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Default Intensity and Expected Recovery of Japanese Banks and "Government":

New Evidence from the CDS Market

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Default Intensity and Expected Recovery of Japanese Banks and "Government": New Evidence from the CDS Market

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Abstract

Using term structure data of Credit Default Swap (CDS) spreads for the four Japanese mega-banks and the government, we jointly estimate the default intensity and expected recovery (loss) given a default. In doing so, we attempt to further identify the difference in the expected recovery ratios between senior and subordinated CDS contracts. Estimation results are summarized as follows. (i) The default intensities for the banks and the government substantially rose in times of a banking crisis since the late 1990s. (ii) The expected recovery ratios for subordinated CDS contracts are significantly smaller than those for senior CDS contracts, ranging from 46 to 85 percent of those for senior CDSs, depending on the banks. (iii) Each bank's default intensity is significantly cointegrated with, and reacts to, the Japanese government's default intensity. This result implies that a systemic risk factor among Japanese major banks is closely related to the default intensity of the Japanese government.

Key Words: Credit Default Swap, Japanese Banks, Sovereign CDS, Subordinated CDS, Loss Given Default

JEL Classifications: G12, G21

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1. Introduction

This paper attempts to evaluate the market perceptions about the creditworthiness of the four Japanese mega-banks and the Japanese government using the term structure data of Credit Default Swaps (CDSs). In doing so, we pay particular attention to the following three issues, each of which is addressed for the first time in this paper to the best of our knowledge. First, we attempt to jointly estimate the default intensity and expected recovery (loss) ratio given a default for all these entities. Second, we attempt to further identify the difference in the expected recovery ratios between senior and subordinated CDS contracts. Third, we attempt to investigate the "systemic" nature of bank credit risk and its relationship to the government, by analyzing the long-term and short-term relationships of the estimated default intensities between the four mega-banks and the government.

In the past several years, the market for credit derivatives has significantly developed. Among them, CDSs are the most commonly used products. In a CDS contract, the protection-buyer pays the seller a fixed premium each period, until either pre-specified credit events, typically a default, occur to the reference entity, or, the swap contract matures. In return, if the credit events occur, the protection-seller is obliged to buy back from the buyer the bond at its face value. Thus, the CDS premiums or spreads provide a direct measure of the default probability perceived by market participants. Since the principal is not needed for trading CDSs, given the nature of derivatives, CDS contracts have traded much more frequently and thus the liquidity of CDS market has been much higher than traditional straight bonds issued by the same reference entities.[1]

As pointed out by Ito and Harada [2004], main reference entities in the Japanese CDS market are the mega-banks. Due to the recent expansion of CDS trading for Japanese

[1] The market liquidity of Japanese straight bonds, not to mention the liquidity of subordinated bonds, is extremely low, compared with CDS contracts. It is because (i) Japanese straight bond investors tend to "buy-and-hold" those bonds, and (ii) there has been no repo market (transactions with repurchase agreements) for corporate bonds in Japan.

banks, coincident with a broadening investor base, CDS spreads are now considered to reflect the credit risks of Japanese banks much more sensitively than straight bond spreads and the so-called Japan premium in the money markets.[2,3]

Another noteworthy feature of the CDS market for Japanese entities is that Japanese sovereign contracts have been traded very actively. As shown by Packer and Suthiphongchai [2003], from 2000 to 2003, total number of CDS quotes for Japanese sovereign bonds amounts to 2,313, which corresponds to third place only after Brazil and Mexico.[4] This fact, along with successive downgrades of the credit rating on Japanese sovereign bonds, shows investors' deep concerns over the financial standing of the Japanese government itself facing prolonged deflation since the bursting of the bubble economy in the early 1990s, and the ensuing structural problems, such as the fragile banking system.[5]

These developments of the CDS market, together with their inherent attractive properties we explain below, provide us with a desirable way of directly measuring the size of the default component in credit spreads for the Japanese mega-banks and the government. This issue has been of great interest to many researchers, particularly from the perspective of the non-performing loan problem, but has not been thoroughly analyzed quantitatively yet. Also, when it comes to the pricing of sovereign CDSs, fewer studies exist. To the best of our knowledge, most of those that do exist analyze emerging countries' CDSs and this paper

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[2] The main protection sellers in the Japanese CDS market were non-Japanese securities companies and hedge funds. Recently, Japanese banks and institutional investors, including insurance companies and pension funds, have entered the CDS or CDS-related markets such as the Collateralized Debt Obligation (CDO) market as protection sellers, that is, credit risk investors.

[3] The Japan premium is the premium Japanese banks must pay in borrowing U.S. dollars from western banks. During the period of financial instability around 1997-98, the Japan premium reached nearly 100 bps. In the period around the end of 2001, Japanese banks again showed vulnerability, but the Japan premium never became apparent. See Ito and Harada [2004] for more details.

[4] Relative to the corporate sector, the concentration of CDS quotes on sovereign is marked. The five leading names are Brazil, Mexico, Japan, the Philippines, and South Africa during this period, which together account for more than 40 percent of listed quotes on sovereign names. See Packer and Suthipongchai [2003] for more details.

[5] Moody's lowered the credit rating on the Japanese sovereign bonds to Aa1 from Aaa in November 1998, to Aa2 in September 2000, to Aa3 in December 2001, and to A2 in May 2002. The countries with the same credit rating A2 at that time were Greece, Israel, and Botswana.

is the first attempt to rigorously analyze the sovereign CDS contracts for the Japanese government.[6]

In doing so, we attempt to estimate the recovery ratios given a credit event expected by market participants, jointly with the default intensities. Conceptually, the use of CDS spreads enables us to separately identify the default intensity and the recovery ratio for the same reference entity, if and only if CDS spread data with more than one maturity are available. This property is called the "fractional recovery of face value."[7] In reality, however, such data are not generally available, particularly for Japanese non-financial entities.

In the case of Japanese mega-banks' CDS contracts, although most of the senior CDS contracts have a 5-year maturity, many subordinated CDS contacts have maturities other than 5 years, such as 10 years. Thus, using both senior and subordinated CDS spreads enables us to estimate the term structure of CDS spreads, which leads to identifying between the default intensity and the expected recovery ratio as mentioned above. On the other hand, maturities of Japanese sovereign CDS contracts are more diverse, including 3, 5, 7 and 10 years.

An issue naturally arising from the use of both senior and subordinated CDS spreads is the identification of the expected recovery ratios between senior and subordinated CDS contracts. Under the assumption that both senior and subordinated CDS spreads for a bank have the same default intensity, but have the different expected recovery ratio, we can extract the difference in the expected recovery ratios between senior contracts and subordinated contracts. One major advantage of this strategy is to enable us to directly test whether Japanese banks' subordinated bonds are properly priced in the CDS market in terms of the difference in the recovery ratios relative to senior CDS contracts.

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[6] For instance, Pan and Singleton [2005] analyze the term structure of CDS spreads for Mexico, Russia, and Turkey, and Zhang [2003] analyzes the CDS pricing for Argentina. Pan and Singleton [2005] also show that the expected recovery ratios for those countries range from 0.56 to 0.84 using unconstrained affine models.

[7] See Duffie and Singleton [2003] for details.

In the United States and Europe, a growing number of proposals that use subordinated bond spreads to discipline banks have been set forth, on the grounds that they should more sensitively reflect the creditworthiness of the banks than senior bond spreads do.[8] By definition, a subordinated bond is unsecured debt that is junior to either other unsecured or secured debt provided by a senior lender. Thus, if a failed bank were liquidated, subordinated bond holders would receive a payment, only if all the depositors and prior debts are paid in full. Thus, it is reasonable to assume that subordinated bonds should send the most sensitive signals to the risk assumed by banks.[9]

Many studies have investigated so far whether subordinated bond holders are sensitive enough to the risks assumed by the banks in the context of U.S. banks and prudential supervisory system. Relatively earlier studies, including Avery, Belton, and Goldberg [1988], and Gorton and Santomero [1990], show that (excessive) risk-taking by bank managers were not priced in the subordinated bond spreads in the 1980s. Flannery and Sorescu [1996] further argue that no pricing of credit risk until the 1980s is due to a rational response of investors to a government's "too-big-to-fail" policy, along with well-established market perceptions of forbearance, but once such institutional framework was eliminated, subordinated bond spreads began to reflect credit risk. Their method is to analyze the responses of subordinated bond spreads to various variables that are likely to reflect credit risks, including equity prices and balance sheet data. No existing studies have used subordinated CDS spreads for banks for such a purpose, however.

Further, using the estimated default intensities for the four mega-banks and the government, we attempt to investigate the systemic nature of credit risks among Japanese major banks and its relationship to the creditworthiness of the Japanese government.

[8] See Covitz, Hancock, and Kwast [2004] for instance.

[9] Although equity is the last in priority to be paid in the event of a bank failure, it may be thought of as the best signal for market discipline. It is a well-known fact, however, that equity-holders prefer banks' managers to take more-than-optimal risk, which may expose the banks to an undesirably high risk of failure.

Japanese banks had been long protected by the government's so-called "convoy policy." Even since the beginning of the "Japanese Big Bang" deregulation in 1996, the government has kept supporting the fragile banking system by injecting capital into major banks using public funds as well as postponing lifting full protection on bank deposits.[10,11] Thus, it is of particular interest to us to investigate how CDS market participantss have evaluated the "implicit (or explicit) guarantee" given by the government in assessing the creditworthiness of Japanese banks.

Specifically, our strategy is to extract a common factor from the banks' default intensities by factor analysis and to compare the factor with the default intensity for the government. Also, we conduct a cointegration analysis between each bank's default intensity (or the common factor) and the Japanese government's default intensity to explore their long-term relationships, and derive the impulse responses of each bank's default intensity to a shock in the government's default intensity or *vice versa*.

The rest of the paper is organized as follows. Section 2 describes our CDS data set. Section 3 explains an overview of the estimation model. Section 4 reports and discusses the estimation results. Section 5 concludes the paper. Last, Appendix shows technical details about the estimation model.

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[10] The government injected capital in March 1998 and March/September 1999. In March 1998, 21 banks received public funds of 1.8 trillion yen. Also, in March 1999, the top 15 banks received public funds of 7.5 trillion yen, and in September 1999, 4 regional banks received public funds of 240 billion yen. See Ito and Harada [2006] for more details.

[11] Initially, the government planned to lift the full protection on bank deposits in March 2001. But, the government decided to postpone it, as the non-performing loan problem reemerged. In April 2002, full protection on time deposits was removed, and in April 2005, guarantees on all bank deposits were capped at 10 million yen, except for non-interest-bearing deposits in payment and settlement accounts.

2. Data

We use data on CDS contracts denominated in U.S. dollars whose reference entities are the four Japanese mega-banks, namely, Bank of Tokyo-Mitsubishi (BTM), Sumitomo-Mitsui Banking Corporation (SMBC), UFJ Bank (UFJ), and Mizuho Bank (MIZUHO), as well as the Japanese sovereign government (Japan).[12,13] Transaction frequency and thus market liquidity are viewed as the highest for these CDS contracts of all the contracts for Japanese entities' CDS contracts. Data were provided by GFI Limited, one of the leading providers of inter-dealer brokerage, market data, and software.

Our sample period is from September 1998 to December 2005 and specific dates vary depending on the availability of CDS data for each entity. The data frequency is daily, and we use all the available traded prices and the mid-prices of bid-ask spreads quoted on the same date. The number of observations and maturity structure by entity and type of reference bonds, senior or subordinated, are shown in Table 1 (i). Note here that most of the senior bank CDS contracts have a 5-year maturity, while subordinated CDS contracts include those with other maturities, typically 10 years. On the other hand, Japanese sovereign CDS contracts have much more diverse maturities, including 3, 5, 7, and 10 years. As will be explained later, identification between the default intensity and the expected recovery (loss) given a default is possible, if and only if CDS spread data with more than one maturity are available. Thus, using both senior and subordinated CDS spreads enables us to jointly estimate the default intensity and the expected recovery (loss) ratio for each bank.

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[12] Most of the CDS contracts for Japanese major banks and the government are denominated in U.S. dollars for the following reasons. First, those CDS contracts were originally signed between non-Japanese counterparties. Second, CDS market participants have been concerned about the correlation between the value of the yen and reference entities. Their perception is that if credit events should happen to Japanese major banks and the government, then its effects would be catastrophic, and thus the yen would be significantly depreciated.

