavior, making them more prone to empire building and other such agency concerns.

This relation, however, is confounded in the cross-border context. These deals, while an intriguing and legally more complicated environment in which to test the existence of value effects, represent changes in investor protection related to both corporate governance and sovereign legal systems, the latter in turn affecting the characteristics and financial markets of the firms themselves. La Porta *et al.* (1998) reports a significant relation between legal protection and the development of financial markets, while subsequent work by Lubrano (2003) further documents that improvements in corporate governance contribute to the maturity of capital markets. Claessens and Laeven (2003) similarly notes that firms in countries with better-enforced property rights enjoy better growth prospects. Indeed, it has been shown that the legal environment potentially impacts the financing decision (Demirgüç-Kunt and Maksimovic (1999)); market efficiency (Mørck *et al.* (2000)); the degree of foreign exchange collapse (Johnson *et al.* (2000)); capital allocation (Wurgler (2000), Beck and Levine (2002), and Claessens and Laeven (2003)); and even firm valuation (La Porta *et al.* (2002) and Himmelberg *et al.* (2002)).

However, it is likewise the case that firms may execute an optimal level of self-imposed governance by applying their own guidelines in addition to extant legal boundaries. So long as contracts are enforceable and abided by, investors can be protected to the same extent by all firms regardless of the legal environment, ergo, the legal environment alone cannot determine the totality of investor protection. Along those lines, Bris and Cabolis (2008) documents that, despite controls for cross-country differences, differences in firm-level corporate governance generate significant value effects for the merged entity, though their work uses accounting standards derived from sovereign-level measures of La Porta *et al.* (1998), as its measure of corporate governance. Likewise, La Porta *et al.* (2002) demonstrates that benefits from better macro-level shareholder protection are more pronounced when CEOs' cash flow rights are relatively small. Indeed, a lengthy literature suggests that even private contracts and ATP specifics are priced, among these Bebchuk *et al.* (2004), Bebchuk and Cohen (2005), and Cremers and Nair (2005). Stated simply, both the literature and common sense suggest that cross-country variations caused by differences in the macro-environment and cross-sectional differences in governance at the firm level both impact investor protection. Our study seeks to control for and separate these two governance factors.

We first document that, for domestic U.S. deals, acquirers with more ATPs in place pay a higher premium. This serves to confirm the intuition that managerial recklessness leads to higher premiums and benefits target shareholders. However, in addition to domestic deals, our study confirms using cross-border deals that, despite controls for sovereign legal systems and investor protection, it is nonetheless the case that U.S. acquirers that are poorly governed tend to pay more than their well-governed counterparts, benefiting overseas target shareholders. Similarly, Starks and Wei (2004) finds acquirers from countries with superior sovereign-level corporate governance pay smaller acquisition premiums, but do not examine the effects of firm-level governance. While we investigate U.S. acquirers of foreign targets, Kuipers *et al.* (2003) shows that foreign acquirers from countries with better investor protection pay more for U.S. targets. However, their study also falls short of investigating firm-level corporate governance.[1] By proxying for both firm-level governance and country-level factors, we conclude that not only geographic location and differences in legal systems but also poor firm-level governance of the acquirer account for the premium paid to foreign target shareholders (henceforth the "dictator premium").

In addition, we also document that poorly governed U.S. acquirers enjoy higher post-announcement returns than their well-governed counterparts in cross-border deals. In contrast, Masulis *et al.* (2007) reports no benefits for

[1] In less closely related work aimed at cross-border deals, Doukas and Travlos (1988) shows that the announcement effect is greater if the acquirer is a first-time entrant into the foreign market. Chari *et al.* (2004) finds the acquirer's return is more positive when the it attains management control of the target.

poorly governed acquirers in U.S. domestic deals. Importantly, our study includes only publicly listed acquirers and targets while Masulis *et al.* (2007) includes private U.S. domestic targets in their sample. Public firms are generally more transparent, are followed by more analysts, and are thus more closely monitored. While this difference in data may explain this disparate finding, we leave full resolution of this puzzle to future work.

The remainder of this paper is organized as follows: section 2 discusses existing theories that illustrate our key intuition and generate testable hypotheses, section 3 describes the data and empirical methodology employed, section 4 presents empirical and estimation results, and section 5 concludes.

2.2 Theory and hypotheses

There is an extensive literature on agency problems and corporate governance. Indeed, the agency literature speaks to the tension between shareholders and managers and can date as far back as Berle and Means (1933) and certainly Jensen and Meckling (1976). More specifically addressing corporate governance, Core, Holthausen and Larcker (1999) provides support for a negative relation between corporate governance and agency concerns, further documenting that poorly governed firms under-perform. Jensen and Ruback (1983), Jarrell *et al.* (1988), and Andrade *et al.* (2001) each extensively examine this topic.

Our focus is on how these findings impact the corporate control market. Mørck *et al.* (1990), Lang *et al.* (1991), and Jensen (1986) all conclude that, while acquisitions may substantially benefit the firm, it generally benefits managers and managers may themselves generally focused on personal interests rather than the good of the shareholders.[2] In response, Mitchell and Lehn (1990) shows that the market for corporate control can also serve to suppress agency problems since managers that make unwise acquisitions are in turn more likely to be taken over themselves. Hence, the extent to which the firms may be taken over may directly relate to the behavior of managers. Those who are well-entrenched may be more prone to agency problems since it is then comparatively difficult to punish them. As Bebchuk *et al.* (2002, 2003) and Field and Karpo (2002) explain, ATPs make takeovers difficult as they significantly slow the takeover process, generate higher transactions costs, and hence curb the incentives of potential bidders to acquire firms in a hostile takeover. Thus, we posit that number of ATPs is a proxy for managerial recklessness and is a negative measure of the firm's commitment to investor protection: the greater the number of ATPs, the worse a firm's corporate governance.

2.2.1 U.S. domestic deals

The impact of firm-level governance on the market for corporate control has been oft studied in the domestic environment, and our goal is to test for a consistent finding in our domestic U.S. dataset. Namely, do U.S. targets respond more favorably to domestic bidders with high ATPs?

In this study, we choose to focus on target firms. In contrast, Masulis *et al.* (2007) shows that acquirers with more ATPs have more negative announcement period returns than their low ATP counterparts. They conclude that acquiring firms pay higher merger premia, which is reflected in a more negative return.

[2] Yermack (2006) shows that corporate size, i.e. empire building, is directly associated with executive perquisites, implying that mergers and acquisitions may be self-serving and generate positive externalities for executives at the cost of shareholder value.

However, a variety of factors affect acquirer returns that are difficult to quantify and control for since the acquirer will be the on-going concern and is often a larger, more complex business entity than the target. We focus on targets and contend that when high ATP bidders acquire targets, target shareholders extract greater returns since acquirers overpay. This is often essentially a cash-out scenario, especially for high-level managers making the decision, so any overpayment transfers directly to shareholders. Investor appraisal of merger effects on the targets have received less than due attention. Although Huang and Walkling (1987) and Song and Walking (1993) relate merger wealth effect on targets with deal characteristics and ownership stake, respectively, there is a dearth of scholarly articles that reason the value of private contracts that foster managerial recklessness, i.e. ATPs, enforced by the acquirers being impounded on the target returns upon merger announcements. In other words, we argue that the target returns are a more visible and less complicated measure of value transfer through mergers. Specifically:

> *H1 [Dictator Premium Hypothesis]: Ceteris paribus, the higher the ATP index of the acquirer, the higher the cumulative abnormal return (*CAR*) of the domestic target.*

2.2.2 Cross-border deals

Cross-border deals are growing in popularity and generate unique and important sources of value. As documented by Alexander (2000), cross-border takeovers may result from *1.* intensive conglomeration as a method of preemptive restructuring or generating economies scale, *2.* response to technological innovation, *3.* need for a global marketing platform, *4.* absence of domestic merger targets, and 5. the desire or need to expand into new markets. As Starks and Wei (2004) finds acquirers from countries with superior sovereign-level corporate governance pay smaller acquisition premiums, but do not examine the effects of firm-level governance, there remains empirical verification of whether recklessness of individual U.S. acquiring firm managers can still, unintentionally, benefit their foreign target shareholders by impulse-driven overpayment. Extending the intuition from the domestic deals scenario, we hypothesize that bidder with more ATPs will likewise overpay in foreign acquisitions such that:

> *H2 [Cross-border Dictator Premium Hypothesis]: Ceteris paribus, the higher the ATP index of the acquirer, the higher the* CAR *of the cross-border target.*

2.3 Data and methodology

The key piece of data for this study is our measure of corporate governance, the ATP index. For each firm, the index is calculated as the total number of ATPs listed in the firm's articles of incorporation, the higher number of ATPs, the higher the index. This method mimics the methodologies of Gompers et al. (2003)'s "*G* Index" and Masulis *et al.* (2007) , which are both based upon 24 total possible ATPs.[3] These papers define democracy (dictatorship) as firms with an index less than or equal to 5 (greater than or equal to 14) and less than or equal to 9 (greater than or equal to 10), respectively. We follow the latter of

[3]Bebchuk, *et al.*'s (2004) index is based on only 6 ATPs while Cremers and Nair's (2005) index is composed of only *3.* Bebchuk and Cohen's (2005) applies a binary variable based on whether a firm has a staggered board.

these papers in our definition.

The data required is held at the Investor Responsibility Research Center (IRRC) which contains firm-level data for publication years 1990, 1993, 1995, 1998, 2000, 2002, 2004, and 2006. Nearly 1,500 firms are represented in the dataset each year, including the S&P 500 and those on the Forbes, Business Week, and Fortune magazines' lists of largest U.S. corporations. On average, the database represents about 90 percent of U.S. stock market capitalization, though more recent years' data are more inclusive. Following the literature, we assume that index does not change between publications years.

2.3.1 M&A data

We acquire deal data from the Securities Data Corporation (SDC). For domestic deals, we collect data on 1456 acquisitions between January 1990 and December 2007 that meet the following criteria: *1.* both target and acquirer are public companies incorporated in the U.S., 2. transaction value exceeds $1 million, *3.* the acquirer controls less than 50% of the target's shares prior to the announcement and owns 100% of the target's shares after the transaction, and *4.* the acquirer has annual financial statement information available from Compustat and stock return data available for at least 210 trading days prior to acquisition announcement. We then match this with our ATP index, reducing the sample to 1439 domestic acquisitions. Figure B.1 illustrates the number of mergers and valuation multiples (price/target earnings) for all deals in our dataset. In Panel A, we see that domestic deal flow and multiples peaked during the Internet Bubble period of the late 1990s.

For cross-border deals, data is available starting in 1984. We apply criteria that mirror the aforementioned with the exception that the target must be a public firm incorporated outside of the U.S. There are 1,024 such cross-border deals with an average deal size of $359.24 million, for an average stake of 59%. The most active U.S. acquirers in these deals are Citigroup, Coca-Cola, and Merrill Lynch (12 deals each), followed by Microsoft (11 deals). In total, 57 countries are represented with Canada having the largest number of deals, accounting for 25.3% (259 deals) of the total, followed by the U.K. (17%; 174 deals), and Australia (8.3%; 85 deals). In terms of industry, using SIC codes, we find that Crude Petroleum and Natural Gas and Prepackaged Software (44 deals each) are most common, followed by Pharmaceutical Preparations (32 deals) and Gold Ores (28 deals). We then reduce the dataset to match the time period for domestic deals and then merge these data with that of our ATP index, reducing the dataset to 599 cross-border deals. Figure B.1 Panel B shows the number of and valuation multiples for cross-border deals. Once again, the number of deals peaks in the late 1990s but multiples are lowest during that period, peaking instead near the beginning and end of our data set.[4]

For sovereign legal system indicators, we source accounting standards from La Porta *et al.* (1998) and anti-director rights-which proxies for the degree of shareholder protection-from Djankov *et al.* (2008). We also calculate the ratio of stock market capitalization to GDP as a relative measure of country-specific equity market development, also suggested by Djankov *et al.* (2008). Table B.1 summarizes this data.

[4]The flattening out of multiples for cross-border deals in the mid and late 1990's may be related to currency devaluations experienced during the Asian financial crisis.

2.3.2 Calculating CAR

Following a sizable literature in this field and others, we apply a short-term event study methodology that analyzes CAR surrounding deal announcements. Previous work applying event study techniques around announcement of ATP adoption or amendment includes DeAngelo and Rice (1983), Linn and McConnell (1983), Malatesta and Walkling (1988), and Ryngaert (1988).[5]

For domestic deals, our tests are relatively straightforward. We measure CAR using a market model adjusted for market risk for the [-2,+2], [-5,+5], and [-10,+10] windows around the announcement date. Specifically, following Masulis *et al.* (2007) we use the CRSP equal-weighted return as our measure of market return and estimate the market model over the 200-day period starting 210 days before the event, ending 11 days before the event to capture stock run-ups.[6] That is:

$$R_{ik\tau} = \alpha_{ik} + \beta_i^m R_{m\tau} + \epsilon_{ik\tau} \quad \forall\, \tau \in [-210, \cdots, -11],$$

where $R_{i\tau}$ is the daily return for domestic target i with domestic acquirer k, and $R_{m\tau}$ is the aforementioned equal-weighted market return. Following the standard practice, we then apply this β_i^m, calculate a predicted abnormal return for the event window, and subtract it from the actual return to arrive at CAR:

$$\text{CAR}_{ijt}[\pm d] \equiv \left\{ \prod_{\tau=-d}^{+d} \left(1 + \widehat{\epsilon_{ij\tau}}\right) \right\} - 1, \quad \forall\, d \in \{2, 5, 10\} \text{ and for annoucement date } t.$$

For cross-border deals, we require some minor revisions. We estimate a dollar-translated market model for predating days $[210, \cdots, -11]$, following Bris and Cabolis (2008), and include both the target's home market index and the

[5]Bhagat and Romano (2002) provides an extensive survey.
[6]Masulis *et al.* (2007) also applies this method.

MSCI world index. Specifically, we have:

$$R_{ijk\tau} = \alpha_{ijk} + \beta_i^m R_{mj\tau} + \beta_i^w R_{w\tau} + \epsilon_{ijk\tau} \quad \forall \tau \in [-210, \cdots, -11],$$

where $R_{ijk\tau}$ is the daily return for foreign target i based in country j with U.S. bidder k, $R_{mj\tau}$ is the market index return in country j, and $R_{w\tau}$ is the return on the MSCI world index. We calculate CARs for the same event windows, then, in an analogous fashion:

$$\text{CAR}_{ijkt}[\pm d] \equiv \left\{ \prod_{\tau=-d}^{+d} \left(1 + \widehat{\epsilon_{ijk\tau}}\right) \right\} - 1, \quad \forall d \in \{2, 5, 10\} \text{ and for announcement date } t.$$

2.4 Results

Our main results use target CAR as the dependent variable and the acquirer's ATP index as the key explanatory variable. We also apply a number of control variables including deal characteristics, acquirer firm characteristics as they may confound governance measures, and M&A market conditions.

For deal characteristics, we include log deal value and binary variables for whether it is a cash deal for whether the acquirer and target are in a high-tech industry.[7] Acquirer firm characteristics include Tobin's (1969) Q, leverage ratio, free cash flow ratio, relative deal size, and whether it is a diversifying acquisition. Finally, our proxy for market conditions is the average premium paid to targets in a given year. The general form of our regression is as follows:

$$\begin{aligned} \text{CAR} &= \beta_1 \cdot G\ Index + \beta_2 \cdot Deal\ Characteristics \\ &\quad + \beta_3 \cdot Acquirer\ \&\ Target\ Characteristics \\ &\quad + \beta_4 \cdot M\&A\ Market\ Condition + error\ term. \end{aligned}$$

[7]The high-tech dummy is as defined by Loughran and Ritter (2004).

2.4.1 Domestic dictator premium

In Table B.2, we present the results of tests run using only domestic deals. In virtually every test, we find that the *G* index is positively related to target CAR, i.e. the weaker the governance, the higher returns to target shareholders, implying expected overpayment on the part of acquirers. This is consistent with hypothesis 1 and echoes the results of the extant literature. The combination of the Masulis *et al.* (2007) finding that poorly governed acquirers lose value and our finding that the targets of these acquirers gain value suggests that these deals serve as a wealth transfer mechanism between respectively shareholders. This finding is always statistically significant, almost always at the 5% level, regardless of the size of the window used or which control variables are included.

The first set of tests results presented in *Model 1* include controls for deal characteristics. Deal size plays a role and is negatively related to returns, statistically significantly so at the 1% level. This is perhaps because, given limited resources, acquirers are unable to pay high premiums if targets are large in an absolute sense. Cash deals are more attractive as the relation between CAR and the cash dummy is positive and again generally significant at the 1% level. Cash is less risky than an equivalent market value in shares so this is not surprising. Whether the target or acquirer is a high-tech firm seems positively related to CAR, though this finding is not consistently significant. A positive relation could be due to improvements in future growth prospects, generally considered critical in high-tech firm valuations.

Results presented in *Model 2* further address acquirer characteristics as controls. The data for these controls is rather limited and reduces our sample size to 526. Results are, however, qualitatively, unchanged. Most importantly, target

CAR is positively related to the *G* index, hence negatively related to governance. Cash and deal size coefficients remain significant, though the remaining controls are only significant sporadically. CAR appears to be negatively related to leverage, possibly a reaction to increased risk of future financial distress, less access to future capital, or limited ability for managers to overpay given capital constraints. On the other hand, when the deal represents entrance into a different industry, i.e. acquirer and target have different SIC codes, CAR generally reacts positively perhaps indicating that greater synergies are likely to result when business units are less redundant. Notably, in *Model 3*, we control for potential irregularities related to low-priced stocks, which we define as a target stock trading at below $10.00 on announcement day, but find no such effects.

Finally, *Model 4* includes controls for M&A market conditions, important in the cyclical market for corporate control. M&A activity has been shown to be related to ease of financing related to the development of debt instruments, such as high-yield bonds in the 1980s and collateralized debt obligations in the mid 2000s. They can also be catalyzed by economic booms. Valuations can likewise be affected by these cycles. We find that, while the market condition factor is positively related to CAR as expected, the *G* index remains significantly positively related, in fact slightly more so.

In summary, controlling for deal, acquirer, and market conditions, we find that more poorly governed acquirers (higher *G* index) are expected to overpay for domestic targets, as reflected in more positive CAR. This suggests that mergers potentially act as a wealth transfer mechanism which takes place, at least partially, during the announcement window.

2.4.2 Cross-border dictator premium

The focus of this study is on cross-border deals, the degree of wealth transfer and the impact of governance at both the sovereign and firm levels on this transfer. First, consider the preliminary results presented in Table B.3 Panel A where we investigate the CAR of both acquirer and target. Consistent with Jensen (1993), CAR, referring as before to the targets' cumulative abnormal returns, are significantly positive at 13.7%, 14.0%, and 18.4% over the event windows [-2, +2], [-5, +5], and [-10, +10], respectively, with significance calculated using the Wilcoxon p-values. However, consistent with Moeller *et al.* (2004), U.S. acquirers' cumulative abnormal return (ACAR) are negative at -53bp, -73bp, and -56bp, over the same event windows, respectively, again statistically significantly so. In Panel B, we see that correlations between target and acquirer returns are likewise negative. As such, it appears that U.S. acquirers execute deals at some expenses to their shareholders, and wealth transfer appears to be substantial.

What, then, is the role of governance factors in this transfer of wealth? In Table B.4 Panel A, we show CAR for cross-border targets, separated into those with well governed acquirers (democracy) and those with poorly governed (dictatorship). We note that, for all horizons, both mean and median CAR are higher when the acquirer is poorly governed. The difference between the means is what we call "dictator premium". Despite a large standard deviation in CARs, this premium is significantly greater than 0 for all horizons, using Wilcoxon p-values, shown in Panel B. Indeed, the result is more significant, the longer is the event window. Although a full multivariate analysis has yet to come, this agrees with hypothesis *H2: targets of more poorly governed acquirers enjoy higher returns in cross-border deals.*

Interestingly, in Panel C, we test only those acquisitions where 100% of equity is transferred. Bris and Cabolis (2008) argues that a full acquisition is unique in that it effectively transfers the governance of the acquirer (as well as investor protection at the sovereign level) directly to the foreign target. In that case, one would expect that the lack of overpayment by well governed acquirers may be offset by the marginal benefit of being acquired by a less risky acquirer. In fact, we find that, in these deals, the targets of well governed firms enjoy higher CAR than those of poorly governed. This statement is true of both mean and median virtually all horizons, though owing to the vastly reduced sample size, none of these differences are statistically significant (Panel D). We continue to explore the role of full acquisitions in subsequent multivariate regressions.

2.4.3 Dictator premium concavity and acquirer returns

To add further granularity, Table B.5 Panel A lists CAR and ACAR for each value of the *G* index from 2 to 19, the respective minimum and maximum in our cross-border dataset. The figures appearing in Figure B.2 Panel A show that, for all three windows, CAR peaks when *G* is between 10 and 12, an important novel finding of this study. While CAR is generally positive and is matched with generally negative ACAR, we find that the relation to the acquirer's *G* index is clearly not monotonic. This finding is novel in the literature that require a further in-depth investigation.

In Table B.5 Panel B, we provide both the linear and quadratic fitted results where the dependent variable is once again CAR as follows:

$$\text{CAR} = Intercept + \beta_1 \cdot G\ Index + \beta_2 \cdot (G\ Index)^2 + error\ term.$$

We find that the linear model is never significant, either in F-statistic or in the t-statistic on the coefficient of the G index. The quadratic model performs better at all three horizons. The G index coefficient is positive in all three cases and the quadratic element is negative in all three cases. These findings are statistically significant for the 5-day ([±2]) and 11-day ([±5]) models, for all relevant t-statistics and F-statistics. In the absence of sound economics of such phenomenon, we can nonetheless facilitate an environment for further constructive exchanges of thoughts among our readers. Unlike U.S. domestic mergers, the cross-border nature of these deals also generates additional concerns. For the foreign target, being acquired by a U.S.-listed corporation creates a de facto cross-listing. There is an increasing interest in the recent international finance literature, including Doidge *et al.* (2004), regarding cross-listing that documents a positive cross-listing premium, i.e. targets enjoy higher returns when they are acquired by U.S. firms, arising from improvements in corporate governance due to more stringent disclosure rules as well as from increased analyst coverage of U.S. stocks. When G is particularly low, the U.S. acquiring firm is well governed, and the foreign target's CAR is low since acquirers do not overpay. However, when G is particularly high, the poor governance of the foreign acquirer is perceived as a risk to target shareholders and CAR is eroded. As a result, the highest acquiree CAR may lie in the median interval of the G index for cross-border deals. In essence, for targets of relatively well governed firms, the marginal benefit of being overpaid appears to outweigh the cost of absorbing the risk associated with a more poorly governed acquirer. Symmetrically, for targets of poorly governed firms, the transfer of poor governance and associated costs can outweigh potential overpayment.

Panel C, Table B.5, reveals that, when the dependent variable is the U.S. acquirer's return (ACAR), the linear model is found to generate a reasonable fit. In terms of each coefficient estimates, the quadratic specification is only dominant over the short 5-day ([±2]) period and it loses overall explanatory power (F-statistic) as the event window widens. For the linear model, the coefficient is positive and significant for all horizons, with significance increasing as the horizon lengthens. Puzzlingly, while previously we concluded that acquirers generally transfer wealth to targets, it appears to be the best governed of these that transfer the most (or expect to benefit the least from the acquisition). The figures appearing in Figure B.2 Panel B presents the plots for ACAR. Acquirer returns generally increase with G index and do so in a linear fashion, consistent with aforementioned results. Again, a developing literature has yet to come to a confluence as to whether and how foreign acquisitions generate value for the acquirer, but our study suggests that the most poorly governed extract the most value perhaps because investors expect the least scrupulous acquirers to extract the largest rents from unwitting foreign targets.[8] Masulis *et al.* (2007) finds the opposite for domestic deals. Poorly governed acquirers perform worse than their well governed counterparts owing to overpayment, perhaps because they are unable to benefit from the additional information asymmetry provided through cross-border deals.[9] This, though, is clearly simple conjecture. There may be hidden variables that we do not account for yet. We will later further investigate whether the concave curvature of foreign target returns survives after controlling for deal, acquirer and target characteristics, macroeconomic factors, M&A market condition, and sovereign measures of corporate governance which

[8]Dennis *et al.* (2002) argues U.S. acquirers, at the aggregate level, trade at a discount, while Doukas and Lang (2003) disagree, though both find that cross-border deals are less value-destroying than domestic ones.

[9]Note that our dataset differs from that of Masulis *et al.* (2007) in two critical ways. First, our dataset does not include private targets. Second, the sample period in that paper includes potential valuation bubbles in the late 1990s that might particularly exascerbate valuations especially of private firms.

are all possible candidates of missing links that may provide clues.

