Bank Credit Risk, Common Factors, and Interdependence of Credit Risk in Money Markets:

Observed vs. Fundamental Prices of Bank Credit Risk

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Abstract

This paper empirically reexamines the role of TIBOR/LIBOR as indicators of bank credit risk and investigates the interdependence of bank credit risk in money markets within and across the border since the 1990s. Empirical results are summarized as follows. (i) Observed risk premiums constructed from TIBOR/LIBOR contain two common factors, global and currency factors, which explain most of the variance of the risk premiums; (ii) thus the generalized impulse response of risk premiums from the shocks of the same currency markets are much larger than the responses from then shocks of the same bank groups; and (iii) the conditional correlations, derived from a Multivariate GARCH model, of the same bank groups' risk premiums between the yen and dollar markets fluctuate around zero, while the correlations between Japanese and foreign banks' risk premiums in the same currency market are very high; (iv) after controlling for these common factors, we successfully derived the fundamental prices of bank credit risk both particularly for Japanese banks using a state space model; (iv) these fundamental prices show plausible time-series properties such as a high degree of impulse response from the shocks of the same bank groups, and a high conditional correlation of the same bank groups' credit risk prices between the yen and dollar two markets; (v) however, the fundamental prices account for only a tiny portion of the total variance of risk premiums.

JEL Classification: E43, G14, G15

Key Words: LIBOR, TIBOR, Credit Risk, Factor Analysis, State Space Model, Kalman Filter, Cointegration, , Generalized Impulse Response, Multivariate GARCH

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I. Introduction

This paper aims to empirically reexamine the role of TIBOR/LIBOR[1] as indicators of bank credit risk as well as investigate the interdependence of bank credit risk in money markets within and across the border. Our main objective lies in shedding light on how money market interest rates functioned as a price discovery tool for bank credit risk since the 1990s.

Credit risks of Japanese and foreign banks are expected to be priced in TIBOR/LIBOR since the majority of referenced banks for TIBOR/LIBOR are Japanese/foreign banks, respectively. Indeed, the so-called "Japan premium", generally defined as the difference between U.S. dollar-TIBOR and LIBOR, rose sharply to nearly 100 bps at the height of the Japanese banking crisis in 1997-98. The Japan premium was considered to reflect western banks' skepticism on opaque Japanese accounting and banking supervision system. Around 2001 to 2002, the vulnerability of Japanese banks became highlighted again mainly due to their low earnings and newly emerging nonperforming loans. This time, however, the Japan premium did not appear. Ito and Harada [2004] assert that the Japan premium lost its role of showing market perception about the vulnerability of Japanese banks.

Specifically, in this paper, we attempt to extract the fundamental credit risk prices for Japanese and foreign banks from the observed risk premiums constructed from the daily yen- and U.S. dollar- TIBOR/LIBOR during the period from 1992/8/3 to 2005/2/2. In this period ,the Japanese banking system experienced unprecedented situations including (a) the instability arising from the non-performing loan problem in the late 1990s and (b) a ultra low interest rate environment put in force by the Bank of Japan (BOJ) under the name of the Quantitative Monetary Easing Policy (QMEP) since March 2001.

[1] TIBOR and LIBOR are abbreviations for Tokyo Interbank Offered Rate and London Interbank Offered Rate, respectively.

During the period of financial instability, which was initiated by a series of failure of Japanese major financial institutions in late 1997, Hanajiri [1999] argues that the arbitrage relationship collapsed between the yen and the U.S. dollar cash markets.[2] Also, under the QMEP, Baba et al. [2005a,b] argue that Japanese money markets almost ceased to function as a pricing mechanism of banks' creditworthiness in that money market interest rates have been so lowered that they hardly reflect the differences in credit risk among individual banks.

There are few studies that investigated the relationship between TIBOR and LIBOR with the notable exceptions of Covrig, Low, and Melvin [2004], Peek and Rosengren [2001], and Nishioka and Baba [2004].[3] Covrig, Low, and Melvin [2004] investigated the determinants of Japan premium and concluded that lower Japanese interest rates, a flatter yield curve, and a decline in stock prices raised the Japan premium. Peek and Rosengren [2001] attributed the Japan premium to Japanese banks' financial soundness, net worth, and default risk. Also, Nishioka and Baba [2004] discussed the equilibrium relationship among yen- and U.S. dollar-TIBOR and LIBOR to explore the cause of the "negative" nominal yen funding costs for foreign banks in the FX swap market.[4] We aim to add another line of research to these studies by rigorously analyzing the fundamental prices of bank credit risk included in yen- and U.S. dollar TIBOR/LIBOR and the interdependence structure among these fundamental risk prices.

To extract the fundamental credit risk prices for Japanese and foreign banks from the

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[2] In November 1997, concern over the financial stability heightened following a series of failures of four financial institutions: Sanyo Securities (November 3), Hokkaido Takushoku Bank (November 17), Yamaichi Securities (November 24), and Tokuyo City Bank (November 26). The concern over the financial instability subsided after the nationalization of Long-Term Credit Bank of Japan (October 23, 1998) and Nippon Credit Bank (December 13, 1998).

[3] In addition, Lo, Fung, and Morse [1995] investigated the relationship between yen-LIBOR and yen interest rate on negotiable certificate of deposits (NCDs).

[4] It is around 1995 when we first observed negative yen funding costs in the FX swap market. In the periods of financial instability and the QMEP, we frequently witnessed negative FX yen funding costs. They have sometimes yielded "negative nominal uncollaterarized call rate" particularly under the QMEP. See Nishioka and Baba [2004] and Baba et al. [2005a] for more details.

observed risk premiums, we first find some common factors that do not reflect creditworthiness of banks and thus pay a role of control variables using factor analysis. Recent empirical studies on the U.S. credit spreads show that the structural (fundamental) factors specific to each referenced entity are important, but can explain only a small portion of the credit spreads. For instance, Collin-Dufresne, et al. [2001] show that structural factors can explain only a quarter of the changes in the U.S. credit spreads, indicating that systematic factors, common to the aggregate corporate bond market, have much more contribution to the changes in credit spreads.[5] Thus, we conjecture that major part of the risk premium variation in short-term money markets are also likely to be accounted for by some common factors as in the case of the U.S. credit spreads. And then, using these common factors as control variables, we extract the fundamental prices of bank credit risk based on the state space model in which shadow prices of bank credit risk govern the fundamental prices of credit risk for Japanese and foreign banks, respectively.

After estimating the fundamental prices of bank credit risk, we investigate their time-series properties to explore the dynamic interdependence structure of bank credit risk within and across the border, using the interdependence structure of the observed risk premiums as a benchmark for comparison. The basic methodologies we use are (a) Johansen's [1991, 1995] conintegration analysis, (b) VAR (Vector Autoregressive) Model or VECM (Vector Error Correction Mode)-based Granger causality test and the generalized impulse response function, and (c) M-GARCH (Multivariate Generalized Autoregressive Conditional Heteroscedasticity) models.[6] Main objective in this part is to clarify the difference between the fundamental prices of bank credit

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[5] Elton, et al. [2001] also show that systematic risk of the equity market is more important determinants of the U.S. credit spreads than expected default loss and tax premium. On the other hand, Driessen [2004] decomposes the term structure of credit spreads, finding a similar result.

[6] The univariate ARCH and GARCH models were developed by Engle [1982] and Bollerslev [1986], respectively. The univariate GARCH model was extended to a multivariate framework by Bollerslev, Engle, and Wooldrige [1988]. Using M-GARCH models, King et al. [1994] analyze the volatility transmission between national stock markets, while Kearney and Patton [2000] investigate the volatility transmission in the EMS. Also, Kim et al. [2005] analyze the volatility transmission between stock and bond markets in the EMS.

risk and the observed risk premiums in terms of the dynamic interdependence structure.

The rest of the paper is organized as follow. Section II derives the equilibrium relationships among money market interest rates based on the foreign currency funding structure of Japanese and foreign banks including the FX swap market. Section III describes risk premium data we use in this paper. Section IV briefly explains the overall empirical strategy and the methodologies adopted in this paper to decompose the risk premiums and extract the fundamental credit risk prices as well as analyze their time-series properties. Section V reports and discusses the empirical results. Section VI concludes the paper.

II. Theoretical Relationships Linking Money Market Interest Rates

(i) Foreign Currency Funding Structure of Japanese and Foreign Banks

Following Nishioka and Baba [2004], we show that active arbitrage transactions in the FX swap markets create a transmission channel of risk premiums for Japanese and foreign banks between the yen and U.S. dollar markets. Specifically, we consider the no-arbitrage conditions for Japanese and foreign banks' foreign currency funding costs. The following three markets are under study: (a) the yen cash market, (b) the U.S. dollar cash market, and (c) the FX swap market.[7]

As shown in Figure 1, FX swap transaction has been active since the early 1990s except the period of financial instability from late 1997 to 1998. The FX swap transaction plays a role of a funding source of foreign currencies for both Japanese and foreign banks, alternative to the direct funding from cash markets. Thus, active FX swap transaction creates two no-arbitrage conditions for yen and U.S. dollar funding, which in turn creates an equilibrium condition linking four risk

[7] A typical FX swap transaction is a contract in which Japanese banks borrow U.S. dollars from, and lend yen to, foreign banks at the same time.

premiums: yen- and U.S. dollar risk premiums for both Japanese and foreign banks.

(ii) No-arbitrage and Equilibrium Conditions

The funding costs in the cash markets can be written as the sum of the risk-free interest rate and the risk premium for Japanese or foreign banks. Let i and i^* denote the yen and dollar risk-free interest rates, JY and JD the risk premiums for Japanese banks in the yen and dollar market, and FY and FD the risk premiums for foreign banks in the yen and dollar market, respectively. Also, let F and S denote the yen-dollar forward and spot rate of foreign exchange.

As shown in Figure 2, Japanese banks have two alternative funding sources of dollars: (a) raising dollars directly from the dollar market, and (b) raising yen from the yen market and exchanging it for dollars in the FX swap market. Then, if these funding sources are perfect substitutes for Japanese banks, the following no-arbitrage condition holds

$$1+i^*+JD=\frac{S}{F}(1+i+JY). \tag{1}$$

The left-hand side of equation (1) is the dollar interest rate for Japanese banks, while the right-hand side is the dollar funding cost for Japanese banks in the FX swap market.

Similarly, foreign banks have two alternative funding sources of yen: (a) raising yen directly from the yen market, and (b) raising dollars raised from the dollar market and exchanging those for yen in the FX swap market. Then, if these two funding sources are perfect substitutes for foreign banks, the following no-arbitrage condition holds

$$1+i+FY=\frac{F}{S}(1+i^*+FD). \tag{2}$$

The left-hand side of equation (2) is the yen interest rate for foreign banks, while the right-hand side is the yen funding cost for foreign banks in the FX swap market. Substituting equation (1) into equation (2) yields the following equilibrium condition:

$$\frac{1+i+JY}{1+i+FY}=\frac{1+i^{*}+JD}{1+i^{*}+FD}. \tag{3}$$

Equation (3) creates a transmission channel for interest rates within and across the border. Approximation of equation (3) enables us to find its significance more intuitively:

$$JY-JD=FY-FD. \tag{4}$$

The left-hand side of equation (4) shows the difference in risk premiums for foreign banks between the yen and dollar markets, while the right-hand side shows the difference in risk premiums for foreign banks between the two markets. The significance of this result is that to achieve equilibrium, we do not need the "parity" of risk premiums for the same bank groups between the yen and dollar markets. In section V, we explore the relationships among these four variables with due attention to the equilibrium condition (4).[8]

III. Data

(i) Data Description

In this paper, we use 90-day yen- and dollar-LIBOR (London Interbank Offered Rate) and TIBOR (Tokyo Interbank Offered Rate) to construct risk premiums for Japanese and foreign banks. Of 16 referenced banks that comprise yen-TIBOR/LIBOR, 14 banks are Japanese banks in yen-TIBOR,

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[8] Another interesting extension of equilibrium condition (3) is to decompose the FX swap yen funding cost for foreign banks, which frequently have moved below zero under the QMEP. See Appendix 1 for more details. Nishioka and Baba [2004] show that this negative FX swap yen funding cost is closely linked to negative nominal money market interest rates that have been observed in Japan since 2001.

while 11 banks are foreign banks in yen-LIBOR.[9] In a similar fashion, of 10 banks referenced by dollar-TIBOR, 8 banks are Japanese banks, while of 16 banks referenced by dollar-LIBOR, 14 banks are foreign banks. Appendix 2 provides more details of the data.

While LIBOR forms the pricing basis for floating rate securities and loans settled during European trading hours, Asia-Pacific issuers or borrowers need settlement during Asia-Pacific trading hours to avoid interest rate risk. TIBOR forms the basis for such settlement. It should be noted, however, that LIBOR is quoted at 11 am London time, while TIBOR is quoted at 11 am Tokyo time. Since the 11 am London time corresponds to 7 or 8 pm Tokyo time, LIBOR reflects the market events that occurred in Japan's afternoon. To accommodate this time difference, Covrig, Low, and Melvin [2004] use the same day quotes for TIBOR and the one-day lag quotes for LIBOR in investigating the determinants of "Japan Premium", which they defined as a yen-TIBOR/LIBOR spread. We tried both versions, the same-day quotes for TIBOR/LIBOR and one-day lag for LIBOR, but no distinct differences were found in estimation results. Thus, in what follows, we report only the results using the same-day quotes.

As risk-free interest rates, we use Japanese and the U.S. Treasury bill rates. Thus, JY/FY that appeared in section II are computed as yen-TIBOR/LIBOR minus Japanese Treasury bill rate, and JD/FD are computed as dollar-TIBOR/LIBOR minus the U.S. Treasury bill rate. Figure 3 shows the risk premiums for Japanese and foreign banks thus constructed. An interesting point to note here is that the dollar risk premiums are almost always higher than the yen risk premiums irrespective of Japanese and foreign banks and the differences are pronounced in the period of financial instability from 1997 to 1998. As shown by equation (4), we do not need the equality of

[9] The relative impreciseness of yen-LIBOR as a proxy for foreign banks' yen interest rate resulted in a poor performance of the extracted yen-market fundamental price of foreign banks' credit risk. We will discuss this issue in session V.

risk premiums for the same bank groups between the yen and dollar markets to attain equilibrium and a casual observation suggests that equilibrium condition (4) holds over the sample period except the period of financial instability around 1998.

