PRICE DISCOVERY FOR SEGMENTED U.S. –LISTED CHINESE STOCKS: LOCATION OR MARKET QUALITY?

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Abstract: This study extends the cross-listing literature by examining how, and to what extent, the trading of cross-listed China-backed ADRs on the New York Stock Exchange contributes to information flows and price discovery for the corresponding stocks traded in China's A-share market. We find that the cross-listed U.S. prices and Chinese prices are not cointegrated in the long-run and the home market plays a far more important role in both price discovery and volatility spillover than does the U.S. market. The home bias hypothesis still holds for the segmented Chinese A-share market and the location where price discovery actually originates is the essential factor in the process of international information transmission.

INTRODUCTION

Cross-border listing has increasingly gained popularity in recent years. As a growing number of companies pursue to cross-list stocks on a foreign stock exchange, the New York Stock Exchange (NYSE) is often a priority choice. As a result, the proportion of non-U.S. listings on the NYSE doubled from 8.5% in 1994 to 17% in 2003 (Chouinard and D'Souza, 2004). While it has been well documented in the literature that cross-listing in the U.S. market contributes benefits, such as better access to capital and product markets, greater liquidity in trading shares, prestige, transparent corporate governance, etc., to the foreign-based companies, an issue that remains controversial is what contributes to price discovery of the listing companies.

Recently, Karolyi (2006), expanding on his early paper (1998), surveys, synthesizes, and critically reviews the most recent literature in international cross-listings, and then identifies several yet unresolved questions for future research. According to Karolyi, one of the questions that remain open is to uncover where the markets are deepest and where price discovery actually originates, in light of the growing breadth of analysis across markets around the world. Therefore, it would be interesting to analyze how the U.S. contribution to price discovery varies across countries.

So far, the empirical evidence regarding where price discovery occurs has been mixed and conflicting. Some studies suggest that the home market continues to play a dominant role in price discovery (e.g., Lieberman, Ben-Zion, and Hauser, 1999; Kim, Szakmary, and Mathur, 2000; Menkveld, Koopman, and Lucas, 2003; Grammig, Melvin, and Schlag, 2004; von Furstenberg and Tabora, 2004; Phylaktis and Korczak, 2005; and Pascual, Pascual-Fuster, and Climent, 2006). Yet, the evidence from other studies lends some support for a significant role in price formation for both at home and the U.S. market (e.g., Lau and Diltz, 1994; Werner and Kleidon, 1996; and Eun and Sabherwal, 2003).

The purpose of this paper is to extend the literature by examining how, and to what extent, the trading of the cross-listed Chinese stocks in the U.S. market contributes to information flows and price discovery for the corresponding stocks traded in China as of March 2007. The present study provides a differentiation from other studies that examine price discovery for cross-listings from developed countries in the U.S., but also from studies that examine U.S. cross-listings from emerging markets with no capital control restrictions. Currently, China has two stock exchanges: the Shanghai Stock Exchange (SHSE) and the Shenzhen Stock Exchange (SZSE), with two types of shares traded on both exchanges. A-shares are denominated in Chinese RMB and can be purchased by domestic investors, whereas B-shares are denominated in either US dollars (SHSE) or HK dollars (SZSE) and could be purchased by foreign investors prior to 2001. Another type of China-backed stocks, H-shares, is listed on the Stock Exchange of Hong Kong (SEHK).

Historically, the A-share market in China was closed for foreign investors and the B-share market for domestic investors. In February 2001, the B-share market was first opened to individual domestic investors with foreign currency holdings. In November 2002, the Qualified Foreign Institutional Investors (QFII) measures were promulgated to allow qualified foreign institutions to invest between USD50 million and USD800 million in the A-share market under certain foreign exchange flow and disclosure requirements. However, the impact of the QFII transactions is quite limited, which only accounted for approximately 5% of the market as of December 2005, due to many undue restrictions imposed on the QFIIs.[1] As a consequence, the Chinese A-share market is still a segmented market that is not easily accessible to most foreign investors.

[1]The restrictions include that: 1) one QFII cannot own more than 10% of shares of one company; 2) one company cannot have more than 20% of its shares owned by QFIIs; 3) most QFIIs can apply for withdrawal after 1 year, at which time maximum withdrawal is 20% at intervals greater than 1 month. Closed-end funds cannot withdraw funds for a 3-year time period; and 4) a QFII must have invested for at least 3 months before they can transfer their investment quota to another QFII or qualifying applicants.

As of today, two readily available investment avenues for U.S. investors to invest in China are through either mutual funds or China-backed stocks listed in the U.S. as American Depository Receipts (ADRs). All China-backed ADRs represent multiples of underlying China-backed equity shares listed on either the SEHK or the SHSE. The ADRs use the U.S. GAAP for reporting purpose, whereas their corresponding Chinese shares adopt the Chinese GAAP. So far, there have been numerous studies on the information transmission between markets for domestic and foreign investors in China, but they all focus on price discount and market segmentation of the Chinese A-shares, Chinese B-shares, or Hong Kong H-shares (see Baily, 1994; Chakravarty, Sakar, and Wu, 1998; Chui and Kwok, 1998; Xu and Liu, 2001; Yang, 2003 and Chan et al, 2007). The only exception is Xu and Fung (2002) who examine the pattern of information flows for China-backed stocks that are cross-listed on exchanges in Hong Kong and New York. Their results indicate that stocks in the domestic market (Hong Kong) appear to play a more significant role of information transmission in the pricing process, whereas stocks listed in the offshore market (NYSE) play a bigger role in volatility spillover.

However, the "domestic" market specified in Xu and Fung (2002) is at best a "pseudo home" market because all HK-listed Chinese companies are headquartered in China. This paper differs from Xu and Fung (2002) in that they study the information flows between the "pseudo home" market (Hong Kong) and the offshore market (NYSE), whereas we focus on the information flows between the "real home" market (i.e., China's A-share market) and the offshore market (NYSE). In addition, having the A-shares traded in the "real home" market adds another link in the price discovery process, which can potentially allow us to extend our analysis by examining how the introduction of A-shares affects the price discovery process in Hong Kong and the U.S.

Another issue of interest is, as Karolyi (2006) points out, which is the first-order factor in price discovery, location or market quality? Owing to the fact that our sample firms are cross-listed on all three exchanges (NYSE, SEHK, and SHSE), our research design allows us to distinguish home market (China) from pseudo home market (Hong Kong) and location from market quality as the essential factor in price discovery. In terms of market quality, the U.S. market is perceived to be superior to the Chinese A-share market in information transmission because of its efficient market microstructure and increased disclosure requirements imposed by the Sarbanes-Oxley Act. Regarding location, the Chinese A-share market is much closer than Hong Kong to corporate headquarters where information is originated. If institutional differences play an important role in the price discovery process, then we can observe the effect of superior market quality versus the effect of location. If superior market quality from New York is more dominant than location (the home bias), then the information transmission from U.S. to China is expected to be highly significant and with higher quality. If location dominates the market quality effect, then the home market should play a more important role than the pseudo home market.

Another contribution of this paper is that we employ a new causality-in-variance test developed by Hong (2001) in the cross-listing literature. So far, most empirical studies on volatility spillover have used either a multivariate GARCH model that simultaneously fits two or more time series or Granger's (1969) causality test that fits each time series independently to a model, with the causality testing done by regressing the squared residuals of one series on the other. One apparent disadvantage to both approaches is the potential interaction between the time series and the corresponding uncertainty surrounding both first- and second-moment dynamics. The other pitfall is that they also suffer from the uncertainty surrounding the asymptotic distribution of the maximum likelihood estimator for a multivariate GARCH process.[2] Kroner and Ng (1998) further suggest that the choice of multivariate volatility model can substantially affect the conclusion of an analysis. In contrast to the multivariate GARCH model, Hong's (2001) causality-in-variance test, however, makes no distributional assumptions, does not rely on simultaneously modeling intra-series and inter-series dynamics, and provides a robust test of Granger's causality relation. In addition, Hong's (2001) test can be applied to the case when relatively long lags are expected in the causation pattern. In this paper, we will present both results based on a multivariate GARCH model and Hong's causality-in-variance test.

