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Abstract

This paper investigates the relative role of price discovery between two long-term swap contracts that exchange between the U.S. dollar and the Japanese yen: cross-currency basis swap and FX (foreign exchange) swap. First, we show that these two swaps should be in a no-arbitrage relationship by allowing for differential risk premiums. Second, we empirically investigate the relative role of price discovery using the structural-form approach based on the state space models. Main finding are as follows. (i) The efficient prices extracted as a common factor of the two swaps show a very similar movement, regardless of model specifications. (ii) The currency swap market plays a much more dominant role in price discovery than the FX swap market. (iii) The FX swap prices tend to under-react to the efficient price changes, while the cross-currency swap prices almost exactly react to them.

Key Words: Currency Swap; FX Swap; Price Discovery, State Space Model, Efficient Price

JEL Classification: G12, G14, G15

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1. Introduction

This paper investigates the relative role of price discovery between two long-term swap contracts that exchange between the U.S. dollar (USD) and the Japanese yen: cross-currency swap and foreign exchange (FX) swap contracts. To that end, we use the structural-form approach based on the state space models.

Cross-currency and FX swap markets have been important markets for raising foreign currencies with terms longer than one year for both Japanese and non-Japanese banks. Although the scheme is somewhat different, they have the same economic function as exchanging between the USD and the yen.[1] Historically, liquidity of the long-term FX swap market had been very low, compared with the cross-currency swap market, and hence most of the studies have used only the cross-currency swap prices to test the long-term covered interest parity thus far. Recently, however, market participants argue that liquidity of the long-term FX swap market has improved substantially, and there are a number of active arbitragers who attempt to take profits when prices deviate each other. Hence, in this paper, we attempt to assess the relative role of price discovery between these two swap markets by extracting the efficient price that is common to both markets. More specifically, we proceed by taking the following steps.

First, we show that the cross-currency swap and the FX swap should be in a no-arbitrage relationship by allowing for differential risk premiums. The literature uses cross-currency swap prices to test the long-term interest rate parity, ignoring the differential risk premiums between counterparties, as well as between currencies.[2] This simplification seems to have been relevant prior to the middle of the late 1990s. Since the late 1990s, however, the creditworthiness of

[1] The price of cross-currency swaps is expressed as the spread over the LIBOR-based interest rates, while the price of the FX swaps is expressed as the spread between FX forward and spot rates.

[2] See Popper [1993], Fletcher and Taylor [1994], and Takezawa [1995], among others. They used the currency swap data in the late 1980s or early 1990s to test the long-term interest parity, and found the non-negligible deviations from the parity, although such deviations diminished over time.

Japanese banks deteriorated so much, due mainly to the non-performing loan problem, that they were obliged to pay the so-called Japan premiums on a global basis.[3] The Japan premiums were observed not only in the short-term money markets, but also in the longer-term markets until recently, as shown by Baba *et al.* [2006].[4] We also have consistently observed non-negligible time-varying differential risk premiums for both Japanese and non-Japanese banks between the USD and the yen markets. In fact, as shown in section 2, we cannot explain the real-life price movement of cross-currency swap and FX swap without allowing for such differential risk premiums.

Second, given that both swap prices should be in a no-arbitrage condition, we investigate the relative role of price discovery between the cross-currency swap and FX swap markets. Empirical methodologies to test price discovery have progressed significantly in the past decade or so. First, Gonzalo and Granger [1995] and Hasbrouck [1995] proposed a price discovery measure, respectively, based on the vector error correction model (VECM). We call this type of methodology a reduced-form approach. On the other hand, Menkveld, Koopman, and Lucas [2007] recently opened up the way to model the unobserved efficient price common to cross-listed stocks using a state space model, and successfully gauged the relative role of price discovery between the two markets under study. We call this type of methodology a structural-form approach following Harvey [1989].[5]

These two approaches have merits and demerits, respectively. The reduced-form approach places fewer *ad hoc* restrictions on the data than the structural-form approach. In terms of the interpretability of estimation results, however, the structural-form approach is more

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[3] For more details of the Japan premium, see Covrig, Low, and Melvin [2004] and Ito and Harada [2000], among others.

[4] Baba *et al.* [2006] showed that the quantitative easing policy by the Bank of Japan contained credit risks for Japanese banks in the short-term money markets, but the non-negligible credit risk premium remained in the long-term credit markets, including the straight bond and credit derivatives markets until 2003.

[5] Harvey [1989] defines a structural time-series model as one that is set up in terms of unobservable components that have a direct interpretation.

straightforward than the reduced-form approach. The structural-form approach can also be used more effectively and easily for various hypothesis-testing including the partial vs. complete adjustment of market prices to the efficient price and under/overreaction of market prices to the efficient price. In this paper, we use the structural-form approach due mainly to the effectiveness and easiness of such hypothesis-testing and higher interpretability. But, taking the above-mentioned merits and demerits into account, we also use the more conventional reduced-form approach to check robustness of the estimation results from the structural approach.

The rest of the paper is organized as follows. Section 2 briefly reviews the basic schemes of and recent developments in the cross-currency swap and FX swap markets. Section 3 constructs the no-arbitrage condition between these two swap markets by allowing for differential risk premiums. Section 4 descries the data. Section 5 reviews the structural models we use. Section 6 reports the estimation results. Section 7 concludes the paper.

2. Basic Schemes of Cross-Currency Basis Swap and FX Swap

2.1 Cross-Currency Basis Swap

There are numerous types of cross-currency swap contracts, among which the most widely used is the following type of contracts named the cross-currency basis swap.[6] A typical cross-currency basis swap (currency swap, hereafter) agreement is a contract in which Japanese banks borrow USD from, and lend yen to non-Japanese banks simultaneously. Figure 1 (i) illustrates the flow of funds associated with this currency swap. At the start of the contract, bank A (a Japanese bank) borrows X USD from, and lends $X \times S$ yen, to bank B (a non-Japanese bank), where S is the FX spot rate. During the contract term, bank A receives yen-LIBOR 3M+α from, and pays USD-LIBOR 3M,

[6] The most traditional cross-currency swaps are the contracts in which fixed interest rates are exchanged between currencies. Another example is the coupon swaps in which interest rates are exchanged between currencies, but there is no exchange of principals at the start and end of the contract.

to bank B, every 3 months.[7] When the contract expires, bank A returns X USD to bank B, and bank B returns $X \times S$ yen to bank A. At the start of the contract, both banks decide α, which is the price of the basis swap. In other words, bank A (B) borrows foreign currency by putting up its home currency as collateral, and hence this swap is effectively a collateralized contract.

These currency swaps have been employed by both Japanese and non-Japanese banks to fund foreign currencies, both for their own and their customers, including multinational corporations engaged in foreign direct investment. Currency swaps have been also used as a hedging tool, particularly for issuers of the so-called *Samurai* bonds, which are yen-denominated bonds issued in Japan by non-Japanese companies. By nature of these transactions, most of them are long-term from one year to 30 years.

