WHERE DO INFORMED TRADERS TRADE CANADIAN SHARES CROSS-LISTED ON U.S. TRADING VENUES?

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ABSTRACT

Based on various theoretical models, the stylized prediction is that the dominant market in terms of trade activity for cross-listed firms will attract all trades. Examining trading and trade behavior for earnings (non)announcement periods for Canadian firms that are cross-listed on the TSX and the main U.S. trade venues shows that the stylized prediction does not hold empirically. Time-varying versions of the Easley et al model for estimating the probability of informed trading (PIN) and a regime-switching model for estimating PINs and decomposing bid-ask spreads are introduced. Its U.S. competitors continue to flourish although the TSX is dominant in terms of trade activity, has no trade cost advantage in terms of spreads and offers higher market depths. Trade venues in both countries not only have similar PINs during nonannouncement days but they react simultaneously to earnings news. Since most (but not all) informed trading occurs on the Canadian market for such announcements, information asymmetry is reduced on the U.S. market but remains unchanged on the TSX. This results in the Canadian market having the greater informativeness during earnings announcement periods.

JEL classifications: G14, G15.

Keywords: Inter-listing, information asymmetry, liquidity, probability of informed trading, regime-switching model.
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1. INTRODUCTION

Theoretical models of multi-trading venues with and without the presence of asymmetric information and in the absence of trapped liquidity, which are formulated by Pagano (1989), Chowdhry and Nanda (1991), Huddart, Hughes and Brunnermeier (1999) among others, show that one market will eventually dominate the other and attract all trades. These theoretical findings have led to the following stylized prediction. Uninformed traders tend to simultaneously trade together on the same venue to reduce the impact of trading on price and to diminish the probability of getting bagged by privately informed traders. If their choice of trade venue is based on lowest quoted trade cost, then this will be an important factor in inter-market competition. Given the beliefs of the informed about the trading behavior of the uninformed, informed traders also will act strategically by trading in the more active trade venue to hide their trades. This delays the informational impact of their cumulative trades on prices and prolongs their informational advantage over the uninformed. In turn, this leads to an equilibrium where one market should completely dominate the other market(s) and attract all trades.

Canadian firms cross-listed on U.S. markets provide a good experimental setting for testing why this stylized prediction and variations thereof do not hold in practice. Thus, the first question (Q1) addressed in this paper is: If the TSX dominants in terms of trade activity for Canadian shares cross-listed in U.S. markets, then why does the TSX not attract all trades of these shares? This can be further decomposed into two subquestions: Is the TSX the lowest trade cost venue with the greatest market depth (Q1a)? If it is, then why do the U.S. trade venues continue both to be viable cross-listing trade venues and meaningful contributors to the price search or discovery process for Canadian cross-listed stocks (Q1b)? The first subquestion is addressed by examining differences in spreads and depths
across trade venues for Canadian firms cross-listed on U.S. markets, while the second sub-question is addressed by examining trade activity by (un)informed investors and trade costs for the same sample.

To address question Q1b, we investigate how and where the (un)informed trade during normal (i.e., nonannouncement) periods and how this differs for information events. Informed trading is of particular importance because it is the portion of all trading activity that has a permanent impact on share prices and reveals privately held firm-specific information to the public. Since informed traders act to eliminate sufficient-sized deviations of market prices from “true” fundamental values (Black, 1986), informed traders make an important and permanent contribution to price formation and discovery (i.e., the informational efficiency of the market).

Thus, question Q1b is addressed by examining the more focused question (Q2): Where do the privately informed trade Canadian shares cross-listed on U.S. trade venues, especially during regular “hot” trading periods associated with information disclosures? This identifies the trade venue with the greatest informativeness about the fundamental values of cross-listed firms. To this end, we analyze the time-series behavior of the probability of informed trading (PIN) and the cost recovery payment required by market makers willing to trade against potentially informed traders on the competing trade venues for already cross-listed Canadian firms. Such an analysis allows for a fuller understanding of where the informed trade first and how price innovations are propagated and impounded into prices over time.

To better examine informational content or informativeness, trades and quotes are analyzed around earnings announcement dates. The choice of this event is not fortuitous. First, not only is trading activity usually amplified for material corporate announcements but public announcements also have multiple and complex effects on the pool of active (un)informed traders and their trade behaviors. In turn, this provides an empirical opportunity to observe changes in these trader pools and their behaviors on the competing venues.\(^1\) Secondly and from a methodological perspective, earnings announcements are recurring events that have the advantage of increasing the sample size beyond the

\(^1\) Krinsky and Lee (1996), among others, examine changes in spreads around earnings announcements.
Given the ongoing existence and resilience of the less dominant U.S.
markets in terms of trade activity for Canadian cross-listed firms, we expect nevertheless that the U.S.
trade venues still play a non-negligible role in terms of price informativeness, and that the
preponderance of information discovery occurs in the Canadian or home trade venue for material
information disclosures.

This paper makes three contributions to the literature. The first (and main) contribution is to
demonstrate that the change in the probability of trading against informed traders (PIN) is asymmetric
between the U.S. and the Canadian markets around the studied information event. The small change in
PIN hides several contradictory effects. First, PIN is reduced because information disclosure
eliminates the usefulness of previously private information on both markets. The presence of a larger
pool of uninformed traders encourages more traders to become privately informed (as Noronha, Sarin
and Saudagarhan (1996) also find). However, informed traders do not split their trades equally on both
competing trade venues. They trade more frequently in the domestic market compared to their pre-
announcement trading patterns on both markets. As a result, PIN declines significantly in the U.S.
market where traders are relatively more patient. The permanent trading cost, which is closely related
to private information trading, increases upon announcement on the Canadian market and is associated
with higher informed trading activity. As a result, the domestic Canadian market becomes much more
informative than the U.S. market for the sample firms. This suggests that investors should follow more
closely the domestic market to infer information about future price direction during announcement
windows than during regular (or non-announcement) periods where both markets contribute
significantly to price formation.

The second contribution is methodological. We believe that we are arguably the first to implement
a time-varying or conditional version of the Easley et al. (1996) model in order to track the behaviors

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2 The cost of this sample enlargement is that statistical tests must account for the loss of independence between
observations for the same firm.
of information asymmetry and PIN around information events, and also the first to use a regime-
switching model to estimate the time-varying or conditional components of the bid-ask spread.3

The third and final contribution is to show that no unexplainable spread cost advantage exists
among the competing trade venues, at least for the firms, events and periods studied herein. This
suggests that a Canadian company willing to cross-list in the U.S. should not use trading costs as a
decision criterion for choosing among the three main U.S. trade venues. Moreover, we find that the
proportional quoted and effective spreads are not higher on average on the U.S. markets compared to
the domestic Canadian market, counter to what was reported earlier in the literature. This partially
explains why the U.S. market still attracts trading activity and does not vanish. However, the Canadian
market offers higher depth (thickness) compared to the U.S. markets, and is dominant in terms of trade
activity. This greater depth attracts both liquidity and privately informed traders to trade on the
Canadian market around material information events. As a result, PIN is similar between the U.S. and
Canadian trade venues for the Canadian cross-listed shares during regular nonannouncement periods.
Since uninformed investors face the same PIN and trade costs, they do not shift all their trades to a
single location. This explains why both markets continue to co-exist for the trade of the shares of the
same firms.

The remainder of the paper is organized as follows. Section 2 briefly reviews the relevant
literature. Section 3 describes the sample, data set and basic statistical tests used herein. Section 4
reports and discusses the results for trading activity and liquidity comparisons between the two
national markets for the Canadian cross-listed shares. Section 5 reports and analyzes the time-varying
PIN estimates from the basic EKOP model. Section 6 infers the level of information asymmetry using
a regime-switching model and estimates the time-varying PINs and differentials in the spread
components. Section 7 concludes the paper.

3 Lei and Wu (2005) recently provide such a framework by assuming that the arrival rate of uninformed buy
orders switches between two levels in a Markov process with endogenous time-varying transition probabilities.
In contrast, Easley et al. (2002) use two GARCH-type specifications for time-varying arrival rates.
2. BRIEF REVIEW OF THE RELEVANT LITERATURE

2.1 Market Fragmentation

Market fragmentation is important since it affects the “public interest” via its impact on price discovery (O’Hara, 1995). Market fragmentation potentially can improve trading operational efficiency through inter-market or venue competition, which encourages the members of a trade venue to achieve a better trading mechanism and to pass the benefits to traders. Fragmentation also can negatively impact market efficiency and price discovery (Hamilton, 1979). Trading can be revealing as a second component of information, as in Blume, Easley and O’Hara (1994), since privately informed traders initiate a portion of all trades. Since the information content of trading is split or even lost with fragmentation, informed traders benefit because they have more time to profit from their nonpublic information. If orders are not subject to an international best bid and offer (IBBO), then traders are not guaranteed best execution. Furthermore, the market maker may not have the monopolistic power needed to face informed traders given order-book splitting.

Pagano (1989) demonstrates that one market will eventually dominate the other and attract all trades based on a model with multi-trading venues where traders transact only for liquidity purposes. Chowdhry and Nanda (1991) develop an equivalent but nonstrategic behavior model and show that one venue again dominates in terms of trading volume for the case of short-lived asymmetric information. In such a market, privately informed traders split their orders strategically to hide but trade aggressively to take advantage of the fragmentation. However, the existence of discretionary liquidity traders leads to an Admati and Pfleiderer (1988) concentrated market. In the absence of trapped liquidity, Huddart, Hughes and Brunnermeier (1999) demonstrate that all firms list on the high disclosure exchange to the preclusion of all other venues.

Using a model based on Glosten and Milgrom (1985), Madhavan (1995) shows that fragmented markets, where fragmentation is related to the degree of order-flow disclosure, are exploited by large liquidity traders and by dealers given reduced price competition. Furthermore, execution costs are higher due to the presence of informed traders and the behavior of large traders to trade usually on the
same side of the market. Davis and Lightfoot (1998) find that stocks trading under rule 19c-3 that permitted off the board trading (market fragmentation) had higher spreads than stocks under rule 390 that prohibited such trading (market concentration). Bennett and Wei (2003) report that listing switches from NASDAQ to NYSE exhibit higher variance reductions and lower execution costs for more fragmented pre-switch markets.

Amihud, Lauterbach and Mendelson (2003) test the value of consolidation using the expiry of deep-in-the-money warrants (close share substitutes). They report increased trading volume, lowered implicit spreads, and positive abnormal returns upon consolidation (i.e., at warrant expiry) due to enhanced liquidity, and greater effects for higher levels of fragmentation.