The main findings of this paper can be summarized as follows. First, ADR prices and China's A-share prices are not cointegrated in the long-run. Our results discover that China's A-share market was still segmented as of March 2007. Second, while Hong's causality-in-mean test results indicate that the U.S. market may exert a temporal (one day) effect on China's A-share stock returns, the traditional Granger's causality test results show a significant price discovery originating from China's A-share market to the U.S. market. Third, in term of causality in volatility, both multivariate GARCH and Hong's test results exhibit some volatility spillovers from China's A-share market to the U.S. market. Therefore, the home bias hypothesis still holds for the segmented Chinese A-share market, and the location where price discovery actually originates is the essential factor in the process of international information transmission. The price discovery process between the U.S. market and the pseudo home market (Hong Kong) in China also supports our findings.

The remainder of the paper is organized as follows. Section 2 reviews the relevant cross-listing literature. Section 3 describes the sample of China-backed companies tri-listed on the NYSE, SEHK, and SHSE. Section 4 outlines the methodology and Section 5 presents the empirical results. Finally, Section 6 provides concluding remarks.

LITERATURE REVIEW

[2] For a detailed analysis, refer to Engle and Kroner (1995).

In the literature of international cross-listings, many studies support the home bias hypothesis that the domestic market plays a dominant role in price discovery. For example, Lieberman et al. (1999), who study six Israeli firms cross-listed in the U.S., find that price discovery mostly occurs in Israel. Grammig et al. (2004) study three U.S.-listed German stocks and find support for the dominance of the home market, i.e., German prices dominate both U.S prices and exchange rate effects in price discovery. von Furstenberg and Tabora (2004) compare the durability of price innovations in two major Mexican stocks traded in the U.S., and they find that the Mexican market has some advantages in information efficiency, even though it is economically inefficient. Their result casts doubt on the common assertion that price discovery for some of the major emerging markets is predominantly done in New York.

However, there are some studies supporting the global center hypothesis. For instance, Kim et al. (2000) study the price transmission dynamics between ADRs and their underlying foreign shares. Their results indicate a small role for U.S. price discovery in the cross-listed firms from Japan, Netherlands, U.K., Sweden, and Australia. Menkveld et al. (2003) examine one year of transactions data on seven Dutch firms and uncover important price and quote activities around the NYSE opening for these stocks. Phylaktis and Korczak (2005), examining 95 U.S.-listed British and French companies, show that the extent of U.S. price discovery is positively related to the liquidity of U.S. trading. Finally, Pascual et al. (2006) investigate six Spanish stocks using two-hour overlapping periods and find that the NYSE has a high contribution to the price discovery process.

Yet, the evidence from other studies lends some support for a significant role in price formation for both at home and the U.S. market. Lau and Diltz (1994) find a two-way causality between the Japanese prices and the prices of seven Japanese firms cross-listed in the U.S. Werner and Kleidon (1996) analyze intraday patterns for U.K. and U.S. trading of British cross-listed stocks. They discover that the two-hour overlapping period is characterized by concentrated trading as private information, originating from New York, gets incorporated in both markets. Recently, Eun and Sabherwal (2003) who study Canada and U.S. trading also find support for significant price discovery in both markets.

In addition, it is a common belief that price discovery will occur on a foreign exchange when its market quality is superior to that of the domestic exchange. Coffee (2002) first demonstrates that the market with higher disclosure requirements and a greater standard of enforcement increases the reliability of information. Doidge, Karolyi, and Stulz (2004) contend that disclosure requirements for both trading and accounting information, as well as the degree of protection of minority shareholders' interest, all have implications for the valuation of the firm. Lang, Lins, and Miller (2003) show that firms cross-listed in the U.S. have increased forecast accuracy relative to firms that are not cross-listed.

So far, abundant studies have examined the segmentation and integration of China's A- and B-share markets. Baily (1994) first provides some preliminary evidence on the Chinese B-share market. He finds that discounts at which B-shares trade relative to A-shares available to Chinese citizens are correlated across firms and related to similar premiums in other Asian markets, and B-shares have considerable diversification value but are not entirely segmented from global financial conditions. Chakravarty et al. (1998) develop a model, which incorporates both information asymmetry and market segmentation, to explain why Chinese B-shares trade at an average discount of about 60% to the prices of A-shares. They argue that one reason for the large price discount of B-shares is because foreign investors have less information on Chinese stocks than domestic investors. Chui and Kwok (1998) demonstrate that returns of B-shares are correlated with those of A-shares and that this correlation depends on the information transmission mechanism at work. They argue that because foreign investors may receive news about China faster than domestic Chinese investors due to information barriers currently existing in China, returns on B-shares lead the returns on A-shares. Yang (2003) examines the market segmentation and information asymmetry patterns in Chinese stock markets. He shows that foreign investors in the Shanghai B-share market are better informed than Chinese domestic investors in the A-share market and foreign investors in both Shenzhen and Hong Kong markets. Chan et al. (2007) study 76 firms that issue both A-shares and B-shares and compare the price discovery role of the two segmented markets in China. Before 2001, the A-share market led the B-share market in price discovery, but after 2001 when domestic investors were allowed to invest in the B-share market, a reverse causality occurred from the B-share to the A-share market.

The most relevant paper related to our study is Xu and Fung (2002), who examine the information flows between China-backed stocks dual-listed on the SEHK and NYSE. Based on a sample of 10 ADRs, they use a bivariate GARCH model to investigate the information flow related to both pricing (first moment) and volatility (second moment) spillover across markets. They discover that Hong Kong appears to play a more important role in influencing the pricing of corresponding companies in New York, whereas the same stocks listed in New York are more influential in terms of volatility spillover across markets. They conclude that the pricing results of information flows are consistent with the home bias hypothesis and the volatility spillover results support the global center hypothesis.

DATA

As of March 2007, there were 18 China-backed ADRs listed on the NYSE, all of which were dually-listed in Hong Kong as H-shares, with only nine cross-listed on the SHSE as A-shares. Table 1 presents the nine Chinese companies cross-listed on all three exchanges (NYSE, SEHK, and SHSE), their ticker symbol and listing date on each exchange, and the conversion ratio of each ADR to the underlying local shares. As shown, two companies, China Life Insurance (LFC) and Guangshen Railway (GSH), were listed on the SHSE just three to four months prior to March 2007. Due to a lack of sufficient price data, both companies are not included in the study on the price dynamics between SHSE and NYSE. The remaining seven ADRs are China Petroleum & Chemical Corporation (SNP), an oil and natural gas company; China Unicom (CHU), a wireless communication firm; China

Eastern Airline (CEA) and China Southern Airline (ZNH), two leading airline companies; Huaneng Power International (HNP), an electric power firm; Shanghai Petrochemical (SHI) and Yanzhou Coal Mining (YZC). Evolving from former state-owned enterprises, all seven Chinese companies are currently the leading company in their respective industry, and they do not have B-shares trading in China.