2.2 FX Swap

A typical FX swap agreement is also a contract in which Japanese banks borrow USD from, and lend yen to non-Japanese banks simultaneously.[8] The main differences from the currency swap are that (i) during the contract term, there are no exchanges of interest between yen and USD rates; and (ii) at the end of the contract, the different amount of funds are returned from the amount exchanged at the start.

Figure 1 (ii) illustrates the flow of funds associated with the FX swap. At the start of the contract, bank A (Japanese bank) borrows X USD from, and lends $X \times S$ yen, to bank B (non-Japanese bank), where S is the FX spot rate. When the contract expires, bank A returns X USD to bank B, and bank B returns $X \times F$ yen to bank A, where F is the FX forward rate as of the start of the contract. As is the case with currency swaps, FX swaps are effectively collateralized contracts.

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[7] LIBOR 3M is the three-month London Interbank Offered Rate.

[8] The explanation here follows Nishioka and Baba [2004].

FX swaps have been employed by both Japanese and non-Japanese banks for funding foreign currencies, both for their own and their customers, including exporters and importers, as well as Japanese institutional investors investing in hedged foreign bonds. FX swaps have also been used as a tool for speculative trading. The most liquid term is shorter than one year, but transactions with longer maturities have been actively conducted from such motives as foreign currency funding for corporate direct investments and arbitrage activities with cross-currency currency swaps. In fact, many market participants point out that liquidity of the FX swaps with maturities longer than one year has improved during the past several years.

3. No-Arbitrage Conditions between Currency Swap and FX Swap Markets

3.1 Basis Setup

In this section, we construct the no-arbitrage condition between currency swap and FX swap markets. The literature uses only currency swap prices to test the long-term covered interest parity, ignoring the differential risk premiums between lenders and borrowers, as well as between currencies, although some noticed its potential importance.[9]

Figure 2 shows a proxy for the risk premiums for Japanese and non-Japanese banks in the USD and the yen markets.[10] Here, we notice substantial differences in the risk premiums between Japanese and non-Japanese banks in the same currency market, as well as between the USD and the yen markets for the same bank group. The differential risk premiums between Japanese and

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[9] For the long-term interest rate parity, see Popper [1993], Fletcher and Taylor [1994, 1996], Fletcher and Sultan [1997] and Takezawa [1995], among others. Regarding the differential risk premiums, Popper [1993] mentions the potential bias from not considering those, particularly in Euromarkets.

[10] Risk premiums for Japanese banks in the USD/yen markets are calculated as 1-year USD-TIBOR/yen-TIBOR minus yields on 1-year U.S./Japanese government bonds. Risk premiums for non-Japanese banks in the USD/yen markets are calculated as 1-year USD-LIBOR/yen-LIBOR minus yields on 1-year U.S./Japanese government bonds. TIBOR is the Tokyo Interbank Offered Rate. As of the end of March 2007, USD-TIBOR /yen-TIBOR is calculated based on rates quoted by a panel of 9/15 banks, of which 7/14 banks are Japanese banks. On the other hand, USD-LIBOR/yen-LIBOR is based on rates quoted by a panel of 16/16 banks, of which 2/4 banks are Japanese banks.

non-Japanese banks have been usually explained by the so-called Japan premium story.[11] Since the late 1990s, deterioration in creditworthiness of Japanese banks relative to other advanced nations' banks has significantly influenced their foreign currency funding, particularly USD funding. The deterioration of creditworthiness was originally caused by the non-performing loan problem triggered by the bursting of the asset bubbles in the early 1990s.

More puzzling is the larger and more persistent differential risk premiums between the USD and the yen markets for the same bank group. Figure 2 shows that risk premiums are much higher in the USD market than in the yen market for both Japanese and non-Japanese banks and fluctuate widely over time. Market participants often cite a difference in main participants and hence in attitudes toward risk evaluation as its main reason.[12] Aside from the reasons, there surely exist the non-negligible differential risk premiums, particularly between the USD and the yen market for the same bank group, and hence in this paper, we explicitly allow for such differential risk premiums to construct the no-arbitrage conditions linking the currency and FX swap market.

Specifically, we start by describing the typical funding structure of Japanese and non-Japanese banks, following Nishioka and Baba [2004]. The funding costs in the yen and USD markets are the sum of the risk-free interest rate and the risk premium for the representative Japanese or non-Japanese bank in each market. Let r_{jpy} (r_{usd}) denote the yen (the USD) risk-free interest rate, ϕ_{jpy} (ϕ_{usd}) the risk premium for the Japanese bank in the yen (USD) market, and θ_{jpy} (θ_{usd}) the risk premium for the non-Japanese bank in the yen (USD) market. The main source of risk premiums is credit or default risk of borrowers, but here we expand the notion of risk premiums to involving price movements caused by *ex ante* supply-demand and

[11] For the Japan premium, see Covrig, Low, and Melvin [2004], Ito and Harada [2000], and Peek and Rosengren [2001], for instance.
[12] Until quite recently, there are non-negligible differences in credit ratings for the same Japanese firms between Japanese credit-rating agencies and the U.S. ones; the ratings by the U.S. agencies were generally lower than those by Japanese agencies for the same firms.

liquidity conditions, not limiting to the conventional stationary component.

As shown in Figure 3, the Japanese (non-Japanese) bank has two alternative funding sources of the USD (yen): (i) raising the USD (yen) directly from the USD (yen) cash market; and (ii) first raising the yen (USD) from the yen (USD) cash market and exchanging it for the USD (yen) in the currency swap or FX swap market.

3.2 Currency Swap Market

Given the basic funding structure above, first, let us look at the no-arbitrage condition for the currency swap market. Because the interest rates in the currency swap are floating rates, for comparison with the FX swap prices, we need to exchange floating rates for fixed rates through interest rate swaps (IRS).[13] After this conversion and ignoring the transaction costs, the no-arbitrage conditions for the currency swap market can be written as[14]

$$1 + r_{usd} + \phi_{usd} = \left(1 + r_{jpy} + \phi_{jpy}\right) + \left[\left(1 + r_{usd}\right) - \left(1 + r_{jpy} + \alpha\right)\right] \tag{1}$$

for Japanese banks

$$1 + r_{jpy} + \theta_{jpy} = \left(1 + r_{usd} + \theta_{usd}\right) + \left[\left(1 + r_{jpy} + \alpha\right) - \left(1 + r_{usd}\right)\right].^{15} \tag{2}$$

for non-Japanese banks

Equilibrium in the currency swap market requires

$$\phi_{jpy} - \phi_{usd} = \theta_{jpy} - \theta_{usd} = \alpha \,. \tag{3}$$

The left-hand side of equation (3) denotes the difference in the risk premiums for the Japanese bank between the yen and USD markets and the right-hand side denotes the difference in the risk premiums for the non-Japanese bank between the yen and USD markets. Note here that without considering the differential risk premiums, the price for the currency swap α should be zero and hence the observed negative α cannot be rationalized. The generalization of the no-arbitrage

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[13] Here, we implicitly assume that the IRS fixed interest rates are equivalent to the risk-free interest rates in the cash markets. This assumption means that risk premiums are measured relative to the IRS rates.
[14] This is merely for simplicity. We consider the potential transaction costs in the empirical analysis.
[15] Note that these conditions hold even in the case that Japanese (non-Japanese) banks raise the yen (USD).

relationship we have proposed above, however, shows that α can take on both positive and negative values without violating the no-arbitrage condition.