### 2.2 Market Behavior of Cross-Listed Firms Already Cross-Listed

Various papers compare the trade costs of cross-listed firms after the cross-listing decision. Kryzanowski and Zhang (2002) find that execution costs depend on the trade venue for Canadian stocks cross-listed on U.S. markets, and that midspread differences help to explain differentials in execution costs.\(^5\) They find a reduction in the total trade cost advantage of the Canadian market over the U.S. markets after the tick size reductions in 1996 and 1997. Ahn, Cao and Choe (1998) find that TSX decimalization only affected (small reduction) spreads of Canadian stocks cross-listed on the NASDAQ, and that U.S. traders did not switch trade venues due to the higher benefits of trading on U.S. markets.

Other papers examine market-maker participation and the price discovery process for already cross-listed firms. Eun and Sabherwal (2003) find that generally prices in the U.S. also contribute to

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\(^4\) Some papers in the rich literature dealing with the impact of initial cross-listing are of relevance herein. The finding by Sarkissian and Schill (2004) that cross-listing only reduces home bias minimally suggests that home bias may be a strong determinant of the ongoing viability of multi-trading venues for cross-listed firms. Noronha, Sarin and Saudagaran (1996) find that quoted spreads on the U.S. market do not decrease initially for U.S. firms cross-listed on the LSE and TSE as the increased competition hypothesis would suggest. They attribute this to an increase in information based trading (i.e., the attraction of additional informed trades to the cross-listing venues). Bailey, Karolyi and Salva (2002) find that volume and trading volatility increase upon cross-listing for firms making earnings announcements, although firms willingly choose to disclose more information to reduce informational asymmetry by cross-listing.

\(^5\) Price differences for Canadian cross-listed stocks should be arbitrated out. In contrast, perfect arbitrage is not possible for ADRs that are not fully fungible at zero cost due to differential trading costs and barriers (e.g., Gagnon and Karolyi, 2004).
the formation of Canadian prices through mutual feedback effects. This raises the unresolved issue of whether the relative informativeness of U.S. and Canadian prices change around material information events. Xu and Fung (2002) find similar results for Chinese shares cross-listed on the NYSE and the Stock Exchange of Hong Kong (SEHK). Grammig, Melvin and Schlag (2005) find that, while the majority of price discovery typically occurs in the home market, price discovery across international markets is not only more complex and richer during trading overlap periods but that the role for U.S. price discovery increases as the ratio of U.S.-to-home-market liquidity increases. Since U.S. and Canadian markets share common opening hours and the ratio of U.S.-to-home-market liquidity is likely to vary between information- and noninformation-event periods, this raises the unresolved issue of how price discovery and informativeness vary between competing national markets for Canadian cross-listed shares. Bacidore and Sofianos (2002) find that specialist participation is the same for U.S. domestic securities and Canadian securities cross-listed on the NYSE. However, they do not provide evidence on how market-maker participation varies across competing national trade venues for the Canadian shares cross-listed in the U.S. This unresolved issue cannot be addressed without access to a richer data set that identifies the trades and quotes of the market makers.

3. SAMPLE, DATA COLLECTION AND BASIC STATISTICAL TESTS

Data on quarterly earnings announcements are collected for 172 Canadian firms cross-listed on the U.S. exchanges for the calendar year 2002. The earliest announcement dates for each event are obtained by searching the press releases in SEDAR, company websites, the CBCA, Lexis Nexis and Bloomberg. All announcement dates with announcement window overlap for the same company are deleted, where each announcement window covers 41 trading days centered on the announcement date. Similarly, announcement dates with windows during which the stock switched U.S. listing venue or stocks traded at a price below a dollar are deleted.6 The resulting sample consists of 493 events for 135 companies (specifically, 58, 187 and 248 events on the AMEX, NASDAQ and NYSE.

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6 The minimum trade price increment and board lot size change at one dollar on the TSX.
respectively). No announcements occur on the weekend, and no day-of-the-week pattern is evident in the earnings announcements, although most occur mid-week (80% on Tuesday through Thursday).

Based on the descriptive company statistics reported in Table 1, Canadian cross-listings on the NYSE are much bigger than those on the AMEX and NASDAQ based on all three size variables. Canadian cross-listings on the NYSE are traded more heavily on the TSX. Comparisons between Canadian cross-listings on the AMEX and NASDAQ are hindered by the more heterogeneous composition of the AMEX cross-listings where some big and liquid outliers shift the mean but not the median upward. Based on the first two digits of the North American Industry Classification System (NAICS), most of these firms are in manufacturing (56 firms), mining (25 firms) and information (21 firms).

[Please insert table 1 about here.]

Trading data are obtained from the TSX’s Equity Trades and Quotes History (ETQH) and the TAQ databases. Record deletion occurs for any quote or trade outside of the regular trading hours (9:30 to 16:00 eastern time), open trades, trades with zero number of shares traded, cancelled or corrected trades, trades with delayed delivery, trades with special settlement conditions, trades representing a tick that exceeds 50%, quotes where the bid exceeds the ask price or where either equals zero or their relative spread exceeds 30%, and quotes posted during a trading halt or on a non-U.S. listing venue (the latter to avoid autoquote problems, as in Chordia, Roll and Subrahmanyam, 2001).7

The t- (Wilcoxon) test is used to determine if the mean (median) of any measure (or estimate) is statistically different from zero. Bootstrapped p-values are computed to deal with the lack of independence problem caused by some firms being represented up to four times in the sample. The bootstrapping procedure begins by running a regression of the appropriate measure on a vector of ones to get an estimate of the mean measure and its associated t-statistic. To generate the empirical

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7 These filters eliminated 2.2067% and 0.4585% of the quotes and trades, respectively, on the TSX, and 0.6121% and 0.0023% of the quotes and trades, respectively, on the U.S. markets.
distribution of this t-statistic under the null hypothesis, the residuals (mean deviations) are computed from the regression and then N samples of pseudo-random residuals are generated by drawing with replacement from the computed residuals. Each sample corresponds to a bootstrapped sample of the dependent variable (the measure) under the null since the latter corresponds to the mean being zero. By regressing each of these N samples on a vector of ones and computing the t-statistic of each intercept, which is the only explanatory variable, we obtain a vector of N t-statistics simulated under the null hypothesis. The original t-statistic is then inserted into this empirical distribution to determine the corresponding p-value from the cumulative empirical distribution (Davidson and MacKinnon, 2004; Greene, 2003). For the choice of N (i.e., the number of replications), both a fixed 999 repetitions and alternatively the three-step procedure suggested by Andrews and Bushinsky (2000) are used.

4. TRADING ACTIVITY AND LIQUIDITY OF SAMPLE FIRMS

Statistics on the cross-sectional distributions of various trading activity and liquidity measures are reported in table 2. As expected, the Canadian market is the trade venue of choice for trades of Canadian cross-listed firms. On average, 302 daily trades [385,703 shares] occur on the U.S. market versus 416 trades [537,350 shares] on the Canadian market. The cross-sectional mean [median] of the ratios of the number of trades in Canada versus the U.S. of 3.95 [1.58] implies that 79.80% [61.24%] of the trades occur in Canada for an average [typical] cross-listed firm. The market share of the TSX is even higher using share volume as the measure of trade activity. Specifically, the mean [median] ratio of Canadian to U.S. share volume of 9.40 [2.01] suggests that trading on the Canadian market represents up to 90% [67%] of total trading. For NASDAQ cross-listed shares, the median [mean] number of trades is lower [higher] in the U.S. The cross-sectional average ratio of 2.50 and 6.25 for

8 The mean ratio differs from the ratio of means due to cross-sectional variation in the ratios. For the number of trades, the ratio of means is 1.38 (4.16/3.02) while the mean ratio (i.e., the cross-sectional mean of the ratios) is 3.95. As the variance of the cross-sectional ratios increases, the difference between the mean ratio and the ratio of means widens. For the share volume, the contrast is even higher with a mean ratio of 9.40 and a ratio of means of 1.39.
number of trades and share volume, respectively, indicates that the Canadian market still captures most of the trades for an average firm cross-listed on NASDAQ.\footnote{Three groups of the NASDAQ cross-listed shares are formed based on their total trading volume. For the thinly traded shares, the Canadian market is highly dominant and represents on average over 91\% of total trading. In contrast, for the highly active issues, the Canadian market only captures 36\% of the volume.}

These trade activity results lead us to address the first question (Q1) posed earlier in the introduction: If the Canadian trade venue dominates the U.S. trade venues in terms of trade activity, then why does the TSX dominance not lead to the elimination of one or more U.S. markets for Canadian cross-listed shares? Is it because the TSX is not the lowest cost trade venue? We answer this latter question by first examining depths and spread differences between the two national markets for the sample of cross-listed firms. In the theoretical models of Pagano (1989) or Chowdhry and Nanda (1991), the dominant market must at least be execution competitive, and must be execution superior to drive out its competitors.\footnote{Other reasons beyond cost may explain the long-run sustainability of a fragmented market. These include trapped liquidity, clientele effect, and differences in investor preferences and holding periods (e.g., Kryzanowski and Rubalcava, 2005).} This condition implies that the spread should be at least competitive and that the depth be available.

As is evident from table 2, the TSX offers a deeper market for traders in the cross-listed firms with much less cross-sectional variability. The mean depth at both quotes of 2,578 shares on the TSX is significantly higher than the 2,335 shares on the U.S. trade venues, and for the two venues (NYSE and NASDAQ) that account for more than 85\% of our sample. The AMEX generates much of the cross-sectional variability in depths for the U.S. trade venues.\footnote{Outliers also exist on NYSE and AMEX. For example, the average depth for Nortel Networks is 162,276 shares during the third quarter of 2002 on the NYSE. This is more than 15 times the standard deviation above the mean of 2,557 shares for all NYSE observations.}

In contrast, the mean proportional quoted [effective] spreads are almost the same at 1.42\% and 1.44\% [1.10\% and 1.15\%] for the U.S. and Canadian markets, respectively. Based on bootstrapped p-values, the mean difference between the U.S. and Canadian markets of 1.8 [5.0] basis points for the quoted [effective] proportional spread is not significant. The mean effective proportional spread is
significantly higher on the domestic Canadian market (2.28%) than on the U.S. market (1.76%) for Canadian shares cross-listed on the AMEX.\textsuperscript{12}

Trade cost differences also exist across U.S. listing venues. The mean proportional quoted spreads of 2.45% and 2.21% for respectively the AMEX and NASDAQ are much higher than the 0.57% for the NYSE. The mean proportional effective spreads exhibit a similar pattern with values close to 1.76% for both AMEX and NASDAQ, and a much lower value of 0.42% for the NYSE.\textsuperscript{13} As Barclay et al. (1999) and Chung, Van Ness and Van Ness (2002) report, trading costs on NASDAQ declined after the introduction of new order handling rules in 1997. However, our findings still show higher costs on the NASDAQ compared to the NYSE for the sample studied herein.