Table 1. China-backed stocks cross-listed on the NYSE, SEHK, and SHSE as of March 31, 2007

Company Name	**NYSE Ticker**	**Sector**	**NYSE Listing Date**[a]	**SHSE Listing Date**[b]	**SEHK Listing Date**[c]	**SHSE Ticker**	**SEHK Ticker**
China Eastern Airlines	CEA	Travel & Leisure	Feb. 04, 1997	Nov. 5, 1997	Feb. 5, 1997	600115	670
China Life Insurance	LFC	Life Insurance	Dec. 17, 2003	Jan. 9, 2007	Dec. 18, 2003	601628	2628
China Petroleum & Chemical	SNP	Oil & Gas Producers	Oct. 18, 2000	Aug. 8, 2001	Oct 19, 2000	600028	386
China Southern Airlines	ZNH	Travel & Leisure	July 30, 1997	July 26, 2003	July 31, 1997	600029	1055
China Unicom	CHU	Mobile Telecom.	June 21, 2000	Oct. 9, 2002	June 22, 2000	600028	762
Guangshen Railway	GSH	Travel & Leisure	May 13, 1996	Dec. 22, 2006	May 14, 1996	601333	525
Huaneng Power International	HNP	Electricity	Oct. 06, 1994	Dec. 6, 2001	Jan. 21, 1998	600011	902
Sinopec Shanghai Petrochemical	SHI	Chemicals	July 26, 1993	Nov. 8, 1993	July 26, 1993	600688	338
Yanzhou Coal Mining	YZC	Mining	Mar. 31, 1998	July 1, 1998	April 1, 1998	600188	1171

[a] Data source: www.nyse.com
[b] Data source: www.yahoo.com
[c] Data source: http://www.hkex.com.hk
[d] The conversion ratio indicates the number of underlying local shares that corresponds to one ADR share in New York.

For each sample firm, we collect daily closing prices adjusted for dividends on three exchanges from the Datastream. To avoid potential IPO pricing and liquidity effects, our sample period covers from one year after the latest listing date to March 31, 2007[3]. For example, the latest IPO date of China Petroleum & Chemical (SNP) occurred on August 8, 2001 in China's A-share market, so its price data include 1,210 observations from August 8, 2002 to March 31, 2007. In addition, both A-share and H-share prices are converted into US-dollar equivalents using the prevailing exchange rates, and then continuously compounded daily returns are calculated.

Table 2 provides the descriptive statistics of daily returns on both NYSE and SHSE for all seven cross-listed Chinese stocks. As shown, ZNH has the fewest return observations with 698 and SHI has the most with 3,233. The majority of ADRs exhibit higher mean returns and standard deviations than their SHSE-listed counterparts. Both skewness and kurtosis indicate that the distribution of returns is not normal for all stocks because of the existence of fat-tails.

[Table 2 is available upon request]

EMPIRICAL FRAMEWORK

This section provides a detailed discussion of the methodologies employed to analyze both long-run and short-run mean and volatility dynamics of information transmission.

Recursive Cointegration Tests

A large body of literature has investigated the price discovery process based on stock return data (e.g., Fung, Lee and Leung, 2000; and Chan et al., 2007). Because stock returns follow a stationary process, an information transmission analysis that is based on stock returns implicitly rules out any possibility of price cointegration in the long run, which eventually leads to either model misspecification or incorrect conclusion for the price discovery dynamics. To address this econometric issue, we first investigate whether a long-run relationship exists between cross-listed prices by estimating a cointegration/error correction model. Let

[3] For HNP, we start from 12/13/2002 because the first observation of HNP's A-share prices in Datastream is 12/13/2001.

X_t denote a log price vector for both U.S. and Chinese stocks. Without loss of generality, we can write the system as an error correction model (ECM) with *k*-1 lags (which is derived from a level VAR with *k* lags) as follows:

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \varepsilon_t \quad (t = 1,....,T) \tag{1}$$

where $X_t = \begin{pmatrix} X_{1t} \\ X_{2t} \end{pmatrix}$ with X_{1t} and X_{2t} representing the log stock price in the China and U.S., respectively; $\Gamma_i = \begin{pmatrix} \Gamma_{11} & \Gamma_{12} \\ \Gamma_{21} & \Gamma_{22} \end{pmatrix}$ are coefficients ; μ is a (2×1) vector of constants; ε_t is a (2×1) vector of white noise disturbances; Π is a (2×2) matrix of parameters that contains information about the long-run relations between $X_t{}'s$, whose rank (*r*) determines the number of independent cointegrating vectors; and *T* is the number of observations.

Two cases are of interest for our study. First, if *r* = 0, then Π contains no long-run information. Next, we apply the least-square method to a VAR in the first difference of (log) price series. The optimal number of lags (*k*) can be determined by the Akaike information criterion (AIC) based on the whole sample period. Second, if *r* = 1, then there exist matrices $a = \begin{pmatrix} a_1 \\ a_2 \end{pmatrix}$ and $\beta = \begin{pmatrix} \beta_1 \\ \beta_2 \end{pmatrix}$ such that $\Pi = a\beta'$, where a represents the speed of the adjustment to the equation and β is the matrix of the long-run coefficients. In this case, $\beta' X_t$ is stationary that indicates a long-run equilibrium relationship between the log prices, even though X_t is not stationary. An efficient estimation of Equation (1) requires the use of the Johansen procedure (1991).

Before estimating Equation (1), we need to make an adjustment of the data due to the time zone difference between the two markets.[4] Since the U.S market always closes before the Chinese market opens on the next day, its closing value cannot affect the Chinese market on the *same* calendar day. Therefore, it is inappropriate to match the daily closing prices on the NYSE with the closing prices on the SHSE on the same day. Instead, we use the one-day lagged returns on the NYSE in line with the daily returns on the SHSE in Equation (1).

To check the robustness of Johansen's cointegration test results, the recursive cointegration analysis developed by Hansen and Johansen (1999) is applied to examine the stability of the cointegration (or non-cointegration) relationship. The recursive cointegration analysis can be implemented under two VAR representations of Equation (1). Under Z-representation, all parameters in the error correction model are re-estimated during recursive estimations, whereas under R-representation, short-run parameters Γ_i are fixed to their full sample values and only long-run parameters in Equation (1) are re-estimated. Following Hansen and Johansen (1999), we conduct parameter constancy tests based on both representations.

Short Run Dynamics: ECM-GARCH Model

In this section, we focus on the short-term dynamic linkage between the home and offshore markets. If the cointegration relationship is confirmed by the recursive cointegration analysis, we can examine short-run information transmission through the return linkage between the two markets by testing the null hypothesis that there is no Granger causality from X_{2t} (the Chinese stock price) to X_{1t} (the U.S. stock price), which is equivalent to jointly testing the following coefficient restrictions in Equation (1):

$$a_1 = (\Gamma_{12})_{-1} = (\Gamma_{12})_{-2} = ... = (\Gamma_{12})_{-k} = 0 \tag{2}$$

If the number of cointegrating vectors is zero which means no long-run relationship, we then rely on the standard Wald test in first-difference VAR to show in-sample evidence of a causal relationship in returns between the two markets. To address the potential bias generated by the unstable VAR, we conduct both recursive- and rolling-modeling approaches suggested by Pesaran and Timmermann (1995), which simultaneously consider model uncertainty and the possible time-varying pattern of information transformation.

Next, we examine the short-run volatility linkage by using the following multivariate GARCH model:

$$\Delta X_t = a\beta' X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \varepsilon_t \quad (t = 1,...,T) \qquad \varepsilon_t \mid \Omega_{t-1} \sim D(0, H_t) \tag{3}$$

where Ω_{t-1} is the conditioning information set at time *t*-1 and $H_t = \begin{pmatrix} h_{11} & h_{12} \\ h_{21} & h_{22} \end{pmatrix}$, which denotes the conditional covariance matrix. The information transmission through the volatility linkage can be examined by estimating the conditional covariance matrix H_t in Equation (3).