3.3 FX Swap Market

Second, let us look at the no-arbitrage conditions for the FX swap market, which can be written as

$$1+r_{usd}+\phi_{usd}=\frac{S}{F}\left(1+r_{jpy}+\phi_{jpy}\right) \qquad \text{for Japanese banks} \tag{4}$$

$$1+r_{jpy}+\theta_{jpy}=\frac{F}{S}\left(1+r_{usd}+\theta_{usd}\right). \qquad \text{for non-Japanese banks} \tag{5}$$

Equilibrium in the FX swap market requires

$$\frac{1+r_{jpy}+\theta_{jpy}}{1+r_{jpy}+\phi_{jpy}}=\frac{1+r_{usd}+\theta_{usd}}{1+r_{usd}+\phi_{usd}}. \tag{6}$$

Equation (6) can be approximated as

$$\phi_{jpy}-\phi_{usd}\approx\theta_{jpy}-\theta_{usd}. \tag{7}$$

To facilitate the interpretation of equation (7), let us implicitly define β as

$$\frac{F}{S}=\frac{1+r_{jpy}+\beta}{1+r_{usd}}, \tag{8}$$

where β is the spread over the yen risk-free interest rate implied by the FX swap price quotation F/S and the USD risk-free interest rate. From equations (4), (5), and (8), we get

$$\phi_{jpy}-\phi_{usd}\approx\theta_{jpy}-\theta_{usd}\approx\beta. \tag{9}$$

Equation (9) means that the implicitly defined FX swap price β is approximately equivalent to the difference in the risk premiums for each bank between the yen and USD markets. Again, without allowing for differential risk premiums, β should be zero.

3.4 No-Arbitrage Condition between the Currency swap and FX Swap Markets

Now, we can combine equations (3) and (9) to derive the no-arbitrage condition between the currency swap and FX swap markets as follows:

$$\alpha \approx \beta \approx \phi_{jpy} - \phi_{usd} = \theta_{jpy} - \theta_{usd} . \tag{10}$$

α and β reflect the same fundamentals, and hence should be in the arbitrage relationship.

4. Data

For α, we use the cross-currency basis swap price reported by Bloomberg under the name of JPY BASIS SW. The price is as of the NY closing time (17:00). On the other hand, we calculate β according to equation (8) using the FX spot rate of the yen against the USD (USD-JPY X-RATE), spot-forward spread (JAPNESE YEN), USD IRS rate (USD SWAP), and the yen IRS rate (JPY SWAP). All of these variables are as of the NY closing time and obtained from Bloomberg. We use the midpoints of the observed bid-ask quotes. Maturities are from 1 year to 5 years. The sample period is from July 1, 1997 to November 30, 2006, and the number of observations is 2,458.

Figure 4 plots α and β and Table 1 reports summary statistics of α and β. From these, we find that regardless of maturity: (i) means of both α and β are negative; (ii) means of β is smaller than those of α; and (iii) β is more volatile in terms of standard deviations than α. Recall that the conventional no-arbitrage condition that does not allow for differential risk premiums predicts that α and β should be zero. Hence, consistently negative α and β indicate the violation of the conventional no-arbitrage condition. We have already proved that this is not the case if we allow for differential risk premiums that have actually been observed in the markets. Table 2 reports the result of the ADF (Augmented Dickey-Fuller) test and shows that α and β are found to be I(1) at the 5% significance level expect 3-year β.

5. Structural Models of Price Discovery

Price discovery is defined by Lehmann [2002] as the efficient and timely incorporation of information implicit in investor trading into market prices. When the same fundamentals are priced into two markets, order flow is fragmented and price discovery is split between these markets. There are various specifications for structural models. In this paper, we adopt the following three models that are extensively investigated in the literature.

5.1 Baseline Model

$$\alpha_t = m_t + s_{\alpha,t} \qquad s_{\alpha,t} \sim N\left(0, \sigma_{s\alpha}^2\right) \tag{11}$$

$$\beta_t = m_t + s_{\beta,t} \qquad s_{\beta,t} \sim N\left(0, \sigma_{s\beta}^2\right) \tag{12}$$

$$m_t = m_{t-1} + \xi_t \qquad \xi_t \sim N\left(0, \sigma_\xi^2\right) \tag{13}$$

This model is similar to the structural-form model used in Lehmann [2002]. Here, there is an unobservable efficient price m_t that follows a random-walk process, which is common to both swap prices.[16] The random walk representation of the efficient price dates back to Samuelson [1965]. The parameters in the state space model are estimated by maximizing the log-likelihood that can be evaluated by the Kalman filter.[17] Throughout the paper, we assume that each residual is mutually independent.

Under this setting, we calculate the following "signal-to-noise" measures to assess price discovery, which is a share of efficient price volatility in the total volatility for each swap price:[18]

$$\mathrm{SIS}_\alpha^{(1)} = \frac{\sigma_\xi^2}{\sigma_\xi^2 + \sigma_{s\alpha}^2} \quad \text{and} \quad \mathrm{SIS}_\beta^{(1)} = \frac{\sigma_\xi^2}{\sigma_\xi^2 + \sigma_{s\beta}^2}. \tag{14}$$

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16 Another popular specification is the local linear trend specification proposed in Harvey[1989]:

$$m_t = m_{t-1} + \delta_{t-1}$$

$$\delta_t = \delta_{t-1} + \omega_t \quad \omega_t \sim N\left(0, \sigma_\omega^2\right),$$

where the efficient price m_t is a random walk process with a stochastic drift δ_t. We also tried this specification, but we found that the estimated δ_t is not significantly different from zero. Hence, we did not choose this specification.

17 For details of the state space models, see Durbin and Koopman [2001], for instance.

18 The conventional definition of the signal-to-noise ratio is the ratio of efficient price volatility to stochastic noise volatility. We use our form of the signal-to-noise ratio due mainly to the ease of the comparison between α and β.

We call this measure the structural information share (SIS) in this paper. In what follows, we conduct the Wald test to test whether these two measures are significantly different from each other.