[13] SMBC, UFJ, and Mizuho Bank were established as a result of their respective mergers during our sample period. Before the mergers, we use the data on Sumitomo Bank for SMBC, Sanwa Bank for UFJ, and Fuji Bank for Mizuho Bank. BTM and UFJ were merged on January 1, 2006. Since our sample period is through December 2005, we do not need to consider this merger.

Table 1 (ii) report summary statistics of CDS spreads with all maturities and Figure 1 displays 5-year CDS spreads. In most of the period, spreads for MIZUHO and UFJ are much wider than others. The rest can be listed in the descending order of SMBC, BTM, and Japan. This order is very realistic in terms of the financial standing of each bank. Also, two large spikes are observed in 1998, and, around 2001 to 2003, for CDS spreads. In both periods, the instability of the Japanese banking system received a great deal of attention. More specifically, in November 1997, concerns over the financial stability mounted following a series of failures of financial institutions.[14] The strong concerns had persisted until two major banks were nationalized.[15] And then, around 2001 to 2002, the vulnerability of Japanese banks became heightened again, due mainly to their low earnings and newly emerging nonperforming loans this time, as described by Ito and Harada [2004].[16] Further, it should be noted that the CDS spreads for all the four banks tend to move together, and have been much tightened since 2004.

Next, Table 1 (iii) reports the summary statistics of the bid-ask spreads with all maturities and Figure 2 shows those for 5-year CDS contracts. Although standard deviations of bid-ask spreads are higher for subordinate CDS contracts than those for senior CDS contracts, their respective means show a far narrower difference. This result suggests that market liquidity is also high for subordinated CDS contracts. Also, Figure 2 shows that bid-ask spreads experienced substantial spikes in 1998 and around 2001 to 2003, as is the case with the CDS spreads shown in Figure 1. Thus, we can infer that in stressful periods, during which the perception of a banking crisis received a great deal of attention, the CDS market liquidity was substantially reduced.

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[14] The following financial institutions failed in this period: Sanyo Securities (November 3), Hokkaido Takushoku Bank (November 17), Yamaichi Securities (November 24), and Tokuyo City Bank (November 26).

[15] They are Long-Term Credit Bank of Japan (October 23, 1998) and Nippon Credit Bank (December 13, 1998), respectively.

[16] Also, in this period, a substantial decline in the overall stock prices in Japan triggered concerns over the vulnerability of banks holding large stock portfolios.

3. Model

3.1 Basic Pricing Structure of CDS Contracts

Let M denote the maturity of the CDS contract, $CDS_t(M)$ the annualized spread at issue, R^Q ($L^Q \equiv 1-R^Q$) the expected constant risk-neutral fractional recovery (loss) ratio of face value on the underlying bond given a credit event, λ^Q the risk-neutral arrival rate of a credit event (default intensity), and r_t the risk-free interest rate. Following Duffie and Singleton [2003], a CDS contact with quarterly premium payments can be priced as

$$\begin{aligned} &\frac{1}{4}CDS_t(M)\sum_{j=1}^{4M} E_t^Q\left[\exp\left\{-\int_t^{t+\frac{1}{4}j}\left(r_s+\lambda_s^Q\right)ds\right\}\right] \\ &\qquad = \left(1-R^Q\right)\int_t^{t+M} E_t^Q\left[\lambda_u^Q \exp\left\{-\int_t^u \left(r_s+\lambda_s^Q\right)ds\right\}\right]du \end{aligned} . \quad (1)$$

The left hand side of equation (1) indicates the present value of the protection-buyer's premiums, which are payable contingent upon a credit event that has not occurred yet. The right hand side of equation (1) is the present value of the contingent payment by the protection-seller upon a credit event.[17] CDS spreads are priced at the point, where these two values are equalized by each other.

3.2 Affine Diffusion Model

Following Pan and Singleton [2005], we use the following square-root diffusion process for the default intensity:

$$d\lambda_t^Q = \kappa^P\left(\theta^P - \lambda_t^Q\right)dt + \sigma^Q\sqrt{\lambda_t^Q}\,dB_t\,, \quad (2)$$

$$\eta_t = \frac{\delta_0}{\sqrt{\lambda_t^Q}} + \delta_1\sqrt{\lambda_t^Q}\,, \quad (3)$$

[17] Ideally, we should treat the recovery ratio as a time-varying parameter in light of the evidence of a negative correlation between the default intensity and the recovery ratio over the business cycle, as argued by Altman, et al. [2003], for instance. This treatment, however, imposes a serious identification problem between the default intensity and the expected recovery ratio. Thus, we follow the convention that treats the expected recovery ratio as a constant parameter.

where B is a standard Brownian motion, and η_t is the market price of risk.[18] Each coefficient has a natural interpretation: θ is the long-run mean of λ, κ is the mean rate of reversion to θ, and σ^Q is a volatility coefficient. Throughout the paper, super-scripts Q and P denote risk-neutral and actual measures, respectively. This process allows for both mean reversion and conditional heteroskedasticity in CDS spreads, and guarantees the non-negativity of the default intensity process. Given this specification for market price of risk, λ_t^Q follows a square-root process under both P and Q measures. Under Q, we can write

$$d\lambda_t^Q = \kappa^Q\left(\theta^Q - \lambda_t^Q\right)dt + \sigma^Q\sqrt{\lambda_t^Q}\,dB_t^Q, \tag{4}$$

where $\kappa^Q = \kappa^P + \delta_1\sigma^Q$ and $\kappa^Q\theta^Q = \kappa^P\theta^P - \delta_0\sigma^Q$ hold.

As emphasized by Pan and Singleton [2005], specification (3) allows for κ and $\kappa\theta$ to differ across P and Q. Under the current setting, η_t can change signs over time.[19] This treatment is motivated by the arguments in Duffee [2002], and Dai and Singleton [2002], who discuss the importance of time-varying signs for market prices of risk in the context of the term structure models of default-free bonds. This specification rules out arbitrage when $\delta_0 \le \left(k^P\theta^P\right)/\sigma^{Q2} - 0.5$ holds.[20] Thus, in what follows, we estimate the model by imposing this condition.

Further, under this setting, we can calculate both risk-neutral and pseudo-actual survival probabilities in a closed form. We will later explore the risk premiums arising from unpredictable variations over time on λ_t^Q by taking advantage of this feature.

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[18] Equation (2) is the so-called CIR process, named after Cox, Ingersoll, and Ross [1985].
[19] Under the standard setting where δ_0 =0, the sign of η_t is fixed by the sign of δ_1.
[20] For details, see Cheridito, Filipovic, and Kimmel [2003] and Collin-Dufresne, Goldstein, and Jones [2004].

3.3 Kalman Filter Setup

Discretizing equation (2) gives the following transition equation:

$$\lambda_{t+h}^{Q} = \mathrm{E}\left[\lambda_{t+h}^{Q} \middle| \lambda_{t}^{Q}\right] + \eta_{t+h}, \quad \mathrm{Var_t}(\eta_{\mathrm{t+h}}) = \mathrm{Var}\left(\lambda_{t+h}^{Q} \middle| \lambda_{t}^{Q}\right). \tag{5}$$

We follow De Jong [2000] for expressions of the conditional expectation and variance in (5).[21] Here, let $CDS_{j,t}$ denote an $N_{j,t}$-dimensional vector of the observed CDS spreads for *j*-th entity at time *t*. The measurement equation for $CDS_{j,t}$ is then given by

$$CDS_{j,t+h} = z\left(\lambda_{j,t+h}^{Q}\right) + \varepsilon_{j,t+h}, \ \mathrm{Var}_t\left(\varepsilon_{j,t+h}\right) = H_{j,t}, \tag{6}$$

where $z\left(\lambda_{j,t+h}^{Q}\right)$ is a function that relates the CDS spreads to the default intensity, and $\varepsilon_{j,t+h}$ is a measurement error vector. $\varepsilon_{j,t+h}$ is assumed to be normally distributed with mean zero and standard deviations $\sigma_{\varepsilon}\left|Bid_{j,t} - Ask_{j,t}\right|$, where $Bid_{j,t} - Ask_{j,t}$ is a bid-ask spread and $\sigma_{j,\varepsilon}$ is a constant to be estimated that measures volatility in units of bid-ask spreads.[22] Using this specification for standard deviations, we aim to control for the time-varying effects of market liquidity on CDS spreads.

As shown in equation (1), $z\left(\lambda_{j,t+h}^{Q}\right)$ is a non-linear function. As proposed by Duffee [1999], a Taylor approximation of $z\left(\lambda_{j,t+h}^{Q}\right)$ around the one-period forecast of $\lambda_{j,t+h}^{Q}$ is used to linearize the model. The likelihood function is constructed following De Jong [2000]. The default intensity process is not assumed to be stationary. As in Duffee [1999], a least squares approach is used to extract an initial distribution. In estimation, the risk-free interest rate, and thus, the term structures are assumed to be constant for computational ease.[23]

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[21] For more details of the Kalman filter setup, see Appendix 1.

[22] This specification is also from Pan and Singleton [2005]. The difference in estimation method between Pan and Singleton [2005] and this paper is that they directly estimate the parameters by a method of maximum likelihood under the assumption that 5-year CDS spreads are perfectly priced without errors, while we use the Kalman filter after linearizing the pricing equation. The use of the Kalman filter in this paper is just for computational ease.

[23] Pan and Singleton [2005] also assume a constant risk-free interest. They report that estimating the model of sovereign CDS spreads under the assumption of variable risk-free interest rate (two-factor model) yields virtually identical results.

3.4 Identifying Default Intensity and Expected Recovery (Loss) Ratio

For pricing CDS contracts, we can adopt the so-called fractional recovery of face value (RFV) as shown by Duffie and Singleton [1999], in which λ^Q and L^Q play distinct roles. Specifically, pricing equation (1) takes the form:

$$CDS_t = L^Q f\left(\lambda_t^Q\right). \tag{7}$$

Thus, we can separately identify λ^Q and L^Q, if and only if the CDS spreads with more than one maturity are available. It is because the linear dependence of CDS_t on L^Q implies that the ratio of two CDS spreads with different maturities does not depend on L^Q, but does contain information about λ^Q. Most of the preceding studies and a convention among market participants give an *ad-hoc* value to L^Q and thus R^Q, in estimating the default intensity from CDS spreads. In the case of Japanese banks' CDS contracts, market participants most often assume that R_{se}^Q is 0.4 for unknown reasons. We formerly test the validity of this market convention in section 4.[24]

Furthermore, we attempt to identify the difference in the expected recovery ratios between senior and subordinated CDS contracts for each bank. The underlying assumption for the identification is that both senior and subordinated CDS contracts for the same bank have the same default intensity, but have the different expected recovery ratios.[25] Specifically, we specify its relationship as

$$R_{su}^Q = \alpha R_{se}^Q, \quad 0 \le \alpha \le 1, \tag{8}$$

where R_{su}^Q and R_{se}^Q are the expected recovery ratios for subordinated and senior CDS contracts, respectively, and α is a constant to be estimated that measures the relative riskiness between senior and subordinated CDS contracts.

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[24] Since no studies that estimate the expected recovery ratios for Japanese banks exist, we choose to test the validity of this market convention.

[25] From interviews with CDS market participants, we found it to be a market practice to assume that once a credit event occurs, it happens to both senior and subordinated CDS contracts. This practice validates our assumption that the default intensity is common to both senior and subordinated CDS contracts.

4. Estimation Results

4.1 Restricted Case: R_{se}^{Q} =0.4

Now, let us take a look at estimation results of the restricted case where R_{se}^{Q} =0.4, which market participants most often assume in pricing CDS contracts for Japanese banks. Table 2 (i) reports the ML (maximum likelihood) estimates of each parameter and the corresponding standard errors.

First, the estimates of k^{Q} and θ^{Q} are significantly positive except for Japan, but the estimates of k^{P} are insignificant for most of the cases. Also, the estimates of θ^{P} for UFJ and MIZUHO are insignificant. This result implies that default intensities under Q, λ^{Q}'s, are significantly mean-reverting for all banks, but they are not under P. Second, α, the proportionality constant between R_{se}^{Q} and R_{su}^{Q}, are estimated to be very close to zero and insignificant for each bank.