There exist many parameterizations of the conditional covariance matrix H_t in the multivariate GARCH model. We us Engle and Kroner's (1995) BEKK specification as our GARCH model specification, which is sufficiently general and guarantees a positive definite conditional covariance matrix. Following Engle and Kroner (1995), a parsimonious bivariate GARCH (1,1) BEKK specification is adopted below:

$$H_t = \begin{bmatrix} h_{11,t} & h_{12,t} \\ h_{21,t} & h_{22,t} \end{bmatrix} = \begin{bmatrix} c_{11} & c_{12} \\ 0 & c_{22} \end{bmatrix}' \begin{bmatrix} c_{11} & c_{12} \\ 0 & c_{22} \end{bmatrix} + \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{bmatrix}' \begin{bmatrix} \varepsilon_{1,t-1}^2 & \varepsilon_{1,t-1}\varepsilon_{2,t-1} \\ \varepsilon_{2,t-1}\varepsilon_{1,t-1} & \varepsilon_{2,t-1}^2 \end{bmatrix} \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{bmatrix} + \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix}' \begin{bmatrix} h_{11,t-1} & h_{12,t-1} \\ h_{21,t-1} & h_{22,t-1} \end{bmatrix} \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix} \tag{4}$$

[4] The core trading sessions on the NYSE are from 9:30 a.m. to 4:00 p.m. The trading hours for both the SHSE and SZSE are from 9:30 a.m. to 11:30 a.m. and from 1:00 p.m. to 3:00 p.m. (the New York equivalent hours are from 9:30 p.m. to 11:30 p.m. of the previous day and from 1:00 a.m. to 3:00 a.m.).

Equation (4) is estimated simultaneously by using the maximum likelihood estimation procedure, where the off-diagonal parameter a_{12} measures the transmission of information originated from the offshore (U.S.) market in the previous period to current period's conditional volatility in the home (China) market, and the off-diagonal parameter b_{12} measures the dependence of the conditional volatility of the home market on that of the offshore market in the previous period. Similar interpretations can be applied to a_{21} and b_{21} as well.

Robust Tests of Causal Relations in Return and Variance between the Two Markets

Because of some potential interactions between the time series and the corresponding uncertainty surrounding both first- and second-moment dynamics, the formulation of a multivariate GARCH model can be very complicated. This makes the task of correctly specifying an adequate multivariate model very challenging. Cheung and Ng (1996) develop a two-stage test called the *S* statistic to solve this problem. While the *S* statistic improves the test efficiency, it inherently possesses one major weakness, namely that it is constructed by uniformly weighting each lag, making no distinction between recent and distant cross-correlations. Recently, Hong (2001) addresses this problem by modifying the *S* statistic with a weighting function that can accommodate a number of weighting kernels. The single-direction (or "one-sided") causality statistic, *Q*, is defined as follows:

$$Q_1 = \left\{ T \sum_{j=1}^{T-1} k^2(j/M)\hat{\rho}^2_{uv}(j) - C_{1T}(k) \right\} / \left[2D_{1T}(k) \right]^{1/2} \qquad (5)$$

where $C_{1T}(k) = \sum_{j=1}^{T-1} (1 - j/T) k^2(j/M)$;

$D_{1T}(k) = \sum_{j=1}^{T-1} (1 - j/T)\{1 - (j+1)/T\} k^4(j/M)$; $k(j/M) = sin(\pi \cdot j/M)/(\pi \cdot j/M), \; -\infty < j/M < +\infty$;

$\hat{\rho}_{\mu\nu}(j) = \{\hat{C}_{\mu\mu}(0)\hat{C}_{\nu\nu}(0)\}^{-1/2} \hat{C}_{\mu\nu}(j)$;

$\hat{C}_{\mu\mu}(0) = T^{-1} \sum_{t=1}^{T} \hat{\mu}_t^2$ and $\hat{C}_{\nu\nu}(0) = T^{-1} \sum_{t=1}^{T} \hat{\nu}_t^2$;

$\hat{C}_{\mu\nu}(j) = \begin{cases} T^{-1} \sum_{t=j+1}^{T} \hat{\mu}_t \hat{\nu}_{t-j}, & j \geq 0 \\ T^{-1} \sum_{t=j+1}^{T} \hat{\mu}_{t+j} \hat{\nu}_t, & j < 0 \end{cases}$; and

$\hat{\mu}_t = \hat{\varepsilon}^2_{1t}/\hat{h}_{1t} - mean(\hat{\varepsilon}^2_{1t}/\hat{h}_{1t})$ and $\hat{\nu}_t = \hat{\varepsilon}^2_{2t}/\hat{h}_{2t} - mean(\hat{\varepsilon}^2_{2t}/\hat{h}_{2t})$.[5]

The above *Q* statistic is robust to any distributional assumption and is asymptotically distributed *N*(0,1), using the single-tailed normal distribution.[6] Moreover, like the *S* statistic, the *Q* statistic can be extended to test causality in mean by substituting standardized residuals $\hat{\varepsilon}_{1t}/\hat{h}_{1t}^{1/2}$ and $\hat{\varepsilon}_{2t}/\hat{h}_{2t}^{1/2}$ into the sample cross-correlation function.

EMPIRICAL RESULTS

Are the U.S and Chinese Price Series Cointegrated in the Long Run?

Before examining whether the cointegration between the prices of seven sample stocks in China and the United States exists, we use the augmented Dickey-Fuller (1981) test, which includes lagged first differences of the log-price series in the equation, to detect the presence of a unit root. The augmented Dickey-Fuller results indicate that no time trends are present for all log stock price series, and that the first differences of the log prices yield no additional unit roots.[7] After confirming that both time series are integrated of order one, denoted as *I*(1), we use the multivariate extension of the AIC to identify the optimal lag length for the cointegration test. The AIC turns out to be minimized at a lag length of 2 or 4 for most stocks, except for CHU that has a system lag of 5.

Next, the long-run cointegration relation between the two price series is examined by using Johansen's (1991) cointegration analysis. Both maximum-eigenvalue and trace tests are conducted to determine the number of cointegration vectors if the prices are indeed cointegrated. The null hypothesis for both maximum-eigenvalue and trace tests is that there is zero cointegrating vector ($r = 0$), which can be interpreted as no long-run relationship between the two price series. Table 3 reports Johansen's (1991) cointegration test results with maximum eigenvalues and trace statistics. As shown, we fail to reject the null hypothesis that there is zero cointegrating vector at the 5% level for all but one (CHU) sample stocks under both maximum-eigenvalue and trace tests. Since Johansen's test statistics may suffer from a finite-sample bias, we also use Reimers' (1992) small sample corrected formulas as a robustness check. The results, however, remain the same.

Table 3. Johanson's cointegration test results: ADR prices and Chinese A-share prices

[5] Hong (2001) subtracts unity from the squared standardized errors when computing sample cross-correlations and finds no difference in statistic size or power.

[6] The *Z* score for the 10%, 5%, and 1% significance levels is 1.282, 1.645, and 2.326, respectively.

[7] The augmented Dickey-Fuller test results are available upon request.

The null hypothesis is zero cointegrating vector. *r* refers to the number of cointegrating vectors. The optimal lag order of VAR is chosen by the criterion of Akaike information criterion (AIC). The corresponding 95 percent critical values for maximum eigenvalue and trace statistics are 15.89 and 20.26, respectively. * denote statistical significance at the 5 percent.

	Maximum Eigenvalue $H_0: r = 0$ $H_1: r = 1$	Trace Statistic $H_0: r = 0$ $H_1: r = 1$	Lag
CEA	5.66	8.63	4
SNP	14.13	20.04	2
ZNH	9.29	9.89	2
CHU	16.44*	17.88	5
HNP	9.25	12.20	2
SHI	12.01	14.49	2
YZC	8.61	10.21	4

Although the concept of cointegration is a very useful tool through which the relationship between non-stationary variables can be analyzed, the power of Johansen's (1991) cointegration test, however, is relatively low, especially for a small sample (Haug, 1996). To check the robustness of the above results, we also conduct the recursive cointegration analysis, which provides the point estimate of stability (or lack of it) for the cointegration relationship and, therefore, is more informative than the standard cointegration test.