2.4.4 Sovereign vs. corporate governance impact

We now turn our attention to separating the impact of corporate governance differences from sovereign legal system variability. As aforementioned, in cross-border deals, both the governance of the firm itself and the impact of the legal systems involved may have an impact on the ability to transfer wealth and the resulting shareholder response. Specifically, we employ La Porta *et al.*'s (1998) and Djankov *et al.*'s (2008) measures for sovereign legal systems as a control in our tests. Presented in Table B.6 Panel A are CARs for dictatorship and democracy acquires separated by the anti-director rights (AD), accounting standards (AS), and stock market cap to GDP ratio (SMCTG) measures. High and low indicate values higher or lower than the median score. Importantly, in this case, we measure the target's sovereign governance since all acquirers are from the U.S. so that a target from a country with a high score implies that the acquirer is relatively weakly governed at the country level.[10]

Through all event windows, CAR is, on average, higher in countries high sovereign governance. That is, when the acquirer's sovereign governance is relatively weak, it is more likely to overpay, a finding consistent with our general conclusions and with the notion that the target receives little or no additional governance protection because of the U.S. legal system . Within low sovereign governance countries, the dictator premium is negligible. However, within high sovereign governance countries, dictatorship CAR always exceeds that of democracy. The difference is statistically and economically significant. We con-

[10]The U.S. is in the high classification for all three classifications.

clude that the dictator premium does not necessarily arise from the fact that the target resided in a country with a qualitatively inferior sovereign system. For high groups of each sovereign governance measure, the dictator premium appears ostensible with, overall, statistically significant differences, in Panel B, over all event periods. When the U.S. acquirer is under dictatorship its foreign target typically and always experiences a higher event study return when the target country well protects its investors than not and/or it has a relatively well-established capital markets than not, as shown in Panel C. This home-country effect also tends to hold for a well-behaving U.S. acquirer but the data softens for accounting standards over 11-day and 21-day event windows.

Panel D shows symmetric results for ACARs, analogous to Panel A for foreign target CARs, and the overall negative estimates suggest there are value transfers from the U.S. acquirer shareholders to the foreign target investors. The effects are both statistically and economically meaningful for dictatorial U.S. bidders making acquisitions in high-category foreign countries over the 5-day event window. We now turn to multivariate regression analyses to seek a richer relation between acquirer corporate governance and cross-border target shareholder value.

2.4.5 Multivariate regressions for cross-border deals

To mirror the domestic deal tests, we re-run multivariate regressions for cross-border deals with additional control variables following the spirit of Bris and Cabolis (2008). This set of cross-border specific controls includes considerations for aforementioned sovereign corporate governance and macroeconomic variables. Specifically, for each cross-border acquisition, we have:[11]

$$\begin{aligned}\text{CAR} \;=\; & \textit{Full Acquire} \cdot \beta_1 \cdot \textit{G Index} + \beta_2 \cdot (\textit{G Index})^2 \\ & + \beta_3 \cdot \textit{Deal Characteristics} + \beta_4 \cdot \textit{Acquirer \& Target Characteristics} \\ & + \beta_5 \cdot \textit{M\&A Market Condition} + \beta_6 \cdot \textit{Macroeconomic Factors} \\ & + \textit{Full Acquire} \cdot \beta_7 \cdot \textit{LLSV Indices} + \textit{error term}.\end{aligned}$$

The key explanatory variable is the *G* index of the acquirer, though we continue to include relevant control variables as in the case of domestic deals.[12] Because cross-border deals are additionally affected by macroeconomic parameters and sovereign differences in the aggregate level of corporate governance, we augment these tests with the per-U.S.-dollar exchange rate (*Foreign Exchange*), the log-difference in per capita GDP between the U.S. and the target country (GDP), the ratio of stock market capitalization to GDP (SMCTG) for the target country from Djankov *et al.* (2008) representing the degree of financial market development, and the differences between La Porta *et al.* (1998)'s indices for the U.S. and the target country. The *Full Acquire* indicator equals one if the deal is a 100% acquisition, and zero otherwise. We interact this control with the *G* index and La Porta *et al.* (1998)'s measures since a complete transfer of ownership may generate a more material change in investor protection.

Results are presented in Table B.7. For both the 11- ([±5]) and 21-day ([±10]) horizons, we once again find that the *G* index is positively related to CAR. In a full acquisition, foreign target shareholders benefit the more the number of ATPs

[11] Bris and Cabolis (2008) additionally constructs a wholesomeness index that reflects antitrust laws and merger controls, citing the White & Case survey "Worldwide Antitrust Merger Notification Requirements." Similarly, Dyck and Zingales (2004) gathers data on statutes requiring additional shares purchases at certain thresholds.

[12] The significance of these controls in general M&A deals is discussed by Eckbo *et al.* (1990) and others.

held by their U.S. acquirers (*Model 1*) and the private contracts are valuable in addition to risk exposure to sovereign legal measures (AD and AS in *Model 2*). In other words, the dictator premium exists, economically and statistically significantly, in the hands of foreign investors and the effect is evident controlling for cross-border differences in the level of country-level protection of investors, and further macroeconomic factors, deal, acquirer and target characteristics and M&A market condition (*Model 3*). Only this full linear model, upon a full acquisition, conserves overall implications intact over the 5-day ([±2]) window. Asynchronous trading and foreign exchange translation effects may delay quick responses across the border, though this conjecture warrants a further investigation. Adding the quadratic *G* index does not alter the inferences from the linear model (*Model 4*), thus its suggests that the concave curvature of target returns against the *G* index shown previously is due to unexplained characteristics that are comprehensively controlled for herein. The grand results in Table B.7 confirm that not only country-level protection of investors but also firm-level commitment of a U.S. bidder matters in determining overseas target shareholder value. In sum, the regression results from domestic and cross-border deals agree with our central thesis: *reckless managers in the acquiring firms are beneficial to their target shareholders, either domestic or foreign, by means of overpayment as anticipated by the market participants.*

2.5 Conclusion

Our contribution to the market for corporate control literature is in two-fold: one in U.S. domestic deals and the other in cross-border deals. We find that the more antitakeover provisions (ATPs) a U.S. acquirer has in-place the higher premiums they pay to their domestic shareholders. We extend the domes- tic finding to the cross-border context and further corroborate that cross-border target shareholders benefit more from high-ATP U.S. acquirers (dictators) than from low-ATP peers (democrats).

These results are the first documentation of seemingly beneficial effects (dictator premiums) on targets—regardless of geographic locations—from ATP-ridden and, thus supposedly, over-spending bidders listed in the U.S. Lastly, quite surprisingly, the markets respond more delightedly on acquirers with more ATPs, upon publicizing cross-border merger deals. This finding is contrary to what Masulis *et al.* (2007) report in U.S. domestic deals, and it warrants a further scholarly attention. Our contribution to the literature is made by focusing on how target shareholders within and across the border are affected by managerial recklessness of U.S. acquirers proxied for by their ATPs.

This study attempts to address only a limited number of agenda in the literature. As noted throughout this paper, a contract-theoretic approach to unraveling the concave cross-border target returns deserves due attention. Thus far, we have only relied on an empirical ground by controlling for various candidate factors that are of import in the cross-border market for corporate control. A notable limitation of our research is that we do not account for the target firms' ATPs. This is to balance the experimental design settings of domestic versus foreign mergers, since ATP databases do not exist in the majority of target countries.

CHAPTER 3

AMERICAN DEPOSITARY RECEIPTS: ASIA-PACIFIC EVIDENCE ON CONVERGENCE AND DYNAMICS

3.1 Introduction

Over the last decades an increasing number of firms have chosen to list their shares on multiple exchanges in order to reduce capital costs and increase liquidity. According to the World Federation of Exchanges, the global market capitalization of cross-listed stocks increased by 16.3% in 2005, reaching the staggering sum of U.S.$5.76 trillion. Approximately 2,000 cross-listings[1] were recorded in the U.S. alone, where investments in American Depositary Receipts (ADRs) represented U.S.$657 billion which is 36% more than in 2004.

The popularity of cross-border listing and trading attracts much attention to this area. Two main issues are studied broadly. First, since an arbitrage opportunity arises by selling high and buying low when the price spread between a home-exchange share and its ADR widens sufficiently,[2] a myriad of academic work has attempted to gauge market efficiency by testing the parity of ADR-underlying prices. However, a number of practical limitations exist like asynchronous trading and short sale restrictions. Thus, some argue that even persistent disparity does not counter-example market efficiency.

The second issue is detecting where the pricing information and price dis-

[1] Including Levels I&II Depositary Receipts (DRs), Level I over-the-Counter (OTC) DRs, Rule 144a private placement DRs, ordinary shares, and Global Registered Shares (GRSs).

[2] In theory, a violation of the law of one price implies an arbitrage. In practice, the spread has to sufficiently exceed transaction costs.

covery are generated for a security traded in a multi-market setting.[3] The market that determines the security price is called the lead market, otherwise the lag market. Intuitively, lag market prices must converge to the lead market's price. If the market is efficient, the convergence to the parity must occur immediately. However, in reality, much evidence suggests that price deviations may exist temporarily or even persistently.[4] Consequently, people question how fast the convergence process is. Answering this question may help us better understand market efficiency.

There are several articles in the literature on the convergence and dynamics of ADR price spread. Koumkwa and Susmel (2005) use the Exponential Smooth Transition Autoregressive model (ESTAR) to investigate the narrowing divergence of ADR spreads. Gagnon and Karolyi (2004) set a linear time-series regression model and find that the convergence speed is positively related to financial market efficiency. Specifically, the convergence process is faster for countries with higher per-capita GDP, stronger investor protections, higher accounting standards, fewer short sale restrictions, and greater institutional ownership.

Following Gagnon and Karolyi (2004), we explore the convergence and dynamic structure of ADR price spreads. Instead of relying on conventional parametric approaches that carry embedded model-specification errors, we adopt a nonparametric method to estimate the convergence speed parameter. This is a first known attempt in the cross-listed shares literature. The rationale behind this is that the dynamic integration of markets can affect the convergence speed over time. This implies that any model assuming a constant convergence speed

[3] See Hasbrouck (1995).

[4] Gagnon and Karolyi (2004) record the existence of sizable price deviations for a significant portion of the 581 ADR-underlying pairs in their study. They report discounts of up to 90% and premia of up to 70%.

parameter may be misspecified. Introducing a nonparametric platform liberates us from such concerns.

In our model, we allow the convergence speed parameter to be a function of time. Since a prior on the functional form is not needed, our model is better suited for describing and studying the dynamic structure of the price spreads. In fact, our results indicate that the convergence speed explicitly evolves over time, across firms, and across countries. The time-explicit dynamics is a result of the idiosyncratic characteristics that are stratified in terms of cross-listing firms, home countries, home exchanges, industries, short sale availability etc.

We present the time-varying dynamics of both firm and country-specific convergence speed parameters. As we test our hypotheses, we empirically verify and visually corroborate the comparative dynamics of convergence with respect to home market efficiency and/or completeness, time-lag effect of home exchanges, short sales feasibility, and ADR-listing types. Our conclusion is that price deviations disappear faster as market efficiency measures improve over time.

In addition, we also examine various risk factors that ADR prices face. Bin *et al.* (2003) document that ADR returns are sensitive to the movements in *1.* the U.S. market; *2.* the underlying home equity market; and *3.* the corresponding foreign exchange market. Analogously, we put these factors into our dynamic structure model. The Beta coefficients on these factors can cast some light upon the sensitivity to the aforementioned risk factors.

Several reasons other than the rapid growth experienced by the Asia-Pacific economies over the last two decades have centered our analysis on this region's

ADRs. First, Asia-Pacific economies are, overall, still emerging markets. As such, they continue to be less integrated to global capital markets. There is a higher probability for emerging market ADRs to over- or under-react to information regarding underlying stocks, thus a higher chance that price corrections will ensue.

Second, in reality there are non-negligible limits to arbitrage between ADRs and their underlying shares in Asian stock markets. This can lead to time-resistant price disparities in some ADR-pairs. Moreover, the time-varying integration and development of the Asia-Pacific markets can further increase the time-varying convergence speed of ADR-pairs.

The remaining sections of this paper are organized as follows. Section 2 presents a literature review; and Section 3 a concise and yet intuitive description of a nonparametric remedy to the existing parametric method.[5] Section 4 describes the data.[6] The critical hypotheses, the associated testing procedures and the subsequent results are given in Sections 5, 6, and Appendix C.4. Section 7 ends with concluding remarks.

3.2 Literature

The existing literature exhibits a wide discussion of the arbitrage opportunities that arise from cross-listed shares. The early phase studies—Maldonado and Saunders (1983), Kato, Linn, and Schallheim (1991), Park and Tavokkol (1994), Miller and Morey (1996), and Karolyi and Stulz (1996)—conclude that no arbi-

[5] Appendix C.3 provides a supplemental explanation of our method.
[6] The tables in Appendix C.2 summarize the data.

trage opportunities exist for cross-listed shares and, thus, that they are priced in accordance to the parity implied by the no-arbitrage condition.

However, a thread of recent work has begun to document a significant divergence from the arbitrage price parity. Wahab, Lashgari, and Cohn (1992) were among the first to report evidence of arbitrage opportunities. Froot and Dabora (1999) study the pricing of a few dually-listed corporations (Royal Dutch and Shell, and Unilever N.V. and Unilever PLC) and find sizable and significant price deviations from the arbitrage parity. For a recent and complete summary of the literature see Gagnon and Karolyi (2004). The latter two authors find sizable price deviations for a significant portion of the 581 ADR-underlying pairs analyzed in their study. They report discounts of up to 90% and premiums of up to 70%.

Accepting the evidence that price deviations may exist temporarily, or even persistently, numerous papers have attempted to explain the phenomenon by adopting novel models that capture the dynamic structure of the spread. For example, Kim, Szakmary, and Mathur (2000) use Vector Autoregressive (VAR) and Seemingly Unrelated Regression (SUR) methods to analyze the adjustment of ADR-implied prices. Eun and Sabberwal (2003) employ an Error Correction Model (ECM) to explore the convergence of arbitrage parity prices for more than 60 Canadian stocks cross-listed on the New York Stock Exchange (NYSE). There are also many studies that try to explain the existence of a large price spread. Melvin (2003) and August *et al.* (2004) document that capital flow restrictions wield significant influence over arbitrage price parity with greater intensity during periods of economic and currency turmoil.

Considering the fact that the speed of convergence may differ depending on how far the prices are from the steady state, nonlinear models have been employed to describe the dynamic structure of price spreads. Surpassing the limitations of linear models, nonlinear formulations make more sense in capturing the time-varying property of price spreads. Early works such as Michael *et al.* (1997) and Taylor *et al.* (2001) apply Amplitude-Dependent Exponential Autoregressive (EXPAR) models to study the speed of adjustment of foreign exchange rates to the purchasing power parity (PPP). Nonlinear models have also been employed to examine the convergence speed of ADR price spreads. Rabinovitch *et al.* (2003) study an ECM model reflecting nonlinear dynamic adjustments for twenty Chilean and Argentine ADR-pairs.

Koumkwa and Susmel (2005) analyze a sample of twenty one Mexican shares dually listed in the U.S., where both cross-border markets lie within the same time zone. They employ the Exponential Smooth Transition Autoregressive (ESTAR) model with appropriate model selection criteria and find that price deviations are more or less short-lived. For 15 of the 21 pairs they studied, it took one trading day for the ADR-implied price spread to reduce by 50% (half-life). For some of the firms, half-lives of greater than four days were attributed to the low average daily trading volume. They report that their migration from the existing linear analytics to a nonlinear alternative reduces the mean half-life by down to circa 60%. Thus, the linear models are deemed to have exaggerated the arbitrage opportunities.

However, it is more art than science to determine the parametric specification of a nonlinear model. That is why we turn to nonparametrics, in which functional forms are not required *a priori*. See Fan and Yao's (2003) treatment of nonparametric time series models.

3.3 Model

As a starting point, we express the ADR price D_t as $D_t = \alpha_t + S_t \cdot E_t \cdot k + \eta_t = \alpha_t + \widetilde{D}_t + \eta_t$, where (1) S_t is the daily-closing price of the home exchange stock known at time t; *2.* E_t is the foreign exchange rate at time t; (3) k is the host-home exchange ratio; thus (4) $\widetilde{D}_t = S_t \cdot E_t \cdot k$ is the price adjusted for the exchange rate and its bundling ratio; (5) η_t is the stochastic term that follows a martingale so that $E[\eta|\mathcal{F}_t] = 0$, where $\mathcal{F}_t$ is the past information $\sigma-$algebra; and (6) α_t is the ADR premium (positive) or discount (negative) and equals zero if the market is efficient.[7]

Gagnon and Karolyi (2004) propose a time series model to see whether cross-border equity market shocks and foreign exchange market shocks can explain movements of the price spread. They incorporate first-order autoregression in the model to determine whether there is a mean reversion component. We extend their model by allowing the coefficient for the first lagged term to be a function of time. The rationale behind this is that the dynamic integration of markets can affect the convergence speed over time. We only consider the first-order lagged term of relative ADR premium under a Markov chain assumption controlling for the return on home/host equity and foreign exchange markets.

In the nonparametric spirit, for an ADR-pair i, our model can be described as

$$DR_i(t) = \alpha_i + \theta_i(t)\,\mathrm{DR}_i(t-1) + \sum_{j=-1}^{1} \beta_j^{US} R_M^{US}(t+j) + \sum_{j=-1}^{1} \beta_j^{H} R_M^{H}(t+j) + \sum_{j=-1}^{1} \beta_j^{FX} R_{FX}(t+j) + \varepsilon_i(t),$$

[7] α_t may reflect a momentum effect, thus it can be modeled in a time series context as a component of the price spread $D_t - \widetilde{D}_t$.

where (1) $DR_i(t) \equiv \frac{D_i(t)-\widetilde{D}_i(t)}{\widetilde{D}_i(t)}$ is the relative price spread, or relative ADR premium[8]; *2.* $\theta_i(t)$ is a coefficient function that measures the convergence speed of the price spread: the closer the absolute value of the parameter to one, the slower the convergence. As the control variables, *3.* R_M^{US} is the return on U.S. host market index; (4) R_M^H is the return on home market index; and (5) R_{FX} is the return on foreign exchange rate. We consider the distributed effect of three consecutive trading days around day t following Gagnon and Karolyi (2004).

Our model is a partial varying-coefficient model. Fan and Huang (2005) provide a local polynomial estimation for constant parameters, β's, and the coefficient function, $\theta_i(t)$. See Appendix C.3 for further details.

3.4 Data

We used the Bank of New York Mellon's ADR directory webpage[9] to locate and collect 400 pairs of ADRs and their underlying stocks for cross-listed firms based in the Asia-Pacific economies. Spanning January 2000 through December 2005, the daily closing prices of these stocks were sourced from Thompson Financial Datastream and The Center for Research in Security Prices (CRSP). In order to maximize the number of overlapping trading days, we chose the latter-half of the data—January 2003 through December 2005—and thus reduced the ADR-pairs down to 320. Tables C.1, C.2, and C.3 summarize the data used in this study.[10]

[8] A further refinement is the minimum relative spread which we actually use in the nonparametric estimation procedure. See Section 3 or Appendix C.3 for further discussion.

[9] URL: `http://www.adrbny.com/dr_directory.jsp`

[10] The hierarchy of the dataset is organized as follows (1) ADR*s Sheet* for ADR prices of Asia-Pacific firms; *2. Underlying Shares Sheet* for the corresponding underlying share prices; *3. Foreign Exchange Sheet* for home-U.S.$ exchange rates defined per U.S. dollar; (4) *Index Sheet* for U.S.

Table C.1 provides descriptive statistics of our sample, which covers 15 Asia-Pacific countries including large economies like China, Japan, India, and Korea. Among them, Japan, Hong Kong and Australia account for the majority of ADRs—up to 70% of the matched sample. We categorized ADRs according to the U.S. host exchanges and the associated stock indices. Additionally, we summarized ADR-pairs according to their industry, short sale availability, MSCI market category, and trading time difference. Table C.2 shows a profile of selected Asia-Pacific ADR-pairs and Table C.3 summarizes the characteristics of the regional stock exchanges.

As our key quantities of interest, *1. Shortsell* is defined as a dummy variable equal to one if short sales are allowed in a specific country and zero otherwise[11]; *2.* MSCI equals to one if a country belongs to the MSCI (Morgan Stanley Capital International) Developed Markets (DM) index, or zero if MSCI Emerging Markets (EM) index: it is used as a proxy for market efficiency or completeness of the home markets in the Asia-Pacific region; and *3. Time Difference* is the positive number of time difference between the closing time of the home exchange and the opening time of host exchanges in New York.[12]

Previous studies on ADR price spread such as Eun and Sabherwal (2003), Grammig *et al.* (2005), and Ding *et al.* (1999) use high frequency data to ensure the simultaneity of observed prices in the cross-border markets. There are overlapping trading hours among the host and home markets in their analyses,

and home market indices; and (5) *Master Sheet* for a fact sheet of all ADRs under our analysis which contains information regarding Datastream codes for ADR-pairs, ADR-listing types, home exchanges, home-market indices, U.S. market indices, countries, foreign exchange rates, time differences, short sale availability and MSCI indices.

[11] We obtained this data from Bris, Goetzmann, and Zhu (2003).

[12] A related *Timezone* variable was first introduced by Gagnon and Karolyi (2004): they obtained it from the World Federation of Exchanges (`http://www.world-exchanges.org`) albeit they use the number of time zones between the home and host exchanges.

and thus inferring from quote behavior is critical in a synchronous trade setting. However, as our data spans across Asia-Pacific markets, there are no overlapping trading hours with U.S. host exchanges. For this reason, we believe that the daily frequency of data is sufficient for the purpose of our analysis.

For a daily closing price $P_i^H(t)$ in the home market, we can define a price spread by either $P_i^{US}(t) - P_i^H(t)$ or $P_i^{US}(t-1) - P_i^H(t)$. Thus, the relative price spread can be defined in the two following ways: *1.* $DR_i(t) \equiv \left(P_i^{US}(t) - P_i^H(t)\right)/P_i^H(t)$, which is the same-calendar day causality from an Asia-Pacific market to a U.S. market; and (2) $DR_i(t)' \equiv \left(P_i^{US}(t-1) - P_i^H(t)\right)/P_i^H(t)$, which is the reversed causality from the U.S. exchange to the Asia-Pacific exchange.

Figure C.1 shows the nonparametric kernel density plots of the two relative price spreads for six randomly selected firms. In the plots, the dotted lines are the kernel probability density functions for $DR_i(t)$ and the solid lines are those of $DR_i(t)'$. The bold solid line is the minimum of the two relative spreads, which is defined as $\widetilde{DR}_i(t) \cong \min\{DR_i(t), DR_i(t)'\}$.[13]

From Figure C.1, we can see that the dotted kernel density plots usually exhibit higher kurtosis with narrower dispersion than the solid counterparts. Thus, we can intuitively infer that the underlying stock prices of Asia-Pacific markets lead ADR prices sequentially, rather than the latter predicts the former. This result confirms the conclusion of Hasbrouck (1995), Lieberman et al. (1999), and Su and Chong (2007).[14] The minimum spread, by definition, reflects the nar-

[13]See Appendix C.3 for a detailed explanation on the minimum relative spread.

[14]Lieberman et al. (1999) examine the price behavior of six firms cross-listed on Israel and U.S. exchanges with daily closing prices. They demonstrate that the effect of Israeli market on the share prices in the U.S. is stronger than the it in the reversed direction. Su and Chong (2007) study the contribution to price discovery for Chinese cross-listed stocks, which is a subset of our sample. They find that the stock prices of two exchanges, New York Stock Exchange (NYSE) and the Hong Kong Stocks Exchange (HKSE) are co-integrated and mutually-adjusting, but the

rowest and yet "pickiest" distribution. Since we study the mutual convergence between pair-wise markets, we use the minimum relative spread, $\widetilde{DR}_i(t)$, which captures symmetric convergence.

3.5 Results

3.5.1 Convergence speed parameters

Using a nonparametric kernel estimation,[15] we obtained the coefficient function $\theta_i(t)$ for each ADR-pair in the sample. Figure C.2 shows the results for six randomly selected pairs. We find that $\theta_i(t)$s change over time within the sample period—which further confirms the justification of our nonparametric setting with time-varying convergence speed parameters. A downward-sloping $\theta_i(t)$ indicates that the convergence accelerates as market efficiency and/or information quality improves. There are some pairs with upward-sloping $\theta_i(t)$s and they show that the associated conditions for market efficiency worsened throughout the period.

What is more intriguing is the result of the market and country-wise average convergence speed. Figure C.3 exhibits the country-level dynamics of ADRs convergence parameters. Most countries—Australia, China, Hong Kong etc.—exhibit decreasing $\theta_i(t)$s where market efficiency gained positive momentum. It appears that, nonetheless, some Southeast Asian countries—Malaysia, Indonesia, and Thailand—show non-decreasing $\theta_i(t)$s. This suggests that over

HKSE contributes about 80% to price discovery, more than the NYSE does.

[15] See Appendix C.3.

the sample period their markets were slow to integrating to the U.S. equity markets. It can be attributed to the shift in economic regimes following the Asian Financial Crisis of 1997-1998. Regional governments implemented market regulatory policies, such as capital flow restrictions, short sale restrictions, and foreign ownership ceilings, that undermined undermine market efficiency.

Figure C.4 (placed after Figure C.5 due to space allocation) shows the overlapping comparative dynamics chart of convergence speed for several selected countries. As the chart indicates, Pakistan is dominated by Thailand, which is in turn dominated by Korea. Australia's dominance over Korea is valid from December 2003 to July 2005. The results are not counter-intuitive. We also compare the convergence speed parameters by home exchanges in Figure C.5.

Figures C.6 and C.7 delineate the convergence speed of different ADR-listing types or exchanges. The 144A private placements are the slowest in convergence, if ever, and they are dominated by the Level I "pink sheet" OTC (over-the-counter) ADRs, then successively by the Level II/III ADRs listed on the NASDAQ and the NYSE as expected—given the latter being relatively superior to the former in terms of information quality and trading volume. The parameters of time-varying convergence speed, categorized by the availability of short sales, are presented in Figure C.8. It is apparent that, the convergence speed of ADR-pairs which are associated with home exchanges/countries where short sales are allowed and practiced significantly dominate those which are not.