4.2 Unrestricted Case

4.2.1 Parameter Estimates and Overall Performance

Next, let us take a look at the unrestricted case where the default intensity and the expected recovery ratio are jointly estimated for each entity. Table 3 (i) reports the ML estimates of the parameters and the corresponding standard errors. Unlike the restricted case, most of the parameter estimates are significant at the 1% level, particularly for BTM, SMBC, and UFJ. Let us look at more details of the parameter estimates.

First, the estimates of k^{Q} are negative for all entities, which implies that default intensities λ^{Q}'s are explosive under Q. Under the actual measure P, on the other hand, the estimates of k^{P} are significantly positive except for MIZUHO. Thus, λ^{P}'s are mean-reverting under actual measure P.[26] The large differences between k^{Q} and k^{P}

[26] Pan and Singleton [2005] report a similar tendency for sovereign CDSs for Mexico, Russia, and Turkey.

are suggestive of the large market risk premiums arising from uncertainty about the future arrival rates of credit events. We closely investigate this issue later in this section. Second, the estimates of θ^Q and θ^P for Japan are very close to zero, and the estimates of σ^Q and σ_ε are significantly positive. These results suggest that the default intensity for the Japanese government is almost negligible in normal times, but significantly fluctuates in times of stress.[27]

Also, two parameters of the expected recovery ratios R_{se}^Q and α are estimated in the range between 0 and 1, significantly different from both 0 and 1, respectively. This result implies that subordinated CDS contracts for Japanese mega-banks are priced significantly higher than senior CDS contracts are in terms of the expected recovery (loss) ratios. Further, for all banks, estimated recovery ratios for senior CDS contracts are significantly larger than the market practice of 0.4. Figure 3 displays the estimates of expected recovery ratios of both senior and subordinated CDS contracts for each bank. Looking at the relative scale of the expected recovery ratios across the banks, the following two tendencies are worth noting. First, relatively high creditworthy banks, BTM and SMBC, have relatively low expected recovery ratios, while relatively low creditworthy banks, UFJ and MIZUHO, have relatively high expected recovery ratios. Second, the higher creditworthy the bank is, the larger difference arises in the expected recovery ratios between senior and subordinated CDS contracts. We try to interpret these results later in this paper.

Next, Table 3 (ii) reports the absolute value of pricing errors. Note here that the means and standard errors of the pricing errors derived from the unrestricted model are very small both in absolute and relative terms. Specifically, the mean pricing errors of the unrestricted model correspond to only 4-7 percent of mean CDS spreads reported in Table

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[27] Note that we assume pricing errors are normally distributed with mean zero and standard deviations $\sigma_\varepsilon \left| Bid_{j,t} - Ask_{j,t} \right|$. In times of stress, bid-ask spreads tend to substantially widen as shown in Figure 2.

1 (ii), and are much smaller than those of the restricted model reported in Table 2 (ii). In particular, note that pricing errors for subordinated CDS contracts are substantially reduced in the unrestricted model. Also, Table 3 (iii) reports the result of Likelihood Ratio (LR) test whose null hypothesis is R_{se}^{Q} =0.4.[28] The null hypothesis is significantly rejected at the 1% level for all entities. Thus, we can statistically confirm that performance of the unrestricted model is much better than the restricted model.

4.2.2 Estimated Default Intensity

Figure 4 exhibits the estimated default intensity under Q and the term structure of CDS spreads for each entity. Evidently, each bank experienced two large spikes in the default intensity, in 1998 and, around 2001 to 2003. In both periods, the instability of the Japanese banking system became heightened as we mentioned in section 2. However, the magnitude of the spikes significantly differs across the banks. BTM shows a much smaller default intensity than the others, while UFJ and MIZUHO show a higher default intensity.

Another interesting point here is that the default intensity for the Japanese government shows a movement very similar to those for each bank, although its level is much smaller. One noticeable difference between the government and each bank is that the default intensity for the government is much higher in the second round of financial instability, around 2001 to 2003, than in the first round in 1998. During the second round, the Japanese financial system became vulnerable again, with the Nikkei 225, one of the major overall stock price indices in Japan, having slumped to around the 10,000 mark. Under such circumstances, the active role of the government in revitalizing the Japanese banking system was pressed hard from the political side, which is likely to create a very close link in the default intensity between the Japanese government and each bank.[29]

[28] The LR test is calculated as minus two times the difference in the likelihood ratio between the restricted and unrestricted models. It is distributed as a chi-square with four degrees of freedom.
[29] As we mentioned in section 1, Moody's lowered credit rating on Japanese sovereign bonds to A2 in

Also, the term structures of CDS spreads are estimated to be upward sloping for all entities. Generally speaking, the term structure of credit spreads based on an exogenously specified risk-neutral default intensity process starts at a non-zero spread and rises gradually.[30] Our estimates follow this pattern. During the periods of financial instability, the curve substantially shifted upward and steepened. Quite recently, the curves have almost completely flattened, reflecting the market sentiments that Japanese mega-banks have almost revived, since they have finished disposing of their non-performing loans.

4.2.3. Risk-Neutral vs. Pseudo-Actual Measures of Survival Probabilities

Another interesting issue lies in the difference in the survival probabilities between the risk-neutral measure Q and actual measure P. First, under the reduced-form pricing framework we adopted in this paper, the risk-neutral survival probability for time horizon M is given by

$$S_t(M) \equiv E_t^Q\left[\exp\left(-\int_t^{t+M} \lambda^Q(u)du\right)\right]. \tag{9}$$

Next, the pseudo-actual survival probability can be calculated as

$$PS_t(M) \equiv E_t^P\left[\exp\left(-\int_t^{t+M} \lambda^Q(u)du\right)\right]. \tag{10}$$

As shown in Longstaff, Mithal, and Neis [2005] and Duffie, Pan, and Singleton [2000], we can calculate both survival probabilities in a closed form, given the square-root dynamics for the default intensity λ^Q.[31] As stated by Duffie and Singleton [2003] and Pan and Singleton [2005], if market participants are, indeed, risk-neutral toward the risk arising from a variation over time in λ^Q, then the pseudo-actual survival probability given by equation (10)

May 2002, which was equivalent to the rating on Greece, Israel, and Botswana.

[30] On the other hand, the term structure of credit spreads implied by structural models like the first-passage model takes a strikingly different shape. The credit spreads are almost zero for short maturities, and then rapidly increases over the first several years of maturity. See Duffie and Singleton [2003] for more details about the comparison of the term structure of model-implied credit spreads.

[31] See Appendix 2 for more details.

coincides with the risk-neutral survival probability (9). Put differently, a comparison between $S_t(M)$ and $PS_t(M)$ gives us an assessment of the quantitative importance of the risk premiums arising from an unpredictable variation over time in λ^Q.[32]

Figure 5 displays $S_t(M)$ and $PS_t(M)$, and the difference between these two measures. As expected, the differences between $S_t(M)$ and $PS_t(M)$ are particularly large in the periods of financial instability for all entities. Also, note that during normal periods other than the periods of financial instability, the differences between $S_t(M)$ and $PS_t(M)$ are substantially narrowed. Toward the end of the sample period, in particular, the differences are even negative for all entities. This observation suggests that the risk premiums associated with the uncertainty about the risk-neutral arrival rate of credit events have existed only in times of stress.

4.3 Interpreting Estimated Recovery Ratios

4.3.1 Relationship between CDS Default Intensity and Expected Recovery Ratio

In what follows, we attempt to interpret the relative magnitude of the estimated default intensities and the expected recovery ratios across the banks shown in Figure 3. More specifically, the questions we address here are two-fold: (i) "why do the banks with relatively low default intensities have low expected recovery ratios, while the banks with relatively high default intensities have high recovery ratios?" and (ii) "why do banks with low default intensities have large differences in the estimated recovery ratios between senior and subordinated CDS contracts?"

Regarding emerging countries' sovereign CDS contracts, Duffie, Pedersen, and Singleton [2003] and Pan and Singleton [2005] interpret the default intensity λ^Q as the sum

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[32] It should be noted, here, that what we do estimate is not the historical survival probabilities $E_t^P\left[\exp\left(-\int_t^s \lambda^P(u)du\right)\right]$. To that end, we need data on the physical intensity λ^P, which cannot be extracted from CDS spread data alone. See Jarrow, Lando, and Yu [2005], and Yu [2002] for details.

of the arrival intensities of the following credit events for sovereign debts specified by the ISDA:[33] (i) acceleration, (ii) failure to pay, (iii) restructuring, and (iv) repudiation. As shown by Lando [1998], in a doubly stochastic setting, the probability that any two of the credit events simultaneously occur is zero. As a result, L^Q can be formulated as the average of the four expected loss ratios corresponding to each credit event, weighted by each arrival intensity.

Since our main focus is on bank CDS contracts, we think it more natural to conceptually break down the default intensities into two risk categories rather than institutionally specified events like ISDA terms of credit events: (i) the risk that is common to all the four banks, and (ii) the risk that is specific to each bank. The first one is closely related to the systemic risk and the second one is to a more structural or idiosyncratic risk.[34] Thus, in our setting, λ^Q and L^Q are decomposed as

$$\lambda^Q = \lambda_c^Q + \lambda_s^Q \text{ , and } L^Q = \frac{\lambda_c^Q}{\lambda^Q} L_c^Q + \frac{\lambda_s^Q}{\lambda^Q} L_s^Q , \qquad (11)$$

where λ_c^Q and λ_s^Q denote the arrival intensity of common and specific risks, respectively, and L_c^Q and L_s^Q denote the corresponding expected loss ratios. λ_c^Q, λ_s^Q, L_c^Q, and L_s^Q are allowed to differ across the banks.

Figure 6 shows a simulated path of L^Q as a function of λ_s^Q in the hypothetical case where λ_c^Q =0.05, L_c^Q =0.50, and L_s^Q =0.25. We assume, here, that the expected loss ratio given a systemic event L_c^Q is larger than the expected loss ratio given a specific event L_s^Q. In such a case, L^Q is found to be a decreasing (increasing) function of the arrival intensity of the default risk specific to each bank λ_s^Q (the ratio of the default risk common to each bank: λ_c^Q / λ^Q). This exercise implies that, as the risk component that is specific to a particular bank λ_s^Q rises relative to the systemic risk component λ_c^Q, the

[33] ISDA is the International Swaps and Derivatives Association, which is the global trade association representing participants in the privately negotiated derivatives industry.
[34] The definition of a "systemic event" is diversified so much in literature that we do not intend to further step into it. See De Bandt and Hartmann [2000] for a detailed survey of systemic risk.

total expected loss ratio L^Q for the bank falls. The only assumption for this result is that the expected loss ratio given a systemic event is larger than the loss ratio given the specific event, thus $L_C^Q > L_S^Q$. Is this assumption realistic enough? Yes, if the systemic event is catastrophic such that, although its probability is very low, once it occurs, the Japanese government itself is likely to be heavily damaged.[35] In such a case, the government is not likely to afford the total cost arising from systemic failure of the four mega-banks, which leads to the higher expected loss ratios than in the case of specific events.

Next, let us explicitly allow for the differences in the expected recovery (loss) ratios between senior and subordinated CDS contracts. In what follows, we ignore the superscript Q for notational ease. The expected loss ratios for senior and subordinated contracts, L^{se} and L^{su}, respectively, can be written as

$$L^{se} = \frac{\lambda_C}{\lambda} L_C^{se} + \frac{\lambda_S}{\lambda} L_S^{se} \text{ and } L^{su} = \frac{\lambda_C}{\lambda} L_C^{su} + \frac{\lambda_S}{\lambda} L_S^{su}. \tag{12}$$

In what follows, we examine the hypothetical case where λ_C=0.05, L_C^{se}=0.50, L_S^{se}=0.25, L_C^{su}=0.90, and L_S^{su}=0.25. This assumption is meant to capture the following market perceptions. In the case of a specific event, the government can afford the cost of redeeming even the subordinated bonds, in addition to the senior bonds, and thus L_S^{se} and L_S^{su} take on the same low values. This assumption is consistent with the Japanese situation where under the current law, both senior and subordinated bonds issued by the banks that received the government support, such as the injection of public funds, are equally treated when it comes to redemption. This assumption is also strongly motivated by the experience that the Japanese government have kept protecting subordinated bondholders in most of the specific bank failures that happened.[36] In the case of a systemic event, on the contrary, the

[35] In such a systemic situation, the values of collaterals, typically land, put up for loans may be substantially lost, which further damages the banking sector.