Figure 1 presents statistics that are normalized by the 10% critical values given in Osterwald-Lenum (1992), wherein any statistic greater than 1.0 signals the rejection of the null hypothesis at the 10% level. There are two graphs for each stock: Z-representation (top) and R-representation (bottom). Figure 1 reveals that the majority of trace statistics are below the critical value of 1.0, except for a few outliers, indicating that the null hypothesis of no long-run relationship cannot be rejected for our sample stocks.

[Figure 1 is available upon request]

Overall, the above results confirm the standard cointegration test, but they are contrary to the common belief that there exists a long-run relationship for cross-listed stocks as evidenced in current literature. Although cross-listed stocks are traded in different national stock markets, they inherently share the same underlying asset and are likely to be affected by the same fundamental information set. Hence, it is natural to expect a long-run and stable relationship between the cross listings. However, our results clearly indicate that the prices of seven Chinese A-shares and their NYSE-listed ADRs are not cointegrated, suggesting that the Chinese A-share (real home) market is segmented from the U.S. (offshore) market. This evidence is new to the cross-listing literature.

Short-Run Information Transmission

Given no evidence of a long-run cointegration relationship between the two price series on NYSE and SHSE, we next focus on short-run information transmission by using stock return data. First, causality in return is examined using the standard Wald statistics calculated from the recursive estimation method as well as the rolling estimation method. Table 4 summarizes the results. Block I reports the proportion of times that the null hypothesis of no causality from China to U.S. is rejected at the 5% level, whereas Block II reports the results for the null hypothesis of no causality from U.S. to China. The Wald rejection rates shown in Panel A are based on recursive n-60, n-120, n-240, n-360, and n-480 estimations, respectively, where n is the number of observations. For example, in CEA's n-60 case, we start with 60 observations, then add one observation and test the null hypothesis. Next, we add two observations and test the null hypothesis, and so on. The Wald rejection rate of 0.91 means that we test the null hypothesis 2,132 (=2,192-60) times and we reject the null hypothesis 2,068 times. As shown in Panel A, five out of seven stocks show high rejection rates of the null hypothesis that there is no causality in return from China to U.S. On the contrary, only three stocks (CEA, SNP, and CHU) reject the null hypothesis that there is no causality from U.S. to China. Similarly, Panel A also provides moderate evidence showing that a bi-directional causal relationship exists for only two ADRs, CEA and CHU.

Table 4. Granger causality-in-return via Wald tests: ADR prices and Chinese A-share prices

The table reports pairwise Granger causality/block exogeneity Wald tests. Lags of VARs are chosen by the criterion of Akaike information criterion (AIC). Numbers in Blocks I and II are Wald test rejection frequencies that reject the null hypothesis at the 5% level. Panel A reports Wald statistics based on recursive n-60, n-120, n-240, n-360 and n-480 estimations, where n is the number of observations. Panel B reports Wald statistics based on the rolling estimation with observations of 60, 120, 240, 360 and 480.

	Block I Pairwise Granger Causality/Block Exogeneity Wald Tests *Ho:* No causality from China to U.S.					Block II Pairwise Granger Causality/Block Exogeneity Wald Tests *Ho:* No causality from U.S. to China				
Panel A: Recursive estimations										
Number of observations	60	120	240	360	480	60	120	240	360	480
CEA	0.91	0.93	0.99	1.00	1.00	0.62	0.63	0.63	0.65	0.70
SNP	0.00	0.00	0.00	0.00	0.00	0.59	0.63	0.70	0.73	0.69
ZNH	0.97	1.00	1.00	1.00	1.00	0.00	0.00	0.00	0.00	0.00
CHU	0.62	0.65	0.76	0.93	1.00	0.48	0.44	0.49	0.59	0.56
HNP	0.80	0.82	0.89	0.96	1.00	0.16	0.16	0.19	0.22	0.26
SHI	0.35	0.34	0.36	0.37	0.39	0.00	0.00	0.00	0.00	0.00
YZC	0.43	0.45	0.46	0.49	0.53	0.02	0.02	0.02	0.02	0.02
Panel B: Rolling estimations										
Number of observations	60	120	240	360	480	60	120	240	360	480
CEA	0.08	0.18	0.21	0.28	0.40	0.07	0.07	0.09	0.05	0.08
SNP	0.11	0.07	0.04	0.00	0.05	0.04	0.07	0.31	0.28	0.35
ZNH	0.45	0.71	1.00	1.00	1.00	0.00	0.00	0.00	0.00	0.00
CHU	0.20	0.44	0.71	0.84	0.96	0.07	0.18	0.12	0.11	0.14
HNP	0.27	0.33	0.64	0.76	0.96	0.02	0.08	0.04	0.05	0.00
SHI	0.20	0.26	0.30	0.31	0.34	0.05	0.08	0.11	0.13	0.06
YZC	0.19	0.25	0.37	0.42	0.56	0.05	0.04	0.16	0.33	0.44

Table 5. BEKK model estimation results: ADR prices and Chinese A-share prices

	CEA		SNP		ZNH		CHU		HNP		SHI		YZC	
Parameters	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*
c_{11}	0.01**	0.00	0.02	0.01	0.01**	0.00	-0.01	0.00	0.00	0.00	0.00	0.00	0.00	0.00
c_{12}	0.00**	0.00	0.01**	0.00	0.01**	0.00	0.01**	0.00	0.00**	0.00	0.00	0.00	0.00**	0.00
c_{22}	0.00	0.00	0.00	0.01	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
a_{11}	0.28**	0.02	0.32**	0.04	0.24**	0.06	-0.32**	0.03	0.29**	0.03	0.00	0.10	0.26**	0.02
a_{12}	0.01	0.05	0.00	0.42	-0.06	0.09	0.00	0.16	0.03	0.14	0.03	0.02	0.03	0.03
a_{21}	-0.03	0.03	0.05	0.05	0.24**	0.05	-0.11**	0.04	0.01	0.03	-0.08	0.12	0.20**	0.02
a_{22}	0.00	0.07	-0.01	0.90	-0.02	0.08	0.03	0.19	0.03	0.15	-0.34**	0.03	-0.08	0.05
b_{11}	0.95**	0.01	0.90**	0.03	-0.85**	0.04	-0.89**	0.02	0.94**	0.01	0.00	0.02	0.94**	0.01
b_{12}	-0.03	0.06	0.01	0.44	-0.02	0.09	0.00	0.18	0.03	0.15	0.07**	0.01	0.02	0.03
b_{21}	-0.10**	0.03	0.04	0.03	0.12	0.09	0.06	0.04	-0.03	0.03	0.03	0.12	-0.01	0.06
b_{22}	0.00	0.02	0.00	0.03	0.00	0.05	0.00	0.03	0.00	0.03	-0.86**	0.03	0.00	0.03

** and * denote statistical significance at the 1 percent and 5 percent, respectively.
s.e. denotes standard error.

Panel B of Table 4 reports the Wald rejection rates based on the rolling estimation of the models with 60, 120, 240, 360, and 480 observations, respectively, which provide additional information about the short-run stock returns dynamics. In Block I, four out of seven ADRs show high rejection rates of the null hypothesis that there is no causality from China to U.S. However, in Block II, all seven ADRs fail to reject the null hypothesis that there is no causality from U.S. to China. Therefore, Panel B demonstrates strong uni-directional causality from China to U.S. In summary, the evidence from both recursive- and rolling-estimation tests implies that a causal relationship exists between ADRs and China's A-shares, and the effects of China's A-shares on their ADRs are more prominent than vice versa.