(ii) Statistical Properties of Risk Premiums

Table 1 shows summary statistics of the risk premiums we use in our empirical analysis. As shown in Table 1(i), means and standard deviations of the yen risk premiums are much smaller than those of the dollar risk premiums. Also, all of the risk premiums have positive skewness and excess kurtosis, which can be jointly confirmed by the Jarque-Bera test. Positive skewness of risk premiums implies that the total return including the capital gain/loss has negative skewness, fixing the underlying risk-free rates, which is consistent with the notion of default risk.[10] And, the high degree of kurtosis suggests a fat-tailed property of the risk premiums. In addition, we tested for serial correlations of both the level of the variables themselves and squared ones up to the 12th order using the Ljung-Box Q test denoted LB(12) and $LB^2(12)$, respectively. Both statistics show a very high degree of serial correlations. These properties of the risk premiums support the use of the GARCH models particularly with the multivariate Student t distribution.

On the other hand, Table 1(ii) reports correlation matrix between each pair of risk premiums. A noteworthy point here is that the correlations between JY and FY (JD and FD) are higher than the correlations between JY and JD (FY and FD). That is, the correlations of risk premiums for the same bank groups between the yen and dollar markets are lower than the correlations of risk premiums in the same currency market between Japanese and foreign banks.

[10] The possibility of extreme negative returns on credit instruments in the case of a credit event creates negatively-skewed distributions. See Chapter 13 of Duffie and Singleton [2002] for more details.

Also, note that correlations between JY and FD (JD and FY) are higher than the correlation between FY and FD. Since the pairs of JY and FD (JD and FY) do not have common attributes in terms of the referenced bank groups and the denominated currency, we can infer that some common factors rather than credit fundamentals of referenced bank groups contribute to moving these risk premiums in the same direction.[11] This finding actually motivated us to decompose the risk premiums into common factors and credit risk fundamentals.

Table 2 reports the results of two unit root tests: the Augmented Dickey-Fuller (ADF) test and the Phillips-Perron (PP) test. The test results show that all of the risk premiums are I(0).[12] Thus, we should not use Johansen's cointegration analysis to test for the equilibrium condition among risk premiums shown by equation (4) and instead use the VAR (Vector Autoregressive)-based models to investigate the dynamic time-series properties and interdependence among the risk premiums.

IV. Empirical Strategy and Methodologies

(i) Empirical Strategy

Our empirical strategy goes as follows. First, we attempt to decompose risk premiums to extract fundamental prices of bank credit risk that are specific either to Japanese or foreign banks. To that end, as a preliminary step, we employ factor analysis to derive common factors and then construct a state space model using common factors as control variables. In the state space model, fundamental prices of bank credit risk is linked to state variables that act as noisy shadow prices of credit risk..

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[11] This tendency is pronounced in terms of the conditional correlations derived from an M-GARCH model. See section V for details.

[12] The lag length is determined by the Schwarz criterion as suggested by Hayashi [2000].

Next, we investigate the time-series properties of the derived fundamental prices of bank credit risk using the risk premiums themselves as a benchmark. As for the risk premiums, we (a) conduct the Granger causality test, (b) derive the generalized impulse response function based on the VAR system and (c) estimate an M-GARCH model to investigate dynamic interdependency among the risk premiums. As for the fundamental prices of bank credit risk, we first conduct Johansen's [1991,1995] cointegrating analysis since the fundamental prices of bank credit risk are I(1) by construction. Then, we explore the properties of dynamic interdependency by conducting the Granger causality test and deriving the generalized impulse response function based on the VECM (Vector Error Correction Model) as well as estimating an M-GARCH model.

(ii) Empirical Methodologies

A. Decomposition of Risk Premiums

a. Preliminary Step: Factor Analysis

As a preliminary step, we derive common factors from four risk premiums. Our approach is to use the following traditional orthogonal factor analysis that does not need any *a priori* assumptions about common factors:[13]

$$\mathbf{R}_t = \boldsymbol{\mu} + \mathbf{B}\mathbf{F}_t + \boldsymbol{\varepsilon}_t , \qquad (5)$$

where $\mathbf{R}_t$ denotes a vector of four risk premiums $\left[JY_t, JD_t, FY_t, FD_t\right]'$, $\mathbf{F}_t$ a vector of K

[13] Driessen, Melenberg, and Nijman [2003] successfully found a linear factor model with five factors that explains 96.5% of the variation of international bond returns using factor analysis. An alternative approach is to assume factor structure in advance. In the context of international bond returns, Barr and Priestley [2004] assume that the world bond index and individual local market bond index are common factors, for instance. In the case of TIBOR/LIBOR, however, such indices do not exist, so we use factor analysis to derive common factors.

common factors to be estimated, $\mathrm{cov}(\mathbf{F}_t, \boldsymbol{\varepsilon}_s) = 0$ for all t and s, $\mathrm{E}[\mathbf{F}_t] = \mathrm{E}[\boldsymbol{\varepsilon}_t] = 0$, $\mathrm{var}(\mathbf{F}_t) = \mathbf{I}_K$, and $\mathrm{var}(\boldsymbol{\varepsilon}_t) = \mathbf{G}$. Here, $\mathbf{G}$ is a diagonal matrix with σ_i^2 along the diagonal. Then, the covariance matrix of risk premiums $\boldsymbol{\Omega}$ can be decomposed as

$$\boldsymbol{\Omega} = \mathbf{BB}' + \mathbf{G}.$$

We use principal factor method with no rotation to extract common factors.[14] As discussed in section V, factor analysis above successfully extracted two relevant common factors: (a) the global factor denoted Fg_t, which is almost equally common to all of the risk premiums, and (b) the currency factor denoted Fc_t, which captures the difference between the yen and dollar markets such that when $Fc_t > 0$, it raises/lowers risk premiums in the dollar/yen market irrespective of the bank groups.

b. Decomposition Framework

Using the common factors derived by factor analysis above, in what follows, we explain the methodology to extract fundamental prices of bank credit risk from risk premium data. Following Blanco, Brennan, and Marsh [2005], suppose that the unobservable shadow prices of credit risk for Japanese banks J_t^*, and foreign banks F_t^*, follows a random walk process, respectively:[15]

$$J_t^* = J_{t-1}^* + e_t^j, \qquad \text{and} \qquad F_t^* = F_{t-1}^* + e_t^f, \tag{6}$$

where e_t^j and e_t^f are the noises with zero mean and constant variance. We assume that the observed risk premiums, JY_t, JD_t, FY_t, and FD_t are equal to the sum of (a) the fundamental

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[14] Researchers often rotate the initial solution for ease of interpretation, but the rotation entails arbitrariness. To avoid such arbitrariness, we use the initial solution obtained by the principal factor method. See Chan, Karceski, and Lakonishok [1998] for a review of factor analysis and factor models.

[15] Blanco, Brennan, and Marsh [2004] investigate the relationship between corporate bond yields and CDS (credit default swap) spreads.

price of bank credit risk in each market, denoted $JYPRICE_t$, $JDPRICE_t$, $FYPRICE_t$, $FDPRICE_t$, respectively, (ii) non-transient common factors, Fg_t and Fc_t, and (iii) stochastic terms including transient microstructural noises, e_t^{jy}, e_t^{jd}, e_t^{fy} and e_t^{fd}. The structure is summarized as follows:

Japanese banks:
$$JY_t = JYPRICE_t + a_t^g Fg_t + a_t^c Fc_t + e_t^{jy} \quad (7)$$

$$JD_t = JDPRICE_t + b_t^g Fg_t + b_t^c Fc_t + e_t^{jd} \quad (8)$$

Foreign banks:
$$FY_t = FYPRICE_t + c_t^g Fg_t + c_t^c Fc_t + e_t^{fy} \quad (9)$$

$$FD_t = FDPRICE_t + d_t^g Fg_t + d_t^c Fc_t + e_t^{fd}, \quad (10)$$

where
$$JYPRICE_t = a_0^{jy} + a_1^{jy} J_t^* + e_t^{jyprice}, \quad JDPRICE_t = a_0^{jd} + a_1^{jd} J_t^* + e_t^{jdprice}$$

$$FYPRICE_t = a_0^{fy} + a_1^{fy} F_t^* + e_t^{fyprice}, \quad FDPRICE_t = a_0^{fd} + a_1^{fd} F_t^* + e_t^{fdprice}.$$

Here, coefficients of the sensitivity to each common factor are allowed to move over time, and the fundamental prices of bank credit risk in each market are assumed to be linearly linked to the fundamental prices of credit risk for Japanese and foreign banks with noises.

c. State Space Model

To express the above decomposition framework and estimate each parameter and fundamental price of bank credit risk, we construct the following state space model. In this model, two random-walk state variables are assumed to govern the fundamental prices of bank credit risk after controlling for the effects of common factors derived by factor analysis, Fg_t and Fc_t:[16]

$$\begin{bmatrix} JY \\ JD \\ FY \\ FD \end{bmatrix}_t = \begin{bmatrix} JYPRICE \\ JDPRICE \\ FYPRICE \\ FDPRICE \end{bmatrix}_t + \begin{bmatrix} s4 & s6 \\ s5 & s7 \\ s14 & s16 \\ s15 & s17 \end{bmatrix}_t \begin{bmatrix} Fg \\ Fc \end{bmatrix}_t + \begin{bmatrix} e1 \\ e2 \\ e11 \\ e12 \end{bmatrix}_t, \tag{11}$$

where

$$\begin{bmatrix} JYPRICE \\ JDPRICE \end{bmatrix}_t = \begin{bmatrix} c3 \\ c4 \end{bmatrix} + \begin{bmatrix} c1 \\ c2 \end{bmatrix} s3_t + \begin{bmatrix} e3 \\ e4 \end{bmatrix}, \quad \begin{bmatrix} FYPRICE \\ FDPRICE \end{bmatrix}_t = \begin{bmatrix} c13 \\ c14 \end{bmatrix} + \begin{bmatrix} c11 \\ c12 \end{bmatrix} s13_t + \begin{bmatrix} e13 \\ e14 \end{bmatrix},$$

$$\begin{bmatrix} s3 \\ \vdots \\ s7 \end{bmatrix}_t = \begin{bmatrix} s3 \\ \vdots \\ s7 \end{bmatrix}_{t-1} + \begin{bmatrix} e5 \\ \vdots \\ e9 \end{bmatrix}_t, \text{ and } \begin{bmatrix} s13 \\ \vdots \\ s17 \end{bmatrix}_t = \begin{bmatrix} s13 \\ \vdots \\ s17 \end{bmatrix}_{t-1} + \begin{bmatrix} e15 \\ \vdots \\ e19 \end{bmatrix}_t.$$

Here, s with numbers denote state variables, c constant coefficients, and e Gaussian noises. Among these state variables, $s3_t$ and $s13_t$ correspond to the random-walk shadow prices J_t^* and F_t^* in equation (6) that govern the fundamental prices of credit risk for Japanese and foreign banks, JYPRICE/JDPRICE and FYPRICE/FDPRICE, respectively. Note, here, that we allow for the difference in fundamental prices of credit risk priced between the yen and dollar markets. This is mainly due to the differences in risk averseness of market participants, which is related to $c1/c2$ and $c11/c12$. We can test the differences in these constant terms using the Wald test. Also note that we allow for time-varying sensitivities to common factors and (constant) covariances across relevant equations.[17] In estimating the model, we use the Kalman filter that is a recursive algorithm for sequentially updating the one-step ahead estimate of the state variables given new information.[18] Marquardt method is used as an optimization algorithm.

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[16] The model can be regarded as an extension of "state space representation of the local level model" in which the observed asset price is assumed to be the sum of a ransom walk fundamental component and a Gaussian error term. See Durbin and Koopman [2001] for details. Extensive surveys of applications of state space models in econometrics are found in Chapter 13 in Hamilton [1994] and Chapters 3 and 4 in Harvey [1989].

[17] See Table 4 for the assumed covariance structure.

[18] We initialize the states and variances using priors and adopt the maximum likelihood estimation techniques

B. Analysis of Time-Series Properties

a. Johansen's Cointegration Test

Since the fundamental prices of bank credit risk are I(1) by construction, we use Johansen's [1991, 1995] cointegration test to investigate the long-term relationships among those prices. Let $\mathbf{R}_t$ denote a vector that includes p non-stationary time series (p =4 in our case), all of which have a property of I(1).[19] Suppose the following VAR (vector autoregression) representation of $\mathbf{R}_t$:[20]

$$\mathbf{R}_t = \mathbf{a}_1\mathbf{R}_{t-1} + \mathbf{a}_2\mathbf{R}_{t-2} + \ldots + \mathbf{a}_k\mathbf{R}_{t-k} + \boldsymbol{\varepsilon}_t = \sum_{i=1}^{k}\mathbf{a}_{t-i}\mathbf{R}_{t-i} + \boldsymbol{\varepsilon}_t\,, \tag{12}$$

where $\mathbf{a}_i$ is a coefficient matrix and $\boldsymbol{\varepsilon}_t$ is a error vector. Equation (12) can be rewritten as a VECM (Vector Error Correction Model):

$$\Delta\mathbf{R}_t = \boldsymbol{\Pi}\mathbf{R}_{t-1} + \sum_{i=1}^{k-1}\boldsymbol{\Gamma}_i\Delta\mathbf{R}_{t-i} + \boldsymbol{\varepsilon}_t\,, \tag{13}$$

where $\boldsymbol{\Pi} = \sum_{i=1}^{k}\mathbf{A}_i - \mathbf{I}$ and $\boldsymbol{\Gamma}_i = -\sum_{j=i+1}^{k}\mathbf{a}_j$.

Granger's representation theorem states that if the coefficient matrix $\boldsymbol{\Pi}$ has reduced rank $r < p$, then there exist $p \times r$ matrices $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ with rank r such that $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}^T$, where $\boldsymbol{\beta}^T\mathbf{R}_t$ is I(0).[21] Here, r is the number of cointegrating relations (cointegrating rank) and each column of $\boldsymbol{\beta}$ is the cointegrating vector. Johansen's method is to estimate the $\boldsymbol{\Pi}$ matrix from an unrestricted VAR and to test whether we can reject the restrictions implied by the reduced rank of $\boldsymbol{\Pi}$. The number of cointegrating relations is determined by the trace statistic and the maximum eigenvalue statistic.

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from Durbin and Koopman [2001].

[19] In fact, yen- and dollar-TIBOR/LIBOR, and the fundamental prices of bank credit risk derived by a state space model are found to be I(1). See Table 2 and session V for details.

[20] For ease of notations, we ignore a constant term and exogenous variables throughout this section.

