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Multimarket Trading and Liquidity: Theory and Evidence

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ABSTRACT

We develop a new model of multimarket trading to explain the differences in the foreign share of trading volume of internationally cross-listed stocks. The model predicts that the trading volume of a cross-listed stock is proportionally higher on the exchange in which the cross-listed asset returns have greater correlation with returns of other assets traded on that market. We find robust empirical support for this prediction using stock return and volume data on 251 non-U.S. stocks cross-listed on major U.S. exchanges.

WITH THE ENHANCED GLOBALIZATION OF FINANCIAL MARKETS, the number of non-U.S. firms choosing to cross-list shares on a U.S. exchange has increased substantially. In 1990 there were only 352 non-U.S. stocks listed on the New York Stock Exchange (NYSE) and Nasdaq, but by the end of 2002 the number had more than doubled to over 850. If one includes stocks trading over the counter (OTC) and as private placement issues, the number of non-U.S. companies with shares trading in the United States now exceeds 2,300. This dramatic increase reflects not only U.S. investors' desire for international diversification, but also foreign companies' desire to access global capital, broaden their shareholder base, and enhance company visibility.

Corporations generally view U.S. cross-listings as value-enhancing decisions, but there is disagreement about the sources of the benefit.[1] One of the benefits

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[1] There are many studies of the economic benefits of international cross-listings for issuers, including the gains from overcoming investment barriers and segmented markets (Alexander, Eun, and Janakiramanan (1988), Foerster and Karolyi (1999), Miller (1999)), from expanding the shareholder base (Foerster and Karolyi (1999)), from enhancing visibility among analysts and the media (Baker, Nofsinger, and Weaver (2002), Lang, Lins, and Miller (2003), Bailey, Karolyi, and Salva (2005)) and from legal "bonding" to the U.S. legal system to protect minority shareholders

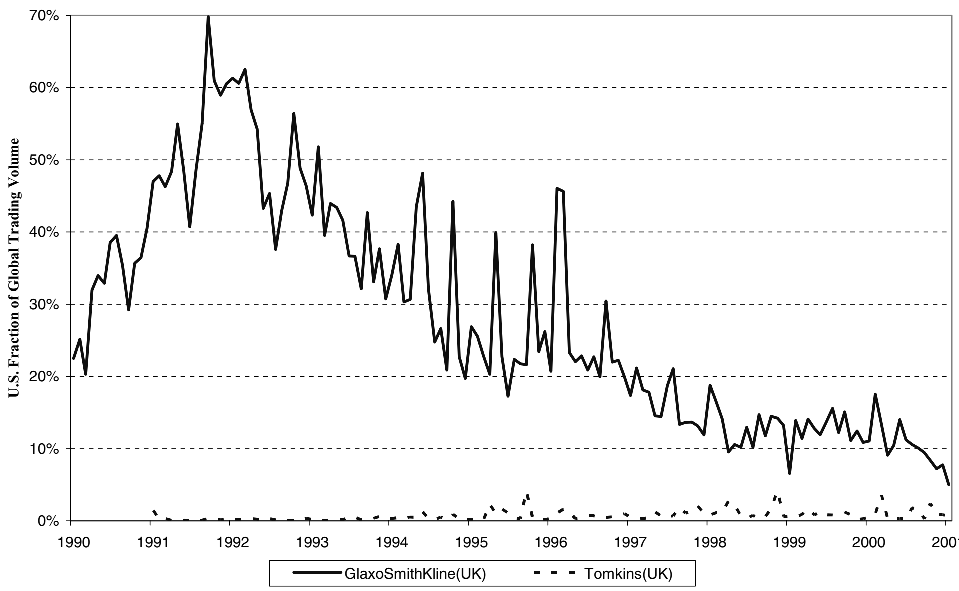

Figure 1. Multimarket trading in shares of Tomkins (United Kingdom) and GlaxoSmithKline (United Kingdom). Proportion of combined monthly trading volume on the home market and in cross-listed shares on the U.S. market represented by that of cross-listed shares on the U.S. market.

most frequently cited in surveys of company managers and investors alike is the increased liquidity in share trading associated with a U.S. listing (Mittoo (1992), Fanto and Karmel (1997)). Consistent with this view, the empirical evidence indicates that a U.S. listing is accompanied by a large 40% to 50% increase in the number and value of shares traded in the combined U.S. and home market compared to that in the home market before the listing (Smith and Sofianos (1997), Foerster and Karolyi (1998)).

However, not all companies experience the benefits of increased liquidity. In fact, a remarkable feature of trading activity in cross-listed stocks on U.S. exchanges is the great variability in the U.S. fraction of global trading. For some stocks, U.S. trading typically represents less than 5% of global trading in any given month, whereas, in other stocks, U.S. trading comprises well over 90% of global trading. Moreover, it is not simply country-level factors, such as regulatory restrictions or the extent of overlapping trading hours that dictate these outcomes, as there is significant cross-sectional diversity in the U.S. fraction of trading even among stocks from the same country. Consider, for example, the different experiences of two firms from the same home market over the same period (Figure 1). Monthly trading volume (in number of shares traded)

(Coffee (1999), Reese and Weisbach (2002), Doidge (2004), Doidge, Karolyi, and Stulz (2004)). Karolyi (1998) surveys dozens of earlier studies of the share price and liquidity effects of international listings up to 1998.

on the NYSE for Tomkins, a UK engineering company, has rarely risen over 2% of its combined global volume (NYSE and the London Stock Exchange) since it listed on the NYSE in 1988, while the U.S. fraction of trading volume for GlaxoSmithKline, a UK pharmaceutical company, has been over 20% for all but the last years of the period illustrated in the figure.

Understanding why such different multimarket trading environments arise is of paramount importance for managers of all cross-listed companies, but especially those with little U.S.-based trading activity, as the outcome potentially reflects on the long-term viability of the listing and its potential as a vehicle for raising capital, broadening the shareholder base and for enhancing the company's visibility and profile. Moreover, understanding how trading activity is apportioned across markets is important for global investors, particularly arbitrageurs that actively trade in both overseas ordinaries and their equivalent cross-listed shares in the U.S., as the distribution of trading affects the feasibility of their strategies.[2] Finally, understanding the factors that affect the distribution of global trading volume is important to stock exchanges, which compete with each other for new listings and for order flow among existing listings.

In this paper, we develop a new model of multimarket trading to explain the variation in the U.S. share of global trading volume across the sample of non-U.S. stocks cross-listed on U.S. exchanges. The model predicts that, under fairly general conditions, the distribution of trading volume across exchanges competing for order flow is related to the correlation of the cross-listed asset returns with the returns of other assets traded in the respective markets. The model is based on a standard Kyle (1985) framework with two stock exchanges and three assets, where one asset trades exclusively in the first market, a second asset trades exclusively in the second market, and the cross-listed asset, which trades on both exchanges. The two exchanges are segmented in that the risk-neutral market makers observe the order flow of assets only on their own exchange. The informed traders, who observe private information about the different assets, and the discretionary liquidity traders are able to trade on either market and even across markets. Because the asset returns on each exchange are correlated, competitive market makers, when pricing an asset, can infer information not only from the asset's own order flow, but also from the order flow of other assets traded on the exchange. Indeed, in equilibrium, the more correlated the returns of the two assets, the more relevant the order flows of one asset are for the pricing of the other asset, and the less sensitive the price of any one asset to its own order flow. This outcome guides the liquidity traders

[2] It is difficult to gauge the size of the arbitrage market in internationally cross-listed shares. However, there are a number of vehicles to facilitate the process, including Bank of New York's DR Converter, which is a proprietary cost analysis model for ADRs versus home market ordinaries with web-based access for global investors. Citibank's ADR division has also initiated a new SWIFT centralized messaging system for brokers to reduce settlement risk and delays for ADR issuances and cancellations on behalf of customers. Similarly, JP Morgan's ADR division has built Cross-Book Maximizer, which is the market's first online automated marketplace for ADR traders and brokers to execute ADR-ordinary share exchange transactions.

in choosing where to trade the cross-listed asset: The higher the correlation of the returns of the cross-listed asset with the domestic asset, the more informative the domestic asset's order flow, which leads both liquidity and informed traders to submit a larger proportion of their orders in the cross-listed asset to that exchange. That is, proportionally more volume takes place on the market in which the cross-listed asset has greater correlation with the other assets traded on that market.

Our new model of multimarket trading represents an important departure from the best-known models to date, including those of Pagano (1989), Chowdhry and Nanda (CN, 1991) and Domowitz, Glen, and Madhavan (DGM, 1998), in that neither exchange design nor assumptions of differential trading costs plays a central role in determining the distribution of trading volume across exchanges.[3] For example, Pagano's two-period model with risk-averse investors abstracts from asymmetric information considerations and focuses on the role of traders' expectations of other traders' actions. His key result is a "knife-edge" Nash-type of equilibrium in which both markets can survive, but only if a number of special assumptions about exchange design, such as equal transactions costs and equal numbers of traders in each market, hold. CN does allow for asymmetric information, as we do, by extending the framework of Kyle (1985) and Admati and Pfleiderer (1988) to allow simultaneous trading in multiple markets. Their informed investors trade strategically to maximize profits from their private information by locating trades across markets according to which markets are "thick" with liquidity traders. Liquidity naturally clusters in a particular market, which CN define as their "winner takes most" equilibrium. We differ from CN because (i) we model the trading decisions of both informed and discretionary liquidity traders, and (ii) the exchanges in our model provide different liquidity in the cross-listed asset by means of the joint distribution of the asset returns traded on each exchange rather than by way of assumptions about the number of "small" liquidity traders that are confined to each market.