Table 5 presents the pattern of information transmission through volatility. As discussed in Section 4, we are interested in estimating the off-diagonal parameters a_{12}, a_{21}, b_{12}, and b_{21} in Equation (4). As shown, for four out of seven stocks (except SNP, HNP, and SHI), either a_{21} or b_{21} is statistically significant, and all but one (SHI) estimates for a_{12} and b_{12} are statistically insignificant. These results reveal that there exists a strong one-way volatility transmission from China to U.S. for YZC, ZNH, CEA, and CHU, but there is no information spillover from U.S. to China.

As there are some potential limitations surrounding the multivariate GARCH model as addressed in Section 4, we report Hong's (2001) robust causality-in-mean and causality-in-variance test results in Tables 6 and 7, respectively. Table 6 presents the causality-in-mean test results for up to 10 lags. For each stock pair, we report one-way causality-in-mean test statistics Q from China to U.S. ($C \rightarrow U$) in the first column, and reverse one-way Q statistics from U.S. to China ($U \rightarrow C$) in the second column.

The evidence from the above uni-directional tests is interesting. Unlike the multivariate Granger causality tests reported in Table 4, we find some evidence of one-lag causality in mean from U.S. to China for four stocks, CEA, SNP, CHU, and HNP. The causality-in-mean effects, however, are transient and limited to just one day. For ZNH and SHI, we find mean spillover effects from China to U.S., which could last up to 10 days.

Table 6. Test statistics for causality-in-mean: ADR prices and Chinese A-share prices

For each stock pair, $C \rightarrow U$ refers to one-way causality from China to U.S. and $U \rightarrow C$ denotes one-way causality from U.S. to China. The test statistics reported in the first column are for the null hypothesis of ***no*** casual relation from China to U.S., and in the second column, for ***no*** casual relation from U.S. to China. ** , *, and # denote statistical significance at the 1 percent, 5 percent, and 10 percent, respectively, using the single-tailed normal distribution.

	CEA		SNP		ZNH		CHU		HNP		SHI		YZC	
Lag	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$
1	-1.12	2.15*	-1.68	2.06*	-0.26	-0.22	-1.79	1.33#	-3.03	3.95**	3.68**	-0.89	-2.61	0.98
2	1.01	-0.74	-0.83	0.20	0.58	-0.51	-0.55	-0.80	-0.70	-0.28	2.20**	-0.56	-0.79	-0.49
3	0.93	-0.84	-0.84	0.32	1.51#	-0.56	-0.66	-0.87	-0.78	-0.14	2.96**	-0.10	-0.81	-0.60
4	0.68	-0.96	-1.00	0.31	1.88*	-0.53	-0.74	-0.98	-0.81	-0.07	3.59**	-0.01	-0.97	-0.69
5	0.52	-1.04	-1.07	0.12	1.99*	-0.52	-0.92	-1.00	-0.95	-0.14	4.13**	0.12	-1.03	-0.84
6	0.37	-1.15	-1.11	0.22	2.32*	-0.55	-0.95	-1.19	-1.04	-0.21	4.44**	0.32	-1.04	-0.97
7	0.27	-1.22	-1.07	0.44	2.68**	-0.53	-1.04	-1.12	-1.07	-0.18	4.61**	0.43	-1.02	-1.01
8	0.15	-1.18	-1.11	0.63	2.95**	-0.48	-1.11	-1.05	-0.97	-0.21	4.73**	0.50	-1.04	-1.10
9	0.11	-1.16	-1.20	0.76	3.12**	-0.43	-1.15	-1.12	-0.87	-0.19	4.73**	0.47	-1.08	-1.15
10	0.12	-1.10	-1.28	0.89	3.21**	-0.45	-1.18	-1.25	-0.83	-0.17	4.70**	0.45	-1.06	-1.15

Table 7 reports the corresponding tests for causality in variance. Consistent with what we find in Table 4, for five stocks, CEA, ZNH, CHU, SHI, and YZC, we find strong volatility spillovers from China to U.S., and the spillover effects could last up to 10 days. The tests also show that only three ADRs (ZNH, CHU, and YZC) exhibit some volatility spillovers from U.S. to China, but their effects (except ZNH) occur most likely within the first lag from U.S. to China. Furthermore, there are only two stocks, SNP and HNP, for which we do not find any significant interaction through volatility.

The above results from short-run dynamics of information transmission provide two notable implications. First, the U.S. investors seem to play a temporal role in price discovery of the cross-listed stocks in China. Despite the absence of a long-run cointegration relationship between the U.S and Chinese markets, some strong, temporal, and causal linkages from China to U.S. still prevail. The uni-directional volatility transmission from China to U.S. also indicates that the price discovery in China is important for the price discovery of the cross-listed ADRs. Our finding is consistent with Xu and Fung (2002) in that the home market appears to play a more significant role of information transmission in the pricing process than the offshore market.

Second, the differences in ability to transmit information shown in this paper indeed reinforce the important role of location (i.e., the "real home" market that is closer to corporate headquarters where information is originated) in the price discovery process for our sample. While our paper provides strong support for the home bias hypothesis, it casts doubt on the common belief that price discovery *should* occur on the foreign exchange (NYSE) whose market quality is superior to that of the domestic exchange (SHSE).

Finally, the pattern of information transmission is shown to exhibit a short-lived phenomenon from the U.S. market to the Chinese market, but causality is rarely detected to exist in longer lags (with the exception of lag 1) except for ZNH. This finding suggests that the market price adjustments in China, responding to news originating from U.S., are speedy and completed in one day, which is analogous to Hamao, Masulis, and Ng's (1990) discovery of a bi-directional volatility spillover in a day or two between the U.S. market and the Japanese market.

Table 7. Test statistics for causality-in-volatility: ADR prices and Chinese A-share prices

For each stock pair, $C \rightarrow U$ refers to one-way causality from China to U.S. and $U \rightarrow C$ denotes one-way causality from U.S. to China. The test statistics reported in the first column are for the null hypothesis of no casual relation from China to U.S., and in the second column, for no casual relation from U.S. to China. ** , *, and # denote statistical significance at the 1 percent, 5 percent, and 10 percent, respectively, using the single-tailed normal distribution.

	CEA	SNP	ZNH	CHU	HNP	SHI	YZC

Lag	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$	$C \rightarrow U$	$U \rightarrow C$
1	1.42$^{\#}$	-0.77	-6.73	0.63	1.79*	-1.58	-5.39	4.68**	-9.79	-4.86	16.39**	-17.10	-10.60	1.52**
2	22.36**	0.05	-0.05	-0.58	1.38$^{\#}$	1.14	4.35**	-0.80	0.29	-0.82	-0.42	-0.75	9.40**	-0.70
3	21.66**	0.12	-0.18	-0.30	2.79**	2.35**	4.53**	-0.77	0.17	-0.91	-0.07	-0.90	11.14**	-0.79
4	19.73**	0.29	-0.22	-0.75	3.46**	3.22**	4.38**	-1.05	-0.07	-1.01	-0.55	-1.00	11.92**	-0.89
5	17.91**	0.21	-0.24	-0.52	3.90**	3.47**	3.96**	-0.21	-0.27	-1.12	-0.50	-1.15	11.80**	-0.94
6	16.27**	0.12	-0.28	-0.07	4.10**	3.40**	3.55**	-0.93	-0.44	-1.25	0.10	-1.26	11.25**	-0.94
7	15.00**	0.12	-0.32	0.34	4.29**	3.23**	3.39**	-0.40	-0.59	-1.33	0.80	-1.32	10.62**	-0.90
8	13.91**	0.19	-0.39	0.57	4.40**	3.05**	3.23**	0.20	-0.70	-1.42	1.37$^{\#}$	-1.38	9.95**	-0.86
9	12.95**	0.29	-0.42	0.69	4.37**	2.87**	3.01**	0.13	-0.79	-1.49	1.76*	-1.42	9.40**	-0.81
10	12.13**	0.38	-0.44	0.76	4.30**	2.65**	2.85**	-0.32	-0.89	-1.53	2.05*	-1.47	8.93**	-0.76

Information Transmission between SEHK and NYSE

So far, our finding strongly indicates that the "real home" market (China) exerts a significant influence on both mean and volatility of returns in the offshore market (NYSE) for most ADRs. As aforementioned, Xu and Fung (2002) have examined the information flows between China-backed stocks dual-listed on both SEHK and NYSE. However, most of the ADRs (7 out of 10) in their sample did not have A-shares listed in China prior to 2000. In the following analysis, we will re-examine the information transmission between Hong Kong and U.S. by restricting our sample to the only nine Chinese companies cross-listed on NYSE, SEHK, and SHSE by March 2007. The sample period will cover the period before and after A-shares are listed in China. As shown in Table 1, all A-shares are introduced at a later date after their listing on either SEHK or NYSE. Having A-shares traded in the "real home" market indeed adds another link in the price discovery process. This unique feature of our sample allows us to investigate the effects of introducing A-shares on the price discovery process between Hong Kong H-shares and their corresponding U.S. ADRs.