[21] See Engel and Granger [1987] for details.

b. Stability Tests of Cointegration Relationships

We test for potential structural breaks of cointegrating relationships among fundamental credit risk preces using the rolling test proposed by Pascual [2003] and others, although Hansen and Johansen's [1999] recursive tests are often used in this context. There are two types of recursive tests under the VECM representation. In the "Z-representation", all of the parameters of the VECM are recursively re-estimated over the sample period. On the other hand, in the "R-representation", the short-run parameters $\mathbf{\Gamma}_i$ are fixed to their full sample values and only the long-run (error correction) parameters $\mathbf{\Pi}$ are recursively re-estimated. Thus, in the recursive tests, the sample size increases one-by-one as the relevant parameters are recursively re-estimated.

The recursive tests have one potential shortcoming by nature: the power of the test becomes higher as the sample size for estimation increases, which will bias toward rejection of the null hypothesis of no cointegration. In fact, our test results show this property.[22] To avoid such a bias, Pascual [2003] proposed a rolling test, in which equation (13) and thus the trace and maximum eigenvalue statistics are re-estimated using the same sample size (fixed rolling window).

c. Generalized Impulse Response Function

We use the "generalized" impulse responses proposed by Pesaran and Shin [1998] instead of the impulse responses derived from the usual "orthogonalized" Cholesky decomposition following Sims [1980]. The generalized impulse responses have an advantage in that they are invariant to the order of the variables in the VAR model. Let us briefly describe the method as follows.

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[22] The same tendency is observed in Pascual [2003] and Rangvid [2001]. We also used a residual-based stability test of cointegration relationships proposed by Gregory and Hansen [1996], but did not find meaningful results. We do not report these results.

Under the assumption that $\mathbf{R}_t$ is covariance-stationary, equation (12) can be rewritten as the infinite moving average representation as follows:

$$\mathbf{R}_t = \mathbf{A}_0\boldsymbol{\varepsilon}_t + \mathbf{A}_1\boldsymbol{\varepsilon}_{t-1} + \ldots + \mathbf{A}_\infty\boldsymbol{\varepsilon}_{t-\infty} = \sum_{i=0}^{\infty}\mathbf{A}_i\boldsymbol{\varepsilon}_{t-i}\ , \tag{14}$$

where $\mathbf{A}_i = \mathbf{a}_1\mathbf{A}_{i-1} + \mathbf{a}_2\mathbf{A}_{i-2} + \ldots + \mathbf{a}_k\mathbf{A}_{i-k}$ and $\mathbf{A}_0 = \mathbf{I}_p$. The conventional approach by Sims [1980] is to apply the Cholesky decomposition to $E\left(\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t^T\right) = \boldsymbol{\Sigma}$ such that $\mathbf{Q}\mathbf{Q}^T = \boldsymbol{\Sigma}$ where $\mathbf{Q}$ is a $p \times p$ lower triangular matrix. Hence, $p \times 1$ vector of the orthogonalized impulse response function of a unit shock to the j th equation on $\mathbf{X}_{t+n}$ is given by

$$\boldsymbol{\psi}_j^o(n) = \mathbf{A}_n\mathbf{Q}\mathbf{e}_j\ , \tag{15}$$

where $\mathbf{e}_j$ is a $p \times 1$ selection vector with unity as its j th element and zeros elsewhere.

On the other hand, the approach by Pesaran and Shin [1998] directly uses the following simplified version of the definition of the generalized impulse response function proposed by Koop et al. [1996]:

$$GI_x\left(n, \delta_j, \boldsymbol{\Omega}_{t-1}\right) = \mathrm{E}\left[\mathbf{X}_{t+n}\middle|\boldsymbol{\varepsilon}_{jt} = \delta_j, \boldsymbol{\Omega}_{t-1}\right] - \mathrm{E}\left[\mathbf{X}_{t+n}\middle|\boldsymbol{\Omega}_{t-1}\right], \tag{16}$$

where $\boldsymbol{\Omega}_{t-1}$ denotes the conditioning information set at time t-1. Note here that instead of shocking all the elements of $\boldsymbol{\varepsilon}_t$, as in Koop et al. [1996], Pesaran and Shin [1998] choose to shock only one element and integrate out the effects of other shocks using the historically observed distribution of the errors. Under the assumption that $\boldsymbol{\varepsilon}_t$ follows a multivariate normal distribution, we get the generalized impulse response function as

$$\boldsymbol{\psi}_j^G(n) = \sigma_{jj}^{-\frac{1}{2}}\Sigma\mathbf{e}_j\ , \tag{17}$$

where σ_{jj} is the jj the element of the residual covariance matrix $\boldsymbol{\Sigma}$.

d. Multivariate GARCH Model

We use multivariate GARCH (M-GARCH) models to derive conditional correlations between each pair of variables. The basic structure can be written as

$$\mathbf{R}_t = \sum_{i=1}^{k} \mathbf{a}_{t-i}\mathbf{R}_{t-i} + \boldsymbol{\varepsilon}_t \qquad \boldsymbol{\varepsilon}_t \big| \boldsymbol{\Omega}_{t-1} \sim D(0, \mathbf{H}_t), \tag{18}$$

where we assume that the mean equation can be described by the same VAR (or VECM) system as in equation (12) (or (13)) and the residuals follow a multivariate Student *t* distribution D that can capture the fat-tailed property of each variable.[23,24]

There exist numerous methods of parameterizations of the conditional covariance matrix $\mathbf{H}_t$ in equation (18).[25] The specification we adopt is the BEKK[26] model proposed by Engle and Kroner [1995]. The BEKK model is sufficiently general and guarantees a positive definite conditional covariance matrix. The BEKK (1,1) model is given by [27]

$$\mathbf{H}_t = \mathbf{C}^T\mathbf{C} + \mathbf{A}^T\left(\boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}_{t-1}^T\right)\mathbf{A} + \mathbf{B}^T\mathbf{H}_{t-1}\mathbf{B}, \tag{19}$$

where $$\mathbf{H}_t = \begin{bmatrix} h_{11,t} & h_{12,t} & h_{13,t} & h_{14,t} \\ h_{21,t} & h_{22,t} & h_{23,t} & h_{24,t} \\ h_{31,t} & h_{32,t} & h_{33,t} & h_{34,t} \\ h_{41,t} & h_{42,t} & h_{43,t} & h_{44,t} \end{bmatrix}, \quad \mathbf{C} = \begin{bmatrix} c_{11} & c_{12} & c_{13} & c_{14} \\ 0 & c_{22} & c_{23} & c_{24} \\ 0 & 0 & c_{33} & c_{34} \\ 0 & 0 & 0 & c_{44} \end{bmatrix},$$

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23 We adopt the following two-step estimation strategy: (a) estimate the VAR (ECM) mean system and (b) apply the M-GARCH model to the residuals derived from the VAR system. This treatment is just for securing efficiency of M-GARCH model estimation, which has 42 parameters to be estimated only in the residual-covariance terms.

24 See Cambell, Lo, and Mackinlay [1997] for the relevance of the use of the multivariate Student *t* assumption.

25 We prefer the BEKK model to the so-called "diagonal vec model" proposed by Bollerslev, Engle, and Wooldridge [1988] since the latter model does not guarantee positive definiteness of the conditional covariance matrix. For a survey of ARCH-type models, see Bollerslev, Chou, and Kroner [1992], Bollerslev, Engle, and Nelson [1994], and Pagan [1996]. For a survey of multivariate GARCH models in particular, see Bauwens, Laurent, and Rombouts [2005].

26 BEKK is the acronym for Baba, Engle, Kraft, and Kroner [1990].

27 In practice, GARCH (1,1) specification suffices since it corresponds to ARCH(∞).

$$\mathbf{A} = \begin{bmatrix} a_{11} & a_{12} & a_{13} & a_{14} \\ a_{21} & a_{22} & a_{23} & a_{24} \\ a_{31} & a_{32} & a_{33} & a_{34} \\ a_{41} & a_{42} & a_{43} & a_{44} \end{bmatrix}, \text{ and } \mathbf{B} = \begin{bmatrix} b_{11} & b_{12} & b_{13} & b_{14} \\ b_{21} & b_{22} & b_{23} & b_{24} \\ b_{31} & b_{32} & b_{33} & b_{34} \\ b_{41} & b_{42} & b_{43} & b_{44} \end{bmatrix}.$$

Equation (19) is estimated under the assumption that the residuals follow the following multivariate Student *t* distribution with υ degree of freedom and the scale matrix $\mathbf{S}_t$:

$$f(\boldsymbol{\varepsilon}_t) = \frac{\Gamma[(\upsilon + k)/2]}{(\pi\upsilon)^{k/2}\Gamma(\upsilon/2)} \frac{|\mathbf{S}_t|^{-1/2}}{\left[1 + \boldsymbol{\varepsilon}_t^T \mathbf{S}_t^{-1} \boldsymbol{\varepsilon}_t / \upsilon\right]^{(\upsilon+k)/2}}, \tag{20}$$

where k is a dimension of $\boldsymbol{\varepsilon}_t$, $\Gamma(\bullet)$ is the gamma function. $\mathbf{S}_t$ is given by

$$\mathbf{S}_t = \frac{\upsilon - 2}{\upsilon} \mathbf{H}_t,$$

where the degree of freedom υ, simultaneously estimated with other parameters, should satisfy $\upsilon > 2$. The Student *t* distribution converges to the normal distribution as υ increases, but has kurtosis $3 + (6/(\upsilon - 4))$, which exists if and only if $\upsilon > 4$.

The time-varying conditional correlation between i th and j th variables is given by

$$\rho_{ij,t} = \frac{h_{ij,t}}{\sqrt{h_{ii,t} h_{jj,t}}}. \tag{21}$$

In the above BEKK model, the off-diagonal parameters are of particular interest in terms of volatility transmission across markets and banks. For instance, a_{ij} measures the transmission of the squared values of the shocks from i th variable in the previous period to the j th variable in the current period. Similarly, b_{ij} measures the transmission of the conditional volatility of the i th variable in the previous period to j th variable in the current period.

IV. Empirical Results

(i) Decomposition of Risk Premiums

A. Factor Analysis

Table 3 and Figure 4 report the estimation results of factor analysis by principal factor method. We did not assume the number of factors in advance. Table 3 shows that three factors were retained in terms of a positive eigenvalue. Eigenvalues of the first two factors exceed one and these two factors account for about 96% of the total variance. In particular, note the importance of the first factor, which account for about 70% of the total variance.

Table 3(ii) and Figure 4(i) show that the first factor has almost equal factor loadings across all of the risk premiums. We call this factor "global factor". On the other hand, the second factor has positive loadings on the dollar risk premiums, JD and FD, and negative loadings on the yen risk premiums, JY and FY. We call this factor "currency factor". Although the relevant factors are limited to these two factors in terms of the magnitude of eigenvalues and explanatory power, the third factor deserves our attention since it has positive loadings on the risk premiums for foreign banks, FY and FD, and negative loadings on the risk premiums for Japanese banks, JY and JD. This third factor is likely to correspond to the relative degree of credit fundamentals between Japanese and foreign banks, which we call "credit factor". Thus, it seems appropriate to control for the effects of the first two factors in extracting fundamental prices of credit risk for Japanese and foreign banks from risk premium data.

Figure 4(ii) shows the time-series movement of these three factors.[28] During the

[28] As a robustness check, we also used the method of independent component analysis. Independent component analysis is a recently developed linear transformation method that can decompose non-Gaussian data into the "statistically independent" factors. Using this method, we derived four factors from four risk premiums. The derived factors show a very similar movement to those derived by traditional factor analysis.

period of financial instability from 1997 to 1998, the global factor and the credit factor experienced two large spikes and dips, respectively, and the currency factor has one dip. Since the credit and currency factors move in the opposite direction regarding the risk premiums for Japanese banks, JY and JD, from the factor loadings, all of these factors, particularly the global and currency factors, are likely to have contributed a substantial rise in risk premiums for Japanese banks in this period.

B. State Space Model

Table 4 reports estimation results of the state space model. First, all of the coefficients, which link between the shadow prices of credit risk for Japanese and foreign banks, common to the yen and dollar markets, and the fundamental prices of bank credit risk, differently priced in the yen and dollar market, are estimated significantly at the 1% level. This result indicates that the relations between both prices are stable, although they contain positive noises. Also, it should be noted that $c2/c12$ are found to be significantly larger than $c1/c11$ as shown by the Wald test, as shown in Table 4(ii). In our interpretation, these coefficients are closely related to risk averseness of market participants in either yen or dollar market. Thus, the Wald test results suggest that the dollar market is significantly more risk averse than the yen market irrespective of the priced bank group, Japanese or foreign banks..

Next, all of the variance and covariance terms are significantly estimated at least at the 5% level. The significant estimates of the covariance terms indicate that cross-equation correlation structure for error terms, both between the yen and dollar markets for the same bank groups and between Japanese and foreign banks in the same currency markets, cannot be ignored.[29]

See Hyvarinen [1999], for details of independent component analysis.

[29] We estimated the same state space model without the covariance structure, but failed to get robust

Figure 5 compares the estimated shadow prices and the fundamental prices in either yen or dollar market.[30] We can see that among the fundamental prices, JYPRICE/ JDPRICE and FDPRICE move almost in parallel with the shadow prices of Japanese and foreign banks' credit risk, while FYPRICE seems quite insensitive to the shadow price of foreign banks' credit risk.

Table 5 reports summary statistics of the fundamental prices of bank credit risk. First, Table 5(i) shows that means of fundamental prices are almost the same as the corresponding risk premiums, but standard deviations are much lower. Also, skewness is negative for FYPRICE and FDPRICE, which was positive for FY and FD. The reason for this result may be that in our sample period, credit risk of foreign banks were not worried about in contrast to the Japanese banks and thus fundamental credit prices did not experience large spikes. Second, Table 5(ii) reports the correlation matrix between the fundamental prices of bank credit risk and the corresponding risk premiums. We find that correlations between fundamental prices for Japanese banks are high, while the correlations for foreign banks are even negative. This result suggests that factors other than credit fundamentals govern the variation of the observed risk premiums for foreign banks. Third, Table 5(iii) shows that the fundamental prices of bank credit risk explain only a small portion of the total variance of risk premiums: 2.6-2.7% for Japanese banks, and 0.5-2.3% for foreign banks. Particularly poor performance of FY is likely to arise from the relative impreciseness of yen-LIBOR as a proxy for foreign banks' yen interest rate.[31]

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parameter estimates of both constant coefficients and variance terms.

[30] We ignored the first five observations from the estimated shadow and fundamental prices of bank credit risk due to instability of the estimates, which is inherent in the Kalman filter setup.

[31] As mentioned in Section III, the number of foreign banks in yen-LIBOR is 11 out of 16 referenced banks.

