Price discovery and threshold cointegration: Theory and empirics on cross-border stock trading

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Abstract

The adjustment to parity can be discontinuous for an original listing and its cross-listing: convergence may be quicker when the price deviation is sufficiently profitable, but otherwise slower. In other words, the dynamics of premiums on cross-listings fall into two regimes: within and beyond the threshold, i.e. the transaction costs and risk premiums of arbitrage. We complement Harris et al.'s (1995, 2002) linear error correction model to estimate the relative extent of market-respective contribution to price discovery (information share) of cross-listed pairs on the New York Stock Exchange (NYSE) and the Toronto Stock Exchange by considering threshold cointegration per Balke and Fomby (1997). Our beyond-threshold (outer-regime) information shares suggest an increasing influence by the NYSE on Canadian stocks over time. We find that the estimated outer-regime information shares and thresholds are typically affected by the relative degree of private information, market friction, and liquidity measures, and idiosyncratic firm-level characteristics.

JEL Classification: C32; G15; G14

Keywords: Price discovery; Information share; Threshold cointegration; Error correction model

1. Introduction

We contribute to the literature by implementing the threshold error correction mechanism in estimating the relative extent of exchange-respective contribution to price discovery of the pairs of cross-listings and their original listings. The existing methods assume linear convergence of price deviations[1] to parity whereas we hinge our premise on the reality that the premiums disappear quicker when it is profitably arbitrageable than otherwise.

Price discovery is a search for an equilibrium price (Schreiber and Schwartz, 1986) and is a key function of a securities exchange. It is the process by which information is priced in the market. When a security is traded in multiple markets, it is often of interest to determine where and how price discovery occurs. Harris et al. (1995) and Hasbrouck (1995) examine the exchange-specific relative contribution to price discovery of fragmented stocks on the NYSE and other U.S. exchanges, and confirm the leadership assumed by the NYSE. As for international cross-listing, Bacidore and Sofianos (2002) and Solnik et al. (1996) suggest that price discovery mostly takes place in the home market where substantial information originates. Eun and Sabherwal (2003) report the U.S. host exchanges determine the prices of Canadian cross-listings, however, to a lesser extent than the Toronto Stock Exchange (TSX) does.

In the literature, there are two broad approaches to estimating the contribution of each market to price discovery of fragmented listings. Hasbrouck's (1995) *innovation variance approach* extracts the information shares by employing variance decomposition based on the vector moving average representation of an error correction model (ECM). Harris et al.'s (1995,

[1]We define the "relative premium" as the percentage premium of a cross-listed stock traded on a foreign exchange against the home market share, adjusted by the exchange rate. 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.

2002) *common factor approach* employs permanent-transitory decomposition of a cointegrated system to estimate the information share of each market. As Eun and Sabherwal (2003) point out, Hasbrouck's (1995) approach involves Cholesky factorization of the covariance matrix of the innovations to prices on various exchanges and yields multiple information shares. This may cause confounding identification of the venue of price discovery. Hasbrouck's (2002) modification can be numerically onerous in implementation.[2] In this paper, we expand Harris et al.'s (1995, 2002) platform and complement Hasbrouck's (1995) idea.

Harris et al. (1995) associate error correction dynamics with price discovery of cross-listed pairs which are cointegrated[3] by the law of one price. The cointegrating vectors of the vector ECM (VECM) represent the long-run equilibrium (near-parity condition), while the error correction terms characterize the convergence mechanism, i.e. "the process whereby markets attempt to find equilibrium." Through representation, we can assess the relative extent of the contribution made by each market to price discovery of fragmented stocks using the estimates of adjustment coefficients. If the price of a Canadian cross-listing on the NYSE responds sensitively to shocks from the TSX whereas the home exchange is largely unaffected by the ripples occurring in the host market, price discovery can be deemed as predominantly taking place on the TSX. Harris et al. (2002) buttress the method earlier formulated in Harris et al. (1995) by incorporating a microstructure model where the price is assumed to be the sum of an efficient (permanent) price component and an error (transitory) term.[4]

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[2]See De Jong (2002), Harris et al. (2002), and Hasbrouck (2002) for further discussion.

[3]A group of multiple random-walk processes is cointegrated if, by definition, there exists a stationary linear combination of the processes. A time series is (weakly) stationary if the probability laws (of up to the second moments) are time-invariant.

[4]In Harris et al. (2002), the efficient price component is unobservable and reflects the underlying fundamentals. Gonzalo and Granger's (1995) permanent-transitory decomposition posits the permanent price as a linear

However, an implicit assumption made by Harris et al.'s (1995, 2002) works is that adjustment to parity, the long-run equilibrium, is continuous and linear. Various economic circumstances challenge such restrictions, particularly where transaction costs and policy intervention are present. Given the complexity of trading rules and indirect transaction costs, nonlinear convergence to parity captures the market to a higher proximity. The rationale of nonlinear modeling is straightforward. A relatively small deviation of the price of an American Depositary Receipt (ADR) from its parity-implied price can be unarbitrageable if the dollar spread is insufficient to cover the fees, commissions, liquidity shortfalls, and other related costs. In this case, the dollar premium or discount behaves like a near-unit root process and will not converge to parity. Arbitrage forces will activate as the spread widens beyond the "threshold."

To date, we find a dearth of articles with a nonlinear framework in the literature. Among them, Rabinovitch et al. (2003) use a nonlinear threshold model to estimate the adjustment dynamics of the return deviations for 20 Chilean and Argentine cross-listings. Koumkwa and Susmel (2008) suggest two nonlinear adjustment models: the exponential smooth transition autoregressive (ESTAR) and the logarithmic smooth transition autoregressive (LSTAR) to delineate the relative premiums of Mexican ADRs. Chung et al. (2005) study the dynamic relationship between the prices of three Taiwanese ADRs and their underlying stocks using a threshold VECM.

For a cross-listed pair, convergence to parity may be quicker when the price spread is profitable, or slower otherwise. In other words, the dynamics of cross-listed pairs fall into two

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combination of the observable prices where the normalized weights can be as market-respective information shares. The higher the normalized weight of an exchange, the bigger the influence on setting the permanent price. It can be shown that the normalized weights are orthogonal to the adjustment coefficient vector, which can be conveniently obtained from an ECM.

regimes: within versus beyond the threshold, which is determined by transaction costs and associated risk premiums of arbitrage. In this regard, we extend Harris et al.'s (1995, 2002) ECM to estimate exchange-respective information shares of Canadian cross-listed pairs traded on the NYSE and the TSX by considering threshold cointegration per Balke and Fomby (1997). Departing from linear modeling, our information share is estimated from the outer-regime adjustment coefficients based on a two-regime threshold ECM.

Our method has many advantages. To list a few of them, first, we theoretically depict and empirically analyze the discrete dynamics of "bumpy" parity-convergences, which are frequently observed in the market due to various risk factors such as information asymmetry and market friction. Second, a large deviation from parity far beyond the threshold (extreme regime), e.g. a very profitable arbitrage opportunity, is more likely to reflect information shocks than a small deviation, which can be due to noise trading. Thus, we believe our method captures relative contribution to price discovery to an enhanced degree compared to the existing linear approaches in the literature, which circumvent the time- and regime-contingent characteristics of information shares.

In addition, we identify and explicate the factors that affect the estimated information share and threshold, and find that they are typically determined by the relative degree of private information, market friction and liquidity measures, and idiosyncratic firm-level characteristics. The remainder of this paper is organized as follows. We provide a theoretical asset pricing model for cross-listings under information asymmetry, with price discovery implications, in Section 2. Section 3 summarizes the standard ECM and proposes our threshold ECM for price discovery of cross-listed pairs. Section 4 describes the data. Discussion of the main estimation results and multivariate regression analyses are presented in Section 5. We conclude in Section 6.

2. Information asymmetry and pricing of cross-listings

In this section, we present a model to illustrate how information asymmetry affects equilibrium prices when a stock cross-lists on a foreign exchange. Chen and Choi (2010) previously extended the noisy rational expectations model of Grossman and Stiglitz (1980) in the following ways: firstly, a stock is fragmented across two cross-border exchanges; secondly another type of market participants, arbitrageurs, is added to the model.[5] We expand Chen and Choi's (2010) model by incorporating a quadratic cost function assumed for all investors so that the market is not friction-free, and by extending their model to a dynamic setup to embrace multi-market price discovery.

Chen and Choi's (2010) model is also closely related to that of Chan et al. (2008), which explains the persistent discounts in Chinese B-shares. However, in their model, A and B-share markets are completely separated: domestic investors trade in the A-share market and foreign investors trade in the B-share market. There are no arbitrageurs in their framework and all foreign investors are assumed to be uninformed traders. Given the background of the Chinese markets during their sample period, these assumptions may be valid. In comparison, Chen and Choi (2010) designs a model that prices cross-listings: Candian stocks that trade simultaneously on the NYSE and the TSX; and these two markets are not completely separated. Moreover, there are informed traders not only on the TSX but also on the NYSE given its important role in the global financial markets.

In our model, we also emphasize the role of arbitrageurs in pricing of the cross-listed stocks.

[5]Chen and Choi (2010) extend Grossman and Stiglitz's (1980) noisy rational expectations model to show that in no-arbitrage equilibrium, under certain conditions, the premium of a Canadian NYSE-listing against its original TSX-listing is due to the informational dominance of the investors on the TSX compared to those on the NYSE. Easley et al.'s (1986) probability of informed trading (PIN) measures the relative degree of private information on a security by estimating the proportion of information-based trades among all trades of the stock.

Our solution of the general equilibrium prices shows that supply shocks in one market can spillover to the other side by cross-border arbitrage ("arbitrage effect"). To some extent, arbitrageurs carry risk across the border, thereby reducing supply shock risk (volatility) in one market and increasing it in the other. In addition, arbitrageurs play a role in information propagation: they can transfer some information from the market with a higher degree of private information to the lower. We expect that the price deviation is negatively related to the intensity of arbitrage activity.

We add dynamics to the static equilibrium model introduced by Chen and Choi (2010), then derive cointegration between the prices of a given cross-listed pair. Based on an implied ECM, we use permanent transitory decomposition per Gonzalo and Granger (1995) to extract the relative contribution (information share) of each market to price discovery. We expect that the information share is not only directly related to the market's relative liquidity, but also to a relative degree of information asymmetry, which can be empirically proxied for by the PIN.

2.1. The model

In order to faciliate trades of the original listing and its cross-listing, there are two stock exchanges: the TSX and the NYSE. We conveniently use $i = 1, 2$ to index the respective market. There are three types of market participants: informed and uninformed traders, and arbitrageurs. Informed and uninformed traders always trade, whereas arbitrageurs only emerge when arbitrage opportunities present themselves. We further assume that there are N_1 and N_2 market participants who only trade on the TSX and the NYSE, respectively, and N_3 arbitrageurs who trade in both markets.

Following Grossman and Stiglitz (1980), our one-market traders are of two types: informed versus uninformed with respective proportions π_i and $(1 - \pi_i)$ in market $i = 1, 2$. The future payoff (υ) of the risky asset (stock) is uncertain: $\upsilon \sim N(\overline{\upsilon}, \sigma_\upsilon^2)$. Informed traders recognize a signal S about υ with random noise: $\varepsilon_s \sim N(0, \sigma_s^2)$, such that $S = \upsilon + \varepsilon_s$. All variances are expressed in precision terms in the following discussion: $\tau_\upsilon \equiv 1/\sigma_\upsilon^2$ and $\tau_s \equiv 1/\sigma_s^2$. 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. Each individual can borrow at the risk-free rate (r) to purchase the risky asset.

The budget constraint of the model is as follow: in each market i, in the beginning, trader j is endowed with 0 shares, and the exchange-specific aggregate supply of shares is Y_i, where $\Delta Y_i \sim N(0, \sigma_i^2)$ is a random net supply from noise traders. All traders share the same constant relative risk aversion (CARA) utility function with a risk aversion coefficient (ρ) or a risk tolerance parameter ($\eta \equiv 1/\rho$):

$$V(W) = e^{-\rho W}, \rho > 0. \tag{1}$$

Since all random variables are assumed to be normally distributed, so is the wealth (W). With the CARA utility function, the investor's objective function can be written as

$$\mathbb{E}(W|\Phi) - \frac{\rho}{2} Var(W|\Phi), \tag{2}$$

where Φ is the information set of the trader.

The market is not friction-free, thus, we introducet a quadratic transaction-cost function:

$$C(x) = \frac{cx^2}{2} \tag{3}$$

where x is the amount of risky asset. To begin with, we suppose that the two markets share the same ratio of transaction cost c.

2.1.1. One-market traders

We first consider the demand function of the one-market traders in each market. For informed traders, they update their belief on the future payoff based on private signals. We denote surprise in earnings signal as $\Delta S \equiv S - \overline{S}$. The prices of a given cross-listed pair are bullish on a positive earnings shock ($\Delta S > 0$). We assume informed traders in both markets observe the same private signal S regardless of their location. Upon receiving a new earnings signal, their updated (posterior) earnings forecast ($\mathbb{E}(v|S)$) and updated earnings forecast precision ($\tau(v|S)$) are given by

$$\mathbb{E}(v|S) = \overline{v} + \left(\frac{\tau_s}{\tau_s+\tau_v}\right)\Delta S, \tag{4}$$

$$Var(v|S) = \frac{1}{\tau_s+\tau_v}. \tag{5}$$

Under the CARA utility function assumption, exchange-specific informed traders j's ($j = 1,2,\cdots,N_i$) demand for shares in the market i is

$$\begin{aligned} x_{i,j}^I(p_i,S) &= \frac{\mathbb{E}(v|S) - p_i(1+r)}{\rho Var(v|S) + c} \\ &= \left\{\overline{v} + \left(\frac{\tau_s}{\tau_s+\tau_v}\right)\Delta S - p_i(1+r)\right\}\Big/\left(\frac{\rho}{\tau_s+\tau_v} + c\right) \\ &= \beta^I(\mu^I - p_i). \end{aligned} \tag{6}$$

where $\beta^I = (1+r)/\left(\frac{\rho}{\tau_s+\tau_v} + c\right)$ is the elasticity of demand and $\mu^I = \left\{\overline{v} + \left(\frac{\tau_s}{\tau_s+\tau_v}\right)\Delta S\right\}/(1+r)$ is the reservation price of the informed trader. Since we assume all informed traders are homogeneous except for the endowment, and the demand function does not depend on the endowment, both β^I and μ^I are the same for all informed traders in both markets.

We now consider the demand function for uninformed traders who observe prices on their respective exchanges and form their expectations of future earnings. Let p_1 and p_2 be the equilibrium prices in markets 1 and 2. We postulate that P_1 and P_2 are linear related to the observables at each time such that

$$p_i = \alpha_{i0} + \alpha_i^S \Delta S - \alpha_{ii}^Y \Delta Y_i - \alpha_{i-i}^Y \Delta Y_{-i}. \tag{7}$$

In the ensuing analysis, we verify that this conjecture is consistent with the equilibrium outcomes we derive. Compared to the original model of Grossman and Stiglitz (1980), our formulation has additional terms, $\alpha_{12}^Y \Delta Y_2$ and $\alpha_{21}^Y \Delta Y_1$, which reflect reponses to the supply shocks from the counterpart market. For example, if a negative supply shock occures in market 2 then it decreases p_2, and the negative price effect can be transferred into market 1 since arbitragers will buy the shares in market 2 and shortsell the same number of shares in market 1. These simulataneous transactions can reduce supply shock risk (volatility) in one market but increase it in the other.