Our model is also different from that of DGM, who extend the model of Glosten and Milgrom (1985) to allow investors to trade in the home market or the new cross-listed market at differential costs of execution due to the bid–ask spread and the different costs of information acquisition in the two markets. Their model relies on incremental information acquisition costs to ensure that local investors find it cheaper to trade locally unless there is perfect transparency in quotes between the two markets, in which case cross-listing will lower volatility due to lower spreads from greater volume overall, and more intense competition for order flow from both exchanges can arise. They conclude that CN's winner takes most equilibrium is complicated by the degree of transparency between the two markets. Our model does not rely on assumptions about information acquisition costs to yield the important differences in trading costs. Rather, the choice of where to trade stems from the liquidity that arises endogenously from

[3] Chapter 8 of O'Hara (1995) provides a useful overview of the theoretical literature on liquidity and the relationship across markets.

the extent of the correlation of the cross-listed asset's returns with the returns of other assets traded on a particular exchange.

We proceed to test the key prediction of our model and find strong empirical support. We examine weekly trading stock price and volume data from the home market and the NYSE/Nasdaq for 251 companies from 24 emerging and developed countries around the world. We compute the average U.S. fraction of global trading for each stock based on available data from January 1995 to December 2004. Using U.S. dollar-denominated home market returns on the stock, we construct a "U.S. information factor," which is our proxy of the correlation of the cross-listed stock's returns with information signals about the values of other assets traded in the home and U.S. market. The ratio is computed from multi-index market model regressions of the individual stock's returns on U.S. dollar-denominated returns of the home market index and a U.S. market index (S&P 500). Specifically, we measure the factor as the incremental explanatory power of the U.S. market in terms of the difference in R-squared of a two-index model including the U.S. index relative to the R-squared of a single-index model with just the home market index, adjusted for degrees of freedom. Even after controlling for potential endogeneity effects as well as a host of other firm-specific (e.g., market capitalization, U.S. institutional ownership, foreign ownership restrictions, and home market analyst coverage), issue-specific (e.g., NYSE vs. Nasdaq), as well as regional- and country-specific factors (e.g., time zone factors), we find that the U.S. fraction of global trading is strongly positively related to the U.S. information factor, which is consistent with our theory.

Our paper makes an important contribution to the existing empirical work on multimarket trading. To date, the most comprehensive empirical study of the distribution of global trading in internationally cross-listed stocks is that of Pulatkonak and Sofianos (1999).[4] Their study examines 1996 trading volume data for 254 NYSE-listed non-U.S. stocks. They estimate an econometric model explaining the variation in the U.S. share of global trading volume drawing from country-specific, company-specific, and issue-specific factors. The most important variable in their model is the time zone factor; 40% of the variation of the U.S. market share can be explained by the hours of overlap in trading between the NYSE and the home market for the stock. Surprisingly, other variables, such as differential trading costs, an industry dummy, market capitalization, whether the issue was associated with raising capital, and what kind of cross-listing is employed (direct ordinary listing, New York Registered Share, American Depositary Receipt (ADR), Global Registered Share), explain relatively little of the cross-sectional variation. Our theoretical model predicts a new "information factor" variable for consideration in the model, for which

[4] A recent paper by Halling et al. (2004) examines the overseas trading volume for a sample of 111 European firms cross-listed on a variety of overseas exchanges. They test and confirm the explanatory power of our information factor. They also show that the magnitude of overseas trading wanes dramatically in the years following the cross-listing, which leads them to call into question whether providing an active foreign marketplace for shares is a reasonable motivation for cross-listings, at least for European firms.

we find strong empirical support, even after controlling for time zone effects and other variables.[5]

The remainder of this paper is organized as follows. Section I develops our multimarket model of trading. In this section we outline the predictions of the model, as well as a series of testable alternative hypotheses. Our empirical analysis follows in Section II, where we describe the sample of firms, data sources, variable construction, econometric model, and main results. Section III concludes the paper.

I. Analytical Framework and Testable Hypotheses

A. The Model

We consider a world with two stock exchanges (denoted by subscripts 1 and 2). Two assets are listed on each exchange. Exchanges are organized as a standard Kyle (1985) market, where competitive risk-neutral market makers make the market for the assets listed on their exchange. To have a meaningful separation between exchanges, we introduce the following segmentation. We assume that before setting the prices, market makers observe the net order flow of all the assets traded on their exchange; they do not observe the orders submitted to the other exchange. After each round of trading, a public announcement is made after which the risk-neutral market makers agree that the prices of the assets traded on the first and second exchanges have changed by the innovations υ_1, $\upsilon_2 \in R^2$, respectively. Specifically,

$$\upsilon_1 = F_1 s + \varepsilon_1$$

and

$$\upsilon_2 = F_2 s + \varepsilon_2,$$

where

$$F_1 = \begin{pmatrix} 1 & 0 \\ a & b \end{pmatrix} \quad \text{and} \quad F_2 = \begin{pmatrix} 0 & 1 \\ a & b \end{pmatrix}. \tag{1}$$

The variables s, ε_1, and ε_2 are zero mean normally distributed random variables that take values in R^2. Furthermore, s is independent from ε_1 and ε_2, and the covariance matrix of s, denoted by P, is diagonal

$$P = \begin{pmatrix} \sigma_1^2 & 0 \\ 0 & \sigma_2^2 \end{pmatrix}. \tag{2}$$

[5] Our paper also provides a useful theoretical backdrop for various new studies of price discovery in stocks with multimarket trading (Ding et al. (1999), Eun and Sabherwal (2003), Grammig, Melvin, and Schlag (2005)) and for other empirical studies of multimarket trading in dual-listed companies (DLCs, or "Siamese twins"), such as Royal Dutch Shell or Unilever (Rosenthal and Young (1990), Froot and Dabora (1999), and, Bedi, Richards, and Tennant (2003)), and ADRs (Gagnon and Karolyi (2004)).

We do not impose a restriction on the covariance matrix of the ε's or on the correlation between ε_1 and ε_2. In particular, we can assume that the second elements of ε_1 and ε_2 are identical, so the innovations in the second asset traded on each exchange are identical. We hereafter refer to the second asset on each exchange as the cross-listed asset. Similarly, we refer to the first asset on each exchange as the local asset.

The innovation in the price of the local asset traded on the first exchange depends on s_1, but not on s_2; while the innovation in the price of the local asset traded on the second exchange depends on s_2 but not on s_1. The parameters a and b measure the exposure of the cross-listed asset to s_1 and s_2, respectively. Without loss of generality, we assume that the value of the assets prior to the innovation is zero and that interest rates are zero.

There are two strategic risk-neutral informed traders. One of them observes s_1, while the other observes s_2. The informed traders may submit orders to both exchanges and for all assets. Let $x_1^i, x_2^i \in R^2$ $(i = 1, 2)$ be the market order that the ith informed trader submits to the first and second exchange, respectively, and $x_1 = x_1^1 + x_1^2$ and $x_2 = x_2^1 + x_2^2$ be the aggregate informed orders submitted to the first and second exchange, respectively.

Next, we consider the uninformed liquidity traders. We take the liquidity demand for each asset as given, and assume that the liquidity demand for both the local asset and the cross-listed asset are uncorrelated standard normal variables. However, to discuss volume patterns of the cross-listed asset, we have to allow the liquidity traders who trade the cross-listed asset to choose to which exchange they prefer to submit their order, or even to allow them to split their order between both exchanges. As in Admati and Pfleiderer (1988), we assume that there are n risk-neutral discretionary liquidity traders, each of whom has to trade a certain amount of the cross-listed stock. Let w_i be the amount the ith discretionary trader has to trade. The market views w_i as a zero mean normal random variable. Furthermore, $\mathrm{cov}(w_i, w_j)$ equals zero for $i \neq j$. Thus,

$$1 = \mathrm{var}\left(\sum_{i=1}^{n} w_i\right) = \sum_{i=1}^{n} \mathrm{var}(w_i). \tag{3}$$

We let $\alpha(w_i)$ be the fraction of the demand the ith discretionary trader submits to the first exchange. When $\alpha_i = 1$ (0), the ith trader submits his entire order to the first (second) exchange. Let R_1 and R_2 be the covariance matrix of liquidity demand on the first and second exchange, respectively. We then have

$$R_1 = \begin{pmatrix} 1 & 0 \\ 0 & \sum_{i=1}^{n} \mathrm{var}(\alpha_i(w_i)w_i) \end{pmatrix} \quad \text{and} \quad R_2 = \begin{pmatrix} 1 & 0 \\ 0 & \sum_{i=1}^{n} \mathrm{var}((1-\alpha_i(w_i))w_i) \end{pmatrix}.$$

We denote by $y_1 \in R^2$ and $y_2 \in R^2$ the aggregate order flow (i.e., the total orders submitted by both informed traders and liquidity traders) submitted to the first and second exchange, respectively. Anonymity of trade and segmentation of markets imply that market makers in the first exchange can only observe y_1 while market makers in the second exchange can only observe y_2. Market makers in each exchange set prices $p_1 \in R^2$ and $p_2 \in R^2$ to clear the markets.

The expected profits of the first and second informed traders, conditional on their information (superscript t denotes the transpose operation), are

$$E\left[(\upsilon_1 - p_1)^t x_1^1 \,\middle|\, s_1\right] + E\left[(\upsilon_2 - p_2)^t x_1^2 \,\middle|\, s_1\right] \tag{4}$$

and

$$E\left[(\upsilon_1 - p_1)^t x_2^1 \,\middle|\, s_2\right] + E\left[(\upsilon_2 - p_2)^t x_2^2 \,\middle|\, s_2\right], \tag{5}$$

respectively. The only information a discretionary liquidity trader possesses is the knowledge of his own liquidity shock. The expected profit of the ith discretionary trader who trades the cross-listed asset is[6]

$$E\left[(\upsilon_1 - p_1)\begin{pmatrix} 0 \\ \alpha_i w_i \end{pmatrix} \,\middle|\, w_i\right] + E\left[(\upsilon_2 - p_2)\begin{pmatrix} 0 \\ (1-\alpha_i) w_i \end{pmatrix} \,\middle|\, w_i\right]. \tag{6}$$

An *equilibrium* is defined as price rules $p_1, p_2 : R^2 \to R^2$, strategies for the informed traders $x_1^i, x_2^i \in R^2$, $i = 1, 2$, and an order splitting decision for the discretionary liquidity traders $\{\alpha_i w_i\}_{i=1}^n$ such that

- The price rules satisfy the condition

$$p_1(y_1) = E[\upsilon_1 \mid y_1] \quad \text{and} \quad p_2(y_2) = E[\upsilon_2 \mid y_2]. \tag{7}$$

- x_1^i and x_2^i maximize equation (4) and equation (5), respectively, for each realization of s_i.
- $\alpha_i(w_i)$ maximizes equation (6) for each realization of w_i.