[36] Recently, when Risona Bank was nationalized in May 2003, and Ashikaga Bank failed in September 2003, the government completely protected the holders of subordinated bonds issued by both banks. In particular, the generous government support to Risona Bank gave many market participants solid grounds that the government would give a full support to mega-banks for certain, once they face a critical situation.

government is not likely to afford to protect all the debts issued by the banks, and thus L_C^{se} and L_C^{su} take on much higher values such that $L_C^{se} < L_C^{su}$.

Figure 7 shows a simulated path of L^{su}/L^{se} as a function of λ_S under this setting. Evidently, L^{su}/L^{se} falls as the specific risk component λ_S rises. This result is consistent with our empirical result of the estimated recovery ratios: the higher creditworthy the bank is, the larger difference in the expected recovery ratios arises between senior and subordinated CDS contracts.

Then, our next task is to empirically examine the relevance of the above setting. The most important underlying assumption here is whether the systemic event is catastrophic enough such that once it occurs, it is likely to heavily damage the financial standing of the Japanese government. Now that we have the default intensity for the Japanese government, we can quantitatively assess the validity of this key assumption.

4.3.2 Some Further Analysis

Specifically, our strategy can be summarized as follows. First, we extract a latent common factor from the estimated default intensities for the four banks by factor analysis, and compare it with the default intensity for the Japanese government. Second, we conduct a cointegration analysis on the relationship between each bank's and the Japanese government's default intensities, after showing that these default intensities are I(1).[37]

First, the result of factor analysis is reported in Table 4. It shows that the first factor whose factor loadings are almost equal across the four banks contributes more than 90 percent of the total variation of these default intensities.[38] Thus, it seems quite natural to call this first factor the "systemic risk (common) factor." Note that, although explanatory power

[37] Below, we replaced the missing data points with the most recent observations to facilitate comparison. As a result, the number of observations is the same (2,630) for all banks.
[38] We used the principal factor method with no rotation. We tried several methods for rotating factors, but we always gained almost identical results.

is very low, the second and third factors are likely to capture the relative differences in the default intensity among the banks. Figure 8 compares between the default intensity for the Japanese government and this common factor. Surprisingly enough, these two default risk indices are almost perfectly related to each other, with the correlation coefficient of higher than 0.95.

Second, the result of the unit root tests is reported in Table 5. We conducted both ADF (Augmented Dickey-Fuller) and PP (Phillips-Perron) tests. Both test results show that all of the default intensities and the common factor are I(1).[39] So, we proceed to a cointegration analysis. Table 6 reports the result of the Johansen cointegration test. All of the pairs between the Japanese government and each bank are found to be cointegrated significantly at the 1% level by both Trace and Max-eigen statistics. The estimated cointegrating vectors show that BTM has the closest relationship to the Japanese government (1, -0.447), and then SMBC (1, -0.197), UFJ (1, -0.100), and MIZUHO (1, -0.089) in that order. This result suggests that the higher creditworthy the bank is, the higher the ratio of systemic or catastrophic default intensity is, which is consistent with the hypothesis we raised above. We also conducted the same analysis between the default intensity for the Japanese government and the common factor. Here, we standardized the default intensity for the Japanese government with mean zero and standard deviation one, since the common factor is a standardized series by definition. The result shows that both series are cointegrated significantly at the 1% level with the cointegrating vector (1, -1.028), which confirms that the systemic risk among the four Japanese banks are very closely related to the market perceptions about the credit risk of the Japanese government.

Further, Figure 9 shows the generalized impulse responses derived from the error correction model with the same cointegrating vectors reported in Table 6.[40] The result

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[39] The specification reported in Table 5 includes a trend term. We also tested the specification without a trend term, and gained the same result.
[40] We derived the generalized impulse responses using the method of Pesaran and Shin [1998].

shows that a shock in the Japanese government's default intensity exerts a long-lasting strong effect on each bank's default intensity, particularly for less creditworthy banks in terms of the level of default intensity, while a shock in each bank's default intensity has a much weaker effect on the Japanese government's default intensity. Also, an interesting point to note here is that the impulse responses of the Japanese government become stronger, as the source of a shock becomes a more creditworthy bank. For instance, the impulse response of the Japanese government to BTM gradually rises and eventually becomes almost equivalent to the impulse response to the Japanese government itself. On the other hand, the impulse response of the Japanese government to MIZUHO is much lower, and stays at almost the same level over time. This result is likely to reflect the market sentiments that once a negative shock occurs to a highly creditworthy bank like BTM, it may cause a systemic effect to other banks, which also endangers the government both economically and politically. And such market sentiments, in fact, existed in the periods of a banking crisis, particularly around 2001 to 2003. Last, the shape of the impulse response of the Japanese government to the common factor lies in between BTM and SMBC, which are relatively higher creditworthy banks.

5. Concluding Remarks

This paper has evaluated market perceptions about the creditworthiness of the four Japanese mega-banks and the Japanese government using the data of newly-evolving CDS market. CDS contracts have some attractive features to this end: higher market liquidity than the straight bond market, a large number of contracts for Japanese sovereign bonds, the fractional recovery of face value, and so on.

Noteworthy points unique to this paper can be summarized as follows. First, we have jointly estimated the default intensities and the expected recovery ratios given a default for the four Japanese mega-banks and the Japanese government. Second, we have further identified the difference in the expected recovery ratios between senior and subordinated CDS contracts for each bank. Third, we have investigated the "systemic" nature of bank credit risk and its role of or the relationship to the government by analyzing the estimated default intensities of the four mega-banks and the Japanese government. All these issues in this paper are new to literature, to the best of our knowledge.

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Appendix 1: Kalman Filter Setup

This appendix describes the setup for the extended Kalman filter estimation used in this paper. Let λ_t^Q be the default intensity with a process under the actual measure P given by

$$d\lambda_t^Q = \kappa^P\left(\theta^P - \lambda_t^Q\right)dt + \sigma^Q\sqrt{\lambda_t^Q}\,dB_t \tag{A-1}$$

$$\eta_t = \frac{\delta_0}{\sqrt{\lambda_t^Q}} + \delta_1\sqrt{\lambda_t^Q} \tag{A-2}$$

With this choice of market price of risk η_t, λ_t^Q follows a square diffusion process under both P and Q. Specifically, under Q,

$$d\lambda_t^Q = \kappa^Q\left(\theta^Q - \lambda_t^Q\right)dt + \sigma^Q\sqrt{\lambda_t^Q}\,dB_t^Q, \tag{A-3}$$

where $\kappa^Q = \kappa^P + \delta_1\sigma^Q$ and $\kappa^Q\theta^Q = \kappa^P\theta^P - \delta_0\sigma^Q$. Discretizing (1) gives the following transition equation:

$$\lambda_{t+h}^Q = \mu + \Phi\lambda_t^Q + \eta_{t+h}, \tag{A-4}$$

where $\mu = \theta^P\left(1-\exp\left(-\kappa^{Ph}\right)\right)$ and $\Phi = \exp\left(-\kappa^P h\right)$. η_t is assumed to be normally distributed with mean zero and standard deviations σ_η, where

$$\sigma_\eta = \sigma^Q\sqrt{\left(\frac{1-\exp\left(-\kappa^P h\right)}{\kappa^P}\right)\left(\frac{\theta^P\left(1-\exp\left(-\kappa^P h\right)\right)}{2} + \lambda_t^Q\exp\left(-\kappa^Q h\right)\right)}. \tag{A-5}$$

Now, let CDS_t denote an N_t-dimensional vector of the observed CDS spreads at time *t*. The measurement equation for CDS_t is then given by

$$CDS_{t+h} = z\left(\lambda_{t+h}^Q\right) + \varepsilon_{t+h}, \quad \mathrm{Var}_t\left(\varepsilon_{t+h}\right) = H_t. \tag{A-6}$$

Here, $z\left(\lambda_{t+h}^Q\right)$ maps the default intensity into N_t CDS spreads. For CDS spreads, this mapping is implicitly given by numerically solving for the CDS spreads implied by the default intensity. The function $z\left(\lambda_{t+h}^Q\right)$ is nonlinear and ε_{t+h} is a measurement error vector. The matrix H_t is an $N_t \times N_t$ diagonal matrix of which *j*-th diagonal element is $\sigma_\varepsilon\left|Bid_{j,t} - Ask_{j,t}\right|$.

As in Duffee [1999], a Taylor approximation of this function around the one-period forecast of λ_t^Q is used to linearize the model and we do not assume that the default intensity processes are not stationary. Therefore, we cannot use the unconditional distribution of λ_t^Q to initiate the Kalman filter recursion. Instead, we use a least-squares approach to extract an initial distribution from the first CDS spread observation. Denote this first date as date 0. Then,

$$z\left(\lambda_0^Q\right) \approx z\left(\lambda_0^Q\right) - Z\theta^Q + Z\lambda_0^Q, \tag{A-7}$$

where Z is the linearization of z around θ^Q:

$$Z = \left.\frac{\partial z\left(\lambda_0^Q\right)}{\partial\lambda_0^Q}\right|_{\lambda_0^Q = \theta^Q} \tag{A-8}$$

Assuming that this linearization is exact, we can write the measurement equation for the first date 0 CDS spreads as

$$CDS_0 = z\left(\theta^Q\right) - Z\theta^Q + Z\lambda_0^Q + \varepsilon_0 . \tag{A-9}$$

This equation can be rewritten in terms of λ_0^Q:

$$\lambda_0^Q = \frac{Z'\left(CDS_0 - z\left(\theta^Q\right) + Z\theta^Q\right)}{Z'Z} - \frac{Z'\varepsilon_0}{Z'Z} . \tag{A-10}$$

Thus, the distribution of λ_0^Q is assumed to have mean $Z'\left(CDS_0 - z_0\left(\theta^Q\right) + Z\theta^Q\right)/(Z'Z)$ and variance $H_0/(Z'Z)$. Following De Jong [2000], given this initial distribution of unobserved default intensity, the extended Kalman filter recursion proceeds as follows.

Model:

$$CDS_{t+h} = A\left(\lambda_t^Q\right) + B\left(\lambda_t^Q\right)\lambda_{t+h}^Q + \varepsilon_{t+h}, \ \mathrm{Var}\left(\varepsilon_{t+h}\right) = H_t , \tag{A-11}$$

$$A\left(\lambda_t^Q\right) = z\left(\mu + \Phi\lambda_t^Q\right) - B\left(\lambda_t^Q\right)\left(\mu + \Phi\lambda_t^Q\right) \tag{A-12}$$

$$B\left(\lambda_t^Q\right) = \left.\frac{\partial z\left(\lambda_{t+h}^Q\right)}{\partial \lambda_{t+h}^Q}\right|_{\lambda_{t+h}^Q = \mu + \Phi\lambda_t^Q} \tag{A-13}$$

$$\lambda_{t+h}^Q = \mu + \Phi\lambda_t^Q + \eta_{t+h} . \tag{A-14}$$

Initial Conditions:

$$\hat{\lambda}_0^Q = Z'\left(CDS_0 - z\left(\theta^Q\right) + Z\theta^Q\right)/(Z'Z), \tag{A-15}$$

$$\hat{q}_0 = H_0/(Z'Z). \tag{A-16}$$

Prediction:

$$\lambda_{t|t-h}^Q = \mu + \Phi\hat{\lambda}_{t-h}^Q, \tag{A-17}$$

$$q_{t|t-h} = \Phi^2\hat{q}_{t-h} + \sigma_\eta^2 . \tag{A-18}$$

Likelihood Contributions:

$$u_t = CDS_t - A\left(\hat{\lambda}_{t-h}^Q\right) - B\left(\hat{\lambda}_{t-h}^Q\right)\lambda_{t|t-h} , \tag{A-19}$$

$$V_t = B\left(\hat{\lambda}_{t|t-h}^Q\right) q_{t|t-h} B\left(\hat{\lambda}_{t-h}^Q\right) + H_t , \tag{A-20}$$

$$-2\ln L_t = \ln\left|V_t\right| + u_t' V_t^{-1} u_t . \tag{A-21}$$

Updating:

$$K_t = q_{t|t-h} B\left(\hat{\lambda}_{t-h}^Q\right)' V_t^{-1}, \tag{A-22}$$

$$L_t = I - K_t B\left(\hat{\lambda}_{t-h}^Q\right) \tag{A-23}$$

$$\hat{\lambda}_t^Q = \lambda_{t|t-h} + K_t u_t , \tag{A-24}$$

$$\hat{q}_t = L_t q_{t|t-h} . \tag{A-25}$$

Appendix 2: Survival Probabilities

The risk-neutral survival probability for time horizon M is given by

$$S_t(M) \equiv E_t^Q\left[\exp\left(-\int_t^{t+M} \lambda^Q(u)du\right)\right]. \quad \text{(A-26)}$$