5.2 Partial Adjustment Model

$$\alpha_t = C_\alpha m_t + (1 - C_\alpha)\alpha_{t-1} + s_{\alpha,t} \qquad s_{\alpha,t} \sim N\left(0, \sigma_{s\alpha}^2\right) \tag{15}$$

$$\beta_t = C_\beta m_t + (1 - C_\beta)\beta_{t-1} + s_{\beta,t} \qquad s_{\beta,t} \sim N\left(0, \sigma_{s\beta}^2\right) \tag{16}$$

$$0 \le C_\alpha, C_\beta \le 2$$

$$m_t = m_{t-1} + \xi_t \qquad \xi_t \sim N\left(0, \sigma_\xi^2\right) \tag{17}$$

This model is a partial adjustment model similar to the model proposed in Amihud and Mendelson [1987] and Hasbrouck and Ho [1987]. $0 < C_i < 1$ represents the case of partial price adjustment, with C_i =0 and C_i =1 as special cases of no price reaction to the efficient price and complete price adjustment, respectively. Also notice that we do not exclude the case of overreaction or overshooting of prices to new fundamental information, which corresponds to $1 < C_i \le 2$. Under this setting, our structural information shares can be written as

$$\mathrm{SIS}_\alpha^{(2)} = \frac{C_\alpha^2 \sigma_\xi^2}{C_\alpha^2 \sigma_\xi^2 + \sigma_{s\alpha}^2} \quad \text{and} \quad \mathrm{SIS}_\beta^{(2)} = \frac{C_\beta^2 \sigma_\xi^2}{C_\beta^2 \sigma_\xi^2 + \sigma_{s\beta}^2}. \tag{18}$$

5.3 Under/Overreaction Model

$$\alpha_t = m_t + D_\alpha (m_t - m_{t-1}) + s_{\alpha,t} \qquad s_{\alpha,t} \sim N\left(0, \sigma_{s\alpha}^2\right) \tag{19}$$

$$\beta_t = m_t + D_\beta (m_t - m_{t-1}) + s_{\beta,t} \qquad s_{\beta,t} \sim N\left(0, \sigma_{s\beta}^2\right) \tag{20}$$

$$m_t = m_{t-1} + \xi_t \qquad \xi_t \sim N\left(0, \sigma_\xi^2\right) \tag{21}$$

This model also allows for possible under- or overreaction to new fundamental information, as emphasized by Amihud and Mendelson [1987]. This model is actually used by Menkveld, Koopman, and Lucas [2007] to investigate round-the-clock price discovery for cross-listed stocks. Here, significantly positive (negative) D_i indicates overreaction (under-reaction) to fundamental information. In this case, structural information shares can be written as

$$\mathrm{SIS}_{\alpha}^{(3)} = \frac{(1+D_{\alpha})^2 \sigma_{\xi}^2}{(1+D_{\alpha})^2 \sigma_{\xi}^2 + \sigma_{s\alpha}^2} \quad \text{and} \quad \mathrm{SIS}_{\beta}^{(3)} = \frac{(1+D_{\beta})^2 \sigma_{\xi}^2}{(1+D_{\beta})^2 \sigma_{\xi}^2 + \sigma_{s\beta}^2}. \tag{22}$$

6. Empirical Results

6.1 Estimation Results of Structural Models

Tables 3-5 report the estimation results of each state space model. In estimation, we added a constant term in the equation of α to adjust for the possible differences in institutional factors including transaction costs between α and β. First, Table 3 shows the result of the baseline model. All the variance coefficients and the constant term are significant at the 1% level. Estimated constant terms are within the range of 5-9 basis points, which is broadly consistent with anecdotal evidence of transaction costs. The structural information shares are much larger for α than β, and $\mathrm{SIS}_{\alpha}^{(1)}$ is found to be significantly higher than $\mathrm{SIS}_{\beta}^{(1)}$ at the 1% level for all the maturities.

Second, Table 4 shows the result of the partial adjustment model. All the coefficients are significantly estimated at the 1% level. The adjustment coefficients for α except 1 year are not significantly different from one, suggesting that the currency swap price α reflects the efficient price almost completely. On the other hand, those for β are within the range of 0.3-0.6. Consistent with these findings, the structural information shares for α is significantly higher than that for β at the 1% level.

Third, Table 5 shows the result of the under/overreaction model. Here, the under/overreaction coefficients for α are not significant for all the maturities, suggesting that α shows an almost exact response to the efficient price changes. On the other hand, those for β is significantly negative for all the maturities, suggesting that β tends to under-react to the efficient price changes. Consistent with this result, the structural information shares significantly favor α over β.

In sum, all of the results from the three models show that the currency swap price α has a significantly more important role of price discovery than the FX swap price β for all the maturities. α almost exactly reacts to the efficient price changes, but β tend to under-react to the efficient price changes. Figure 5 shows the efficient prices we estimated as the filtered state variables from these models.[19] From this figure, we can see a very similar movement of the efficient price regardless of the model specifications.

6.2 Robustness Check: Reduced-Form Analysis

As an attempt to check robustness of the above estimation results, we also estimate the price discovery measures based on the reduced-form approach. There are two approaches that attracted academic attention in this regard. One is the permanent-transitory (PT) model developed by Gonzalo and Granger [1995], and the other is the information share (IS) model developed by Hasbrouck [1995]. Both models rely on the estimation of the VECM of market prices.

The PT model decomposes the common factor itself and attributes superior price discovery to the market that adjusts least to price movements in the other market. On the other hand, the IS model decomposes the variance of the common factor based on the assumption that price volatility reflects new information flows and hence the market that contributes most to the

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[19] We used the Kalman filtering algorithm instead of the Kalman smoothing algorithm. Since the Kalman filtering is a forward recursion that computes one-step ahead estimates of the state variables, it seems more suitable to express the dynamic process of incorporating new information into market prices. On the other hand, the Kalman smoothing is a backward recursions based on the full data.

variance of the innovations to the common factor is considered to contribute most to price discovery.[20]

Table 6 reports the result of the cointegration test and the corresponding price discovery measures for the PT and IS models. First, Table 6 (i) shows that both Johansen trace and maximum eigenvalue tests suggest the significant existence of one cointegrating vector between α and β for each maturity at least at the 5% level. Second, Table 6 (ii) shows that both the PT and IS measures of price discovery are very close to one, which means that currency swap price has a significantly more important role of price discovery than the FX swap price. Hence, the result of the reduced-form model confirms our findings from the structural models.

7. Concluding Remarks

This paper has investigated the relative role of price discovery between two long-term swaps that exchange between the U.S. dollar and the Japanese yen: cross-currency (basis) swap and FX swap. First, we have shown that we should consider differential risk premiums, particularly between the yen and USD markets for the same bank group, to explain the negativity of the prices of .these two swaps using the no-arbitrage argument.

Second, we have empirically investigated the relative role of price discovery using three structural models. Main findings are as follows. (i) The efficient prices extracted as a common factor of the two swaps show a very similar movement, regardless of model specifications. (ii) The currency swap market plays a much more dominant role in price discovery than the FX swap market. (iii) The FX swap prices tend to under-react to the efficient price changes, while the currency swap prices almost exactly react to them. These results are broadly consistent with perceptions of market participants.

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[20] See Appendix for more details of each model.