The price function is not fully revealed to uninformed traders due to existence of unobervervable supply shocks and earnings surprises. We assume uninformed traders extract information from the price on their respective exchanges only, which is reasonable since uninformed investors cannot tell the informativeness of prices so they only refer to the familiar listings. Uninformed traders' price-contingent updated (posterior) payoff forecast ($\mathbb{E}(v|p_i)$), updated payoff precision ($\tau(v|p_i)$) and demand functions are, respectively, given by

$$\begin{aligned}
\mathbb{E}(v|p_i) &= \mathbb{E}(v) + \frac{Cov(v, p_i)}{Var(p_i)}(p_i - \mathbb{E}(p_i)) \\
&= \bar{v} + \left\{\frac{Cov(v, \alpha_{i0} + \alpha_i^S \Delta S - \alpha_{ii}^Y \Delta Y_i - \alpha_{i-i}^Y \Delta Y_{-i})}{Var(\alpha_{i0} + \alpha_i^S \Delta S - \alpha_{ii}^Y \Delta Y_i - \alpha_{i-i}^Y \Delta Y_{-i})}\right\} \Delta p_i \\
&= \bar{v} + \left\{\frac{\alpha_i^S (1/\tau_v)}{(\alpha_i^S)^2 (1/\tau_s + 1/\tau_v) + (\alpha_{i1}^Y)^2/\tau_1 + (\alpha_{i2}^Y)^2/\tau_2}\right\} \Delta p_i \\
&= \bar{v} + \frac{1}{\alpha_i^S}\left(\frac{\tau_s \tau_1 \tau_2}{\tau_v \tau_1 \tau_2 + \tau_s \tau_1 \tau_2 + h_{i2}^2 \tau_s \tau_2 \tau_v + h_{i3}^2 \tau_s \tau_1 \tau_v}\right) \Delta p_i \\
&= \bar{v} + \frac{1}{\alpha_i^S}\left(\frac{\phi_i \tau_s}{\tau_v + \phi_i \tau_s}\right) \Delta p_i,
\end{aligned} \tag{8}$$

where

$$\phi_i = \frac{\tau_1\tau_2}{\tau_1\tau_2 + h_{i2}^2\tau_2\tau_s + h_{i3}^2\tau_1\tau_s} \text{ and } h_{i1} = \alpha_{i1}^Y/\alpha_i^S, h_{i2} = \alpha_{i2}^Y/\alpha_i^S \tag{9}$$

and

$$\begin{aligned} Var(v|p_i) &= Var(v) - \frac{Cov(v,p_i)^2}{Var(p_i)} \\ &= \frac{1}{\tau_v}\left\{1 - \frac{(\alpha_i^S)^2(1/\tau_v)}{(\alpha_i^S)^2(1/\tau_s + 1/\tau_v) + (\alpha_{i1}^Y)^2/\tau_1 + (\alpha_{i2}^Y)^2/\tau_2}\right\} \\ &= \frac{1}{\tau_v}\left\{1 - \frac{1/\tau_v}{1/\tau_v + 1/\tau_s + h_{i2}^2/\tau_1 + h_{i3}^2/\tau_2}\right\} \end{aligned} \tag{10}$$

or

$$\tau(v|p_i) \equiv \frac{1}{Var(v|p_i)} = \tau_v + \frac{\tau_1\tau_2\tau_s}{\tau_1\tau_2 + h_{i2}^2\tau_2\tau_s + h_{i3}^2\tau_1\tau_s} = \tau_v + \phi_i\tau_s. \tag{11}$$

Under the CARA utility function assumption, the demand function for uninformed traders is

$$\begin{aligned} x_i^U(p_i) &= \frac{\mathbb{E}(v|p_i) - p_i(1+r)}{\rho Var(v|p_i) + c} \\ &= \left\{\overline{v} + \frac{1}{\alpha_i^S}\left(\frac{\phi_i\tau_s}{\tau_v + \phi_i\tau_s}\right)\Delta p_i - p_i(1+r)\right\}\Big/\left(\frac{\rho}{\tau_v + \phi\tau_s} + c\right) \\ &= \beta_i^U(\mu_i^U - p_i). \end{aligned} \tag{12}$$

where

$$\beta_i^U = \left\{1 + r - \frac{1}{\alpha_i^S}\left(\frac{\phi_i\tau_s}{\tau_v + \phi_i\tau_s}\right)\right\}\Big/\left(\frac{\rho}{\tau_v + \phi\tau_s} + c\right) \tag{13}$$

and

$$\begin{aligned} \mu_i^U &= \left\{\overline{v} - \frac{1}{\alpha_i^S}\left(\frac{\phi_i\tau_s}{\tau_v + \phi_i\tau_s}\right)E(p_i)\right\}\Big/\left\{1 + r - \frac{1}{\alpha_i^S}\left(\frac{\phi_i\tau_s}{\tau_v + \phi_i\tau_s}\right)\right\} \\ &= \frac{\overline{v}}{1+r} - \lambda_i \end{aligned} \tag{14}$$

with

$$\lambda_i = \left(\alpha_{i0} - \frac{\bar{v}}{1+r}\right)\left(\frac{1}{\alpha_i^S}\right)\left(\frac{\phi_i\tau_s}{\tau_v+\phi_i\tau_s}\right)\Big/\left\{1+r-\frac{1}{\alpha_i^S}\left(\frac{\phi_i\tau_s}{\tau_v+\phi_i\tau_s}\right)\right\}. \tag{15}$$

λ_i can be shown to be zero in equilibrium, thus the reservation price for uninformed traders is $\frac{\bar{v}}{1+r}$. When the price is $p_i = \bar{v}/(1+r)$, there is no useful information regarding the signal S for uninformed traders, thus they think there is no earnings surprise ($\Delta S = 0$). In that case, uninformed traders will demand zero risky asset. Alternatively, if $p_i < (>)\ \bar{v}/(1+r)$, uninformed traders will conjecture a negative (positive) signal from the observed price and they will choose to sell (buy) the risky asset. We establish the following proposition of the demand elasticity of informed and uninformed traders.

Proposition 1. $\beta^I > \beta_i^U$ *for* $i = 1,2$.

Proof. See Appendix B.

Proposition 1 tells us that the demand elasticity of informed traders is larger than that of uninformed traders.

2.1.2. Arbitrageurs

We subsequently consider the demand function of arbitrageurs, who are able to go long and short in both markets. Suppose an arbitrageur holds a portfolio $(\{B, x_1^A, x_2^A\})$, where B is the amount of the risk-free asset and (x_1^A, x_2^A) are the amount of a given cross-listed pair held in respective markets 1 and 2, subject to the initial wealth $\overline{B}$, then we have

$$\overline{B} = B + p_1 x_1^A + p_2 x_2^A + \frac{1}{2}c(x_1^A)^2 + \frac{1}{2}c(x_2^A)^2. \tag{16}$$

The future wealth will be

$$\begin{aligned} W &= (1+r)B + v(x_1^A + x_2^A) \\ &= v(x_1^A + x_2^A) + (1+r)\left\{\overline{B} - P_1 x_1^A - P_2 x_2^A - \frac{1}{2}c(x_1^A)^2 - \frac{1}{2}c(x_2^A)^2\right\}. \end{aligned} \tag{17}$$

Under the CARA utility function assumption, the arbitrageur's objective function can be written as

$$\mathbb{E}(W) - \frac{\rho}{2}Var(W) = (1+r)\left\{\overline{B} - P_1 x_1^A - P_2 x_2^A - \frac{1}{2}c(x_1^A)^2 - \frac{1}{2}c(x_2^A)^2\right\}$$

$$+E(v)(x_1^A + x_2^A) - \tfrac{\rho}{2}Var(v)(x_1^A + x_2^A)^2. \quad (18)$$

We also assume that arbitrageurs use a perfectly hedged strategy so that their short position equals their long position: $x_1^A(P_1, P_2) + x_2^A(P_1, P_2) = 0$. Under this condition, the demand function of the arbitrageur is

$$x_1^A(P_1, P_2) = -x_2^A(P_1, P_2) = \tfrac{(p_2 - p_1)}{2c}. \quad (19)$$

The aggregate demand of arbitrageurs is

$$X_1^A(P_1, P_2) = -X_2^A(P_1, P_2) = N_3 \tfrac{(p_2 - p_1)}{2c} = \beta^A(p_2 - p_1), \quad (20)$$

where $\beta^A = \frac{N_3}{2c}$ is the aggregate demand elasticity of arbitrageurs.

2.1.3. Market equilibrium

The market clearing condition for each exchange i is given by

$$\textstyle\sum_{j=1}^{N_1\pi_1} \Delta x_{1,j}^I(p_1, S) + \sum_{j=N_1\pi_1+1}^{N_1} \Delta x_{1,j}^U(p_1, S) + \beta^A(p_2 - p_1) = \Delta Y_1, \quad (21)$$

$$\textstyle\sum_{j=1}^{N_2\pi_2} \Delta x_{2,j}^I(p_2, S) + \sum_{j=N_2\pi_2+1}^{N_2} \Delta x_{2,j}^U(p_2, S) + \beta^A(p_1 - p_2) = \Delta Y_2. \quad (22)$$

By plugging demand functions of informed and uninformed traders, we have

$$\pi_1 N_1 \beta^I(\mu^I - p_1) + (1-\pi_1)N_1\beta_1^U(\mu_1^U - p_1) + \beta^A(p_2 - p_1) = \Delta Y_1, \quad (23)$$

$$\pi_2 N_2 \beta^I(\mu^I - p_2) + (1-\pi_2)N_2\beta_2^U(\mu_2^U - p_2) + \beta^A(p_1 - p_2) = \Delta Y_2. \quad (24)$$

We define

$$D_1 \equiv \beta^I\pi_1 N_1 + \beta_1^U(1-\pi_1)N_1 \text{ and } D_2 \equiv \beta_2^U(1-\pi_2)N_2 + \beta^I\pi_2 N_2, \quad (25)$$

$$\mu_1 \equiv \{\mu_1^U\beta_1^U(1-\pi_1)N_1 + \mu^I\beta^I\pi_1 N_1\}/\{\beta^I\pi_1 N_1 + \beta_1^U(1-\pi_1)N_1\}, \quad (26)$$

$$\mu_2 \equiv \{\mu_2^U\beta_2^U(1-\pi_2)N_2 + \beta^I\mu^I\pi_2 N_2\}/\{\beta_2^U(1-\pi_2)N_2 + \beta^I\pi_2 N_2\}. \quad (27)$$

Notice that D_1 and D_2 are the aggregate demand elasticities in markets 1 and 2, respectively. μ_1 and μ_2 are weighted respective averages of the reservation prices of informed and uninformed traders.

Since $\mu_i^U = \overline{v}/(1+r) - \lambda_i$ and $\mu^I = \left\{\overline{v} + \left(\frac{\tau_s}{\tau_s+\tau_v}\right)\Delta S\right\}/(1+r)$, we have

$$\mu_1 = \overline{v}/(1+r) + \Delta S \frac{1}{(1+r)} \frac{\tau_s}{(\tau_s+\tau_v)} \left\{\frac{\beta^I \pi_1 N_1}{\beta^I \pi_1 N_1 + \beta_1^U N_1 (1-\pi_1)}\right\} - \lambda_1 \left\{\frac{\beta_1^U (1-\pi_1) N_1}{\beta^I \pi_1 N_1 + \beta_1^U N_1 (1-\pi_1)}\right\}, \tag{28}$$

$$\mu_2 = \overline{v}/(1+r) + \Delta S \frac{1}{(1+r)} \frac{\tau_s}{(\tau_s+\tau_v)} \left\{\frac{\beta^I \pi_2 N_2}{\beta^I \pi_2 N_2 + \beta_2^U N_2 (1-\pi_2)}\right\} - \lambda_2 \left\{\frac{\beta_2^U (1-\pi_2) N_2}{\beta^I \pi_2 N_2 + \beta_2^U N_2 (1-\pi_2)}\right\}. \tag{29}$$

According to Easley et al. (1997a,b), the PIN is the probability of a random-chosen trader being information based, thus we define

$$PIN_1 \equiv \frac{\beta^I \pi_1 N_1}{\beta^I \pi_1 N_1 + \beta_1^U N_1 (1-\pi_1)}, \tag{30}$$

$$PIN_2 \equiv \frac{\beta^I \pi_2 N_2}{\beta^I \pi_2 N_2 + \beta_2^U N_2 (1-\pi_2)}. \tag{31}$$

Let $h = \frac{\tau_s}{(\tau_s+\tau_v)}$ measure the precision of the signal. Larger h implies a more informative signal. With the above notations, we simplify μ_1 and μ_2 as

$$\mu_1 = \frac{1}{1+r}(\overline{v} + hPIN_1 \Delta S) - \lambda_1 (1 - PIN_1), \tag{32}$$

$$\mu_2 = \frac{1}{1+r}(\overline{v} + hPIN_2 \Delta S) - \lambda_2 (1 - PIN_2). \tag{33}$$

The equilibrium prices for two markets are derived as

$$p_1 = \frac{\mu_2 \beta^A / D_1 + \mu_1 (1 + \beta^A / D_2)}{(1 + \beta^A / D_1) + \beta^A / D_2} - \frac{(\beta^A + D_2)\Delta Y_1 + \beta^A \Delta Y_2}{D_2 D_1 + \beta^A (D_1 + D_2)}, \tag{34}$$

$$p_2 = \frac{\mu_1 \beta^A / D_2 + \mu_2 (1 + \beta^A / D_1)}{(1 + \beta^A / D_1) + \beta^A / D_2} - \frac{(\beta^A + D_1)\Delta Y_2 + \beta^A \Delta Y_1}{D_2 D_1 + \beta^A (D_1 + D_2)}. \tag{35}$$

By plugging in the expressions for μ_1 and μ_2, we get

$$p_1 = \overline{v}/(1+r) - \frac{\lambda_2 (1 - PIN_2) \beta^A / D_1 + \lambda_1 (1 - PIN_1)(1 + \beta^A / D_2)}{(1 + \beta^A / D_1) + \beta^A / D_2}$$

$$+\Delta S\frac{h}{(1+r)}\left\{\frac{PIN_2\beta^A/D_1+PIN_1(1+\beta^A/D_2)}{(1+\beta^A/D_1)+\beta^A/D_2}\right\}-\frac{(\beta^A+D_2)\Delta Y_1+\beta^A\Delta Y_2}{D_2D_1+\beta^A(D_1+D_2)} \tag{36}$$

and

$$p_2=\bar{v}/(1+r)-\frac{\lambda_1(1-PIN_1)\beta^A/D_2+\lambda_2(1-PIN_2)(1+\beta^A/D_1)}{(1+\beta^A/D_1)+\beta^A/D_2}$$

$$+\Delta S\frac{h}{(1+r)}\left\{\frac{PIN_1\beta^A/D_2+PIN_2(1+\beta^A/D_1)}{(1+\beta^A/D_1)+\beta^A/D_2}\right\}-\frac{(\beta^A+D_1)\Delta Y_2+\beta^A\Delta Y_1}{D_2D_1+\beta^A(D_1+D_2)}. \tag{37}$$

Letting the above two equations be equal to the linear conjecture model,

$$p_i=\alpha_{i0}+\alpha_i^S\Delta S-\alpha_i^Y\Delta Y_i-\alpha_{-i}^Y\Delta Y_{-i}, \tag{38}$$

yields

$$\alpha_{10}=\bar{v}/(1+r)-\frac{\lambda_2(1-PIN_2)\beta^A/D_1+\lambda_1(1-PIN_1)(1+\beta^A/D_2)}{1+\beta^A/D_1+\beta^A/D_2}, \tag{39}$$

$$\alpha_1^S=\frac{h}{(1+r)}\left\{\frac{PIN_2D_2\beta^A+D_1PIN_1(D_2+\beta^A)}{D_2D_1+\beta^A(D_1+D_2)}\right\}, \tag{40}$$

$$\alpha_{11}^Y=\frac{(\beta^A+D_2)}{D_2D_1+\beta^A(D_1+D_2)}, \tag{41}$$

$$\alpha_{12}^Y=\frac{\beta^A}{D_2D_1+\beta^A(D_1+D_2)}, \tag{42}$$

$$\alpha_{20}=\bar{v}/(1+r)-\frac{\lambda_1(1-PIN_1)\beta^A/D_2+\lambda_2(1-PIN_2)(1+\beta^A/D_1)}{1+\beta^A/D_1+\beta^A/D_2}, \tag{43}$$

$$\alpha_2^S=\frac{h}{(1+r)}\left\{\frac{PIN_1D_1\beta^A+D_2PIN_2(D_1+\beta^A)}{D_2D_1+\beta^A(D_1+D_2)}\right\}, \tag{44}$$

$$\alpha_{21}^Y=\frac{\beta^A}{D_2D_1+\beta^A(D_1+D_2)}, \tag{45}$$

$$\alpha_{22}^Y=\frac{(\beta^A+D_1)}{D_2D_1+\beta^A(D_1+D_2)}. \tag{46}$$

Since

$$\lambda_i=\left(\alpha_{i0}-\frac{\bar{v}}{1+r}\right)\left(\frac{1}{\alpha_i^S}\right)\left(\frac{\phi_i\tau_s}{\tau_v+\phi_i\tau_s}\right)\Big/\left\{1+r-\frac{1}{\alpha_i^S}\left(\frac{\phi_i\tau_s}{\tau_v+\phi_i\tau_s}\right)\right\}$$

$$=\left(\alpha_{i0}-\frac{\bar{v}}{1+r}\right)\theta_i, \tag{47}$$

where

$$\theta_i = \left(\frac{1}{\alpha_{i1}}\right)\left(\frac{\phi_i \tau_s}{\tau_v + \phi_i \tau_s}\right) \Big/ \left\{1 + r - \frac{1}{\alpha_{i1}}\left(\frac{\phi_i \tau_s}{\tau_v + \phi_i \tau_s}\right)\right\}, \tag{48}$$

thus

$$\lambda_1 = -\left[\frac{\{\lambda_2(1-PIN_2)\}(\beta^A/D_1) + \{\lambda_1(1-PIN_1)\}(1+\beta^A/D_2)}{1+\beta^A/D_1+\beta^A/D_2}\right]\theta_1, \tag{49}$$

$$\lambda_2 = -\left[\frac{\{\lambda_1(1-PIN_1)\}(\beta^A/D_2) + \{\lambda_2(1-PIN_2)\}(1+\beta^A/D_1)}{1+\beta^A/D_1+\beta^A/D_2}\right]\theta_2. \tag{50}$$

Solving $\lambda_1 = \lambda_2 = 0$ gives

$$\mu_1 = \frac{1}{1+r}(\overline{v} + hPIN_1\Delta S), \tag{51}$$

$$\mu_2 = \frac{1}{1+r}(\overline{v} + hPIN_2\Delta S). \tag{52}$$

μ_1 and μ_2 are the market-expected values of the cross-listed pair traded on respective exchanges 1 and 2; they are the sums of the present values of unconditional expected fundamental value and a premium/discount due to earnings surprise (ΔS) magnified by the relative degree of private information (PIN). When the signal is more informative, with a higher PIN, we have a higher (lower) market expected value for a positive (negative) signal.

