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# Price Discovery of Credit Spreads for Japanese Mega-Banks: Subordinated Bond and CDS

**Naohiko Baba* and Masakazu Inada****

**Abstract**

This paper empirically investigates the determinants of credit spreads for Japanese mega-banks with emphasis on comparing subordinated CDS spreads with the subordinated bond spreads from the viewpoint of price discovery in both credit markets. The main findings are summarized as follows. First, subordinated CDS and subordinated bond spreads are significantly cointegrated for most banks, and price discovery measures suggest that the CDS spread plays a more dominant role in price discovery than the bond spread. Second, although both CDS and bond spreads significantly react to the Japanese sovereign CDS spread, only the CDS spread reacts significantly to other financial market variables including its own volatility and equity return. Third, both spreads are responsive to the changes in fundamental accounting variables such as the capital–asset ratio and the nonperforming loan ratio. These accounting variables are likely to constitute common factors that are behind the cointegration relationship. Last, significant volatility spillovers are detected from the CDS to bond spreads. This result implies that new information flows more in this direction.

**Keywords:** Subordinated Bond; Credit Default Swap; Japanese Banks; Price Discovery; Volatility Spillover; Bivariate GARCH

**JEL classification:** G12, G14, G15

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## 1. Introduction

This paper attempts to empirically investigate the determinants of credit spreads on subordinated bonds for Japanese mega-banks, paying particular attention to detecting the relative role of price discovery for banks' credit risk between subordinated credit default swaps (CDS) and subordinated bond spreads. Subordinated bonds are unsecured fixed-income instruments that are senior only to common equity when a failed bank is liquidated. Thus, the spreads on subordinated bonds over government bonds should be more sensitive to the banks' credit risk than the spreads on straight bonds, because investors in the subordinated bond are likely to lose at least part of their principal and interest in the case of a failure. On the other hand, a subordinated CDS contract is a swap contract with a subordinated bond as a reference bond (entity). The CDS also mirrors the identical credit risk because it refers to the same asset (entity).

Many studies have investigated whether subordinated bond investors are sensitive to the credit risk as asset price theory suggests using the spreads of subordinated bonds issued by U.S. banks over Treasury yields. Their main objective is to search for the possible role of subordinated bonds as a tool for disciplining banks' risk-taking behavior. First, Avery, Belton, and Goldberg [1988], and Gorton and Santomero [1990] show that (aggressive) risk taking by bank managers was not priced in the subordinated bond spreads in the 1980s. Studies using data from 1991 onwards, which corresponds to the post FDICIA (Federal Deposit Insurance Corporation Improvement Act) period, however, show that credit risk is priced in U.S. subordinated bond spreads. In particular, Flannery and Sorescu [1996] argue that credit risk was not priced until the 1980s because of a rational response of investors to a "too-big-to-fail" policy along with well-established perceptions of forbearance by governments, and once such a perception on the institutional framework vanished, subordinated bond investors began to price credit risk.[1]

In contrast to the empirical studies on US banks, there have been no works on Japanese banks except Kobayashi [2003] and Imai [2007]. Using weekly data of subordinated bond spreads for 13 Japanese banks in the secondary market from 2000 to 2002, Kobayashi [2003] finds that the spreads are not significantly sensitive to both bank-specific risk measured by the market-based leverage ratio and market conditions such as the 10-year Japanese government bond yield and excess stock returns. She concludes that her results are consistent with the previous studies using U.S. data, suggesting a nonnegligible influence of the implicit too-big-to-fail guarantee by the Japanese government. On the other hand, Imai [2007] uses subordinated bond spreads in the primary market from 1993 to 2004 and finds that spreads are significantly higher for banks with weaker financial standing reflected in credit ratings, as well as accounting variables including nonperforming loan ratio and loans to asset ratio.

It should be noted that in recent years, credit markets have created and expanded trading of credit derivatives, the CDS being the most popular product. In a CDS contract, the protection buyer pays the seller a fixed premium in each period until either prespecified 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 must buy back from the buyer the valueless bond at its face value. Thus, the CDS spreads should provide a direct measure of the default probability and the recovery rate from the defaulted bond perceived by credit market participants. Because the principal is not needed for trading CDS contracts, given the nature of derivatives, CDS contracts have been traded

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[1] DeYong et al. [2001] confirm the result of Flannery and Sorescu [1996] using extended data covering 1986–1995, finding that subordinated bond spreads are significantly correlated with accounting and market risk measures. Also, Covitz, Hancock, and Kwast [2004] find that U.S. subordinated bond spreads were significantly risk sensitive after 1985 using a two-step Heckman procedure. Furthermore, Covitz and Harrison

much more frequently, and thus the liquidity of the CDS market has been much higher than traditional straight bonds issued by the same reference entities.