A *linear equilibrium* is an equilibrium in which (i) there exist two 2×2 matrices, Λ_1 and Λ_2, such that

$$p_1(y_1) = \Lambda_1 y_1 \quad \text{and} \quad p_2(y_2) = \Lambda_2 y_2,$$

(ii) the informed traders' strategies are linear in the signals, that is, there exist two 2×2 matrices, β_1 and β_2, such that

$$x_1^1 + x_1^2 = \beta_1 s \quad \text{and} \quad x_2^1 + x_2^2 = \beta_2 s,$$

and (iii) the order splitting decisions of all discretionary liquidity traders are identical. In particular, they are independent of their liquidity shocks, that is, $\alpha_i = \alpha$ for some constant α. In the following, we focus on the properties of a linear equilibrium.

Let the scalar λ_1 be the second row–second column entry of the matrix Λ_1. Then, λ_1 is the price impact of a market order for the cross-listed asset on the first exchange. Indeed, a change in the demand for the cross-listed asset on the first exchange by the amount Δ results in a price change of the cross-listed

[6] In equilibrium, these profits are negative, that is, they are payment for immediacy.

asset on the first exchange by $\lambda_1\Delta$. Similarly, λ_2, the second row–second column entry of the matrix Λ_2, is the price impact of a market order for the cross-listed asset on the second exchange.

THEOREM 1: *A linear equilibrium exists. Let the matrices F_1, F_2, and P be given by equations (1) and (2). Let the 2×2 matrices $R_1, R_2, \Lambda_1, \Lambda_2, \beta_1$, and β_2 and the scalar α, be a solution to the system of equations*

$$\begin{aligned} R_1 &= \begin{pmatrix} 1 & 0 \\ 0 & \alpha^2 \end{pmatrix} \\ \beta_1 &= \left(\Lambda_1 + \Lambda_1^t\right)^{-1} F_1 \\ \Lambda_1 &= F_1 P \beta_1^t \left(R_1 + \beta_1 P \beta_1^t\right)^{-1}, \\ R_2 &= \begin{pmatrix} 1 & 0 \\ 0 & (1-\alpha)^2 \end{pmatrix} \\ \beta_2 &= \left(\Lambda_2 + \Lambda_2^t\right)^{-1} F_2 \\ \Lambda_2 &= F_2 P \beta_2^t \left(R_2 + \beta_2 P \beta_2^t\right)^{-1} \\ \alpha &= \frac{\lambda_2}{\lambda_1 + \lambda_2}, \end{aligned} \tag{8}$$

subject to the second-order conditions that $-(\Lambda_1^t + \Lambda_1)$ and $-(\Lambda_2^t + \Lambda_2)$ are negative semidefinite matrices. Then the matrices $R_1, R_2, \Lambda_1, \Lambda_2, \beta_1$, and β_2 and the scalar α, form a linear equilibrium.

The proof of Theorem 1 is given in the Appendix. Notice that the second-order conditions imply that $\alpha \in [0, 1]$. To solve the system, we use the elimination method. For an arbitrary positive α, we solve the first three matrix equations in equation (8) subject to the second-order condition. To emphasize that the solution we obtain depends on the arbitrary choice of α, we write the solution as a function of α. In particular, we get

$$\lambda_1(\alpha) = \frac{1}{2\alpha} \frac{\alpha\left(a^2\sigma_1^2 + b^2\sigma_2^2\right) + |b|\sigma_1\sigma_2}{\sqrt{\left(\sigma_2 + \alpha|b|\sigma_2\right)^2 + \alpha^2 a^2 \sigma_1^2}} \tag{9}$$

and

$$\lambda_2(\alpha) = \frac{1}{2(1-\alpha)} \frac{(1-\alpha)\left(a^2\sigma_1^2 + b^2\sigma_2^2\right) + |a|\sigma_1\sigma_2}{\sqrt{(\sigma_2 + (1-\alpha)|a|\sigma_1)^2 + (1-\alpha)^2 a^2 \sigma_2^2}}. \tag{10}$$

To find the equilibrium α, we solve the last equation in equation (8) for α using equations (9) and (10). We have four possible solutions for α:

$$\left\{ \frac{\sigma_1(|a|\sigma_1 + \sigma_2)}{|a|\sigma_1^2 - |b|\sigma_2^2}, 0, 1, \frac{\sigma_1\left(|a|^3\sigma_1^3 + a^2\sigma_1^2\sigma_2 + b^2\sigma_2^2|a|\sigma_1 - \sigma_2^3 b^2\right)}{a^2\sigma_1^2|b|\sigma_2^2 - |a|^3\sigma_1^4 + \sigma_2^4|b|^3 + b^2\sigma_2^2|a|\sigma_1^2} \right\}. \tag{11}$$

The first solution does not satisfy the condition $\alpha \in [0, 1]$.[7] The solutions $\alpha = 0$ and $\alpha = 1$ are the trivial equilibria. If a discretionary liquidity trader conjectures that all other liquidity traders trade at the first (second) exchange, then regardless of what the values of a and b are, this trader also submits his orders to the first (second) exchange. The fourth possible solution in equation (11) is the interesting one, because this solution depends on the values of a and b. Note that the fourth solution is valid in a certain region. Nonetheless, there is a family of equilibria with properties that change continuously with the parameters a and b. We use that family of equilibria to present our comparative statics related to trading volume on the two exchanges. We find it simpler to fix b and let a vary. Because of the symmetry of the model, analogous results hold when we fix a and let b vary. The following corollary formalizes this result.

COROLLARY 1: *Let b, σ_1, and σ_2 be given. Then, for each a there is a linear equilibrium $\Lambda_1(a)$, $\Lambda_2(a)$, $\beta_1(a)$, $\beta_2(a)$, and $\alpha(a)$, such that each element of the matrices, $\Lambda_1(a)$, $\Lambda_2(a)$, $\beta_1(a)$, and $\beta_2(a)$ and the scalar $\alpha(a)$ are continuous in a. Moreover, $\alpha(a)$, the portion of trade submitted to the first exchange, increases with $|a|$. Furthermore, if $b^2\sigma_2^2 < \sigma_1^2 (a^2\sigma_1^2 < \sigma_2^2)$ and if*

$$a^2 \geq \frac{b^2\sigma_2^2}{\sigma_1^2}\frac{\sigma_1 + |b|\sigma_2}{\sigma_1 - |b|\sigma_2}\left(b^2 \geq \frac{a^2\sigma_1^2}{\sigma_2^2}\frac{\sigma_2 + |a|\sigma_1}{\sigma_2 - |a|\sigma_1}\right), \tag{12}$$

then all trade of the cross-listed asset takes place on the first (second) exchange.

The proof of the corollary is straightforward. If $b^2\sigma_2^2 < \sigma_1^2$ and $a^2\sigma_1^2 < \sigma_2^2$, then there is a linear equilibrium with

$$\alpha = \frac{\sigma_1\left(|a|^3\sigma_1^3 + a^2\sigma_1^2\sigma_2 + b^2\sigma_2^2|a|\sigma_1 - \sigma_2^3 b^2\right)}{a^2\sigma_1^2|b|\sigma_2^2 - |a|^3\sigma_1^4 + \sigma_2^4|b|^3 + b^2\sigma_2^2|a|\sigma_1^2}. \tag{13}$$

Differentiation with respect to $|a|$ shows that α is increasing in $|a|$. However, if $b^2\sigma_2^2 \geq \sigma_1^2$ and $a^2\sigma_1^2 \geq \sigma_2^2$, then equation (13) may be smaller than zero or greater than one. In the former case, we use the trivial equilibria in which $\alpha = 0$, and in the latter, we use the trivial equilibria in which $\alpha = 1$, so that the family of equilibria has properties that vary continuously in a.

It is worth contrasting the equilibrium solution of our model with those of the well-known multimarket trading models of Pagano (1989) and Chowdhry and Nanda (1991). Pagano shows that in his model an equilibrium exists in which both markets can survive, but only if a number of restrictive assumptions, such as equal transactions costs and an equal number of traders in each market, are satisfied. If these restrictions are not satisfied then only a single market remains open in equilibrium. In Chowdhry and Nanda's equilibrium, informed traders allocate more of their trades to the market that is "thick"

[7] Indeed, for the first solution to be positive, we must have $|a|\,\sigma_1^2 - |b|\,\sigma_2^2 > 0$, but this implies that the solution is greater than one.

with liquidity traders. However, the extent to which liquidity trade clusters in a given market is driven by exogenous assumptions about the number of "small" liquidity traders in each market. In contrast, we make no assumption regarding the distribution of small liquidity traders and we endogenize the choice of where to trade for both the informed and the discretionary liquidity traders. Two of the possible equilibria in our model are such that only a single market survives. However, in contrast to Pagano, these single market equilibria are not stable in the following sense: If we were to introduce in our model traders who are constrained to trade in one market (small liquidity traders), then this market would also attract informed and discretionary trading, that is, the trivial equilibria with concentration would disappear. Thus, the natural equilibrium outcome in our model is such that trade occurs in both markets. Moreover, the fact that the properties of the family of equilibria that we derive change continuously as a function of the parameters a and b allow us to derive cross-sectional predictions about differences in trading volume across stocks while holding other parameters constant.