We can subsequently solve the expressions of α_1^S α_2^S α_{i1}^Y and α_{i2}^Y based on the following six equations:

$$\alpha_1^S = \frac{h}{(1+r)}\left\{\frac{\beta^I \pi_2 N_2 \beta^A + \beta^I \pi_1 N_1 (D_2 + \beta^A)}{D_2 D_1 + \beta^A (D_1 + D_2)}\right\}, \tag{53}$$

$$\alpha_{11}^Y = \frac{(\beta^A + D_2)}{D_2 D_1 + \beta^A (D_1 + D_2)}, \tag{54}$$

$$\alpha_{12}^Y = \frac{\beta^A}{D_2 D_1 + \beta^A (D_1 + D_2)}, \tag{55}$$

$$\alpha_2^S = \frac{h}{(1+r)}\left\{\frac{\beta^I \pi_1 N_1 \beta^A + \beta^I \pi_2 N_2 (D_1 + \beta^A)}{D_2 D_1 + \beta^A (D_1 + D_2)}\right\}, \tag{56}$$

$$\alpha_{21}^Y = \frac{\beta^A}{D_2 D_1 + \beta^A (D_1 + D_2)}, \tag{57}$$

$$\alpha_{22}^{Y} = \frac{(\beta^A + D_1)}{D_2 D_1 + \beta^A (D_1 + D_2)}. \tag{58}$$

Firstly, α_{12}^{Y} and α_{21}^{Y} are symmetric and they reflect the interactive price impact of supply shocks from each other. Also notice that $\alpha_{12}^{Y} < \alpha_{11}^{Y}$ and $\alpha_{21}^{Y} < \alpha_{22}^{Y}$, which implies that the interaction effect is smaller than the price impact of a supply shock by itself. Secondly, α_1^S and α_2^S measure the price sensitivity to private information. It can be shown that the degree of price sensitivity is directly related to a weighted average of the PINs of the two markets. See Appendix B for numerical solutions for all coefficients of the price function.

Given the expressions of two equilibrium prices, we can calculate the dollar premium of the cross-listing on the NYSE against its original listing on the TSX as

$$p_2 - p_1 = \frac{\mu_2 - \mu_1}{1 + \beta^A/D_1 + \beta^A/D_2} + \frac{\Delta Y_1/D_1 - \Delta Y_2/D_2}{1 + \beta^A/D_1 + \beta^A/D_2}$$

$$= \Delta S \left(\frac{h}{1+r}\right) \left(\frac{PIN_2 - PIN_1}{1 + \beta^A/D_1 + \beta^A/D_2}\right) + \frac{\Delta Y_1/D_1 - \Delta Y_2/D_2}{1 + \beta^A/D_1 + \beta^A/D_2}. \tag{59}$$

Next, we discuss the dollar premium $(p_2 - p_1)$ with respect to the value of β^A.

Case I: $\beta^A = \infty$, i.e. the market is perfect and there is no transaction cost $(c = 0)$, or arbitragers have an infinite demand elasticity. We can show $p_2 - p_1 = 0$, thus the efficient market price (p_e) is

$$p_1 = p_2 = p_e \equiv \frac{\bar{v}}{1+r} + \Delta S \left(\frac{h}{1+r}\right) \left\{\frac{PIN_2 D_2 + PIN_1 D_1}{D_1 + D_2}\right\} - \frac{\Delta Y_1 + \Delta Y_2}{(D_1 + D_2)}. \tag{60}$$

Case II: $0 < \beta^A < \infty$, i.e. there are limits to arbitrage, thus

$$p_2 - p_1 = \Delta S \left(\frac{h}{1+r}\right) \left(\frac{PIN_2 - PIN_1}{1 + \beta^A/D_1 + \beta^A/D_2}\right) + \frac{\Delta Y_1/D_1 - \Delta Y_2/D_2}{1 + \beta^A/D_1 + \beta^A/D_2}$$

$$= \frac{\Delta S\{h/(1+r)\}(PIN_1 - PIN_2) + \Delta Y_1/D_1 - \Delta Y_2/D_2}{1 + \beta^A/D_1 + \beta^A/D_2}. \tag{61}$$

Case III: $\beta^A = 0$, i.e. two markets are completely separated so that there is no cross-border arbitrage, thus

$$p_2 - p_1 = \Delta S\left(\frac{h}{1+r}\right)(PIN_2 - PIN_1) + \Delta Y_1/D_1 - \Delta Y_2/D_2. \tag{62}$$

In conclusion, we can see the dollar premium on the NYSE-cross-listing against its original TSX-listing is largest in the absence of arbitrage. The price spread is negatively related to the demand elasticity of arbitrageurs, or positively related to the transaction cost, and these predictions are consistent with the empirical findings of Gagnon and Karolyi (2010).

2.2. Price discovery and information asymmetry

The noisy rational expectations model presented in Section 2.1 describes a static equilibrium relationship between the two prices of a given cross-listed pair. However, in order to faciliate a better framework in the time-series analytic context, the equilibrium model must be supplemented with evolution of the market reservation prices $\mu_{1,t}$ and $\mu_{2,t}$. Following Garbade and Silber (1983) and Kyle (1985),[6] we assume that the dynamics of $\mu_{1,t}$ and $\mu_{2,t}$ are determined as

$$\mu_{1,t} = \mu_{1,t-1} + hPIN_1\Delta S_t + \varepsilon_{1t}, \tag{63}$$

$$\mu_{2,t} = \mu_{2,t-1} + hPIN_1\Delta S_t + \varepsilon_{2t}, \tag{64}$$

where ΔS_t reflects new information signal arrival between periods $t-1$ and t. ΔS_t and ε_{it} are assumed to be stationary processes, and $\mu_{1,t}$ and $\mu_{2,t}$ are random walks. Moreover, after market clearance at the end of the period $t-1$, p_{it-1} is the reservation price for every trader in market i, thus we have

$$\mu_{1,t} = p_{1,t-1} + hPIN_1\Delta S_t + \varepsilon_{1t}, \tag{65}$$

$$\mu_{2,t} = p_{2,t-1} + hPIN_1\Delta S_t + \varepsilon_{2t}. \tag{66}$$

[6]Kyle (1985) assumes the expected value of security conditional on the information set at time t:

$$m_t = m_{t-1} + \gamma Q_t V_t + e_t,$$

where γ is the price impact (inverse market depth) parameter; V_t is trade size and Q_t is trade sign (+1 if buy, -1 sell); e_t is the public information signal; γ is used to capture the effects of asymmetric information.

According to the previously worked out solutions for equilibrium prices, we have

$$p_{1,t} = \frac{\mu_{2,t}(\beta^A/D_1)+\mu_{1,t}(1+\beta^A/D_2)}{(1+\beta^A/D_1)+\beta^A/D_2} - \frac{(\beta^A+D_2)\Delta Y_{1t}+\beta^A\Delta Y_{2t}}{D_2D_1+\beta^A(D_1+D_2)}, \tag{67}$$

$$p_{2,t} = \frac{\mu_{1,t}(\beta^A/D_2)+\mu_{2,t}(1+\beta^A/D_1)}{(1+\beta^A/D_1)+\beta^A/D_2} - \frac{(\beta^A+D_1)\Delta Y_{2t}+\beta^A\Delta Y_{1t}}{D_2D_1+\beta^A(D_1+D_2)}. \tag{68}$$

By plugging, they arrive at

$$\begin{pmatrix} p_{1,t} \\ p_{2,t} \end{pmatrix} = \begin{pmatrix} 1-a, a \\ b, 1-b \end{pmatrix}\begin{pmatrix} p_{1,t-1} \\ p_{2,t-1} \end{pmatrix} + \begin{pmatrix} 1-a, a \\ b, 1-b \end{pmatrix}\begin{pmatrix} \Delta Y_{1t}/D_1 + hPIN_1\Delta S_t + \varepsilon_{1t} \\ \Delta Y_{2t}/D_2 + hPIN_1\Delta S_t + \varepsilon_{2t} \end{pmatrix}, \tag{69}$$

where

$$a = \frac{\beta^A/D_1}{1+\beta^A/D_1+\beta^A/D_2}, \tag{70}$$

$$b = \frac{\beta^A/D_2}{1+\beta^A/D_1+\beta^A/D_2}. \tag{71}$$

A resulting VECM can be derived by subtracting $(p_{1,t-1}, p_{2,t-1})^T$ from both sides:

$$\begin{aligned}\begin{pmatrix} \Delta p_{1,t} \\ \Delta p_{2,t} \end{pmatrix} &= \begin{pmatrix} -a, a \\ b, -b \end{pmatrix}\begin{pmatrix} p_{1,t-1} \\ p_{2,t-1} \end{pmatrix} + \begin{pmatrix} 1-a, a \\ b, 1-b \end{pmatrix}\begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix} \\ &= (-a, b)^T\begin{pmatrix} 1, -1 \\ 1, -1 \end{pmatrix}\begin{pmatrix} p_{1,t-1} \\ p_{2,t-1} \end{pmatrix} + \begin{pmatrix} 1-a, a \\ b, 1-b \end{pmatrix}\begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix} \\ &= (-a, b)^T(p_{1,t-1} - p_{2,t-1}) + \begin{pmatrix} 1-a, a \\ b, 1-b \end{pmatrix}\begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix}. \end{aligned} \tag{72}$$

Since $\mu_{1,t}$ and $\mu_{2,t}$ are assumed to be random walks, it can be shown that both prices, $p_{1,t}$ and $p_{2,t}$, are unit root processes, thereby the above VECM describes short term dynamics toward the long run equilibrium given the cointegrating vector $(1, -1)$. The short term adjustment coefficients are a and b for prices $p_{1,t}$ and $p_{2,t}$, respectively, which reflect response to a deviation from the long run equilibrium in each market.

When $\beta^A \neq 0$, we can apply the permanent transitory decomposition, per Gonzalo and Granger (1995), to the above VECM. The permanent component is a linear combination of the two prices, $(p_{1,t},\ p_{2,t})$, formed by a scaled orthogonal vector of the adjustment coefficient vector, (a, b):

$$f_t = \frac{b}{a+b} p_{1,t} + \frac{a}{a+b} p_{2,t}. \tag{73}$$

Now that

$$\frac{b}{a+b} = \frac{\beta^A/\{\beta_2^U N_2(1-\pi_2) + \beta^I \pi_2 N_2\}}{\beta^A/\{\beta_2^U N_2(1-\pi_2) + \beta^I \pi_2 N_2\}) + \beta^A/\{\beta^I \pi_1 N_1 + \beta_1^U N_1(1-\pi_1)\}}$$

$$= \frac{\beta^I \pi_1 N_1 + \beta_1^U N_1(1-\pi_1)}{\{\beta_2^U N_2(1-\pi_2) + \beta^I \pi_2 N_2\} + \{\beta^I \pi_1 N_1 + \beta_1^U N_1(1-\pi_1)\}}$$

$$= \frac{D_1}{D_1 + D_2} \tag{74}$$

and

$$\frac{a}{a+b} = \frac{D_2}{D_1 + D_2}, \tag{75}$$

we have

$$f_t = \frac{b}{a+b} p_{1,t} + \frac{a}{a+b} p_{2,t}$$

$$= \overline{v}_t/(1+r) + \Delta S_t \frac{h}{(1+r)} \frac{PIN_2 D_2 + PIN_1 D_1}{D_1 + D_2} - \frac{\Delta Y_{1t} + \Delta Y_{2t}}{(D_1 + D_2)}$$

$$\equiv p_t^e. \tag{76}$$

Notice that the derived permanent component (p_t^e) is the price under the condition that $\beta^A \to \infty,$ i.e. under the perfect market assumption. According to Harris et al. (1997, 2002), the information share for each market is:

$$IS_1 \equiv \frac{b}{a+b} = \frac{D_1}{D_1 + D_2}, \tag{77}$$

$$IS_2 \equiv \frac{a}{a+b} = \frac{D_2}{D_1 + D_2}, \tag{78}$$

where IS_1 and IS_2 reflect the relative contribution share of each exchange-specific price to the permanent component.[7] Hasbrouck (1995, 2002) and Figuerola-Ferretti and Gonzalo (2009) find $b = N_1$ and $a = N_2$, thus they view information shares as fractions of the number of participants in respective markets: the information share is positively related to the relative trading volume. However, they elicit their results asuming that all market participants are informed, $\pi_1 = \pi_2 = 1$, which is a special case of our general solutions laid out as the following propositions:

Proposition 2. *Given $PIN_1 > PIN_2$ and $N_1 \geq N_2$, we have $IS_1 > IS_2$.*

Proof. See Appendix B.

Lastly, we discuss the convergence speed of the two market prices to their long run equilibrium. One can conjecture an AR(1) model using the price spread and the first order coefficient is taken as the measure of convergence speed to the equilibrium parity.

$$\Delta p_t = \delta \Delta p_{t-1} + \varepsilon_t, \tag{79}$$

where $\Delta p_t \equiv p_{1t} - p_{2t}$. According to Garbade and Silber (1983), δ can be expressed as

$$\begin{aligned}\delta &= 1 - a - b \\ &= 1 - \frac{\beta^A/D_1}{1 + \beta^A/D_1 + \beta^A/D_2} - \frac{\beta^A/D_2}{1 + \beta^A/D_1 + \beta^A/D_2} \\ &= \frac{1}{1+\beta^A/D_1+\beta^A/D_2}.\end{aligned} \tag{80}$$

We present the following Proposition 3 on the convergence speed without proof.

Proposition 3. *The demand elasticity of arbitrageurs β^A has no direct relationship with the information share of each market, but it is positively related to the short term convergence speed δ.*

[7] See Figuerola-Ferretti and Gonzalo (2009) for a detailed procedure in extracting the information share for two closely related markets.

3. Error correction models

In order to empirically implement the noisy rational expectation model presented in Section 2, it has to be re-rendered in the time-series context by establishing an appropriate econometric platform. We begin with the existing linear ECM for price discovery, followed by a bivariate threhold ECM to reflect the regime-switching pattern of cross-border arbitrage on cross-listed pairs.

3.1. Standard error correction model

For any firm i at time t, let p_{it}^{N} and p_{it}^{T} be the prices of its listings on the NYSE and the TSX, respectively. The law of one price enforces parity in the long run:

$$p_{it}^{\mathrm{N}} = \omega_i \chi_t p_{it}^{\mathrm{T}}, \tag{81}$$

where ω_i is the home-host bundling ratio and χ_t is the foreign exchange rate at time t. As it is the case for most TSX-NYSE cross-listed pairs, let $\omega_i = 1$. The U.S. \$-denominated parity-implied price of the cross-listing on the NYSE is

$$\tilde{p}_{it}^{\mathrm{T}} \equiv \omega_i \chi_t p_{it}^{\mathrm{T}} = \chi_t p_{it}^{\mathrm{T}}. \tag{82}$$

In reality, we almost always observe a spread between the two prices due to various market forces. Cointegration is an appropriate concept to describe the long-run relationship between the pair of prices. The two security prices are cointegrated if, by definition, their long run linear relationship is stationary, i.e. a significant deviation is shortlived.[8] Empirically, cointegration can be verified by testing the deviation time series for stationarity.[9] Given a cointegrating vector $((1, -1)^T)$, define the dollar premium on the NYSE-listing against its original TSX-listing as

$$\kappa_{it} \equiv p_{it}^{\mathrm{N}} - \tilde{p}_{it}^{\mathrm{T}}. \tag{83}$$

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[8]A group of multiple random-walk processes is cointegrated if, by definition, there exists a stationary linear combination of the processes.

[9]A time series is (weakly) stationary if the probability laws (of up to the second moments) are time-invariant.