Following Longstaff, Mital, and Neis [2005], $S_t(M)$ can be calculated as

$$S_t(M) = R(M)\exp\left(T(M)\lambda_t^Q\right), \quad \text{(A-27)}$$

where

$$R(M) = \exp\left(\frac{\kappa^Q\theta^Q\left(\kappa^Q + \omega\right)}{\sigma^{Q2}}M\right)\left(\frac{1-\nu}{1-\nu\exp(\omega M)}\right)^{\frac{2\kappa^Q\theta^Q}{\sigma^{Q2}}} \quad \text{(A-28)}$$

$$T(M) = \frac{\kappa^Q - \omega}{\sigma^{Q2}} + \frac{2\omega}{\sigma^{Q2}\left(1-\nu\exp(\omega M)\right)}, \quad \text{(A-29)}$$

$$\omega = \sqrt{2\sigma^{Q2} + \kappa^{Q2}}\text{, and} \quad \text{(A-30)}$$

$$\nu = \frac{\kappa^Q + \omega}{\kappa^Q - \omega}. \quad \text{(A-31)}$$

Pseudo-actual measure of survival probabilities $PS_t(M) \equiv E_t^P\left[\exp\left(-\int_t^{t+M} \lambda^Q(u)du\right)\right]$ can be calculated in the same manner by replacing κ^Q and θ^Q with κ^P and θ^P.

Table 1: Summary Statistics

(i) Number of Observations

	Senior CDS			Subordinated CDS			Total
	5-year	Others	Sub-total	5-year	Others	Sub-total	
BTM	749	80	829	439	58	497	1,326
SMBC	699	48	747	260	299	559	1,306
UFJ	673	84	757	427	106	533	1,290
MIZUHO	571	90	661	437	60	497	1,158
Japan	706	1,009	1,715	—	—	—	1,715

Source: GFI Limited

(ii) CDS Spreads (bps)

	Senior CDS				Subordinated CDS			
	Mean	Std	Max	Min	Mean	Std	Max	Min
BTM	51.74	31.84	185.00	12.00	65.16	43.28	200.00	14.00
SMBC	65.21	38.84	202.50	12.00	91.75	72.31	330.00	23.00
UFJ	78.49	52.96	240.00	13.50	91.16	73.58	397.50	26.50
MIZUHO	86.95	58.42	305.00	11.00	88.32	81.26	407.50	13.50
Japan	21.55	10.65	59.00	3.25	—	—	—	—

Source: GFI Limited

(iii) Bid-Ask Spreads (bps)

	Senior CDS				Subordinated CDS			
	Mean	Std	Max	Min	Mean	Std	Max	Min
BTM	10.66	7.96	55.00	1.00	11.67	10.45	80.00	1.00
SMBC	12.93	9.02	60.00	0.50	17.40	19.75	110.00	1.00
UFJ	13.73	10.90	75.00	0.50	12.97	15.15	115.00	1.00
MIZUHO	16.56	15.37	160.00	1.00	12.99	18.74	140.00	0.50
Japan	3.23	1.40	12.00	0.50	—	—	—	—

Source: GFI Limited

Table 2: Estimation Results: Restricted Case (R_{se}^{Q} =0.4)

(i) Parameter Estimates

Parameter	BTM	SMBC	UFJ	MIZUHO	Japan
κ^{Q}	0.00018***	0.00012***	0.00028***	0.00011***	-0.15592*
	[0.00001]	[0.00001]	[0.00004]	[0.00000]	[0.08476]
θ^{Q}	6.00029**	9.16520**	4.48504***	10.23816*	0.00000
	[2.99103]	[4.25287]	[0.27161]	[5.29294]	[0.00223]
σ^{Q}	0.07718***	0.07335***	0.08159***	0.08416***	0.02501***
	[0.01810]	[0.02810]	[0.00455]	[0.03037]	[0.00017]
κ^{P}	0.77023	0.61207	0.49179	0.33716*	0.25880
	[0.47814]	[1.90150]	[3.22539]	[0.19578]	[0.19904]
θ^{P}	0.00409*	0.00662***	0.00715	0.01116	0.00146**
	[0.00233]	[0.00068]	[0.05286]	[0.01051]	[0.00060]
σ_{ε}	0.57620**	0.65772***	0.96166***	0.82245***	0.70254***
	[0.27085]	[0.14621]	[0.38775]	[0.23195]	[0.19177]
α	0.00000	0.00000	0.00000	0.00000	———
	[0.58143]	[0.44494]	[0.16061]	[0.59064]	———
Log likelihood	6.839	6.529	6.284	6.335	7.836
Observations	1,326	1,306	1,290	1,158	1,715

Note: Figures in parentheses are standard errors. ***, **, and * denote significance level at the 1%, 5%, and 10% level, respectively. Log likelihood is the sample average.

(ii) Absolute Value of Pricing Error (bps)

	Senior CDS				Subordinated CDS			
	Mean	Std	Max	Min	Mean	Std	Max	Min
BTM	3.28	4.60	37.99	0.00	5.64	6.84	41.97	0.00
SMBC	4.73	6.05	44.25	0.01	12.00	17.91	113.51	0.00
UFJ	6.61	9.86	111.42	0.01	11.36	18.28	130.58	0.00
MIZUHO	7.05	10.39	95.93	0.01	9.73	19.13	195.45	0.00
Japan	1.38	1.72	44.10	0.00	——	——	——	——

Table 3: Estimation Results: Unrestricted Case

(i) Parameter Estimates

Parameter	BTM	SMBC	UFJ	MIZUHO	Japan
κ^Q	-0.14706***	-0.15232***	-0.27278***	-0.21296***	-0.16988***
	[0.02837]	[0.02021]	[0.01251]	[0.04327]	[0.01126]
θ^Q	-0.00593***	-0.01172***	-0.00307***	-0.01215**	0.00000
	[0.00177]	[0.00067]	[0.00040]	[0.00545]	[0.00026]
σ^Q	0.08076***	0.10673***	0.10790***	0.12072***	0.04100***
	[0.00159]	[0.00420]	[0.00094]	[0.02347]	[0.00014]
κ^P	0.98254***	0.72368***	0.32046***	0.60957	0.53125***
	[0.00756]	[0.15330]	[0.08053]	[1.95225]	[0.00949]
θ^P	0.00387***	0.00852**	0.02004***	0.01301	0.00296
	[0.00080]	[0.00428]	[0.00181]	[0.03871]	[0.00687]
σ_ε	0.36108***	0.35766***	0.56058***	0.52194***	0.67471***
	[0.00118]	[0.00068]	[0.00011]	[0.00429]	[0.00211]
R_{se}^Q	0.62701***	0.78550***	0.87363***	0.85205***	0.77400***
	[0.05931]	[0.04295]	[0.01029]	[0.05686]	[0.00927]
α	0.46032***	0.73774***	0.85260***	0.82571***	———
	[0.13753]	[0.06714]	[0.01377]	[0.07936]	———
Log likelihood	7.163	6.993	6.704	6.740	7.850
Observations	1,326	1,306	1,290	1,158	1,715

Note: Figures in parentheses are standard errors. ***, **, and * denote significance level at the 1%, 5%, and 10% level, respectively. Log likelihood is the sample average.

(ii) Absolute Value of Pricing Error (bps)

	Senior CDS				Subordinated CDS			
	Mean	Std	Max	Min	Mean	Std	Max	Min
BTM	2.09	3.80	34.22	0.00	2.26	3.63	31.43	0.00
SMBC	2.68	3.75	36.78	0.01	3.76	6.30	59.50	0.00
UFJ	3.89	8.06	103.45	0.00	3.92	7.09	67.68	0.00
MIZUHO	4.69	8.76	90.00	0.01	3.77	9.61	154.67	0.00
Japan	1.36	1.43	16.95	0.00	——	——	——	——

(iii) LR Test Result $H_0 : R_{se}^Q = 0.4$

	BTM	SMBC	UFJ	MIZUHO	Japan
χ^2 statistic	858.8***	1.212.7***	1,081.9***	937.4***	46.4***

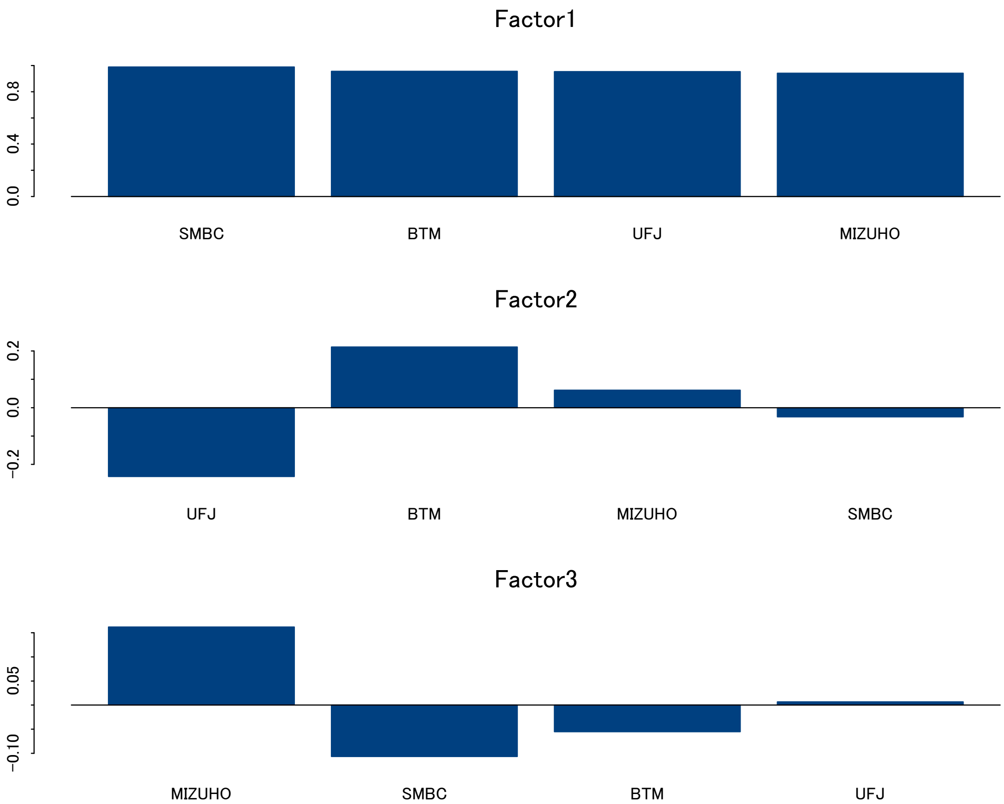
Note: *** denotes significance level at the 1% level.

Table 4: Factor Analysis

(i) Importance of Factors

	Factor 1	Factor 2	Factor 3
Eigenvalue	3.689	0.110	0.041
Explained proportion of total variance	0.922	0.027	0.010

(ii) Factor Loadings



Note: The method of principal component is used. The result is the initial solution without a rotation.

Table 5: Unit Root Tests

Sample Period:1998/10/20 to 2005/12/31 (Number of Observations: 2,630)

	ADF (Augmented Dickey-Fuller) Test		PP (Phillips-Perron) Test	
	Test Statistic	Lags	Test Statistic	Bandwidth
BTM	-2.375	1	-2.628	9
SMBC	-2.069	0	-2.156	11
UFJ	-1.320	1	-1.350	15
MIZUHO	-2.511	1	-2.505	12
Common	-1.917	1	-2.073	22
Japan	-1.541	1	-1.674	18
Δ BTM	-45.652***	0	-45.835***	5
Δ SMBC	-48.573***	0	-48.635***	9
Δ UFJ	-47.694***	0	-47.797***	14
Δ MIZUHO	-47.840***	0	-48.060***	11
Δ Common	-44.566***	0	-46.189***	20
Δ Japan	-47.808***	0	-48.337***	17

Notes:
1. Test statistics are based on the specification including a constant term and a linear trend.
2. The number of lags is chosen by Schwarz Criterion. The bandwidth is based on the Newrey-West bandwidth.
3. *, **, and *** show that the null hypothesis of existence of a unit root is rejected at the 10%, 5% and 1% significance level, respectively.