Figure 2 illustrates the results presented in the corollary. The figure shows how the equilibrium portion of trading volume in the first exchange varies with the sensitivity, a, of the cross-listed asset's value to s_1. The graph displays three representative families of equilibria for three different ratios of the variance of s_1 relative to the variance of s_2, which is normalized to one. The figure demonstrates that for a sufficiently small, all trade takes place on the second exchange (i.e., $\alpha = 0$), while for a sufficiently large all trade occurs on the first exchange. The intuition for this result, which forms the basis of our empirical analysis, is as follows. When pricing an asset, market makers can infer information not only from the asset's own order flow, but also from the order flow of other assets traded on the same exchange. In equilibrium, the more correlated the returns of the two assets, the more relevant the order flows of one asset are for the pricing of the other asset, and the less sensitive the price of any one asset to its own order flow. This outcome increases the liquidity of the cross-listed asset, which guides the liquidity traders in choosing where to trade. That is, proportionally more volume takes place on the market in which the cross-listed asset has greater correlation with the other assets traded on that market. In all but extreme cases both markets remain open, however, because the discretionary liquidity traders split their orders across the two exchanges in order to minimize their average trading costs.[8]

B. Model Extension

One important assumption in our model is that before setting prices, the competitive risk-neutral market makers observe the net order flow of the assets

[8] Note that splitting orders remains an optimal strategy, even if one market has an absolute advantage in terms of overall liquidity. Menkveld (2004) provides evidence of order-splitting behavior in terms of a positive correlation of signed volume during the overlapping hours of trading in the Amsterdam and NYSE markets for Dutch cross-listed stocks.

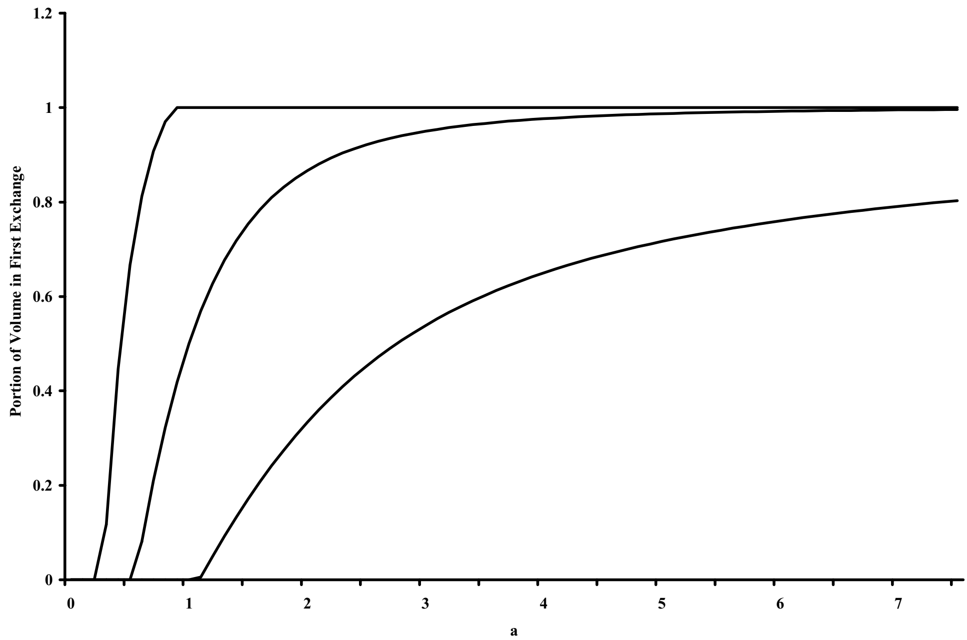

Figure 2. Volume share as a function of the sensitivity of the innovation in the value of the cross-listed asset to s_1. This figure shows how, for different values of σ_1, the portion of liquidity traders' orders submitted to the first exchange varies with the sensitivity of the cross-listed asset to s_1 as measured by a. The figure is generated assuming that $b = 1$ and $\sigma_2 = 1$, where b is the sensitivity of the cross-listed asset to s_1, σ_1 is the standard deviation of the innovation of the value of the local asset on the first exchange, and σ_2 is the standard deviation of the innovation of the value of the local asset on the second exchange. The upper line in the graph shows results for $\sigma_1 = 2$ and then in decreasing order $\sigma_1 = 1$ and $\sigma_1 = 0.5$.

traded on their own exchange, but not the order flow on the other exchange. This matters because local market makers are thus conferred a comparative advantage in understanding local innovations by virtue of their exclusive view of the local asset's order flow. How reasonable this assumption is in reality is unclear, as there may be competitive multimarket-trading scenarios in which the markets are so close (say, geographically or in the same time zone) that this complete unobservability restriction is false.[9]

In order to investigate the robustness of our model's predictions, we develop an extension to the model that allows for market makers to completely observe the order flow in the other market while maintaining some information asymmetry between the two groups of market makers. To keep the model tractable, we make two simplifying assumptions: (i) There exists only the cross-listed asset, and (ii) liquidity traders can choose how much to trade, but are constrained

[9] In fact, the model and empirical results in Domowitz et al. (1998) demonstrate how less-than-perfect transparency in quotes between two competing markets can impact liquidity and trading activity in both markets.

to do so only in their own market (no order splitting). Now, market makers in the two competing exchanges can observe demand for the cross-listed asset in both exchanges, but the special role previously played by the purely domestic assets in understanding the local innovation is replaced with an exogenous signal. The resulting equilibrium defines price rules that are linear in the exogenous signal that the market makers observe and in the order flows to *both* exchanges. More importantly, so long as the exogenous signals observed by the market makers are not identical, the key comparative statics obtain. That is, the differential price impact of market orders for the cross-listed asset on the two exchanges is directly related to the relative magnitude of the exposures of the asset's price innovations to the signals in the two markets. Thus, the higher the exposure to the signal on a given exchange, the greater the liquidity that exchange provides and the more aggressively discretionary liquidity traders trade on that exchange.

The model extension makes clear that it is not the unobservability of order flow per se that drives our predictions about volume. Rather, the necessary condition for our result is that the local market maker in each market has some comparative advantage relative to his counterpart in the other market in understanding the local innovation to which the cross-listed stock is partially exposed. Of course, there may be more complex generalizations not considered here for which these inferences may be less clear. The details of the extended model are furnished in an unpublished appendix (which is available from the authors upon request).

C. *Testable Alternative Hypotheses*

Overseas investors may wish to trade shares of cross-listed stocks in a particular market for reasons that are not related to informational motives at all. Additionally, trading can stem from informational advantages favoring one investor over another that are unrelated to our central hypothesis about market makers' inferring news from the order flows of other stocks trading in a given market. Within the framework of our model, these additional factors can be thought of as market frictions that keep traders captive in a specific trading venue. In this subsection, we propose a number of alternative testable hypotheses that we exploit in the empirical analysis to render additional power to the tests.

Among noninformation-related factors that might influence the extent of the U.S. share of cross-listed stocks' trading volume, differences in trading costs (commissions, transactions taxes) or regulatory restrictions (foreign ownership limits, currency convertibility constraints) are among the most important. Empirical evidence indicates that U.S. markets hold an absolute advantage over most other markets in terms of transactions costs,[10] and emerging markets clearly are subject to much higher transactions costs and regulatory barriers than developed markets, so it is important to control for these country-level

[10] See, among others, Table V of Chiyachantana et al. (2004).

differences in our empirical analysis. Because most of these factors manifest themselves at the country level, they cannot explain cross-sectional differences in the proportion of U.S. trading among stocks from the same country, which is the focus of this paper. However, there are also firm-specific barriers in some emerging markets that relate to the maximum number of home market shares that are allowed to become eligible to trade overseas in cross-listed form. This is the case, for example, in several Korean, Taiwanese, and Indian companies trading on major U.S. exchanges with explicit ceilings on the number of shares that can become ADRs.[11]

Cost considerations associated with the breadth and composition of the ownership base of the firm might also influence the extent of trading in different markets. For example, differences in transactions costs may matter less for retail than institutional investors, who are more strategic in their trade execution decisions. Institutional traders are more likely to split their orders across markets to minimize the net price impact of their trades, even if there is a perceived absolute advantage in terms of lower overall transactions costs in the United States. If this is true, a larger U.S. ownership base will be associated with a higher fraction of overall trading in the United States, but this effect will be even greater if the U.S. ownership base is dominated by retail investors more than strategically motivated institutions that actively seek to minimize their overall trading costs. These inferences will reverse, however, if institutions are constrained by prudential policies by their plan sponsors or oversight boards that limit them from directly holding stock abroad and that require them to hold ADRs or other forms of cross-listed shares instead.

Another noninformational factor is related to the diversification opportunities that cross-listed stocks present for overseas investors. Cross-listed stocks that have low return correlations with the U.S. market should be good candidates for diversification purposes.[12] This prediction is particularly useful to us as it represents a specific testable alternative to the central hypothesis of our model; that is, cross-listed stocks whose returns are weakly correlated with other U.S. stocks furnish market makers lower quality information about order flow dynamics and are *less* likely to attract U.S. trading volume.

Informational motives that are not necessarily related to the correlation structure of returns may also be an important factor in determining the global distribution of trading volume for cross-listed stocks. For example, it has been argued that local investors have an informational advantage relative to foreign

[11] These ceilings are usually stipulated in the Securities and Exchange Commission annual filings of Form 20-F in the section entitled "Description of American Depositary Shares" outlined with the conditions under which the ceiling can change over time.