The first time-differences of the two price series are

$$\Delta p_{it}^{\mathrm{N}} = p_{it}^{\mathrm{N}} - p_{it-1}^{\mathrm{N}} \text{ and } \Delta \tilde{p}_{it}^{\mathrm{T}} = \tilde{p}_{it}^{\mathrm{T}} - \tilde{p}_{it-1}^{\mathrm{T}}. \tag{84}$$

Engle and Granger (1987) and Engle and Yoo (1987) show that if a pair of time-indexed random variables, say p_{it}^{N} and $\tilde{p}_{it}^{\mathrm{T}}$, are cointegrated, the short term dynamics of the two time series can be represented by a bivariate ECM. The error correction mechanism assumes that a fraction of the deviation of a period will be subsequently corrected. A standard ECM for the bivariate cointegrated system of the cross-listed pair can be structured as

$$\Delta p_{it}^{\mathrm{N}} = \beta_0^{\mathrm{N}} + \alpha^{\mathrm{N}} \kappa_{it-1} + \textstyle\sum_{j=1}^{m_1} \beta_j^{\mathrm{N}} \Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_j^{\mathrm{N}} \Delta \tilde{p}_{it-j}^{\mathrm{T}}, \tag{85}$$

$$\Delta \tilde{p}_{it}^{\mathrm{T}} = \beta_0^{\mathrm{T}} + \alpha^{\mathrm{T}} \kappa_{it-1} + \textstyle\sum_{j=1}^{m_1} \beta_j^{\mathrm{T}} \Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_j^{\mathrm{T}} \Delta \tilde{p}_{it-j}^{\mathrm{T}}, \tag{86}$$

where κ_{it-1} gives the remaining cross-listing dollar premium or cointegrating residual. α^{N} and α^{T} are the adjustment coefficients of the NYSE and tsx, respectively, that describe how much deviation will be subsequently adjusted to restore the long run equilibrium in each series. Per the *Granger Representation Theorem*, if p_{it}^{N} and $\tilde{p}_{it}^{\mathrm{T}}$ are cointegrated, then at least one of α^{N} and α^{T} must be nonzero. In other words, p_{it}^{N} or $\tilde{p}_{it}^{\mathrm{T}}$, or both, will adjust fractionally to restore parity in the long run.

Harris et al. (1995, 2000) propose to use the ECM adjustment coefficients to estimate the relative extent of exchange-respective contribution to price discovery (information share) of shares whose order purchases are fragmented across multiple markets. For a Canadian company originally listed on the TSX and cross-listed on the NYSE, the proportion of the adjustments that took place on the TSX out of the total adjustments occurred on both exchanges is the share of the home exchange in contribution to setting the long-run equilibrium price as a result of synchronous cross-border stock trading. In an extreme case where there is no feedback from the NYSE such that $\alpha^{\mathrm{T}} = 0$, then the NYSE has no contribution to price discovery of the cross-listed pair. Eun and Sabherwal (2003) further define the respective information shares of the NYSE and the TSX as

$$\mathrm{IS}^{\mathrm{N}} \equiv \frac{|\alpha^{\mathrm{T}}|}{|\alpha^{\mathrm{T}}|+|\alpha^{\mathrm{N}}|} \text{ and } \mathrm{IS}^{\mathrm{T}} \equiv \frac{|\alpha^{\mathrm{N}}|}{|\alpha^{\mathrm{T}}|+|\alpha^{\mathrm{N}}|}. \quad (87)$$

Suppose $p_{it-1}^{\mathrm{N}} < \tilde{p}_{it-1}^{\mathrm{T}}$ in the previous period $(t-1)$, then a likely scenario to reduce the gap between the two prices is: p_{it}^{N} increases or $\tilde{p}_{it}^{\mathrm{T}}$ decreases, or both. In this case one can conjecture that α^{T} is non-positive and α^{N} is non-negative. There are two other possibilities: (1) p_{it-1}^{n} decreases but $\tilde{p}_{it-1}^{\mathrm{T}}$ decreases more; or (2) p_{it-1}^{N} increases but $\tilde{p}_{it-1}^{\mathrm{T}}$ increases less.[10] As Eun and Sabherwal (2003) mention the latter two outcomes are very unlikely, so they are excluded from our study. One can analogously design a similar adjustment mechanism to show that α^{T} is non-positive and α^{N} is non-negative for the symmetric situation when $p_{it-1}^{\mathrm{N}} > \tilde{p}_{it-1}^{\mathrm{T}}$. Therefore, we define the exchange-respective information shares of the NYSE and the TSX as

$$\mathrm{IS}^{\mathrm{N}} \equiv \frac{|\alpha^{\mathrm{T}}|}{|\alpha^{\mathrm{T}}|+\alpha^{\mathrm{N}}} \text{ and } \mathrm{IS}^{\mathrm{T}} \equiv \frac{\alpha^{\mathrm{N}}}{|\alpha^{\mathrm{T}}|+\alpha^{\mathrm{N}}}. \quad (88)$$

An implicit assumption is that convergence to parity is linear and present in all periods. That is very restrictive given the various market perturbing factors in practice. Adjustment may neither be linearly smooth nor be occurring at every moment. Only when the deviation parity exceeds the transaction costs plus other risk premiums, will arbitrageurs act to take a short position on the dearer side and a long position on the other end. Otherwise, the two prices are unleashed to diverge: the relative premium can follow a near-unit root process. As Krugman (1991) notes, the long run parity relationship can remain inactive within a range of disequilibrium before it becomes active when the system crosses the boundaries of allowed

[10]These odds may reflect the underreaction to the information share of the market. When information incorpration takes multiple periods, the price adjustment should persist unilaterally during then.

fluctuations (thresholds). We subsequently develop a threshold ECM to describe the discrete adjustment mechanism of relative premiums.

3.2. Threshold error correction model

In reality, the market is imperfect due to various sources of market friction such as transaction costs, direct and indirect trading barriers, etc. We let γ_i measure the sum of transaction costs and risk premiums required from arbitrageurs. Arbitrage opportunities exist when

$$\kappa_{it} \equiv p_{it}^{\mathrm{N}} - \tilde{p}_{it}^{\mathrm{T}} < -\gamma_i \text{ or } \kappa_{it} > \gamma_i, \tag{89}$$

which becomes $|\kappa_{it}| > \gamma_i$. Transaction costs of cross-border arbitrage consist of the bid-ask spreads of the prices on both exchanges and the foreign exchange rate, fixed costs, and liquidity shorfalls. Chen and Choi (2010) find the relative premium of a Canadian cross-listing on the NYSE, on average, includes an adverse-selection risk premium due to the cross-border imbalance in private information on the issuing firm. Along with the asymmetric information component, macroeconomic factors, such as GDP growth rates and interest rates, may also affect the determining of the threshold.

Now, cointegration between p_{it}^{N} and $\tilde{p}_{it-1}^{\mathrm{T}}$ is dormant with a range of disequilibrium but the error correction dynamics becomes active once the cross-listing dollar premium sufficiently digresses from parity beyond the threshold. Balke and Fomby (1997) propose this regime-switching mechanism as *threshold cointegration*. Accordingly, κ_{it} is factored in the following threshold autoregressive (TAR) framework:

$$\kappa_{it} = \begin{Bmatrix} \alpha_{\mathrm{out}} + \rho_{\mathrm{out}}\, \kappa_{it-1} + \varepsilon_t, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \alpha_{\mathrm{in}} + \rho_{\mathrm{in}}\, \kappa_{it-1} + \varepsilon_t, \text{ if } |\kappa_{it-1}| \le \gamma_i \end{Bmatrix}, \tag{90}$$

where $\rho_{\mathrm{in}} = 1$ and $-1 < \rho_{\mathrm{out}} < 0$, i.e. the dollar premium (κ_{it}) of the cross-listing is a unit root in the (unprofitable) middle regime when $|\kappa_{it-1}| \leq \gamma_i$, otherwise a mean-reverting process in the (profitable) outer regime when $|\kappa_{it-1}| > \gamma_i$ with the presence of arbitrageurs. Although the cross-listed pair with a TAR feature is cointegrated, the implied error correction dynamics is neither linear nor time-continuous:

$$\Delta p_{it}^{\mathrm{N}} = \begin{cases} \beta_{10}^{\mathrm{N}} + \alpha_{\mathrm{out}}^{\mathrm{N}} \kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{1j}^{\mathrm{N}} \Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{1j}^{\mathrm{N}} \Delta \tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \beta_{20}^{\mathrm{N}} + \alpha_{\mathrm{in}}^{\mathrm{N}} \kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{2j}^{\mathrm{N}} \Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{2j}^{\mathrm{N}} \Delta \tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| \leq \gamma_i \end{cases}, \tag{91}$$

$$\Delta \tilde{p}_{it}^{\mathrm{T}} = \begin{cases} \beta_{10}^{\mathrm{T}} + \alpha_{\mathrm{out}}^{\mathrm{T}} \kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{1j}^{\mathrm{T}} \Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{1j}^{\mathrm{T}} \Delta \tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \beta_{20}^{\mathrm{T}} + \alpha_{\mathrm{in}}^{\mathrm{T}} \kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{2j}^{\mathrm{T}} \Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{2j}^{\mathrm{T}} \Delta \tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| \leq \gamma_i \end{cases}. \tag{92}$$

In the middle regime when $|\kappa_{it-1}| \leq \gamma_i$, there are neither market forces nor arbitrageurs to sustain cointegration of the pair of prices. In other words, unless the pair shows a significant price gap exceeding the threshold minimum profit, the adjustment coefficients are zero ($\alpha_{\mathrm{in}}^{\mathrm{N}} = \alpha_{\mathrm{in}}^{\mathrm{T}} = 0$) and, thus, neither price (p_{it}^{N} nor $\tilde{p}_{it-1}^{\mathrm{T}}$) appropriately reflects the risks. Given that the outer regime typically determines stationarity of a TAR process, we define the information share, or the relative measure of contribution to price discovery, for respective market using the outer regime coefficient estimates[11] ($\alpha_{\mathrm{out}}^{\mathrm{N}}$ and $\alpha_{\mathrm{out}}^{\mathrm{T}}$):

$$\mathrm{IS}^{\mathrm{N}} \equiv \frac{|\alpha_{\mathrm{out}}^{\mathrm{T}}|}{|\alpha_{\mathrm{out}}^{\mathrm{T}}| + \alpha_{\mathrm{out}}^{\mathrm{N}}} \text{ and } \mathrm{IS}^{\mathrm{T}} \equiv \frac{\alpha_{\mathrm{out}}^{\mathrm{N}}}{|\alpha_{\mathrm{out}}^{\mathrm{T}}| + \alpha_{\mathrm{out}}^{\mathrm{N}}}. \tag{93}$$

A large deviation (outer regime) is believed to be more susceptible to new information, either public or private. In contrast, a small deviation (inner regime) can be due to noise trading and, consequently, there is little connection between price discovery and error correction dynamics. Our threshold ECM ideally incorporates such a dichotomy while the predecessor

[11]Eun and Sabherwal (2003) estimate the adjustment coefficients in every period using a linear ECM following Harris et al. (1995).

linear ECMs may overestimate the information share when there is no cointegration in the unprofitable inner regime. See Appendix A for a detailed estimation procedure and significance testing of the parameters of interest in the threshold ECM.

4. Data

56 TSX-NYSE pairs are identified through the sample period: January 1, 1998, through December 31, 2000. In order to estimate asymmetric-information and market-friction measures, 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 was purchased from Olson & Associates.

4.1. Cointegration analysis

We first examine whether pairs of times series on the TSX and the NYSE price series are unit roots or not. We use the augmented Dickey and Fuller's (1981) (ADF) test, which considers lagged first differences of time series in the specification. If the test statistic is too large, then we reject the null hypothesis of unit root and conclude that the time series is stationary. As a result, the null hypothesis was rejected only for four out of 168 firm-years, at a five percent significance level. Thus, we conclude that both price series in our sample are, overall, first-order integrated ($I(1)$) or unit units.

We subsequently examined, using Johansen's (1991) test, to see if there was any cointegration between the two price series. We did not include the S&P TSX Composite and the S&P 500 indices (market indices of the TSX and the NYSE, respectively) in the cointegration system since Eun and Sabherwal (2003) find that the estimated coefficients of the two index series are statistically insignificant. Since we have two price series in each regression equation, there is at most one cointegrating vector. We estimated the cointegrating vector for each cross-listed pair in each year. Our results show that most of the estimated cointegrating vectors are $(1,-1)^T$, which is the expected values according to the law of one price. Table 1 reports summary statistics of the normalized estimation of the cointegrating vector[12] for p_{it}^{N} and $\tilde{p}_{it}^{\mathrm{T}}$, and the *t*-statistics for the null hypothesis attest that the cointegrating vector equals $(1,-1)^T$.

[Insert Table 1 about here.]

In Table 1, we see that the median of the normalized estimates throughout the sample is $(1,-1^T)$ which confirms that the Canadian cross-listed pairs tend to follow the law of one price and are, therefore, cointegrated. Given the estimated cointegrating vector $(b^{\mathrm{N}},-1)^T$, the estimated cross-listing dollar premium is $\kappa_{it} \equiv b^{\mathrm{N}} p_{it}^{\mathrm{N}} - \tilde{p}_{it}^{\mathrm{T}}$. We, then, test κ_{it}, for stationarity per the ADF test and find that only 3 out of 92 samples do not reject the null hypothesis of unit root. Thus, we conclude that there the TSX-NYSE cross-listed pairs are cointegrated.

4.2. Nonlinearity test

The law of one price suggests that two market prices for the same stock should not drift far from each other. This relationship is confirmed by the cointegration analysis in the previous section. However, linear adjustment dynamics is not necessarily prescribed by market efficiency assumptions. In this section, we examine possible nonlinearity in the course of short-run adjustment dynamics to long-run parity equilibrium.

Given various market frictions, such as transactions costs and short sale limitations, arbitrage forces achieving a long run equilibrium depend on the magnitude of price deviation between two prices. Thus, it is more likely that a nonlinear model, such as the threshold

[12]Normalized such that $b^{\mathrm{T}} = -1$.

cointegration model provides a better description of a practical trading environment. We begin our analysis by considering the symmetric bivariate threshold ECM model (introduced in Section 3.2.) by normalizing the cointegrating vector at 1 ($b^{\mathrm{N}} = 1$). We use Akaike's (1974) and Schwart's (1978) Bayesian information criteria to choose the number of lags, and consistently choose the lag length of 1 ($m_1 = m_2 = 1$). The model is estimated by the maximum likelihood method described in Appendix A. We estimate the above model in each quarter for each pair and the results are reported in Table 2 Panels A, B, and C.

[Insert Table 2 Panels A, B, and C about here.]

Table 2 Panel A reports summary statistics of the error correction parameter estimates $(\alpha^{\mathrm{N}}_{\mathrm{out}}, \alpha^{\mathrm{N}}_{\mathrm{in}}, \alpha^{\mathrm{T}}_{\mathrm{out}}, \alpha^{\mathrm{T}}_{\mathrm{in}})$ and the associated *t*-statistics. In general, we find that $\alpha^{\mathrm{T}}_{\mathrm{out}}$ is larger than $\alpha^{\mathrm{T}}_{\mathrm{in}}$, which implies a faster convergence rate in the outer regime. Moreover, it appears that the threshold effect is more likely to take place on the TSX.

Panel B exhibits summary statistics of the threshold estimates. To assess evidence of the threshold effect, we apply the super-Lagrangian multiplier (supLM) test for both cases of cointegrating vectors. The *p*-values are computed by the parametric bootstrap method suggested by Hansen and Seo (2002). From the table, we find that the respective means of supLM for both cointegrating vectors (22.321 and 21.983) are very close to their respective 95% critical values (22.075 and 22.090), which implies that we can almost reject the null hypothesis of no threshold effect. It is not surprising that we did not find significant threshold effect in some quarters for some stocks since it is possible that we did not observe any price deviation exceeding the threshold value, for certain cross-listed pairs, required by arbitrageurs in those quarters. Alternatively, even though the price deviation is very large, there can still be an absence of arbitrageurs surrounding such pairs due to poor liquidity or high transaction costs. In sum, the supLM test further shows that there are threshold effects in the short-run adjustment procedure.

We further tested the threshold effect in the long run in Panel C using the Wald statistics. $\mathrm{Wald}_{\mathrm{ECM}}$ gives the Wald statistic for the joint null hypothesis: $H_0: \alpha_{\mathrm{out}}^{\mathrm{T}} = \alpha_{\mathrm{in}}^{\mathrm{T}}$ and $\alpha_{\mathrm{out}}^{\mathrm{N}} = \alpha_{\mathrm{in}}^{\mathrm{N}}$, while $\mathrm{Wald}_{\mathrm{DC}}$ gives the wald-statistic for: $H_0: \beta_{1j}^{\mathrm{N}} = \beta_{2j}^{\mathrm{N}}$, $\beta_{1j}^{\mathrm{N}} = \beta_{2j}^{\mathrm{N}}$, $\beta_{1j}^{\mathrm{T}} = \beta_{2j}^{\mathrm{T}}$, and $\beta_{1j}^{\mathrm{T}} = \beta_{2j}^{\mathrm{T}}$. The results show that there are, on average, threshold effects in both the error correction and short dymanic terms.