The market liquidity of Japanese straight bonds, not to mention the liquidity of subordinated bonds, is low compared with the CDS market. Probable reasons for this are: (i) Japanese 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.

In the case of the Japanese CDS market, Ito and Harada [2004] mention that the main reference entities are the Japanese mega-banks, and that CDS spreads reflect the credit risks of Japanese banks more sensitively than other credit instruments including bond spreads. This is due mainly to the above-mentioned higher liquidity of CDS trading for Japanese banks, as well as the difference in investor types between CDS and bond markets.

The main investors in the Japanese CDS market are hedge funds and non-Japanese investment banks, while those in the Japanese corporate bond market are Japanese institutional investors including life insurance companies and pension funds. In general, hedge funds and investment banks trade actively, but Japanese institutional investors tend to follow a buy-and-hold strategy in investing in corporate bonds.

Taking account of the development of the CDS market as an important venue for pricing the credit risk of each reference entity, in this paper we pay particular attention to the comparison between the subordinated CDS and subordinated bond spreads in monitoring the credit risk of Japanese mega-banks. The rest of the paper is organized as follows. Section 2 first describes the empirical methodology and data, and then reports the results. Section 3 concludes the paper.

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[2004] report a "positive selection" attribute in that issuance tends to be timed with the announcement of positive news like ratings upgrades.

## 2. Empirical Analysis

### 2-1 Overall Empirical Strategy

Figure 1 shows the subordinated CDS and bond spreads for the four Japanese mega-banks, Bank of Tokyo Mitsubishi (BTM), Sumitomo Mitsui Banking Corporation (SMBC), MIZUHO Corporate Bank (MIZUHO), and UFJ Bank (UFJ).[2] Prior to 2004, CDS spreads were much higher than bond spreads for all banks, but after that, these spreads tend to converge.

The empirical analysis in this paper takes the following steps. First, we explore the common factors explaining CDS and bond spreads, which are likely to create the long-term relationship between these spreads, as well as the other idiosyncratic factors that induce the short-term deviation of each spread from the long-term relationship. We do so by examining how and to what extent both spreads react to changes in the related financial market variables and fundamental accounting variables for the same banks. The financial market variables that we test include their own volatilities, the Japanese sovereign CDS spread, and equity returns on a weekly basis, while the accounting variables include the capital asset ratio and the nonperforming loan ratio on a semiannual basis.

Second, we investigate which of the two credit markets primarily evaluates the credit risk on a daily basis using the price discovery measures proposed by Gonzalo and Granger [1995] and Hasbrouck [1995]. More specifically, we first examine the cointegration relationship between both spreads, and then we investigate their individual adjustment processes toward the long-term cointegration relationship using a vector error correction model (VECM). Two price discovery measures are computed from the VECM estimation results.

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[2] BTM and UFJ merged into the Bank of Tokyo-Mitsubishi UFJ on January 4, 2006. Because our sample period ends on the last trading date in 2005, this merger does not influence our analysis.

Finally, we explore the volatility spillovers between subordinated CDS and bond spreads on a daily basis using the bivariate GARCH (Generalized Autoregressive Conditional Heteroskedasticity) model. The motivation here comes from the conventional view that the direction of volatility spillovers between markets reveals the flow of new information between them.

Before proceeding to the empirical analysis, let us elaborate on the properties of the CDS spreads that we use in this paper. First, credit events for CDS contracts are wider than straight bonds. They typically include (i) bankruptcy, (ii) failure to pay, and (iii) restructuring.[3] The larger number of credit events for CDS contracts works to raise the CDS spreads over the bond spreads in general. The CDS spreads that we use in this paper are the spreads priced in the CDS contracts where the underlying reference bonds are subordinated bonds issued by the Japanese four mega-banks.[4] Both CDS and bond spreads are mid-prices of the observed bid-ask quotes, denominated in basis points, released by Bloomberg. The bond spreads are computed as the yield on the subordinated bond minus the yield on the Japanese government bond with the same maturity. We use the subordinated CDS with 5-year maturity and bond spreads with 10-year maturity when issued. This difference in maturity inevitably arises from the market practice that most of the CDS contracts are made with 5-year maturity. However, the actual difference in maturities is smaller for the following reasons: (i) CDS is contracted with existing subordinated bonds as a reference, and (ii) most of the subordinated bonds issued by Japanese banks are callable in that they give the banks the option of early redemption from the holder at stipulated call prices and actually are called 2–3 years prior to the original maturity. Also note that CDS spreads are as of 5 p.m. in London, while bond spreads are as of 5 p.m. in Tokyo. Thus, we should consider this

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[3] See Blanco, Brenna, and Marsh [2005] for more details.