[12] Many theories abound to explain the "home-bias puzzle" whereby portfolio holdings of investors unduly favor local over foreign stocks in spite of attractive risk diversification opportunities. Such explanations parallel the alternative hypotheses we present for the geographic location of trading in cross-listed stocks, including market frictions (Black (1974), Stulz (1981)), home assets as hedges against home country-specific risks (Adler and Dumas (1983), Cooper and Kaplanis (1994)) or differences across investors in prior beliefs about expected returns or risks (Pástor (2000), Glassman and Riddick (2001)).

investors.[13] The information hierarchy that favors local investors may arise from geographic proximity to the company's main business operations (Coval and Moskowitz (2001)), language barriers for foreign investors (Grinblatt and Keloharju (1999), Hau (2001)), or a greater familiarity with a firm's operations, accounting practices, or capital market environment. While such advantages need not affect the location of trading volume, in many cases it will be cheaper and quicker to trade in the home market, especially if the information is short-lived and speed of execution is critical. Based on this argument, we predict that firms with certain attributes will mitigate this informational disadvantage for foreigners. For example, the fact that larger companies are more visible than smaller firms should lead to a positive correlation between market capitalization and foreign trading volume (Merton (1987), Kang and Stulz (1997)). If a familiarity bias exists among investors (Grinblatt and Keloharju (1999), Hau (2001)), then cross-listed companies with a larger fraction of foreign sales should be more likely to develop an active foreign market for the trading of their shares. Similarly, greater familiarity may arise from a larger contingent of industry peers trading in one venue relative to another, a factor that some have suggested as a rationale for cross-listings in the first place (Pagano, Roell, and Zechner (2002), Sarkissian and Schill (2004)). If the total market capitalization of peer firms from the same global industry sector as the cross-listed firm is higher in the foreign market than in the firm's home market, a higher foreign trading volume can be expected to arise.[14] Finally, more analyst coverage can increase the amount of public information available to all investors (Lang et al. (2003), Bailey et al. (2006)), so we predict that the number of analysts following cross-listed stocks is positively related to foreign trading volume.

In the next section, we describe the empirical tests of our theory. We first identify the sample of U.S. cross-listed stocks; we then outline how our information factor variable is constructed. In addition, we discuss various proxy variables that we use to evaluate the testable alternative hypotheses defined above.

II. Empirical Analysis

A. Data and Sample

The primary prediction of the model, as expressed in Corollary 1, is that when the value of the cross-listed stock is more sensitive to information in the U.S. market relative to information in the home market, the U.S. will have a higher share of the overall trading volume in the stock. To examine this prediction we collect an initial sample of all stocks cross-listed on U.S. exchanges with

[13] This assumption is key to a number of recent explanations and findings related to the home bias phenomenon in international portfolio holdings (Kang and Stulz (1997), Grinblatt and Keloharju (1999), Ahearne, Griever, and Warnock (2004)).

[14] This specific alternative hypothesis is a significant challenge for our model's prediction, as the returns correlation of the cross-listed stock with other stocks trading in the new venue should be influenced by the presence of a greater number of industry peers among those other stocks.

listing dates prior to January 1, 1999 that also have stock price data available for both the U.S. and the home market from Datastream International. This initial sample consists of 384 cross-listed stocks with daily data on stock price, trading volume, and the market value of equity over the period beginning January 1991 and ending December 2004. Cross-listed securities, which trade in the United States in the form of ADRs, are often bundled in a ratio different from one-to-one. For example, the ADR for Telefonos de Mexico, a Mexican company, represents 20 underlying Mexican shares, while the ADR for Norsk Hydro, a Norwegian company, represents a single underlying Norwegian share. To compute the U.S. share of trading volume in each stock we require an accurate measure of the bundling ratio between the home market and cross-listed securities. We obtain the most recent bundling ratio for each stock from Citibank's Universal Issuance Guide. Unfortunately, many cross-listed securities alter their bundling ratios over time. To ensure that we obtain an accurate representation of the bundling ratios for the stocks in the sample we compare the bundling ratios from the Citibank Universal Issuance Guide to the yearly average of the ratio of the month-end price in the U.S. market to the month-end price in the home market (converted to U.S. dollars). If the price ratio differs from the reported bundling ratio by more than 20% we attempt to reconcile the discrepancy by examining other sources (such as the ratios reported along with the firm names in Datastream). After eliminating stocks for which we are unable to obtain an accurate bundling ratio we are left with a final sample of 251 cross-listed securities representing 24 different countries. By way of comparison, Pulatkonak and Sofianos (1999) employ a sample of 254 stocks cross-listed on the NYSE. Pulatkonak and Sofianos also provide a detailed description of the different types of cross-listed securities.

B. Variable Construction and Econometric Model

To test the prediction of the model we require a measure of the sensitivity of the stock's value to information in the United States relative to information in the home market. Eun and Sabherwal (2003) compute a measure of the U.S. information share using an error correction model and intradaily data for a sample of 52 Canadian stocks. While their measure is intuitively appealing, it cannot be employed in our setting because we lack intradaily data, and furthermore, because many of the stocks in our sample are not traded during the same hours in both the U.S. and home markets due to differences in time zones across countries. Instead we take a more macro view in measuring the sensitivity of the stock's value to information in the U.S. market. Our underlying assertion is that stock prices respond primarily to new information, and we focus on the relative informativeness of market movements in the United States and the home country. Specifically, for each stock, we perform a variance decomposition of returns converted to U.S. dollars to estimate the contribution of information contained in U.S. market index returns relative to the information content of index returns in the home market. To perform the variance decomposition we first estimate the following two time-series regressions for each stock:

$$R_{it} = \alpha_i + \sum_{k=-1}^{+1} \beta_{i,H,t+k} R_{\mathrm{Home},t+k} + \varepsilon_{it}$$

and (14)

$$R_{it} = \alpha_i + \sum_{k=-1}^{+1} \beta_{i,H,t+k} R_{\mathrm{Home},t+k} + \sum_{k=-1}^{+1} \beta_{i,\mathrm{US},t+k} R_{\mathrm{US},t+k} + \varepsilon_{it},$$

where R_{it} is the return (measured in U.S. dollars) of stock i in period t, $R_{\mathrm{Home},t+k}$ is the return denominated in U.S. dollars on the market index in the stock's home country in period $t + k$, and $R_{\mathrm{US},t+k}$ is the dollar-denominated return on the U.S. market index in period $t + k$. The lead and lag terms in the regressions are used to account for nonsynchronous trading across markets in different time zones. In addition, to deal with the possibility of infrequent trading at the stock level we base our analysis on weekly returns. The first regression in equation (14) is considered the restricted regression and the second regression in equation (14) is considered the unrestricted regression. Assuming n observations for the stock, six regressors in the unrestricted model, and three restrictions, we compute an F-statistic for each stock that measures the explanatory power of the unrestricted model relative to the explanatory power of the restricted model as follows:

$$\frac{(R_{UR}^2 - R_R^2)/3}{(1 - R_{UR}^2)/(n-6)}. \tag{15}$$

By construction, this measure captures the incremental contribution of U.S. market movements in explaining variation in the firm's stock price over and above the information about the firm's stock price contained in movements in the firm's home market index. Our theory therefore predicts that higher values of this U.S. "information factor" correspond to higher U.S. shares of overall trading volume. The U.S. information factor is computed for each stock and each sample year using a minimum of 36 months and a maximum of 48 months of past data. This estimation period ensures that we obtain reasonably precise estimates of the U.S. information sensitivity of the stocks.[15] To the extent that the information factor contains measurement error, we expect an attenuation bias toward zero for the coefficient estimate on this variable in our cross-sectional regression analysis, which biases against finding support for our hypothesis.

For each year in the sample, the U.S. share of overall trading volume for each stock is computed as the weekly average across the year of the ratio of the number of shares traded in the U.S. market, adjusted using the bundling ratio of U.S. shares to home market shares, divided by the total trading volume in both the U.S. and the home market. In computing the U.S. share of trading volume we include only trading volume in the U.S. and the home market. Some of the

[15] Similar results obtain in our empirical analysis if we use a 3-year estimation period.

securities are also cross-listed in London or on other global exchanges, such as Frankfurt. We do not include the trading volume from the other exchanges in our measure of the U.S. share of trading volume.

Our information factor is computed using returns and it is possible that returns and volume are contemporaneously correlated for reasons outside of those predicted by our model. Hence, it is important that our information factor is computed using data from a different period than from the period in which we measure trading volume. To do so, we employ a panel approach and lag the U.S. information sensitivity measure 1 year. Specifically, the measure of the stock's U.S. information sensitivity computed using returns from years 1991 to 1994 is used to predict the stock's U.S. volume share in 1995 and so on through the end of the sample period. This experimental design ensures that the information factor is predetermined at the time that it is included as an independent variable in the regression. Stocks are allowed to enter and exit the panel over time. Our panel contains 1,416 observations for 251 cross-listed stocks covering the years 1995 through 2004.

In addition to the U.S. information factor we also control for several other factors that could affect the U.S. share of overall trading volume. These other factors are related to the testable alternative hypotheses discussed in the previous section. First, there are a number of country-level factors that might affect the U.S. share of trading volume at the country level. For example, there are systematic differences in market size and development, trading commissions, and the legal environment across countries. In addition, Pulatkonak and Sofianos (1999) find that a primary factor affecting differences in the U.S. share of trading volume across stocks is that of time zone effects: Stocks from countries with a time zone closer to that in the U.S. exhibit a higher share of U.S. trading volume. In some of our specifications, we control for time zone effects using a set of indicator variables for different time zones classified in 3-hour increments relative to New York. For example, Japan and Korea (which are 10 hours behind New York) and Australia (which is 9 hours behind New York) are classified in time zone −3, while Hong Kong and the Philippines, which are both 13 hours ahead of New York, are classified in time zone 4. In the specifications using time zone indicators we also include an indicator for emerging markets versus developed markets, where we use the emerging market definition supplied by the International Finance Corporation (IFC). We predict that emerging markets are more likely to have binding regulatory constraints and higher overall trading costs compared to developed markets, thus leading to proportionally more U.S. trading volume for these stocks. In other specifications, we control directly for country-level differences using country fixed effects. The all-encompassing use of country fixed effects allows us to focus on differences across stocks within countries while removing the effect of all time-invariant factors that are common to all firms within a country.