4.3. Dataset construction

4.3.1. Microstructure measures

Unlike the NYSE, which is a specialist-based auction exchange, the TSX is an electronic exchange, which uses a Central Limit Order Book (CLOB) system, where orders are required to be posted in the book to be valid.[13] By studying decrements in the inside depth on one side of the quote that correspond to uncommon trade sizes (such as a trade of 1,300 shares), matching trades with prevailing quotes with a 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.[14]

We construct the preliminary datasets for estimation of the PIN following Easley et al. (1996, 2002). The NYSE-resident specialists are central to the theory of the PIN (Easley et al., 2001; Duarte and Young, 2008). There are official market makers, known as registered traders, on the TSX whose function is akin to that of the NYSE specialists. Thus, a comparison of trade

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

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

informedness on the two exchanges by the PIN is deemed appropriate.[15]

4.3.2. Panel data for regression analyses

We construct a panel data for regression analyses of the estimates of information shares and thresholds with columns of various indices, dependent variables, explanatory variables, and control variables. *Symbol* is the NYSE ticker of a TSX-NYSE cross-listed pair. *Year* is the year index of an estimated value. *IsLin* is the information share estimate of the NYSE per Harris et al. (1995, 2002). *ISIn* is the inner-regime information share estimate of the NYSE.

- *Dependent variables. IsOut* is the outer-regime information share of the NYSE. *Threshold* is the U.S.$-denominated threshold estimate.
- *Explanatory variables. PinRatio* is the ratio of the PIN of the NYSE over that of the TSX. *PINAvg* is the average PIN of the NYSE and the TSX. *PinDiff* is the difference between the PIN of the NYSE and that of the TSX. *SpreadRat* is the ratio of the relative quoted bid-ask spread of the NYSE over that of the TSX. *SpreadAvg* is the average relative quoted bid-ask spread of the NYSE and the TSX. *SpreadDiff* is the difference of the quoted bid-ask spread of the NYSE over that of the TSX.
- *Control variables. USVol* is the average daily trading volume of the NYSE out of both the NYSE and the TSX following Eun and Sabherwal (2003). *VolAvg* is the average of the log-transformations of average daily trading volume measures of the NYSE and the TSX. *VolDiff* is the difference of the log-transformation of average daily trading volume of the NYSE over that of the TSX. *UsDollarVol* is the average daily dollar trading volume of the NYSE out of both of the NYSE and the TSX. *DollarVolAvg* is the sum of log-transformations of average daily dollar trading volume measures of the NYSE and

[15]We owe this comment to Lawrence Kryzanowski. See Fuller, Van Ness, and Van Ness (2008) for difficulties in estimating the PIN for Nasdaq trades.

the TSX. *DollarVolDiff* is the difference of the log-transformation of average daily dollar trading volume of the NYSE over that of the TSX. *NoteAvg* and *NoteDiff* are the average and difference of the U.S. and Canadas' 10-year Treasury Note yields, respectively. *BillAvg* and *BillDiff* are the average and difference of the U.S. and Canadas' 90-day Treasury bill discounts, respectively. *VolatAvg* and *VolatDiff* are the average and difference of the U.S. and Canadas' market index return volatility, respectively. *GdpAvg* and *GdpDiff* are the average and difference of the U.S. and Canadas' GDP growth rates, respectively. *Governance* is the Report on Business governance index of Canadian firms published by *Globe and Mail* (McFarland, 2002). *Industry* equals one if the cross-lister is a manufacturing firm, and zero otherwise. *Size* is the normalized average market capitalization on the TSX and the NYSE.

5. Empirics

5.1. Estimation

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,b). Further, we adopt Easley, Engle, O'Hara, and Wu's (2008) log-likelihood function specification for improved numerical stability in computing the the PIN. The bid-ask spreads 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 previous minutes' exchange rate quote mid-points.

Based on, unreported, ten-minute frequency relative premiums of 56 cross-listed pairs traded throughout the sample period, the arithmetic mean, the median, and the standard deviation are

0.00306, 0.00004, and 0.03031, respectively. The average relative premium of 30.6 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. Chen and Choi (2010) report 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: the positive but small average daily relative premium is a result of cross-border imbalance in private information.

We first employ a linear ECM to estimate adjustment coefficients following Eun and Sabherwal (2003). The estimated coefficients are summarized in the first column, Table 3. Eun and Sabherwal's (2003) sample period is February through July, 1998. Their estimates of the information share of the NYSE (IS^N) range from 0.2% to 98.2%, with an average of 38.1%. They conclude that price discovery for most cross-listed pairs occurs on the TSX, but there is also significant feedback from the NYSE. Our results based on a longer sample period are consistent with their results: the estimated information shares of the NYSE (IS^N) range from 1% to 97.5%, with a mean of 40.7%. There is no discernable trend over the sample period as the yearly estimates of IS^N in 1998, 1999, and 2000 are 0.45, 0.48, and 0.387, respectively.

As we emphasized in Section 3.2, the threshold ECM model purports to yield less biased estimates of information shares. Based on estimates, the outer regime adjustment coefficients via the bivariate threshold ECM (Table 2 Panel B) and their associated information shares are reported in the subsequent columns, Table 3. The estimated information shares of the NYSE (IS^N) range from 2% to 94%, with an average of 43.5%. Compared to the results from the linear ECM, overall, the NYSE makes larger contributions to the price discovery. There appears to be an upward trend effect through the sample period as the median estimates of IS^N in 1998, 1999, 2000 are 0.435, 0.51 and 0.54, respectively. The data reveals that, over time, the NYSE gained influence on setting equilibrium prices of the cross-listed pairs. In Table 2 Panel B, the estimated thresholds (γ_i) range from 0.009 to 0.545 with a mean of 0.146: that is, when the cross-listing dollar premium/discount records more than 14.6 cents, respectively, arbitrageurs begin to take positions on both sides and drive the deviation back into the “no-arbitrage” band.

After considering the threshold effect, we find some evidence that, the information share of the NYSE in the outer regime is larger than in the inner regime. One possible explanation is that informed traders choose to trade on the TSX when the price deviation is small but they migrate to the NYSE if the deviation is large enough to compensate for the cost of changing trading venues.

5.2. Regression analyses

Figuerola-Ferretti and Gonzalo (2009) model and measure price discovery on the NYMEX and IPE crude oil markets. The two contract prices co-move relatively closely, but transportation costs and grade differences pose potential difficulties in determining arbitrage opportunities. They investigate two interesting questions: (1) How does arbitrage ensure adjustment to the long run path given location and grade differences ?; and (2) Which of the markets is the market leader, or the most important contributor to price discovery?

5.2.1. Regression of the information share

We construct a panel data to analyze the factors that affect the relative extent of the NYSE’s contribution to price discovery. The estimated outer-regime information shares are regressed onto the panel of explanatory and control variables with and without intercept in Panel A and Panel B of Table 4, respectively. It turns out that the contribution of the NYSE increases relatively against that of the TSX as the NYSE-based trades become more informative (PIN). This is cross-border evidence that informed trades contribute to fostering price discovery, in line with Chen and Choi (2010). Either in quantity or value, the higher the liquidity on the NYSE the more it leads in price discovery. This is consistent with Eun and Sabherwal's (2003) findings: they estimate the information share of the NYSE by using Harris et al.'s (1995, 2002) approach. They find that the information share is directly related to the U.S.'s share of total trading (*USVol*), the proportion of informative trades on U.S. exchanges and the TSX (confirmed as proxied by the PIN), and the inversely related to the ratio of bid-ask spreads on U.S. exchanges and the TSX, which is not discernable in Table 4.[16] A better investor-protecting (*Governance*) and larger (*Size*) Canadian firm tends to lead price setting on the TSX as seen in Models 1 through 22 in Panels A and B. The overall explanatory power is significantly higher with models without intercept.

[Insert Tables 5 and 6 about here.]

We conduct analoguous panel regressions for the inner-regime and linear information shares in Tables 5 and 6, respectively. Neither alternative measure of exchange-specific contribution to price dicovery has a higher explanatory power (adjusted R^2) and economically and statistically meaningful implications. From this end, the outer-regime information shares (Table 4) have not only proved heuristically appealing but also economically reasonable and statistically robust.

5.2.2. Regression of the estimated threshold

For each cross-listed pair, the threshold includes transactions costs, which consist of bid-ask price spreads on both exchanges and the foreign exchange rate, fixed costs, and liquidity shorfalls. Implicit risk premiums, including those from information asymmetry and macroeconomic uncertainty, can also affect the determination of the threshold. Accordingly,

[16]Hasbrouck (1995) finds that there is a positive and significant correlation between contribution to price discovery made by the NYSE and its market share by trading volume using the U.S. domestic data. Using the same data, Harris et al. (2002) finds evidence that the information share increases when its bid-ask spreads decline relative to the regional exchange.

Table 7 provides the results of panel regressions of the estimated thresholds onto average (Panel A) and difference (Panel B) measures of asymmetric information component (PIN) and the inverse of market depth (spread), controlling for liquidity, either in quantity (*UsVol*) or value (*UsDollarVol*), firm-level idiosyncratic characteristics (*Industry*, *Governance*, and *Size*), and interest rates (yields of 90-day bills and 10-year notes).

[Insert Table 7 about here.]

As expected, our measure of market friction (relative quoted spread) significantly increases required dollar return of cross-border arbitrage as 8 out of 16 models using average measures (Panel A) and all models using difference measures (Panel B) agree with it. The better the firm is governed at home, the lower the minimum required profit as all models with the *Governance* control variable show. Manufacturing firms (when *Industry* equals 1) tend to require larger relative premiums to be exploited. Overall, difference measures turn out to have a greater determination on the threshold level than the average measures do as the adjusted R^2's of Panel B dominate those of Panel A through all specifications. In sum, the effective break-even point (threshold) of cross-border arbitrage appears to be affected by the relative degree of private information, market friction, and liquidity measures, and idiosyncratic firm-level characteristics. These, much economically appealing, empirical results lend support to the findings of Gagnon and Karolyi (2010).

6. Conclusion

For a pair of the original listing and its cross-listing, the adjustment to parity can be discontinuous: convergence may be quicker when the relative premium is profitable, or slower otherwise. In other words, the dynamics of cross-listed pairs fall into two regimes: within and beyond the threshold, e.g. transaction costs and associated risk premiums of arbitrage. This paper extends Harris et al.'s (1995, 2002) ECM to estimate the extent of contribution to price discovery (information share) by considering threshold cointegration per Balke and Fomby (1997).

The existing methods assume linear convergence of relative premiums to parity whereas we hinge our premise on the reality that premiums disappear faster when it is profitably arbitrageable than otherwise. A large deviation (outer regime) is believed to be more susceptible to new information, either public or private. In contrast, a small deviation (inner regime) can be due to noise trading and, therefore, there is little connection between price discovery and error correction dynamics.

Our threshold ECM ideally incorporates such a dichotomy while the predecessor linear ECMs may overestimate the information share when there is no cointegration in the unprofitable inner regime. Also, we find that the estimated information share and threshold are typically affected by the relative degree of private information, market friction and liquidity measures, and idiosyncratic firm-level characteristics. Unlike Grammig et al. (2005), we do not account for exchange-rate market friction in our threshold ECM framework. We invite readers to augment additional sources of randomness to the modelling of the nonlinear dynamics of cross-listed stocks.

References

Akaike, H., 1974. A new look at the statistical model identification. IEEE Transactions on Automatic Control 19, 716–723.

Bacidore J., Sofianos G., 2002. Liquidity provision and specialist trading in NYSE-listed non-U.S. stocks. Journal of Financial Economics 63, 133-158.

Balke, N.S., Fomby, T.B., 1997. Threshold cointegration. International Economic Review 38, 627-645.

Chan, K., Menkveld, A.J., Yang, Z., 2008. Infomration asymmetry and asset prices: Evidence from the China foreign share discount. The Journal of Finance 63, 159-196.

Chen, H., Choi, P.M.S., 2010. A trans-Niagara tale of informed traders. Working paper, Cornell University.

Chung, H., Ho, T.-W., Wei, L.-J., 2005. The dynamic relationship between the prices of ADRs and their underlying stocks: Evidence from the threshold vector error correction model. Applied Economics 37, 2387-2394.

De Jong, F., 2002. Measures of contributions to price discovery: A comparison. Journal of Financial Markets 5, 323-327.

Dickey, D.A., Fuller, W.A., 1979. Distribution of the estimators for autoregressive time series with a unit root. Journal of the American Statistical Association 74, 427–431.

Doidge, C., Karolyi, G.A., Stulz, R.M., 2009. Has New York become less competitive in global markets? Evaluating foreign listing choices over time. Journal of Financial Economics 91, 253-287.

Duarte, J., Young, L., 2009. Why is PIN priced? Journal of Financial Economics 91, 119-138.

Easley, D., Engle, R.F., O' Hara, M., Wu, L., 2008. Time-varying arrival rates of informed and uninformed trades. Journal of Financial Ecnometrics 6, 171-207.

Easley, D., Hvidkjaer, S., O'Hara, M., 2002. Is information risk a determinant of asset returns? The Journal of Finance 57, 2185-2221.

Easley, D., Kiefer, N.M., O'Hara, M., 1997. One day in the life of a very common stock. Review of Financial Studies 10, 805-835.

Easley, D., Kiefer, N.M., O'Hara, M., 1997. The information content of the trading process. Journal of Empirical Finance 4, 159-186.

Easley, D., Kiefer, N.M., O'Hara, M., Paperman, J.B., 1996. Liquidity, information, and infrequently traded stocks. The Journal of Finance 51, 1405-1436.

Easley, D., O'Hara, M., Saar, G., 2001. How stock splits affect trading: A microstructure approach. Journal of Financial and Quantitative Analysis 36, 25-51.

Engle, R.F., Granger, C.W.J., 1987. Co-integration and error correction: Representation, estimation and testing. Econometrica 35, 251-276.

Engle, R.F., Yoo, B.S., 1987. Forecasting and testing in co-integrated systems. Journal of Econometrics 35, 143-159.

Eun, C., Sabherwal, S., 2003. Cross-border listings and price discovery: Evidence from U.S.-listed Canadian stocks. The Journal of Finance 58, 549-574.

Figuerola-Ferretti, I., Gonzalo, J., 2009. Modelling and measuring price discovery in commodity markets. Journal of Econometrics, forthcoming.

Fuller, K., Van Ness, B., Van Ness, R. Is information risk priced for NASDAQ-listed securities? Working paper, University of Mississippi.

Gagnon, L., Karolyi, G. A., 2010. Multi-market trading and arbitrage. Journal of Financial Economics, forthcoming.

Garbade, K.D., Silber, W.L., 1983. Price movements and price discovery in futures and cash markets. The Review of Economics and Statistics 65, 289-297.

Gonzalo, J., Granger, C.W.J., 1995. Estimation of common long-memory components in cointegrated systems. Journal of Business and Economic Statistics 13, 1-9.

Grammig, J., Melvin, M., Schlag, C., 2005. Internationally cross-listed stock prices during overlapping trading hours: Price discovery and exchange rate effects. Journal of Empirical Finance 12, 139-164.

Grossman, S., Stiglitz, J, 1980. On the impossibility of informationally efficient markets. American Economic Review 70, 393-408.

Hansen, B.E., Seo, B., 2002. Testing for two-regime threshold cointegration in vector error correction models. Journal of Econometrics 110, 293-318.

Harris, F.H.deB., McInish, T.H., Shoesmith, G., Wood, R.A., 1995. Cointegration, error correction, and price discovery on informationally linked security markets. Journal of Financial and Quantitative Analysis 30, 563-579.

Harris, F.H.deB., McInish, T.H., Wood, R.A., 1997. Common long-memory components of intraday stock prices: A measure of price discovery. Working paper, Wake Forest University.

Harris, F.H.deB., McInish, T.H., Wood, R.A., 2002. Security price adjustment across exchanges: An investigation of common factor components for Dow stocks. Journal of Financial Markets 5, 277-308.

Hasbrouck, J., 1995. One security, many markets: determining the contributions to price discovery. Journal of Finance 50, 1175-1201.

Hasbrouck, J., 2002. Stalking the efficient price in market microstructure specifications: An overview. Journal of Financial Markets 5, 329-339.

Johansen, S., 1991. Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models. Econometrica 59, 1551-1580.

Koumkwa, S., Susmel, R., 2005. Arbitrage and convergence: evidence from Mexican ADRs. Working paper, University of Houston.

Krugman, P.R., 1991. Target zones and exchange rate dynamics. Quarterly Journal of Economics 106, 669-682.

Kyle, A., 1985. Continuous auctions and insider trading. Econometrica 53, 1315-1335.

Lee, C.M.C., Ready, M.J., 1991. Inferring trade direction from intraday data. The Journal of Finance 46, 733-746.

McFarland, J., 2002. Report of Business. Globe and Mail, 7 October, B6.

Rabinovitch, R., Silva, A.C., Susmel, R., 2003. Impact of capital controls and transaction costs on the return distribution of dually traded securities: Evidence from Chile and Argentina. Working paper, University of Houston.

Schreiber, P.S., Schwartz, R.A., 1986. Price discovery in securites markets. Journal of Portfolio Management 12, 43-48.

Schultz, P., Shive, S., 2008. Mispricing of dual class shares: Profit opportunities, arbitrage, and trading. Working paper, University of Notre Dame.

Schwarz, G., 1978. Estimating the dimension of a model. The Annals of Statistics 6, 461-464.