Table 6: Johansen Cointegration Test

(i) Japan vs. BTM

Sample Period:1998/10/20 to 2005/12/31 (Number of Observations: 2,630)

Cointegration Rank Test					
H0	H1	Eigenvalue	Trace Statistic	Max-Eigen Statistic	Lags
$r \le 0$	$r = 1$	0.009	26.688***	22.583***	2
$r \le 1$	$r = 2$	0.002	4.105	4.105	

Cointegrating Vector		
Japan	BTM	Constant
1.000	-0.447***	-0.001***

(ii) Japan vs. SMBC

Cointegration Rank Test					
H0	H1	Eigenvalue	Trace Statistic	Max-Eigen Statistic	Lags
$r \le 0$	$r = 1$	0.019	55.141***	51.498***	1
$r \le 1$	$r = 2$	0.001	3.643	3.643	

Cointegrating Vector		
Japan	SMBC	Constant
1.000	-0.197***	-0.002***

(iii) Japan vs. UFJ

Cointegration Rank Test					
H0	H1	Eigenvalue	Trace Statistic	Max-Eigen Statistic	Lags
$r \le 0$	$r = 1$	0.012	34.951***	32.210***	1
$r \le 1$	$r = 2$	0.001	2.741	2.741	

Cointegrating Vector		
Japan	UFJ	Constant
1.000	-0.100***	-0.002***

Notes: 1. r denotes the number of cointegration ranks.
2. The number of lags is chosen by Schwarz Criterion.
3. *, **, and *** denote the 10%, 5%, and 1% significance level, respectively.

(iii) Japan vs. MIZUHO

Sample Period:1998/10/20 to 2005/12/31 (Number of Observations: 2,630)

Cointegration Rank Test					
H0	H1	Eigenvalue	Trace Statistic	Max-Eigen Statistic	Lags
$r \leq 0$	$r = 1$	0.015	43.850***	40.699***	1
$r \leq 1$	$r = 2$	0.001	3.150	3.151	

Cointegrating Vector		
Japan	MIZUHO	Constant
1.000	-0.089***	-0.002***

(iv) Japan vs. Common

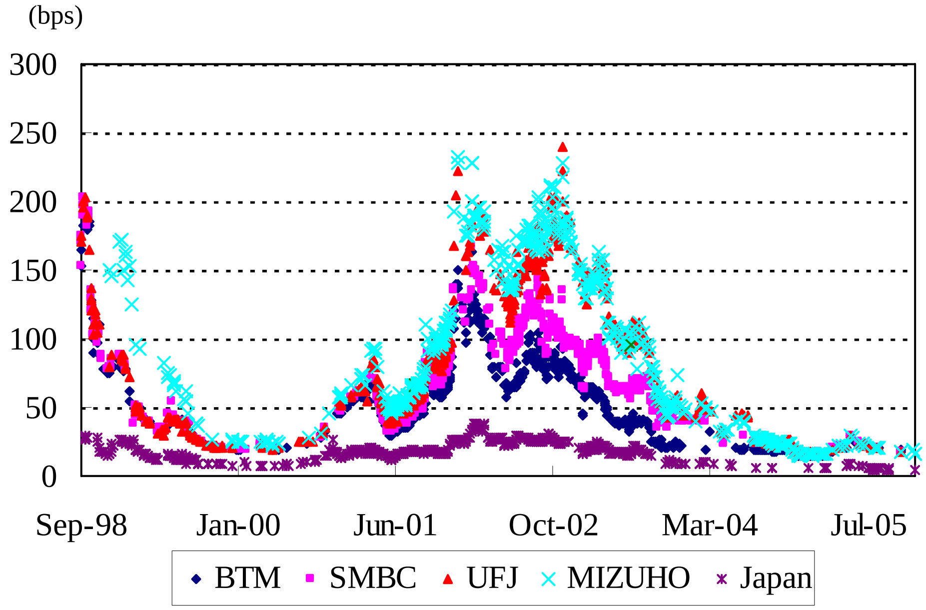
Cointegration Rank Test					
H0	H1	Eigenvalue	Trace Statistic	Max-Eigen Statistic	Lags
$r \leq 0$	$r = 1$	0.014	40.155***	36.421***	2
$r \leq 1$	$r = 2$	0.001	3.733	3.733	

Cointegrating Vector		
Japan	Common	Constant
1.000	-1.028***	-0.038

Notes: 1. r denotes the number of cointegration ranks.
2. The number of lags is chosen based on Schwarz Criterion.
3. *, **, and *** denote the 10%, 5%, and 1% significance level, respectively.
4. The data of Japan for (iv) are standardized with mean zero and standard deviation one.

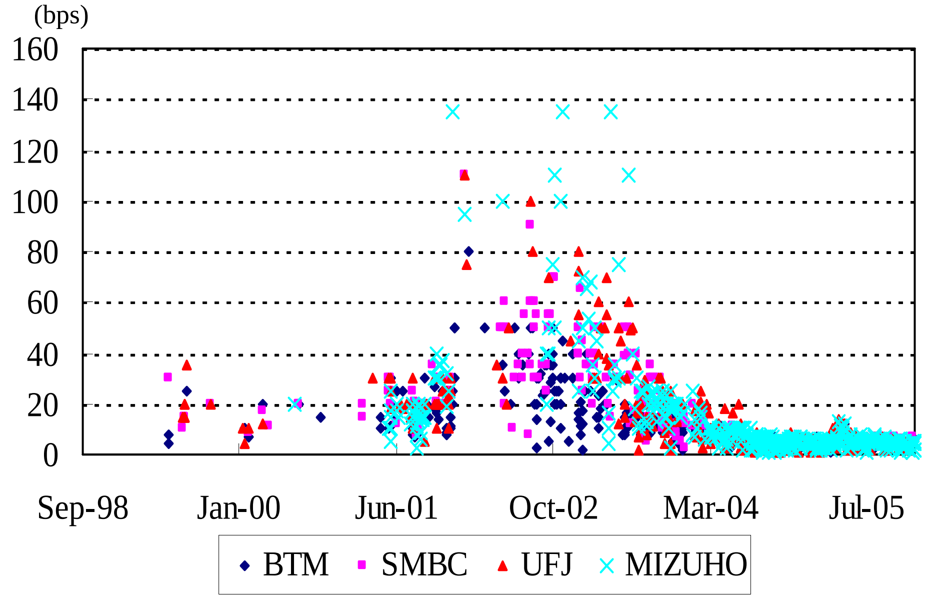
Figure 1: CDS Spreads (5-year Maturity)

(i) Senior CDS Spreads



Source: GFI Limited

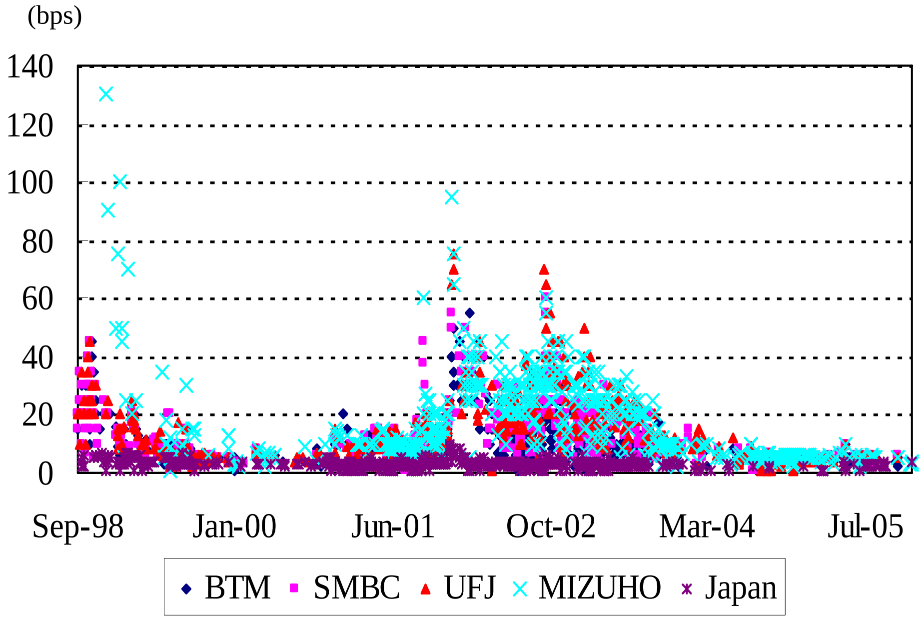
(ii) Subordinated CDS Spreads



Source: GFI Limited

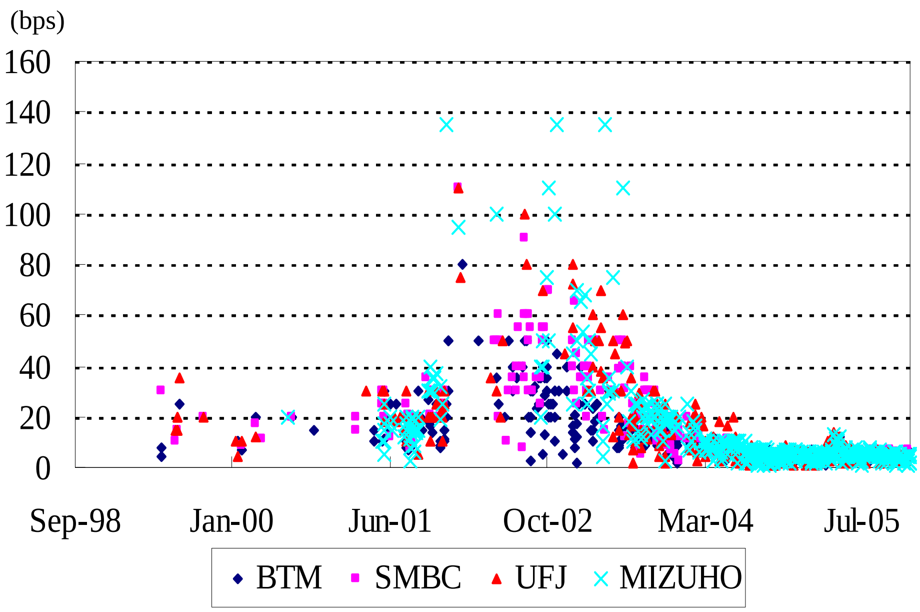
Figure 2: Bid-Ask Spreads (5-year Maturity)

(i) Senior CDS Spreads



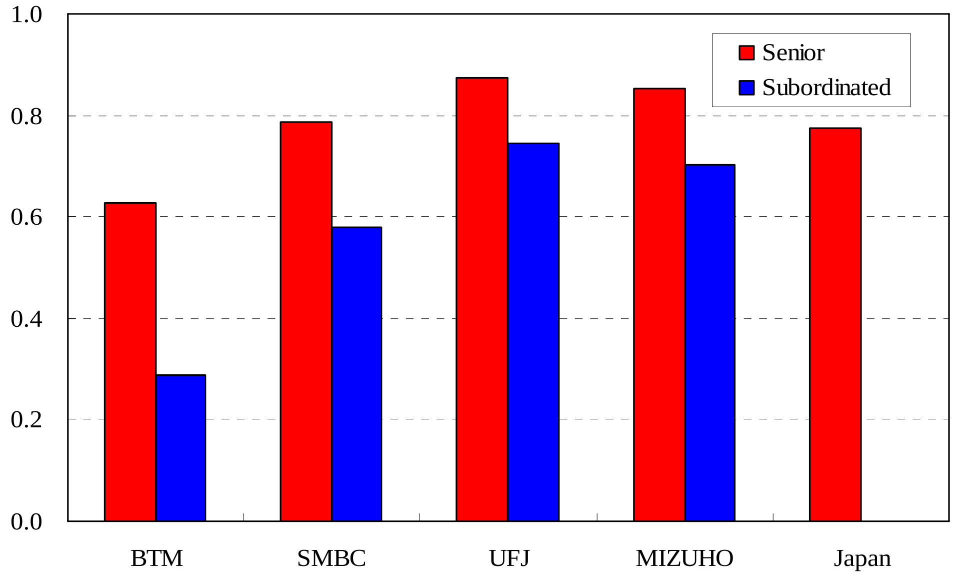
Source: GFI Limited

(ii) Subordinated CDS Spreads



Source: GFI Limited

Figure 3: Estimates of Expected Recovery Ratios

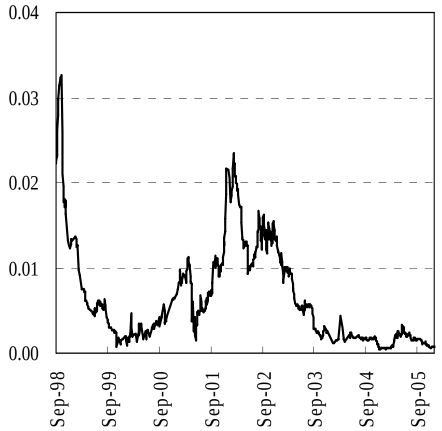


Note: Estimated recovery ratios are based on the estimation results reported in Table 3.