Our regression specifications also include a variety of firm-specific measures that could potentially affect the stock's U.S. share of trading volume. First, because all of the stocks in our sample are listed on the NYSE or Nasdaq/Amex, we include an indicator equal to one for firms listed on Nasdaq/Amex in order to account for systematic differences in either actual or reported trading volume across listing exchanges. Next, we control for firm size using the natural logarithm of the stock's average market value of equity converted to U.S. dollars. Firm size might proxy for the familiarity of the firm to U.S. investors. If this is the case, then we expect firm size to be positively related to the U.S. share of trading volume. Another reason to control for firm size is that large firms are likely to make up a substantial proportion of the market capitalization of the market index in the home country. This could induce a spurious positive correlation between the firm's return and the return on the home market index and bias the U.S. information factor for the stock downward. As additional proxies for the visibility of the firm to U.S. investors, we use the percentage of foreign sales reported by the firm, and the difference in the percentage of the global market capitalization of the firm's industry located in the U.S. and the percentage of global industry market capitalization for the firm's industry in the home country. The data on foreign sales are drawn annually from the Worldscope database (Item WC08731) based on the fiscal year-end of December 31. The relative industry capitalization variable comes from Datastream International and is based on Level 3 industry grouping codes for the home markets (which include 10 global sectors). We expect both foreign sales and the U.S. share of the firm's industry to be positively associated with the U.S. share of trading volume. To capture firm-specific constraints on the U.S. share of trading volume due to regulatory restrictions on foreign ownership, we collect the annual investable weights from Standard and Poor's Emerging Market Database. These weights measure the fraction of the total market capitalization of the firm that is free from explicit foreign investment restrictions.

Higher analyst coverage in the home market potentially reduces the costs of information acquisition for foreign investors trading in the local market. We measure analyst coverage using the natural logarithm of one plus the number of analysts that report earnings estimates for the stock in the home country. These data are from the International Summary database from Thomson Financial's Institutional Brokers' Estimate System (I/B/E/S). Following Lang et al. (2003), we define analyst coverage for a cross-listed share as the number of estimates that comprise the consensus 1-year-ahead earnings estimate for that fiscal year. The data are obtained at a monthly frequency as of the third Thursday of each month, and we use the December values to measure analyst coverage for the following year. We expect higher analyst coverage in the home market to be negatively related to the U.S. share of trading volume; however, this relationship could be complicated if the information generated by the home market analysts works to overcome, rather than exacerbate, the informational advantage of locals over foreigners.

Finally, we control for the fraction of the firm's shares owned by U.S. institutions. We would ideally like a breakdown of the full ownership base of these cross-listed firms, but such detailed data are not readily available. Thomson Financial's 13F database provides quarterly data on the percentage of shares held by U.S. institutions based on filings to the Securities and Exchange Commission. We collapse the quarterly data into annual horizons to align the data with our panel regression analysis. To the extent that institutional ownership is correlated with the overall size of the U.S. investor base, and if the U.S. market offers absolute advantages in trading costs relative to the home market, then the share of U.S. trading volume will be positively correlated with U.S. institutional ownership. This prediction is less clear, however, if the U.S. institutional ownership base dominates the U.S. retail base. If the more strategically motivated institutions are more inclined to split their orders across markets to minimize the overall net price impact of their trades, then the U.S. share of trading volume will be less positively correlated with U.S. institutional ownership. If however these U.S. institutions are bound by prudential policies that preclude them from directly holding stock abroad and that require them to hold the cross-listed shares instead, then the U.S. share of trading volume will be more positively correlated with U.S. institutional ownership.

C. Summary Statistics

Table I reports summary statistics for the stocks in our sample for each country within subsamples of developed (Panel A) and emerging (Panel B) markets. The sample consists of 201 stocks from the 14 countries classified as developed markets and 50 stocks from the 10 countries classified as emerging markets. Cross-listed stocks from Canada (78) and the United Kingdom (40) make up a substantial portion of the sample. However, several other countries, including Australia (13), France (14), Japan (19), the Netherlands (14), Chile (12), and Mexico (14), are also well represented. The U.S. share of trading volume for cross-listed firms from emerging markets averages 33%, which is nearly twice the average U.S. share of trading volume of cross-listed firms from developed markets (17%). There is also substantial variation in the average U.S. share of trading volume across countries, varying from a low of about 2% for Italy and Switzerland to a high of around 60% for Chile and Mexico. The average value of the U.S. information factor in emerging markets is 2.02, slightly higher than the average of 1.83 for cross-listed firms in developed markets. This suggests that, on average, movements in the U.S. market provide more incremental price-relevant information for cross-listed stocks from emerging markets relative to those from developed markets.

The average market value of equity for stocks from developed markets is \$8.43 billion compared to \$7.79 billion for the average emerging market stock, although these values are highly skewed and vary widely across countries. In terms of the U.S. venue of listing, 33% of cross-listed firms in developed markets are listed on Nasdaq compared to only 22% of cross-listed firms in emerging markets. On average, approximately 20 analysts follow firms in the home country in developed markets compared to an average of about 14 analysts in emerging markets. The number of home country analysts varies greatly across countries, however. The percentage of shares held by U.S. institutions averages 8.88% in developed markets and 6.74% in emerging markets. China, Hong Kong, Taiwan, and Italy exhibit the lowest levels of U.S. institutional ownership, while Canada, Finland, and Israel exhibit high levels of ownership

Table I

Sample Summary Statistics

The table reports sample summary statistics by country for our sample of 251 cross-listed firms from 24 countries using a panel of annual data from 1995 through 2004. Individual country means are reported along with overall means for developed and emerging market countries as defined by the IFC. The table reports the number of cross-listed firms in each country; the U.S. share of trading volume; the U.S. information factor computed using equations (14) and (15); market value of equity (in USD millions); the fraction of firms listed on Nasdaq; the number of analysts following the firm in the home country (Source: I/B/E/S International Summary Database); U.S. institutional ownership as a fraction of shares outstanding (Source: Thomson Financial's 13F Database); the investable weight, which is the fraction of the market capitalization of the individual firm accessible for foreign investors (Source: Standard & Poor's Emerging Market Database); the U.S. industry relative, which is the difference between the share of global market capitalization of the firm's industry in the U.S. and that in the home country (based on Datastream's Level 3 sector definitions); and the percentage of foreign sales (Source: Worldscope item WC08731).

Panel A. Developed Markets

	Number of Firms	U.S. Volume Share $N = 1{,}416$	U.S. Information Measure $N = 1{,}416$	Market Value of Equity (USD) $N = 1{,}416$	Nasdaq Indicator $N = 1{,}416$	Home Analyst Coverage $N = 1{,}073$	U.S. % Institutional Ownership $N = 1{,}282$	Investable Weight (%) $N = 1{,}312$	U.S. Industry Relative (%) $N = 1{,}063$	Foreign Sales (%) $N = 822$
Australia	13	0.15	2.92	8.12	0.32	10.92	1.38	100.0	35.40	33.52
Canada	78	0.42	1.85	6.64	0.47	11.55	23.58	100.0	42.24	67.68
Finland	1	0.40	1.11	11.34	0.00	42.17	30.98	100.0	68.28	78.51
France	14	0.07	1.60	8.56	0.49	18.65	7.07	100.0	45.62	59.56
Hong Kong	1	0.42	3.77	4.99	0.00	8.75	0.60	100.0	26.38	90.86
Italy	6	0.02	0.95	9.00	0.03	29.88	0.50	100.0	26.81	48.99
Japan	19	0.03	2.10	9.54	0.49	15.64	2.48	100.0	23.50	40.48
Netherlands	14	0.24	2.87	9.16	0.34	22.92	11.51	100.0	48.24	64.50
New Zealand	1	0.12	1.15	5.41	1.00	9.00	5.11	100.0	51.90	0.00
Norway	4	0.21	1.12	7.77	0.11	19.89	15.26	100.0	40.79	79.84
Spain	3	0.06	1.08	10.28	0.00	33.38	5.68	100.0	32.45	NA
Sweden	5	0.13	1.41	9.07	0.97	26.77	7.79	100.0	43.67	85.06
Switzerland	2	0.02	1.84	9.46	0.00	18.57	10.19	100.0	44.37	59.01
United Kingdom	40	0.09	1.86	8.69	0.37	12.23	2.22	100.0	36.49	55.65
Developed Mean		0.17	1.83	8.43	0.33	20.02	8.88	100.0	40.44	58.74

(continued)

Table I—*Continued*

Panel B. Emerging Markets										
	Number of Firms	U.S. Volume Share $N = 1{,}416$	U.S. Information Measure $N = 1{,}416$	Market Value of Equity (USD) $N = 1{,}416$	Nasdaq Indicator $N = 1{,}416$	Home Analyst Coverage $N = 1{,}073$	U.S. % Institutional Ownership $N = 1{,}282$	Investable Weight (%) $N = 1{,}312$	U.S. Industry Relative (%) $N = 1{,}063$	Foreign Sales (%) $N = 822$
Chile	12	0.58	1.20	7.01	0.00	8.91	1.99	79.2	41.46	47.31
China	5	0.13	2.80	7.54	0.00	4.60	0.10	85.6	32.80	11.68
Indonesia	2	0.30	5.32	7.73	0.00	21.36	0.46	NA	32.88	0.00
Israel	5	0.46	1.97	6.83	0.73	2.68	21.14	89.2	54.95	67.43
Korea	1	0.16	1.08	9.34	0.00	19.00	9.59	32.9	41.21	0.00
Mexico	14	0.61	0.97	6.95	0.00	14.52	5.47	90.5	43.05	34.69
Philippines	1	0.49	0.69	7.69	0.00	21.50	18.64	NA	33.71	0.00
Portugal	2	0.03	2.13	9.07	0.00	21.38	5.67	NA	36.35	22.94
South Africa	6	0.35	2.09	6.90	0.90	8.32	3.74	96.5	33.27	30.12
Taiwan	2	0.20	1.94	8.84	0.56	16.57	0.60	NA	67.63	11.80
Emerging Mean		0.33	2.02	7.79	0.22	13.88	6.74	78.9	41.73	22.60

by U.S. institutions. The investable weight is 100% for all developed market firms, but averages only 79% for emerging market firms (interestingly, Korea is the country with the lowest fraction of accessibility at 33% over this period). The difference in the percentage of the global market capitalization of a firm's industry located in the U.S. and that in the home market (U.S. industry relative) is about 40% on average across both developed and emerging markets. The average percentage of foreign sales reported by firms is nearly 59% in developed markets compared to only 22% for firms in emerging markets. Finally, we also note that several of our variables are not uniformly available for all of the observations in our panel. Investable weights (1,312 observations) and U.S. institutional ownership (1,282 observations) have the broadest coverage, and the percentage of foreign sales the least (822 observations). We report all of our subsequent results using the largest number of observations possible.