Our contention is that these differences are share- and not exchange-specific since a similar pattern emerges for the Canadian-based trades differentiated by U.S. cross-listing venue for both absolute and relative spread measures. To illustrate, the posted proportional spread on the TSX of 0.52% for cross-listings on NYSE is significantly lower than the 2.89% and 2.18% for cross-listings on AMEX and NASDAQ respectively. A similar conclusion follows based on a comparison of the effective spreads.

A cross-sectional regression approach in the spirit of Stoll (2000) and others is used to test the following hypothesis $H_0$: The differences in the various spread measures across the U.S. trade venues are share and not exchange specific. The following cross-sectional regression is run to test this hypothesis (as in Harris, 1994; Bessembinder, 1999; Stoll, 2000):

$$\text{spread}_i = \alpha_1 + \alpha_2 \times AX_i + \alpha_3 \times NSQ_i + \beta_1 \times \text{size}_i + \beta_2 \times \text{volat}_i + \beta_3 \times \text{vol}_i + \epsilon_i$$ \hspace{1cm} (1)

where spread, is the proportional quoted (or effective) spread for observation $i$, and AX and NSQ are dummy variables taking the value of one if the cross-listing is on AMEX and NASDAQ, respectively, and is zero otherwise. The rationale for the choice of these regressors, already used by others in the

\textsuperscript{12}This is based on all trades regardless of their order size, trade-side initiator, time of the day and implied probability of informed trading.

\textsuperscript{13}As expected, not reflecting the relation between trade costs and price level gives a different inference. Specifically, the unreported mean posted absolute spreads are 7.77, 7.36 and 14.14 U.S. cents on the AMEX, NYSE and NASDAQ, respectively, and the corresponding unreported mean effective absolute spreads are 5.80, 5.46 and 11.08 U.S. cents. These absolute spreads suggest a much higher cost on the NASDAQ compared to the AMEX and NYSE. The patterns are similar for the medians for both types of spread measures.
literature, is now discussed. Size is the natural logarithm of total assets or total market capitalization (both from Compustat), and proxies for adverse information. Since the number of financial analysts following a firm and public investor knowledge should increase with increasing firm size, $\hat{\beta}_1$ is expected to be negative. Volat is the volatility as measured by the standard deviation of daily stock returns. Since volatility captures the price risk of carrying excess or low inventories by market makers, a positive $\hat{\beta}_2$ is expected as suggested by the inventory models. Vol is the average daily dollar trading volume (or number of trades). Since volume is inversely related to the order processing cost (basically fixed per transaction), the processing cost per share or per traded dollar is lower with higher volumes. Therefore, a negative relation is expected between volumes and spreads, as suggested by the literature. Both Volat and Vol are measured over the same period as the spread measure. Our primary interests are $\hat{\alpha}_1$ and $\hat{\alpha}_3$, which measure the specific additional cost of the AMEX and NASDAQ markets over the NYSE.

The results for various formulations of regression (1) are reported in table 3. Proportional quoted and effective spreads are respectively the dependent variables in models {1} – {4} and {5} – {8} for the TSX and correspondingly in models {9} – {12} and {13} – {16} for the U.S. trade venues. Each block of four regressions uses a mix between the two size (measured at calendar year-end) and the two volume explanatory variables. To reduce the dependence problem by limiting each firm to a solo appearance, only fourth quarter earnings announcements are used.

[Please insert table 3 about here.]

The R-square values of all 16 specifications are high ranging from 66.7% to 75.1%. The $\beta_1$ coefficient estimate on the size variable is not significantly different from zero for Canadian trades using either total assets or market cap. However, $\hat{\beta}_1$ is significant in all eight specifications with the predicted negative sign for U.S. trades. The estimated $\beta_2$ coefficient is significant in all 16 [14] specifications at the 10% [5%] level with the predicted positive sign. This confirms the conjecture that market makers widen spreads to self-protect against higher volatility and hence higher probability of
ruin. The estimated $\beta_3$ coefficients are significantly different from zero with the predicted negative sign in all specifications. This suggests that spreads are closely related to trading volumes, and that the order processing cost which is partly linked to trading volume is an important component of trading costs.

The $\hat{\alpha}_2$ and $\hat{\alpha}_3$ provide evidence against exchange-specific effects, as we conjectured earlier.

First, the AMEX impact on posted proportional spreads is significant and positive for the U.S. based trades, and ranges from 4.96 to 5.45 basis points for the four specifications. This represents the excess trading cost on the AMEX in comparison to the NYSE after controlling for share-specific factors. However, this extra cost is absent when the effective proportional cost is considered. In contrast, and like cross-listings on NASDAQ, cross-listings on AMEX have nonsignificant differences in trading costs for Canadian based trades and quotes. Second, no significant exchange effect occurs for shares cross-listed on NASDAQ for 13 out of the 16 specifications at the 5% confidence level. Thus, the higher overall trading costs for cross-listings on NASDAQ and to a lesser extent AMEX are due primarily to smaller firms that are subject to higher adverse selection, less liquid companies with higher per trade order processing costs, and riskier companies leading to higher inventory risk for market makers.

We now have the following partial answer to the first question ($Q1$) posed in the introduction. The TSX dominance in terms of trade activity does not lead to the elimination of one or more U.S. markets for Canadian cross-listed shares because all markets offer similarly “low” trade costs for average trade sizes for (especially liquidity) traders. This follows from the spread and depth comparisons where the two national markets are generally comparable in terms of spreads while the Canadian market offers greater depth. Nevertheless, the lower depth available on the U.S. market appears to be sufficient to fill a large number of transactions therein.

To provide some initial insights that the trade behavior of the trader pool differs between national markets and changes around earnings announcements, the proportion of executions against limit
orders (PEALO) is now examined. Given the longer holding periods (higher share turnover) documented by Kryzanowski and Rubalcava (2005) for U.S. trade venues, a lower PEALO is expected on U.S. trade venues. Since more patient investors are more likely to submit limit orders when they do decide to trade, they are less likely to have their trades executed against the standing BBO. As in Easley, O’Hara and Saar (2001), PEALOs are inferred using the Greene (1997) algorithm. Based on the entire 41-day (38-day nonannouncement) window, PEALOs are higher in the Canadian than the U.S. markets (respectively, 77% and 41% for number of trades and 58% and 34% for share volume), probably because the representative trader is less patient on the TSX, as noted above. This can result, for example, from a greater proportion of the executed orders in U.S. markets resulting from more “aggressive” limit order submissions (i.e. within the outstanding BBO) that are matched rapidly, and thus, are not counted as executions against (BBO) limit orders using the Greene algorithm.

PEALO changes that occur upon earnings announcement also are examined since such material events affect the price level and contain an element of surprise that increases volatility and thereby may affect the relative proportion of submitted limit orders (Handa and Schwartz, 1996; Foucault, 1999). The event-induced change in PEALO will be upwards, nil or downwards depending upon whether these new limit orders are at, away or within the national BBO and are subsequently executed against market orders. The PAELO drops by an insignificant 72 basis points for the TSX and by a significant 151 basis points on U.S. trade venues for the three-day announcement window. The significant decline in the U.S. implies that uninformed traders in that market become more aggressive as they are less likely to trade against informed traders upon earnings announcement, as is shown in the next two sections.

5.  **PIN CHANGES FROM A TIME-VARYING EKOP MEASURE**

We now address the second question posed in the introduction; namely: Where do the privately informed trade Canadian shares cross-listed on U.S. trade venues, especially during “hot” trading periods associated with information disclosures. To this end, the following two hypotheses are tested:
\( H_0^a \): The PINs are the same in the U.S. and Canadian markets for the cross-listed shares.

\( H_0^b \): The PINs change around quarterly earnings announcement dates in an asymmetric manner for the competing national trade venues.

Our expectation for \( H_0^a \) is that the competing cross-listing trade venues can co-exist if the PINs are the same across venues or that they can slightly differ provided the effective spreads do not. Furthermore, if they do differ, we expect that the PIN for the Canadian market is lower given its wider pool of uninformed traders due to its market dominance. Our expectation for \( H_0^b \) is that the respective PINs are likely to change upon announcement due to a changing mix of (un)informed traders on the different national trade venues, and that the direction of this change is indeterminate and downwards for the Canadian and U.S. trade venues respectively. The indeterminate PIN change for the Canadian market is because the increase in informed trading on this market may be more or less offset by an increase in uninformed trading. The downward PIN change for the U.S. markets is due to the migration of most informed trading to the Canadian market upon earnings announcement to take advantage of the deeper market.

The PINs are obtained using a conditional version of the Easley, Kiefer, O’Hara and Paperman or EKOP (1996) methodology.\(^{14}\) To obtain a time–varying daily PIN measure, each day is split into 78 successive intervals of five-minutes length, and the trades in each interval are treated as a single block. Ten samples are constructed where each sample contains eight trading blocks except for the ninth and the tenth that contain seven trading blocks each. Block assignment is conducted systematically on a rolling basis. The first block that covers the time period of 9:30 to 9:35 a.m. is assigned to the first sample. The second block that follows immediately is assigned to the second sample and so on. This design suffers from fewer problems than other alternatives in that not only are blocks of trades kept

---

\(^{14}\) EKOP use maximum likelihood to estimate the probability of an event occurring over a given number of days based on the argument that this probability is either one or zero for each day and cannot be inferred directly. EKOP also assume that, if an event takes place, the event occurs before trading starts for that trading day.
together but also every sample contains blocks spread throughout the trading day. Assigning individual trades randomly to samples would ignore the fact that the market maker may be able to infer potential private information from the time-series behavior of buy and sell orders and adjust quotes accordingly. Furthermore, the interval needs to be long enough to capture divided orders. This is supported by the observation that successive trades with the same sign and trader ID for our sample are executed within 15 seconds for over 97% of the trades.