Specifically, we re-examine the information transmission between SEHK and NYSE based on two subsamples. The first subsample focuses on the impact of A-shares trading in China on the price discovery between SEHK and NYSE. The subsample includes seven ADRs (CEA, SNP, ZNH, CHU, HNP, SHI, and YZC) with the sample period covering from one year after the SHSE listing date to March 31, 2007. The second subsample investigates the price discovery between SEHK and NYSE without A-share trading and includes only five ADRs (ZNH, CHU, HNP, GSH, and LFC) with the sample period covering from one year after the later listing date on either NYSE or SEHK to the date immediately before the introduction of A-share trading in China. The reason why we exclude four ADRs (CEA, SNP, SHI, and YZC) from our second subsample is mainly because their respective sample period is shorter than one year, which is not long enough for us to draw any reliable inference on the information transmission between SEHK and NYSE. Overall, the above empirical design can not only shed light on the effects of A-share introduction but also allow us to compare our results with those reported in Xu and Fung (2002)[9].

Table 8 presents Johansen's (1991) cointegration test results. Tables 9 and 10 report the BEKK estimation results and Hong's (2001) causality-in-volatility test results, respectively.

The results for both subsamples show that Hong Kong H-share prices are cointegrated with ADR prices in the long run. We also implement the recursive cointegration analysis, which confirms the stability of a cointegration relationship for both subsamples.[8] When we apply the ECM-GARCH model to examine volatility spillovers between Hong Kong and U.S., the results reported in Panel A and Panel B of Table 9 strongly suggest that there are volatility spillovers from the Hong Kong market to the U.S. market. Specifically, for our first seven-ADR subsample in Panel A, six companies are associated with significant a_{21} or b_{21}. Only one ADR (SHI) shows information transmission from U.S. to Hong Kong with significant a_{12} or b_{12}. Panel B indicates that even before the introduction of A-share trading, there is strong evidence on volatility spillovers from the Hong Kong market to the U.S. market (for four out of five ADRs). Hong's causality-in-variance Q test results presented in Table 10 confirm the consistent evidence.

In summary, Xu and Fung (2002) document that stocks in Hong Kong appear to play a more significant role of information transmission in the pricing process, whereas stocks listed in the U.S. play a bigger role in volatility spillover. However, the evidence reported in Tables 8-10 indicate that although stocks in Hong Kong are cointegrated with their respective ADRs in the long run, volatility, however, has spilled over from Hong Kong to the U.S. both before and after the introduction of A-share trading in China. While our results on the information dynamics between Hong Kong and U.S. are somewhat different from Xu and Fung (2002), the evidence reported here consistently supports that location could be the essential factor in price discovery between the U.S. market and the pseudo home market (Hong Kong) in China.

[8] The results are available upon request.

Table 8. Johanson's cointegration tests: ADR prices and Hong Kong H-share prices

The null hypothesis is zero cointegrating vector. *r* refers to the number of cointegrating vectors. The optimal lag order of VAR is chosen by the criterion of Akaike information criterion (AIC). The corresponding 95 percent critical values for maximum eigenvalue and trace statistics are 15.89 and 20.26, respectively. * denotes rejection of the null hypothesis at the 5% level.

Panel A	Maximum Eigenvalue $H_0: r = 0$ $H_1: r = 1$	Trace Statistic $H_0: r = 0$ $H_1: r = 1$	Lag	Sample Period
CEA	141.28*	145.64*	7	11/05/98-03/31/07
SNP	28.01*	34.37*	3	08/08/02-03/31/07
ZNH	36.34*	37.64*	9	07/26/04-03/31/07
CHU	217.52*	218.82*	4	10/09/03-03/31/07
HNP	164.70*	174.72*	4	12/06/02-03/31/07
SHI	136.34*	138.77*	14	11/08/94-03/31/07
YZC	80.45*	81.99*	10	07/01/99-03/31/07

Panel B	Maximum Eigenvalue $H_0: r = 0$ $H_1: r = 1$	Trace Statistic $H_0: r = 0$ $H_1: r = 1$	Lag	Sample Period
ZNH	231.13*	237.06*	3	07/31/98-07/24/03
CHU	98.97*	104.54*	4	06/22/01-10/08/02
HNP	123.39*	125.72*	3	01/21/99-12/05/01
GSH	111.81*	112.70*	9	05/14/97-12/21/06
LFC	36.26*	57.00*	12	12/18/04-01/08/07

Table 9. BEKK model estimation results: ADR prices and Hong Kong H-share prices

Panel A	CEA		SNP		ZNH		CHU		HNP		SHI		YZC	
Sample Period	11/05/98-03/31/07		08/08/02-03/31/07		07/26/04-03/31/07		10/09/03-03/31/07		12/06/02-03/31/07		11/08/94-03/31/07		07/01/99-03/31/07	
Parameters	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*	*Estimate*	*s.e.*
c_{11}	0.02**	0.00	0.01**	0.00	0.01**	0.00	-0.03**	0.00	0.00	0.00	0.00	0.00	0.02**	0.00
c_{12}	0.00	0.00	-0.01**	0.00	-0.01**	0.00	0.00	0.00	0.00**	0.00	-0.01**	0.00	-0.01**	0.00
c_{22}	0.00	0.01	0.00	0.00	0.00	0.00	0.00	0.00	-0.02**	0.00	-0.01**	0.00	0.00	0.01
a_{11}	-0.07	0.05	-0.32**	0.03	-0.27**	0.07	0.29**	0.11	-0.14**	0.05	-0.19**	0.04	0.22**	0.04
a_{12}	0.17	0.40	0.02	0.14	-0.01	0.16	-0.05	0.36	0.00	0.18	0.22**	0.05	-0.03	0.33
a_{21}	0.29**	0.02	0.13**	0.04	0.18**	0.04	-0.28**	0.06	-0.25**	0.04	0.00	0.03	-0.30**	0.03
a_{22}	-0.09	0.20	0.00	0.13	-0.05	0.10	0.00	0.22	-0.01	0.15	-0.36**	0.03	0.06	0.20
b_{11}	-0.66**	0.04	0.90**	0.02	-0.72**	0.08	-0.26	0.20	0.75**	0.06	0.01	0.02	-0.74**	0.05
b_{12}	-0.13	0.42	-0.02	0.13	0.00	0.15	0.02	0.24	-0.01	0.16	-0.51**	0.16	0.03	0.35
b_{21}	0.02**	0.01	-0.09**	0.03	0.09	0.06	0.15**	0.05	-0.16**	0.03	-0.02	0.02	0.05*	0.02
b_{22}	0.01	0.04	0.00	0.03	0.00	0.05	0.00	0.07	0.00	0.04	0.90**	0.02	0.00	0.03