Table II reports the correlation matrix for the variables. Because our interest is primarily in whether the model can explain differences in the share of U.S. trading volume across firms within countries, the correlations are reported after subtracting the country mean from each variable. Consistent with the prediction of our model, the U.S. information factor and the U.S. share of trading volume are positively correlated. The signs of the correlations on the other variables are also generally consistent with intuition. Larger firms have a lower share of U.S. trading volume. The U.S. share of trading volume is lower when more analysts follow the firm in the home country, and is higher when U.S. institutions hold a larger fraction of the firms' shares and when the firms have higher investable weights (fewer firm-level foreign ownership constraints). The U.S. share of trading volume is also higher for firms with more foreign sales, when the firm's industry is more overrepresented (by market capitalization) in the U.S., and when the investable weight is higher.

D. Multivariate Regression Analysis

To further investigate the factors that affect the U.S. share of trading volume, Table III reports results of a multivariate regression analysis. The dependent variable in the regressions is the logistic transformation of the U.S. share of trading volume. We employ the logistic transformation to account for the fact that the U.S. volume share is bounded between zero and one. All of the specifications include the U.S. information factor from equation (15), firm size, and the indicator variable for cross-listed firms listed on Nasdaq/Amex. In some specifications we also include various combinations of the other control variables. To control for country-level differences, all of the regressions include either the indicators for time zone and emerging markets or country fixed effects. The table reports White (1980) heteroskedasticity-consistent t-statistics adjusted for clustering at the firm level. All of the independent variables are lagged 1 year.

The results reported in Table III provide strong support for the prediction of the model. In model (1), which employs the broadest possible sample and controls for country effects using the time zone and emerging market indicators, the coefficient on the U.S. information factor is 0.144 (t-statistic =

Table II

Correlation Matrix of Variables

The table reports the correlation matrix for our sample of 251 cross-listed firms from 24 countries. The variables include the U.S. share of trading volume; the U.S. information factor computed using equations (14) and (15); market value of equity (in USD millions); the fraction of firms listed on Nasdaq; the number of analysts following the firm in the home country (Source: I/B/E/S International Summary Database); the U.S. institutional ownership as a fraction of shares outstanding (Source: Thomson Financial's 13F Database); the investable weight, which is the fraction of the market capitalization of the individual firm accessible for foreign investors (Source: Standard & Poor's Emerging Market Database); U.S. industry relative, which is the difference between the share of global market capitalization of the firm's industry in the U.S. and that in the home country (based on Datastream's Level 3 sector definitions); and the percentage of foreign sales (Source: Worldscope item WC08731). Correlations are measured after subtracting the country mean from each variable, and p-values are reported in parentheses.

	U.S. Volume Share	U.S. Information Factor	Market Value of Equity	Home Analyst Coverage	U.S. Institutional Ownership	Investable Weight	U.S. Industry Relative	Foreign Sales
U.S. volume share	1.000							
U.S. information factor	0.125	1.000						
	(0.000)							
Market value of equity	−0.147	0.121	1.000					
	(0.000)	(0.000)						
Home analyst coverage	−0.317	0.151	0.630	1.000				
	(0.000)	(0.000)	(0.000)					
U.S. institutional ownership	0.230	0.103	0.257	0.231	1.000			
	(0.000)	(0.000)	(0.000)	(0.000)				
Investable weight	0.079	−0.098	−0.062	0.022	0.055	1.000		
	(0.004)	(0.000)	(0.026)	(0.498)	(0.057)			
U.S. industry relative	0.126	0.106	−0.188	−0.183	−0.070	0.080	1.000	
	(0.000)	(0.001)	(0.000)	(0.000)	(0.032)	(0.012)		
Foreign sales	0.085	−0.002	0.013	0.037	0.116	0.000	0.064	1.000
	(0.014)	(0.958)	(0.716)	(0.284)	(0.002)	(0.995)	(0.069)	

Table III

OLS Regressions of the U.S. Share of Trading Volume

The sample consists of 251 cross-listed firms from 24 countries for the period 1995 to 2004. The dependent variable in the regressions is the logistic transformation of the U.S. share of trading volume. Independent variables include the U.S. information factor, the natural log of the market value of the firm's equity in USD, an indicator equal to one for Nasdaq listings, home analyst coverage measured as the natural log of one plus the number of analysts reporting earnings forecasts for the firm in the home country, the percentage of outstanding shares owned by U.S. institutions, the difference between the share of the global market capitalization of the firm's industry in the U.S. and the industry share in the home country, the investable weight as measured by the fraction of the firm's equity market capitalization accessible to foreign investors, and the percentage of foreign sales. Some regression specifications include an indicator equal to one for emerging markets and the time zone of the country (classified in 3-hour increments relative to New York) and some specifications include indicator variables for each of the 24 countries in the sample. Robust heteroskedasticity consistent t-statistics adjusted for clustering at the firm level are reported in parentheses.

	Dependent Variable: Logistic Transformation of U.S. Share of Trading Volume							
Variable	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)	Model (6)	Model (7)	Model (8)
U.S. information factor	0.144	0.102	0.084	0.067	0.121	0.113	0.100	0.086
	(4.15)	(2.80)	(2.15)	(1.74)	(3.36)	(2.88)	(2.35)	(2.02)
Market value of equity	−0.138	−0.051	−0.0582	−0.006	−0.116	0.020	−0.004	0.033
	(−1.79)	(−0.49)	(−0.54)	(−0.06)	(−1.60)	(0.21)	(−0.04)	(0.31)
Nasdaq indicator	0.269	0.204	0.267	0.373	0.309	0.410	0.420	0.393
	(0.93)	(0.67)	(0.86)	(1.17)	(1.08)	(1.35)	(1.38)	(1.18)
Home analyst coverage	NA	−0.477	−0.465	−0.475	NA	−0.672	−0.670	−0.681
		(−1.88)	(−1.75)	(−1.53)		(−2.54)	(−2.46)	(−2.17)
U.S. institutional ownership	NA	0.046	0.045	0.054	NA	0.043	0.044	0.050
		(5.40)	(5.28)	(5.86)		(5.01)	(5.02)	(5.48)
U.S. industry relative	NA	0.010	0.009	0.010	NA	0.004	0.001	0.004
		(1.37)	(1.22)	(1.30)		(0.57)	(0.18)	(0.52)
Investable weight		NA	0.914	0.369	NA	NA	3.247	4.312
			(0.76)	(0.19)			(3.57)	(3.04)
Foreign sales	NA	NA	NA	0.001	NA	NA	NA	0.001
				(0.79)				(0.91)
Emerging market indicator	1.353	1.226	1.463	0.982	NA	NA	NA	NA
	(3.54)	(2.79)	(2.51)	(1.33)				
Time zone indicators	Yes	Yes	Yes	Yes	No	No	No	No
Country indicators	No	No	No	No	Yes	Yes	Yes	Yes
N	1,416	958	888	720	1,416	958	888	720
Adj. R-squared	0.4197	0.4997	0.4854	0.4890	0.4710	0.5349	0.5313	0.5318

4.15), indicating that firms with a higher value of the U.S. information factor have higher U.S. shares of trading volume. The coefficient estimates on the control variables also appear reasonable. Larger firms exhibit a lower share of trading volume in the U.S., and firms from emerging markets have a higher U.S. share of trading volume, all else equal. The Nasdaq indicator is not statistically different from zero. Model (2) adds additional control variables for home analyst coverage, U.S. institutional ownership, and the relative share of the firm's industry located in the U.S. Model (3) adds the investable weight and model (4) adds foreign sales to the specification as well. Because of missing values, the number of observations drops from 1,416 in model (1) to 958 in model (2), 888 in model (3), and 720 in model (4). In all four specifications, the coefficient on the U.S. information factor remains positive and statistically significant at the 10% level or better. Models (5) through (8) repeat the regressions in models (1) through (4) but drop the time zone and emerging market indicators in favor of country fixed effects. The results are very similar. The coefficient estimate on the U.S. information factor in model (5) is 0.121 (t-statistic $= 3.36$) indicating a slightly smaller effect of U.S. information on the U.S. share of trading volume compared to model (1) after controlling for country-level differences. Compared to model (1) the negative effect of firm size on the U.S. volume share is slightly smaller. Models (6) to (8) yield results similar to those in models (2) to (4), with one notable exception: The explanatory power of the investable weight variable is much larger and positive with country dummies and without the emerging markets indicator. This result suggests that there is strong clustering in the investable weight variable among emerging markets.

Overall, the multivariate analysis reported in Table III provides support for the idea presented in our model that cross-listed firms for which price-relevant information contained in U.S. market movements is larger will also exhibit a larger share of their trading volume in the U.S. market. In addition, firms with more U.S. institutional ownership are associated with more U.S. trading volume and firms with more home country analyst coverage exhibit less trading volume in the United States. Moreover, the higher is the investable weight for a cross-listed firm, the higher the U.S. share of trading volume. The estimates on the U.S. share of the firm's industry and foreign sales are not significant in any of the specifications. To assess the magnitude of these effects we compute the percentage change in trading volume implied by a movement from the 25th to the 75th percentile in each of the variables of interest based on the regression coefficients in model (6).[16] The effect of the U.S. information factor on the U.S. share of trading volume is economically meaningful, but is somewhat smaller in magnitude than the effects of analyst coverage and institutional ownership. Moving from the 25th to the 75th percentile of the U.S. information

[16] To compute the economic effects, we set the value of all the continuous variables in the regression to their sample mean values. The Nasdaq indicator is set to zero, and the constant term in the regression is set such that the predicted U.S. share of trading volume for this "average" stock is 25% (which is approximately the sample average). The variable of interest is then varied between the 25th and 75th percentiles (based on within country variation) and the implied U.S. volume share is computed by inverting the logistic transformation.

factor implies an increase in the U.S. share of trading volume of over 15%. This compares to a 34% decline in the U.S. share of trading volume for a change of similar magnitude in home market analyst coverage and a 26% increase associated with a change of a similar magnitude in institutional ownership.