[4] Credit events are generally the same between the senior and subordinated CDS contracts.

time difference between CDS and bond spreads in investigating the relative importance of price discovery.

### 2-2 Credit Spreads and Financial Market Variables

First, we investigate how financial market variables exert influence on subordinated spreads. We follow Blanco, Brennan, and Marsh [2005] in choosing the variables. Specifically, they use the changes in (i) the 10-year bond yield, (ii) the slope of the yield curve, (iii) equity price, and (iv) implied equity volatility. We also tried implied volatilities for both equity (TOPIX and NIKKEI 225) and Japanese government bond futures. However, we omitted the estimation results because no meaningful results were found. To focus on the directional movement of each spread rather than daily fluctuations, we use weekly data as in Blanco, Brennan, and Marsh [2005]. Our sample banks are four mega-banks, and the sample period is from April 2, 2004 to December 30, 2005. The number of observations is 92 for each bank.

The financial market variables that we test here are listed as follows. [5] Definitions of each variable are described in Table 1.

(i) Volatility of subordinated CDS and bond spreads. We use this variable to capture the risk-return relationship in that a higher spread might be needed to compensate for high-volatility risk.

(ii) Japanese sovereign CDS spread. It is pointed out that the Japanese banking system has been heavily protected by the government. Thus, anecdotal evidence suggests that credit market participants have perceived that the financial condition of the Japanese government influences the credit risk of banks. We test this perception using the Japanese sovereign CDS

[5] The data is sourced from Bloomberg.

spread. By definition, it directly measures the market perception of the credit risk of Japanese government bonds.

(iii) Equity returns. We use the equity returns as measured by the rate of change of equity prices as a proxy for bank default risk.

(iv) 10-year JGB yield. We use this variable as a proxy for the baseline current interest rate. The spreads depend on the economic situation, which is reflected in the interest rate.

(v) The slope of the yield curve. We use this variable because it is sometimes argued that the slope of the yield curve has predictive power for the future state of real economic activity, which might be linked to the financial standing of banks.

(vi) Merger Dummy. We include a dummy variable that is equal to one for UFJ on July 16, 2004, when the merger with the BTM was reported, and zero otherwise.

Estimation is done by cross-sectional time-series regression, where we chose the most appropriate estimation method using the following three specification tests: (i) the $F$ test between the fixed-effects and pooled models, (ii) Breusch and Pagan's Lagrange multiplier (LM) test between the random-effects and pooled models, and (iii) the Hausman test between the random-effects and fixed-effects models. It turns out that the pooled model was accepted for all the cases. Because of the high correlation between the 10-year JGB yield and the slope of the yield curve, we use one of these variables in each regression.

Table 2 reports the results of the unit root tests. Both the Augmented Dickey–Fuller test and the Phillips–Perron test show that all the variables except the bond volatility and equity returns are I(1), and thus we use first-differenced data for these variables. Table 3 reports the estimation results. The table exhibits the following similarities and differences between subordinated CDS and bond spreads.

First, the Japanese sovereign CDS spread has a significantly positive effect on both subordinated CDS and bond spreads, seemingly serving as one of the factors common to both spreads. This is suggestive of the market participants' strong perception about the decisive role of the government in protecting the banking system, consistently with the anecdotal evidence. The CDS spread, however, has a much higher sensitivity to the Japanese sovereign CDS spread than the bond spread.

Second, riskiness as measured by volatility of the spreads is significantly priced in for only the subordinated CDS spread. This result implies that the subordinated bond spread does not price its own riskiness.

Third, equity returns have a significantly negative effect only for the subordinated CDS spread. Anecdotal evidence suggests that the main players in CDS trading, non-Japanese investors including hedge funds and investment banks, tend to react more to the equity return because they conduct cross-market trading more actively on a global basis than Japanese investors. On the other hand, the main investors in subordinated bonds, Japanese institutional investors, tend to buy and hold them, not tending to react to developments in other markets. These differences in investor types between CDS and bond markets are likely to cause short-term deviations from the long-term relationship.