(ii) Time-Series Analysis

A. Risk Premiums

a. Granger Causality Test and Generalized Impulse Response Function

Now, let us move on to the time-series analysis of both risk premiums. Table 6 reports the estimation results of the VAR model and the corresponding Granger causality test. The lag length is determined by the Schwarz Criterion. As is evident from Table 6(ii), we can observe a high degree of informational interdependence between risk premiums except between the dollar risk premiums for Japanese banks (JD) and the yen risk premiums for foreign banks (FY).

Figure 6 shows the generalized impulse response function. As a general tendency, the impulse responses from the shocks of the same currency markets are much larger than the responses from the shocks of the same bank groups. Also, the impulse responses from the shocks of the same currency markets respond faster and are exponentially decayed compared with those from the shocks of the same bank groups. This result suggests that the global and currency factors are more important determinants in pricing banks' risk than the credit fundamentals themselves, particularly in the short term.

b. M-GARCH Model

Table 7 reports estimation results of the M-GARCH model. The shape parameter υ is significantly larger than 2, indicating a much higher degree of fat tails than normal distribution. The Ljung-Box Q tests applied to the standardized residuals show that serial correlation of the risk premiums remains. However, our BEKK model did a fairly good job to eliminate the heteroscedasticity in squared standardized residuals.

Now, let us look at the estimation results of both ARCH and GARCH terms. All of the diagonal parameters are significant, which implies a high degree of persistence in conditional standard deviations. Regarding the estimation results of off-diagonal parameters, which measure the degree of volatility spillovers, 4 parameters out of 12 ARCH parameters and 8 parameters out of 12 GARCH parameters are significant. In particular, insignificance of the parameters a12, a21, a34 of ARCH parameters is of interest since they measure the interdependence of volatility between the same bank groups.[32]

Figure 7 shows the conditional correlations derived from the M-GARCH model. We can see that most conditional correlations widely fluctuate, which supports the use of time-varying correlations instead of usual constant correlations. Notable properties of the estimated conditional correlations are as follows. First, correlations of the same bank groups' risk premiums between the yen and dollar markets (JY vs. JD and FY vs. FD) fluctuate around zero. This is rather a surprising result since if the risk premiums properly reflect credit fundamentals of the bank groups, the correlations between JY and JD and between FY and FD should be high enough.

Second, throughout the whole period, correlations between Japanese and foreign banks' risk premiums in the same currency markets (JY vs. FY and JD vs. FD) are very high. This result is suggestive of the importance of a currency or global factor in decomposing the risk premiums.

[32] Note here that in assessing the volatility spillover, not the sign but only the significance level of the parameters is important since only squared ARCH and GARCH terms enter into the volatility spillover paths. Signs of ARCH and GARCH terms are important, however, in computing the conditional correlations.

B. Fundamental Prices of Credit Risk

a. Cointegrating Relationships

Since our fundamental prices of bank credit risk are I(1) by construction, we first analyze cointegrating relationships. Figure 8 shows the trace and maximum eigenvalue statistics from the stability test of cointegrating relationships among the four fundamental prices, JYPRICE, JDPRICE, FYPRICE, and FDPRICE.[33] Throughout the sample periods, three cointegrating relationships were found in a very stable manner.[34] Thus, we use full sample period to derive the cointegrating vectors, which is shown in Table 8. As is easily expected, the cointegration rank test shows that there are three cointegrating vectors among JYPRICE, JDPRICE, FYPRICE, and FDPRICE. LR test for the equilibrium relationship [1,-1,-1,1,C] shows, however, that the cointegration restriction is rejected at the 1% significance level.

b. Granger Causality Test and Generalized Impulse Response Function

Table 9 reports the estimation results of VECM and the Granger causality test. Lag length of the VECM is determined by the Schwarz Criterion. As shown in Table 9(ii), a higher degree of informational interdependence are found than in the case of risk premiums reported in Table 6(ii). Indeed, each of the four fundamental prices significantly Granger-causes other three prices.

Figure 9 shows the generalized impulse response function of each fundamental price. In contrast to the risk premiums, each fundamental price of bank credit risk responds larger from

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[33] We use 1,000 observations for the size of the rolling window. We followed Banerjee, Lumsdaine, and Stock [1992], who recommend that the size of the rolling window should be one-third of total number of observations in the context of the stability of unit-root tests. Since we have 3,086 observations in total, the choice of 1,000 observations correspond to their recommendation.

[34] The sole exception is around September 1999. But, if we adopt the 10% significance level, three cointegrating relationships are found.

the shocks of the same bank groups than from the shocks of the same currency markets. This result suggests that estimated fundamental prices of bank credit risk properly reflect credit fundamentals.

c. M-GARCH Model

Table 10 reports estimation results of the M-GARCH model consisting of four fundamental prices. The shape parameter υ is significantly larger than 2 as in the case of risk premiums. The Ljung-Box Q tests that serial correlation remains in the fundamental prices of Japanese banks' credit risk, JYPRICE and JDPRICE. However, our BEKK model did a fairly good job to eliminate the heteroscedasticity in the fundamental prices of foreign banks' credit risk, FYPRICE and FDPRICE.

Next, let us look at the estimation results of ARCH and GARCH terms. First, all of the diagonal parameters are significant. Second, regarding the estimation results of off-diagonal parameters, 8 parameters out of 12 ARCH parameters and 8 parameters out of 12 GARCH parameters were significant. In particular, it is noteworthy that the parameters that measure the interdependence of volatility between the same bank groups, a12, a21, a34 of ARCH parameters are significant unlike the case of risk premiums.

Figure 10 shows the conditional correlations between four fundamental prices. There are several points to note here, as compared to the case of risk premiums shown in Figure 7. First, the fundamental prices for the same bank groups, JYPRICE/JDPRICE and FYPRICE/FDPRICE are highly correlated almost throughout the sample period. In particular, the correlation between JYPRICE and JDPRICE is found to be almost unity. Second, the fundamental prices of different bank groups in the same currency market, JYPRICE/FYPRICE and JDPRICE/FDPRICE, are negatively correlated in marked contrast to the case of risk premiums in which JY/FY and JD/FD shows a very high correlations. Also, the fundamental prices of difference banks in different currency markets, JYPRICE/FDPRICE and JDPRICE/FYPRICE, are negatively correlated in many phases. This result indicates that credit risks of Japanese banks and foreign banks have moved in an opposite direction during the sample period. This is consistent with our experience in that Japanese banks have struggled hard to dispose of their non-performing loans until quite recently , while the U.S. banks recovered from the S&L crisis and Latin American crisis occurred in the 1980s from the 1990s.

Put these results together, we infer that we successfully derived the fundamental prices of each bank group. And the global and currency common factors are likely to create spurious correlations between observed risk premiums for different bank groups between the different currency markets.

V. Concluding Remarks

This paper has investigated the role of TIBOR/LIBOR as indicators of bank credit risk and the interdependence structure of bank credit risk in the money markets within and across the border. In doing so, we decomposed the risk premiums for Japanese and foreign banks constructed from TIBOR/LIBOR to extract fundamental prices of credit risk. Our findings can be summarized as follows.

(i) Observed risk premiums constructed from TIBOR/LIBOR contain two common factors, global and currency factors, which explain most of the variation of the observed risk premiums.

(ii) Thus, the generalized impulse response of risk premiums from the shocks of the same currency markets are much larger than the responses from the shocks of the same bank groups. And the conditional correlations of the same bank groups' risk premiums between the yen and dollar markets fluctuate around zero, while the correlations between Japanese and foreign banks' risk premiums in the same currency market are very high.

(iii) After controlling for these common factors, we successfully derived the fundamental prices of bank credit risk both for Japanese and foreign banks using the state space model. These fundamental prices show plausible time-series properties such as a high degree of impulse response from the shocks of the same bank groups, and a high correlation of the same bank groups' credit risk between the two markets. However, the fundamental prices account for only a tiny portion of the total variance of risk premiums.

Put these results together, although TIBOR/LIBOR have played the role of indicators of bank credit risk since the 1990s, the importance has been substantially reduced, as asserted by Ito and Harada [2004]. We conclude this paper by mentioning three possible causes of this result. The first one is that Japanese banks have been required to put up cash collaterals to raise dollars in the money markets since around 2000-2001. The second one is that weaker banks have already exited from the international money markets. These possibilities are pointed out by Ito and Harada [2004]. The third one is that money markets ceases to properly function as a price discovery mechanism in a very low interest rate environment, particularly in Japan. Baba, et al. [2005] supports this view by analyzing the dispersion and credit curves of interest rates on NCDs issued by individual Japanese banks. To determine the relative importance of these hypotheses is beyond the scope of this paper. This is one of our future tasks.

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Appendix 1: Negative FX Swap Yen Funding Cost for Foreign Banks

In this Appendix, we try to decompose the FX swap yen funding cost for foreign banks, which has been frequently negative in recent years. Let us restate the yen funding costs in net terms c for foreign banks in the FX swap market as

$$1+c=\frac{F}{S}\left(1+i^{*}+FD\right) \tag{1a}$$

Equilibrium condition (3) implies that

$$1+c=\frac{F}{S}\left(1+i^{*}+FD\right)=\frac{1+i+JY}{1+i^{*}+JD}\left(1+i^{*}+FD\right)$$

$$\Leftrightarrow\ c\approx i+FD-\left[JD-JY\right] \tag{2a}$$

Equation (2a) shows that the yen funding costs for foreign banks in the FX swap market can be decomposed into the following three factors: (a) the yen risk-free interest rate, (b) the risk premium for foreign banks in the dollar market, and (c) the difference in the risk premiums for Japanese banks between the dollar and yen markets.

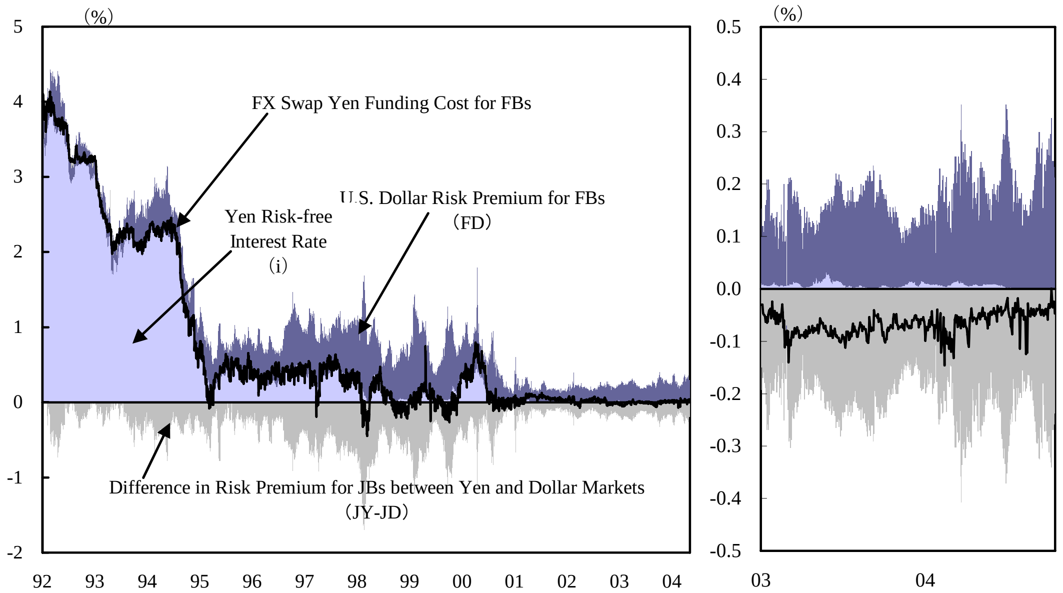
If the difference in the risk premiums for Japanese banks between the two markets is zero, that is, $JD=JY$, the yen funding costs for foreign banks boil down to the usual form of funding costs: the sum of the yen risk-free interest rate and the risk premium for foreign banks.[35] Put differently, the fact that the yen funding costs differ from the usual form of funding costs stems from the difference in the risk evaluation of Japanese banks between the dollar and the yen markets.

Appendix Figure shows the decomposition result based on this asymmetry in risk evaluation between the two markets. This figure reveals that in a quite low interest rate environment

[35] Note that equation (4) yields $FY=FD=\overline{\theta}$ when $JY=JD$.

in recent years, the difference in risk premiums for Japanese Banks between yen and dollar markets causes negative FX swap yen funding cost for foreign banks.

Appendix Figure: Decomposition of FX Swap Yen Funding Cost for Foreign Banks



Note: Left figure uses TIBOR/LIBOR as the proxy for the interest rates for Japanese banks (JBs) and foreign banks (FBs), respectively. Right figure uses yen and U.S. dollar bid interest rates exclusively for Japanese and foreign banks in the Euro markets. The bid rates are available only from May 2004.

Source: Meitan Tradition Co. (left figure). For the data source of right figure, see Appendix 2.

Appendix 2: Data Details

We use 90-day TIBOR/LIBOR defined as the average of the interest rates offered by reference banks as the proxy for the interest rates for Japanese banks and foreign banks, respectively. Data sources are as follows:

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

i +JY	Yen Interest Rate for Japanese Banks	Japanese Bankers' Association (Yen-TIBOR)
i +FY	Yen Interest Rate for Foreign Banks	British Bankers Association (Yen-LIBOR)
i^* +JD	U.S. Dollar Interest Rate for Japanese Banks	QUICK (U.S. Dollar-TIBOR)
i^* +FD	U.S. Dollar Interest Rate for Foreign Banks	British Bankers Association (U.S. Dollar-LIBOR)
i	Treasury Bill Rate in Japan	Bloomberg
i^*	Treasury Bill Rate in the U.S.	FRB, *FRED*
S	Domestic Currency Value of the Spot Exchange Rates.	Bank of Japan
$F - S$	Forward Premium	Bank of Japan

The reference banks of TIBOR and LIBOR are as follows:

Yen-TIBOR	**Mizuho Bank**, **Sumitomo Mitsui Banking Co**., JP Morgan Chase, **the Bank of Tokyo Mitsubishi**, **Saitama Resona Bank**, **UFJ Bank**, **Shinsei Bank**, **the Chuo Mitsui Trust and Banking Co**., **the Mitsubishi Trust and Banking Co**., **the Sumitomo Trust and Banking Co**., **Mizuho Corporate Bank**, **Mizuho Trust and Banking Co**., **the Shoko Chukin Bank**, UBS AG, **Shinkin Central Bank**, **the Norinchukin Bank**
U.S. Dollar-TIBOR	**Sumitomo Mitsui Banking Co**., **the Bank of Tokyo Mitsubishi**, **UFJ Bank**, **Mizuho Corporate Bank**, **the Norinchukin Bank**, **the Mitsubishi Trust and Banking Co**., **the Sumitomo Trust and Banking Co**., **the Chuo Mitsui Trust and Banking Co**., Citibank NA, UBS AG
Yen-LIBOR	Bank of America, Barclays Bank Plc, Citibank NA, Deutsche Bank AG, HSBC, JP Morgan Chase, Lloyds TSB Bank Plc, Rabobank, The Royal Bank of Scotland Group, UBS AG, Westdeutsche Landesbank AG, **the Bank of Tokyo Mitsubishi**, **Sumitomo Mitsui Banking Co**., **Mizuho Corporate Bank**, **UFJ Bank**, **the Norinchukin Bank**
U.S. Dollar-LIBOR	Abbey National Plc, Bank of America, Barclays Bank Plc, Citibank NA, Credit Suisse First Boston, Deutsche Bank AG, HBOS, HSBC, JP Morgan Chase, Lloyds TSB Bank Plc, Rabobank, The Royal Bank of Scotland Group, UBS AG, Westdeutsche Landesbank, **the Bank of Tokyo Mitsubishi**, **the Norinchukin Bank**

Note: Bold letters indicate Japanese banks.