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Appendix: Reduced-Form Approach of Price Discovery

The reduced-form approach starts from estimation of the following conventional VECM:

$$\Delta\alpha_t = \lambda_1(\alpha_{t-1} - \omega\beta_{t-1} - C) + \sum_{j=1}^{p} g_{1j}\Delta\alpha_{t-j} + \sum_{j=1}^{p} h_{1j}\Delta\beta_{t-j} + \varepsilon_{1t} \tag{A1}$$

$$\Delta\beta_t = \lambda_2(\alpha_{t-1} - \omega\beta_{t-1} - C) + \sum_{j=1}^{p} g_{2j}\Delta\alpha_{t-j} + \sum_{j=1}^{p} h_{2j}\Delta\beta_{t-j} + \varepsilon_{2t}\,, \tag{A2}$$

where $\alpha_{t-1} - \omega\beta_{t-1} - C$ denotes the error correction term, and ε_{1t} and ε_{2t} are i.i.d shocks. An underlying assumption here is that there is an unobservable efficient price common to both prices.

Based on the VECM above, the PT model decomposes the common factor itself and attributes superior price discovery to the market that adjusts least to price movements in the other market. As stated in Engle and Granger [1987], the existence of cointegration ensures that at least one market has to adjust. Price discovery measure for the first market can be measured by

$$\mathrm{PT} = \frac{\lambda_2}{\lambda_2 - \lambda_1} \tag{A3}$$

in the Gonzalo and Granger model

On the other hand, the IS model decomposes the variance of the common factor based on the assumption that price volatility reflects new information flows. Hence, the market that contributes most to the variance of the innovations to the common factor is considered to contribute most to price discovery. The price discovery for the first market can be measured as

$$\mathrm{IS}_1 = \frac{\lambda_2^2\left(\sigma_1^2 - \dfrac{\sigma_{12}^2}{\sigma_2^2}\right)}{\lambda_2^2\sigma_1^2 - 2\lambda_1\lambda_2\sigma_{12} + \lambda_1^2\sigma_2^2} \quad \text{and} \quad \mathrm{IS}_2 = \frac{\left(\lambda_2\sigma_1 - \lambda_1\dfrac{\sigma_{12}}{\sigma_1}\right)^2}{\lambda_2^2\sigma_1^2 - 2\lambda_1\lambda_2\sigma_{12} + \lambda_1^2\sigma_2^2}\,, \tag{A4}$$

where σ_1^2, σ_2^2 and σ_{12} are factors in the covariance matrix of ε_{1t} and ε_{2t}. Note here that IS_1 and IS_2 measure the lower and upper bounds of Hasbrouck's measure of price discovery, where the difference between these two bounds is positively related to the degree of correlation between residuals.[21] Baillie *et al.* [2002] argue that the average of these two bounds provides a sensible estimate of price discovery when the data frequency is high. Also note here that PT ignores the correlation between the markets and hence if the residuals are strongly correlated, then both models can provide substantially different results.[22]

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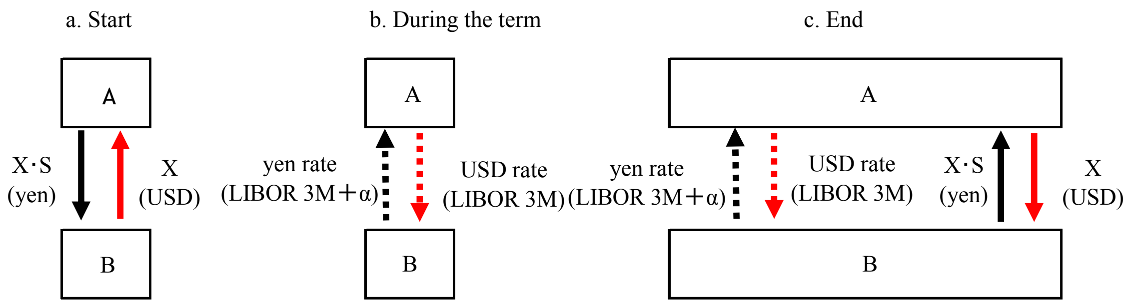
[21] When the residuals are not correlated, that is, the variance-covariance matrix of residuals is diagonal, the information share is identified. When the residuals are correlated, on the other hand, it is not identified because the result depends on the ordering of variables in the Cholesky factorization of the variance-covariance matrix. Hence, all one can do is to compute the lower and upper bounds.

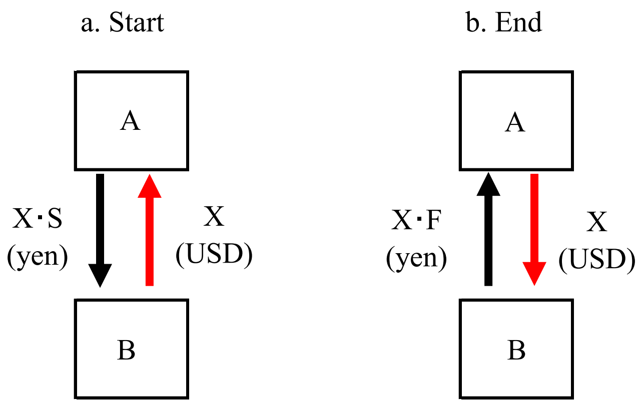
[22] Yan and Zivot [2006] rigorously analyze the determinants of these two price discovery measures under some structural model settings, in which both permanent (fundamental) and transitory shocks are identified, and the correlation between residuals from the VECM and each fundamental and transitory shock are explicitly taken into consideration. As a result, they find some inconsistency in the interpretation of these two price discovery measures, based on the reduced-from approach.

Figure 1: Basic Schemes of Cross-Currency Basis Swap and FX Swap

(i) Cross-Currency Basis Swap



(ii) FX Swap

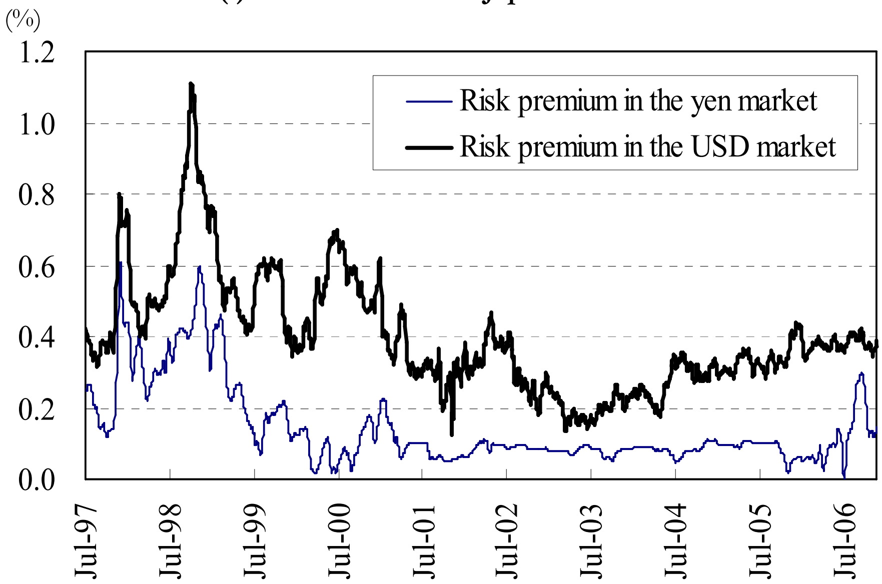


S : FX spot rate (yen/USD)