Fourth, the news of the merger between BTM and UFJ on July 16, 2004 significantly lowered both CDS and bond spreads for UFJ, lowering the CDS/bond spread by about 37/20 basis points, respectively.

The above results suggest that the financial market variables such as the volatility of the CDS spread and equity returns tend to let CDS and bond spreads deviate from each other at least in the short run. On the other hand, the Japanese sovereign CDS spread tends to comove with both spreads, but the sensitivity of the CDS spread is much higher than the bond spread.

### 2-3 Credit Spreads and Accounting Variables

Second, we investigate the relationship between credit spreads and fundamental accounting variables particularly to search for the factors behind the long-term relationship between subordinated CDS and bond spreads. Here, the frequency of accounting variables is semiannual, and the sample period is from the first half of fiscal 2002 to the first half of fiscal 2005 because of the availability of spread data.[6] Considering the time lag until these accounting variables are available to investors, we compare the accounting variables as of the first (second) half of each fiscal year with the credit spreads as of the end of November (May) when Japanese banks release their financial statements.

The accounting variables that we test here are as follows.[7] Definitions of each variable are described in Table 4.

(i) Capital ratio. We use the capital ratio as a proxy for the expected default probability of each bank.

(ii) Adjusted capital ratio.[8] This variable is defined as the adjusted capital divided by total assets, where the adjusted capital is calculated by subtracting the amount of the potential shortage of the loan-loss reservation, net unrealized losses/gains on securities and land, deferred tax assets and several items from the "nominal" equity capital reported by each bank. The use of the adjusted capital ratio instead of the ordinary capital ratio is meant to capture the assured capital of each bank.

(iii) Nonperforming loan ratio. Japanese banks had suffered from the nonperforming loan problem since the bursting of the asset bubble in the early 1990s. Thus, investors have been

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[6] As for the CDS spreads, the sample period starts from the second half of fiscal 2002. However, the number of observations is larger for CDS spreads than bond spreads, because there are periods during which some banks did not have subordinated bonds outstanding.

[7] The data is sourced from financial statements.

very conscious about the nonperforming loan ratio in gauging bank risk. Because of high correlations between each pair of the capital ratio, the adjusted capital ratio, and the nonperforming loan ratio, we use one of each variable in each regression.

(iv) Size. In general, assets of larger banks are considered to be more diversified. Also, the government might protect larger banks under the "too-big-to-fail" policy. Thus, investors are likely to perceive that the larger the bank, the smaller the expected default risk.

(v) Lending growth. We use this variable to control for the growth factor of each bank.

Our sample banks include the four Japanese mega-banks as before. Estimation is done by panel regression, where we chose the most appropriate estimation method using the same specification tests as in the preceding analysis. The specification results show that for most cases, the pooled model was the most appropriate method.

Table 5 reports the estimation results. It shows that the adjusted capital ratio, nonperforming loan ratio, and size matter significantly as determinants of both subordinated CDS and bond spreads. A bank with a higher adjusted capital ratio, a lower nonperforming ratio, and a larger size tends to have significantly lower subordinated CDS and bond spreads. Put differently, these variables are likely to constitute the common factors in the long run. The overall estimation result suggests that investors of subordinated CDS and bonds properly price in bank risk using fundamental accounting variables such as the adjusted capital ratio and the nonperforming loan ratio.[9] On the other hand, the "normal" capital ratio has a significantly negative coefficient for the

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[8] We follow the definition of the Japan Center for Economic Research, which releases this variable on a regular basis. The original name of this variable is the "real capital ratio."

[9] The overall result here is consistent with Imai [2007], but not with Kobayashi [2003]. Kobayashi [2003] finds that the subordinated bond spread is not sensitive to bank-specific risk such as leverage. One of the possible reasons for this difference is that her measure of leverage is calculated on a weekly basis using equity prices. Because the frequency of the accounting variables such as the leverage on a book-value basis is semiannual, the weekly variation of her measure of leverage mainly stems from the weekly variation of equity prices, of which changes are not always consistent with the financial variable that is updated infrequently. Meanwhile, she does not analyze the effects of the size and nonperforming loan ratio on the credit spreads.

subordinated CDS spread but is not significant for the subordinated bond spread. Another thing to note here is that the CDS spread is more sensitive than the bond spread to all the variables except size. This is likely to be associated with the higher sensitivity of the CDS spread than the bond spread to the implicit efficient price as suggested by the cointegration analysis that follows.