Table 1: Summary Statistics (i)

(i) Basic Statistics

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

(%)	Mean	Std.Dev.	Skewness	Kurtosis	Jarque-Bera	LB(12)	$LB^2(12)$
JY	0.169	0.139	2.047	7.974	5330.08***	30023***	21107***
JD	0.483	0.311	1.729	6.960	3549.75***	32020***	30681***
FY	0.129	0.093	1.475	5.432	1877.01***	24999***	21107***
FD	0.398	0.216	1.324	5.579	1754.67***	29536***	26392***

(ii) Correlation Matrix

	JY	JD	FY	FD
JY	1.000			
JD	0.664	1.000		
FY	0.918	0.529	1.000	
FD	0.349	0.895	0.294	1.000

Notes: 1. LB(12) and $LB^2(12)$ are Ljung-Box *Q* test statistics for serial correlations of the variables themselves and squared variables up to the 12^{th} order.
2. *** denotes the 1% significance level.

Table 2: Unit Root Test

Specification: $\Delta y_t = \mu + \alpha_0 y_{t-1} + \alpha_1 \Delta y_{t-1} + \cdots \alpha_N \Delta y_{t-N} + bt + \varepsilon_t$

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

	ADF (Augmented Dickey-Fuller) Test		PP (Phillips-Perron) Test	
	Test Statistic	Lags	Test Statistic	Bandwidth
JY	-5.462***	0	-5.405***	9
JD	-4.363***	2	-4.334***	15
FY	-7.958***	2	-8.847***	4
FD	-5.132***	3	-6.361***	1

Notes: 1. The number of lags is chosen based on Schwarz Criterion.
2. *, **, and *** show that the null hypothesis of the existence of a unit root is rejected at the 10%, 5% and 1% significance level, respectively.

JY : Japanese Banks' Yen Risk Premium JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium FD: Foreign Banks' Dollar Risk Premium

Table 3: Estimation Results of Factor Analysis (i)

(i) Importance of Factors

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

	Factor 1	Factor 2	Factor 3
Eigenvalue	2.820	1.000	0.085
Proportion of the total variance	0.705	0.250	0.021
Cumulative proportion of the total variance	0.705	0.955	0.976

(ii) Factor Loadings

	Factor 1	Factor 2	Factor 3
JY	0.883	-0.448	-0.131
JD	0.921	0.372	-0.136
FY	0.807	-0.511	0.179
FD	0.736	0.633	0.132

Notes: 1. The method of principal factor is used.
2. The result is before rotation.

JY : Japanese Banks' Yen Risk Premium JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium FD: Foreign Banks' Dollar Risk Premium

Table 4: Estimation Results of State Space Model

(i) Specification

$$\begin{bmatrix} JY \\ JD \\ FY \\ FD \end{bmatrix}_t = \begin{bmatrix} JYPRICE \\ JDPRICE \\ FYPRICE \\ FDPRICE \end{bmatrix}_t + \begin{bmatrix} s4 & s6 \\ s5 & s7 \\ s14 & s16 \\ s15 & s17 \end{bmatrix}_t \begin{bmatrix} Fg \\ Fc \end{bmatrix}_t + \begin{bmatrix} e1 \\ e2 \\ e11 \\ e12 \end{bmatrix}_t ,$$

where

$$\begin{bmatrix} JYPRICE \\ JDPRICE \end{bmatrix}_t = \begin{bmatrix} c3 \\ c4 \end{bmatrix} + \begin{bmatrix} c1 \\ c2 \end{bmatrix} s3_t + \begin{bmatrix} e3 \\ e4 \end{bmatrix}, \quad \begin{bmatrix} FYPRICE \\ FDPRICE \end{bmatrix}_t = \begin{bmatrix} c13 \\ c14 \end{bmatrix} + \begin{bmatrix} c11 \\ c12 \end{bmatrix} s13_t + \begin{bmatrix} e13 \\ e14 \end{bmatrix},$$

$$\begin{bmatrix} s3 \\ \vdots \\ s7 \end{bmatrix}_t = \begin{bmatrix} s3 \\ \vdots \\ s7 \end{bmatrix}_{t-1} + \begin{bmatrix} e5 \\ \vdots \\ e9 \end{bmatrix}_t , \text{ and } \begin{bmatrix} s13 \\ \vdots \\ s17 \end{bmatrix}_t = \begin{bmatrix} s13 \\ \vdots \\ s17 \end{bmatrix}_{t-1} + \begin{bmatrix} e15 \\ \vdots \\ e19 \end{bmatrix}_t .$$

(ii) Parameter Estimates and Wald Test Results

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

	Parameter	Std. error
c1	0.069***	5.01E-07
c2	0.150***	9.47E-09
c3	-0.349***	0.010
c4	-0.650***	0.022
c11	0.020***	3.49E-08
c12	0.105***	1.46E-05
c13	-0.442***	0.108
c14	-2.847***	0.564
lnVar(e1)	-18.505***	3.44E-06
lnVar(e2)	-14.543***	0.001
lnVar(e3)	-17.037***	6.29E-05
lnVar(e4)	-14.465***	0.026
lnVar(e5)	-4.789***	4.23E-05
lnVar(e6)	-16.008***	0.009
lnVar(e7)	-15.946***	1.47E-04
lnVar(e8)	-15.209***	0.001
lnVar(e9)	-14.545***	4.58E-06
lnVar(e11)	-9.820***	0.024
lnVar(e12)	-9.210***	0.009
lnVar(e13)	-10.203***	0.001
lnVar(e14)	-10.991***	0.001
lnVar(e15)	-5.954***	0.002
lnVar(e16)	-10.747***	0.013
lnVar(e17)	-10.144***	0.003
lnVar(e18)	-9.875***	0.041
lnVar(e19)	-9.901***	0.001
cov(e1,e2)*10^3	3.54E-05***	2.67E-06
cov(e3,e4)*10^3	2.35E-05***	6.00E-06
cov(e6,e7)*10^3	6.59E-05***	3.20E-09
cov(e8,e9)*10^3	2.20E-04***	1.62E-07
cov(e11,e12)*10^3	-2.67E-04***	1.03E-07
cov(e13,e14)*10^3	0.001***	8.90E-10
cov(e16,e17)*10^3	0.009***	6.31E-07
cov(e18,e19)*10^3	-7.67E-05***	1.77E-09
cov(e1,e11)*10^3	0.001***	1.75E-05
cov(e2,e12)*10^3	2.80E-04***	3.05E-05
cov(e3,e13)*10^3	0.001***	2.27E-08
cov(e4,e14)*10^3	0.008***	4.22E-05
cov(e5,e15)*10^3	-0.060**	0.028
cov(e6,e16)*10^3	0.001***	9.10E-06
cov(e7,e17)*10^3	0.001***	2.22E-05
cov(e8,e18)*10^3	0.002***	4.94E-05
cov(e9,e19)*10^3	0.003***	1.35E-05
Log likelihood	42617.07	

Wald Test	
Null Hypothesis (H0)	χ^2
c2-c1=0	2.74E+10***
c12-c11=0	3.38E+07***

Notes: 1. Marquardt method is used as an optimization algorithm.
2. *, ***, and *** denote the 10%, 5%, and 1% significance level, respectively.

Table 5: Summary Statistics (ii)

(i) Basic Statistics

Sample Period: 1992/8/7 to 2005/2/2 (Number of Observations: 3,081)

	Mean	Std.Dev.	Skewness	Kurtosis	Jarque-Bera	LB(12)	$LB^2(12)$
JYPRICE	0.002	0.001	2.043	7.956	5303.88***	30023***	21107***
JDPRICE	0.005	0.003	1.731	6.968	3565.92***	32020***	30681***
FYPRICE	0.001	0.001	1.470	5.409	1856.78***	24999***	21107***
FDPRICE	0.004	0.002	1.325	5.582	1760.55***	29536***	26392***

(ii) Correlation Matrix

	JY	JD	FY	FD
JYPRICE	1.000			
JDPRICE	0.664	1.000		
FYPRICE	0.918	0.529	1.000	
FDPRICE	0.349	0.895	0.294	1.000

Notes: 1. LB(12) and $LB^2(12)$ are Ljung-Box Q test statistics for serial correlations of the variables themselves and squared variables up to the 12th order.
2. *** denotes significance at the 1% level.

(iii) Importance of Factors: Proportion of the Total Variance

	Global Factor	Currency Factor	Fundamental Price	Three Factors
JY	0.724	0.250	0.027	0.999
JD	0.805	0.169	0.026	0.999
FY	0718	0.278	0.005	0.999
FD	0.681	0.293	0.023	0.997

JYPRICE : Fundamental Price of Credit Risk for Japanese Banks in the Yen Market
JDPRICE : Fundamental Price of Credit Risk for Japanese Banks in the Dollar Market
FYPRICE : Fundamental Price of Credit Risk for Foreign Banks in the Yen Market
FDPRICE : Fundamental Price of Credit Risk for Foreign Banks in the Dollar Market

JY : Japanese Banks' Yen Risk Premium JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium FD: Foreign Banks' Dollar Risk Premium

Table 6: Estimation Results of VAR Model

(i) Estimation Results

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

	JY	JD	FY	FD
Constant	0.005***	0.006**	0.003***	0.011***
	(0.001)	(0.003)	(0.001)	(0.002)
JY(-1)	0.850***	0.139***	0.261***	-0.052
	(0.026)	(0.056)	(0.028)	(0.051)
JY(-2)	-0.063**	-0.150**	-0.016***	-0.076
	(0.031)	(0.069)	(0.034)	(0.062)
JY(-3)	0.087***	-0.016	0.007	0.039
	(0.025)	(0.054)	(0.027)	(0.049)
JD(-1)	0.090***	0.781***	0.004	0.097***
	(0.016)	(0.035)	(0.017)	(0.031)
JD(-2)	0.038**	0.193***	-0.005	-0.003
	(0.019)	(0.041)	(0.020)	(0.037)
JD(-3)	-0.060***	0.021	-0.022	-0.047
	(0.016)	(0.035)	(0.017)	(0.032)
FY(-1)	0.143***	-0.043	0.686***	-0.002
	(0.023)	(0.050)	(0.025)	(0.046)
FY(-2)	-0.017	0.096*	0.027	0.107**
	(0.026)	(0.058)	(0.028)	(0.053)
FY(-3)	-0.054**	-0.007	0.116***	-0.016
	(0.023)	(0.050)	(0.025)	(0.046)
FD(-1)	-0.098***	0.119***	-0.016	0.750***
	(0.017)	(0.038)	(0.019)	(0.034)
FD(-2)	-0.026	-0.175***	0.032	0.051
	(0.020)	(0.044)	(0.022)	(0.040)
FD(-3)	0.059	0.044	0.011	0.125***
	(0.017)	(0.038)	(0.019)	(0.035)
Adj. R-squared	0.968	0.969	0.915	0.947

(ii) Granger Causality Test Statistics ($\chi 2$ Statistics)

	JY	JD	FY	FD
Excluded				
JY		8.629**	93.541***	7.597*
JD	102.221***		6.201	16.427***
FY	48.262***	3.633		8.370**
FD	82.461***	16.978***	8.274**	
ALL	147.017***	27.650***	137.078***	21.510**

Notes: 1. Figures in parentheses are standard errors.
2. ***, **, and * denote the 1%, 5%, and 10% significance level, respectively.
3. Lag length is chosen based on Schwarz Criterion.
4. Figures in (ii) denote the $\chi 2$ test statistics.

JY : Japanese Banks' Yen Risk Premium JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium FD: Foreign Banks' Dollar Risk Premium

Table 7: Estimation Result of M-GARCH Model (i)

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

		Parameter	std. error	*t*-value
		ARCH		
a11		0.408***	0.026	15.840
a12		0.080	0.058	1.390
a13		-0.310***	0.027	-11.580
a14		-0.055	0.054	-1.027
a21		0.007	0.009	0.870
a22		0.419***	0.040	10.036
a23		0.041***	0.011	3.726
a24		0.030	0.034	0.859
a31		-0.103***	0.018	-5.788
a32		-0.045	0.055	-0.815
a33		0.632***	0.037	17.120
a34		0.042	0.053	0.805
a41		0.002	0.009	0.261
a42		-0.026	0.038	-0.679
a43		-0.035**	0.011	-3.162
a44		0.351***	0.039	8.901
		GARCH		
b11		0.938***	0.004	220.000
b12		-0.030**	0.016	-1.896
b13		0.064***	0.005	11.930
b14		0.011	0.014	0.800
b21		-0.011***	0.002	-5.118
b22		0.911***	0.009	99.630
b23		-0.012***	0.003	-4.083
b24		-0.022***	0.008	-2.665
b31		0.032***	0.005	7.170
b32		0.021	0.017	1.219
b33		0.892***	0.006	151.300
b34		-0.017	0.015	-1.098
b41		0.006***	0.002	3.097
b42		-0.015*	0.009	-1.575
b43		0.006***	0.003	2.433
b44		0.934***	0.009	109.600
		Diagnostic Statistics		
υ (Student *t*)		2.786***	0.098	7.984
LB(12)	JY	63.66***		
	JD	23.36**		
	FY	221.16***		
	FD	29.71**		
LB^2 (12)	JY	0.09		
	JD	15.03		
	FY	5.55		
	FD	12.16		

Notes: 1. υ is the shape parameter (degree of freedom) of the Student *t* distribution for the four joint error processes. *t*-values are computed based on the null and alternative hypotheses υ =2 and υ >2, respectively.