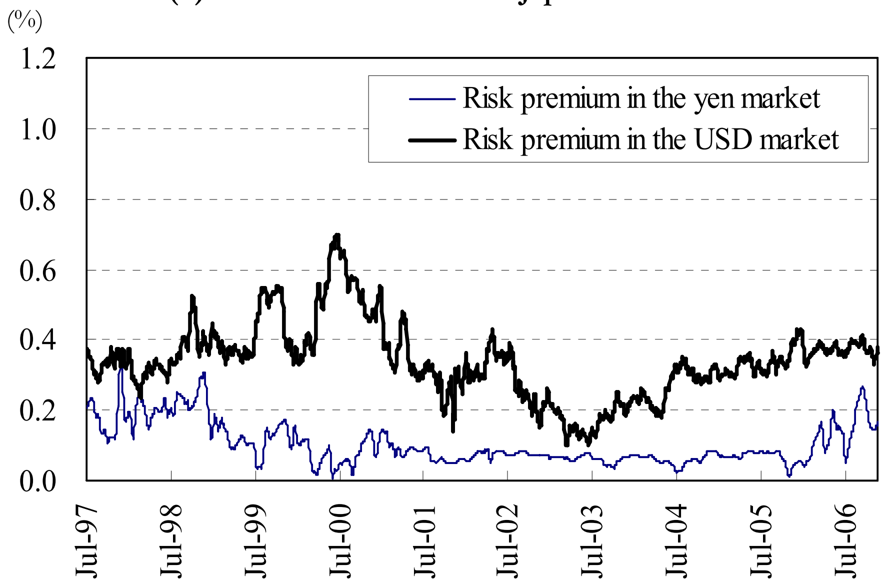
F : FX forward rate (yen/USD)

Figure 2: Differential Risk Premiums between USD and Yen Markets

(i) Risk Premium for Japanese Banks



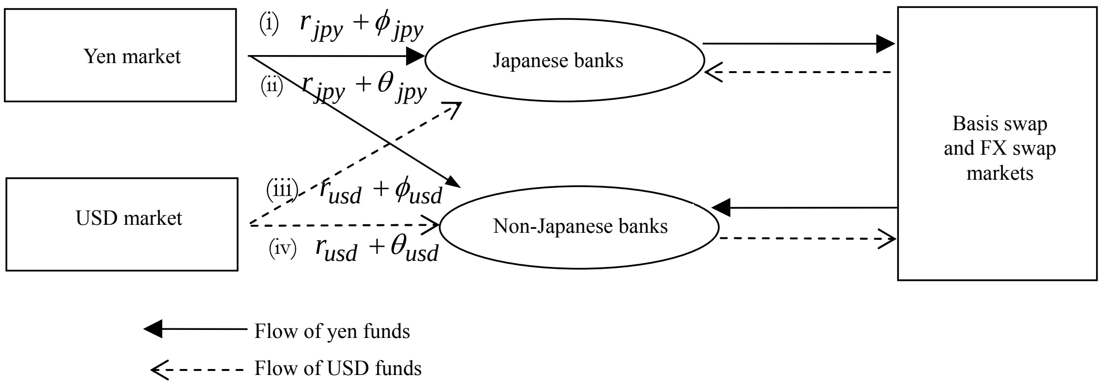
(ii) Risk Premiums for Non-Japanese Banks



Notes: 1. Risk premiums for Japanese banks in the USD/yen markets are calculated as 1-year USD-TIBOR/yen-TIBOR minus yields on 1-year U.S./Japanese government bonds. Risk premiums for non-Japanese banks in the USD/yen markets are calculated as 1-year USD-LIBOR/yen-LIBOR minus yields on 1-year U.S./Japanese government bonds.
2. The data shown above is the 10-day moving average of the original data.