Solnik, B., Boucrelle, C., Le Fur, Y., 1996. International market correlation and volatility. Financial Analysts Journal 52, 17-34.

Appendix A. Estimation and testing of parameters

For convenience, the firm indicator (i) is selectively omitted in the following discussion. The threshold ecm aforementioned in Section 2 can be represented as follows:

$$\Delta x_t = A_1^{'} X_{t-1} d_{1t}(\gamma) + A_2^{'} X_{t-1} d_{2t}(\gamma) + u_t, \quad (A.94)$$

where $\Delta x_t = (p_{it}^{\mathrm{n}}, \tilde{p}_{it}^{\mathrm{t}})$, $X_{t-1} = [1, \kappa_{t-1}, \Delta x_{t-1}, \Delta x_{t-2}, .. \Delta x_{t-m}]^{'}$, $d_{1t}(\gamma) = \mathbf{1}(|\kappa_{it-1}| \leq \gamma_i)$ and $d_{1t}(\gamma) = \mathbf{1}(|\kappa_{it-1}| > \gamma_i)$; $A_1^{'}$ and $A_2^{'}$ contain the parameters to be estimated; and γ is the threshold parameter to be estimated.

The threshold VECM can be estimated using the MLE method proposed by Hansen and Seo (2002). Assuming that the error term (u_t) is *i.i.d.* Gaussian, the likelihood function is

$$\mathcal{L}_n(A_1, A_2, \Sigma, \gamma) = -\frac{n}{2}\ln|\Sigma| - \frac{1}{2}\sum_{t=1}^{n} u_t(A_1, A_2, \gamma)^{'} \Sigma^{-1} u_t(A_1, A_2, \gamma), \quad (A.95)$$

where $u_t(A_1, A_2, \gamma) = \Delta x_t - A_1^{'} X_{t-1} d_{1t}(\gamma) - A_2^{'} X_{t-1} d_{2t}(\gamma)$. The covariance matrix (Σ) is an identity matrix due to the *i.i.d.* Gaussian assumption of the error term. For a fixed c, A_1 and A_2 can estimated by an OLS regression, thus

$$\hat{A}_1(\gamma) = \left(\sum_{t=1}^{n} X_{t-1} X_{t-1}^{'} d_{1t}(\gamma)\right)^{-1} \sum_{t=1}^{n} X_{t-1} \Delta x_t^{'} d_{1t}(\gamma), \quad (A.96)$$

$$\hat{A}_2(\gamma) = \left(\sum_{t=1}^{n} X_{t-1} X_{t-1}^{'} d_{2t}(\gamma)\right)^{-1} \sum_{t=1}^{n} X_{t-1} \Delta x_t^{'} d_{2t}(\gamma), \quad (A.97)$$

and then $\hat{u}_t(\gamma) = \Delta x_t - \hat{A}_1^{'} X_{t-1} d_{1t}(\gamma) - \hat{A}_2^{'} X_{t-1} d_{2t}(\gamma)$. By plugging $\hat{u}_t(\gamma)$, the likelihood function $(\mathcal{L}_n(A_1, A_2, \Sigma, \gamma))$ becomes a univariate function of γ:

$$\mathcal{L}_n(\gamma) = \frac{-n}{2}\ln\left(\frac{1}{n}\sum_{t=1}^{n} \hat{u}_t(\gamma)\hat{u}_t(\gamma)^{'}\right) - \frac{n(m+2)}{2}. \quad (A.98)$$

Following Hansen (2000), the *grid search* method can be used to estimate γ within an preset interval $[\underline{\gamma}, \overline{\gamma}]$. The mle estimators for A_1 and A_2 can be obtained by inserting $\hat{\gamma}$. To further confirm the threshold effect, we need to test the following null hypothesis:

$H_0: A_1 = A_2$ for any $\gamma \in [\underline{\gamma}, \overline{\gamma}]$ (A.99)

against

$H_1: A_1 \neq A_2$ for some $\gamma \in [\underline{\gamma}, \overline{\gamma}]$. (A.100)

We use the super-Lagrangian multiplier (supLM) test (Hansen and Seo, 2002) to test the above hypotheses. The LM statistic is

$$\mathcal{LM}(\gamma) = \left(\hat{A}_1(\gamma) - \hat{A}_2(\gamma)\right)' \left(\hat{V}_1(\gamma) + \hat{V}_2(\gamma)\right)^{-1} \left(\hat{A}_1(\gamma) - \hat{A}_2(\gamma)\right), \tag{A.101}$$

where $\hat{V}_1(\gamma) = M_j(\gamma)^{-1}\Omega_j(\gamma)M_j(\gamma)^{-1}, M_j(\gamma) = I_{m+2} \otimes \Pi_j(\gamma)'\Pi_j(\gamma)$; and$\Omega_j(\gamma) = \Gamma_j(\gamma)'\Gamma_j(\gamma)$, and $\Pi_j(\gamma), \Gamma_j(\gamma)$ are matrices of the stacked rows of $X_{t-1}d_{jt}(\gamma)$ and $\hat{u}_t(\gamma) \otimes X_{t-1}d_{jt}(\gamma)$, respectively. Define

$\sup\mathcal{LM} = \sup_{\gamma \in [\underline{\gamma}, \overline{\gamma}]} \mathcal{LM}(\gamma)$. (A.102)

A bootstrap method is used to generate the critical value since the asymptotic distribution is non-standard.

Appendix B. Proofs

Proof of the adjustment coefficients

The solutions of the adjustment coefficients are:

$$\alpha_1^S = \frac{h}{(1+r)}\left\{\frac{\beta^I\pi_2 N_2\beta^A+\beta^I\pi_1 N_1(D_2+\beta^A)}{D_2D_1+\beta^A(D_1+D_2)}\right\}, \tag{B.103}$$

$$\alpha_{11}^Y = \frac{(\beta^A+D_2)}{D_2D_1+\beta^A(D_1+D_2)}, \tag{B.104}$$

$$\alpha_{12}^Y = \frac{\beta^A}{D_2D_1+\beta^A(D_1+D_2)}, \tag{B.105}$$

$$\alpha_2^S = \frac{h}{(1+r)}\left\{\frac{\beta^I\pi_1 N_1\beta^A+\beta^I\pi_2 N_2(D_1+\beta^A)}{D_2D_1+\beta^A(D_1+D_2)}\right\}, \tag{B.106}$$

$$\alpha_{21}^Y = \frac{\beta^A}{D_2D_1+\beta^A(D_1+D_2)}, \tag{B.107}$$

$$\alpha_{22}^Y = \frac{(\beta^A+D_1)}{D_2D_1+\beta^A(D_1+D_2)}, \tag{B.108}$$

thus

$$h_{11} = \frac{\alpha_{11}^Y}{\alpha_1^S} = \left(\frac{1+r}{h}\right)\frac{(\beta^A+D_2)}{\beta^I\pi_2 N_2\beta^A+\beta^I\pi_1 N_1(D_2+\beta^A)}, \tag{B.109}$$

$$h_{12} = \frac{\alpha_{12}^Y}{\alpha_1^S} = \left(\frac{1+r}{h}\right)\frac{\beta^A}{\beta^I\pi_2 N_2\beta^A+\beta^I\pi_1 N_1(D_2+\beta^A)}, \tag{B.110}$$

$$h_{21} = \frac{\alpha_{21}^Y}{\alpha_2^S} = \left(\frac{1+r}{h}\right)\frac{\beta^A}{\beta^I\pi_1 N_1\beta^A+\beta^I\pi_2 N_2(D_1+\beta^A)}, \tag{B.111}$$

$$h_{22} = \frac{\alpha_{22}^Y}{\alpha_2^S} = \left(\frac{1+r}{h}\right)\frac{(\beta^A+D_1)}{\beta^I\pi_1 N_1\beta^A+\beta^I\pi_2 N_2(D_1+\beta^A)}. \tag{B.112}$$

Since

$$\phi_i = \frac{\tau_1\tau_2}{\tau_1\tau_2+h_{i1}^2\tau_2\tau_s+h_{i2}^2\tau_1\tau_s}, \tag{B.113}$$

we have

$$\frac{1}{\phi_1} = 1+\frac{h_{11}^2\tau_2\tau_s+h_{12}^2\tau_1\tau_s}{\tau_1\tau_2}$$

$$= 1 + \left(\frac{\tau_s}{\tau_1}\right)\left[\left(\frac{1+r}{h}\right)\left\{\frac{(\beta^A + D_2)}{\beta^I\pi_2 N_2\beta^A + \beta^I\pi_1 N_1(D_2 + \beta^A)}\right\}\right]^2$$

$$+ \left(\frac{\tau_s}{\tau_2}\right)\left[\left(\frac{1+r}{h}\right)\left\{\frac{\beta^A}{\beta^I\pi_2 N_2\beta^A + \beta^I\pi_1 N_1(D_2 + \beta^A)}\right\}\right]^2$$

$$= 1 + \left(\frac{\tau_s}{\tau_1\tau_2}\right)\left(\frac{1+r}{h}\right)^2\left[\frac{\tau_2(\beta^A + D_2)^2 + \tau_1(\beta^A)^2}{\{\beta^I\pi_2 N_2\beta^A + \beta^I\pi_1 N_1(D_2 + \beta^A)\}^2}\right]$$

$$\equiv g(D_2) \tag{B.114}$$

and

$$\frac{1}{\phi_2} = 1 + \frac{h_{21}^2\tau_2\tau_s + h_{22}^2\tau_1\tau_s}{\tau_1\tau_2}$$

$$= 1 + \left(\frac{\tau_2\tau_s}{\tau_1\tau_2}\right)\left[\left(\frac{1+r}{h}\right)\left\{\frac{\beta^A}{\beta^I\pi_1 N_1\beta^A + \beta^I\pi_2 N_2(D_1 + \beta^A)}\right\}\right]^2$$

$$+ \left(\frac{\tau_1\tau_s}{\tau_1\tau_2}\right)\left[\left(\frac{1+r}{h}\right)\left\{\frac{(\beta^A + D_1)}{\beta^I\pi_1 N_1\beta^A + \beta^I\pi_2 N_2(D_1 + \beta^A)}\right\}\right]^2$$

$$= 1 + \left(\frac{\tau_s}{\tau_1\tau_2}\right)\left(\frac{1+r}{h}\right)^2\frac{\tau_1(\beta^A + D_1)^2 + \tau_2(\beta^A)^2}{(\beta^I\pi_1 N_1\beta^A + \beta^I\pi_2 N_2(D_1 + \beta^A))^2}$$

$$\equiv g(D_1). \tag{B.115}$$

Now, recall the definitions $D_1 \equiv \beta^I\pi_1 N_1 + \beta_1^U(1-\pi_1)N_1$ and $D_2 \equiv \beta_2^U(1-\pi_2)N_2 + \beta^I\pi_2 N_2$, where

$$\beta_i^U = \left\{1 + r - \frac{1}{\alpha_i^S}\left(\frac{\phi_i\tau_s}{\tau_v + \phi_i\tau_s}\right)\right\}\Big/\left(\frac{\rho}{\tau_v + \phi_i\tau_s} + c\right), \tag{B.116}$$

$$\beta^I = (1+r)\Big/\left(\frac{\rho}{\tau_s + \tau_v} + c\right), \tag{B.117}$$

$$\phi_i = \frac{\tau_1\tau_2}{\tau_1\tau_2 + h_{i2}^2\tau_2\tau_s + h_{i3}^2\tau_1\tau_s}, h_{i2} = \frac{\alpha_{i2}}{\alpha_{i1}}, \text{and } h_{i3} = \frac{\alpha_{i3}}{\alpha_{i1}}, \tag{B.118}$$

thus

$$D_1 \equiv \beta^I\pi_1 N_1 + \beta_1^U(1-\pi_1)N_1$$

$$= \beta^I \pi_1 N_1 + \frac{\tau_v + \phi_1 \tau_s}{(\rho + c\tau_v + c\phi_1 \tau_s)} \left(1 + r - \frac{1}{\alpha_1^S} \left(\frac{\phi_1 \tau_s}{\tau_v + \phi_1 \tau_s} \right) \right) (1 - \pi_1) N_1$$

$$= \beta^I \pi_1 N_1 + \frac{\tau_v + \phi_1 \tau_s}{(\rho + c\tau_v + c\phi_1 \tau_s)} (1 + r)(1 - \pi_1) N_1 - \eta \frac{\phi_1 \tau_s}{\alpha_1^S} (1 - \pi_1) N_1, \qquad \text{(B.119)}$$

hence

$$\phi_1 = f(D_1, \alpha_1^S), \qquad \text{(B.120)}$$

and similarly

$$\phi_2 = f(D_2, \alpha_2^S). \qquad \text{(B.121)}$$

Combining above equations, we can numerically solve $\widehat{D}_1$ and $\widehat{D}_2$ from the following equations:

$$1 = g(D_2) f(D_1, \alpha_1^S) \text{ and } 1 = g(D_1) f(D_2, \alpha_2^S). \qquad \text{(B.122)}$$

Proof of Proposition 1. Since

$$\beta_i^U = \left\{ 1 + r - \frac{1}{\alpha_i^S} \left(\frac{\phi_i \tau_s}{\tau_v + \phi_i \tau_s} \right) \right\} \Big/ \left(\frac{\rho}{\tau_v + \phi_i \tau_s} + c \right), \qquad \text{(B.123)}$$

$$\beta^I = (1 + r) \Big/ \left(\frac{\rho}{\tau_s + \tau_v} + c \right) \qquad \text{(B.124)}$$

and

$$\phi_i = \frac{\tau_1 \tau_2}{\tau_1 \tau_2 + h_{i2}^2 \tau_2 \tau_s + h_{i3}^2 \tau_1 \tau_s} \text{ and } h_{i2} = \alpha_{i2}/\alpha_{i1}, h_{i3} = \alpha_{i3}/\alpha_{i1}, \qquad \text{(B.125)}$$

we can reach the conclusion from $0 < \phi_i < 1$ and $1 + r - \frac{1}{\alpha_{i1}} \left(\frac{\phi_i \tau_s}{\tau_v + \phi_i \tau_s} \right) < (1 + r)$.

Proof of Proposition 2. From Proposition 1, we see $\beta^I > \beta_i^U$. Since $D_1 \equiv \beta^I \pi_1 N_1 + \beta_1^U (1 - \pi_1) N_1$ and $D_2 \equiv \beta_2^U (1 - \pi_2) N_2 + \beta^I \pi_2 N_2$, $PIN_1 = \frac{\beta^I \pi_1 N_1}{D_1}$, and $PIN_2 = \frac{\beta^I \pi_2 N_2}{D_2}$, thus we can show that $D_1 > D_2$. It follows that $IS_1 > IS_2$.

Table 1

Estimated cointegrating vector.

	b^{N}	$t-$ stat.
5 %-ile	0.90	-5.25
25 %-ile	0.995	-1.29
Median	0.999	0.25
75 %-ile	1.002	0.99
95 %-ile	1.011	2.94

Notes: The prices of the sample TSX-NYSE Cross-listed pairs ($\{p_{it}^{\mathrm{N}}, \tilde{p}_{it}^{\mathrm{T}}\}$) are tested for cointegration per Johansen (1991), where p_{it}^{N} and $\tilde{p}_{it}^{\mathrm{T}}$ are the actual trade and parity-implied prices of the cross-listing on the NYSE. Since we have two price series in each regression equation in the cointegrated system, there is at most one cointegrating vector. We estimate the cointegrating vector for each stock in each year while normalizing $b^{\mathrm{T}} = -1$. Our results show that most of the estimated cointegrating vectors are $(1, -1)^T$, which is of the expected values according to the law of one price. The *t*-statistics for the null hypothesis attests that the cointegrating vector equals $(1, -1)^T$. The observations are in firm-years.

Table 2
Parameter estimates and nonlinearity test statistics.
Panel A: Two-regime threshold ECM parameter estimates.

	α_{in}^{N}	*t*-stat.	α_{in}^{T}	*t*-stat.	α_{out}^{N}	*t*-stat.	α_{out}^{T}	*t*-stat.	α_{out}^{N} - α_{in}^{N}	*t*-stat.	α_{out}^{T} - α_{in}^{T}	*t*-stat.
Mean	-0.241	-6.069	0.129	2.967	-0.216	-5.415	0.173	3.541	0.025	0.236	0.044	0.510
St. Dev.	0.247	3.606	0.279	2.868	0.127	3.251	0.143	2.562	0.248	1.997	0.275	1.535
1%-ile	-1.291	-12.884	-0.709	-1.948	-0.636	-15.309	-0.061	-1.152	-0.528	-3.737	-0.815	-3.003
25%-ile	-0.288	-8.659	0.037	0.983	-0.275	-6.823	0.067	1.670	-0.064	-1.039	-0.044	-0.598
50%-ile	-0.212	-6.381	0.106	2.477	-0.212	-5.041	0.150	3.473	0.006	0.087	0.045	0.631
75%-ile	-0.133	-3.256	0.214	4.638	-0.131	-3.080	0.247	5.164	0.083	1.327	0.109	1.545
99%-ile	0.207	1.010	0.974	10.741	-0.001	-0.041	0.568	10.157	1.052	6.327	0.940	3.967

Panel B: Threshold estimates and supLM test statistics.