2. a_{ij} and b_{ij} measure the volatility transmission from *i*-th to *j*-th risk premiums (1:JY, 2:JD, 3:FY, 4:FD).

3. LB(12) and LB^2(12) are Ljung-Box *Q* tests for white noise in the linear and squared standardized residuals up to the 12th order.

4. *, ***, and *** denote the 10%, 5%, and 1% significance level, respectively.

5. Estimation results of constant terms are omitted due to the limitation of space.

Table 8: Cointegration Test

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

Cointegration Rank Test					
H0	H1	Eigenvalue	Trace Statistic	Max-Eigen Statistic	Lags
$r \leq 0$	$r = 1$	0.201	1098.999***	689.867***	3
$r \leq 1$	$r = 2$	0.099	409.132***	323.981***	
$r \leq 2$	$r = 3$	0.020	85.151***	63.814***	
$r \leq 3$	$r = 4$	0.007	21.337***	21.337***	

Cointegrating Vectors				
JYPRICE	JDPRICE	FYPRICE	FDPRICE	Constant
1.000	0.000	0.000	0.677***	-0.454***
0.000	1.000	0.000	1.476***	-1.102***
0.000	0.000	1.000	-0.195***	-0.103***

Test of Cointegration Restrictions				
1.000	-1.000	-1.000	1.000	-26.734*** -34.833*** 8.033***
LR test : $\chi 2$ (1)=557.725***				

Notes 1. JBs and FBs denote Japanese banks and foreign banks, respectively.
2. We took logarithm of interest rates.
3. The number of lags is chosen based on Schwarz Criterion.
4. r denotes the number of cointerating ranks.
5. LR test denotes the Log-Likelihood Ratio test for the equilibrium relationship [1,-1,-1,1,C], where C is constant.
6 *, **, and *** denote the 10%, 5%, and 1% significance level, respectively.

Table 9: Estimation Results of VECM

(i) Estimation Results

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

	Δ (JYPRICE)	Δ (JDPRICE)	Δ (FYPRICE)	Δ (FDPRICE)
Error 1	-1.536**	-1.378	-0.746	15.161***
	(0.723)	(1.589)	(0.481)	(1.159)
Error 2	0.693**	0.608	0.342	-6.960***
	(0.330)	(0.726)	(0.220)	(0.530)
Error 3	0.174***	0.352***	-0.365***	-0.504***
	(0.042)	(0.091)	(0.028)	(0.067)
Δ (JYPRICE(-1))	-0.785	-1.780	0.126	-9.938***
	(0.766)	(1.685)	(0.510)	(1.229)
Δ (JYPRICE(-2))	1.565***	3.494***	-0.820**	-3.722***
	(0.599)	(1.318)	(0.399)	(0.961)
Δ (JYPRICE(-3))	-0.043	-0.149	-0.007	-3.773***
	(0.377)	(0.830)	(0.251)	(0.605)
Δ (JDPRICE(-1))	0.233	0.540	-0.014	4.560***
	(0.352)	(0.775)	(0.235)	(0.566)
Δ (JDPRICE(-2))	-0.805**	-1.798***	0.422	1.687***
	(0.277)	(0.609)	(0.184)	(0.444)
Δ (JDPRICE(-3))	0.019	0.064	0.014	1.729***
	(0.174)	(0.383)	(0.116)	(0.280)
Δ (FYPRICE(-1))	0.003	0.018	-0.074**	0.526***
	(0.050)	(0.110)	(0.033)	(0.080)
Δ (FYPRICE(-2))	-0.131***	-0.285***	0.039	0.226***
	(0.047)	(0.102)	(0.031)	(0.075)
Δ (FYPRICE(-3))	0.040	0.097	-0.022	0.272***
	(0.035)	(0.078)	(0.024)	(0.057)
Δ (FDPRICE(-1))	-0.016	-0.038	-0.005	-0.233***
	(0.018)	(0.040)	(0.012)	(0.029)
Δ (FDPRICE(-2))	0.050***	0.111***	-0.023***	-0.035*
	(0.012)	(0.025)	(0.008)	(0.019)
Δ (FDPRICE(-3))	-0.020**	-0.045**	0.005	-0.060***
	(0.009)	(0.019)	(0.006)	(0.014)
Adj. R-squared	0.299	0.299	0.275	0.133

(ii) Granger Causality Test Statistics ($\chi 2$ Statistics)

	JYPRICE	JDPRICE	FYPRICE	FDPRICE
Excluded				
JYPRICE		24.222***	10.227**	77.841***
JDPRICE	22.599***		10.731**	76.870***
FYPRICE	15.147***	15.716***		52.538***
FDPRICE	28.422***	29.066***	10.798**	
ALL	43.257***	47.153***	102.239***	88.699***

Notes: 1. Figures in parentheses are standard errors.
2. ***, **, and * denote significance at the 1%, 5%, and 10% level, respectively.
3. Lag length is chosen based on Schwarz Criterion.
4. Figures in B denote the test statistics following $\chi 2$ distribution with degree of freedom 2 when one variable is excluded and 6 when all the variables are excluded.

Table 10: Estimation Results of M-GARCH Model (ii)

Sample Period: 1992/8/3 to 2005/2/2 (Number of Observations: 3,086)

		Parameter	std. error	*t*-value
		ARCH		
	a11	0.891***	0.030	30.150
	a12	0.799***	0.105	7.642
	a13	-0.381**	0.212	-1.798
	a14	-5.885***	0.732	-8.041
	a21	-0.115***	0.020	-5.664
	a22	0.270***	0.071	3.831
	a23	0.148***	0.098	1.520
	a24	2.613***	0.336	7.772
	a31	-0.001	0.016	-0.049
	a32	0.004	0.035	0.116
	a33	0.414***	0.026	15.690
	a34	0.132***	0.048	2.736
	a41	0.001	0.007	0.873
	a42	0.001	0.015	0.124
	a43	-0.002***	0.007	-2.730
	a44	0.207***	0.025	8.277
		GARCH		
	b11	0.749***	0.015	49.320
	b12	-0.320***	0.028	-11.590
	b13	0.001	0.044	0.148
	b14	2.152***	0.149	14.450
	b21	0.039***	0.006	6.881
	b22	0.983***	0.009	11.110
	b23	0.007	0.020	0.358
	b24	-0.957***	0.069	-13.840
	b31	-0.010**	0.004	-2.139
	b32	-0.024***	0.010	-2.445
	b33	0.943***	0.004	23.490
	b34	-0.051***	0.012	-4.415
	b41	0.001	0.002	0.371
	b42	0.005*	0.004	1.361
	b43	0.002	0.002	1.002
	b44	1.004***	0.005	206.500
		Diagnostic Statistics		
	υ (Student *t*)	2.587***	0.078	7.517
LB(12)	JYPRICE	15.33		
	JDPRICE	10.63		
	FYPRICE	41.00***		
	FDPRICE	30.17***		
$LB^2(12)$	JYPRICE	0.42		
	JDPRICE	16.32		
	FYPRICE	65.73***		
	FDPRICE	454.04***		

Notes: 1. υ is the shape parameter (degree of freedom) of the Student *t* distribution for the four joint error processes. *t*-values are computed based on the null and alternative hypotheses υ =2 and υ >2, respectively.

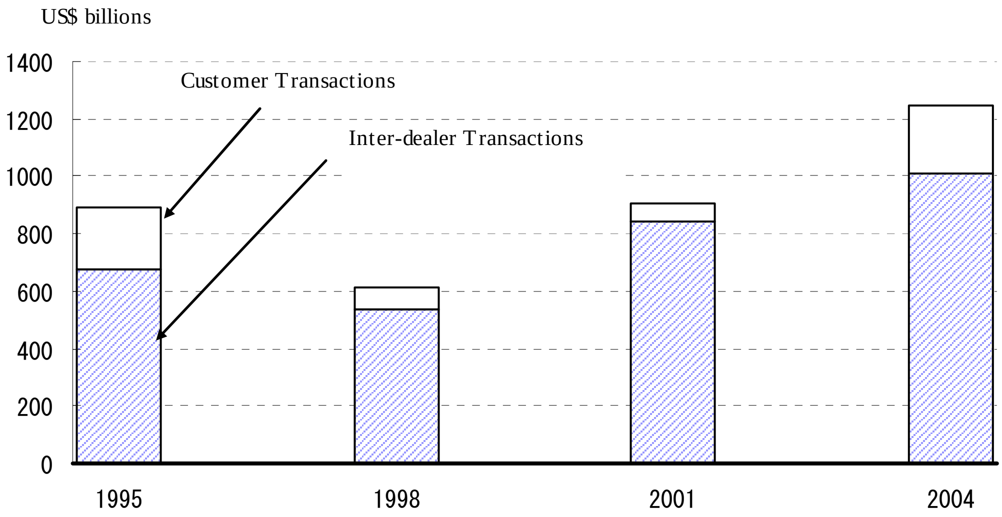
2. a_{ij} and b_{ij} measure the volatility transmission from *i*-th to *j*-th risk premiums (1:JYPRICE, 2:JDPRICE, 3:FYPRICE, 4:FDPRICE).

3. LB(12) and $LB^2(12)$ are Ljung-Box *Q* tests for white noise in the linear and squared standardized residuals up to the 12^{th} order.

4. *, ***, and *** denote the 10%, 5%, and 1% significance level, respectively.

5. Estimation results of constant terms are omitted due to the limitation of space.

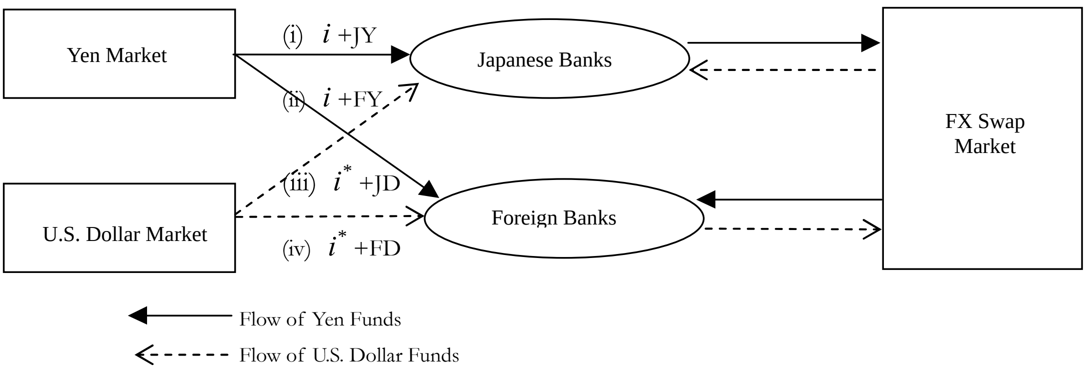
Figure 1: Transaction Volume of FX Swap Market



Note: The data is as of April.

Source: "Central Bank Survey of Foreign Exchange and Derivatives Market Activity", Bank of Japan

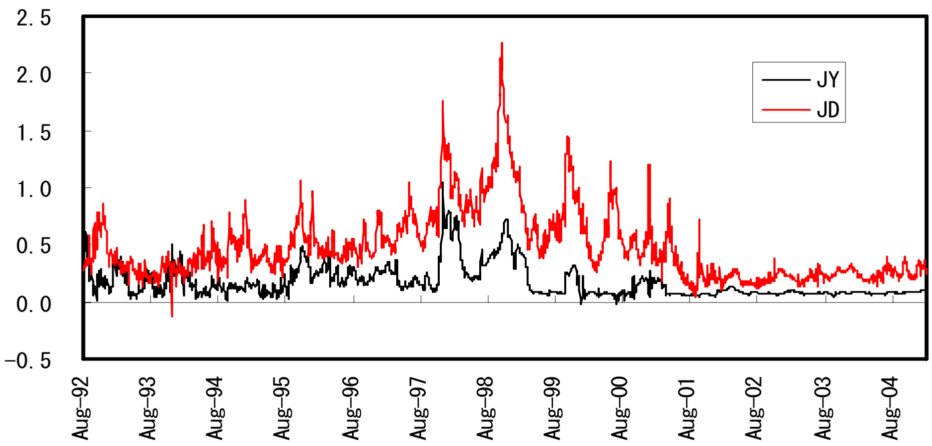
Figure 2: Foreign Currency Funding Structure of Japanese and Foreign Banks



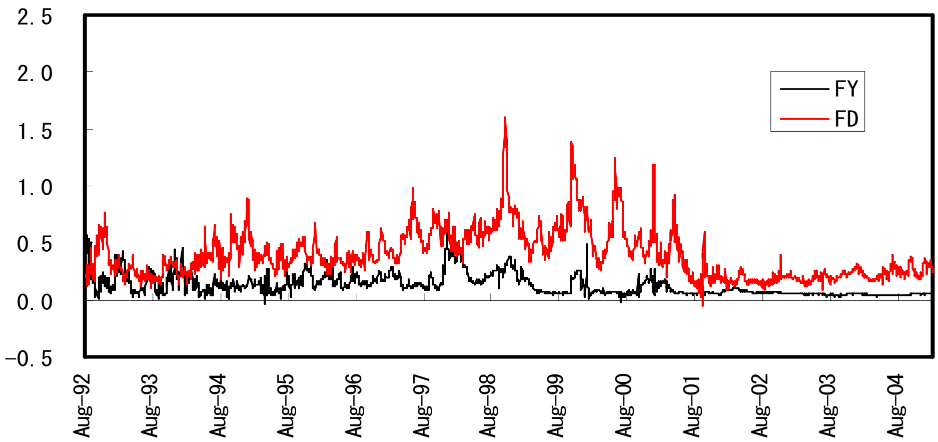
JY : Japanese Banks' Yen Risk Premium JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium FD: Foreign Banks' Dollar Risk Premium