The logic behind the EKOP model is that the arrival of informed and uninformed traders to the market place on day $t$ follows a Poisson process with intensity parameters of $\mu_t$ and $\epsilon_t$, respectively. For every sample $i$ on day $t$, we observe $B_i$, buyer-initiated trades and $S_i$, seller-initiated trades using the Lee and Ready (1991) algorithm, the probability of event occurrence is $\alpha_i$, and the conditional probability that the event has a negative impact on the stock is $\delta_i$ for day $t$. Dropping the $t$ subscript for convenience, the likelihood of observing $B_i$ and $S_i$ conditional on information set $\Theta$ is given by:

$$L((B,S)|\Theta) = (1-\alpha)e^{-\mu} \frac{\mu^B}{B!} \frac{\epsilon^S}{S!} + \alpha \delta e^{-\mu} \frac{\epsilon^B}{B!} \frac{(\mu+\epsilon)^S}{S!}$$

The first term in (2) corresponds to the case of a sample with no news events (i.e., all traders are uninformed). The second and third terms are linked to bad and good event samples, respectively. The parameter vector to be estimated using a data set consisting of the numbers of buys and sells is $\Theta=(\alpha, \delta, \epsilon, \mu)$. In EKOP, $\alpha$ and $\delta$ are constrained to be inside the interval $[0,1]$ through a logit transformation, and $\epsilon$ and $\mu$ are restricted to be positive by a logarithmic transformation.

Over the ten daily samples that are assumed to be independent, the likelihood function is:

$$L((B,S)|\Theta) = \prod_{i=1}^{10} L((B_i,S_i)|\Theta)$$

15 Samples formed successively during the trading day also are used. The trade clustering phenomenon well documented in the literature with periods of high trading volume (such as the morning session and the end of the trading day) and periods of low trading volume (like the midday) cause the samples from this sampling design to be quite different. In turn, this leads to erroneous inferences.
This function is maximized using several numerical methods and several starting points to avoid being 
trapped in a local maximum. The parameter of interest is the conditional probability of informed 
(PIN) trade, which is given by:

\[ \text{PIN}_t = \frac{\alpha_t \mu_t}{\alpha_t \mu_t + 2\epsilon_t} \]  

(4)

Before discussing the results, note that the estimates are likely to be poor if a stock is thinly 
traded. To illustrate, assume that a single daily trade occurs that is buyer initiated so that all the five 
minute intervals but one contains an observed pair of (0,0) and the sole exception contains the entry 
(1,0). Maximizing the logarithm of (3) in this case is trivial, and will force the delta parameter to a 
value of zero. For this reason, 312 observations are retained for further analysis because they exhibit at 
least five trades per day for each period.

The empirical estimates for the intra-daily EKOP model and a more standard inter-daily version 
are reported in Table 4. The median probability of an event occurring as measured by \( \alpha \) is 34.33% and 
35.00% based on U.S. and Canadian trades, respectively. These values are not statistically different 
based on the bootstrapped p-value. This result conforms to a priori expectations since no differences in 
event occurrence or their probabilities of occurrence should exist for the same stocks over the same 
time periods. The significant difference for AMEX trades is caused by a few outliers, especially for 
Bema Gold during the fourth quarter of the year 2002. Thus, the median alphas of 31.86% and 33.01% 
based on the Canadian- and U.S.-based trades, respectively, for the AMEX cross-listings are not 
significantly different.

[Please insert table 4 about here.]

The probability of a negative event given by the estimated \( \delta \) parameter is significantly lower by 
3.89% for U.S. trades than for Canadian trades. Furthermore, the estimate is significantly different 
(lower) only based on trades from NASDAQ and NYSE. As is shown subsequently, this result is

\[ \text{PIN}_t = \frac{\alpha_t \mu_t}{\alpha_t \mu_t + 2\epsilon_t} \]  

(4)

\[ \text{PIN}_t = \frac{\alpha_t \mu_t}{\alpha_t \mu_t + 2\epsilon_t} \]  

(4)

[Please insert table 4 about here.]
consistent with differences in informed trader participations as reflected in the differences in the PINs between various sets of competing trade venues.

Based on the median daily number of trades, uninformed investors trade more intensively on the Canadian (60.10) versus the U.S. trade venues (28.94). However, the mean of the ratios of numbers of trades on the Canadian versus U.S. trade venues of 4.33 reported in table 4 suggests that cases exist where practically all uninformed trading takes place on the Canadian market. To illustrate, Noranda Inc, which is cross-listed on the NYSE, has a ratio of 45.49 that indicates that 98% of its uninformed trading occurs on the TSX.

Similar results occur for informed trading intensity. Based on the second row of table 4, the median number of daily informed trades conditional on event occurrence is significantly higher on the Canadian trade venue (47.57 versus 29.07 in the U.S.). Their respective mean and median ratios of 2.63 and 1.19 are significantly different from one overall, and for the NASDAQ and NYSE and not AMEX. Thus, according to expectations, informed traders trade more intensively on the Canadian trade venue.

Given that the probability of event occurrence is the same for the U.S. and Canadian markets and that both liquidity and informed trading activity are higher on the Canadian trade venue and that the Canadian-to-U.S. market ratios are larger for liquidity- versus information-based trades, we expect that PIN is lower on the Canadian market according to expression (4). Based on table 4, a small but statistically significant lower mean PIN parameter difference of 256 basis points exists between the Canadian and U.S. trading venues (18.39% versus 20.95%). This is consistent with our a priori expectations for $H_{ln}^{I}$ in that the difference is very small. Listing venue-specific differences also occur for this metric. For example, the mean PIN is higher on the TSX than the AMEX (27.20% and 20.75%, respectively). This is most likely due to greater market thinness on the AMEX, which reduces the ability of the informed to hide their trades.
The trading patterns of both trader types around the quarterly earnings announcements now are examined further. Our expectation is that the informed start trading on the main market, which is usually the Canadian trade venue, and then split their trades between the fragmented markets to extend their advantage. Based on the findings of Eun and Sabherwal (2003), we expect the Canadian market to be most informative and the U.S. market to make a material (but more muted) contribution to price discovery since the privately informed also trade on the U.S. market.

An analysis of the plot of the intra-daily EKOP estimates given in Figure 1 supports our expectation and helps to explain the origins of the coexistence of the two national markets for cross-listed shares. Since the expectation is that pre-announcement days are characterized by informed trading, when informed trade first begins and how it moves over time for the two competing trade venues is of interest. Panel A of figure 1 depicts the time-series of mean and median \( \alpha \) estimates and the 95% confidence interval around the mean for both the U.S. and Canadian trades. All the x-axes are centered on day zero or the earnings announcement date.\(^{17}\) The graph for the U.S. trades documents a clear increase in the probability of event occurrence on day zero to a mean [median] peak value of 60.55% [60.20%]. In contrast, a steady decline in \( \alpha \) begins about four days prior to the announcement for the Canadian trades. The intra-daily mean is significantly higher in value than its inter-day counterpart essentially due to the added noise caused by greater data frequency.\(^{18}\)

\[\text{[Please insert figure 1 about here.]}\]

Based on panel B in figure 1, a very short-lived decline occurs on both the Canadian and U.S. markets at the announcement date for the delta parameter (i.e., the probability of a negative event). No difference exists between the inter-daily and intra-daily estimates for delta. Based on panel C, the most dramatic changes occur in the trading intensity parameters. Informed trading intensity increases by different magnitudes on U.S. (more muted) and Canadian trade venues around the earnings

\(^{17}\) The confidence interval is computed as the mean plus and minus twice its standard deviation. If the distribution is skewed and the variance around the mean is low as the number of observations is high, the median can lie outside the confidence interval for the mean. Panels C, D and E of figure 1 exhibit this case.

\(^{18}\) The cross-sectional time-series average alpha parameters over the entire time period are 58.92% and 54.70% for the U.S. and Canadian markets, respectively.
announcement days. The increase in the trading intensity of informed traders starts about three days prior to the announcement for the Canadian market and peaks on day one at 245.3 trades.

Based on panel D, no differences exist in the time series pattern of the intensity parameters (ε) of the uninformed between the two national markets. Both markets exhibit a clear and sudden jump in uninformed trading that corresponds with the earnings announcement. Uninformed trades peak at daily values of 187.0 and 241.2 on the U.S. and Canadian markets, respectively. If the peak period values are excluded, the respective mean ε estimates become 126.9 and 174.4. These trades are due to portfolio rebalancing after information release, to trades based on the belief that prices do not completely reflect the information released, to trades based on the observation of higher trading volumes, and to trades based on noise perceived to be information (Black, 1986).

Given that the alpha parameter increases slightly upon earnings announcement and that the uninformed traders increase their trading more than the informed, the expectation is for a lower PIN because of trader pool dilution created by the relatively larger increase of uninformed than informed traders. Based on Panel E of figure 1, this conjecture is confirmed as the mean PIN on the U.S. market drops from 41.05% two days before the earnings announcement to 37.25% on the day after the earnings announcement. Similarly, the PIN on the Canadian market falls from 37.20% to 34.96% over the same period. As for the inter-daily estimations, the PIN estimates for the intra-daily EKOP are slightly lower on the Canadian versus U.S. markets. Thus, even if the number of informed traders or their trading activity is more intense on the Canadian market, the mere presence of more uninformed trader activity reduces the probability that the counterpart is indeed a privately informed trader for a specific trade initiated by an uninformed trader. Since the PIN decline can be traced to the day before the earnings announcement for both competing markets, we cannot conclude that either market is first in terms of price discovery. This issue is revisited in the next section of the paper.

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19 The daily figures are higher than those estimated from the inter-daily version of 140 and 196 liquidity motivated trades per day for U.S. and Canadian market venues, respectively, which were presented earlier in table 4, because the intra-daily analysis is based on more liquid shares.

20 Figure 1 shows the results per intraday time interval. Since 10 samples are constructed daily, the daily trading intensities are equal to the corresponding intraday values multiplied by 10.
6. PIN AND SPREAD COMPONENT CHANGES FROM A REGIME-SWITCHING APPROACH

In this section, the daily probability of informed trading (PIN) is inferred from U.S. and Canadian based trades for Canadian shares cross-listed on the main U.S. exchanges using an alternate estimation method to better identify where the informed trade first at the announcement and where most of the information is first embedded into the price. The implied quoted spread also is decomposed into its temporary and permanent components. An increase [decrease] in the permanent cost (i.e., compensation for trading against informed traders) implies that the loss to informed traders is higher [lower] and that there are relatively more [less] informed traders in the marketplace.

6.1 Models and Methodology

Our approach is similar to that used by Nyholm (2003) for 20 NYSE-listed shares for August 1995 in that our model is based on Glosten and Harris (1988). However, unlike Nyholm, trading costs are directly dependent on trading volumes for three trade size categories. To develop the model, we first assume that the “true” asset price, \( m_t \), follows the motion equation given by:

\[
m_t = m_{t-1} + I_t \times Z_t \times s_t + e_t
\]

where \( I_t \) is the trade indicator at time \( t \) (-1 for seller initiated, and +1 for buyer initiated),

\( Z_t \) is the adverse information cost given that the trade is information based, which has a permanent impact on the “true” price of the asset,

\( s_t \) is a state variable related to the nature of the trade, which is perfectly observable to the market maker and not to the econometrician, and is equal to 1 for information-based trades and to 0 otherwise (i.e., for purely liquidity-based trades), and

\( e_t \) is related to the public news released during the interval between \( t-1 \) and \( t \).