Source: Bloomberg

Figure 3: Typical Funding Structure of Japanese and Non-Japanese Banks

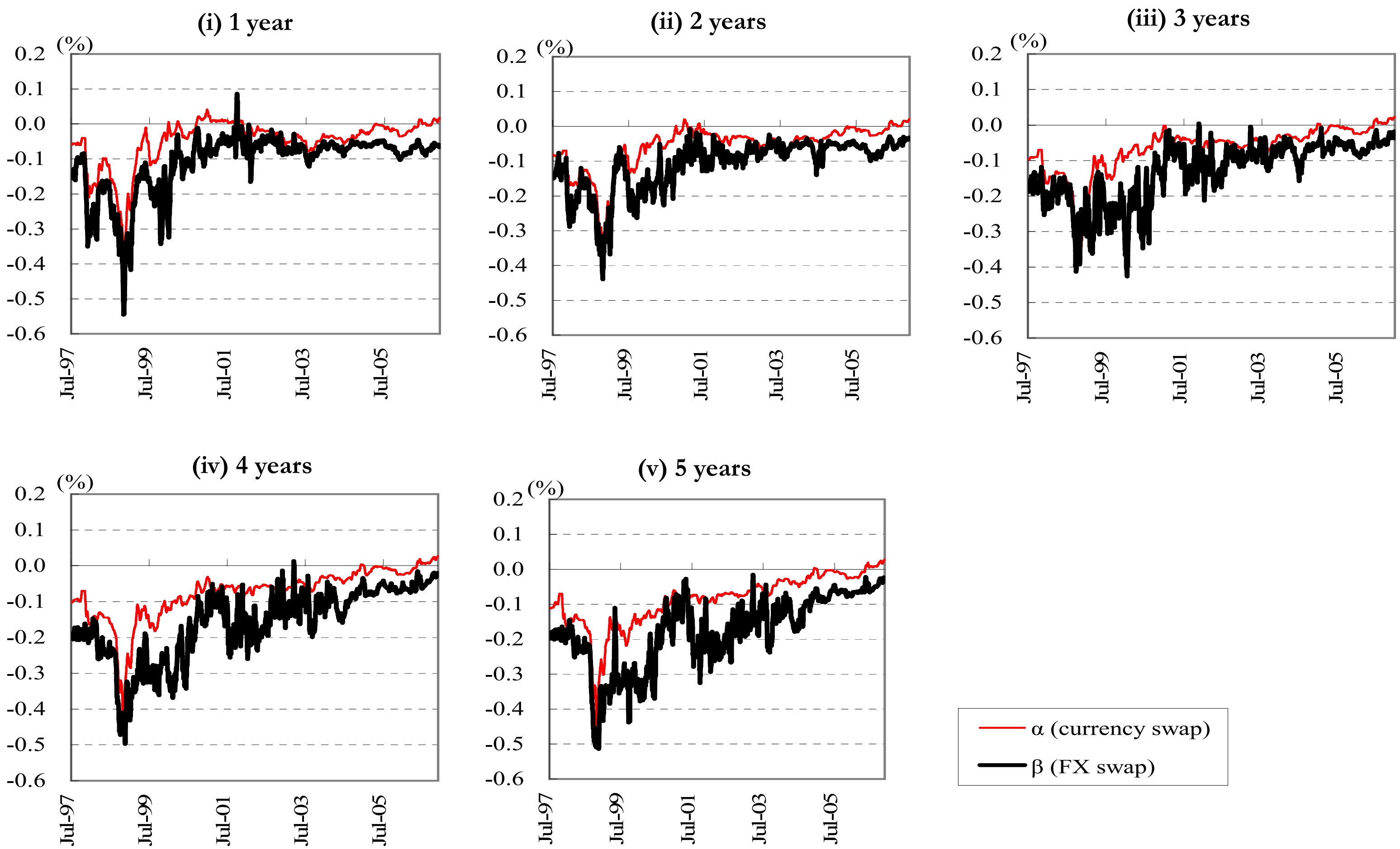


r_{jpy} / r_{usd} : Risk-free yen/USD interest rate

ϕ_{jpy} / ϕ_{usd} : Risk premium for Japanese bank in the yen/USD market

$\theta_{jpy} / \theta_{usd}$: Risk premium for non-Japanese bank in the yen/USD market

Figure 4: Time-Series Movement of α and β



Note: The data shown above is the 10-day moving average of the original data.

Source: Bloomberg

Figure 5: Efficient Prices Extracted from Each Structural Model

(i) Baseline Model

(ii) Partial Adjustment Model

(iii) Under/Overreaction Model

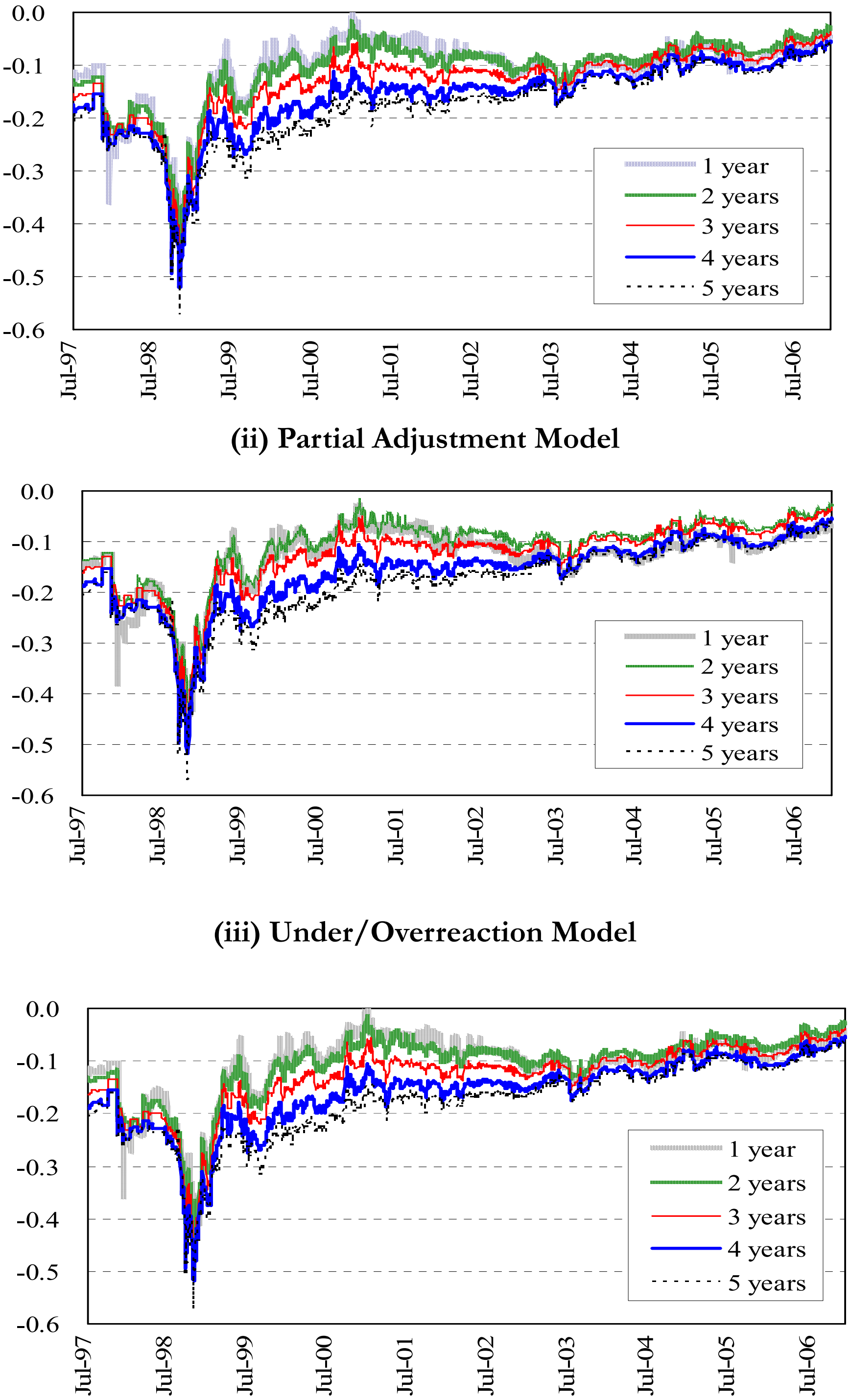


Table 1: Summary Statistics

Sample Period (daily) : July 1, 1997 to November 30, 2006 (Number of Observations: 2,458)

Maturity		Mean	Std. Dev.	Maximum	Minimum
1 year	α	-0.048	0.066	0.060	-0.380
	β	-0.109	0.095	0.308	-0.775
2 year	α	-0.058	0.066	0.038	-0.405
	β	-0.109	0.086	0.202	-0.588
3 year	α	-0.066	0.067	0.028	-0.420
	β	-0.131	0.109	0.336	-0.667
4 year	α	-0.076	0.071	0.031	-0.435
	β	-0.160	0.116	0.405	-0.688
5 year	α	-0.087	0.076	0.033	-0.480
	β	-0.178	0.123	0.143	-0.852

Table 2: Unit Root Test

Sample Period (daily) : July 1, 1997 to November 30, 2006 (Number of Observations: 2,458)

		Augmented Dickey-Fuller Test	
Maturity		Level	1st difference
1 year	α	-2.368	-56.066**
	β	-2.812	-73.151**
2 years	α	-1.744	-24.887**
	β	-2.522	-75.445**
3 years	α	-1.950	-12.140**
	β	-2.960*	-75.989**
4 years	α	-1.683	-23.965**
	β	-1.995	-75.472**
5 years	α	-1.636	-24.299**
	β	-2.136	-76.084**

Notes: 1. * and ** denote significance at the 5% and 1% level, respectively. Test statistics are based on the specification with a constant term.
2. The number of lags is chosen by the modified AIC recommended by Ng and Perron [2001].