Figure 3: Risk Premiums

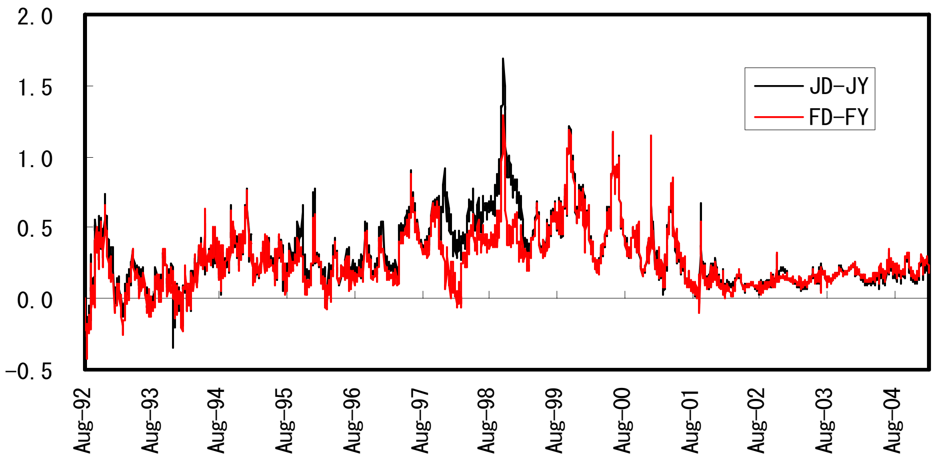
(i) Risk Premiums for Japanese Banks



(ii) Risk Premiums for Foreign Banks



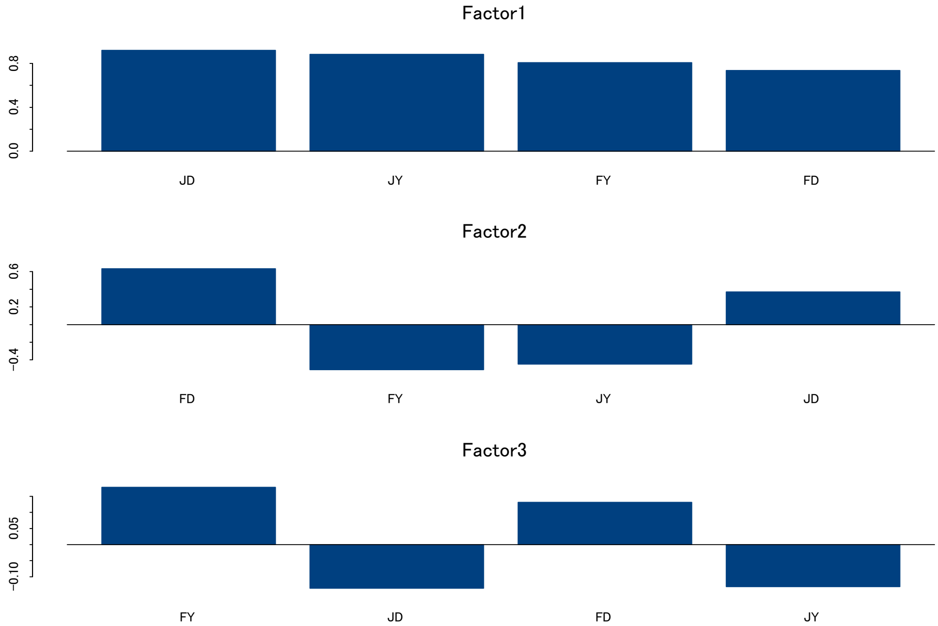
(iii) Difference in Risk Premiums between Dollar and Yen Markets



JY : Japanese Banks' Yen Risk Premium	JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium	FD: Foreign Banks' Dollar Risk Premium

Figure 4: Estimation Results of Factor Analysis (ii)

(i) Factor Loadings



(ii) Time-Series Movement of Each Factor

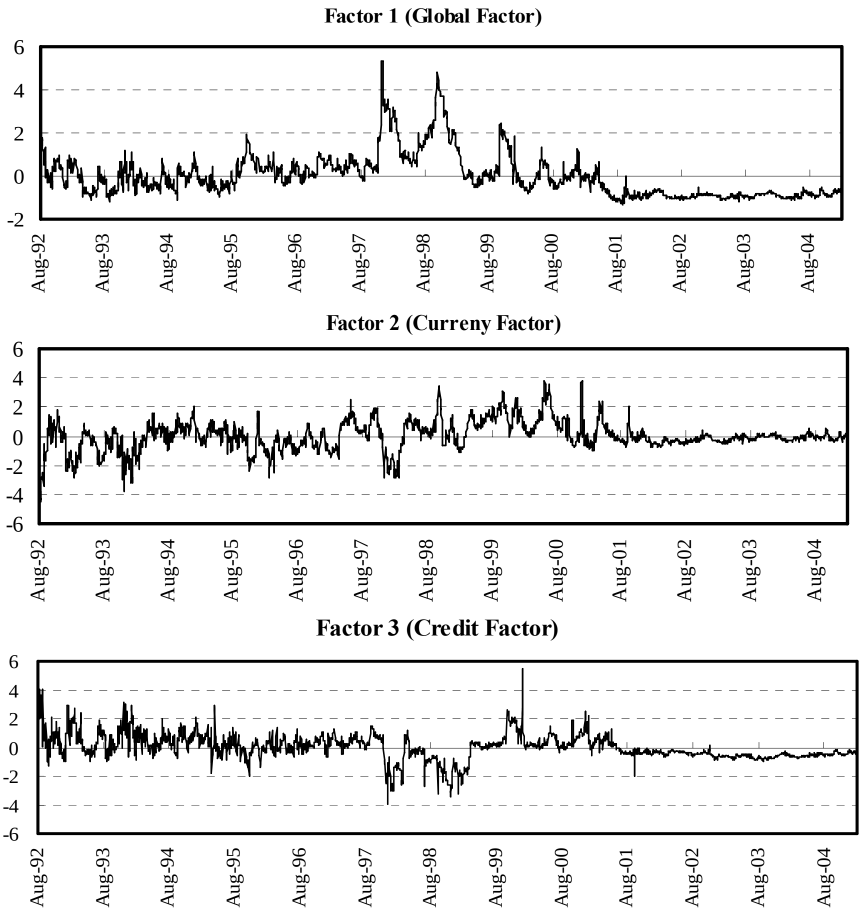
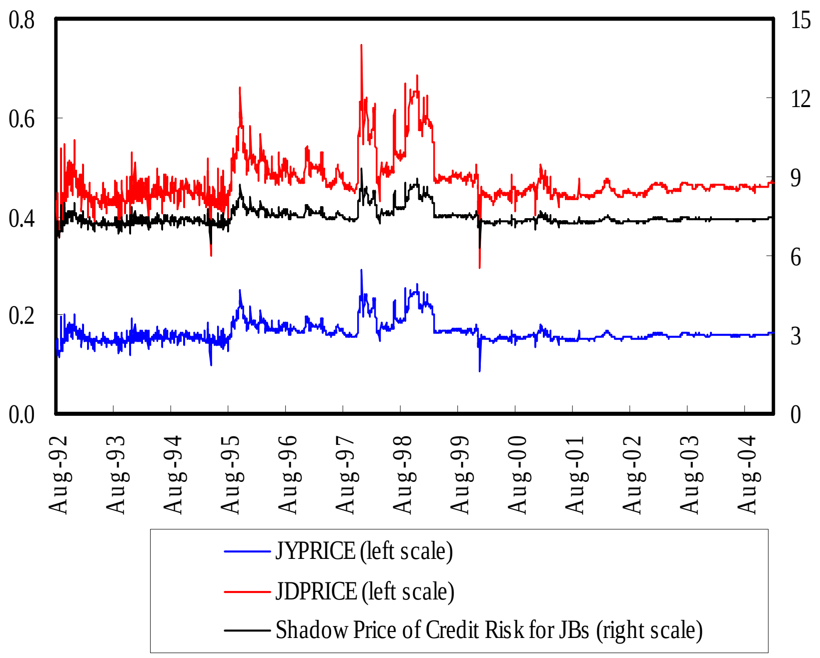
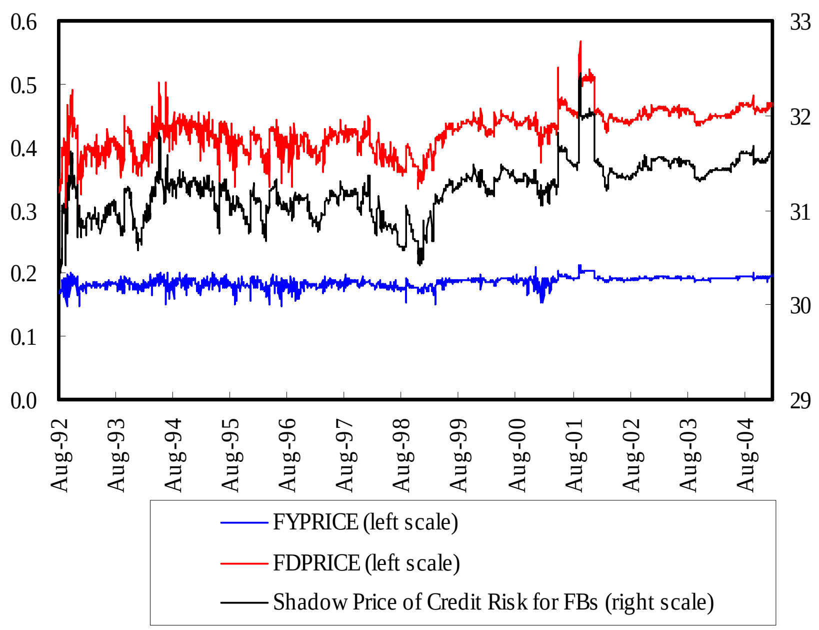


Figure 5: Shadow and Fundamental Prices of Credit Risk

(i) Japanese Banks (JBs)

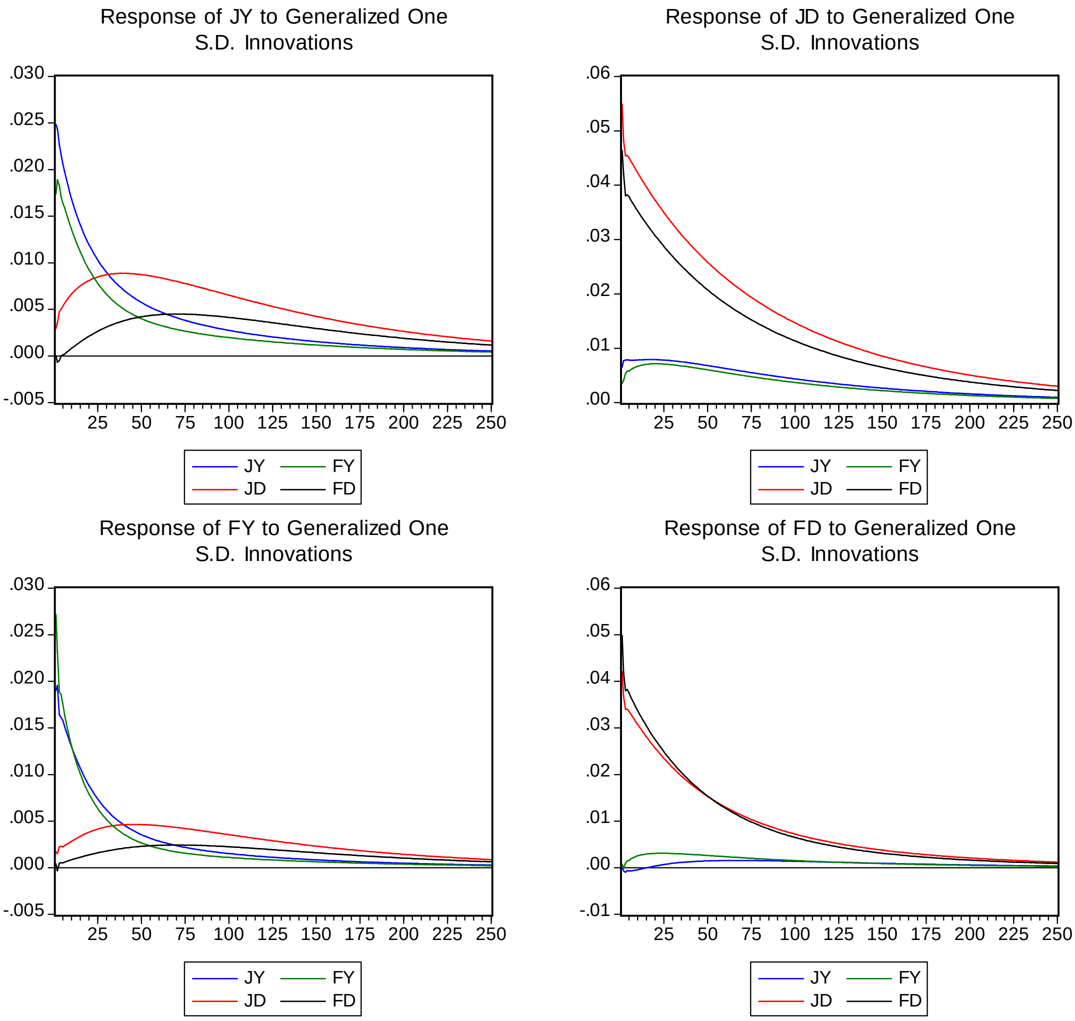


(ii) Foreign Banks (FBs)



JYPRICE	: Fundamental Price of Credit Risk for Japanese Banks in the Yen Market
JDPRICE	: Fundamental Price of Credit Risk for Japanese Banks in the Dollar Market
FYPRICE	: Fundamental Price of Credit Risk for Foreign Banks in the Yen Market
FDPRICE	: Fundamental Price of Credit Risk for Foreign Banks in the Dollar Market