Once the “true price” is set, the market maker sets transaction prices to recover her own order processing fees using the following pricing mechanism:
\[ P_t = m_t + I_t \times C_t \]  \hspace{1cm} (6)

where \( C_t \) is the order processing cost (i.e., the temporary component of the bid-ask spread), and all other terms are as defined previously.

Given the true price motion (5) and the pricing function (6), the change in price between time \( t-1 \) and \( t \) is computed using:

\[ dp_t = I_t \times Z_t \times s_t + I_t \times C_t - I_{t-1} \times C_{t-1} + e_t \]  \hspace{1cm} (7)

In (7), \( dp_t \) depends on the state variable \( s_t \). The change in the transaction price is \( dp_t = I_t \times Z_t \times s_t + I_t \times C_t - I_{t-1} \times C_{t-1} + e_t \) if an informed trader initiates the transaction, otherwise the market maker changes the transaction price by \( dp_t = I_t \times C_t - I_{t-1} \times C_{t-1} + e_t \). Since Glosten and Harris (1988) assume that both \( C_t \) and \( Z_t \) are related to the size of the transaction at time \( t \), size is used as a determinant of the trading cost.\(^{21}\)

Based on the findings of Barclay and Warner (1993), Chakravarty (2001), among others, on stealth trading and the association of medium-size trades with informed trades, trades are divided into three categories that are expected to have different marginal price impacts. Small trades are those involving 500 or less shares, as in Barclay and Warner (1993), Chakravarty (2001), and Koski and Michaely (2000). Medium-size trades are transactions of 501 to 5,000 shares, and large trades are transactions of 5,001 shares and more (e.g., as in Bessembinder, 2003; Koski and Michaely, 2000; SEC rule 11Ac 1-5). Since block trades on the NYSE and NASDAQ are defined as trades of 10,000 shares and more, this would be a natural definition of a large trade. This break point is not adopted herein because block trade delineation should depend on share price and trading activity (Bessembinder and Venkataraman, 2005). For the sample under investigation, no transactions of more than 10,000 shares occur for some firms, especially for trades on NASDAQ.

Hence, we have:

\(^{21}\) Nyholm uses a fixed cost per size category. We argue that since the cut off points are arbitrary, allowing size variability within each class can alleviate the classification problem, especially for the medium size trades that are considered as the ones with the highest proportion of informed trading.
\[ C_t = c_0 + c_1 V_t,DS_t + c_2 V_t,DM_t + c_3 V_t,DL_t \]  
(8a)
\[ Z_t = z_0 + z_1 V_t,DS_t + z_2 V_t,DM_t + z_3 V_t,DL_t \]  
(8b)

where \( DS_t \), \( DM_t \), and \( DL_t \) are dummy variables equal to zero or one, where a one occurs for trade sizes below 500 shares, between 501 and 5,000 shares, and over 5,000 shares for the respective dummies. Therefore, the change in the trading price is:

\[
dp_t = \alpha_1 + c_6 \times dI_t + c_1 \times d(I_t,V_t,DS_t) + c_2 \times d(I_t,V_t,DM_t) + c_3 \times d(I_t,V_t,DL_t) + \epsilon_t,
\]
if state 1; and

\[
dp_t = \alpha_2 + c_6 \times dI_t + c_1 \times d(I_t,V_t,DS_t) + c_2 \times d(I_t,V_t,DM_t) + c_3 \times d(I_t,V_t,DL_t) + z_0 \times I_t + z_1 \times I_t,V_t,DS_t + z_2 \times I_t,V_t,DM_t + z_3 \times I_t,V_t,DL_t + \epsilon_t,
\]
if state 2,  
(9a)
(9b)

where \( d(x) \) is the first order difference of the variable \( x \).

Following Nyholm (2003) and assuming that the state of nature follows a Markov chain, the parameters of interest are estimated using a Hamilton type filter similar to the Kalman filter and Maximum Likelihood Estimation. First, the density vectors of the residuals are formed assuming that they are normally distributed. Using Hamilton’s notation, this vector is \( \eta \) where:

\[
\eta = \begin{pmatrix} \eta_1 \\ \eta_2 \end{pmatrix} = \left( \frac{1}{\sqrt{2\pi \sigma^2_1}} \exp \left( -\frac{1}{2\sigma^2_1} \left( \begin{array}{c} dp_t - \alpha_1 - c_6 \times dI_t - c_1 \times d(I_t,V_t,DS_t) - c_2 \times d(I_t,V_t,DM_t) \\ - c_3 \times d(I_t,V_t,DL_t) \end{array} \right) \right) \right) - \left( \frac{1}{\sqrt{2\pi \sigma^2_2}} \exp \left( -\frac{1}{2\sigma^2_2} \left( \begin{array}{c} dp_t - \alpha_2 - c_6 \times dI_t - c_1 \times d(I_t,V_t,DS_t) - c_2 \times d(I_t,V_t,DM_t) \\ - c_3 \times d(I_t,V_t,DL_t) - z_0 \times I_t - z_1 \times I_t,V_t,DS_t \\ - z_2 \times I_t,V_t,DM_t - z_3 \times I_t,V_t,DL_t \end{array} \right) \right) \right) \}
\]
(10)

Two different intercepts are included in (10) to account for differences not explained by the regressors. Model (10) is unrestricted with unequal variances, although a restricted model with \( \sigma^2_1 = \sigma^2_2 \) also is estimated.\(^{22}\)

The state vector \( \xi_t \) is updated according to the motion equation:

\[
\xi_t = P_\xi \xi_{t-1}
\]

(11)

where \( P \) is the transition matrix that is given by:

---

\(^{22}\) While allowing for state-dependent variances is appealing, it reduces the number of cases with convergence. For this reason, only the restricted model results are reported herein.
In (12), $P_{11}$ and $P_{22}$ correspond to the probability that the state variable remains at states 1 and 2, respectively. $\xi_t$ is the conditional probability of the state vector based on the data up to time $t$. $\xi_{t+1|t}$ is the conditional expectation of the one-step-ahead state vector, which is a posterior update on the state vector given data up to time $t$.

By the conditional probability definition, $\xi_t$ can be inferred from:

$$\xi_t = \frac{\xi_{t-1} \ast \eta_t}{\Gamma(\xi_{t-1} \ast \eta_t)}$$

where $\ast$ denotes the element-by-element multiplication. The model is estimated using a quasi maximum likelihood approach where the likelihood function to be maximized is the intertemporal sum of the natural logarithms of the numerator in equation (13). Specifically, the estimation is based on:

$$\max_{\theta} \sum_{t=1}^{T} \log \left[ 1(\xi_{t-1} \ast \eta_t) \right]$$

where the vector of parameters is $\theta = (\alpha_1, c_0, c_1, c_2, c_3, z_0, z_1, z_2, z_3, \alpha_2, \sigma^2, p_{11}, p_{22})$. The model is estimated using all the filtered microstructure data, and also using equally time-spaced data that are sampled every 30 minutes to yield 14 observations per trading day. The former approach is not only very cumbersome and time consuming but often does not converge.

The smoothed estimates of the state vectors using the Kim (1993) algorithm also are generated. These inferences are not one-step-ahead “forecasts” but rather use all the data up to time $T$. Technically, these smoothed estimates, which use all data, are computed as backward iterations starting from $t = T-1$ and based on the expression:

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23 The same optimization techniques and algorithms are used here as were used earlier for the EKOP estimations. Specifically, the BFGS, DFP and Hill Climbing with different starting points are used.

24 To use a free maximization procedure and to avoid the constraints, an exponential transformation is made for $\sigma$ and a logistic transformation is made for both probability parameters $p_{11}$ and $p_{22}$.}

24
\[ \xi_{it} = \xi_{0i} + (\xi_{1i} \xi_{2i} \xi_{3i} \xi_{4i}) \]  

(15)

6.2 Results on Informed Trading

The results from estimating system (5) to (14) are reported in table 5. Separate estimations are run using trades on the U.S. and Canadian markets over the 41 trading days centered on the earnings announcement for each share observation. Out of the entire sample of 493 earnings announcements, convergence is achieved for 234 and 261 cases for the U.S. and the Canadian markets, respectively. Out of these, 144 cases correspond to the common sample whose results are reported in table 5 and in the remainder of section 6 of the paper.

[Please insert table 5 about here.]

Glosten and Harris (1988) argue that only the parameters \( c_0 \) and the \( \xi \)'s should be significant. Since \( c_0 \) corresponds to the temporary cost, it should be related only to the per-transaction fixed order-processing cost. The expectation is that this parameter estimate should be positive. As for the permanent cost, the argument is that the higher the trading volume the more likely the trader who acts aggressively is informed. Thus, the estimated \( \xi \) parameters are expected to be positive and significant. However, Barclay and Warner (1993), amongst others, argue that the \( \xi \) estimates depend on trade size. Thus, our expectation is that \( \xi_2 \) will be significantly positive indicating that average size trades are hiding orders from privately informed traders, while \( \xi_3 \) is probably of smaller magnitude and insignificant as large trades are normally not information motivated but are arranged before execution.

Based on table 5, all [most] \( \hat{c}_0 \) estimates are positive for the Canadian [U.S.] based trades. The mean \( \hat{c}_1 \) and \( \hat{c}_2 \) are negative for the Canadian trades, which indicates that trading costs decreases with higher trading volumes. This indicates that it is less costly for the market maker to match the higher sizes within the medium and small order-size categories by most likely using automated order-filling services. The \( \hat{c}_3 \) are positive (i.e., for large trades) for the U.S. trades. This indicates that the order processing cost is higher for large transactions because dealers usually need to manually allocate trades and spend more effort to fill large trades.
Since the $\hat{z}_1$ and $\hat{z}_2$ are positive, as expected, the information asymmetry cost increases with higher trade sizes for small and medium sized trades on both markets. However, the mean and median $\hat{z}_1$ are negative for Canadian trades, which indicate that the information asymmetry cost decreases with larger trade size for large trades in this market. This finding is consistent with the results reported in the literature that large trades are not usually motivated by private information. The average $\hat{z}_1$, which is positive for the U.S. trades, is inconsistent with expectations and can be explained as follows. A large, liquidity-motivated trader is more likely to submit an order to the deeper (Canadian) market since such an order is likely to work up or down the book to less attractive limit orders before being filled. If such traders submit their orders to the thinner U.S. market, there is a higher probability that the order is information motivated.