One additional robustness test we investigate relates to the concern about the restrictive assumption about the unobservability of order flow in the competing market for market makers. In Section II.B we develop an extended model that demonstrates that the key prediction of our model holds when observability of order flow across markets is allowed so long as the local market maker maintains some informational advantage relative to the foreign market maker in understanding the local portion of the innovation in the value of the cross-listed asset. Alternatively, if the observability of order flow eliminates the comparative advantage of the local market maker, then the explanatory power of our information factor for the U.S. fraction of trading volume should be greatly diminished. We perform supplementary regression analysis to test this possibility. As a proxy for the degree of observability among market makers in competing exchanges we create an indicator variable for countries that have a substantial overlap in trading hours with the United States. The overlap variable is equal to one for cross-listed stocks from Canada, Mexico, and Latin America, and to zero for cross-listed stocks from Asia and Europe. Using specifications similar to model (1) in Table III, but including the overlap variable (in lieu of time zone indicators and when country indicators are not included) and an interaction term between the overlap variable and the U.S. information factor, we find that the overlap variable is positively and significantly related to the U.S. trading volume, indicating that stocks from countries in a similar time zone to the United States exhibit a greater share of U.S. trading volume. Nevertheless, the information factor remains statistically significant and positively related to the fraction of trading volume. The interaction variable is positive but not significant, indicating that if anything, the effect of the information factor on the share of U.S. trading volume is higher in markets with more overlap in trading hours with the United States.

III. Conclusions

In considering whether to cross-list their stock on a foreign exchange such as the NYSE or Nasdaq in the United States, an issue of concern for managers is the trading activity that the cross-listing will attract. The distribution of trading volume potentially reflects on the long-term viability of the listing and its potential as a vehicle for raising capital, broadening the shareholder base, and enhancing its visibility and profile. In addition, understanding the factors that affect the distribution of trading volume in cross-listed securities is also important for stock exchanges as they attempt to attract new listings and compete with one another for order flow. The issue is particularly relevant today as many non-U.S. firms are threatening to deregister and delist from U.S. markets following the passage of the 2002 Sarbanes–Oxley Act given its increased costs of compliance, requirements of greater auditing independence, and penalties for false reporting.[17] Indeed, in part to alleviate these concerns, the Securities and Exchange Commission proposed in December 2005 a new Rule 12h-6 that would ease the burden on non-U.S. firms to arrange termination of registration in the United States, one important condition of which is based on the fraction of their overall trading volume that takes place in the United States.[18]

In this paper, we develop and test a theoretical model of multimarket trading to explain the differences in the foreign share of trading volume of internationally cross-listed stocks. The model predicts that, under fairly general conditions, the distribution of trading volume across exchanges competing for order flow is related to the correlation of the cross-listed asset returns with returns of other stocks in the respective markets. That is, volume is proportionally higher on the exchange in which the cross-listed asset returns have greater correlation with returns of other assets traded on that market. We test this prediction using monthly stock returns and volume data on 251 non-U.S. stocks cross-listed on major U.S. exchanges. We find strong empirical support for the prediction, even after controlling for possible endogeneity effects and for other firm-specific, issue-specific, and country-level factors.

Our work contributes to both the theoretical and the empirical literature on multimarket trading and provides an explanation for the great variability that is observed in the distribution of trading activity across exchanges in cross-listed securities. Nevertheless, there are many questions that we leave unanswered. For instance, identifying the fundamental forces underlying the information factor that we introduce as an important explanatory variable for the global distribution of trading in cross-listed securities remains an open question. One possibility is that the correlation of the returns of the cross-listed and domestic assets arise from the nature of the firm's business activities related to geography (important U.S.-based operations and assets) or industrial membership (globally competitive industry) that we have not captured fully in our foreign sales and industry-peer proxy variables. A more detailed empirical analysis of this question would be worthwhile, especially given the evidence in Pagano et al. (2002) and Sarkissian and Schill (2004) that trade, colonial ties, common language and culture, and similar industrial structure play important roles in the selection of overseas trading venues for international firms. Indeed, a significant body of new research reexamines the international cross-listing decision of firms in the context of these and other corporate governance-related explanations (Karolyi (2006)). It is still unclear how patterns in the distribution of trading activity across global exchanges may be linked to these explanations. For example, is it possible for a firm to rationalize a decision to cross-list on an

[17] See the February 9, 2004 open letter to then SEC Chairman William Donaldson signed by the Alain Joly, President of the European Association for Listed Companies (EALIC) and several other business associations.

[18] See Termination of a Foreign Private Issuer's Registration of a Class of Securities Under Section 12(g) and Duty to File Reports Under Section 15(d) of the Securities Exchange Act of 1934 (December 23, 2005) and Condition No. 4 ("Level of U.S. Interest") for equity securities in which the average daily trading volume in the U.S. during a recent 12-month period must not be greater than 5% of the daily trading volume average in the primary trading market and U.S. residents do not hold more than 10% of the worldwide public float of equity securities.

overseas exchange in spite of a low information factor and poor prospects for any kind of U.S. market share of trading?

Another open question is whether the dealers and specialists that make markets in non-U.S. stocks on U.S. exchanges perform their function in a way that is likely to influence the proportion of trading activity in the U.S. Bacidore and Sofianos (2002) document with proprietary data that the inventory management behavior (positions closer to zero) and participation and stabilization rates (higher) of NYSE specialists in non-U.S. stocks are very different from those for U.S. stocks. More recent papers by Bacidore et al. (2005) and Moulton and Wei (2005) show that the specialist's market-making activities are significantly different across the trading day depending on when the home-market for the cross-listed shares is open and when it is closed. Our effort also makes only modest progress in understanding the joint dynamics of information factors and the global distribution of trading volume over time in general and around important events like the cross-listings themselves. For example, Halling et al. (2004) show that the foreign volume shares decline significantly in the first 3 years following a listing.

Finally, the problem of explaining differences in trading volume market shares exists in domestic market settings also, and it remains to be seen whether the intuition underlying our model has applicability in such a context. One paper that examines this issue is Baruch and Saar (2006), who employ a model similar to ours to explain why some firms list their shares on one market over another (Nasdaq vs. NYSE) and why some firms switch listings.

Appendix

Proof of Theorem 1: If all the liquidity traders submit the same fraction α of their order to the first exchange, then the covariance matrices R_1 and R_2 are given by the first and fourth equation in equation (8). Let $i \in \{1, 2\}$. Assume that the aggregate order of the informed traders submitted to exchange i is given by β_i, and that the liquidity traders use a linear strategy defined by α. The theory of linear filtering implies that (See Bensoussan, 1992, Theorem 1.1.1)

$$E[\upsilon_i \mid y_i] = \lambda_i y_i,$$

where

$$\lambda_i = F_i P \beta_i^t \left(R_i + \beta_i P \beta_i^t\right)^{-1},$$

which are the third and fifth equations in equation (8). Given a linear price rule $p_i = \lambda_i y_i$, the jth ($j \in \{1, 2\}$) informed trader's problem is

$$\max_{x_1 \in R^2, x_2 \in R^2} E\left[x_1^t(\upsilon_1 - \lambda_1 y_1) + x_2^t(\upsilon_2 - \lambda_2 y_2) \,\middle|\, s_j\right].$$

Given the linear strategy of the other traders (the other informed trader and the discretionary liquidity traders) and market segmentation, the problem can be written as

$$\max_{x_1 \in R^2} E\left[x_1^t(E[\upsilon_1 \mid s_j] - \lambda_1 y_1) \mid s_j\right] + \max_{x_2 \in R^2} E\left[x_2^t(E[\upsilon_2 \mid s_j] - \lambda_2 y_2) \mid s_j\right].$$

The first-order condition then implies that

$$x_1 = \left(\lambda_1 + \lambda_1^t\right)^{-1} E\left[\tilde{\upsilon}_1 \mid s_j\right]$$

and

$$x_2 = \left(\lambda_2 + \lambda_2^t\right)^{-1} E\left[\tilde{\upsilon}_2 \mid s_j\right].$$

The first-order condition is sufficient if the problem is concave, that is, if the Hessians of the objective functions are negative semidefinite matrices. The Hessians are $-(\lambda_1^t + \lambda_1)$ and $-(\lambda_2^t + \lambda_2)$. These are constraints the system (8) has to satisfy.

We also have $E[\upsilon_1 \mid s_1] = s_1$, $E[\upsilon_2 \mid s_1] = as_1$, $E[\upsilon_1 \mid \mathrm{s}_2] = 0$, and $E[\upsilon_2 \mid s_2] = bs_2$. Thus, we can write the aggregate informed demand submitted to the ith exchange as β_i, where β_i is given by

$$\beta_i = \left(\lambda_1 + \lambda_1^t\right)^{-1} F_i,$$

which explains the second and fourth equations in equation (8).

To end the proof, we need to derive the last equation in equation (8). Consider the problem of a discretionary liquidity trader who needs to buy an amount w. Given the linear structure of the price rule and traders' strategies, his problem can be reduced to

$$\max_{\alpha} E\left(as_1 + bs_2 - \lambda_1 \alpha w\right) \alpha w + \left(as_1 + bs_2 - \lambda_2(1-\alpha)w\right)(1-\alpha)w.$$

Because s_1 and s_2 have zero mean, we can write the problem as

$$\min_{\alpha} \left(\lambda_1 \alpha_w^2 + \lambda_2(1-\alpha_w)^2\right) w^2.$$

The optimal fraction is independent of the liquidity shock, and is given by

$$\alpha = \frac{\lambda_2}{\lambda_1 + \lambda_2},$$

which is the last equation in equation (8). This completes the proof. Q.E.D.

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