The visual evidence presented hitherto may be insufficient draw a final conclusion; thus, we corroborate our argument with statistical significance tests based on the key quantities of interest. We hypothesize that:

Hypothesis I. The lack of short sales in some exchanges decelerates the convergence speed.[16]

Hypothesis II. The ADRs *of emerging or less developed economies offer more arbitrage opportunities*[17] *than those of developed economies based on* MSCI *categories.*

Hypothesis III. The higher the time difference between the home exchange and the host U.S. exchange is, the faster the convergence speed.[18]

We use t-test and Wilcoxon signed-rank test.[19] We can formalize the first hypothesis as $H_0 : \theta_{\text{no short}}(t) = \theta_{\text{short}}(t) \quad vs. \quad H_1 : \theta_{\text{no short}}(t) > \theta_{\text{short}}(t)$. Define the difference between nonparametric parameter estimates as $d(t) \equiv \widehat{\theta}_{\text{no short}}(t) - \widehat{\theta}_{\text{short}}(t)$. The t-test statistic is $t_0 \equiv \frac{\overline{d(t)} - 0}{\text{s.e.}(d(t))} \cong \frac{0.1570 - 0}{0.0080} = 19.7040$. Moreover, the Wilcoxon-test statistic[20] is of 1325 with a 2.813×10^{-10} *p-value*. Both the t-test and the Wilcoxon test significantly confirm that the availability of short sales can speed up convergence.

We now test the second hypothesis as $H_0 : \theta_{\text{EM}}(t) = \theta_{\text{DM}}(t) \quad vs. \quad H_1 : \theta_{\text{EM}}(t) > \theta_{\text{DM}}(t)$. Define the difference between nonparametric parameter estimates as $d(t) \equiv \widehat{\theta}_{\text{EM}}(t) - \widehat{\theta}_{\text{DM}}(t)$. The t-test result is $t_0 \equiv \frac{\overline{d(t)} - 0}{\text{s.e.}(d(t))} \cong \frac{0.2085 - 0}{0.0070} = 29.6419$, and yet another noticeably significant Wilcoxon-test statistic is obtained as $V_0 = 1326$ with a $p\text{-}value = 2.651 \times 10^{-10}$.

[16]That is, increase the convergence parameter.

[17]That is, the absolute value of convergence parameters will be higher and closer to one.

[18]In other words, the earlier the home market closes, the quicker the ADR finds price discovery during the trading hours of host exchanges in New York City. This hypothesis is based on the belief that as information chronologically accumulates for an underlying share, its ADR must reflect more precisely the implied-value, yielding a fewer arbitrage opportunities by the closing time of the host market in N.Y.C.

[19]It is a nonparametric test to compare two samples. See Wilcoxon (1945) and Appendix C.4.

[20]The *Wilcoxon* test-statistic is defined as $V_0 \equiv \sum_{\{t\}} \mathbf{1}_{\{d(t)>0\}} \cdot \rho_t$, where ρ_t is the rank of $\{|d(t)|\}_{\{t\}}$.

Both tests show that the developed markets have a higher speed of convergence. The time-varying parameters of convergence speed, categorized by MSCI Developed (DM) and Emerging Market (EM), resemble those classified according to the availability of short sales. (See Figure C.9.) This is because most constituent countries of the MSCI DM index allow short sale. This supports empirically and graphically our conviction that the countries with enhanced market efficiency or completeness , as proxied by MSCI indices, exhibit lower convergence parameters, which in turn yield quicker convergence to the parity.

Next, we want to see the effect of trading time differences. Figure C.10 summarizes the time-difference effect on convergence speed. The trading-time difference between the U.S. and home markets affects the ability of arbitrageurs forcing convergence between an ADR and the underlying share prices. When trading sessions overlap, cross-listed pairs are priced almost synchronously making it easier for investors to implement pairs-trade effectively. Overall, the earlier the home exchange closes prior to the ADR market in New York City, the faster the convergence speed becomes—measured within the trading hours of ADR exchanges. We test the third hypothesis as $H_0 : \theta_{\text{later}}(t) = \theta_{\text{earlier}}(t) \;\; vs. \;\; H_1 : \theta_{\text{later}}(t) > \theta_{\text{earlier}}(t)$. Define the difference between nonparametric parameter estimates as $d(t) \equiv \widehat{\theta}_{\text{later}}(t) - \widehat{\theta}_{\text{earlier}}(t)$.

Table C.4 summarizes the cross-test statistics of the parametric differences between home markets of later and earlier closing times prior to the opening of N.Y.C. host markets. The home exchanges in the horizontal time intervals close earlier than the home exchanges in the vertical slots. All t-test statistics are significant at a 1% right-side significance level denoted by two asterisks. Again, the exceedingly significant Wilcoxon-test statistics are presented with their corresponding *p-value*s.

3.5.2 Betas

Bin *et al.* (2003) document that ADR returns are sensitive to fluctuations in *1.* the U.S. market, *2.* the underlying home equity market, and *3.* the corresponding foreign exchange market. Let us invoke our model specification again.

$$\widetilde{DR}_i(t) = \alpha_i + \theta_i(t)\,\widetilde{DR}_i(t-1) + \sum_{j=-1}^{1} \beta_j^{US} R_M^{US}(t+j) + \sum_{j=-1}^{1} \beta_j^{H} R_M^{H}(t+j) + \sum_{j=-1}^{1} \beta_j^{FX} R_{FX}(t+j) + \varepsilon_i(t).$$

As the plots in Figure C.11 show, most countries exhibit positive and negative risk exposure toward home and U.S. indices respectively. These results are in line with those of Gagnon and Karolyi (2004).[21]

As expected, the foreign exchange risk exposure of the ADR spread, β^{FX}, depends on exchange rate regimes. For the countries whose foreign exchange beta paths are nearly flat about zero, three out of four—China (pegged), Hong Kong (pegged), Singapore (managed float) and Taiwan (fully float)—maintain either pegged or managed-float exchange rate regime as shown in Table C.3.

For the fully-float regimes, the signs of the foreign exchange risk exposure are mixed. Australia and Japan exhibit pronounced positive effects from exchange rates exceeding the effects from home and host exchange indices. A positive beta of foreign exchange exposure (β^{FX}) implies that a positive return of the exchange rate (R_{FX})—depreciation in the home currency against the U.S. dollar—poses an upside for the relative ADR spread.[22]

[21]This could provide us ground to applying the difference between R_M^{US} and R_M^{H} in the model specification rather than separating them.

[22]$DR_i(t) \equiv \frac{P_i^{US}(t) - P_i^{H}(t)}{P_i^{H}(t)}$.

Should the market expect that the currency continues to depreciate such that the positive sign momentum will persist, the relative spread will carry a premium. The ADR will then out-value the price implied by the underlying share. This excess spread—cross-listing premium—is rational since the market reflects the view that the U.S. dollar will appreciate against home currency.

Yet, India shows a negative exchange rate beta even though it is a fully-float currency regime. This is an interesting phenomenon that may be related to currency strengthening. Investors expect the Indian rupee to appreciate, thus moderate exchange rate perturbations do not significantly and negatively affect on the spreads. They are *dominantly* considered temporary jitters that are expected to revert to the trend quickly.

Thus, a depreciation will be considered short-lived, to be followed by a quick trend-reversion. That is why there is a discount component in Indian ADRs from a currency risk perspective. The Chinese renminbi (RMB) may have behaved similarly as a member of the *Chindia Leagues*, had China not pegged its currency against the Greenback.

3.6 Conclusion

In this paper, we explored the convergence of the prices of American Depositary Receipts (ADRs) listed by Asia-Pacific firms and their original shares listed on home exchanges in various time zones. Instead of conventional parametric approaches that carry embedded model-specification errors, we adopted a nonparametric technique to estimate the convergence speed parameter of the existing parametric specification.

We presented the dynamics of both firm and country-specific time-varying convergence speed parameters. We empirically verified and visually corroborated the comparative dynamics of the convergence with respect to the home market efficiency/completeness, time-lag of home exchanges, availability of short sales on the home exchanges, and ADR-listing types. Given these circumstances and due to practical constraints, the limits of arbitrage cannot serve as a counterexample of market efficiency. We conclude that, as an alternative answer, the speed of convergence accelerates as proposed market efficiency measures are enhanced.

APPENDIX A

CHAPTER 1 OF APPENDIX

A.1 Tables

Table A.1: Sample of Canadian firms listed on both the TSX and the NYSE

The PIN is the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996). The bid-ask spreads are defined: *1.* $\text{SPREAD}_{\text{NYSE}} \equiv \frac{\text{ask}_{\text{NYSE}} - \text{bid}_{\text{NYSE}}}{(\text{ask}_{\text{NYSE}} + \text{bid}_{\text{NYSE}})/2}$; and 2. $\text{SPREAD}_{\text{TSX}} \equiv \frac{\text{ask}_{\text{TSX}} - \text{bid}_{\text{TSX}}}{(\text{ask}_{\text{TSX}} + \text{bid}_{\text{TSX}})/2}$. The information share (IS) is exchange-specific relative contribution to price discovery of a security traded on multiple exchanges, following Hasbrouck (1995, 2007). All values are arithmetic means of monthly estimates through the sample period: January 1, 1998, through December 31, 2000.

Company	*PIN NYSE*	*Spread NYSE*	*IS NYSE*	*PIN TSX*	*Spread TSX*	*IS TSX*
Abitibi-Consolidated, Inc.	0.151	0.018	41.7%	0.184	0.005	58.3%
Advantage Energy Income Fund	0.372	0.117	50.0%	0.482	0.131	50.0%
Agnico-Eagle Mines Limited	0.188	0.026	50.0%	0.421	0.080	50.0%
Agrium Inc.	0.190	0.020	43.0%	0.202	0.007	57.0%
Alcan Inc.	0.147	0.006	40.7%	0.169	0.003	59.3%
Bank of Nova Scotia	0.234	0.063	49.9%	0.188	0.003	50.1%
Barrick Gold Corporation	0.190	0.008	38.6%	0.215	0.003	61.4%
BCE Inc.	0.112	0.006	49.1%	0.174	0.002	50.9%
Biovail Corporation	0.181	0.008	49.5%	0.220	0.006	50.5%
BMO Financial Group	0.160	0.007	41.4%	0.204	0.002	58.6%
Brookfield Properties Corporation	0.267	0.020	45.3%	0.226	0.016	54.7%
Cameco Corporation	0.223	0.020	38.0%	0.197	0.009	62.0%
Canadian Imperial Bank of Commerce	0.308	0.017	49.9%	0.160	0.002	50.1%
Canadian National Railway Company	0.139	0.007	48.4%	0.215	0.003	51.6%
Canadian Pacific Railway Limited	0.206	0.007	43.8%	0.173	0.003	56.2%
Canwest Global Communications	0.287	0.023	47.3%	0.295	0.016	52.7%
Celestica Inc.	0.186	0.010	44.4%	0.225	0.005	55.6%
CGI Group Inc.	0.195	0.028	49.9%	0.280	0.018	50.1%
Compton Petroleum Corporation	0.110	0.010	50.0%	0.253	0.023	50.0%
Corus Entertainment, Inc.	0.311	0.016	46.0%	0.210	0.012	54.0%

Table A.1 (Continued)

Company	PIN_{NYSE}	$Spread_{NYSE}$	IS_{NYSE}	PIN_{TSX}	$Spread_{TSX}$	IS_{TSX}
Cott Corporation	0.147	0.012	50.0%	0.223	0.014	50.0%
Domtar Corporation	0.199	0.010	50.0%	0.206	0.007	50.0%
Encana Corporation	0.311	0.054	39.3%	0.291	0.019	60.7%
Energy Metals Corporation	0.203	0.059	50.0%	0.274	0.047	50.0%
Enerplus Resources Fund	0.261	0.020	46.2%	0.286	0.019	53.8%
Fairfax Financial Holdings Limited	0.308	0.012	50.0%	0.254	0.007	50.0%
Four Seasons Hotels Inc.	0.202	0.009	46.5%	0.214	0.009	53.5%
Gildan Activewear Inc.	0.239	0.019	83.6%	0.800	0.018	16.4%
Goldcorp Inc.	0.354	0.072	41.4%	0.178	0.011	58.6%
Intertape Polymer Group Inc.	0.246	0.020	40.8%	0.277	0.014	59.2%
IPSCO Inc.	0.301	0.027	48.7%	0.215	0.010	51.3%
Kinross Gold Corporation	0.247	0.059	44.8%	0.231	0.012	55.2%
Magna International Inc.	0.153	0.006	42.6%	0.179	0.004	57.4%
Manulife Financial Corp.	0.223	0.011	41.4%	0.222	0.031	58.6%
MDS Inc.	0.218	0.024	33.8%	0.323	0.038	66.2%
Meridian Gold Inc.	0.205	0.042	42.7%	0.267	0.019	57.3%
Nexen, Inc.	0.168	0.014	44.9%	0.160	0.004	55.1%
Nortel Networks Corporation	0.205	0.006	47.9%	0.188	0.002	52.1%
NOVA Chemicals Corporation	0.245	0.015	38.8%	0.275	0.006	61.2%
Pengrowth Energy Trust	0.247	0.025	49.2%	0.183	0.007	50.8%
Petro-Canada	0.238	0.017	42.5%	0.196	0.004	57.5%
Potash Corporation of Saskatchewan Inc.	0.128	0.007	44.2%	0.170	0.005	55.8%
Precision Drilling Trust	0.167	0.011	36.5%	0.190	0.005	63.5%
Quebecor World, Inc.	0.230	0.013	45.4%	0.183	0.004	54.6%
RBC Financial Group	0.163	0.007	46.3%	0.173	0.002	53.7%
Rogers Communications Inc.	0.179	0.017	37.4%	0.238	0.006	62.6%
Shaw Communications Inc.	0.195	0.012	49.2%	0.187	0.007	50.8%
Stantec Inc.	0.158	0.010	50.0%	0.394	0.020	50.0%
Suncor Energy Inc.	0.185	0.010	47.4%	0.184	0.004	52.6%
Talisman Energy Inc.	0.190	0.013	39.4%	0.164	0.005	60.6%
TELUS Corporation	0.199	0.014	43.0%	0.228	0.005	57.0%
The Thomson Corporation	0.290	0.034	49.6%	0.175	0.005	50.4%
Tim Hortons Inc.	0.202	0.017	50.0%	0.536	0.124	50.0%
Toronto-Dominion Bank	0.152	0.010	26.0%	0.203	0.002	74.0%
TransAlta Corporation	0.308	0.081	49.9%	0.180	0.005	50.1%
TransCanada Corporation	0.157	0.012	48.6%	0.211	0.004	51.4%
Mean	0.214	0.022	45.6%	0.242	0.015	54.4%
Median	0.202	0.015	46.1%	0.213	0.007	53.9%
Standard Deviation	0.060	0.022	7.2%	0.107	0.025	7.2%

Table A.2: Relative premiums of cross-listings on the NYSE

For a TSX-NYSE cross-listed pair, the relative premium ($\equiv (p_{\text{NYSE}} - p_{\text{TSX}})/p_{\text{TSX}}$) is the percentage premium earned on the NYSE-listed stock against the original listing traded on the TSX, adjusted for the U.S.-Canada exchange rate. The summary statistics in Panel A are based on ten-minute frequency prices 56 cross-listed pairs through the sample period: January 1, 1998, through December 31, 2000. The PIN is the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996). ($\text{PIN}_{\text{TSX}} - \text{PIN}_{\text{NYSE}}$) is the difference in the monthly PINs on the TSX- and the NYSE-listed pairs, respectively. The observations in Panel B are in firm-months through the sample period: January 1, 1998, through December 31, 2000. Standard errors are corrected for heteroskedasticity and autocorrelation, following Newey and West (1987, 1994). The numerical value in the parentheses below the estimate is a *t*-statistic. ***, **, and * stand for statistical significance based on two-sided student-*t* tests at the 1%, 5%, and 10% level, respectively.

Panel A: Summary statistics of relative premiums

	Mean	Median	Standard Deviation
$(p_{\text{NYSE}} - p_{\text{TSX}})/p_{\text{TSX}}$	0.00306	0.00004	0.03031

Panel B: Relative premiums against cross-border difference in the PIN

$$(p_{\text{NYSE}} - p_{\text{TSX}})/p_{\text{TSX}} = \beta(\text{PIN}_{\text{TSX}} - \text{PIN}_{\text{NYSE}}) + \epsilon$$

	Estimate	*No. of Obs.*	*Adj.* R^2
($\text{PIN}_{\text{TSX}} - \text{PIN}_{\text{NYSE}}$)	1.087***	1,591	0.176
	(3.259)		

Table A.3: Wilcoxon signed-rank tests

The PIN is the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996). The bid-ask spreads are defined as: *1.* $\text{SPREAD}_{\text{NYSE}} \equiv \frac{\text{ask}_{\text{NYSE}} - \text{bid}_{\text{NYSE}}}{(\text{ask}_{\text{NYSE}} + \text{bid}_{\text{NYSE}})/2}$; and *2.* $\text{SPREAD}_{\text{TSX}} \equiv \frac{\text{ask}_{\text{TSX}} - \text{bid}_{\text{TSX}}}{(\text{ask}_{\text{TSX}} + \text{bid}_{\text{TSX}})/2}$. The information share is exchange-specific relative contribution to price discovery of a security traded on multiple exchanges, following Hasbrouck (1995, 2007). The Wilcoxon signed-rank test is a non-parametric pair-wise comparison test, following Wilcoxon (1945). The coordinates (i, t) denote each firm and each month, respectively. d is a differential measure defined for the estimates of each quantity of interest. They are defined as: *1.* $d(i,t) \equiv \text{PIN}_{\text{TSX}}(i,t) - \text{PIN}_{\text{NYSE}}(i,t)$; *2.* $d(i,t) \equiv \text{SPREAD}_{\text{NYSE}}(i,t) - \text{SPREAD}_{\text{TSX}}(i,t)$; and *3.* $d(i,t) \equiv \text{IS}_{\text{TSX}}(i,t) - \text{IS}_{\text{NYSE}}(i,t)$. The Wilcoxon test-statistic is defined as: $V_0 \equiv \sum_{\{(i,t)\}} \mathbf{1}_{\{d(i,t)>0\}} \cdot \rho_{it}$, where ρ_{it} is the rank of $\{|d(i,t)|\}$.

	PIN	Spread	Information Share
H_0	$\text{PIN}_{\text{TSX}} = \text{PIN}_{\text{NYSE}}$	$\text{SPREAD}_{\text{TSX}} = \text{SPREAD}_{\text{NYSE}}$	$\text{IS}_{\text{TSX}} = \text{IS}_{\text{NYSE}}$
H_1	$\text{PIN}_{\text{TSX}} > \text{PIN}_{\text{NYSE}}$	$\text{SPREAD}_{\text{NYSE}} > \text{SPREAD}_{\text{TSX}}$	$\text{IS}_{\text{TSX}} > \text{IS}_{\text{NYSE}}$
d	$\text{PIN}_{\text{TSX}}(i,t) - \text{PIN}_{\text{NYSE}}(i,t)$	$\text{SPREAD}_{\text{NYSE}}(i,t) - \text{SPREAD}_{\text{TSX}}(i,t)$	$\text{IS}_{\text{TSX}}(i,t) - \text{IS}_{\text{NYSE}}(i,t)$
V_0	424250	680698	2926092
p-value	0.001458	$< 2.2 \times 10^{-16}$	$< 2.2 \times 10^{-16}$

Table A.4: Fixed-effect panel regressions

The panel dataset is constructed with columns of company symbol, monthly date, TSX indicator, and monthly estimates of the PIN, spread, information share, and volume, following Dempster, Laird, and Rubin (1977), and van Dyk and Meng (2001). On the TSX and the NYSE, for each cross-lister (i) and in each month (t), January 1998 through December 2000, *1.* PIN is the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996); *2.* SPREAD is the relative quoted spread; *3.* the information share (IS) is exchange-specific relative contribution to price discovery of a security traded on multiple exchanges, following Hasbrouck (1995, 2007); *4.* VOLUME is the log of total daily trading volume; and *5.* TSX equals one if the estimated numerical value is of the TSX, or zero if the NYSE. Standard errors are corrected for heteroskedasticity and autocorrelation, following Newey and West (1987, 1994). The numerical values in the parentheses below the estimates are t-statistics. ***, **, and * stand for statistical significance based on two-sided student-t tests at the 1%, 5%, and 10% level, respectively. The observations are in firm-months.

Panel A: $\text{PIN} = \beta_0 + \beta_1 \text{SPREAD} + \beta_2 \text{IS} + \beta_3 \text{VOLUME} + \beta_4 \text{TSX} + \epsilon$

	Model 1	*Model 2*	*Model 3*
(Intercept)	0.161^{***}		0.110^{***}
	(10.120)		(3.301)
SPREAD	1.938^{***}	2.043^{***}	2.003^{***}
	(22.908)	(23.632)	(20.848)
IS	0.043^{***}	0.017^{*}	
	(5.550)	(1.655)	
VOLUME	−0.000	0.021^{***}	0.003
	(−0.086)	(33.774)	(1.234)
TSX Dummy		0.026^{***}	0.083^{***}
		(6.525)	(4.124)
Company & Month Effects	No	No	Yes
No. of Obs.	3,960	3,960	3,960
Adj. R^2	0.118	0.855	0.184

Table A.4 (Continued)

Panel B: $\text{SPREAD} = \beta_0 + \beta_1 \text{PIN} + \beta_2 \text{IS} + \beta_3 \text{VOLUME} + \beta_4 \text{TSX} + \epsilon$

	Model 1	*Model 2*	*Model 3*
(Intercept)	−0.008***		0.022***
	(−2.723)		(4.219)
PIN	0.064***	0.061***	0.050***
	(22.908)	(23.632)	(20.848)
IS	−0.016***	−0.002	
	(−12.191)	(−1.231)	
VOLUME	0.003***	0.001***	0.002***
	(7.229)	(10.019)	(5.314)
TSX Dummy		−0.008***	−0.013***
		(−12.374)	(−3.983)
Company & Month Effects	No	No	Yes
No. of Obs.	3,960	3,960	3,960
Adj. R^2	0.155	0.582	0.369

Panel C: $\text{IS} = \beta_0 + \beta_1 \text{PIN} + \beta_2 \text{SPREAD} + \beta_3 \text{VOLUME} + \beta_4 \text{TSX} + \epsilon$

	Model 1	*Model 2*	*Model 3*
(Intercept)	0.484***		0.461***
	(15.162)		(9.696)
PIN	0.179***	0.098*	
	(5.550)	(1.655)	
SPREAD	−2.196***	−0.175	0.038
	(−12.191)	(−1.231)	(0.280)
VOLUME	0.002	0.049***	−0.026***
	(0.651)	(62.986)	(−7.150)
TSX Dummy		0.260***	0.350***
		(57.344)	(12.167)
Company & Month Effects	No	No	Yes
No. of Obs.	3,960	3,960	3,960
Adj. R^2	0.036	0.919	0.561

Table A.5: Cross-sectional regressions

For each cross-listed pair (i), SPEED$_{\text{CONV}}$ ($\equiv \theta_i$) measures the reciprocal speed of the parity-convergence of relative premium, following Gagnon and Karolyi's (2009a) empirical model:

$$\text{DR}_i(t) = \alpha_i + \theta_i \, \text{DR}_i(t-1) + \sum_{j=-1}^{1} \beta_j^{US} R_M^{US}(t+j) + \sum_{j=-1}^{1} \beta_j^{C} R_M^{C}(t+j) + \sum_{j=-1}^{1} \beta_j^{FX} R_{FX}(t+j) + \varepsilon_i(t).$$

The daily relative premium $\left(\text{DR}_i(t) \equiv \left(P_i^{US}(t) - P_i^{C}(t)\right)/P_i^{C}(t)\right)$ can be explained by *1.* its own lag ($\text{DR}_i(t-1)$) associated with *2.* the *convergence speed parameter* (θ_i): the closer the absolute value to zero, the faster the convergence to parity; and lag-distributed (yesterday ($j = -1$), today ($j = 0$), and tomorrow ($j = +1$)) returns on *3.* the S&P 500 Index $\left(R_M^{US}(t+j)\right)$, *4.* the S&P TSX Composite Index $\left(R_M^{C}(t+j)\right)$, and *5.* the Canada-U.S. exchange rate return ($R_{FX}(t+j)$), a positive R_{FX} implies a depreciation in the Canadian dollar. The forward-lag is due to information leakages and market impact. The remaining variables are: *1.* PIN$_{\text{AVG}}$ is the arithmetic average of the PINs of the pair on the TSX and the NYSE; *2.* SPREAD$_{\text{AVG}}$ is the arithmetic average of the bid-ask spreads of the pair on the TSX and the NYSE; *3.* SIZE is the proxy of normalized firm size and defined as the average log market capitalization on the TSX and the NYSE; *4.* INDUSTRY equals one if the cross-lister is a manufacturing firm, or zero otherwise; *5.* VOLUME is the log of total daily trading volume; *6.* GOVERNANCE is the Report on Business governance index of Canadian firms published by *Globe and Mail* (McFarland (2002)); *7.* $(p_{\text{NYSE}} - p_{\text{TSX}})/p_{\text{TSX}}$ is the relative premium on the NYSE-listed stock; and *8.* $(\text{PIN}_{\text{TSX}} - \text{PIN}_{\text{NYSE}})$ is the difference of the PINs of the pair on the TSX and the NYSE. Standard errors are corrected for heteroskedasticity and autocorrelation, following Newey and West (1987, 1994). The numerical values in the parentheses below the estimates are t-statistics. ***, **, and * stand for statistical significance based on two-sided student-t tests at the 1%, 5%, and 10% level, respectively. The observations are in firm-months.