The continuation probabilities of states one and two are reported in table 5 where $P_{11}$ [$P_{22}$] is the probability that the next trade by the market maker will be against an uninformed [informed] trader given that the present trade is against an uninformed [informed] trader. The continuation of state 1 which corresponds to successive trades by uninformed traders is the more likely based on its higher means and medians for both Canadian and U.S. trades. The unreported cut-off points of the cross-sectional distribution at the first and the third quartiles also confirm this finding. The distribution of the $P_{11}$ parameter for the U.S. trades dominates that from the Canadian trades, and its mean of 82.58% is significantly higher than the corresponding mean for Canadian trades of 77.98%. The $P_{22}$ parameter estimate exhibits a similar but less pronounced pattern for the two trade venues. While the parameter estimate for the U.S. trades dominates that for Canadian trades in terms of the first quartile, the median and the third quartile, it has a wider distribution than for the Canadian trades. This is supported by the matched-sample t-test but not the Wilcoxon test. The mean of $P_{22}$ of 30% for the Canadian trades indicates that market makers expect an informed trade to be followed by an uninformed trade with a probability of 70%, almost 2.5 times the likelihood of trading again with an informed trader. This contrasts with the likelihood of trading with an informed trader after dealing with an uninformed
trader of 17.42% and 22.02% for the U.S. and Canadian trade venues, respectively. The corresponding odds ratios are respectively 4.74 and 3.54 times. These odds ratios provide some initial evidence using this alternative approach that informed trading is less likely on the U.S. market.

To investigate this issue further, the unconditional estimates of trading against uninformed and informed traders (i.e., the probabilities that states 1 and 2 continue) are computed for both trading venues. The unconditional probabilities are given by the limit of the transition matrix, whose first element is equal to:

\[ P_1 = \frac{(1 - P_{22})}{(2 - P_{11} - P_{22})} \]  

(16)

The results reported in table 5 provide weak support for our hypothesis that informed traders are more likely to trade on the domestic Canadian market and hence supports our hypothesis \( H_{u2} \). While the mean unconditional state 1 probability of 82.34% for U.S. trades is significantly higher than the 77.61% mean estimate for Canadian trades, the medians are not significantly different.

The above methodology cannot be used to examine changes around the earnings announcements dates, and whether any changes are symmetrical between trading venues. To compute a probability of trading against (un)informed traders for both markets at both announcement and non-announcement dates, an inference on the realization of the state vector \( s_t \) is required. To that end, the vector \( \xi_{t|\xi} \) is built, which corresponds to the conditional probability of the state vector for each trade based on smoothing all the available data, and a sub-vector from \( \xi_{t|\xi} \) is built that only uses trades from the three-day announcement windows. The probability of state 2 over the earnings announcement window is the sum of the second column of the sub-vector \( \xi_{t|\xi} \) divided by the corresponding number of trades. By analogy, the probability for the days that exclude the earnings announcement uses the \( \xi_{t|\xi} \) components that do not correspond to trades that occur during the earnings announcement window.\(^{25}\)

\(^{25}\) The probability of state 1 (i.e., trading against the uninformed) is obtained using the same methodology used to obtain the values reported in the first column of the vector \( \xi_{t|\xi} \).
The resulting estimates for both trading venues are reported in table 6 for PIN on the Canadian (U.S.) market in the announcement (non-announcement) windows. The mean and median PIN are not significantly different between the two windows for the Canadian market. In contrast, the PIN declines for U.S.-based trades (from 19.64% to 14.18% for the mean). As reported in section five and consistent with our a priori expectations for \( H_{26} \), the PIN declines only for U.S.-based trades during the earnings announcement window. Furthermore, the magnitudes of the PIN estimates are within the 19% to 21% range for both the regime-switching and EKOP models.\(^{26}\)

[Please insert table 6 about here.]

Based on the first column in panel B of table 6, no significant difference exists in the likelihood of information trading between the two trade venues for the nonannouncement period. In contrast to the findings reported in table 4 using the EKOP methodology where the PIN was slightly lower in the U.S. markets compared to the Canadian market for nonannouncement periods, we now find that the PIN is similar for the two national trade venues although informed traders are more likely to trade in the domestic Canadian market. Since both trader types maintain a relatively constant presence during regular (or nonannouncement) periods so as not to change the equilibrium of either market, both venues can survive in a fragmented market. This provides an additional reason why the U.S. trade venues continue both to be viable cross-listing trade venues and meaningful contributors to the price search or discovery process for Canadian cross-listed stocks (as posed as question Q1b in the introduction to this paper).

Based on the second column of panel B in table 6, PIN does not change significantly on the Canadian market upon announcement. In contrast, the PIN on the U.S. market declines significantly by 558 basis points upon announcement and becomes lower than that of its Canadian counterpart (see column 4 in the same panel). Column 3 in panel B documents that the average PIN from the Canadian trades is higher than that of the U.S. trades by 652 basis points during the three days earnings

\(^{26}\) These values are similar to those reported by Hasbrouck (2004) for U.S. trade venues (median and medians of 19.61% and 18.09% for both all U.S. listing venues and AMEX/NYSE).
announcement window. These findings suggest that informed traders trade more intensively on the domestic Canadian market than they do on the U.S. market upon earnings announcements for the Canadian cross-listed shares in the U.S.

6.3 Trading Cost and Spread Components

An implied bid-ask spread and its permanent and transitory components can be computed from the regime-switching estimates. By examining how these components change between nonannouncement and announcement periods, we can infer if market makers have altered the compensation that they require for asymmetric risk exposure. Since the permanent cost component relates to both the intensity of informed trading and the extent of the information asymmetry, an increase [decrease] in the permanent component indicates higher [lower] informed trading and may confirm the above reported finding that informed traders do trade more on the Canadian markets relatively to the U.S. markets upon announcement compared to the nonannouncement periods.

Based on equations (9a) and (9b), the half quoted spread is estimated by:

$$\hat{S}_t = \hat{C}_t + \hat{Z}_t = \hat{c}_0 + \hat{z}_0 + (\hat{c}_1 + \hat{z}_1) \times V_t \times DS_t + (\hat{c}_2 + \hat{z}_2) \times V_t \times DM_t + (\hat{c}_3 + \hat{z}_3) \times V_t \times DL_t$$

(17)

where the hat over the variables indicates the estimated value from the regime-switching model. The half permanent variable and half temporary fixed costs are given respectively by:

$$\hat{Z}_t = \hat{c}_0 + \hat{z}_0 + \hat{c}_1 \times V_t \times DS_t + \hat{c}_2 \times V_t \times DM_t + \hat{c}_3 \times V_t \times DL_t$$

(18)

$$\hat{C}_t = \hat{c}_0 + \hat{c}_1 \times V_t \times DS_t + \hat{c}_2 \times V_t \times DM_t + \hat{c}_3 \times V_t \times DL_t$$

(19)

One drawback of this Glosten-Harris-like approach is that it can yield negative estimates of the implied spread or its components, which is usually caused by one dominant (usually permanent) component being negatively related to volume.\(^{27}\) However, by relating the \(z\) parameter to the trade size itself in a piecewise linear manner, this problem is greatly alleviated.

---

\(^{27}\)This is more common for small trades. Theoretically, when the estimated \(z_0\) parameter is negative and the \(z_1\) parameter is a small positive value, the permanent component is dominated by the negative intercept. In turn, this results in a negative estimate of the permanent trading cost.
The cross-sectional distribution of the half spreads and their components for both the Canadian and the U.S. trades and for the (non)event windows are reported in table 7. The total implied half spread is on average 0.29% [0.26%] for the earnings announcement window during [non]announcement periods for the Canadian based trades. As expected, this value is well below the posted half spread of 0.71% and is closer to the effective half spread of 0.54%. The reason is that the implied spreads are driven by transaction prices and account for any trades that occur inside the posted quotes. The estimates from the U.S. trades are similar with a mean implied half spread of 0.22%.

The statistical significance of the spread differential between Canadian and U.S. implied trades in the announcement window are reported in Panel B of table 7. The change in the proportional temporary component is not significant based on the t-test, and is significant and higher for the Canadian trade venue based on the Wilcoxon test. Both tests agree that the proportional permanent cost component is higher for Canadian- versus U.S.-based trades. The mean [median] differentials for the half spread and the half permanent cost are a significant 8 [5] basis points above the U.S. estimates. This supports our earlier conclusion that more informed trades occur on the domestic Canadian market at the announcement date. Also based on table 7, the reverse occurs during the non-event windows where the permanent spread cost is higher on the U.S. market. This further supports our earlier conclusion that the impact of the announcement is mostly on the U.S. permanent cost component. In other words, market makers expect most of the informed traders to trade on the domestic Canadian market upon announcement. Therefore, market makers increase [decrease] the compensation required to bear this higher [lower] risk on the Canadian [U.S.] market.

With regard to the announcement-induced change, the permanent cost component increases [remains unchanged] and the temporary component exhibits no change on the Canadian [U.S.] market.

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The implied absolute spreads and components are not reported because they are expressed in different currencies. The estimates on the Canadian market are naturally higher because they differ on average by the average U.S./Canadian FX rate of approximately 1.57 during 2002.

---
As a result, the overall implied spread increases significantly on average (by 4.17 basis points) only on the Canadian market. The increase in the permanent cost for Canadian-based trades is a reaction by the market makers to the intensified presence of informed traders, primarily on the Canadian trade venue. Therefore, most informed traders shift their trades to the Canadian domestic market during the announcement window. This also explains the more marked decline in the PIN measure on the U.S. markets that was reported in section five. The informed investors take advantage of the higher uninformed trading intensity of impatient traders who are rebalancing or are reacting to the news announcements. Since these investors are more likely to be present on the same side of the market, the depth that the Canadian market offers is very valuable.

7. CONCLUDING REMARKS

The liquidity of the Canadian shares that are cross-listed on the major U.S. markets was investigated in this paper. Cross-listing is an example of market fragmentation where several trading platforms survive together. For this market situation to hold, minimum trading conditions must be met so that one market does not become dominant and force the others to close. The dynamics of the trading game between informed and uninformed traders and the market makers allows the survival of all markets only if all market participants are present to a “sufficient” extent on all of the markets. As documented in this paper, this appears to be the general case for the cross-listed sample of Canadian stocks examined herein.

To better examine information-based trading, trading across periods with changing information structures was analyzed. Since material corporate announcements are typical examples of changing intertemporal private information structures, this paper examined earnings announcements. The expectation was that such announcements are information revealing and subject to informed trading, and thus, would impact information-related trades by changing the pools of informed and uninformed traders and their trade behaviors.
The cost advantage argument for trading on the Canadian market plays a lesser role in attracting trades since all markets offer similar spread cost structures. However, the domestic Canadian market still offers a deeper market compared to the U.S. market for Canadian cross-listed shares. Therefore, the Canadian market is more appealing for larger market orders because it can absorb these trades within the currently posted BBO.