## 2-4 Measuring Price Discovery of CDS and Bond Spreads

Third, we explore which of the two credit markets primarily evaluates the credit risk for banks. Theoretically, arbitrage activities should keep spreads observed in CDS and bond markets within a certain range at least in the long run. Thus, the CDS and bond spreads for the same entity should be cointegrated and driven by a certain common factor. Previous studies consider the factor to be an implicit efficient price that plays a key role in price discovery. For example, Lehman [2002] defines price discovery as the process by which efficient and timely incorporation of new information implicit in investors' trades is priced into market prices. In this context, the implicit efficient price can be assumed to be driven only by new information because it already fully reflects all information available at the current time. The efficient price, therefore, can be modeled to follow the random walk process (see Samuelson [1965]). When different assets that priced in information on the same entity are traded in different two markets, order flow is fragmented and price discovery is split between these markets. This implies that an individual market price can deviate from its efficient price, which would be formed in an integrated market if all information were incorporated into the single market.

There are two approaches that attracted academic attention for investigating the mechanism of price discovery. One is the permanent–transitory model developed in Gonzalo and Granger [1995]. This model is based on the idea that asset prices will consist of permanent and transitory components. The permanent component corresponds to the random walk process of the efficient price representing developments in fundamental factors, which generates comovement between asset prices. Suppose that the efficient price is primarily discovered in a market, the price in the other market tends to converge to the price in the primary market, and thus the adjustment of the main market price is slower than the other price. This mechanism can be described by a VECM, and the intensities of the price adjustments are measured by the error correction coefficients. Gonzalo and Granger [1995] introduce the measurement of the relative intensity.

The second model of price discovery is the information share model developed in Hasbrouck [1995]. This is also based on a VECM, consisting of permanent and transitory components. Hasbrouck [1995] focuses on the variance of the permanent component in an individual price, the development of which is driven by permanent shocks occurring in both its own market and the other markets. The model measures the relative contribution of the permanent shock in its own market to the total variance in price. The market that contributes most to this variance is the primary market for price discovery.

Both models rely on a VECM of market prices as follows.

$$\Delta \mathrm{CDS}_t = \lambda_1 \left(\mathrm{CDS}_{t-1} - \alpha_1 \mathrm{Bond}_{t\text{-}1} - C\right) + \sum_{j=1}^{p} \beta_{1j} \Delta \mathrm{CDS}_{t\text{-}j} + \sum_{j=1}^{p} \gamma_{1j} \Delta \mathrm{Bond}_{t\text{-}j} + \varepsilon_{1t} \qquad (1)$$

$$\Delta \mathrm{Bond}_t = \lambda_2 \left(\mathrm{CDS}_{t-1} - \alpha_1 \mathrm{Bond}_{t\text{-}1} - C\right) + \sum_{j=1}^{p} \beta_{2j} \Delta \mathrm{CDS}_{t\text{-}j} + \sum_{j=1}^{p} \gamma_{2j} \Delta \mathrm{Bond}_{t\text{-}j} + \varepsilon_{2t} \qquad (2)$$

Here, $\varepsilon_{1t}$ and $\varepsilon_{2t}$ are i.i.d shocks. As mentioned above, an underlying assumption is that there is an unobservable efficient price of credit risk, which follows a random walk, and observable prices contain transitory noise, together with this efficient price. This can be confirmed by the econometric idea that a VAR (Vector Autoregressive) model of I(1) processes with cointegration, the basis of the VECM above, is identically transformed into a VMA (Vector Moving

Average) model of the first differences of the I(1) processes. In this setting, if the two spreads respond to the efficient price to the same degree, they form a cointegrating vector [1, –1, C], where C denotes a constant term representing institutional factors such as the difference in transaction costs.[10] If the CDS (bond) market contributes to price discovery in the sense of the Gonzalo and Granger model, then $\lambda_2 (\lambda_1)$ is positive (negative). As stated in Engle and Granger [1987], the existence of cointegration assures that at least one market has to adjust.

More specifically, the price discovery measure in the Gonzalo and Granger model can be given by:

$$\mathrm{GG} = \frac{\lambda_2}{\lambda_2 - \lambda_1}.$$

Following the preceding studies, we judge that market 1 (CDS) has a dominant role in price discovery when this GG measure for market 1 is larger than 0.5. The measure, however, has a shortcoming that the absolute value of the measure does not necessarily stand for the intensity of the price adjustment because the variances of the right-hand side in equations (1) and (2) might vary in size because of the difference in variances of the shock terms in equations (1) and (2).