Table A.5 (Continued)

Panel A: Cross-sectional determinants of the convergence speed parameter of cross-listed pairs

$$\text{SPEED}_{\text{CONV}} = \gamma_1\,\text{PIN}_{\text{AVG}} + \gamma_2\,\text{SPREAD}_{\text{AVG}} + \gamma_3\,\text{SIZE} + \gamma_4\,\text{INDUSTRY} + \gamma_5\,\text{VOLUME} + \gamma_6\,\text{GOVERNANCE} + \eta$$

	Model 1	*Model 2*	*Model 3*	*Model 4*
PIN	1.281***	0.919***	1.138***	1.147***
	(4.845)	(3.604)	(5.162)	(3.060)
SPREAD	4.606*	3.034		1.828
	(1.821)	(1.487)		(0.466)
SIZE	0.021	−0.006	−0.074	0.073
	(0.207)	(−0.049)	(−0.537)	(0.706)
INDUSTRY	−0.165***	−0.205***	−0.227***	−0.195***
	(−3.436)	(−2.952)	(−3.203)	(−3.254)
VOLUME		0.406**	0.483**	0.324
		(2.134)	(2.568)	(1.200)
GOVERNANCE				−0.001
				(−0.445)
No. of Obs.	1,591	1,591	1,591	1,591
Adj. R^2	0.606	0.635	0.629	0.557

Panel B: Cross-sectional determinants of cross-border average spread

$$\text{SPREAD}_{\text{AVG}} = \delta_1\,\text{PIN}_{\text{AVG}} + \delta_2\,\text{SPEED}_{\text{CONV}} + \delta_3\,\text{SIZE} + \delta_4\,\text{INDUSTRY} + \epsilon$$

	PIN_{AVG}	$\text{SPEED}_{\text{CONV}}$	SIZE	INDUSTRY
Estimate	0.074**	0.018**	−0.021*	−0.002
	(2.359)	(2.548)	(−1.876)	(−0.289)
No. of Obs.	1,591			
Adj. R^2	0.625			

Panel C: Cross-sectional determinants of relative premiums

$$(p_{\text{NYSE}} - p_{\text{TSX}})\,/\,p_{\text{TSX}} = \beta_1\,(\text{PIN}_{\text{TSX}} - \text{PIN}_{\text{NYSE}}) + \beta_2\,\text{SPEED}_{\text{CONV}} + \beta_3\,\text{GOVERNANCE} + \epsilon$$

	$\text{PIN}_{\text{TSX}} - \text{PIN}_{\text{NYSE}}$	$\text{SPEED}_{\text{CONV}}$	GOVERNANCE
Estimate	0.849***	0.018**	−0.0003
	(4.125)	(2.669)	(−0.980)
No. of Obs.	1,591		
Adj. R^2	0.380		

Table A.6: Test of exchange-specific liquidity skewness

$\widetilde{\text{SPREAD}}_{\text{NT}}$ is the percentage cross-border arbitrage profit from buying on the TSX and selling on the NYSE, and $\widetilde{\text{SPREAD}}_{\text{TN}}$ is from buying on the NYSE and selloing on the TSX. They are defined as:

- $\widetilde{\text{SPREAD}}_{\text{NT}} \equiv \{\text{ask}_{\text{NYSE}} - \text{bid}_{\text{TSX}} \cdot (\text{US\$/CAN\$})_{\text{ask}}\} / \{\text{bid}_{\text{TSX}} \cdot (\text{US\$/CAN\$})_{\text{ask}}\}$,
- $\widetilde{\text{SPREAD}}_{\text{TN}} \equiv \{\text{ask}_{\text{TSX}} \cdot (\text{US\$/CAN\$})_{\text{bid}} - \text{bid}_{\text{NYSE}}\} / \text{bid}_{\text{NYSE}}$.

The Wilcoxon signed-rank test is a non-parametric pair-wise comparison test, following Wilcoxon (1945). The coordinates (i, t) denote each firm and each month, respectively. d is a differential measure defined as: $d(i, t) \equiv \text{SPREAD}_{\text{TN}}(i, t) - \text{SPREAD}_{\text{NT}}(i, t)$. The Wilcoxon test-statistic is defined as: $V_0 \equiv \sum_{\{(i,t)\}} \mathbf{1}_{\{d(i,t)>0\}} \cdot \rho_{it}$, where ρ_{it} is the rank of $\{|d(i, t)|\}$.

H_0	$\widetilde{\text{SPREAD}}_{\text{NT}} = \widetilde{\text{SPREAD}}_{\text{TN}}$
H_1	$\widetilde{\text{SPREAD}}_{\text{NT}} \neq \widetilde{\text{SPREAD}}_{\text{TN}}$
d	$\widetilde{\text{SPREAD}}_{\text{NT}}(i, t) - \widetilde{\text{SPREAD}}_{\text{TN}}(i, t)$
V_0	507568
p-value	0.9407

Table A.7: Cross-listings on the NYSE by TSX-listed firms, 1998 through 2000

The PIN is the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996). The bid-ask spreads are defined as: *1.* $\text{SPREAD}_{\text{NYSE}} \equiv \frac{\text{ask}_{\text{NYSE}} - \text{bid}_{\text{NYSE}}}{(\text{ask}_{\text{NYSE}} + \text{bid}_{\text{NYSE}})/2}$; and *2.* $\text{SPREAD}_{\text{TSX}} \equiv \frac{\text{ask}_{\text{TSX}} - \text{bid}_{\text{TSX}}}{(\text{ask}_{\text{TSX}} + \text{bid}_{\text{TSX}})/2}$. When estimating the cumulative abnormal return (CAR) around a cross-listing on the NYSE, *1.* the market model uses the S&P TSX Composite Index as the market return through the pre-run-up period ($[-250, -11]$) prior to the cross-listing; then *2.* the product of "gross" residuals within an event window is subtracted by one to yield the CAR.

Company	*Industry*	*TSX Code*	*TSX Listing*	*NYSE Code*	*NYSE Listing*	*Listing Sequence*
Celestica Inc.	Electrical and Electronic Products	CLS	7 07, 1998	CLS	6 30, 1998	NYSE → TSX
Shaw Communications Inc.	Communications & Media	SJR.B	3 25, 1983	SJR	7 01, 1998	TSX → NYSE
NOVA Chemicals Corporation	Chemicals	NCX	7 03, 1998	NCX	7 06, 1998	TSX → NYSE
CGI Group Inc.	Consulting	GIB.A	4 21, 1992	GIB	10 07, 1998	TSX → NYSE
Brookfield Properties Corporation	Property Management and Investment	BPO	6 27, 1985	BPO	6 02, 1999	TSX → NYSE
Intertape Polymer Group Inc.	Packaging and Containers	ITP	1 06, 1993	ITP	8 16, 1999	TSX → NYSE
Gildan Activewear Inc.	Household Goods	*GIL*	*6 24, 1998*	*GIL*	*9 01, 1999*	*TSX → NYSE*
Manulife Financial Corp.	Insurance	MFC	9 30, 1999	MFC	9 24, 1999	NYSE → TSX
Sun Life Financial, Inc.	Insurance	SLF	3 29, 2000	SLF	3 23, 2000	NYSE → TSX
MDS Inc.	Medical Services	MDS	6 25, 1973	MDZ	4 07, 2000	TSX → NYSE
Corus Entertainment, Inc.	Entertainment Services	CJR.B	9 03, 1999	CJR	5 10, 2000	TSX → NYSE
Canadian Natural Resources, Ltd.	Oil and Gas Producers	CNQ	5 14, 1976	CNQ	7 31, 2000	TSX → NYSE
TELUS Corporation	Telephone Utilities	T.A	2 01, 1999	TU	10 17, 2000	TSX → NYSE
Nexen, Inc.	Oil and Gas Producers	NXY	7 14, 1971	NXY	11 14, 2000	TSX → NYSE
*Enerplus Resources Fund****	Oil and Gas Producers	*ERF.UN*	*3 11, 1987*	*ERF*	*11 17, 2000*	*TSX → NYSE*

	PIN		*PIN*		*PIN**		*Spread*		*Spread*		*Spread***	
Company	*-3M*	*+3M*	*-6M*	*+6M*	*Before*	*After*	*-3M*	*+3M*	*-6M*	*+6M*	*Before*	*After*
Celestica Inc.						0.186						
Shaw Communications Inc.				0.237		0.164						
NOVA Chemicals Corporation	0.329	0.326	0.329			0.268	0.006	0.007	0.006	0.007	0.006	0.007
CGI Group Inc.	0.183	0.283	0.176	0.277	0.256	0.226	0.268	0.151	0.181	0.123	0.150	0.055
Brookfield Properties Corporation		0.223		0.206	0.068	0.194		0.020		0.017	0.069	0.016
Intertape Polymer Group Inc.	0.218	0.247	0.209	0.209	0.266	0.262	0.034	0.025	0.026	0.025	0.027	0.031
Gildan Activewear Inc.												
Manulife Financial Corp.					0.035	0.150		0.003		0.004	0.221	0.003
Sun Life Financial, Inc.												
MDS Inc.	0.156	0.192	0.156	0.154	0.102	0.238	0.008	0.009	0.008	0.008	0.097	0.008
Corus Entertainment, Inc.	0.098	0.212	0.134	0.180	0.067	0.201	0.029	0.051	0.028	0.042	0.025	0.036
Canadian Natural Resources, Ltd.	0.142	0.127	0.159	0.348	0.152	0.128	0.004	0.003	0.004	0.004	0.007	0.004
TELUS Corporation	0.120	0.338		0.559	0.047	0.336	0.004	0.005	0.006	0.005	0.006	0.005
Nexen, Inc.	0.100	0.163	0.100	0.163	0.018	0.134	0.009	0.005	0.009	0.005	0.009	0.005
*Enerplus Resources Fund****												
Average	***0.168***	***0.235***	***0.180***	***0.259***	***0.112***	***0.207***	***0.045***	***0.028***	***0.033***	***0.024***	***0.062***	***0.017***

	Cumulative Abnormal Return			
Company	*[-2,+2]*	*[-5,+5]*	*[-10,+10]*	*[-10,+250]*
Celestica Inc.				
Shaw Communications Inc.	-0.002	0.156	0.242	0.024
NOVA Chemicals Corporation				
CGI Group Inc.	-0.204	-0.269	-0.204	-0.757
Brookfield Properties Corporation	-0.045	-0.041	-0.075	-0.358
Intertape Polymer Group Inc.	0.031	0.040	0.083	-0.740
Gildan Activewear Inc.	0.046	-0.029	-0.124	-0.477
Manulife Financial Corp.				
Sun Life Financial, Inc.				
MDS Inc.	0.018	-0.006	-0.037	-0.341
Corus Entertainment, Inc.	-0.033	-0.087	-0.047	-0.684
Canadian Natural Resources, Ltd.	0.042	-0.011	-0.019	-0.287
TELUS Corporation	-0.027	0.033	-0.011	-0.615
Nexen, Inc.	-0.025	-0.012	0.001	-0.364
*Enerplus Resources Fund****	0.014	0.007	-0.027	-0.287
Average	***-0.017***	***-0.020***	***-0.020***	***-0.444***

** Arithmetic mean of monthly PIN estimates. For derivation and estimation algorithm of PIN, see Appendix A3.*

*** Arithmetic mean of monthly spread estimates*

**** Prior to June of 2001, Enerplus Resources Fund traded under ERF.G. Upon the merger with EnerMark, the symbol became ERF.UN.*

Table A.8: Wilcoxon signed-rank tests

The PIN is the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996). The bid-ask spreads are defined as: *1.* $\text{SPREAD}_{\text{NYSE}} \equiv \frac{\text{ask}_{\text{NYSE}} - \text{bid}_{\text{NYSE}}}{(\text{ask}_{\text{NYSE}} + \text{bid}_{\text{NYSE}})/2}$; and *2.* $\text{SPREAD}_{\text{TSX}} \equiv \frac{\text{ask}_{\text{TSX}} - \text{bid}_{\text{TSX}}}{(\text{ask}_{\text{TSX}} + \text{bid}_{\text{TSX}})/2}$. VOLUME is the log of total daily trading volume. d is a differential measure defined for the estimates of each quantity of interest. The Wilcoxon signed-rank test is a non-parametric pair-wise comparison test, following Wilcoxon (1945). The Wilcoxon test-statistic is defined as: $V_0 \equiv \sum_{\{(i,t)\}} \mathbf{1}_{\{d(i,t)>0\}} \cdot \rho_{it}$, where ρ_{it} is the rank of $\{|d(i,t)|\}$, and the coordinates (i,t) denote each firm and each period, respectively.

Panel A: Cross-listing effect on the PIN on the TSX

	[-3M,+3M]	[-6M,+6M]	Threshold
H_0	$\text{PIN}_{+3\text{M}} = \text{PIN}_{-3\text{M}}$	$\text{PIN}_{+6\text{M}} = \text{PIN}_{-6\text{M}}$	$\text{PIN}_{\text{after}} = \text{PIN}_{\text{before}}$
H_1	$\text{PIN}_{+3\text{M}} > \text{PIN}_{-3\text{M}}$	$\text{PIN}_{+6\text{M}} > \text{PIN}_{-6\text{M}}$	$\text{PIN}_{\text{after}} > \text{PIN}_{\text{before}}$
d	$\text{PIN}_{+3\text{M}} - \text{PIN}_{-3\text{M}}$	$\text{PIN}_{+6\text{M}} - \text{PIN}_{-6\text{M}}$	$\text{PIN}_{\text{after}} - \text{PIN}_{\text{before}}$
V_0	33	14	30
p-value	0.01953	0.05282	0.05469

Panel B: Cross-listing effect on bid-ask spread on the TSX

	[-3M,+3M]	[-6M,+6M]	Threshold
H_0	$\text{SPREAD}_{+3\text{M}} = \text{SPREAD}_{-3\text{M}}$	$\text{SPREAD}_{+6\text{M}} = \text{SPREAD}_{-6\text{M}}$	$\text{SPREAD}_{\text{after}} = \text{SPREAD}_{\text{before}}$
H_1	$\text{SPREAD}_{+3\text{M}} < \text{SPREAD}_{-3\text{M}}$	$\text{SPREAD}_{+6\text{M}} < \text{SPREAD}_{-6\text{M}}$	$\text{SPREAD}_{\text{after}} < \text{SPREAD}_{\text{before}}$
d	$\text{SPREAD}_{-3\text{M}} - \text{SPREAD}_{+3\text{M}}$	$\text{SPREAD}_{-6\text{M}} - \text{SPREAD}_{+6\text{M}}$	$\text{SPREAD}_{\text{before}} - \text{SPREAD}_{\text{after}}$
V_0	45	48	72
p-value	0.34820	0.25740	0.05260

Panel C: Cross-listing effect on volume on the TSX

	[-3M,+3M]	[-6M,+6M]	Threshold
H_0	$\text{VOLUME}_{+3\text{M}} = \text{VOLUME}_{-3\text{M}}$	$\text{VOLUME}_{+6\text{M}} = \text{VOLUME}_{-6\text{M}}$	$\text{VOLUME}_{\text{after}} = \text{VOLUME}_{\text{before}}$
H_1	$\text{VOLUME}_{+3\text{M}} > \text{VOLUME}_{-3\text{M}}$	$\text{VOLUME}_{+6\text{M}} > \text{VOLUME}_{-6\text{M}}$	$\text{VOLUME}_{\text{after}} > \text{VOLUME}_{\text{before}}$
d	$\text{VOLUME}_{+3\text{M}} - \text{VOLUME}_{-3\text{M}}$	$\text{VOLUME}_{+6\text{M}} - \text{VOLUME}_{-6\text{M}}$	$\text{VOLUME}_{\text{after}} - \text{VOLUME}_{\text{before}}$
V_0	42	39	58
p-value	0.7293	0.6285	0.9433

Table A.9: Fixed-effect panel regressions of abnormal returns of TSX-listed stocks

12 TSX-listed firms cross-listed on the NYSE through the sample period: January 1, 1998 through December 31, 2000. For each firm (i) and in each month (t), *1.* the abnormal return ($\text{RETURN}_{\text{AB}}$) is the monthly cumulative return, following Binder (1998), using the S&P TSX Composite Index to obtain the market return; *2.* PIN is the monthly estimate of the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996); *3.* CROSS-LIST is a dummy variable which equals one in the month of cross-listing on the NYSE, or zero otherwise; *4.* SPREAD is the monthly average relative quoted spread; *5.* VOLUME is the monthly average of the log of daily total trading volume; and *6.* VOLATILITY is the standard deviation of daily returns multiplied by 250/12. Standard errors are corrected for heteroskedasticity and autocorrelation, following Newey and West (1987, 1994). The numerical values in the parentheses below the estimates are t-statistics. ***, **, and * stand for statistical significance based on two-sided student-t tests at the 1%, 5%, and 10% level, respectively. The observations are in firm-months.

Table A.9 (Continued)

$$\text{RETURN}_{\text{AB}} = \beta_0 + \beta_1\,\text{PIN} + \beta_2\,\text{CROSS-LIST} + \beta_3\,\text{SPREAD} + \beta_4\,\text{VOLUME} + \beta_5\,\text{VOLATILITY}$$
$$+ \beta_6\,\text{PIN} \times \text{CROSS-LIST} + \beta_7\,\text{SPREAD} \times \text{CROSS-LIST}$$
$$+ \beta_8\,\text{VOLUME} \times \text{CROSS-LIST} + \beta_9\,\text{VOLATILITY} \times \text{CROSS-LIST} + \epsilon$$

	Model 1	*Model 2*	*Model 3*	*Model 4*	*Model 5*	*Model 6*
(Intercept)	0.024	0.025*	0.025*	0.020	0.017	0.020
	(1.636)	(1.670)	(1.652)	(1.300)	(1.148)	(1.248)
PIN	−0.050	−0.037	−0.034	−0.065	−0.020	0.005
	(−0.456)	(−0.320)	(−0.294)	(−0.566)	(−0.175)	(0.041)
CROSS-LIST Dummy	−0.035*	−0.037*	−0.037*	−0.028	−0.024	−0.025
	(−1.886)	(−1.911)	(−1.898)	(−1.447)	(−1.278)	(−1.297)
SPREAD		−0.280	−0.281	−0.296		−0.548
		(−0.356)	(−0.355)	(−0.382)		(−0.559)
VOLUME			0.000	0.000	0.000	0.000
			(0.204)	(−0.570)	(1.491)	(1.427)
VOLATILITY				−0.124***	−0.077	−0.073
				(−3.185)	(−0.969)	(−0.913)
PIN×CROSS-LIST	−0.255*	−0.264*	−0.266*	−0.251*	−0.306**	−0.336**
	(−1.716)	(−1.747)	(−1.754)	(−1.686)	(−2.056)	(−2.138)
SPREAD×CROSS-LIST						0.874
						(0.544)
VOLUME×CROSS-LIST					0.000**	0.000*
					(−2.013)	(−1.957)
VOLATILITY×CROSS-LIST					−0.073	−0.077
					(−0.805)	(−0.834)
No. of Obs.	218	218	218	218	218	218
Adj. R^2	0.039	0.035	0.031	0.071	0.086	0.079

A.2 Figures

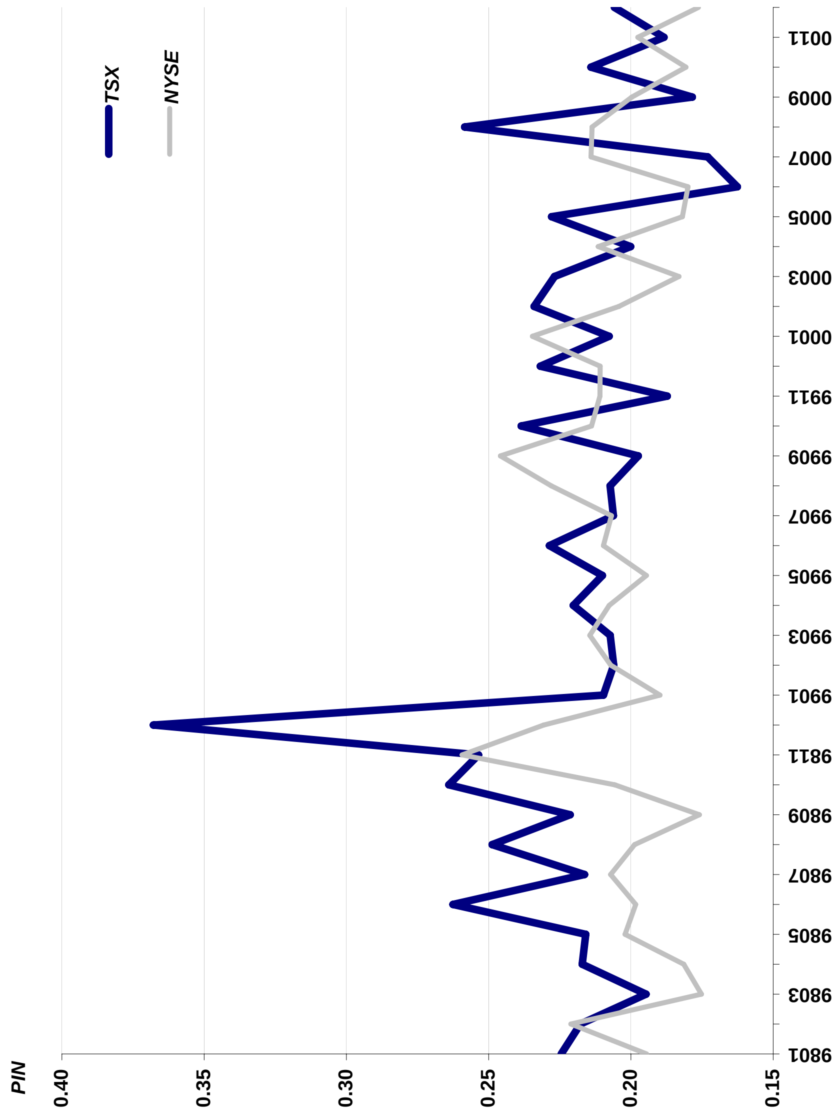
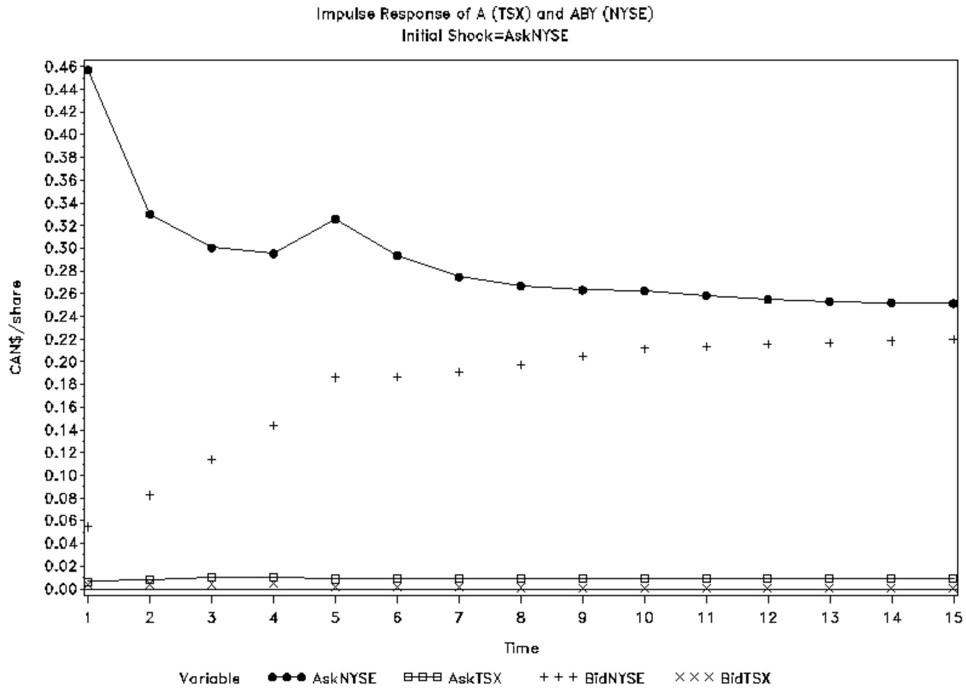
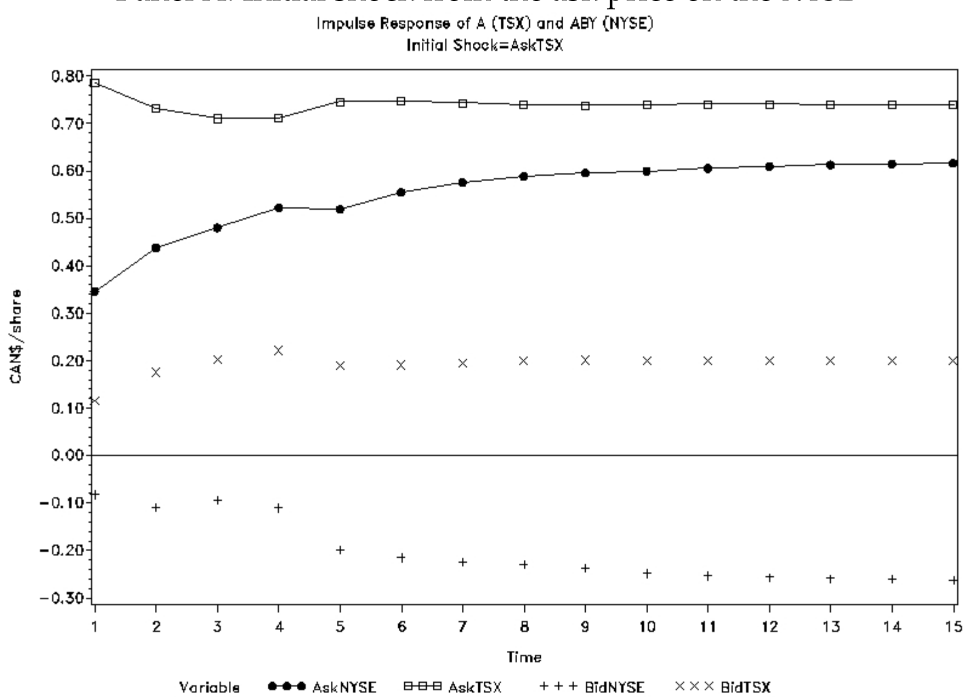


Figure A.1: Monthly estimates of PIN on TSX and NYSE

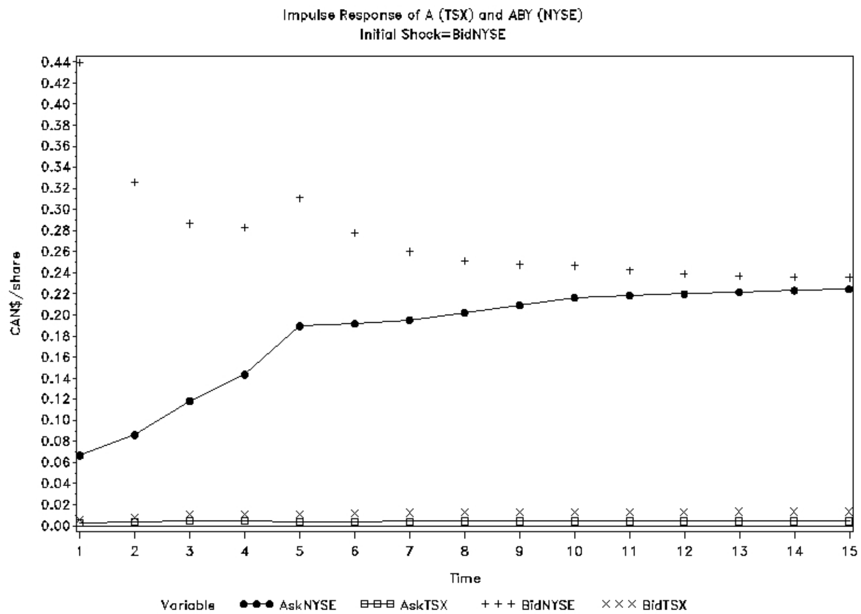
The PIN is the probability of informed trading, following Easley, Kiefer, O'Hara, and Paperman (1996). Figure A.1 shows the average monthly PIN of the sample firms co-listed on the TSX and the NYSE. The annual estimates for the PIN on the TSX are {0.242, 0.213, 0.206} in 1998, 1999, and 2000, respectively, while the corresponding estimates for the NYSE are {0.204, 0.212, 0.196}, over the same period.