Informed traders intensify their trades on the domestic compared to the foreign market after an earnings announcement, as manifested in the probability of informed trading (PIN) and behavior of the bid-ask spread components. The case of Canadian cross-listed shares is special in that several papers find that both markets (domestic Canadian and foreign U.S.) participate in making the price of a stock through feedback effects since investors infer the price and update their beliefs by observing both markets simultaneously. This is counter to other cross-listing situations where all (or most) price discovery occurs in the domestic market, as Grammig, Melvin and Schlag (2005) report.

This paper presents an argument for mainly following the Canadian market during earnings announcements periods. Since this is the market where most of the informed investors trade upon information announcement, its aggregation power and informativeness are stronger. Thus, market participants should concentrate primarily on the domestic Canadian market to infer the market values of the underlying cross-listed firms at announcement times, and monitor both markets during non-announcement periods. This is a very important extension to the findings reported by Eun and Sabherwal (2003). Further investigations of the (very) short-term feedback effects between the two national markets for Canadian cross-listed shares need to be conducted in future research to support or refute these initial results.
REFERENCES


Table 1. Descriptive statistics

<table>
<thead>
<tr>
<th></th>
<th>All venues (493 observations)</th>
<th>AMEX (58 observations)</th>
<th>NASDAQ (187 observations)</th>
<th>NYSE (248 observations)</th>
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<td>Total assets</td>
<td>Trading vol.</td>
<td>Equity</td>
<td>Market cap</td>
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<td>0.03</td>
<td>5.22</td>
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<td>197.47</td>
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<td>17238.00</td>
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<td>9.92</td>
<td>3.24</td>
<td>3.95</td>
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</tbody>
</table>

“All venues” corresponds to all cross-listed shares in our sample. AMEX, NASDAQ and NYSE correspond to the samples of Canadian companies cross-listed on the AMEX, NASDAQ and NYSE, respectively. Market cap is the total market capitalization. Total assets (data item 6 in S&P Compustat database) and total capitalization (data item 24 times data item 25 in Compustat) are as the end of the calendar year 2002 and are in Canadian millions of dollars. The trading volume is the number of shares traded on the TSX for the calendar year 2002 and corresponds to data item 28 from the Compustat database. Equity is in millions of Canadian dollars.
Table 2. Trading activity and liquidity measures for Canadian shares cross-listed on U.S. exchanges

<table>
<thead>
<tr>
<th></th>
<th>Shares volume</th>
<th>Nb. of trades</th>
<th>Percent spread</th>
<th>Percent effective spread</th>
<th>Shares depth</th>
</tr>
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<tr>
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<td>U.S. Canada</td>
<td>U.S. Canada</td>
<td>U.S. Canada</td>
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<td>1.6024</td>
<td>0.7799</td>
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<td>4.1549</td>
<td>1.4220</td>
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<td>0.0501</td>
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<td>0.0000</td>
<td>0.7300</td>
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<td>1.4932</td>
<td>2.4537</td>
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<td>0.4338</td>
<td>0.5212</td>
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<td>0.0260</td>
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<tr>
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<td>0.0000</td>
<td>0.7970</td>
<td>0.3000</td>
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<td><strong>NYSE listed</strong></td>
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<td>0.0626</td>
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<td>4.2648</td>
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<td>5.6658</td>
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<td>0.0000</td>
<td>0.0200</td>
<td>0.1510</td>
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</table>

Share volume is in millions of shares. Number of trades is in lots of 100. Percent spread and percent effective spread are the quoted and effective spreads expressed as percentages. Share depth is in thousands of shares. Only admissible trades are included to compute volume and number of trades. All, AMEX listed, NASDAQ listed and NYSE listed correspond respectively to the entire sample of Canadian shares-observations cross-listed on all the U.S. exchanges, and to the shares-observations cross-listed on the AMEX, NASDAQ and NYSE only. Each shares-observation corresponds to one Canadian share cross-listed on the US market for one earnings announcement period. Mean_d_r (mean difference or ratio) corresponds to the Canada-U.S. matched [difference] ratio for [spread measures] share volume, number of trades and quoted depth. For the volume and the depth variables, the statistics of each company from the US and the Canadian market are matched. The ratio of the mean statistic inferred from the Canadian trades over the mean statistic inferred from the US trades is then computed. The cross-sectional mean ratio over the cross-listed shares-observations is reported as Mean_d_r. For the percentage spread variables, the difference between the Canadian based mean statistic and the US based mean statistic is computed for each cross-listed share. The cross-sectional average over all cross-listed shares-observations is then reported as Mean_d_r. P-values are bootstrapped values.
Table 3. Determinants of trading costs for cross-listed shares

<table>
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<tr>
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<td>Intercept</td>
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<td>0.043</td>
<td>0.047</td>
<td>0.084</td>
<td>0.086</td>
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<td>0.094</td>
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<td>0.912</td>
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<td>0.266</td>
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<td>t-stat</td>
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<td>4.949</td>
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<td>0.447</td>
<td>1.696</td>
<td>1.377</td>
<td>2.069</td>
<td>1.693</td>
</tr>
</tbody>
</table>

Adjusted $R^2$ 0.686 0.688 0.670 0.667 0.751 0.748 0.700 0.702 0.673 0.671 0.694 0.691 0.730 0.741 0.725 0.738

The following regression is run: $spread = \alpha + \beta_1 \times \text{size} + \beta_2 \times \text{volat} + \beta_3 \times \text{vol} + \alpha_1 \times \text{AX} + \alpha_2 \times \text{NSQ} + \epsilon$. The dependent variable is the quoted spread on the Canadian exchange for models 1-4, the proportional effective spread on the Canadian exchange for models 5-8, the quoted spread on the U.S. exchanges for models 9-12, and the proportional effective spread on the U.S. exchanges for models 13-16. The sample contains 130 observations on Canadian shares cross-listed on the U.S. market. Ln(assets) is the natural logarithm of total assets as determined from the S&P Compustat database in Canadian dollars. Ln(MC) is the natural logarithm of the market capitalization computed as the number of shares outstanding multiplied by the price per share, as in Table 1. Volat Can (U.S.) is the standard deviation of the daily returns from Canadian (U.S.) closing prices during the period under study. Ln(Vol Can) and Ln(Vol U.S.) correspond to the natural logarithms of the average daily trading volumes on the Canadian and the U.S. exchanges, respectively. The former is in Canadian dollars and the latter is in U.S. dollars. Ln(nb T. Can) and Ln(nb T. U.S.) correspond respectively to the natural logarithms of the average daily number of trades on the Canadian exchange and the U.S. exchanges for the cross-listed shares. AX and NSQ are dummy variables equal to one when the U.S. listing venue is respectively AMEX and NASDAQ. The coefficients $\beta_1, \beta_2, \alpha_1$, and $\alpha_2$ are multiplied by 1000. White corrected t-statistics are reported in the table. $^a$, $^b$ and $^c$ indicate significance at the 0.10, 0.05 and 0.01 levels, respectively.
Table 4. EKOP results for the inter-day data for the Canadian cross-listed firms

<table>
<thead>
<tr>
<th>Sample</th>
<th>Statistic</th>
<th>U.S. venues</th>
<th>Canada</th>
</tr>
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<tr>
<td></td>
<td></td>
<td>α</td>
<td>δ</td>
</tr>
<tr>
<td>ALL</td>
<td>Median</td>
<td>34.328</td>
<td>39.193</td>
</tr>
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<td></td>
<td>Mean</td>
<td>35.630</td>
<td>41.040</td>
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<td>AMEX</td>
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<td>Mean</td>
<td>34.106</td>
<td>42.314</td>
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<td></td>
<td>T-stat</td>
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<td>Mean_d_r</td>
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</tbody>
</table>

This table presents the EKOP model estimates by maximizing expression (3) for the U.S. and the Canadian based trades, respectively. The sample is separated into three sub-samples to conform to the following three cross-listing venues: the AMEX, the NASDAQ and the NYSE. The estimation is based on interdaily observations of buyer and seller initiated trades. α is the probability of daily event occurrence; δ is the probability that the event, conditional on its occurrence, has a negative impact on the stock; µ is the trading intensity of the informed traders as measured by the number of trades conditional on the occurrence of an event; ε is the trading intensity of the uninformed traders; and PIN is the probability of informed trading. Mean_d_r correspond to the matched differences [ratios] for the α, δ and PIN [µ and ε] parameters based on the Canadian and U.S. based trades. The reported p-values are the bootstrap statistics under the null that the difference is zero for α, δ and PIN, and the ratio is one for µ and ε. a, b and c indicate significance at the 0.10, 0.05 and 0.01 levels, respectively.
Table 5. Regime switching estimates for the Canadian cross-listed firms

<table>
<thead>
<tr>
<th>Statistic</th>
<th>10^5*c_0</th>
<th>10^6*c_1</th>
<th>10^5*c_2</th>
<th>10^5*c_3</th>
<th>10^6*z_0</th>
<th>10^6*z_1</th>
</tr>
</thead>
<tbody>
<tr>
<td>Median</td>
<td>11.328</td>
<td>7.302</td>
<td>1.908</td>
<td>-1.704</td>
<td>0.858</td>
<td>-0.001</td>
</tr>
<tr>
<td>Mean</td>
<td>19.597</td>
<td>23.150</td>
<td>2.405</td>
<td>-2.720</td>
<td>1.242</td>
<td>0.116</td>
</tr>
<tr>
<td>t-stat</td>
<td>8.523</td>
<td>5.559</td>
<td>5.544</td>
<td>1.082</td>
<td>2.428</td>
<td>8.614</td>
</tr>
<tr>
<td>Mean (paired)</td>
<td>-3.804</td>
<td>-8.300</td>
<td>-3.980</td>
<td>-0.508</td>
<td>-11.842</td>
<td>-10.613</td>
</tr>
<tr>
<td>t-stat</td>
<td>-1.706^a</td>
<td>-6.668^b</td>
<td>-1.812^b</td>
<td>-0.544</td>
<td>-0.523</td>
<td>-0.765</td>
</tr>
<tr>
<td>Wilcoxon</td>
<td>2.233^b</td>
<td>7.047^b</td>
<td>9.930^b</td>
<td>1.082</td>
<td>2.428</td>
<td>8.614</td>
</tr>
</tbody>
</table>