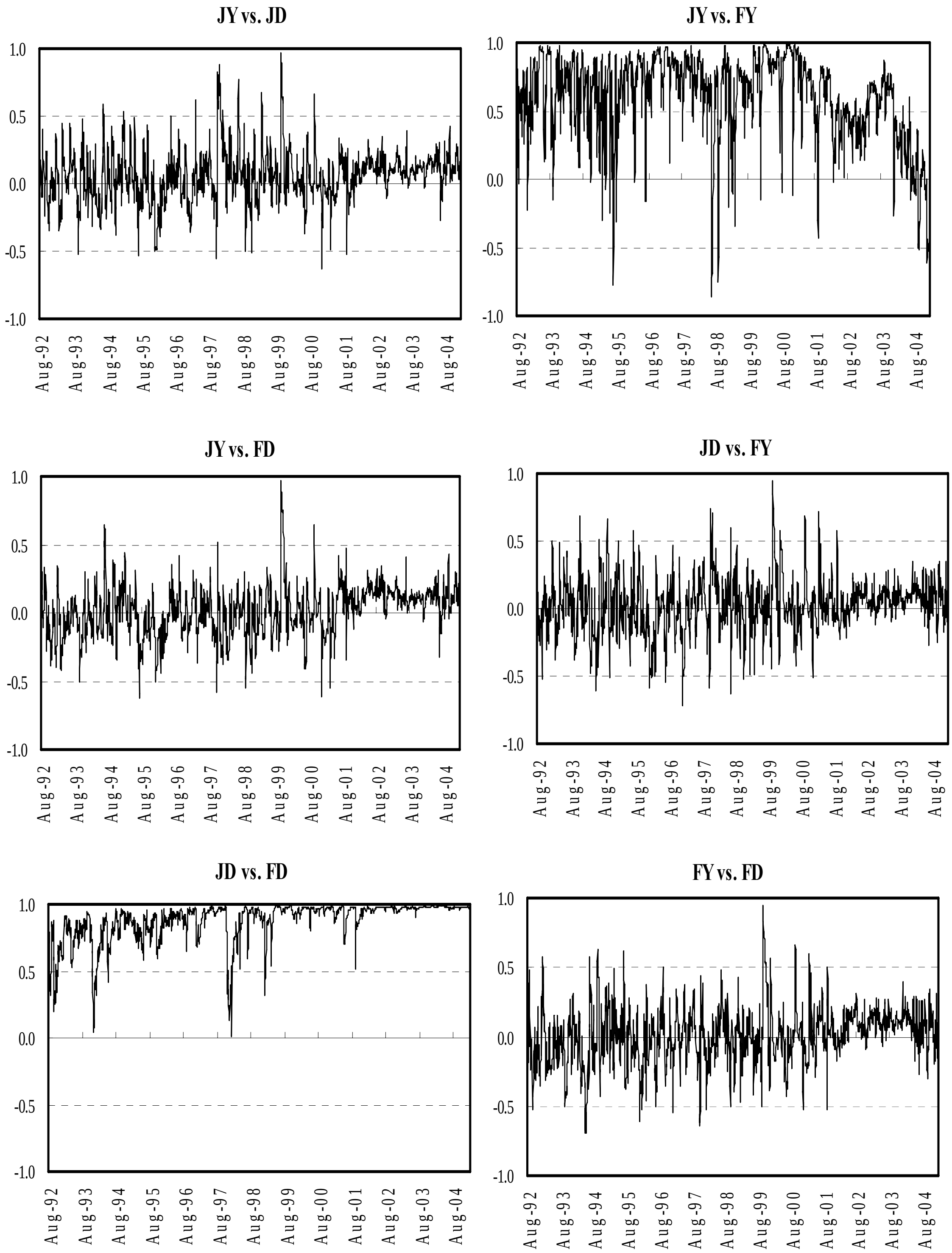
Figure 6: Generalized Impulse Response Functions (i)



JY : Japanese Banks' Yen Risk Premium JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium FD: Foreign Banks' Dollar Risk Premium

Note: Impulse response functions are based on the estimation results of the VAR model reported in Table 4(i).

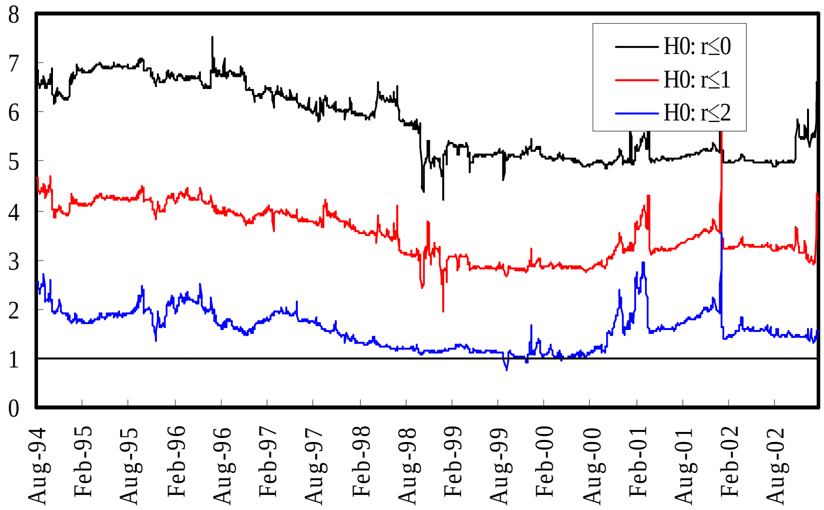
Figure 7 : Conditional Correlations by M-GARCH Model (i)



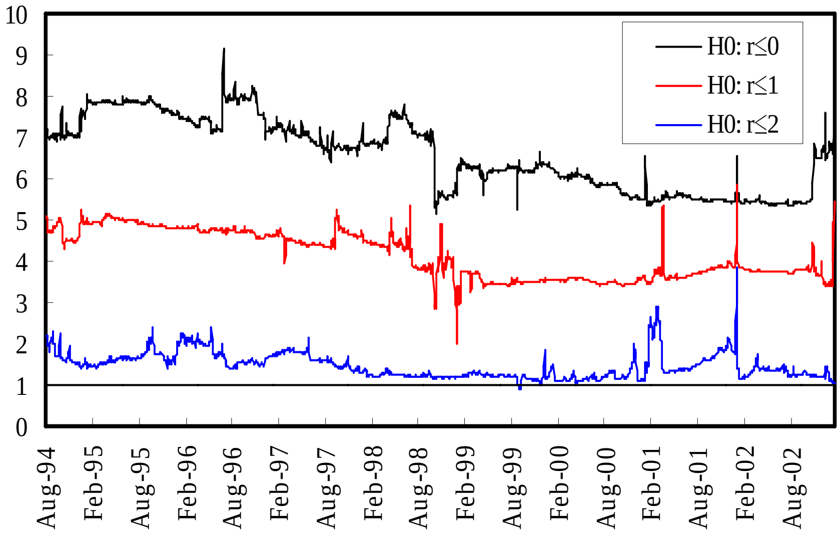
JY : Japanese Banks' Yen Risk Premium JD : Japanese Banks' Dollar Risk Premium
FY: Foreign Banks' Yen Risk Premium FD: Foreign Banks' Dollar Risk Premium

Figure 8: Stability Test of Cointegrating Relationships

(i) Trace Statistic



(ii) Maxmum Eigenvalue Statistic

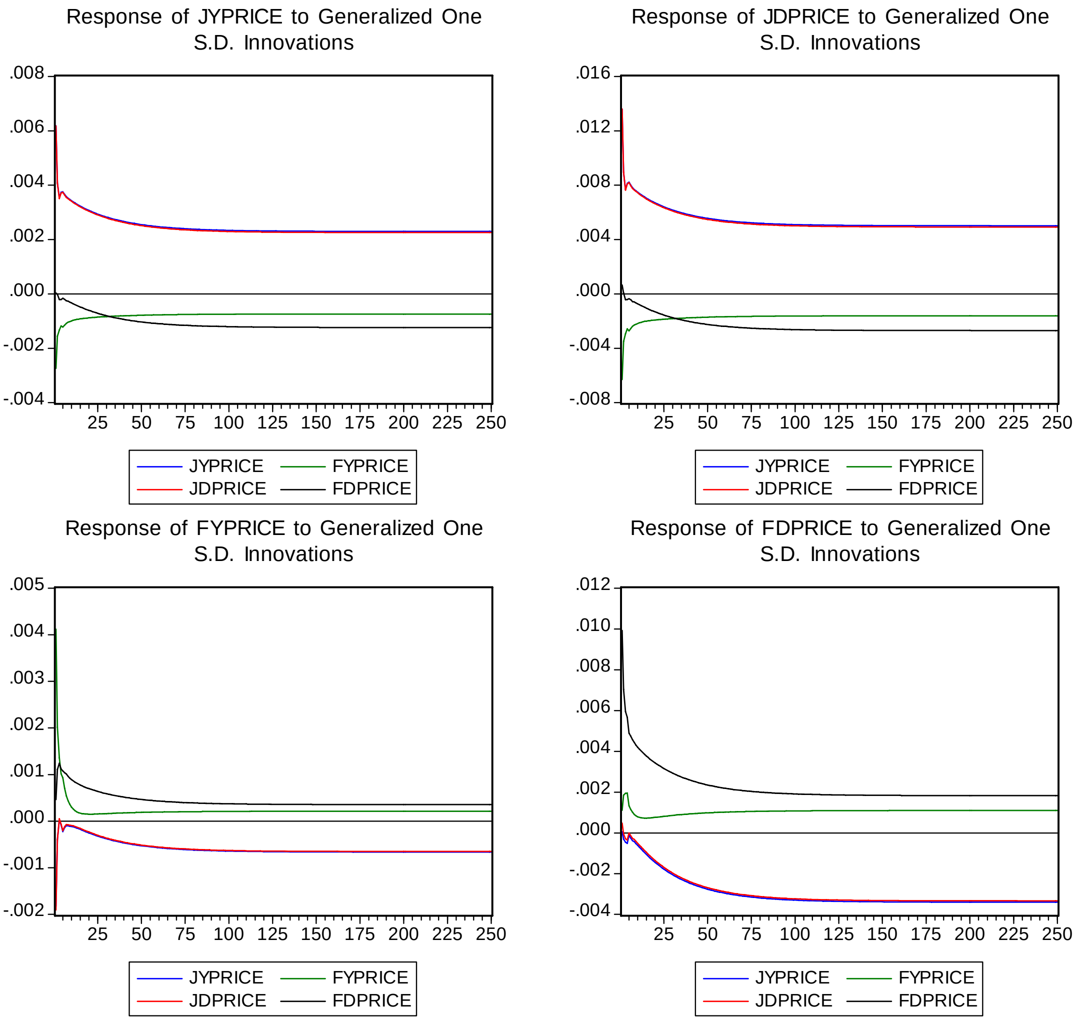


Notes: 1. r denotes the number of cointegrating ranks.

2. Test statistics are divided by the critical values that correspond to the 5% significance level. Thus, the statistics above 1 means that null hypotheses H0 can be rejected at the 5% significance level.

3. Time scale corresponds to the mid-period of the rolling window (1,000 observations).

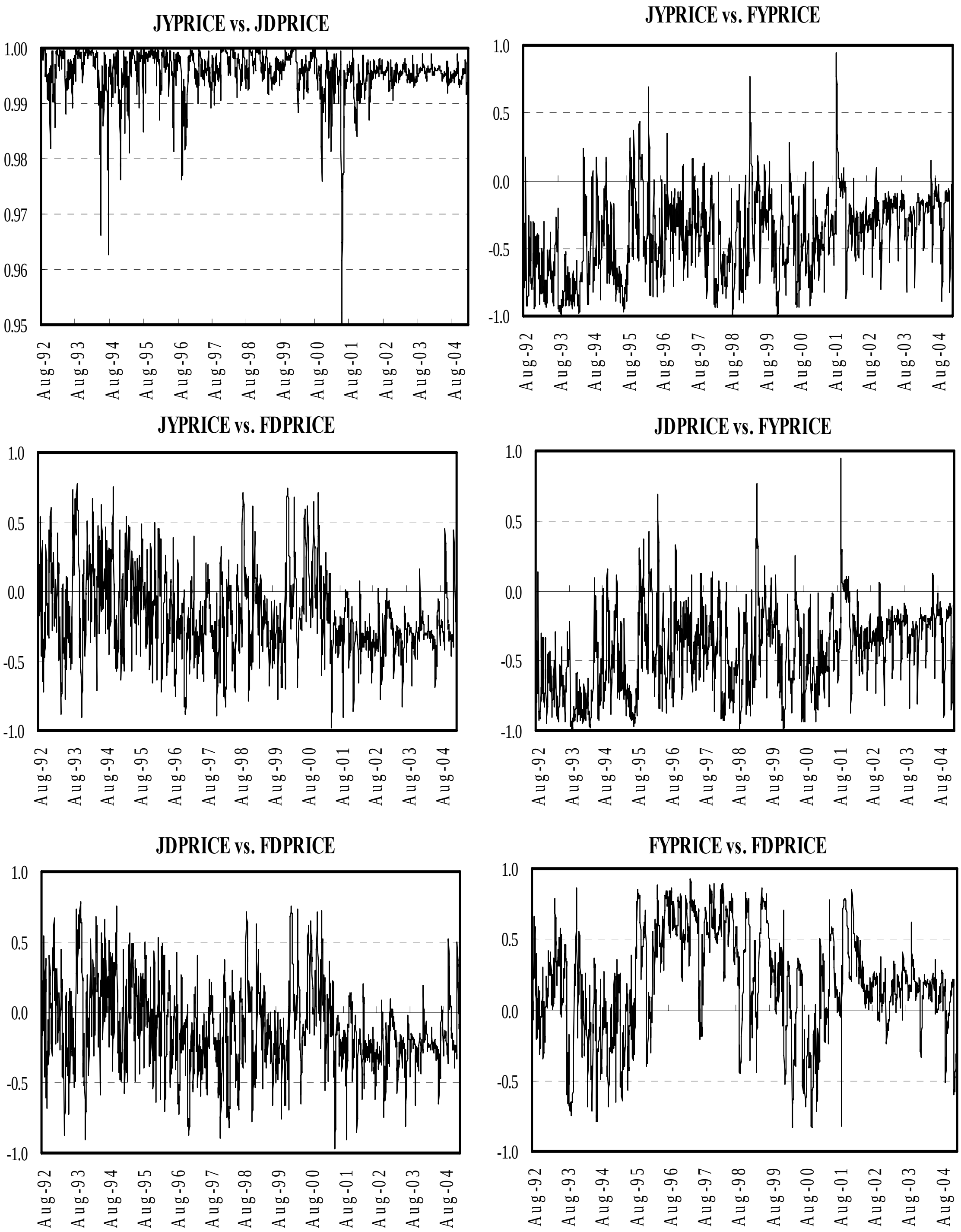
Figure 9: Generalized Impulse Response Functions (ii)



JYPRICE	: Fundamental Price of Credit Risk for Japanese Banks in the Yen Market
JDPRICE	: Fundamental Price of Credit Risk for Japanese Banks in the Dollar Market
FYPRICE	: Fundamental Price of Credit Risk for Foreign Banks in the Yen Market
FDPRICE	: Fundamental Price of Credit Risk for Foreign Banks in the Dollar Market

Note: Impulse response functions are based on the estimation results of the error correction model reported in Table 9(i).

Figure 10 : Conditional Correlations by M-GARCH Model (ii)



JYPRICE	: Fundamental Price of Credit Risk for Japanese Banks in the Yen Market
JDPRICE	: Fundamental Price of Credit Risk for Japanese Banks in the Dollar Market
FYPRICE	: Fundamental Price of Credit Risk for Foreign Banks in the Yen Market
FDPRICE	: Fundamental Price of Credit Risk for Foreign Banks in the Dollar Market