Table 3: Baseline Model

Sample Period (daily) : July 1, 1997 to November 30, 2006 (Number of Observations: 2,458)

	i	1 year	2 years	3 years	4 years	5 years
$\ln \sigma_i^2$	s_α	-11.393**	-11.299**	-11.581**	-11.675**	-11.689**
		[0.053]	[0.048]	[0.031]	[0.018]	[0.016]
	s_β	-5.637**	-5.660**	-4.874**	-5.070**	-4.949**
		[0.003]	[0.018]	[0.012]	[0.018]	[0.018]
	ξ	-9.830**	-10.345**	-10.380**	-10.356**	-10.287**
		[0.002]	[0.026]	[0.016]	[0.011]	[0.010]
Constant		0.061**	0.052**	0.065**	0.084**	0.091**
		[0.001]	[0.001]	[0.002]	[0.002]	[0.002]
Log likelihood		11,659	12,178	11,299	11,524	11,225
$\mathrm{SIS}_\alpha^{(1)}$		0.827	0.722	0.769	0.789	0.803
$\mathrm{SIS}_\beta^{(1)}$		0.015	0.009	0.004	0.005	0.005
$\mathrm{SIS}_\alpha^{(1)} - \mathrm{SIS}_\beta^{(1)}$		0.812**	0.713**	0.765**	0.784**	0.798**
		[0.010]	[0.014]	[0.008]	[0.004]	[0.004]

Notes: 1. The figures in square brackets are standard errors. * and ** denote significance at the 5% and 1% level, respectively.
2. The Wald test is conducted to test the difference between SIS_α and SIS_β.
3. Constant denotes the constant term in α equation.

Table 4: Partial Adjustment Model

Sample Period (daily) : July 1, 1997 to November 30, 2006 (Number of Observations: 2,458)

	i	1 year	2 years	3 years	4 years	5 years
C_i	α	0.929**	1.025**	0.962**	1.018**	1.000**
		[0.015]	[0.052]	[0.076]	[0.066]	[0.004]
	β	0.343**	0.569**	0.592**	0.572**	0.568**
		[0.004]	[0.012]	[0.006]	[0.014]	[0.013]
$\ln \sigma_i^2$	s_α	-10.965**	-12.403**	-12.385**	-12.000**	-12.948**
		[0.034]	[0.497]	[0.623]	[0.509]	[0.001]
	s_β	-6.290**	-5.921**	-5.721**	-5.281**	-5.147**
		[0.008]	[0.001]	[0.008]	[0.017]	[0.019]
	ξ	-9.372**	-10.451**	-10.468**	-10.256**	-10.217**
		[0.015]	[0.024]	[0.018]	[0.024]	[0.011]
Constant		0.078**	0.052**	0.058**	0.085**	0.091**
		[0.003]	[0.003]	[0.005]	[0.006]	[0.003]
Log likelihood		12,106	12,352	11,153	11,769	11,428
$\mathrm{SIS}_\alpha^{(2)}$		0.831	0.876	0.872	0.851	0.939
$\mathrm{SIS}_\beta^{(2)}$		0.044	0.011	0.009	0.007	0.006
$\mathrm{SIS}_\alpha^{(2)} - \mathrm{SIS}_\beta^{(2)}$		0.787**	0.865**	0.863**	0.844**	0.933**
		[0.003]	[0.056]	[0.071]	[0.067]	[0.001]

Notes: 1. The figures in square brackets are standard errors. * and ** denote significance at the 5% and 1% level, respectively.
2. The Wald test is conducted to test the difference between SIS_α and SIS_β.
3. Constant denotes the constant term in α equation.

Table 5: Under/Overreaction Model

Sample Period (daily) : July 1, 1997 to November 30, 2006 (Number of Observations: 2,458)

	i	1 year	2 years	3 years	4 years	5 years
$D_i \times 10^5$	α	-0.001	-0.001	0.000	0.043	0.011
		[0.002]	[0.020]	[0.037]	[0.170]	[1.618]
	β	-99.940**	-45.212**	-152.7**	-48.636**	-62.091**
		[0.200]	[5.896]	[0.001]	[0.779]	[6.582]
$\ln \sigma_i^2$	s_α	-11.436**	-11.314**	-11.736**	-11.710**	-11.385**
		[0.056]	[0.001]	[0.032]	[0.038]	[0.035]
	s_β	-5.654**	-5.709**	-4.870**	-5.071**	-4.957**
		[0.001]	[0.017]	[0.002]	[0.002]	[0.002]
	ξ	-9.805**	-10.412**	-10.395**	-10.272**	-10.237**
		[0.015]	[0.015]	[0.015]	[0.016]	[0.016]
Constant		0.061**	0.052**	0.065**	0.084**	0.091**
		[0.003]	[0.001]	[0.002]	[0.002]	[0.002]
Log likelihood		11,657	12,180	11,300	11,524	11,227
$\mathrm{SIS}_\alpha^{(3)}$		0.836	0.711	0.793	0.809	0.810
$\mathrm{SIS}_\beta^{(3)}$		0.016	0.009	0.004	0.005	0.005
$\mathrm{SIS}_\alpha^{(3)} - \mathrm{SIS}_\beta^{(3)}$		0.821**	0.702**	0.789**	0.803**	0.805**
		[0.010]	[0.003]	[0.007]	[0.008]	[0.007]

Notes: 1. The figures in square brackets are standard errors. * and ** denote significance at the 5% and 1% level, respectively.
2. The Wald test is conducted to test the difference between SIS_α and SIS_β.
3. Constant denotes the constant term in α equation.

Table 6: Reduced-Form Model

(i) Johansen Cointegration Test

Sample Period (daily) : July 1, 1997 to November 30, 2006 (Number of Observations: 2,458)

	Number of Cointegrating Vectors			
	Trace Test		Maximum Eigenvalue Test	
Maturity	None	At Most 1	None	At Most 1
1 year	22.25*	5.49	16.75*	5.49
2 years	24.13*	4.22	19.90*	4.22
3 years	20.71*	4.18	16.53*	4.18
4 years	45.01**	3.11	41.90**	3.11
5 years	25.11*	2.39	22.72**	2.39

(ii) Price Discovery Measures

	PT	IS	
Maturity		Lower	Upper
1 year	0.97	0.94	0.97
2 years	0.96	0.88	0.93
3 years	0.98	0.93	0.95
4 years	0.99	0.99	0.99
5 years	0.99	0.99	0.99

Notes: 1. * and ** denote significance at the 5% and 1% level, respectively. The test of restriction on cointegrating vector is conducted by the LR (Likelihood Ratio) test.
2. The number of lags is chosen by AIC.