Prices are assumed to follow the motion equation (7): \( dp = I_t \times Z_t \times x_t + I_t \times C_t + I_t \times C_s + I_t \times e_t \), where \( I_t \) is the usual trade indicator variable, \( Z_t \) is the information asymmetry trading cost, \( C_t \) is the temporary trading cost, \( x_t \) is a state variable and \( e_t \) is the innovation induced by new public information. The state variable is zero (state one) when the trader is privately informed and one (state two) if not. Both trading cost components are assumed to be linearly dependent on the trading volume. The change in price would be given by the state variable as:

\[
\begin{align*}
dp_1 &= \alpha_1 + c_1 \times dI + c_2 \times dI(V, DS) + c_3 \times dI(V, DM) + c_4 \times dI(V, DL) + e_1, \\
\text{if state 1, and} \\
\end{align*}
\]

\[
\begin{align*}
dp_2 &= \alpha_2 + c_1 \times dI + c_2 \times dI(V, DS) + c_3 \times dI(V, DM) + c_4 \times dI(V, DL) + z_1 \times I_t + z_2 \times I_t \times DS + z_3 \times I_t \times DM + z_4 \times I_t \times DL + e_2, \\
\text{if state 2.}
\end{align*}
\]

Estimation is based on maximizing equation (14): \( \max \sum T \log \left[ \Lambda Z_t, \eta_t \right] \). Two vector parameter sets \( \theta = \left( \alpha_1, c_1, c_2, c_3, z_1, z_2, z_3, z_4, \ldots, \alpha_2, \sigma^2, p_{11}, p_{12} \right) \) are estimated for the Canadian and the U.S. trades of the Canadian cross-listed firms. \( p_{11} \) and \( p_{22} \) are respectively the transition probabilities of state one and two. Lim state 1 is the limiting probability of state 1 and is given by \( p = \left( 1 - p_{12} \right) / \left( 1 - p_{11} \right) \). Since observations are dependent by nature, the p-values are simulated under the null of equality. ^a, ^b and ^c indicate significance at the 0.10, 0.05 and 0.01 levels, respectively.
Table 6. Probability of informed trading from the regime-switching model for the Canadian cross-listed firms

<table>
<thead>
<tr>
<th>Panel A. PIN estimates</th>
<th>Announcement window PIN</th>
<th>Nonannouncement Window PIN</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Canada</td>
<td>U.S.</td>
</tr>
<tr>
<td>Min</td>
<td>0.0000</td>
<td>0.0000</td>
</tr>
<tr>
<td>Median</td>
<td>0.0542</td>
<td>0.0432</td>
</tr>
<tr>
<td>Mean</td>
<td>0.2070</td>
<td>0.1418</td>
</tr>
<tr>
<td>Maxi</td>
<td>1.0000</td>
<td>1.0000</td>
</tr>
<tr>
<td>SD</td>
<td>0.3392</td>
<td>0.2352</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Panel B. Significance tests</th>
<th>Test</th>
<th>(i)</th>
<th>(ii)</th>
<th>(iii)</th>
<th>(iv)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Mean*</td>
<td>0.0091</td>
<td>0.0004</td>
<td>0.0652</td>
<td>-0.0558</td>
</tr>
<tr>
<td></td>
<td>t-statistic</td>
<td>0.2075</td>
<td>0.1071</td>
<td>2.7207</td>
<td>-3.1069</td>
</tr>
<tr>
<td></td>
<td>Median**</td>
<td>0.0042</td>
<td>0.0004</td>
<td>0.0105</td>
<td>-0.0152</td>
</tr>
<tr>
<td></td>
<td>Wilcoxon</td>
<td>0.3443</td>
<td>0.5128</td>
<td>1.9735</td>
<td>1.7852</td>
</tr>
</tbody>
</table>

PIN is the probability of trading against an informed trader. It is computed by summing the smoothed vector of the state variable \( \hat{\xi}_{it} \) generated by using the Kim (1993) algorithm, and then dividing by the number of trades over each period. The regime-switching model is estimated using the methodology described in table 5. Smoothed state vector \( \hat{\xi}_{it} \) is given by \( \hat{\xi}_{it} = (P^{\frac{1}{2}}) \xi_{t+1} \) where \( P \) is the transition matrix. The four tests correspond to (i) the difference between the PIN from the Canadian trades less the PIN from the U.S. trades for the periods excluding the earnings announcement three-day window; (ii) the difference between the PIN for the earnings announcement windows less the PIN for the periods excluding the earnings announcement windows for the Canadian trades; (iii) the difference between the PIN from the Canadian trades less the PIN from the U.S. trades for the earnings announcement three-day windows; and (iv) the difference between the PIN for the earnings announcement days less the PIN for the periods excluding the earnings announcement windows for the U.S. trades. * and ** correspond respectively to the mean and median of these differences, and they are tested for equality with zero. Because of the dependent nature of the observations, the p-values are computed by bootstrapping from the estimates. * , ** , and c indicate significance at the 0.10, 0.05 and 0.01 levels, respectively.
Table 7. Spread components from the regime-switching model for the Canadian cross-listed firms

<table>
<thead>
<tr>
<th>Panel A. Spread and spread components estimates</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Canada</strong></td>
<td><strong>U.S.</strong></td>
</tr>
<tr>
<td><strong>Earnings announcement window</strong></td>
<td><strong>Earnings announcement window</strong></td>
</tr>
<tr>
<td>Statistics</td>
<td></td>
</tr>
<tr>
<td>Median</td>
<td></td>
</tr>
<tr>
<td>(1-π) total temporary permanent (1-π) total temporary permanent</td>
<td></td>
</tr>
<tr>
<td>0.1333</td>
<td></td>
</tr>
<tr>
<td>0.1645</td>
<td></td>
</tr>
<tr>
<td>0.1356</td>
<td></td>
</tr>
<tr>
<td>0.1519</td>
<td></td>
</tr>
<tr>
<td>Mean</td>
<td></td>
</tr>
<tr>
<td>0.2544</td>
<td></td>
</tr>
<tr>
<td>0.1721</td>
<td></td>
</tr>
<tr>
<td>0.3122</td>
<td></td>
</tr>
<tr>
<td>0.1535</td>
<td></td>
</tr>
<tr>
<td>Non announcement window</td>
<td></td>
</tr>
<tr>
<td>Median</td>
<td></td>
</tr>
<tr>
<td>(1-π) total temporary permanent (1-π) total temporary permanent</td>
<td></td>
</tr>
<tr>
<td>0.1452</td>
<td></td>
</tr>
<tr>
<td>0.1093</td>
<td></td>
</tr>
<tr>
<td>0.1398</td>
<td></td>
</tr>
<tr>
<td>0.122</td>
<td></td>
</tr>
<tr>
<td>Mean</td>
<td></td>
</tr>
<tr>
<td>0.3583</td>
<td></td>
</tr>
<tr>
<td>0.2039</td>
<td></td>
</tr>
<tr>
<td>0.2888</td>
<td></td>
</tr>
<tr>
<td>0.1537</td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Panel B. Spread and spread components difference tests</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Canada-U.S.</strong></td>
<td></td>
</tr>
<tr>
<td><strong>Announcement</strong></td>
<td><strong>No announcement</strong></td>
</tr>
<tr>
<td>Mean</td>
<td>Total</td>
</tr>
<tr>
<td>0.0887</td>
<td>0.0091</td>
</tr>
<tr>
<td>t-stat</td>
<td>1.7146a</td>
</tr>
<tr>
<td>Median</td>
<td>0.0491</td>
</tr>
<tr>
<td>Wilcoxon</td>
<td>3.7101c</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Announcement - No announcement</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Canada</strong></td>
<td><strong>U.S.</strong></td>
</tr>
<tr>
<td>Mean</td>
<td>Total</td>
</tr>
<tr>
<td>0.0417</td>
<td>-0.0322</td>
</tr>
<tr>
<td>t-stat</td>
<td>2.4800b</td>
</tr>
<tr>
<td>Median</td>
<td>0.0709</td>
</tr>
<tr>
<td>Wilcoxon</td>
<td>1.4875</td>
</tr>
</tbody>
</table>

Panel A of the table reports results for half-spreads and half-spread components, which are derived from the Glosten and Harris (1988) model using a regime-switching framework. Dollar spreads are multiplied by 10 and the proportional spreads are in percentages. The regime-switching model is estimated using the methodology described in table 5. The quoted implied half-spread is equal to $S_t^c = \hat{C}_t + \hat{Z}_t$. The temporary component is given by $\hat{C}_t = \hat{c}_0 + \hat{c}_1 \times V_t \times DS_t + \hat{c}_2 \times V_t \times DM_t + \hat{c}_3 \times V_t \times DL_t$, and the half permanent component is given by $\hat{Z}_t = \hat{z}_0 + \hat{z}_1 \times V_t \times DS_t + \hat{z}_2 \times V_t \times DM_t + \hat{z}_3 \times V_t \times DL_t$. The fitted values $\hat{c}_0$, $\hat{c}_1$, $\hat{c}_2$, $\hat{c}_3$, $\hat{z}_0$, $\hat{z}_1$, $\hat{z}_2$ and $\hat{z}_3$ are the estimates whose cross-sectional statistics appear in table 5. (1-π) is the proportion in percentage of the implied spread that is attributed to the permanent component. Panel B presents results of paired matched sample tests of the differential half spreads and components. The three-day announcement and the 38-day nonannouncement periods or windows are compared, as are the estimates from the U.S. and the Canadian trades. P-values are computed using a bootstrapping method under the null that the mean or the median is zero. $a$, $b$ and $c$ indicate significance at the 0.10, 0.05 and 0.01 levels, respectively.
Figure 1. Time series of intraday EKOP estimates

Panel A. $\alpha$ estimates
Figure 1 continued

Panel B. $\delta$ estimates
Figure 1 continued

Panel C. µ estimates
Figure 1 continued

Panel D. $\epsilon$ estimates
Panel E. PIN estimates

The continuous line shows the mean. ––––––– corresponds to the 95% confidence interval around the mean. ––.––.–– corresponds to the median. The above graphs show the time series of the EKOP estimates.
Intradaily estimates of five parameters (i.e., $\alpha$, $\delta$, $\mu$, $\epsilon$ and PIN) are obtained for EKOP by maximizing expression 3 using intradaily samples of trades for each cross-listed firm. Cross-sectional averages are then computed using event days. $\alpha$ is the probability of event occurrence; $\delta$ is the probability that the event, conditional on its occurrence, has a negative impact on the stock; $\mu$ is the trading intensity of the informed traders as measured by the number of trades conditional on the occurrence of an event; $\epsilon$ is the trading intensity of the uninformed traders; and PIN is the probability of informed trading.