Panel B	ZNH		CHU		HNP		GSH		LFC	
Sample Period	07/31/98-07/24/03		06/22/01-10/08/02		01/21/99-12/05/01		5/14/97-12/21/06		12/18/04-01/08/07	
Parameters	Estimate	s.e.	Estimate	s.e.	Estimate	s.e.	Estimate	s.e.	Estimate	s.e.
c_{11}	0.00	0.00	0.03**	0.00	-0.03**	0.00	-0.10**	0.00	-0.09**	0.00
c_{12}	0.01**	0.00	0.00	0.00	-0.02**	0.00	0.00**	0.00	0.00	0.00
c_{22}	0.00**	0.00	0.00	0.01	0.00	0.01	0.00	0.00	0.00	0.00
a_{11}	-0.43**	0.04	0.42**	0.14	0.48**	0.09	-0.37**	0.11	0.47**	0.19

a_{12}	-0.50**	0.08	0.11	0.30	0.03	0.34	0.09	0.14	-0.04	0.20
a_{21}	-0.16**	0.02	-0.39**	0.08	-0.21**	0.05	0.27**	0.11	-0.41	0.34
a_{22}	-0.19**	0.04	0.00	0.17	-0.05	0.18	-0.14	0.15	0.06	0.37
b_{11}	-0.80**	0.03	-0.61**	0.20	0.00	0.38	0.84	0.07	0.78**	0.06
b_{12}	0.22	0.14	0.03	0.30	0.00	0.57	-0.05	0.10	0.00	0.05
b_{21}	0.12**	0.05	0.26**	0.05	0.00	0.12	-0.62	0.06	-0.72	0.06
b_{22}	0.23**	0.07	-0.01	0.10	0.00	0.11	0.04	0.07	0.00	0.05

** and * denote statistical significance at the 1 percent and 5 percent levels, respectively. *s.e.* denotes standard error.

Table 10. Test statistics for causality-in-volatility: ADR prices and Hong Kong H-share prices

For each stock pair, $H \rightarrow U$ refers to one-way causality from HK to US and $U \rightarrow H$ denotes one-way causality from US to HK. The test statistics reported in the first column are for the null hypothesis of no casual relation from HK to US, and in the second column, for no casual relation from US to HK. **, *, and # denote statistical significance at the 1 percent, 5 percent, and 10 percent, respectively, using the single-tailed normal distribution.

Panel A	CEA		SNP		ZNH		CHU		HNP		SHI		YZC	
Sample Period	11/05/98-03/31/07		08/08/02-03/31/07		07/26/04-03/31/07		10/09/03-03/31/07		12/06/02-03/31/07		11/08/94-03/31/07		07/01/99-03/31/07	
Lag	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$
1	10.12**	-4.41	3.83**	-2.99	5.01**	-4.29	7.77**	-6.72	3.13**	-4.75	7.25**	-2.07	10.53**	-9.22
2	284.37**	-0.44	1.27	-0.54	-0.76	0.01	2.04*	2.54**	8.59**	1.09	173.20**	2.54**	70.22**	-0.60
3	276.97**	-0.75	1.18	-0.49	-0.49	0.24	1.89*	1.85*	9.03**	1.01	168.80**	2.65**	68.67**	-0.87
4	254.32**	-0.73	0.88	-0.16	-0.12	1.19	1.47	1.99*	8.77**	0.73	155.20**	2.49**	63.17**	-0.90
5	231.79**	-0.65	0.68	-0.02	-0.02	2.19*	0.70	2.24*	8.20**	0.47	141.90**	2.23*	57.66**	-0.84
6	212.76**	-0.42	0.55	0.12	-0.21	2.77**	1.07	2.50**	7.57**	0.30	130.28**	1.94*	52.91**	-0.83
7	197.22**	-0.52	0.46	0.13	0.00	3.07**	1.93*	2.72**	6.94**	0.10	120.94**	1.66*	48.91**	-0.86
8	184.39**	-0.74	0.37	0.16	0.16	3.23**	2.83**	2.74**	6.39**	-0.03	113.15**	1.37	45.64**	-0.89
9	173.73**	-0.85	0.26	0.14	0.15	3.32**	3.50**	2.69**	5.94**	-0.15	106.48**	1.25	42.79**	-0.97
10	164.59**	-0.85	0.21	0.12	-0.05	3.20**	4.04**	2.67**	5.52**	-0.29	100.67**	1.11	40.36**	-1.04

Panel B	ZNH		CHU		HNP		GSH		LFC	
Sample Period	07/31/98-07/24/03		06/22/01-10/08/02		01/21/99-12/05/01		5/14/97-12/21/06		12/18/04-01/08/07	
Lag	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$	$H \rightarrow U$	$U \rightarrow H$
1	9.25**	-7.11	0.53	-0.94	3.43**	-1.76	2.85**	-2.91	-2.43	2.54**
2	134.76**	-0.60	5.52**	-0.44	7.87**	-0.78	6.29**	12.40**	-0.75	-0.70
3	135.02**	-0.18	5.86**	-0.14	7.77**	-0.88	3.83**	12.18**	-0.52	-0.84
4	126.07**	0.28	5.72**	0.02	7.41**	-0.93	5.39**	12.03**	-0.27	-1.00
5	117.20**	0.51	5.36**	0.11	6.97**	-1.11	7.74**	11.64**	-0.28	-1.05
6	109.59**	0.61	4.95**	0.01	6.52**	-1.24	9.36**	11.28**	-0.23	-1.20
7	103.07**	0.52	4.71**	-0.12	6.01**	-1.27	10.27**	10.88**	-0.22	-1.28
8	97.42**	0.43	4.39**	-0.17	5.57**	-1.22	10.72**	10.47**	-0.40	-1.33
9	92.48**	0.45	4.07**	-0.28	5.22**	-1.14	10.82**	10.06**	-0.49	-1.38
10	88.13**	0.44	3.81**	-0.41	4.96**	-1.06	10.81**	9.65**	-0.53	-1.41

CONCLUSION

Does price discovery occur in the cross-listed company's home market or on the exchange abroad? We attempt to address this issue by examining seven U.S.-listed Chinese stocks. The main empirical findings are summarized as follows. First, the Chinese A-share market was still segmented from the U.S. market for its cross-listed stocks as of March 2007. This finding extends the arguments of previous studies on the segmented Chinese A-share and B-share markets (e.g., Fung et al., 2000; Yang, 2003).

Second, the volatility in the Chinese A-share market significantly influences cross-listed Chinese stocks in the U.S. market. It is shown that all the information accrued during the trading on the SHSE is effectively transmitted into the trading on the NYSE the following day. In other words, foreign price adjustments occur sequentially in response to their counterparts in the domestic market. In addition, the information flows from U.S. to China tend to be temporal and short-lived.

Based on our empirical design, the relative importance of location and market quality in price discovery origination is examined. We find that the “real home” market plays a more important and consistent role in both price discovery and volatility spillover. Although this finding is intuitive, it sheds new light on the cross-listing literature that location *per se* is essentially the most important factor in price discovery, and the effect of superior market quality could be dominated by the home bias.

Our findings also raise an interesting issue on why cross-listed stocks, with the same underlying stocks and linked by the same set of economic fundamentals, have constantly shown long-run separated prices in different national markets. Given the fact that China’s stock markets have been gradually opened to foreign investors, the market force through arbitrage is clearly not fully at work at this stage because there is no cointegration relationship in the long run and cross-listing in the U.S. has exerted little short-run impact on the price formation process for the Chinese cross-listed stocks. Furthermore, trading restrictions, the regulations governing the Chinese stock market, and strict foreign exchange control are possible factors that may prevent arbitrage and cointegration from happening. We thus leave it for further research.

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