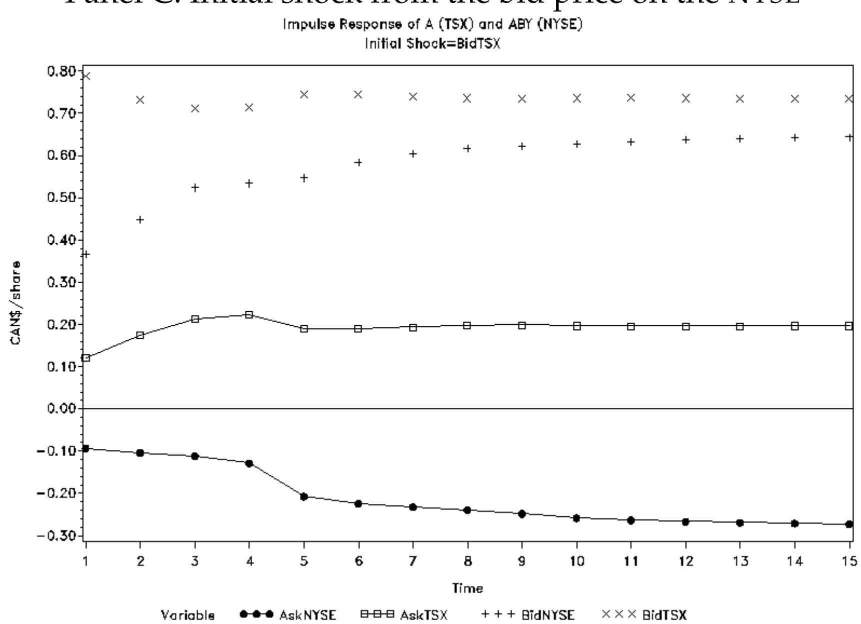
Panel A: Initial shock from the ask price on the NYSE



Panel B: Initial shock from the ask price on the TSX



Panel C: Initial shock from the bid price on the NYSE



Panel D: Initial shock from the bid price on the TSX

Figure A.2: Impulse response function plots: cross-border responses of quote changes

Each of the above four consecutive impulse response function plots of Aibiti Consolidate (co-listed on the TSX and on the NYSE) specifies the source of innovation by two standard deviations. The quotes on the NYSE rarely affect the quotes on the TSX. To the contrary, positive increases in ask and bid prices on the TSX are followed by changes in ask and bid prices on the NYSE, respectively.

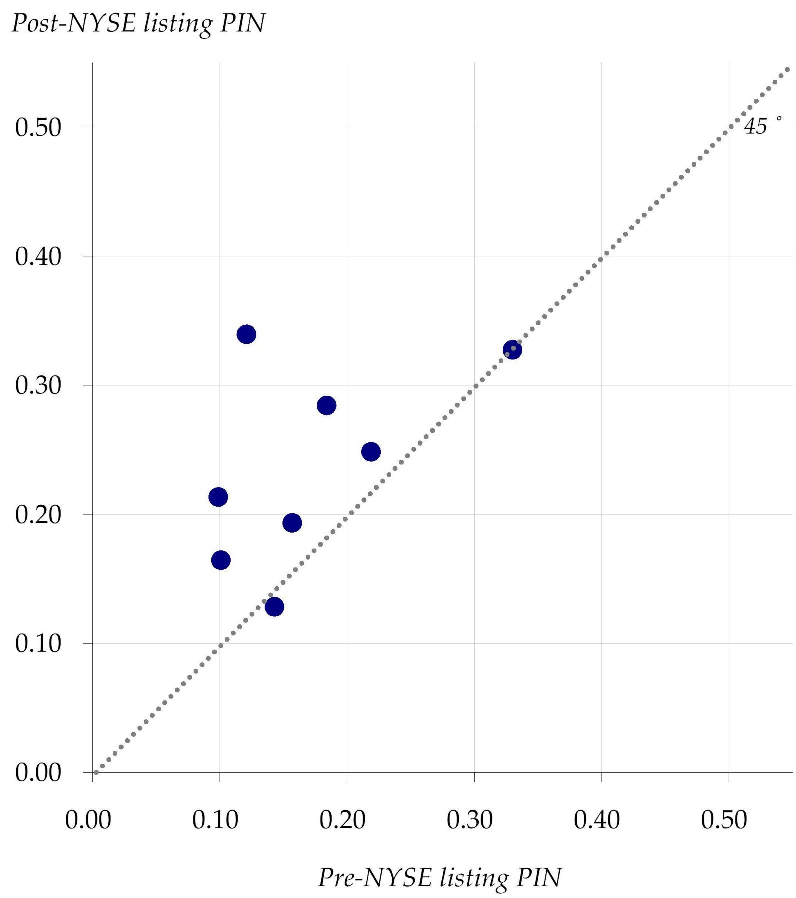


Figure A.3: Cross-listing effect on the PIN on the TSX, six-month ([-3M,+3M]) window

Above scatter plot describes various coordinates of the PIN on the TSX before (horizontal axis) and after (vertical axis) NYSE-listing. A coordinate in the upper 45°-line region denotes a rise in the PIN, whereas one in the lower region a decline.

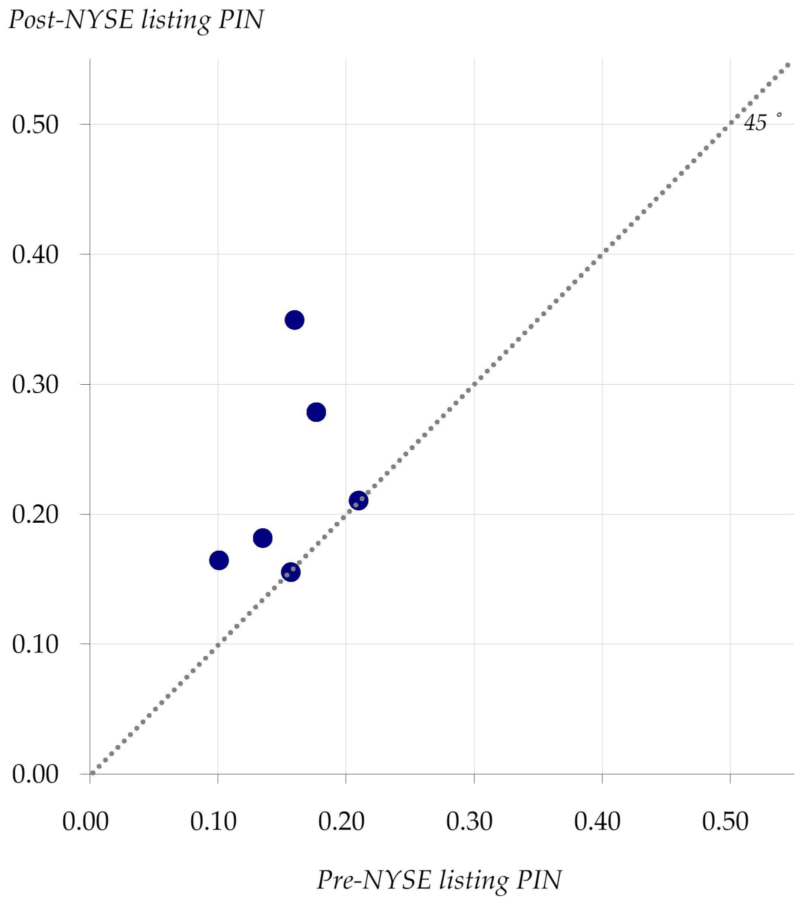


Figure A.4: Cross-listing effect on the PIN on the TSX, twelve-month ([-6M,+6M]) window

Above scatter plot describes various coordinates of the PIN on the TSX before (horizontal axis) and after (vertical axis) NYSE-listing. A coordinate in the upper 45°-line region denotes a rise in the PIN, whereas one in the lower region a decline.

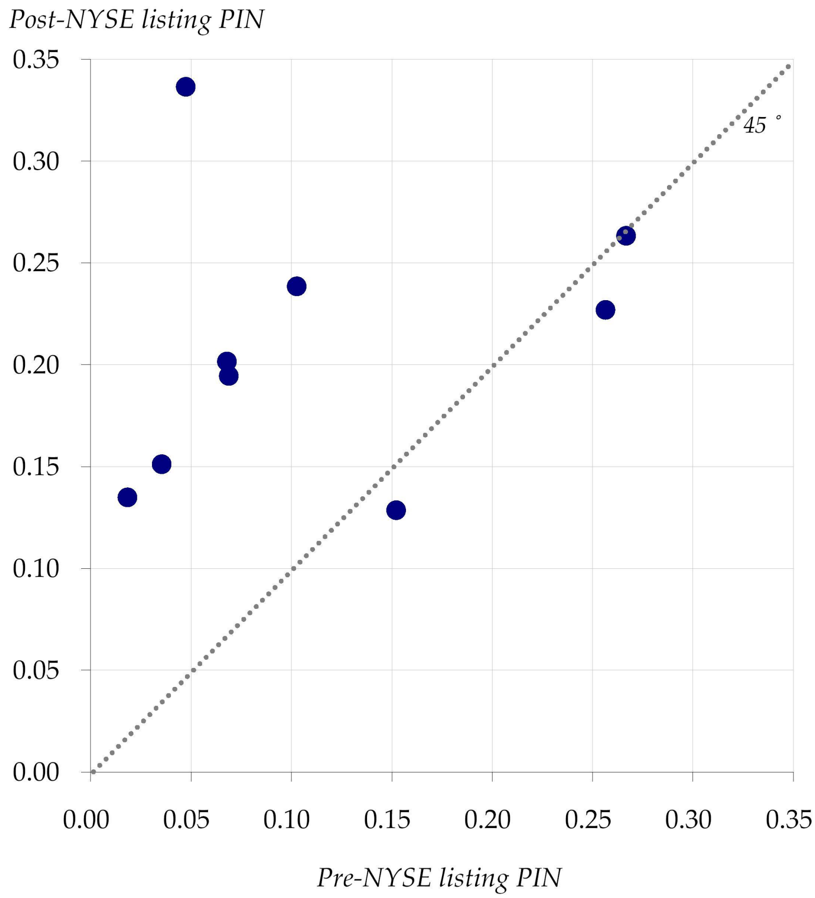


Figure A.5: Cross-listing effect on the PIN on the TSX, threshold monthly

Above scatter plot describes various coordinates of the PIN on the TSX before (horizontal axis) and after (vertical axis) NYSE-listing. A coordinate in the upper 45°-line region denotes a rise in the PIN, whereas one in the lower region a decline.

A.3 PIN estimation algorithm

The PIN estimation algorithm is based on a symmetric Poisson intensity η for arrivals of both uninformed buyers and sellers. Information events occur at the market open with a probability α and, on a realization of such event, informed traders who arrive with an intensity μ perceive a binary signal with a probability either $\delta \equiv \mathbb{P}\{\text{share price falls}\}$ or $1 - \delta = \mathbb{P}\{\text{share price rises}\}$.

The probability of informed trading (PIN) is the relative degree of private information (adverse selection) weighed on a randomly chosen transaction executed by an informed trader

$$\text{PIN} \equiv \frac{\alpha\,\mu}{\mathbb{E}\left[\text{B(uy)} + \text{S(ell)}\right]} = \frac{\alpha\,\mu}{\alpha\,\mu + \eta_{\text{B}} + \eta_{\text{S}}} = \frac{\alpha\,\mu}{\alpha\,\mu + 2\,\eta},$$

assuming symmetric intensity in uninformed trader arrivals, either buyers or sellers (see Figure A.6). Empirically, a trade is considered buyer-initiated if it is higher than the five-second earlier mid-quote, or seller-initiated if lower (Lee and Ready (1991)).

I adopt a log-likelihood factorization from Easley, Engle, O'Hara, and Wu (2008) as follows

$$\begin{aligned}
\mathscr{L} &\equiv \ln \mathbb{P}\left(\{B_t, S_t\}_{t=1}^{T} \mid \alpha, \delta, \eta, \mu\right) \\
&= \sum_{t=1}^{T} \left[-2\,\eta + M \ln(x) + (B_t + S_t) \ln(\mu + \eta)\right] \\
&\quad + \sum_{t=1}^{T} \ln\left[\alpha\,(1-\delta)\exp(-\mu)\,x^{S_t - M_t} + \alpha\,\delta\exp(-\mu)\,x^{B_t - M_t} + (1-\alpha)\,x^{B_t + S_t - M_t}\right],
\end{aligned}$$

where *1.* $M_t \equiv \frac{\min(B_t, S_t) + \max(B_t, S_t)}{2}$; and *2.* $x \equiv \frac{\eta}{\mu + \eta}$. Thus, the parameters are estimated by maximum likelihood method such that

$$\hat{\Theta} \equiv \left(\hat{\alpha}, \hat{\delta}, \hat{\eta}, \hat{\mu}\right) = \arg\max_{\Theta} \left\{\mathscr{L} \mid (\eta, \mu) > \mathbf{0},\ (\alpha, \delta) \in [0,1]^2\right\},$$

hence the resulting PIN estimator is

$$\widehat{\text{PIN}} = \frac{\hat{\alpha}\,\hat{\mu}}{\hat{\alpha}\,\hat{\mu} + 2\,\hat{\eta}}.$$

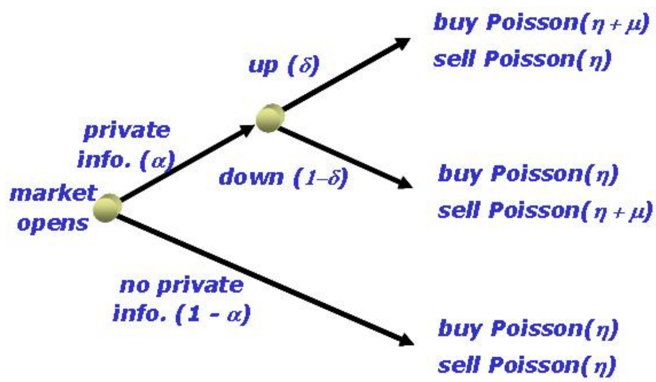


Figure A.6: Derivation of PIN

A.4 Information shares of stock exchanges

Consider a Canadian cross-listed pair $(p_{\mathrm{T}}, p_{\mathrm{N}})$ traded on both the TSX (T) and the NYSE (N). The time series of the pair has a common efficient price[1] (m_t) such that

$$\begin{bmatrix} p_{\mathrm{T},t} \\ p_{\mathrm{N},t} \end{bmatrix} = \begin{bmatrix} 1 \\ 1 \end{bmatrix} m_t + \begin{bmatrix} c_{\mathrm{T}}\, q_{\mathrm{T},t} \\ c_{\mathrm{N}}\, q_{\mathrm{N},t} \end{bmatrix},$$

where c_{T} and c_{N}, and q_{T} and q_{N} are market-specific cost coefficients and their associated trade volumes, respectively. Trade directions in the two markets may be contemporaneously associated as

$$\mathrm{Var}\left(\begin{bmatrix} q_{\mathrm{T},t} \\ q_{\mathrm{N},t} \end{bmatrix}\right) = \begin{bmatrix} 1 & \rho_q \\ \rho_q & 1 \end{bmatrix}.$$

An attractive trait of the common efficient price is that the securities with same underlying assets traded on distinct exchanges are linked by no-arbitrage condition in an equilibrium. An implied-vector moving average (VMA) formulation for the differences of prices is

$$\begin{bmatrix} \Delta p_{\mathrm{T},t} \\ \Delta p_{\mathrm{N},t} \end{bmatrix} = \begin{bmatrix} \epsilon_{\mathrm{T},t} \\ \epsilon_{\mathrm{N},t} \end{bmatrix} + \begin{bmatrix} q_{\mathrm{T},t} \\ q_{\mathrm{N},t} \end{bmatrix} = \begin{bmatrix} \theta_{\mathrm{TT}} & \theta_{\mathrm{TN}} \\ \theta_{\mathrm{NT}} & \theta_{\mathrm{NN}} \end{bmatrix} \begin{bmatrix} \epsilon_{\mathrm{T},t-1} \\ \epsilon_{\mathrm{N},t-1} \end{bmatrix},$$

then

$$\mathbb{E}_t \begin{bmatrix} p_{\mathrm{T},t+1} \\ p_{\mathrm{N},t+1} \end{bmatrix} = \begin{bmatrix} p_{\mathrm{T},t} \\ p_{\mathrm{N},t} \end{bmatrix} + \begin{bmatrix} \theta_{\mathrm{TT}} & \theta_{\mathrm{TN}} \\ \theta_{\mathrm{NT}} & \theta_{\mathrm{NN}} \end{bmatrix} \begin{bmatrix} \epsilon_{\mathrm{T},t-1} \\ \epsilon_{\mathrm{N},t-1} \end{bmatrix},$$

[1]A security price time-series $\left(\{m_t\}_{t=0}^{\infty}\right)$ is *efficient* if, by definition, the conditional expectation of the first-order difference is zero. In other words, an efficient price is unpredictable given the presently available information. Equivalently, the increment of the price follows a martingale difference sequence: $m_t = m_{t-1} + u_t \Rightarrow \mathbb{E}\left(\Delta m_t | \{m_{s-1}\}_{s=1}^{t}\right) = \mathbb{E}\left(u_t | \{m_{s-1}\}_{s=1}^{t}\right) = 0$. See Lee, White, and Granger (1993).

thus

$$\mathbb{E}_t \begin{bmatrix} p_{\mathrm{T},t+1} \\ p_{\mathrm{N},t+1} \end{bmatrix} - \mathbb{E}_{t-1} \begin{bmatrix} p_{\mathrm{T},t} \\ p_{\mathrm{N},t} \end{bmatrix} = \begin{bmatrix} \Delta p_{\mathrm{T},t} \\ \Delta p_{\mathrm{N},t} \end{bmatrix} + \begin{bmatrix} \theta_{\mathrm{TT}} & \theta_{\mathrm{TN}} \\ \theta_{\mathrm{NT}} & \theta_{\mathrm{NN}} \end{bmatrix} \begin{bmatrix} \Delta\epsilon_{\mathrm{T},t-1} \\ \Delta\epsilon_{\mathrm{N},t-1} \end{bmatrix}.$$

given that the two prices share the same efficient underlying price $(1+\theta_{\mathrm{TT}},\ \theta_{\mathrm{TN}}) = (\theta_{\mathrm{NT}},\ 1+\theta_{\mathrm{NN}})$.

Following Eun and Sabherwal (2003), the augmented Dickey-Fuller (1981) unit root test is conducted to each daily-price time series of the 56 TSX-NYSE cross-listed pairs with appropriate lag length, per Akaike (1974), to verify first-order integration ($I(1)$). Applying Johansen's (1991) either trace test or eigen-value test yielded one "cointegrating"[2] equation for each TSX-NYSE pair.

As a result, an econometric impasse is that since the cross-listed pairs are cointegrated, a vector moving average (VMA) representation cannot be recovered by Sims's (1980) vector autoregressive (VAR) structural formulation. Subsequently, in the absence of accounting for sources of shocks to fragmented shares, decomposing exchange-specific relative contribution to price discovery of the TSX-NYSE pairs poses an unwieldy task.

A breakthrough is introduced by Engle and Granger (1987) and Engle and Yoo (1987), and Hasbrouck (1995) adopts their error correction model (ECM) to arrive at the "information share": the percentage share of an exchange in price discovery of shares whose orders are executed from many markets. The vector error correction model (VECM) for the cointegrated trade-level quote prices is

$$\Delta p_t = \phi(L)\Delta p_t + \gamma(\alpha - z_{t-1}) + \epsilon_t,$$

[2]Security prices are cointegrated if there exists a linear combination of the non-stationary prices that can be toned stationary. A time series is strongly stationary if its probability distribution is time-invariant, and weakly stationary if up to its second moments: mean, variance, and covariance.

where *1.* $\phi(L)\Delta p_t$ are vector autoregressive terms; *2.* γ is a vector of cointegrating coefficients; *3.* $\alpha > 0$ is a vector of long-run cross-border bid-ask dollar spreads; and *4.* z_t is a vector of cross-border dollar spreads in ask $\left(p^a_{\mathrm{T},t}, p^a_{\mathrm{N},t}\right)$ and bid $\left(p^b_{\mathrm{T},t}, p^b_{\mathrm{N},t}\right)$ prices on the TSX and the NYSE, respectively, as:

$$z_t \equiv \begin{bmatrix} 1 & -1 & 0 & 0 \\ 1 & 0 & -1 & 0 \\ 1 & 0 & 0 & -1 \end{bmatrix} \begin{bmatrix} p^a_{\mathrm{T},t} \\ p^a_{\mathrm{N},t} \\ p^b_{\mathrm{T},t} \\ p^b_{\mathrm{N},t} \end{bmatrix} = \begin{bmatrix} p^a_{\mathrm{T},t} - p^a_{\mathrm{N},t} \\ p^a_{\mathrm{T},t} - p^b_{\mathrm{T},t} \\ p^a_{\mathrm{T},t} - p^b_{\mathrm{N},t} \end{bmatrix}.$$

A resulting VMA generalization is $\Delta p_t = \Theta(L)\,\epsilon_t$, where $\Omega \equiv \mathrm{Var}(\epsilon_t)$. Define $\sigma^2_\omega \equiv \beta\,\Omega\,\beta'$, where $\beta = (\beta_\mathrm{T}, \beta_\mathrm{N}) = (1 + \theta_\mathrm{TT},\ \theta_\mathrm{TN}) = (\theta_\mathrm{NT},\ 1 + \theta_\mathrm{NN})$. According to Hasbrouck (1995, 2007),

1. if Ω is *diagonal*, the *information share* of a market i (= T, N) is defined as

$$\mathrm{IS}_i \equiv \frac{\beta_i^2\ \mathrm{Var}(\epsilon_{i,t})}{\sigma^2_\omega}$$

   which is market i's proportional contribution to price discovery of a cross-listed pair.

2. If Ω is *non-diagonal*, the lower and upper bounds of information share can be obtained by re-ordering the sources of innovation (shock) with orthogonalized impulse response functions following Hasbrouck (2007). Given four quote prices $\left(p^a_{\mathrm{T},t}, p^a_{\mathrm{N},t}, p^b_{\mathrm{T},t}, p^b_{\mathrm{N},t}\right)$, there are 24 (= 4!) orderings in terms of Cholesky exogeneity. In other words, for each TSX-NYSE pair, there are 24 pairs of information shares of the TSX and the NYSE, respectively. Averaging across varying exogeneity reduces them to a single pair of information shares for each cross-listed pair.

The estimated information shares for 56 cross-listed pairs are listed in Table A.1. The impulse response function plots of bid and ask quotes for Abitibi Consolidated, Inc., are shown in Figure A.2. Each of the four consecutive charts specifies the source of innovation by two standard deviations. The quotes on the NYSE rarely affect the quotes on the TSX. To the contrary, positive increases in ask and bid prices on the TSX are followed by changes in ask and bid prices on the NYSE, respectively. This pattern does not hold for all cross-listed pairs, and the degree to which an exchange responds to the other side is reflected in the relative magnitude of information share.