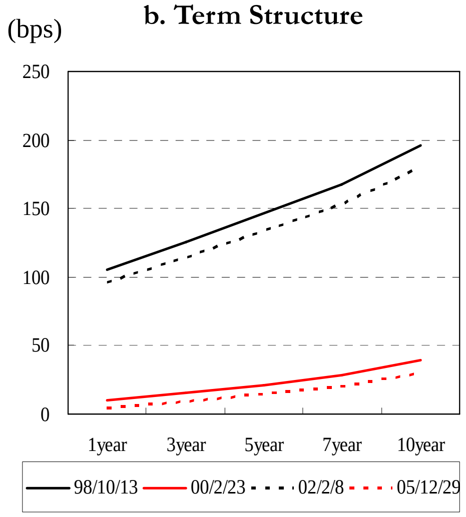
Figure 4: Estimated Default Intensities and Term Structures of CDS Spreads

(i) BTM

a. Default Intensity



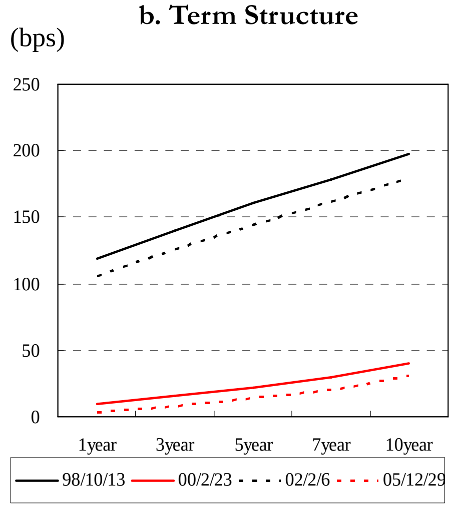
b. Term Structure



(ii) SMBC

a. Default Intensity

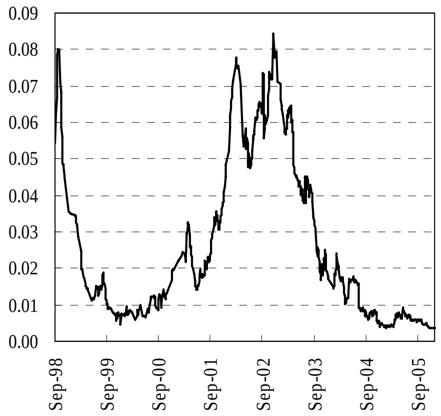
b. Term Structure



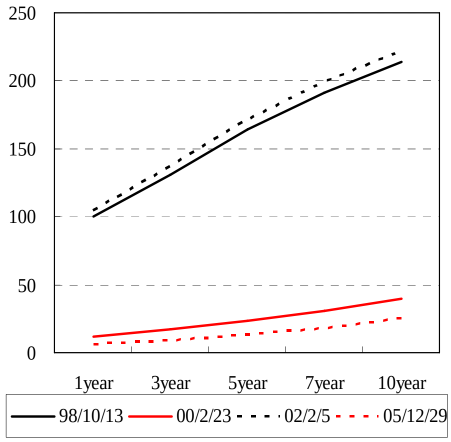
Note: Default intensities and term structures of CDS spreads are calculated based on the estimation results reported in Table 3.

(iii) UFJ

a. Default Intensity

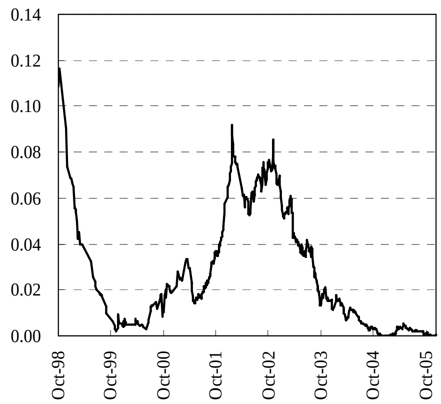


b. Term Structure

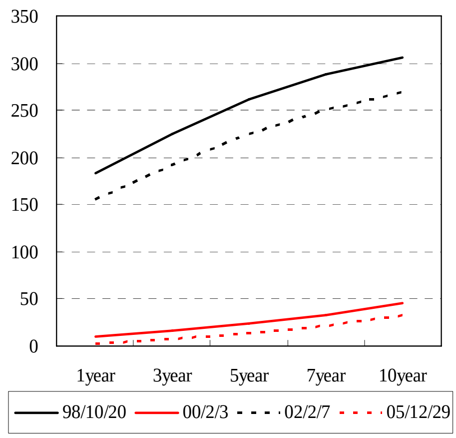


(iv) MIZUHO

a. Default Intensity

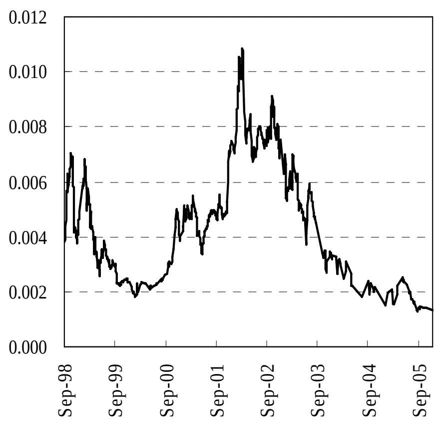


b. Term Structure

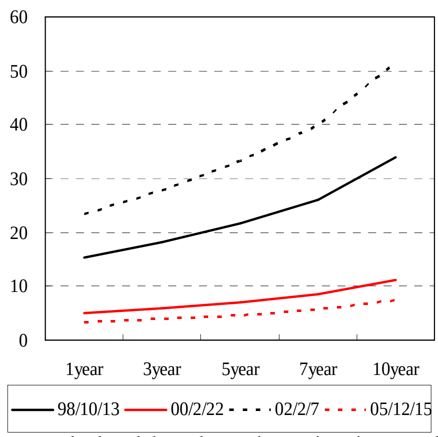


(v) Japan

a. Default Intensity

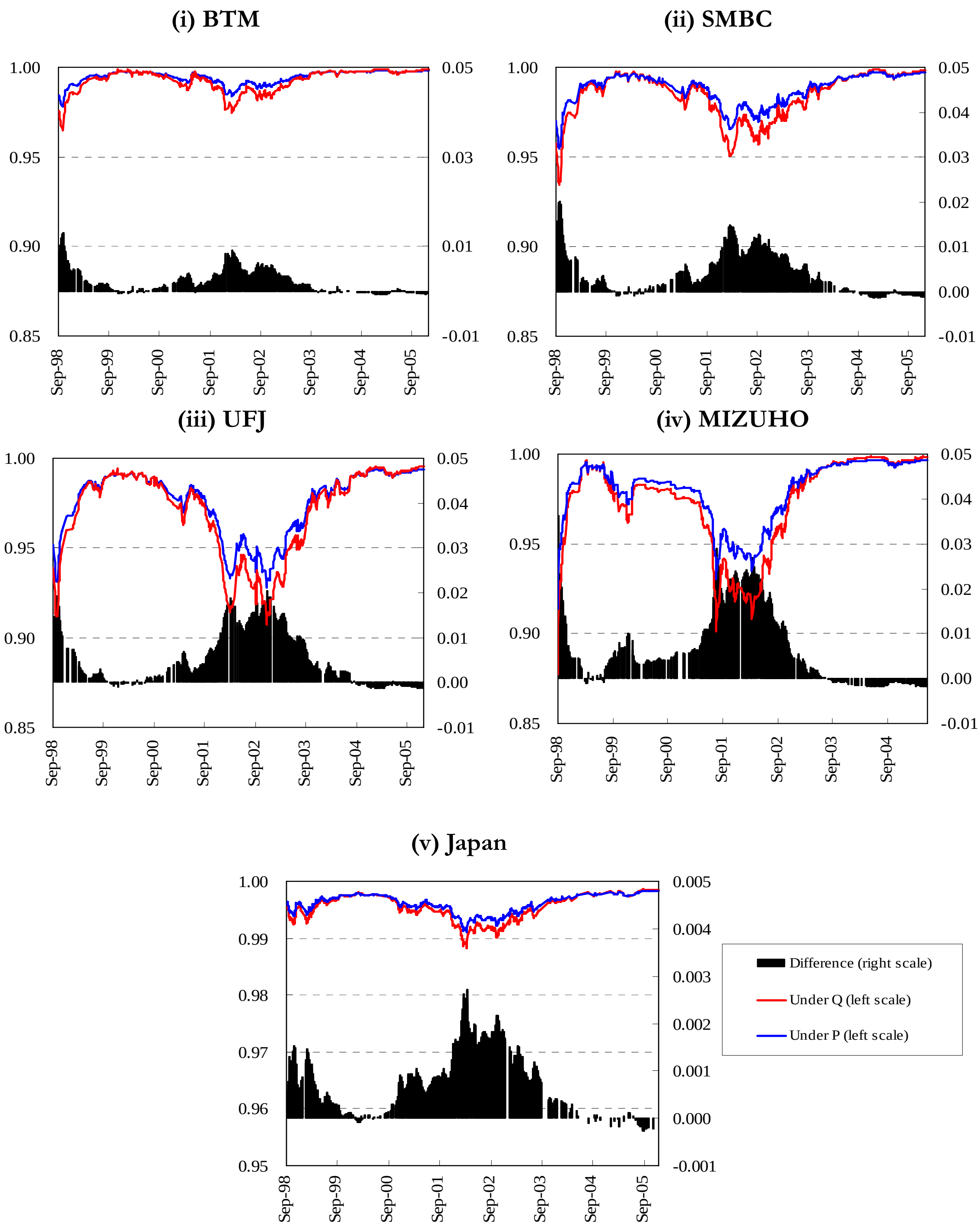


b. Term Structure



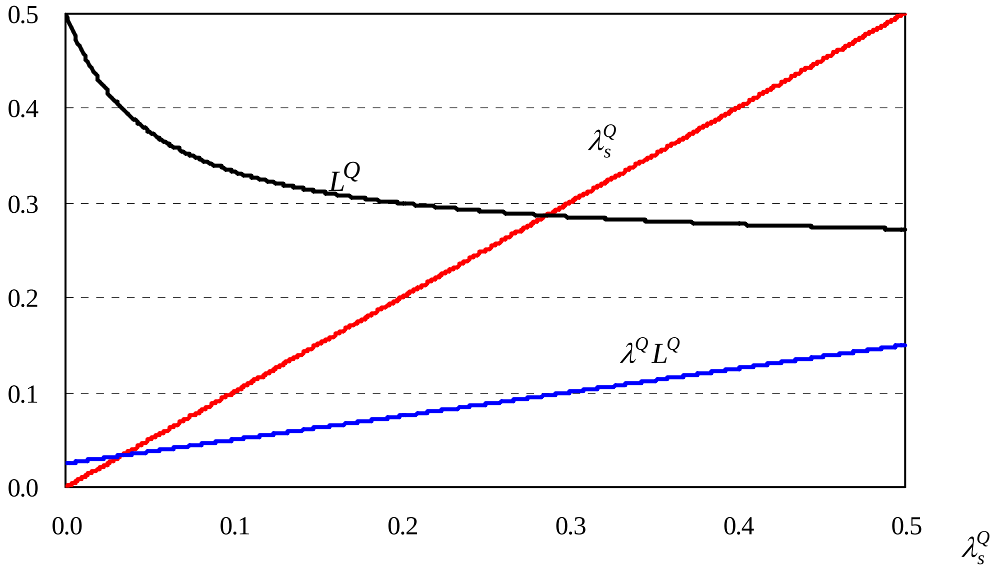
Note: Default intensities and term structures of CDS spreads are calculated based on the estimation results reported in Table 3.