	Threshold	supLM	95%-ile critical value.	*p*-value
Mean	0.146	22.321	22.075	0.259
St. Dev.	0.109	16.328	2.291	0.270
1%-ile	0.009	8.730	18.197	0.000
25%-ile	0.078	15.208	20.961	0.020
50%-ile	0.118	18.876	21.709	0.170
75%-ile	0.198	24.461	22.502	0.420
99%-ile	0.545	69.782	29.065	0.965

Panel C: Wald statistics.

	Wald_{ECM}	*p*-value	Wald_{DC}	*p*-value
Mean	7.683	0.884	11.962	0.928
St. Dev.	10.173	0.182	17.161	0.128
1%-ile	0.056	0.187	1.257	0.467
25%-ile	2.117	0.854	5.478	0.934
50%-ile	4.586	0.968	8.928	0.988
75%-ile	9.540	0.998	13.832	0.999
99%-ile	44.615	1.000	96.717	1.000

Notes: We estimate the adjustment coefficients based on our threshhold ECM framework following Balke and Fomby (1997) and extended from Harris et al. (1995, 2002):

$$\Delta p_{it}^{\mathrm{N}} = \begin{cases} \beta_{10}^{\mathrm{N}} + \alpha_{\mathrm{out}}^{\mathrm{N}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{1j}^{\mathrm{N}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{1j}^{\mathrm{N}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \beta_{20}^{\mathrm{N}} + \alpha_{\mathrm{in}}^{\mathrm{N}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{2j}^{\mathrm{N}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{2j}^{\mathrm{N}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| \le \gamma_i \end{cases},$$

$$\Delta \tilde{p}_{it}^{\mathrm{T}} = \begin{cases} \beta_{10}^{\mathrm{T}} + \alpha_{\mathrm{out}}^{\mathrm{T}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{1j}^{\mathrm{T}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{1j}^{\mathrm{T}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \beta_{20}^{\mathrm{T}} + \alpha_{\mathrm{in}}^{\mathrm{T}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{2j}^{\mathrm{T}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{2j}^{\mathrm{T}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| \le \gamma_i \end{cases}.$$

α_{in} measures the adjustment coefficient of an exchange when the cross-listing dollar premium is within the range of thresholds (γ_i), thus cross-border arbitrage is unprofitable; and α_{out} when beyond the range of thresholds, thus arbitrage forces will activate to drive the premium within the range. The threshold, κ_{it}, is factored in the following threshold autoregressive (TAR) framework:

$$\kappa_{it} = \begin{cases} \alpha_{\mathrm{out}} + \rho_{\mathrm{out}}\,\kappa_{it-1} + \varepsilon_t, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \alpha_{\mathrm{in}} + \rho_{\mathrm{in}}\,\kappa_{it-1} + \varepsilon_t, \text{ if } |\kappa_{it-1}| \le \gamma_i \end{cases},$$

where $\rho_{\mathrm{in}} = 1$ and $-1 < \rho_{\mathrm{out}} < 0$, i.e. the dollar premium (κ_{it}) of the cross-listing is a unit root in the (unprofitable) middle regime when $|\kappa_{it-1}| \le \gamma_i$, otherwise a mean-reverting process in the (profitable) outer regime when $|\kappa_{it-1}| > \gamma_i$ with presence of arbitrageurs. Wald_{ECM} gives the Wald statistic for the joint null hypothesis: $H_0 : \alpha_{\mathrm{out}}^{\mathrm{T}} = \alpha_{\mathrm{in}}^{\mathrm{T}}$ and $\alpha_{\mathrm{out}}^{\mathrm{N}} = \alpha_{\mathrm{in}}^{\mathrm{N}}$, while Wald_{DC} gives the wald-statistics for: $H_0 : \beta_{1j}^{\mathrm{N}} = \beta_{2j}^{\mathrm{N}}$, $\beta_{1j}^{\mathrm{N}} = \beta_{2j}^{\mathrm{N}}$, $\beta_{1j}^{\mathrm{T}} = \beta_{2j}^{\mathrm{T}}$, and $\beta_{1j}^{\mathrm{T}} = \beta_{2j}^{\mathrm{T}}$. All estimates have cointegrating vectors given as 1 ($b^{\mathrm{N}} = 1$). The observations are in firm-quarters.

Table 3
Estimates of information shares.

	Linear ECM	Threshold ECM		
	IS^N	$p^N - p^T$	IS_{in}^N	IS_{out}^N
Mean	0.407	-0.001	0.380	0.435
St. Dev.	0.258	0.020	0.254	0.259
1%-ile	0.010	-0.054	0.007	0.020
25%-ile	0.188	-0.013	0.166	0.215
50%-ile	0.409	0.000	0.356	0.418
75%-ile	0.582	0.009	0.551	0.626
99%-ile	0.975	0.055	0.975	0.980

Notes: We estimate the adjustment coefficients based on the linear ECM framework following Harris et al. (1995, 2002) and Eun and Sabherwal (2003):

$$\Delta p_{it}^{\mathrm{N}} = \beta_0^{\mathrm{N}} + \alpha^{\mathrm{N}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_j^{\mathrm{N}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_j^{\mathrm{N}}\Delta\tilde{p}_{it-j}^{\mathrm{T}},$$
$$\Delta\tilde{p}_{it}^{\mathrm{T}} = \beta_0^{\mathrm{T}} + \alpha^{\mathrm{T}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_j^{\mathrm{T}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_j^{\mathrm{T}}\Delta\tilde{p}_{it-j}^{\mathrm{T}},$$

where κ_{it-1} gives the remaining relative premium or cointegrating residual. α^{N} and α^{T} are the adjustment coefficients of the NYSE and TSX, respectively, that describe how much deviation will be subsequently adjusted to restore the long run equilibrium in each series. Subsequently, we estimate the adjustment coefficients based on our threshhold ECM framework following Balke and Fomby (1997) and extended from Harris et al. (1995, 2002):

$$\Delta p_{it}^{\mathrm{N}} = \begin{cases} \beta_{10}^{\mathrm{N}} + \alpha_{\mathrm{out}}^{\mathrm{N}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{1j}^{\mathrm{N}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{1j}^{\mathrm{N}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \beta_{20}^{\mathrm{N}} + \alpha_{\mathrm{in}}^{\mathrm{N}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{2j}^{\mathrm{N}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{2j}^{\mathrm{N}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| \le \gamma_i \end{cases},$$
$$\Delta\tilde{p}_{it}^{\mathrm{T}} = \begin{cases} \beta_{10}^{\mathrm{T}} + \alpha_{\mathrm{out}}^{\mathrm{T}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{1j}^{\mathrm{T}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{1j}^{\mathrm{T}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \beta_{20}^{\mathrm{T}} + \alpha_{\mathrm{in}}^{\mathrm{T}}\kappa_{it-1} + \sum_{j=1}^{m_1} \beta_{2j}^{\mathrm{T}}\Delta p_{it-j}^{\mathrm{N}} + \sum_{j=1}^{m_2} \tilde{\beta}_{2j}^{\mathrm{T}}\Delta\tilde{p}_{it-j}^{\mathrm{T}}, \text{ if } |\kappa_{it-1}| \le \gamma_i \end{cases}.$$

α_{in} measures the adjustment coefficient of an exchange when the cross-listing dollar premium is within the range of thresholds (γ_i), thus cross-border arbitrage is unprofitable; and α_{out} when beyond the range of thresholds, thus arbitrage forces will activate to drive the premium within the range. The threshold, κ_{it}, is factored in the following threshold autoregressive (TAR) framework:

$$\kappa_{it} = \begin{cases} \alpha_{\mathrm{out}} + \rho_{\mathrm{out}}\,\kappa_{it-1} + \varepsilon_t, \text{ if } |\kappa_{it-1}| > \gamma_i \\ \alpha_{\mathrm{in}} + \rho_{\mathrm{in}}\,\kappa_{it-1} + \varepsilon_t, \text{ if } |\kappa_{it-1}| \le \gamma_i \end{cases},$$

where $\rho_{\mathrm{in}} = 1$ and $-1 < \rho_{\mathrm{out}} < 0$, i.e. the dollar premium (κ_{it}) of the cross-listing is a unit root in the (unprofitable) middle regime when $|\kappa_{it-1}| \le \gamma_i$, otherwise a mean-reverting process in the (profitable) outer regime when $|\kappa_{it-1}| > \gamma_i$ with the presence of arbitrageurs. All estimates have cointegrating vectors given as 1 ($b^N = 1$). The observations are in firm-quarters.

Table 4

Panel regression results of outer-regime information shares.

Panel A: Panel regression with intercept

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9	Model 10	Model 11	Model 12	Model 13	Model 14	Model 15	Model 16
Intercept	0.651 ***	0.702 ***	0.632 ***	0.683 ***	0.262 ***	0.307 ***	0.206 ***	0.242 ***	0.639 ***	0.690 ***	0.621 ***	0.672 ***	0.264 ***	0.308 ***	0.209 ***	0.245 ***
	5.473	5.961	5.339	5.844	3.531	4.006	2.920	3.283	5.415	5.917	5.292	5.809	3.659	4.161	3.078	3.480
PinRatio	0.127 **	0.122 **	0.133 **	0.127 **	0.151 ***	0.136 **	0.179 ***	0.168 ***	0.132 **	0.128 **	0.137 **	0.132 **	0.157 ***	0.143 **	0.182 ***	0.173 ***
	2.303	2.156	2.412	2.246	2.661	2.294	3.273	2.938	2.425	2.289	2.525	2.373	2.794	2.450	3.385	3.070
SpreadRatio	0.001	0.002	0.002	0.002	0.000	-0.001	-0.002	-0.002	0.001	0.001	0.001	0.001	-0.001	-0.002	-0.002	-0.003
	0.340	0.367	0.572	0.573	-0.093	-0.214	-0.419	-0.555	0.125	0.168	0.353	0.370	-0.300	-0.394	-0.582	-0.688
UsVol	0.386 ***		0.358 ***		0.414 ***		0.454 ***		0.393 ***		0.366 ***		0.420 ***		0.455 ***	
	4.200		3.998		4.600		5.668		4.262		4.080		4.664		5.691	
UsDollarVol		0.300 ***		0.277 ***		0.282 ***		0.336 ***		0.308 ***		0.286 ***		0.290 ***		0.339 ***
		3.673		3.486		3.572		4.627		3.771		3.600		3.692		4.693
Industry	-0.054	-0.050							-0.051	-0.048						
	-1.282	-1.175							-1.221	-1.125						
Governance	-0.005 ***	-0.005 ***	-0.005 ***	-0.005 ***					-0.005 ***	-0.005 ***	-0.005 ***	-0.005 ***				
	-3.538	-3.833	-3.717	-3.980					-3.457	-3.760	-3.633	-3.905				
Size	-0.390 **	-0.403 **	-0.353 **	-0.368 **	-0.443 **	-0.473 **			-0.354 **	-0.368 **	-0.319 *	-0.335 *	-0.409 **	-0.439 **		
	-2.256	-2.295	-2.063	-2.122	-2.502	-2.585			-2.056	-2.104	-1.876	-1.941	-2.321	-2.417		
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	115	115	115	115	115	115	115	115	115	115	115	115
Adjusted R^2	0.277	0.252	0.273	0.249	0.207	0.154	0.203	0.144	0.271	0.247	0.268	0.246	0.204	0.154	0.205	0.149

Panel B: Panel regression without intercept

	Model 17	Model 18	Model 19	Model 20	Model 21	Model 22	Model 23	Model 24	Model 25	Model 26	Model 27	Model 28	Model 29	Model 30	Model 31	Model 32
Intercept																
PinRatio	0.127 **	0.122 **	0.133 **	0.127 **	0.151 ***	0.136 **	0.179 ***	0.168 ***	0.299 ***	0.313 ***	0.299 ***	0.313 ***	0.331 ***	0.351 ***	0.325 ***	0.343 ***
	2.303	2.156	2.412	2.246	2.661	2.294	3.273	2.938	5.934	5.938	5.960	5.961	10.721	11.072	11.506	11.858
SpreadRatio	0.001	0.002	0.002	0.002	0.000	-0.001	-0.002	-0.002	-0.002	-0.002	-0.001	-0.001	-0.001	-0.001	-0.001	-0.002
	0.340	0.367	0.572	0.573	-0.093	-0.214	-0.419	-0.555	-0.363	-0.356	-0.275	-0.315	-0.126	-0.160	-0.317	-0.356
UsVol	0.386 ***		0.358 ***		0.414 ***		0.454 ***		0.629 ***		0.614 ***		0.646 ***		0.623 ***	
	4.200		3.998		4.600		5.668		6.912		7.211		9.366		10.310	
UsDollarVol		0.300 ***		0.277 ***		0.282 ***		0.336 ***		0.514 ***		0.507 ***		0.514 ***		0.512 ***
		3.673		3.486		3.572		4.627		6.070		6.398		8.365		9.323
Industry	-0.054	-0.050							-0.023	-0.012						
	-1.282	-1.175							-0.480	-0.253						
Governance	-0.005 ***	-0.005 ***	-0.005 ***	-0.005 ***					0.001	0.001	0.001	0.001				
	-3.538	-3.833	-3.717	-3.980					0.957	0.938	0.845	0.909				
Size	-0.390 **	-0.403 **	-0.353 **	-0.368 **	-0.443 **	-0.473 **			-0.298	-0.315	-0.283	-0.306	-0.248	-0.251		
	-2.256	-2.295	-2.063	-2.122	-2.502	-2.585			-1.546	-1.574	-1.493	-1.560	-1.380	-1.339		
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	115	115	115	115	115	115	115	115	115	115	115	115
Adjusted R^2	0.872	0.868	0.871	0.867	0.854	0.845	0.858	0.848	0.837	0.825	0.838	0.827	0.838	0.823	0.849	0.836

Notes : The dependent variable is IsOut which is the outer-regime information share of the NYSE which is a relative measure of contribution made by the NYSE to price discovery of TSX-NYSE cross-listed pairs. Explanatory variables are: PinRatio is the ratio of the PIN of the NYSE over that of the TSX. SpreadRatio is the ratio of the relative quoted bid-ask spread of the NYSE over that of the TSX. UsVol is the average daily trading volume of the NYSE out of both of the NYSE and the TSX following Eun and Sabherwal (2003). UsDollarVol is the average daily dollar trading volume of the NYSE out of both of the NYSE and the TSX. Control variables. Governance is the Report on Business governance index of Canadian firms published by Globe and Mail (McFarland, 2002). Industry equals one if the cross-lister is a manufacturing firm, and zero otherwise. Size is the normalized average market capitalization on the TSX and the NYSE. The numerical values below the estimates are t-statistics. ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively. The observations are in firm-years.

Table 5

Panel regression results of inner-regime information shares.