A.5 Proofs

From the model in Subsection 1.2, I have

$$\beta_i^0 = \frac{\bar{v}}{1+r} - \frac{\bar{y}_i}{(1+r)(\omega_{\mathrm{T}}^I + \omega_{\mathrm{T}}^U)},$$

$$\beta_i^S = \frac{1}{(1+r)(\omega_i^I + \omega_i^U)} \left\{ \omega_i^I \left(\frac{\tau_\epsilon}{\tau_\epsilon + \tau_v} \right) + \omega_i^U \left(\frac{\phi_i \tau_\epsilon}{\phi_i \tau_\epsilon + \tau_v} \right) \right\},$$

$$\beta_i^Y = \frac{\omega_i^I \{\tau_\epsilon/(\tau_\epsilon + \tau_v)\}}{(1+r)(\omega_i^I + \omega_i^U)} \left\{ \omega_i^I \left(\frac{\tau_\epsilon}{\tau_\epsilon + \tau_v} \right) + \omega_i^U \left(\frac{\phi_i \tau_s}{\phi_i \tau_\epsilon + \tau_v} \right) \right\},$$

$$\beta_i^A = \frac{1}{(1+r)\left(\omega_i^I + \omega_i^U\right)},$$

where $\phi_i \equiv \frac{\pi_i^2 \eta^2 \tau_s \tau_y}{1+\pi_i^2 \eta^2 \tau_\epsilon \tau_y}$, $\omega_i^I \equiv \pi_i \eta (\tau_\epsilon + \tau_v)$, $\omega_i^U \equiv (1-\pi_i)\eta(\phi_i \tau_\epsilon + \tau_v)$, for all $i =$ T(SX), N(YSE). For brevity, I omit the exchange subscript i in the following proofs.

Proposition 1. $\partial\beta^0(\pi)/\partial\pi > 0$, for all $\pi \in [0,1]$.

Proof. Note that $\partial\beta^0(\pi)/\partial(\omega^I + \omega^U) > 0$, and

$$\frac{\partial(\omega^I + \omega^U)}{\partial\pi} = \frac{\eta\tau_\epsilon}{\left(\tau_y\tau_\epsilon\pi^2\eta^2 + 1\right)^2} \left(-\tau_y\tau_\epsilon\pi^2\eta^2 + 2\tau_y\tau_\epsilon\pi\eta^2 + 1\right),$$

where the quadratic solutions for $-\tau_y\tau_\epsilon\pi^2\eta^2 + 2\tau_y\tau_\epsilon\pi\eta^2 + 1 = 0$ are

$$\pi = \frac{1}{\eta\tau_y\tau_\epsilon}\left(\sqrt{\tau_y\tau_\epsilon + \eta^2\tau_y^2\tau_\epsilon^2} + \eta\tau_y\tau_\epsilon\right) > 1,$$

$$\pi = -\frac{1}{\eta\tau_y\tau_\epsilon}\left(\sqrt{\tau_y\tau_\epsilon + \eta^2\tau_y^2\tau_\epsilon^2} - \eta\tau_y\tau_\epsilon\right) < 0,$$

thus $\pi \in [0,1]$ implies $-\tau_y\tau_\epsilon\pi^2\eta^2 + 2\tau_y\tau_s\pi\eta^2 + 1 > 0$, hence $\partial(\omega^I + \omega^U)/\partial\pi > 0$. Therefore,

$$\partial\beta^0(\pi)/\partial\pi = \left\{\partial\beta^0(\pi)/\partial(\omega^I + \omega^U)\right\}\{\partial(\omega^I + \omega^U)/\partial\pi\} > 0 \text{ for all } \pi \in [0,1].$$

Proposition 2. $\partial\beta^S(\pi)/\partial\pi > 0$, for all $\pi \in [0,1]$.

Proof. An analogous argument to the proof of *Proposition 1* leads to

$$\frac{\partial\beta^S(\pi)}{\partial\pi} = \frac{(\tau_\upsilon\tau_\epsilon)\left(-\tau_y\tau_\epsilon\pi^2\eta^2 + 2\tau_y\tau_\epsilon\pi\eta^2 + 1\right)}{\left(\tau_y\pi^2\eta^2\tau_\epsilon^2 + \tau_\upsilon\tau_y\pi^2\eta^2\tau_\epsilon + \pi\tau_\epsilon + \tau_\upsilon\right)^2} > 0 \text{ for all } \pi \in [0,1].$$

Proposition 3. $\partial\beta^Y(\pi)/\partial\pi < 0$, for some large π.

Proof. A direct partial differentiation gives

$$\frac{\partial\beta^Y(\pi)}{\partial\pi} = -\frac{(\tau_\epsilon/\eta)\left(\pi^2\eta^4\tau_y^2\tau_\epsilon^2 + \tau_\upsilon\pi^2\eta^4\tau_y^2\tau_\epsilon + 2\pi\eta^2\tau_y\tau_\epsilon + 2\tau_\upsilon\pi\eta^2\tau_y - \tau_\upsilon\eta^2\tau_y + 1\right)}{\left(\tau_y\pi^2\eta^2\tau_\epsilon^2 + \tau_\upsilon\tau_y\pi^2\eta^2\tau_\epsilon + \pi\tau_\epsilon + \tau_\upsilon\right)^2},$$

where the solutions for $\pi^2\eta^4\tau_y^2\tau_\epsilon^2 + \tau_\upsilon\pi^2\eta^4\tau_y^2\tau_\epsilon + 2\pi\eta^2\tau_y\tau_\epsilon + 2\tau_\upsilon\pi\eta^2\tau_y - \tau_\upsilon\eta^2\tau_y + 1 = 0$ are

$$\pi = -\tfrac{1}{\eta^2\tau_y\tau_\epsilon(\tau_\upsilon+\tau_\epsilon)}\left\{\tau_\upsilon + \tau_\epsilon + \sqrt{\tau_\upsilon(\tau_\upsilon+\tau_\epsilon)\left(\eta^2\tau_y\tau_\epsilon+1\right)}\right\} < 0,$$

$$\pi = -\tfrac{1}{\eta^2\tau_y\tau_\epsilon(\tau_\upsilon+\tau_s)}\left\{\tau_\upsilon + \tau_\epsilon - \sqrt{\tau_\upsilon(\tau_\upsilon+\tau_\epsilon)\left(\eta^2\tau_y\tau_\epsilon+1\right)}\right\} \lesseqgtr 0 \ \text{ if } \tau_\upsilon\eta^2\tau_y \lesseqgtr 1.$$

Thus, if $\tau_\upsilon\eta^2\tau_y < 1$, there exists some constant $c \in [0,1]$ such that $\partial\beta^Y(\pi)/\partial\pi \gtreqless 0$ for $\pi \lesseqgtr c$; and if $\tau_\upsilon\eta^2\tau_y > 1$, then $\partial\beta^Y(\pi)/\partial\pi < 0$ for all $\pi \in [0,1]$. Therefore, $\partial\beta^Y(\pi)/\partial\pi < 0$ for some large π.

Proposition 4. $\partial\beta^A(\pi)/\partial\pi < 0$, for all $\pi \in [0,1]$.

Proof. Note that $\partial\beta^A(\pi)/\partial(\omega^I + \omega^U) > 0$ and, from the proof of *Proposition 1*,

$$\partial\left(\omega^I + \omega^U\right)/\partial\pi > 0 \text{ for all } \pi \in [0,1].$$

Therefore, $\partial\beta^A(\pi)/\partial\pi = \left\{\partial\beta^A(\pi)/\partial(\omega^I + \omega^U)\right\}\{\partial(\omega^I + \omega^U)/\partial\pi\} > 0$ for all $\pi \in [0,1]$.

Proposition 5. There exists no arbitrage in an equilibrium if $p_{\mathrm{N}} - p_{\mathrm{T}} = \beta^0_{\mathrm{N}} - \beta^0_{\mathrm{T}}$.

Proof. The prices of a TSX-NYSE cross-listed pair are, respectively,

$$p_{\mathrm{T}} = \beta^0_{\mathrm{T}} + \beta^S_{\mathrm{T}}\Delta S - \beta^Y_{\mathrm{T}}\Delta Y_{\mathrm{T}} - \beta^A_{\mathrm{T}} x^A_{\mathrm{T}},$$

$$p_{\mathrm{N}} = \beta^0_{\mathrm{N}} + \beta^S_{\mathrm{N}}\Delta S - \beta^Y_{\mathrm{N}}\Delta Y_{\mathrm{N}} - \beta^A_{\mathrm{N}} x^A_{\mathrm{N}}.$$

In a disequilibrium, arbitrageurs' profit in excess of the required cross-listing dollar premium is

$$(p_{\mathrm{N}} - p_{\mathrm{T}}) - \left(\beta^0_{\mathrm{N}} - \beta^0_{\mathrm{T}}\right) = \left(\beta^S_{\mathrm{T}} - \beta^S_{\mathrm{N}}\right)\Delta S + \beta^Y_{\mathrm{N}}\Delta Y_{\mathrm{N}} - \beta^Y_{\mathrm{T}}\Delta Y_{\mathrm{T}} - \beta^A_{\mathrm{N}} x^A_{\mathrm{T}} + \beta^A_{\mathrm{T}} x^A_{\mathrm{N}},$$

then given perfect hedging ($\mu \equiv x^A_{\mathrm{T}} = -x^A_{\mathrm{N}}$), arbitrageurs' short (long) position on the TSX (NYSE) is

$$\mu = \frac{\left(\beta^0_{\mathrm{N}} - \beta^0_{\mathrm{T}}\right) - (p_{\mathrm{N}} - p_{\mathrm{T}}) + \left(\beta^S_{\mathrm{T}} - \beta^S_{\mathrm{N}}\right)\Delta S + \beta^Y_{\mathrm{N}}\Delta Y_{\mathrm{N}} - \beta^Y_{\mathrm{T}}\Delta Y_{\mathrm{T}}}{\beta^A_{\mathrm{N}} + \beta^A_{\mathrm{T}}},$$

thus, in an equilibrium ($\Delta S = \Delta Y_{\mathrm{T}} = \Delta Y_{\mathrm{N}} = 0$), the no-arbitrage ($\mu = 0$) condition must be

$$p_{\mathrm{N}} - p_{\mathrm{T}} = \beta^0_{\mathrm{N}} - \beta^0_{\mathrm{T}}.$$

APPENDIX B

CHAPTER 2 OF APPENDIX

B.1 Tables

Table B.1: Sovereign corporate governance measures: cross-border target countries

Accounting Standards (AS) is from La Porta, et al. (1998), and Antidirector Rights (AD)—which proxies for the degree of shareholder protection—is from Djankov, et al. (2008). As a relative measure of country-specific equity market development, Stock Market Capitalization to GDP (SMCTG) is suggested by Djankov, et al. (2008).

Country	AD	AS	SMCTG	Country	AD	AS	SMCTG
Argentina	2	45	4.062	Jordan	1		4.352
Australia	4	75	4.625	Kenya	2		2.728
Austria	3	54	2.797	Luxembourg	2		4.974
Belgium	3	61	4.208	Malaysia	5	76	5.000
Bolivia	2		2.747	Mexico	3	60	3.086
Brazil	5	54	3.648	Netherlands	3	64	4.881
Canada	4	74	4.665	New Zealand	4	70	3.691
Chile	4	52	4.496	Norway	4	74	3.681
China	1		3.768	Peru	4	38	3.127
Colombia	3	50	2.660	Philippines	4	65	3.871
Croatia	3		2.803	Poland	2		2.815
Czech Republic			3.006	Portugal	3	36	3.833
Denmark	4	62	4.071	Romania	5		1.705
Ecuador	2		1.758	Russia	4		3.503
Finland	4	77	5.177	Singapore	5	78	5.105
France	4	69	4.494	South Africa	5	70	5.049
Germany	4	62	4.002	South Korea	5	62	3.991
Greece	2	55	4.515	Spain	5	64	4.381
Hong Kong	5	69	5.889	Sweden	4	83	4.721
Hungary	2		3.178	Switzerland	3	68	5.517
India	5	57	3.520	Taiwan	3	65	4.624
Indonesia	4		3.207	Thailand	4	64	3.802
Ireland	5		4.214	Turkey	3	51	3.564
Israel	4	64	3.970	United Kingdom	5	78	5.061
Italy	2	62	3.967	United States	3	71	4.957
Japan	5	65	4.237	Venezuela	1	40	1.705

Table B.2: Panel Regression analysis of U.S. Domestic target returns

The sample consists of 1,439 completed U.S. mergers and acquisitions (listed in SDC) between 1990 and 2007 made by firms covered by the IRRC ATP provision database. The dependent variable is the target's 5-day ([±2]), 11-day ([±5]), 21-day ([±10]) windows of cumulative abnormal return around announcement dates. G Index is the number of ATPs of U.S. acquirers provided by Gompers, et al. (2003). Deal Size is the log transaction Value recorded on SDC. Acquirer High-Tech equals 1 if the bidder is from high tech industries defined by Loughran and Ritter (2004), 0 otherwise. Target High-Tech equals 1 if the target is from high tech industries defined by Loughran and Ritter (2004), 0 otherwise. Acquirer Assets is the log of book value of total assets. Tobin's (1969) Q is the market value over the book value of the acquirer's assets. Leverage is the book value of debts over market value of total assets. Free Cash Flow (FCF) is operating income before depreciation minus interest expenses minus income taxes minus capital expenditures, scaled by Acquirer Assets. Diversifying Dummy equals 1 if bidder and target do not share a Fama-French industry, 0 otherwise. M&A Market Condition is the average premium paid for all deals in a given year, computed as the average of premium paid based on the target stock price four weeks prior to merger announcement in a given year for all announced mergers in our sample. Penny Dummy equals 1 if the target stock trades below $10 in a given year. Standard errors are corrected for heteroskedasticity and autocorrelation, following Newey and West (1987, 1994). ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively. All regressions control for year fixed effects, whose coefficient estimates are suppressed.

	(1)			(2)			(3)			(4)		
Variable	[±2]	[±5]	[±10]	[±2]	[±5]	[±10]	[±2]	[±5]	[±10]	[±2]	[±5]	[±10]
G Index	0.002 **	0.003 **	0.003 **	0.005 **	0.005 **	0.006 **	0.005 *	0.005 *	0.006 **	0.005 **	0.006 **	0.006 **
Deal Size	-0.010 ***	-0.010 ***	-0.071 ***	-0.016 ***	-0.020 ***	-0.020 ***	-0.016 ***	-0.019 ***	-0.024 ***	-0.011 *	-0.013 **	-0.092 ***
Cash Dummy	0.050 ***	0.050 ***	0.042 ***	0.069 ***	0.073 ***	0.070 *	0.069 ***	0.072 ***	0.069 ***	0.076 ***	0.080 ***	0.076 ***
Target High-tech	0.040 ***	0.030 **	0.040 ***	0.053 **	0.026	0.030	0.053 **	0.027	0.025	0.063 **	0.036	0.034
Acquirer High-tech	0.004	0.020	0.020	-0.065 ***	-0.050 *	-0.050 *	-0.065 ***	-0.050 *	-0.047 *	-0.060 **	-0.045 *	-0.043 *
Acquirer Leverage				-0.080	-0.152 **	-0.140 **	-0.079	-0.149 ***	-0.140 **	-0.117 **	-0.189 ***	-0.140 ***
Acquirer Tobin's Q				0.002	0.004	0.005	0.002	0.003	0.005	0.000	0.000	0.002
Acquirer FCF				0.274 *	0.108	0.083	0.294 **	0.116	0.086	0.294 **	0.128	0.101
Acquirer Asset				-0.002	-0.006	-0.004	-0.002	-0.006	-0.004	-0.002	-0.006	-0.004
Deal Size × Target High-tech				0.028 **	0.023 **	0.016	0.022 **	0.023 **	0.016 *	0.024 ***	0.025 ***	0.018 *
Diversifying Dummy				0.021	0.039 ***	0.034 **	0.021	0.024 **	0.034 **	0.025 *	0.042 ***	0.037 **
Low Price Dummy							-0.007	-0.025	-0.011			
M&A Market Condition										0.003 ***	0.003 ***	0.003 ***
Number of Obs.	1,439	1,439	1,439	526	526	526	526	526	526	526	526	526
Multiple R^2	0.060	0.060	0.070	0.092	0.103	0.111	0.090	0.100	0.110	0.130	0.130	0.140

Table B.3: Cumulative abnormal returns (CARs) of foreign targets and U.S. acquirers

In Panel A, cross-border acquiree returns (CARs) are cumulative abnormal returns of cross-border targets through 5-day, 11-day, and 21-day event study windows. Bidder cumulative abnormal returns (ACARs) are of the U.S. acquirers through the same respective periods. Wilcoxon test is a statistical significance test for nonparametric pairwise comparison. In Panel B, CAR[±d] is the cumulative abnormal returns of cross-border targets through (2d+1)-day event study window. ACAR[±d] is the cumulative abnormal returns of U.S. acquirers through the same period. ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively.

Panel A: Sample means of CARs with Wilcoxon-test p-values—cross-border deals

	CAR	H_1 : CAR > 0	ACAR	H_1 : ACAR < 0
[+2,-2]	0.137	0.000	-0.0053	0.000
[-5,+5]	0.140	0.000	-0.0073	0.002
[-10,+10]	0.184	0.000	-0.0056	0.029

Panel B: Correlation matrix

	CAR[±5]	CAR[±10]	ACAR[±2]	ACAR[±5]	ACAR[±10]
CAR[±2]	0.847 ***	0.732 ***	-0.056 ***	-0.018 **	-0.041 **
CAR[±5]		0.846 ***	-0.083 ***	-0.050 ***	-0.065 ***
CAR[±10]			-0.063 ***	-0.046 ***	-0.044 ***
ACAR[±2]				0.764 **	0.557 **
ACAR[±5]					0.726 ***

Table B.4: Cumulative abnormal returns (CARs) of foreign targets

We follow the classification of "dictators" by Masulis et al. (2007) if the acquirer has a G Index score–number of ATPs–higher than or equal to ten, or "democrats" if less than or equal to nine. Gompers, et al. (2003) previously classified dictators with firms with more than or equal to 14 ATPs and democrats with firms with less than or equal to five ATPs. CAR[±d]|dict is the cumulative abnormal returns of cross-border targets acquired by dictatorial U.S. acquirers through (2d+1)-day event study window. CAR[±d]|demo is that of democratic U.S. acquirers.

Panel A: Representative statistics

	Mean	Median	Standard dev.	No. of deals
CAR[-2,+2]|dict	0.155	0.029	0.575	235
CAR[-2,+2]|demo	0.121	0.015	0.267	243
CAR[-5,+5]|dict	0.141	0.044	0.361	235
CAR[-5,+5]|demo	0.127	0.013	0.277	243
CAR[-10,+10]|dict	0.217	0.054	1.418	236
CAR[-10,+10]|demo	0.151	0.011	0.341	243

Panel B: H_1 : CAR|dict > CAR|demo—cross-border deals

	Wilcoxon p-value
H_1: CAR[-2,+2]|dict > CAR[-2,+2]|demo	0.0661
H_1: CAR[-5,+5]|dict > CAR[-5,+5]|demo	0.0358
H_1: CAR[-10,+10]|dict > CAR[-10,+10]|demo	0.0194

Panel C: Cross-border acquiree returns (CARs) in 100% acquisitions

	Mean	Median	Standard dev.	No. of deals
CAR[-2,+2]|dict	0.213	0.117	0.290	84
CAR[-2,+2]|demo	0.239	0.122	0.366	82
CAR[-5,+5]|dict	0.232	0.154	0.342	84
CAR[-5,+5]|demo	0.245	0.136	0.363	82
CAR[-10,+10]|dict	0.233	0.174	0.272	84
CAR[-10,+10]|demo	0.298	0.179	0.432	82

Panel D: H_1 : CAR|demo > CAR|dict in 100% acquisitions—cross-border deals

	Wilcoxon p-value
H_1: CAR[-2,+2]|demo > CAR[-2,+2]|dict	0.486
H_1: CAR[-5,+5]|demo > CAR[-5,+5]|dict	0.404
H_1: CAR[-10,+10]|demo > CAR[-10,+10]|dict	0.493

Table B.5: Cumulative abnormal returns (CARs) of foreign targets and U.S. acquirers

In Panel A, G Index is the number of ATPs of U.S. acquirers provided by Gompers, et al. (2003). CAR[±d] is the cumulative abnormal returns of cross-border targets through (2d+1)-day event study window. ACAR[±d] is the cumulative abnormal returns of U.S. acquirers through the same period. In Panel B, the sample consists of 599 completed cross-border takeover deals (listed in SDC) for public foreign targets by U.S. acquirers covered by the IRRC antitakeover provision database between October 31, 1984, and October 15, 2007. The dependent variable is the sample means of cross-border target's 5-day, 11-day, 21-day windows of cross-cumulative abnormal returns around announcement dates. Panel C shows the same results for foreign takeover announcement returns made by U.S. acquirers. Standard errors are corrected for heteroskedasticity and autocorrelation, following Newey and West (1987, 1994). ***, **, and * stand for statistical significance based on two-side tests at the 1%, 5%, and 10% level, respectively. The G Index (Gompers, et al. (2003)) is the number of antitakeover provisions of U.S. acquirers.

Panel A: Sample means of CARs per number of ATPs—cross-border deals

	[-2,+2]		[-5,+5]		[-10,+10]	
G Index	CAR	ACAR	CAR	ACAR	CAR	ACAR
2	0.061	-0.309	-0.022	-0.354	0.140	-0.239
3	-0.001	0.012	-0.001	0.042	-0.001	0.035
4	0.091	0.034	0.088	0.048	0.116	-0.012
5	0.110	-0.012	0.113	-0.009	0.177	-0.018
6	0.070	0.002	0.087	-0.004	0.124	0.005
7	0.153	-0.017	0.160	-0.005	0.180	0.029
8	0.092	-0.018	0.063	-0.019	0.061	-0.024
9	0.109	-0.012	0.132	-0.009	0.147	-0.004
10	0.137	-0.020	0.158	-0.023	0.140	-0.017
11	0.103	-0.002	0.113	0.009	0.110	0.002
12	0.123	-0.005	0.118	0.006	0.123	0.009
13	0.175	-0.004	0.121	-0.002	0.127	-0.002
14	0.070	0.023	0.066	0.022	0.067	0.052
15	0.054	-0.017	0.040	0.051	0.047	0.035
16	0.069	0.023	0.306	0.027	0.329	0.032
18	-0.002	-0.003	0.008	-0.010	-0.007	0.057
19	-0.001	-0.065	0.087	0.047	0.093	0.093

Panel B: Cross-border CAR of foreign targets against G Index of U.S. acquirers

	[-2,+2]		[-5,+5]		[-10,+10]	
Variable	Linear	Quadratic	Linear	Quadratic	Linear	Quadratic
Intercept	0.106 **	-0.054	0.106 **	-0.068	0.144 **	0.001
G Index	0.000	0.040 ***	0.000	0.037 ***	0.000	0.036
$(G\ Index)^2$		-0.002 ***		-0.002 **		-0.001
F-statistic	0.018	8.381 ***	1.442	5.362 **	0.002	1.351
Adj. R^2	-0.065	0.480	-0.065	0.353	-0.067	0.042

Panel C: Foreign acquisition announcement returns of U.S. Acquirers against their G Index

	[-2,+2]		[-5,+5]		[-10,+10]	
Variable	Linear	Quadratic	Linear	Quadratic	Linear	Quadratic
Intercept	-0.067 **	-0.185 **	-0.086 *	-0.177 *	-0.082 ***	-0.113 *
G Index	0.004 **	0.034 ***	0.007 **	0.031 **	0.008 ***	0.016 **
$(G\ Index)^2$		-0.001 ***		-0.001 **		0.000
F-statistic	1.419	3.125 *	3.257 *	2.608	9.673 ***	4.874 **
Adj. R^2	0.025	0.210	0.124	0.167	0.352	0.326

Table B.6: Sample means of CARs and ACARs by U.S. acquirers' and target countries' corporate governance

We follow the classification of "dictators" by Masulis, et al. (2007) if the acquirer has a G Index score—number of ATPs—higher than or equal to ten, or "democrats" if less than or equal to nine. Gompers, et al. (2003) previously classified dictators with firms with more than or equal to 14 ATPs and democrats with firms with less than or equal to five ATPs. In Panel A, CAR[±d]|dict is the cumulative abnormal returns of cross-border targets acquired by dictatorial U.S. acquirers through (2d+1)-day event study window. CAR[±d]|demo is that of democratic U.S. acquirers. In Panel B, ACAR is the cumulative abnormal returns of U.S. acquirers with analogous definitions. Accounting Standards (AS) is from La Porta, et al. (1998), and Antidirector Rights (AD)—which proxies the degree of shareholder protection—is from Djankov, et al. (2008). As a relative measure of country-specific equity market development, Stock Market Capitalization to GDP (SMCTG) is suggested by Djankov, et al. (2008). Any "high" dummy variable equals one if a country's sovereign corporate governance score is higher than the median, and zero otherwise. ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively.