Figure 5: Risk-Neutral vs. Pseudo-Actual Measures of Survival Probabilities



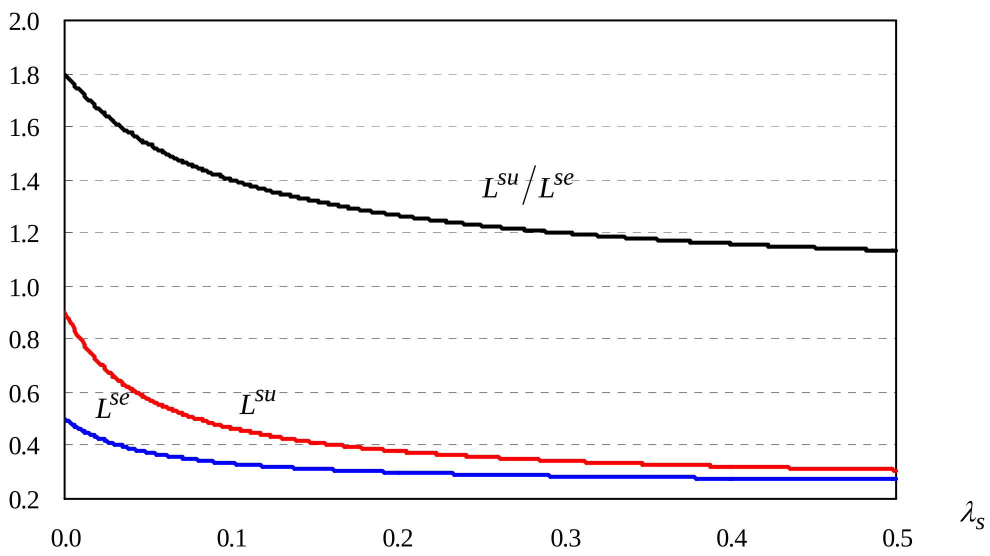
Note: Survival probabilities are for one-year ahead, calculated based on the estimation results reported in Table 3. Survival probabilities under P are the pseudo-actual survival probabilities defined in section 4.

Figure 6: L^Q as a function of λ_s^Q



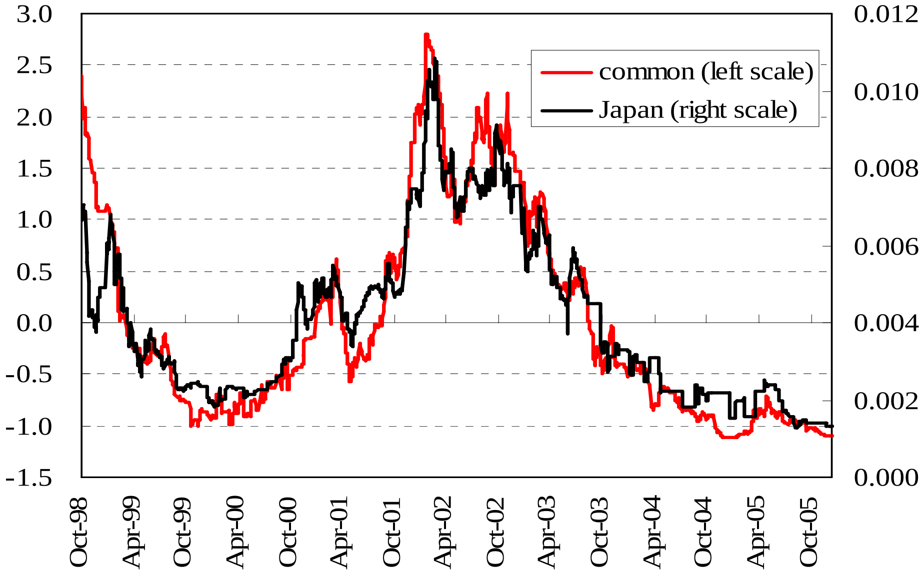
Note: λ_c^Q =0.05, L_c^Q =0.50, and L_s^Q =0.25.

Figure 7: L^{su}/L^{se} as a function of λ_s



Note: λ_c =0.05, L_c^{se} =0.50, L_s^{se} =0.25, L_c^{su} =0.90, and L_s^{su} =0.25.

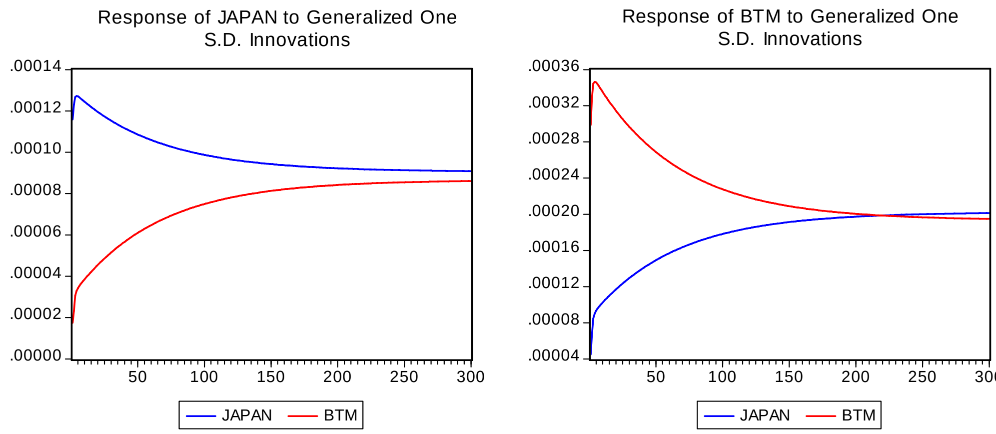
Figure 8: Default Intensity for Japan and Common Factor



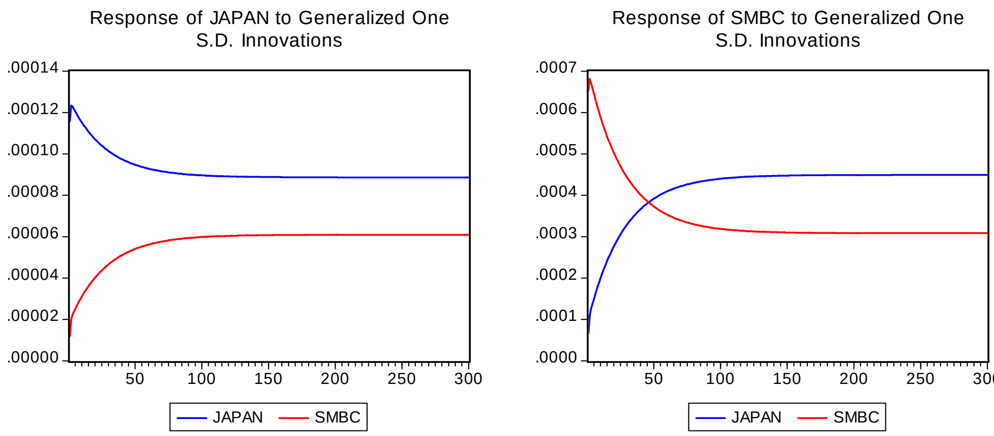
Note: Default intensity for Japan is calculated, based on the estimation results reported in Table 3. Common factor corresponds to Factor 1 reported in Table 4.

Figure 9: Generalized Impulse Reponses

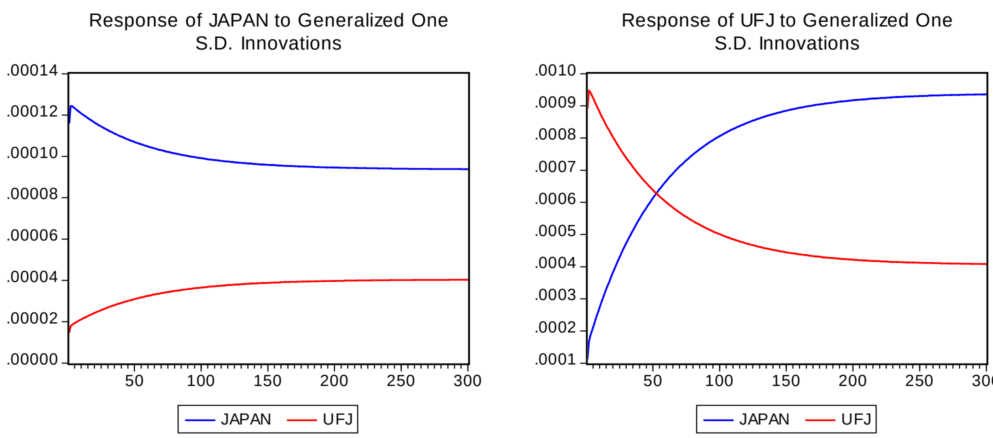
(i) Japan vs. BTM



(ii) Japan vs. SMBC

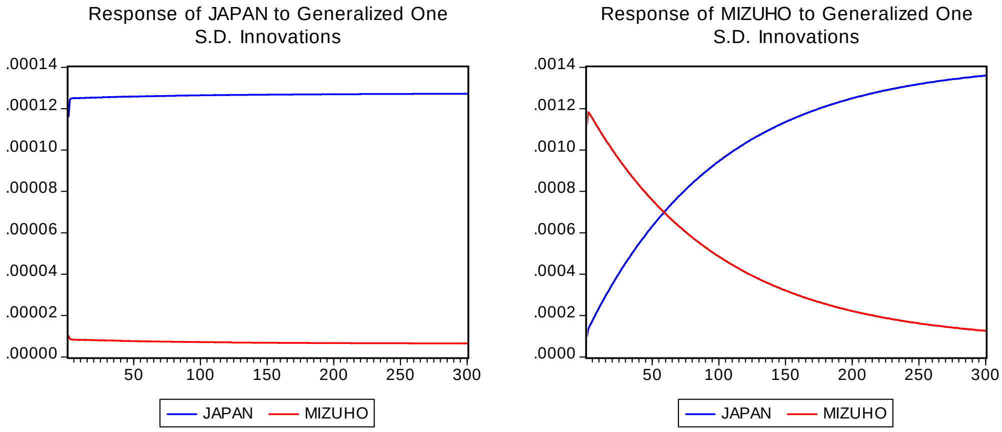


(iii) Japan vs. UFJ

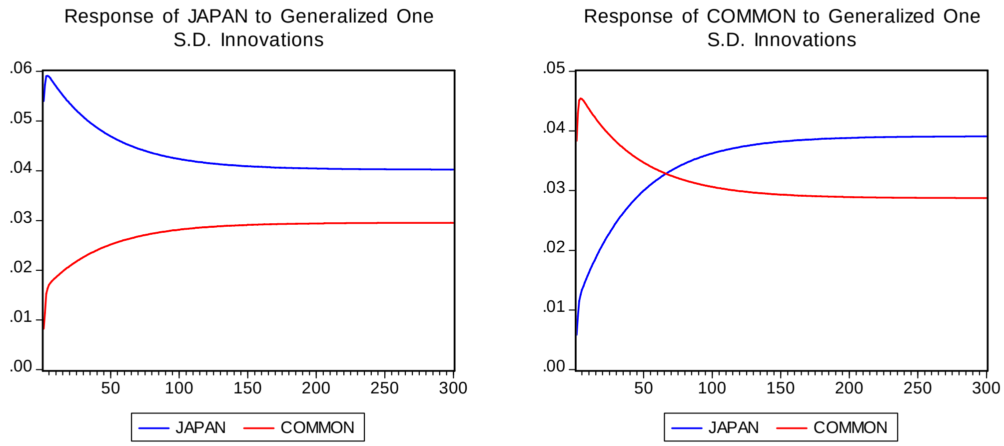


Note : Generalized impulse responses are calculated based on error correction models. Cointegrating vectors are reported in Table 6.

(iv) Japan vs. MIZUHO



(v) Japan vs. Common



Note : Generalized impulse responses are calculated based on error correction models. Cointegrating vectors are reported in Table 6.