Panel A: Panel regression with intercept

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9	Model 10	Model 11	Model 12	Model 13	Model 14	Model 15	Model 16
Intercept	-0.027	-0.026	-0.027	-0.026	-0.020	-0.018	-0.022 *	-0.021	0.001	0.001	0.000	-0.001	0.001	0.002	-0.001	-0.001
	-0.649	-0.631	-0.676	-0.661	-1.367	-1.275	-1.672	-1.633	0.032	0.025	-0.010	-0.022	0.146	0.164	-0.137	-0.161
PinRatio	-0.015	-0.016	-0.015	-0.016	-0.007	-0.008	0.008	0.007	0.000	-0.003	0.000	-0.003	0.003	0.000	0.016	0.013
	-0.370	-0.399	-0.372	-0.401	-0.190	-0.221	0.222	0.184	-0.012	-0.084	-0.012	-0.085	0.077	0.002	0.441	0.373
SpreadRatio	0.000	0.000	0.000	0.000	0.000	0.000	-0.001	-0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001
	-0.129	-0.086	-0.130	-0.087	-0.116	-0.070	-0.213	-0.165	0.376	0.441	0.378	0.444	0.342	0.411	0.323	0.395
UsVol	0.225		0.225		0.200		0.234 *		0.242		0.243		0.219		0.250 **	
	1.516		1.525		1.458		1.886		1.617		1.636		1.587		2.017	
UsDollarVol		0.222		0.222		0.213		0.247 *		0.249		0.250		0.241		0.263 **
		1.473		1.482		1.467		1.911		1.632		1.648		1.653		2.017
Industry	0.000	-0.001							-0.003	-0.003						
	-0.017	-0.034							-0.171	-0.192						
Governance	0.000	0.000	0.000	0.000					0.000	0.000	0.000	0.000				
	0.042	0.066	0.042	0.065					0.039	0.061	0.036	0.057				
Size	-0.030	-0.036	-0.029	-0.035	-0.033	-0.036			-0.032	-0.039	-0.029	-0.035	-0.032	-0.035		
	-0.387	-0.471	-0.395	-0.477	-0.464	-0.512			-0.411	-0.503	-0.385	-0.475	-0.439	-0.497		
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	121	121	131	131	115	115	115	115	121	121	131	131
Adjusted R^2	0.015	0.014	0.025	0.023	0.018	0.018	0.044	0.044	-0.023	-0.022	-0.014	-0.013	-0.006	-0.004	0.012	0.012

Panel B: Panel regression without intercept

	Model 17	Model 18	Model 19	Model 20	Model 21	Model 22	Model 23	Model 24	Model 25	Model 26	Model 27	Model 28	Model 29	Model 30	Model 31	Model 32
Intercept																
PinRatio	-0.015	-0.016	-0.015	-0.016	-0.007	-0.008	0.008	0.007	0.000	-0.003	0.000	-0.003	0.003	0.000	0.016	0.013
	-0.370	-0.399	-0.372	-0.401	-0.190	-0.221	0.222	0.184	-0.012	-0.084	-0.012	-0.085	0.076	0.000	0.443	0.375
SpreadRatio	0.000	0.000	0.000	0.000	0.000	0.000	-0.001	-0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001
	-0.129	-0.086	-0.130	-0.087	-0.116	-0.070	-0.213	-0.165	0.378	0.443	0.380	0.446	0.344	0.414	0.324	0.396
UsVol	0.225		0.225		0.200		0.234 *		0.242		0.243		0.221		0.249 **	
	1.516		1.525		1.458		1.886		1.624		1.644		1.621		2.023	
UsDollarVol		0.222		0.222		0.213		0.247 *		0.249		0.250		0.244 *		0.263 **
		1.473		1.482		1.467		1.911		1.639		1.656		1.684		2.020
Industry	0.000	-0.001							-0.003	-0.003						
	-0.017	-0.034							-0.169	-0.192						
Governance	0.000	0.000	0.000	0.000					0.000	0.000	0.000	0.000				
	0.042	0.066	0.042	0.065					0.196	0.240	0.107	0.148				
Size	-0.030	-0.036	-0.029	-0.035	-0.033	-0.036			-0.032	-0.039	-0.029	-0.035	-0.026	-0.029		
	-0.387	-0.471	-0.395	-0.477	-0.464	-0.512			-0.412	-0.505	-0.386	-0.477	-0.427	-0.483		
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	121	121	131	131	115	115	115	115	121	121	131	131
Adjusted R^2	0.007	0.006	0.016	0.015	0.010	0.010	0.036	0.037	-0.022	-0.022	-0.013	-0.013	-0.006	-0.004	0.012	0.012

Notes: The dependent variable is IsIn which is the inner-regime information share of the NYSE which is a relative measure of contribution made by the NYSE to price discovery of TSX-NYSE cross-listed pairs. Explanatory variables are: PinRatio is the ratio of the PIN of the NYSE over that of the TSX. SpreadRatio is the ratio of the relative quoted bid-ask spread of the NYSE over that of the TSX. UsVol is the average daily trading volume of the NYSE out of both of the NYSE and the TSX following Eun and Sabherwal (2003). UsDollarVol is the average daily dollar trading volume of the NYSE out of both of the NYSE and the TSX. Control variables. Governance is the Report on Business governance index of Canadian firms published by Globe and Mail (McFarland, 2002). Industry equals one if the cross-lister is a manufacturing firm, and zero otherwise. Size is the normalized average market capitalization on the TSX and the NYSE. The numerical values below the estimates are t-statistics. ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively. The observations are in firm-years.

Table 6

Panel regression results of linear information shares.

Panel A: Panel regression with intercept

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9	Model 10	Model 11	Model 12	Model 13	Model 14	Model 15	Model 16
Intercept	0.015	0.014	0.013	0.011	0.014	0.014	0.019	0.019	-0.001	-0.001	-0.001	-0.001	0.001	0.002	0.000	0.001
	0.257	0.234	0.219	0.190	0.693	0.669	1.006	1.045	-0.010	-0.009	-0.014	-0.018	0.058	0.128	0.023	0.053
PinRatio	0.049	0.055	0.049	0.055	0.052	0.057	0.065	0.071	0.054	0.060	0.054	0.060	0.055	0.061	0.064	0.069
	0.868	0.974	0.871	0.978	0.952	1.051	1.258	1.370	0.966	1.095	0.971	1.100	1.023	1.148	1.265	1.366
SpreadRatio	0.002	0.001	0.002	0.001	0.002	0.001	0.001	0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.002	-0.001	-0.001
	0.346	0.248	0.344	0.245	0.353	0.248	0.324	0.248	-0.219	-0.315	-0.220	-0.316	-0.259	-0.360	-0.257	-0.323
UsVol	-0.153		-0.151		-0.128		-0.034		-0.205		-0.205		-0.182		-0.078	
	-0.716		-0.712		-0.658		-0.192		-0.957		-0.962		-0.929		-0.442	
UsDollarVol		-0.350		-0.348		-0.330		-0.189		-0.386 *		-0.386 *		-0.371 *		-0.221
		-1.626		-1.627		-1.618		-1.024		-1.783		-1.791		-1.799		-1.190
Industry	-0.004	-0.005							0.000	-0.001						
	-0.186	-0.211							-0.014	-0.035						
Governance	0.000	0.000	0.000	0.000					0.000	0.000	0.000	0.000				
	0.018	0.046	0.014	0.042					0.035	0.057	0.035	0.056				
Size	-0.030	-0.045	-0.025	-0.040	-0.019	-0.034			-0.014	-0.025	-0.013	-0.024	-0.009	-0.021		
	-0.269	-0.412	-0.236	-0.377	-0.191	-0.340			-0.122	-0.232	-0.122	-0.230	-0.089	-0.207		
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	121	121	131	131	115	115	115	115	121	121	131	131
Adjusted R^2	-0.014	0.006	-0.005	0.014	0.010	0.029	0.017	0.025	-0.038	-0.017	-0.029	-0.008	-0.018	0.002	-0.010	0.000

Panel B: Panel regression without intercept

	Model 17	Model 18	Model 19	Model 20	Model 21	Model 22	Model 23	Model 24	Model 25	Model 26	Model 27	Model 28	Model 29	Model 30	Model 31	Model 32
Intercept																
PinRatio	0.049	0.055	0.049	0.055	0.052	0.057	0.065	0.071	0.054	0.060	0.054	0.060	0.055	0.061	0.064	0.069
	0.868	0.974	0.871	0.978	0.952	1.051	1.258	1.370	0.971	1.100	0.975	1.105	1.027	1.151	1.270	1.371
SpreadRatio	0.002	0.001	0.002	0.001	0.002	0.001	0.001	0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.002	-0.001	-0.001
	0.346	0.248	0.344	0.245	0.353	0.248	0.324	0.248	-0.220	-0.317	-0.221	-0.318	-0.260	-0.361	-0.258	-0.324
UsVol	-0.153		-0.151		-0.128		-0.034		-0.205		-0.205		-0.181		-0.078	
	-0.716		-0.712		-0.658		-0.192		-0.961		-0.966		-0.932		-0.443	
UsDollarVol		-0.350		-0.348		-0.330		-0.189		-0.386 *		-0.386 *		-0.368 *		-0.220
		-1.626		-1.627		-1.618		-1.024		-1.791		-1.799		-1.802		-1.193
Industry	-0.004	-0.005							0.000	-0.001						
	-0.186	-0.211							-0.017	-0.039						
Governance	0.000	0.000	0.000	0.000					0.000	0.000	0.000	0.000				
	0.018	0.046	0.014	0.042					0.072	0.137	0.087	0.158				
Size	-0.030	-0.045	-0.025	-0.040	-0.019	-0.034			-0.014	-0.025	-0.013	-0.024	-0.006	-0.014		
	-0.269	-0.412	-0.236	-0.377	-0.191	-0.340			-0.122	-0.233	-0.122	-0.231	-0.070	-0.165		
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	121	121	131	131	115	115	115	115	121	121	131	131
Adjusted R^2	-0.023	-0.003	-0.014	0.006	0.002	0.021	0.009	0.017	-0.038	-0.017	-0.028	-0.008	-0.018	0.002	-0.010	0.000

Notes : The dependent variable is IsLin which is the linear information share (Harris et al. (1995, 2002)) of the NYSE which is a relative measure of contribution made by the NYSE to price discovery of TSX-NYSE cross-listed pairs. Explanatory variables are: PinRatio is the ratio of the PIN of the NYSE over that of the TSX. SpreadRatio is the ratio of the relative quoted bid-ask spread of the NYSE over that of the TSX. UsVol is the average daily trading volume of the NYSE out of both of the NYSE and the TSX following Eun and Sabherwal (2003). UsDollarVol is the average daily dollar trading volume of the NYSE out of both of the NYSE and the TSX. Control variables. Governance is the Report on Business governance index of Canadian firms published by Globe and Mail (McFarland, 2002). Industry equals one if the cross-lister is a manufacturing firm, and zero otherwise. Size is the normalized average market capitalization on the TSX and the NYSE. The numerical values below the estimates are t-statistics. ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively. The observations are in firm-years.

Table 7
Panel regression results of threshold values.

Panel A: Panel regression with average measures

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9	Model 10	Model 11	Model 12	Model 13	Model 14	Model 15	Model 16	Model 17	Model 18
Intercept	1.275	2.488 *	1.085	2.591 *	0.373	1.852	0.625	1.742 *	1.343	2.545 **	1.128	2.655 **	0.495	1.940	0.730	1.823 *	-0.010	-0.010
	1.381	1.949	1.134	1.949	0.412	1.531	0.766	1.718	1.490	2.026	1.208	2.030	0.558	1.628	0.908	1.815	-0.382	-0.360
PINAvg	-1.419	-2.152	-1.087	-2.082	-0.053	-0.945	-0.410	-1.131	-1.554	-2.315	-1.124	-2.166	-0.341	-1.244	-0.631	-1.354	0.488	0.397
	-0.917	-1.377	-0.678	-1.279	-0.034	-0.611	-0.280	-0.788	-1.052	-1.538	-0.736	-1.382	-0.227	-0.831	-0.445	-0.968	0.427	0.344
SpreadAvg	15.217 ***	11.387 *	15.419 ***	11.923 *	3.959	0.782	2.789	-0.214	15.508 ***	11.918 *	15.339 ***	11.958 *	4.299	1.337	3.117	0.259	33.638 ***	33.698 ***
	2.735	1.681	2.667	1.690	0.861	0.138	0.657	-0.042	2.910	1.847	2.770	1.779	0.956	0.243	0.752	0.052	5.329	4.812
VolAvg	0.003		0.032		0.024		0.008		0.003		0.032		0.024		0.009		0.147 *	
	0.049		0.531		0.393		0.157		0.051		0.530		0.400		0.167		1.683	
DollarVolAvg		-0.066		-0.060		-0.067		-0.056		-0.065		-0.060		-0.064		-0.054		0.024
		-0.981		-0.851		-0.981		-1.020		-0.977		-0.876		-0.955		-0.990		0.323
Industry	0.366 ***	0.370 ***							0.365 ***	0.368 ***								
	3.090	3.180							3.116	3.197								
Governance	-0.010 ***	-0.011 ***	-0.010 **	-0.010 **					-0.010 ***	-0.011 ***	-0.010 **	-0.010 **						
	-2.627	-2.704	-2.487	-2.533					-2.677	-2.760	-2.516	-2.570						
Size	0.458	0.789	0.013	0.411	-0.290	0.126			0.464	0.782	0.032	0.427	-0.300	0.098				
	0.785	1.257	0.022	0.640	-0.487	0.195			0.806	1.270	0.055	0.676	-0.511	0.154				
BillAvg																	-14.832 *	-11.163
																	-1.819	-1.348
NoteAvg																	10.920	11.521
																	1.035	1.080
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	121	121	131	131	115	115	115	115	121	121	131	131	131	131
Adjusted R^2	0.118	0.126	0.048	0.052	-0.034	-0.027	-0.029	-0.021	0.133	0.141	0.064	0.068	-0.022	-0.015	-0.019	-0.011	0.167	0.148

Panel B: Panel regression with difference measures

	Model 17	Model 18	Model 19	Model 20	Model 21	Model 22	Model 23	Model 24	Model 25	Model 26	Model 27	Model 28	Model 29	Model 30	Model 31	Model 32	Model 33	Model 34
Intercept	1.031 ***	1.007 ***	1.278 ***	1.268 ***	0.567 ***	0.574 ***	0.589 ***	0.590 ***	1.106 ***	1.080 ***	1.351 ***	1.335 ***	0.651 ***	0.655 ***	0.679 ***	0.678 ***	-0.012	-0.013
	3.743	3.605	4.424	4.363	4.248	4.592	5.064	5.409	4.233	4.090	4.938	4.866	6.751	7.544	8.512	9.632	-0.408	-0.452
PINDiff	-1.553 *	-1.427 *	-1.731 *	-1.462	-1.206	-1.212	-1.067	-1.048	-1.385 *	-1.262	-1.553 *	-1.296	-1.037	-1.050	-0.904	-0.889	-0.678	-0.670
	-1.875	-1.664	-1.947	-1.594	-1.374	-1.348	-1.389	-1.323	-1.724	-1.519	-1.802	-1.458	-1.215	-1.197	-1.208	-1.148	-1.021	-1.013
SpreadDiff	10.461 ***	9.386 ***	10.091 ***	10.050 ***	9.299 ***	9.115 ***	7.959 ***	8.064 ***	10.334 ***	9.379 ***	9.828 ***	9.875 ***	9.043 ***	8.886 ***	7.635 ***	7.748 ***	12.852 ***	13.757 ***
	3.411	2.840	3.064	2.846	2.822	2.630	2.862	2.664	3.456	2.913	3.063	2.870	2.818	2.628	2.815	2.620	3.822	4.005
VolDiff	-0.093 **		-0.051		-0.013		0.002		-0.089 **		-0.047		-0.010		0.004		0.097	
	-2.208		-1.164		-0.315		0.064		-2.129		-1.077		-0.246		0.114		1.429	
DollarVolDiff		-0.065 *		-0.019		-0.011		0.004		-0.061		-0.016		-0.009		0.004		0.108 *
		-1.666		-0.485		-0.296		0.107		-1.584		-0.402		-0.246		0.138		1.783
Industry	0.495 ***	0.491 ***							0.495 ***	0.490 ***								
	4.183	4.053							4.214	4.078								
Governance	-0.013 ***	-0.011 ***	-0.011 ***	-0.010 **					-0.013 ***	-0.011 ***	-0.011 ***	-0.010 **						
	-3.346	-3.060	-2.707	-2.516					-3.336	-3.061	-2.690	-2.510						
Size	0.194	0.192	-0.170	-0.132	-0.315	-0.318			0.221	0.217	-0.133	-0.096	-0.279	-0.283				
	0.389	0.380	-0.323	-0.247	-0.594	-0.595			0.450	0.436	-0.255	-0.182	-0.532	-0.536				
BillDiff																	-12.800	-14.078
																	-0.708	-0.781
NoteDiff																	30.297	31.246
																	1.287	1.333
Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No	No	No	No	No	No
No. of Obs.	115	115	115	115	121	121	131	131	115	115	115	115	121	121	131	131	131	131
Adjusted R^2	0.208	0.193	0.086	0.076	0.031	0.031	0.035	0.036	0.216	0.202	0.096	0.088	0.040	0.040	0.042	0.042	0.079	0.087

Notes: The dependent variable is Threshold which is the U.S.$-denominated threshold estimate. Explanatory variables: PinDiff is the difference of the PIN of the NYSE over that of the TSX. PinAvg is the average PIN of the NYSE and the TSX. SpreadDiff is the difference of the quoted bid-ask spread of the NYSE over that of the TSX. SpreadAvg is the average relative quoted bid-ask spread of NYSE and the TSX. Control variables: VolAvg is the average of the log-transformations of average daily trading volume measures of the NYSE and the TSX. VolDiff is the difference of the log-transformation of average daily trading volume of the NYSE over that of the TSX. DollarVolAvg is the sum of log-transformations of average daily dollar trading volume measures of the NYSE and the TSX. DollarVolDiff is the difference of the log-transformation of average daily dollar trading volume of the NYSE over that of the TSX. Governance is the Report on Business governance index of Canadian firms published by Globe and Mail (McFarland, 2002). Industry equals one if the cross-lister is a manufacturing firm, and zero otherwise. Size is the normalized average market capitalization on the TSX and the NYSE. NoteAvg and NoteDiff are the average and difference of US and Canada's 10-year Treasury Note yields, respectively. BillAvg and BillDiff are the average and difference of US and Canada's 90-day Treasury bill discounts, respectively. VolatAvg and VolatDiff are the average and difference of US and Canada's market index return volatility, respectively. GdpAvg and GdpDiff are the average and difference of US and Canada's GDP growth rates, respectively. The numerical values below the estimates are t-statistics. ***, **, and * stand for statistical significance based on two-sided tests at the 1%, 5%, and 10% level, respectively. The observations are in firm-years.