Panel A: Sample means of CARs by U.S. acquirers' and target countries' corporate governance

	AD		AS		SMCTG	
	Low	High	Low	High	Low	High
CAR[±2]|dict	0.043 **	0.166 ***	0.013	0.173 ***	0.012 **	0.179 ***
CAR[±2]|demo	0.048 **	0.131 ***	0.030	0.134 ***	0.018	0.135 ***
CAR[±5]|dict	0.048 **	0.151 ***	0.007	0.159 ***	0.010	0.164 ***
CAR[±5]|demo	0.053 *	0.136 ***	0.043	0.139 ***	0.022	0.141 ***
CAR[±10]|dict	0.041 **	0.235 **	-0.006	0.246 **	0.001	0.254 **
CAR[±10]|demo	0.119 **	0.155 ***	0.088 *	0.160 ***	0.032	0.167 ***

Panel B: H_1: $CAR|_{dict} > CAR|_{demo}$—p-values of Wilcoxon difference tests

	AD		AS		SMCTG	
	Low	High	Low	High	Low	High
CAR[±2]	0.412	0.114	0.713	0.080	0.162	0.089
CAR[±5]	0.243	0.094	0.671	0.054	0.272	0.061
CAR[±10]	0.571	0.043	0.483	0.042	0.103	0.062

Panel C: H_1: $CAR|_{High} > CAR|_{Low}$—p-values of Wilcoxon difference tests

	AD	AS	SMTCG
CAR[±2]|dict	0.041	0.000	0.000
CAR[±2]|demo	0.058	0.006	0.001
CAR[±5]|dict	0.073	0.000	0.000
CAR[±5]|demo	0.122	0.020	0.003
CAR[±10]|dict	0.015	0.000	0.000
CAR[±10]|demo	0.407	0.076	0.006

Panel D: Sample means of ACARs by U.S. acquirers' and target countries' corporate governance

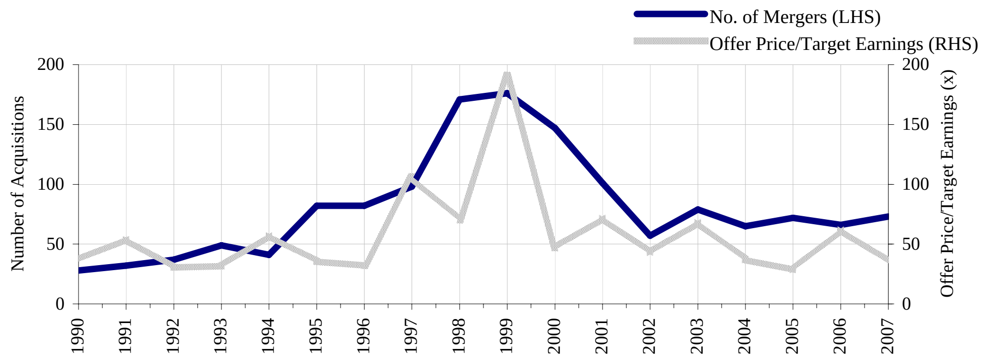
	AD		AS		SMCTG	
	Low	High	Low	High	Low	High
ACAR[±2]|dict	0.007	-0.010 **	-0.001	-0.008 *	0.000	-0.009 **
ACAR[±2]|demo	0.004	-0.007	0.001	-0.007	0.001	-0.007
ACAR[±5]|dict	-0.009	-0.009	-0.006	-0.010	-0.011	-0.009
ACAR[±5]|demo	0.011	-0.006	0.001	-0.005	0.005	-0.006
ACAR[±10]|dict	-0.007	-0.003	-0.004	-0.003	0.007	-0.005
ACAR[±10]|demo	-0.001	-0.009	-0.004	-0.009	-0.002	-0.010

Table B.7: Panel Regression analysis of cross-border target returns

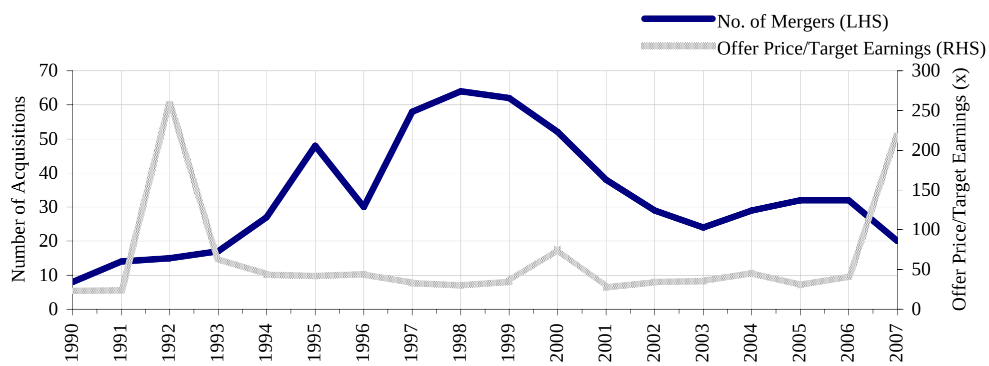
The depedent variable is the cross-border acquiree's cumulative abnormal return over 5-day ([±2]), 11-day ([±5]), and 21-day ([±10]) event windows around the merger announcement date, respectively. Full Acquire equals one when the U.S. bidder acquires the full ownership of its foreign target shares. G Index is the G Index per Gompers, et al. (2003). Deal Size is the log transaction Value recorded on SDC. Acquirer High-Tech equals 1 if the bidder is from high tech industries defined by Loughran and Ritter (2004), 0 otherwise. Target High-Tech equals 1 if the target is from high tech industries defined by Loughran and Ritter (2004), 0 otherwise. Acquirer Assets is the log of book value of total assets. Tobin's (1969) Q is the market value over the book value of the acquirer's assets. Leverage is the book value of debts over market value of total assets. Free Cash Flow (FCF) is operating income before depreciation minus interest expenses minus income taxes minus capital expenditures, scaled by Acquirer Assets. Diversifying Dummy equals 1 if bidder and target do not share a Fama-French industry, 0 otherwise. M&A Market Condition is the average premium paid for all deals in a given year, computed as the average of premium paid based on the target stock price four weeks prior to merger announcement in a given year for all announced mergers in our sample. Foreign Exchange is the annual average foreign exchange rate of the domestic currency per U.S.$ in the year of announcement. GDP is the log-difference in per capita GDP between the U.S. and target country. Antidirector is the difference of Djankov, et al.'s (2008) Antidirector Rights indices of the U.S. acquirer and its cross-border target. Accounting is the difference of La Porta, et al.'s (1998) Accounting Standards indices of the U.S. acquirer and its cross-border target. Stock Market is the stock market capitalization to GDP ratio per Djankov, et al. (2008). Standard errors are corrected for heteroskedasticity and autocorrelation, following Newey and West (1987, 1994). ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively. All regressions control for target country and year fixed effects, whose coefficient estimates are suppressed.

	(1)			(2)			(3)			(4)		
Variable	[±2]	[±5]	[±10]	[±2]	[±5]	[±10]	[±2]	[±5]	[±10]	[±2]	[±5]	[±10]
Full Acquire	0.290	-0.368	0.018	-0.130	-0.496	-0.173	0.125	-0.182	0.020	0.035	-0.731	-0.658
G Index	-0.005	0.006	0.024 *	0.000	0.011	0.031 **	-0.001	0.018 *	0.044 ***	-0.027	-0.139	-0.151
Full Acquire × G Index	-0.022	0.053 *	0.037	-0.015	0.058 **	0.043	0.064 *	0.029	0.046	0.057	0.071 *	0.098 **
Full Acquire × AD				-0.965	-1.496 ***	-2.370 ***	-0.748	-1.778 **	-3.151 ***	-0.812	-2.170 **	-3.635 ***
Full Acquire × AS				0.215	0.413 ***	0.654 ***	0.098	0.489 *	0.914 ***	0.111	0.568 **	1.012 ***
$(\text{G Index})^2$										0.001	0.008	0.010
Deal Size							-0.009	0.004	0.011	-0.009	-0.002	0.004
Cash Dummy							-0.078 *	-0.039	-0.034	-0.076	-0.026	-0.017
Acquirer High-tech							-0.105	0.014	0.132	-0.110	-0.018	0.094
Acquirer Leverage							4.081 **	3.085 **	2.488 *	4.070 **	3.021 **	2.409 *
Acquirer Tobin's Q							0.100	0.099 *	0.132 **	0.098	0.091	0.122 *
Acquirer FCF							4.699 **	5.379 ***	6.826 ***	4.761 **	5.760 ***	7.296 ***
Acquirer Asset							-0.093 **	-0.072 **	-0.061 *	-0.093 **	-0.071 **	-0.060 *
Deal Size × Target High-tech							-0.036	-0.037	-0.026	-0.034	-0.026	-0.013
Diversifying Dummy							-0.021	-0.070	-0.139 **	-0.019	-0.058	-0.124 *
Low Price Dummy							-0.082	-0.045	-0.048	-0.082	-0.042	-0.044
M&A Market Condition							0.005 **	0.005 ***	0.005 ***	0.006 **	0.006 ***	0.006 ***
Foreign Exchange							0.000	0.000	0.000	0.000	0.000	0.000
GDP							-0.164	-0.222	-0.200	-0.178	-0.303 **	-0.300 *
No. of Obs.	457	457	457	441	441	441	436	436	436	436	436	436
Multiple R^2	0.035	0.197	0.368	0.108	0.308	0.474	0.472	0.680	0.851	0.472	0.701	0.863

B.2 Figure



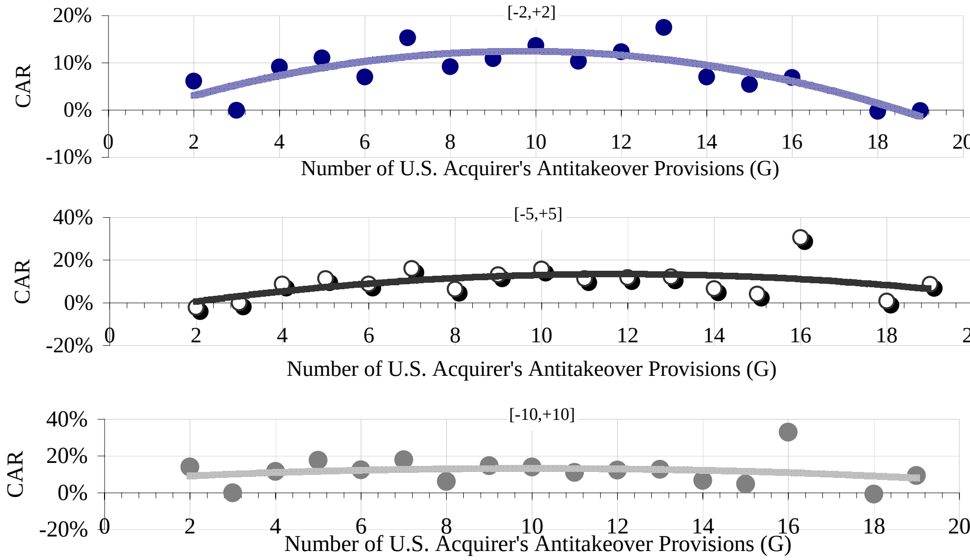
Panel A: U.S. domestic deals



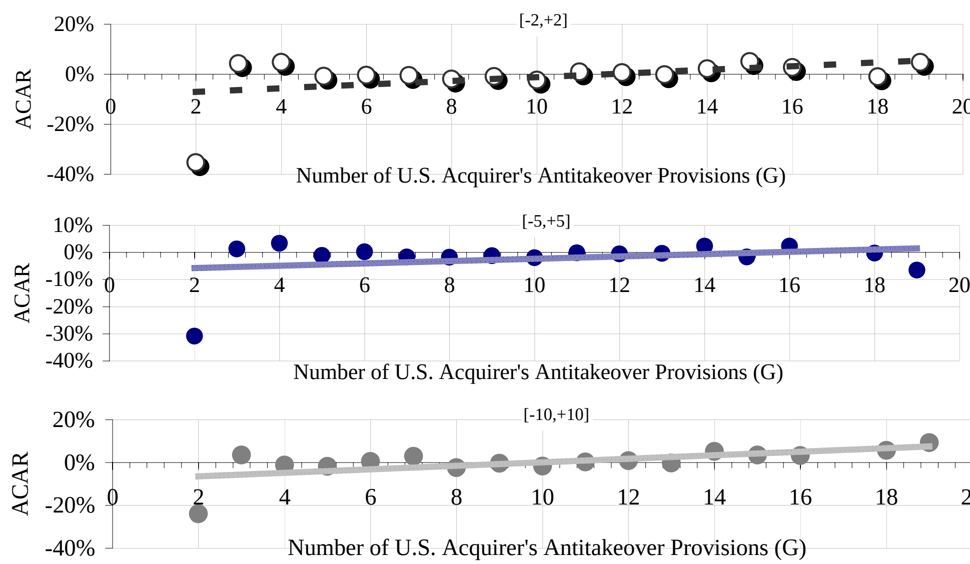
Panel B: U.S. cross-border deals

Figure B.1: Sample distribution by announcement year: U.S. domestic deals

This figures provide summary statistics of 1456 and 499 completed deals of U.S. acquirers' domestic and foreign takeovers, respectively, as given by SDC between 1990 and 2007. All firms are covered by the IRRC ATP database.



Panel A: Foreign target returns against U.S. acquirers' G Index



Panel B: U.S. acquirer returns against their G Index

Figure B.2: Cumulative abnormal return plots of cross-border targets and U.S. acquirers

Figures in Panel A plot foreign targets' CARs against their U.S. acquirers' *G* Index over three event windows, [-2,+2], [-5,+5], and [-10,+10], respectively. Figures in Panel B plot U.S. acquirers' CARs (ACARs) against their *G* Index over the same respective event periods. The *G* Index is the number of ATPs per Gompers *et al.* (2003).

APPENDIX C

CHAPTER 3 OF APPENDIX

C.1 Tables

Table C.1: Summary statistics of ADR pairs

When we assigned ADR's with U.S. indices, NYSE- and NASDAQ-listed ADR's were associated with Dow Jones Industrial Average and NASDAQ Composite indices, respectively, and Over-the-Counter (OTC) and 144A ADRs with S&P 500 index. Industry categorization information is obtained from Bank of New York's ADR website. "Hours Away from N.Y.C." is the positive number of time difference between the closing time of the home exchange and the opening time of host exchanges in New York. A related Timezone variable was first introduced by Gagnon and Karolyi (2005) which they obtained from the World Federation of Exchanges (http://www.world-exchanges.org) albeit they use the number of time zones between the home and host exchanges. Availability of shortsell information was obtained from Bris, Goetzmann, and Zhu (2003). MSCI (Morgan Stanley Capital International) market indices of developed versus emerging market is used as a proxy for market efficiency/completeness of the home markets in Asia-Pacific region. Papua New Guinea was omitted from the sample when we assigned dummies for both shortsell and MSCI categories.

Home country	Count	%
Australia	52	16.3%
China	19	5.9%
Hong Kong	64	20.0%
India	13	4.1%
Indonesia	3	0.9%
Japan	116	36.3%
Korea	8	2.5%
Malaysia	4	1.3%
New Zealand	2	0.6%
Pakistan	1	0.3%
Philippines	6	1.9%
Papua Nea Guinea	1	0.3%
Singapore	13	4.1%
Taiwan	11	3.4%
Thailand	7	2.2%
Total	320	100.0%

Associated U.S. index	Count	%
Dow Jones Industrial Average	59	18.4%
Nasdaq Composite	28	8.8%
S&P 500	233	72.8%
Total	320	100.0%

ADR exchange / listing type	Count	%
NASDAQ	28	8.8%
NYSE	59	18.4%
OTC	216	67.5%
Portal (144A)	17	5.3%
Total	320	100.0%

Industry	Count	%
Aerospace & Defense	1	0.3%
Automobiles & Parts	12	3.8%
Banks	21	6.6%
Beverages	6	1.9%
Chemicals	10	3.1%
Construction & Materials	12	3.8%
Electricity	6	1.9%
Electronic & Electric Equipment	21	6.6%
Fixed Line Telecommunication	12	3.8%
Food Producers	11	3.4%
Forestry & Paper	1	0.3%
Gas, H20 & Multiutility	1	0.3%
General Finance	12	3.8%
General Industrials	12	3.8%
General Retailers	6	1.9%
Health Care Equipment & Services	1	0.3%
Household Goods	4	1.3%
Industrial Engineering	14	4.4%
Industrial Metals	10	3.1%
Industrial Transportation	7	2.2%
Leisure Goods	13	4.1%
Media	7	2.2%
Mining	12	3.8%
Mobile Telecommunication	10	3.1%
Nonlife Insurance	2	0.6%
Oil & Gas Producers	9	2.8%
Personal Goods	7	2.2%
Pharmaceuticals & Biotechnology	8	2.5%
Real Estate	16	5.0%
Software & Computer Services	9	2.8%
Support Services	7	2.2%
Technological Hardware & Equipment	22	6.9%
Travel & Leisure	18	5.6%
Total	320	100.0%

Hours away from N.Y.C.	Count	%
4 - 5 hours	41	12.8%
6 - 7 hours	83	25.9%
8 - 9 hours	187	58.4%
10+ hours	9	2.8%
Total	320	100.0%

MSCI market index	Count	%
Emerging	58	18.2%
Developed	261	81.8%
Total	319	100.0%

Shortsell allowed / practiced	Count	%
No	49	15.4%
Yes	270	84.6%
Total	319	100.0%

Table C.2: Profiles of selected Asia-Pacific ADR-Pairs

When we assigned ADR's with U.S. indices, NYSE- and NASDAQ-listed ADR's were associated with Dow Jones Industrial Average and NASDAQ Composite indices, respectively, and Over-the-Counter (OTC) and 144A ADRs with S&P 500 index. Industry categorization and ADR-underlying exchange ratio information is obtained from Bank of New York's ADR website.

ADR DS code	Underlying Stock DS code	Company	Industry	Home country	Associated home market index	ADR exchange / listing type	Associated US index	ADR : underlying
510918	902201	Amcor	General Industrials	Australia	All Ordinaries	NASDAQ	Nasdaq Composite	1:4
133223	320394	Shanghai Tyre and Rubber	Automobiles & Parts	China	Shanghai Composite	OTC	S&P 500	1:10
133222	320359	Shanghai Erfangji	Industrial Engineering	China	Shanghai Composite	OTC	S&P 500	1:10
134334	140518	PT Indosat	Mobile Telecomm.	Indonesia	Jakarta Composite	NYSE	Dow Jones Indus. Avg.	1:50
154541	148190	DCW	Chemicals	India	S&P CNX 500	144A	S&P 500	1:5
25552J	938939	Advantest Corporation	Tech. Hardware & Equip.	Japan	Nikkei 225	NYSE	Dow Jones Indus. Avg.	1:1
14855H	14855F	Aluminum Corporation of China	Industrial Metals	China	Shanghai Composite	NYSE	Dow Jones Indus. Avg.	1:25
689000	316998	Hyundai Motor	Automobiles & Parts	Korea	KOSPI Composite	144A	S&P 500	2:1
151498	314092	Silverstone	Automobiles & Parts	Malaysia	KLSE Composite	OTC	S&P 500	1:1
151760	315187	Manila Electric Co. (Meralco)	Electricity	Philippines	Philippine SE All Shares	OTC	S&P 500	1:1
264386	133898	Pakistan Telecommunications	Fixed Line Telecomm.	Pakistan	Karachi 100	144A	S&P 500	1:1
132715	315053	Advanced Info Service	Mobile Telecom.	Thailand	SET Index	OTC	S&P 500	1:1
511768	280368	BIL International	Travel & Leisure	New Zealand	NZX 50	OTC	S&P 500	1:10
277293	691869	Chartered Semicond. Mfg.	Tech. Hardware & Equip.	Singapore	Straits Times	NASDAQ	Nasdaq Composite	1:10
25557E	264287	Au Optronics	Tech. Hardware & Equip.	Taiwan	Taiwan Weighted	NYSE	Dow Jones Indus. Avg.	1:10

Table C.3: Asia-Pacific Securities Exchanges

The overall information here is available at the World Fededation of Exchanges website (http://www.world-exchanges.org). "Hours Away from N.Y.C." is the positive number of time difference between the closing time of the home exchange and the opening time of host exchanges in New York. A related Timezone variable was first introduced by Gagnon and Karolyi (2005) which they obtained from the World Federation of Exchanges albeit they use the number of time zones between the home and host exchanges. Availability of shortsell information was obtained from Bris, Goetzmann, and Zhu (2003). MSCI (Morgan Stanley Capital International) market indices of developed versus emerging market is used as a proxy for market efficiency/completeness of the home markets in Asia-Pacific region.

Home country	Foreign exchange regime	Exchange	Hours away from N.Y.C.	Shortsell allowed / practiced	MSCI market index	Stock index	Stock options	Stock index options	Stock futures	Stock index futures
Australia	Fully float	Australian Securities Exchange	9:30	Yes	Developed	All Ordinaries	Yes	Yes	Yes	Yes
HK	Pegged	Hong Kong Exchanges and Clearing	6:30	Yes	Developed	Hang Seng	Yes	Yes	Yes	Yes
Indonesia	Managed float	Jakarta Stock Exchange	5:30	No	Emerging	Jakarta Composite	No	No	No	No
Japan	Fully float	Tokyo Stock Exchange	8:30	Yes	Developed	Nikkei 225	Yes	Yes	No	Yes
		Osaka Securities Exchange	8:30	No	N/A	N/A	Yes	Yes	No	Yes
Malaysia	Managed float	Bursa Malaysia	5:30	Yes	Emerging	KLSE Composite	No	No	No	Yes
New Zealand	Fully float	New Zealand Exchange	11:00	No	Developed	NZX 50	No	No	No	No
Pakistan	Managed float	Karachi Stock Exchange	4:15	No	Emerging	Karachi 100	No	No	No	No
Papua New Guinea		Port Moresby Stock Exchange	15:00	No	N/A	N/A	No	No	No	No
Philippines	Fully float	Philippine Stock Exchane	10:20	No	Emerging	Philippine SE All Shares	No	No	No	No
Singapore	Managed float	Singapore Exchange	5:24	Yes	Developed	Straits Times	No	Yes	No	Yes
Sri Lanka	Managed float	Colombo Stock Exchange	5:30	No	Emerging	All Share	No	No	No	No
Taiwan	Fully float	Taiwan Stock Exchange	9:00	No	Emerging	Taiwan Weighted	No	No	No	No
		Taiwan Futures Exchange		No	N/A	N/A	Yes	Yes	No	Yes
Thailand	Fully float	Stock Exchange of Thailand	5:00	Yes	Emerging	SET Index	No	No	No	No
China	Pegged	Shanghai Stock Exchange	7:30	No	Emerging	Shanghai Composite	No	No	No	No
		Shenzhen Stock Exchange	7:30	No	Emerging	SSE Component	No	No	No	No
India	Fully float	Bombay Stock Exchange	4:30	No	Emerging	India BSE National	No	Yes	No	No
		National Stock Exchange of India	4:30	No	Emerging	S&P CNX 500	Yes	Yes	Yes	Yes
Korea	Fully float	Korea Exchange	8:30	No	Emerging	KOSPI Composite	Yes	Yes	No	Yes
		KOSDAQ	8:30	No	Emerging	KOSDAQ Composite	No	No	No	No

Table C.4: $H_0 : \theta_{\text{later}}(t) = \theta_{\text{earlier}}(t)$ $vs.$ $H_1 : \theta_{\text{later}}(t) > \theta_{\text{earlier}}(t)$

The vertical labels are of later closings and the horizontal labels earlier.

		6-7 hours	8-9 hours	10-11 hours
4-5 hours	t_0	32.246**	57.396**	44.534**
	V_0	1326	1326	1326
	p-value	2.651×10^{-10}	2.651×10^{-10}	2.651×10^{-10}
6-7 hours	t_0		25.189**	18.320**
	V_0		1326	1326
	p-value		2.651×10^{-10}	2.651×10^{-10}
8-9 hours	t_0			7.078**
	V_0			1188
	p-value			4.408×10^{-7}

C.2 Figures

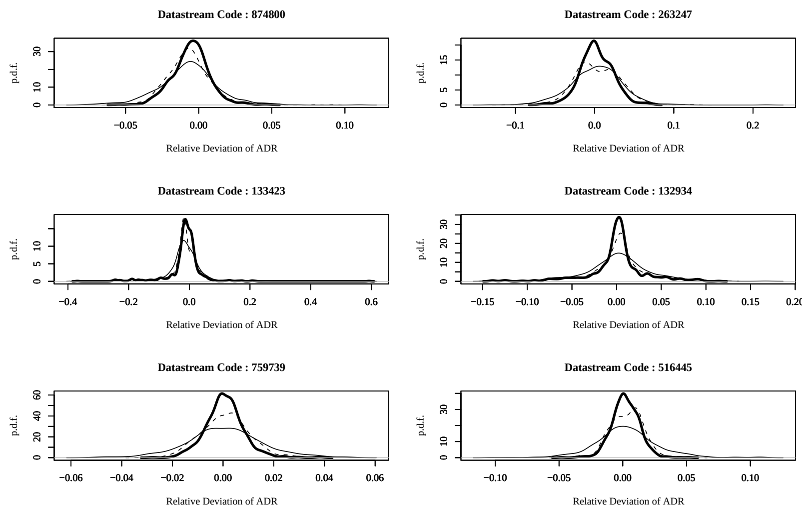


Figure C.1: Percentage deviation from implied prices of ADR-pairs—kernel density plots

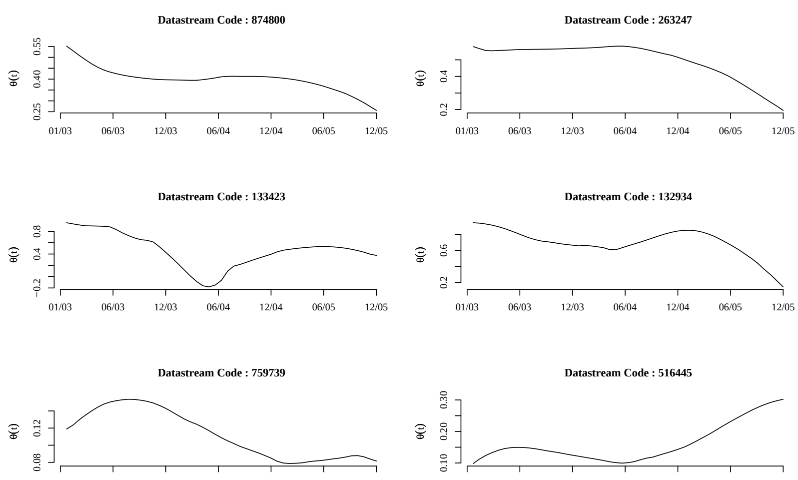


Figure C.2: Convergence speed parameters of ADR-pairs

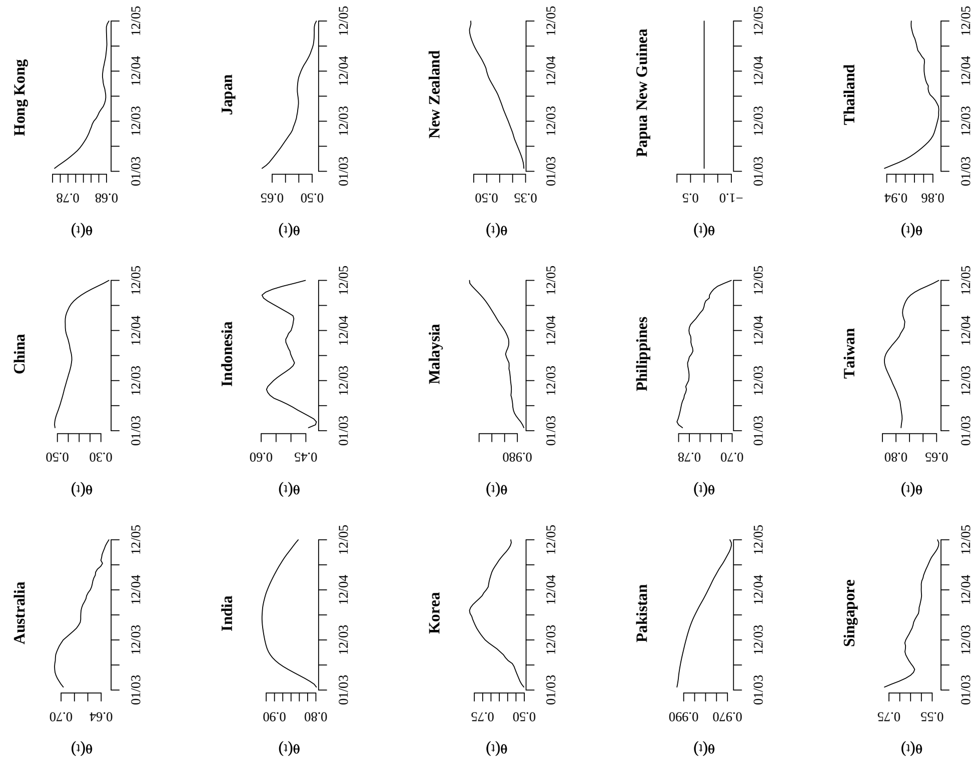


Figure C.3: Convergence speed by home countries

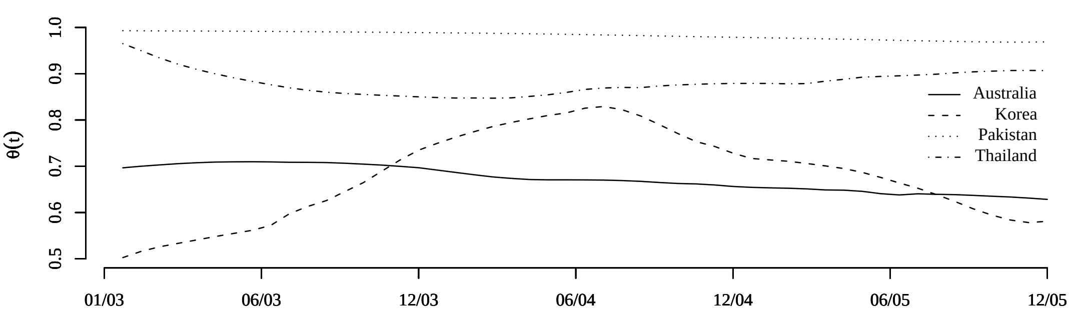


Figure C.4: Comparative dynamics of convergence speed

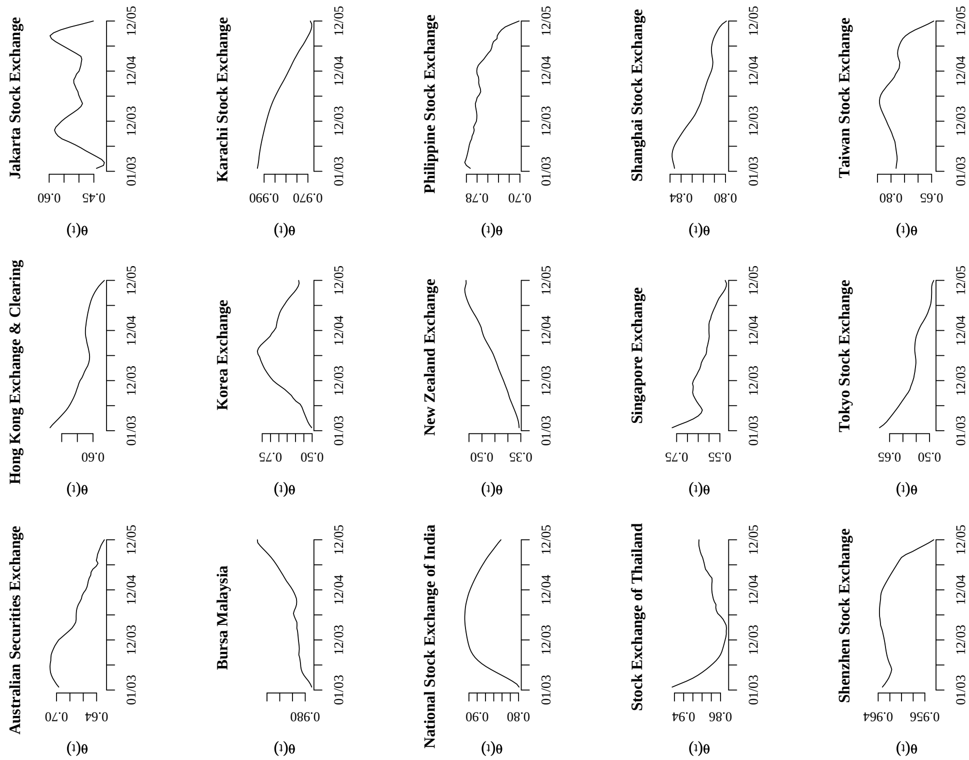


Figure C.5: Convergence speed by home exchanges

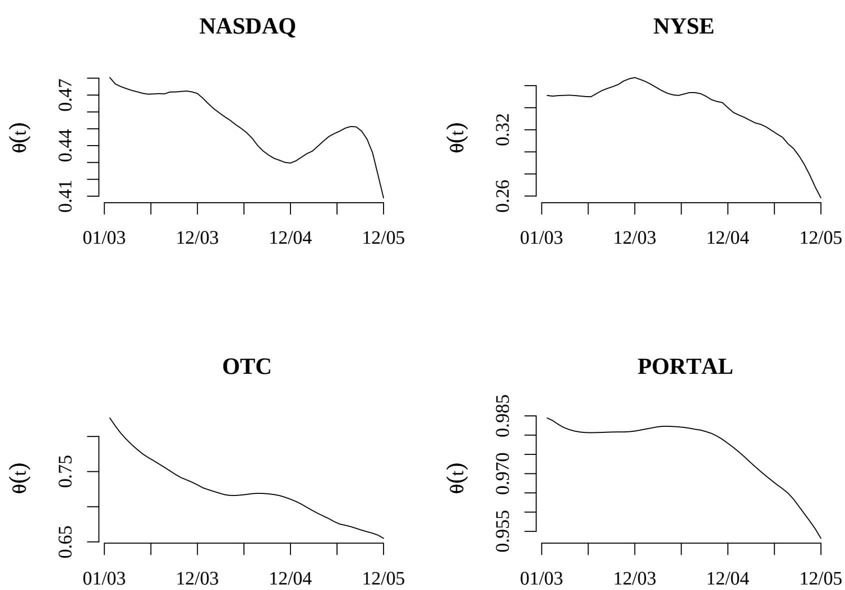


Figure C.6: Convergence speed by various listing types in the U.S.

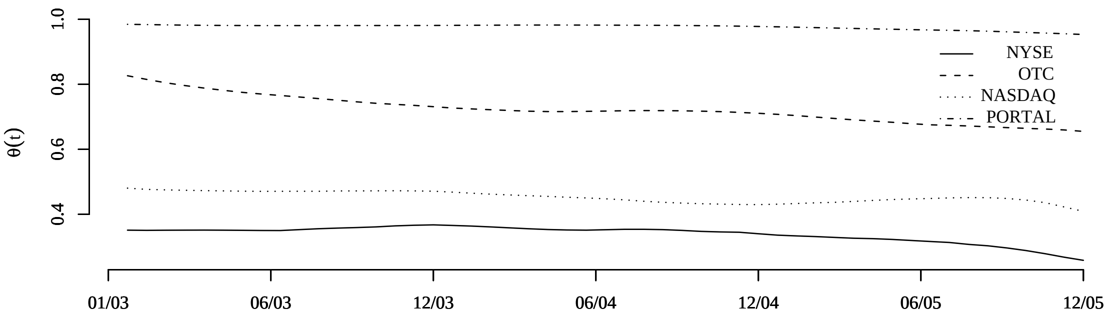


Figure C.7: Comparative dynamics of convergence speed by listing types

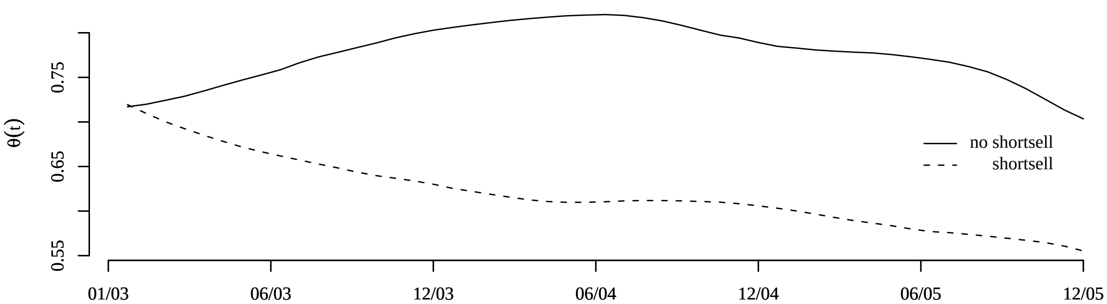


Figure C.8: Comparative dynamics with respect to short sales

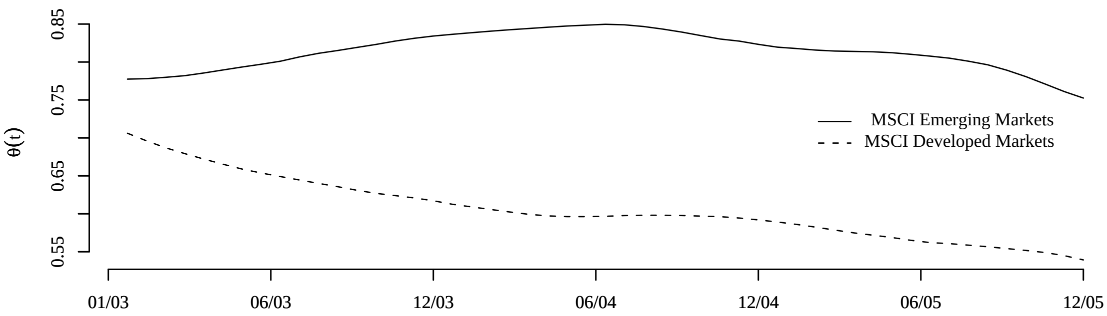


Figure C.9: Comparative dynamics with respect to MSCI indices

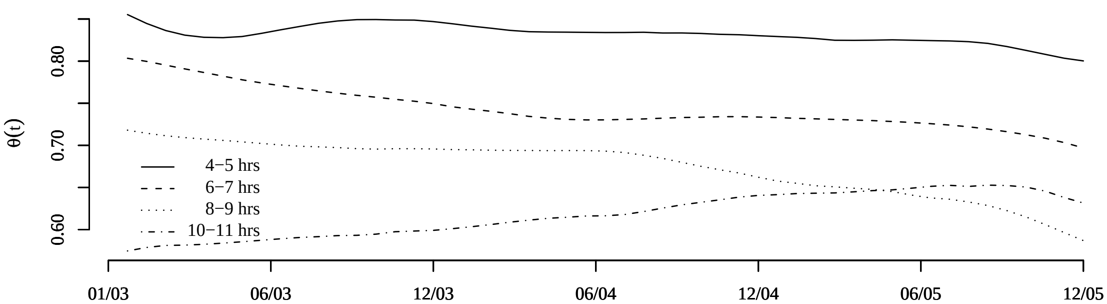


Figure C.10: Comparative dynamics with respect to time differences

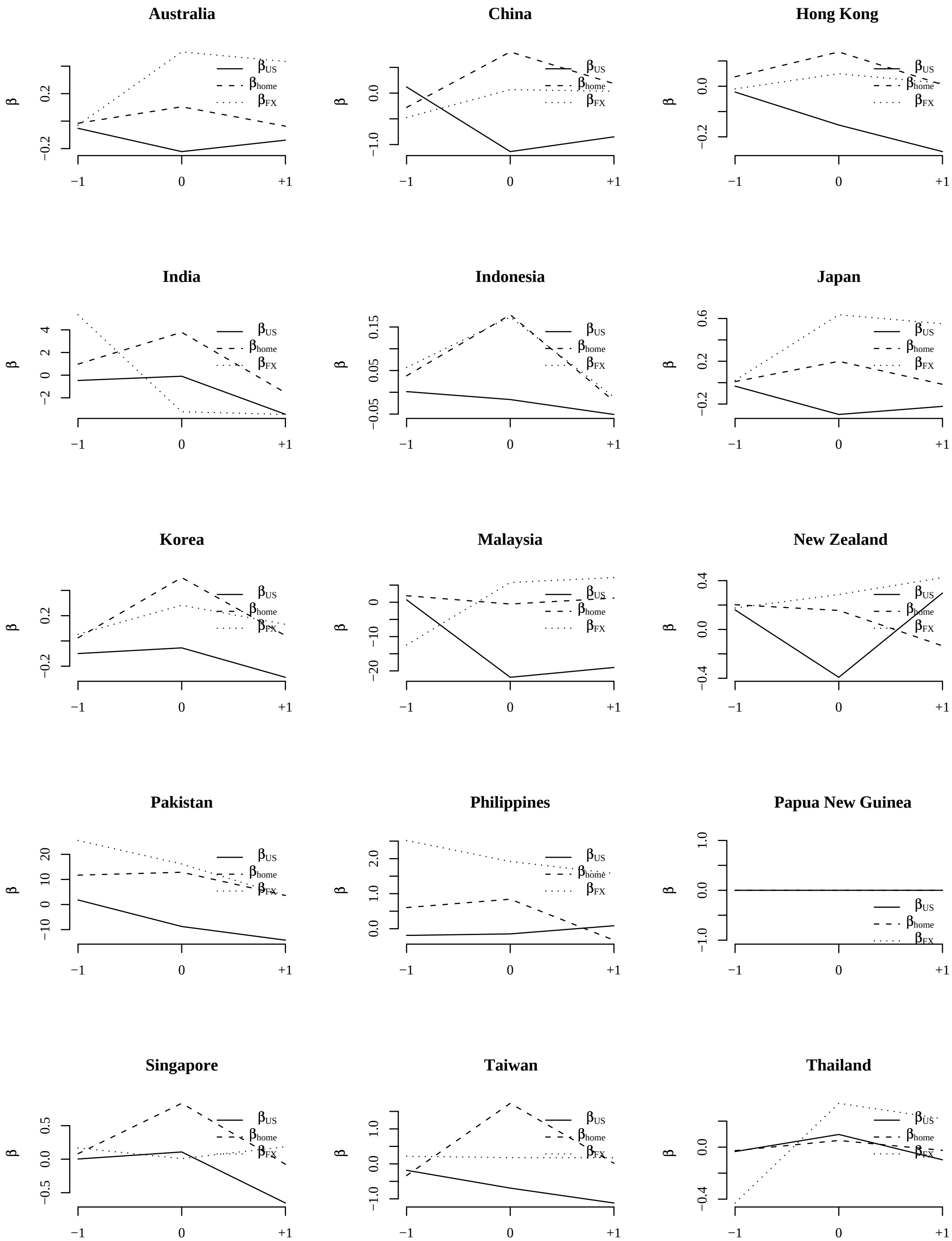


Figure C.11: Betas of home countries

C.3 Nonparametric varying-coefficient model

Consider the following varying-coefficient model: $y_{it} = X_{it}\beta_i(z_{it}) + \varepsilon_{it} \quad \forall\, (i,t) \in [1,...,N] \times [1,...,T]$, where (1) y_{it} is the dependent variable and X_{it} is a regressor variable or vector; (2) $\beta_i(z_{it})$ is a coefficient function on the smoothing variable z_{it}; and (4) ε_{it} is the error term that follows a martingale difference sequence adapted to the information algebra $\mathcal{F}_t = \sigma(X_{it}, z_{it})$. Our objective is to estimate the coefficient function $\beta_i(z_{it})$ based on data $\{(y_{it},\, X_{it},\, z_{it})\,|(i,t) \in [1,...,N] \times [1,...,T]\}$.

Writing the model in compact form, for each i, we have $Y_i = X_i\beta_i(z_i) + \varepsilon_i$, where (1) $Y_i = (y_{i1}, \cdots, y_{iT})^T$ is a $T \times 1$ matrix; and (2) X is a $T \times q$ matrix. Using the local linear kernel estimation method,[1] we have $\widehat{\beta_i}(z) = (X_i^T W_i X_i)^{-1} X_i^T W_i Y_i$, where (1) $K(\cdot) \equiv \mathbf{1}_{\{|(\cdot)|<1\}} \cdot \left\{1 - (\cdot)^2\right\}$ is a kernel of our choice; and (2) W_i is the weighting matrix based on $K((z_{it} - z)/h)$.

An important extension of the above model is to allow some regressors to behave linearly with the dependent variable. In this case, we arrive to a partially varying-coefficient linear model , studied by Fan and Huang (2005). The model is defined as follows: $Y = Xa(U) + Z\beta + \varepsilon$, where (1) Y is the response variable; (2) (U, X, Z) are the covariates; (3) ε is a martingale difference sequence adapted to $\mathcal{F}_t = \sigma(U, X, Z)$; (4) $a(U)$ is an unknown coefficient functions; and (5) β is a constant parameter.

Fan and Huang (2005) provide a breakthrough in estimating the parameters and the coefficient functions: $\widehat{\beta} = [Z^T(I - S)^T(I - S)Z]^{-1} Z^T(I - S)^T(I - S)Y$, where (1) $W_{u_T} = \operatorname{diag}\{K_h(u_1 - u), \cdots, K_h(u_T - u)\}_{\,T\times T}$;

[1]See Fan and Yao (2003).

$$(2)\ D_u = \begin{bmatrix} X_1^T, \frac{u_1-u}{h}X_1^T \\ \vdots \\ X_T^T, \frac{u_T-u}{h}X_T^T \end{bmatrix}_{T\times 2} \text{; and (3) } S = \begin{bmatrix} [X_1^T, 0]\{D_{u_1}^T W_{u_1} D_{u_1}\}^{-1} D_{u_1}^T W_{u_1} \\ \vdots \\ [X_T^T, 0]\{D_{u_T}^T W_{u_T} D_{u_T}\}^{-1} D_{u_T}^T W_{u_T} \end{bmatrix}_{T\times T.}$$

After obtaining $\widehat{\beta}$, we can conduct a nonparametric estimation for the coefficient function $a(U)$ as

$$\widehat{a}(U) = \begin{bmatrix} [1, 0]\{D_{u_1}^T W_{u_1} D_{u_1}\}^{-1} D_{u_1}^T W_{u_1} \\ \vdots \\ [1, 0]\{D_{u_T}^T W_{u_T} D_{u_T}\}^{-1} D_{u_T}^T W_{u_T} \end{bmatrix}_{T\times T} \left(Y - Z\widehat{\beta}\right)_{T\times 1.}$$

Finally, in order to alleviate the model error in the parametric specification, our nonparametric extension[2] from Gagnon and Karolyi (2004) is

$$\widetilde{DR}_i(t) = \alpha_i + \theta_i(t)\widetilde{DR}_i(t-1) + \sum_{j=-1}^{1}\beta_j^{US} R_M^{US}(t+j) + \sum_{j=-1}^{1}\beta_j^{H} R_M^{H}(t+j) + \sum_{j=-1}^{1}\beta_j^{FX} R_{FX}(t+j) + \varepsilon_i(t),$$

where (1) $\theta_i(t)$ is the firm-specific intra-day *convergence speed parameter* which varies over time; the higher the absolute value of the parameter to one, the lower the convergence speed; (2) $\widetilde{DR}_i(t) \equiv \left\{\frac{P_i^{US}(t-\tilde{j})-P_i^H(t)}{P_i^H(t)} \ :\ \tilde{j} = \arg\min_j \left\{\left|\frac{P_i^{US}(t-j)-P_i^H(t)}{P_i^H(t)}\right| \ :\ j = 0, 1\right\}\right\}$ is the minimum of the relative price deviations of the same- and previous-calendar days ADRs $\left(P^{US}\right)$ from its underlying share $\left(P^H\right)$ in the home market, adjusted for the host-home exchange ratio and the foreign exchange rate (see Figure C.12)[3]; (3) $R_M^{US}(t+j)$ is the return on the U.S. market index; (4) $R_M^H(t+j)$ is the return on the home market

[2]In the sprit of "*Let the data reveal the fact...*," it is the first-known attempt in the cross-listed shares literature.

[3]A conventional definition of relative spread would be $\mathrm{DR}_i(t) \equiv \frac{P_i^{US}(t)-P_i^H(t)}{P_i^H(t)}$ following Gagnon and Karolyi (2004). $\widetilde{DR}_i(t)$ as defined above seeks to better capture the dynamic convergence towards the parity. The reason for bringing the minimum spread definition is that the asynchronous trading between the Asia-Pacific exchanges and the U.S. host exchanges will further confound the true convergence of an ADR towards its implied value. Our definition may be needless in synchronous trades as on Canadian and Mexican cross-listed shares.

index; (5) $R_{FX}(t+j)$ is the return on foreign exchange rate at time $t+j$ with exchange rates defined per U.S. dollar, that is a positive R_{FX} implies a depreciation in the home currency of the ADR.

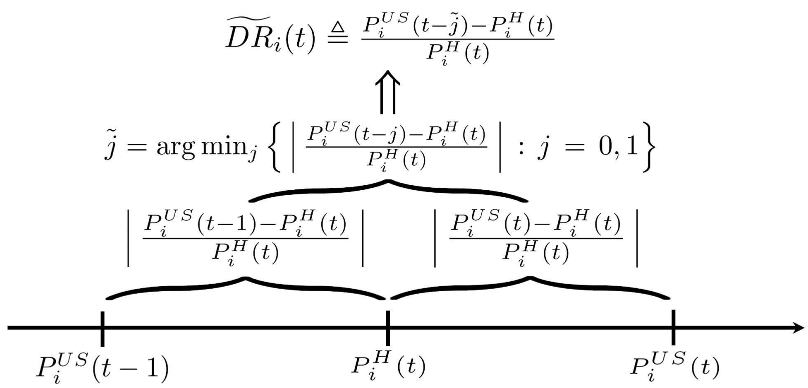


Figure C.12: The definition of minimum relative spread $\widetilde{DR}_i(t)$

C.4 Hypothesis testing methods

When we test the significance of the difference between two regime-dependent nonparametric estimates of time-varying parameters such that: H_0 : $\theta_{\text{regime }1}(t) = \theta_{\text{regime }2}(t)$ $vs.$ H_1 : $\theta_{\text{regime }1}(t) \neq \theta_{\text{regime }2}(t)$, our procedure is as follows: (1) *obtain the difference over the time period,* $d(t) \equiv \widehat{\theta}_{\text{regime }1}(t) - \widehat{\theta}_{\text{regime }2}(t) \quad \forall\, t \in$ $[01/01/03, 12/31/05]$, where $\widehat{\theta}_{\text{regime }j}(t) \equiv \sum_{\{i|\text{regime }j\}} \widehat{\theta}_{i|\text{regime }j} / |\{i|\text{regime }j\}|$; (2) *calculate the standard error of the difference,* $\text{s.e.}\,(d(t)) \equiv \sigma\,(d(t)) \,/\, \sqrt{\#\,(\{t\})}$; and (3) *test the null t-statistic at a given significance level (α),* $t_0 \equiv \left(\overline{d(t)} - 0\right)/\text{s.e.}\,(d(t)) \;\sim\; t_\alpha\,(df)$.

We also use the *Wilcoxon signed-rank test*[4] as an alternate which is a nonparametric test to compare the two samples. The estimators are assumed to follow the same distribution if they are the results of a nonparametric method. The *Wilcoxon*-test statistic is defined as $V_0 \equiv \sum_{\{t\}} \mathbf{1}_{\{d(t)>0\}} \cdot \rho_t$, where ρ_t is the rank of $\{|d(t)|\}_{\{t\}}$.

[4]See Wilcoxon (1945).

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