Hasbrouk's measure gives an alternative to consider the difference in variances, which is calculated by:

$$\mathrm{Has}_1 = \frac{\lambda_2^2 \left( \sigma_1^2 - \frac{\sigma_{12}^2}{\sigma_2^2} \right)}{\lambda_2^2 \sigma_1^2 - 2\lambda_1 \lambda_2 \sigma_{12} + \lambda_1^2 \sigma_2^2}, \quad \mathrm{Has}_2 = \frac{\left( \lambda_2 \sigma_1 - \lambda_1 \frac{\sigma_{12}}{\sigma_1} \right)^2}{\lambda_2^2 \sigma_1^2 - 2\lambda_1 \lambda_2 \sigma_{12} + \lambda_1^2 \sigma_2^2},$$

where $\sigma_1^2$, $\sigma_2^2$, and $\sigma_{12}$ are factors in the covariance matrix of $\varepsilon_{1t}$ and $\varepsilon_{2t}$. We also judge that market 1 (CDS) has a dominant role in price discovery when this Hasbrouck measure for

[10] Note here that the difference in credit events between the CDS, bond contracts and maturities may

market 1 is larger than 0.5. Note here that the Hasbrouck measure is given by the lower bound of $\text{Has}_1$ and the upper bound of $\text{Has}_2$ because it uses a Cholesky factorization for the covariance matrix of $\varepsilon_{1t}$ and $\varepsilon_{2t}$. The bivariate model, therefore, has two measures depending on the order of the two markets in the factorization. The difference between the two bounds is positively related to the degree of correlation between $\varepsilon_{1t}$ and $\varepsilon_{2t}$. Hasbrouck names the measure information share in the context of price discovery. Baillie et al. [2002] argue that the average of these two bounds provides a sensible estimate of price discovery when the data frequency is high. Also note here that GG ignores the correlation between the markets as well as the size of the variance, and thus if the residuals in the VECM are strongly correlated, then both measures can provide substantially different results.

It should be noted here, however, that quite recently, Yan and Zivot [2006] rigorously analyze the determinants of these two price discovery measures under some structural models in which both permanent (fundamental) and transitory shocks are identified.[11] Moreover, they point out that the reduced form of the VECM is not identical to the decomposed form of the VMA($\infty$) for the accumulation of permanent shocks and the stationary VMA($\infty$) of transitory shocks. Their main findings can be summarized as follows. First, contrary to Hasbrouck's original claim that his measure effectively ignores transitory shocks, they argue that the measure actually consists of contemporaneous responses to transitory shocks as well as permanent shocks. Second, the GG measure is less appropriate in that it actually contains only contemporaneous responses to transitory shocks, while the Gonzalo and Granger model originally intended to measure the effect of fundamental factors on the efficient price of a random walk process. Following the arguments in

influence the cointegrating vector itself (rather than through the constant term).

[11] The economic interpretation of these two measures has attracted close attention from academic researchers. For instance, the *Journal of Financial Markets* released a special issue about this issue in 2002 (Issue 3). Among others, Lehmann [2002] expresses his concern about the interpretation of both measures in that

Yan and Zivot [2006], the GG measure for market 1 is higher when the contemporaneous response of market 2 to transitory shocks is higher than market 1. Third, even without a correlation between residuals in the VECM, the Hasbrouck measure cannot avoid ambiguity in interpreting the result. A high value for the Hasbrouck measure for market 1 can be caused by a high value for the contemporaneous response of market 1 to fundamental shocks, a high value for the contemporaneous response of market 2 to transitory shocks, or both. In this regard, Yan and Zivot [2006] argue that combining information from the GG measure is likely to enable us to remove some ambiguity associated with the Hasbrouck measure. More specifically, a high Hasbrouck measure together with a low GG measure for market 1 is likely to indicate that market 1 has a more significant response to fundamental shocks; that is, it has a dominant role in price discovery.

We use daily spread data as in Blanco, Brennan, and Marsh [2005] because the use of the high-frequency data is essential for detecting price discovery. Our sample period is from April 1, 2004 to December 30, 2005, and the number of observations is 430.[12] Table 6 reports the results of the unit root tests. Both the Augmented Dickey–Fuller test and the Phillips–Perron test show that the CDS and bond spreads are all I(1), and thus we proceed to the cointegration analysis. Table 7 reports the results of the Johansen cointegration test. Both the trace and max eigenvalue tests show that the subordinated CDS and bond spreads have a significant cointegrating relationship for SMBC, MIZUHO, and UFJ at least at the 10% level but not for BTM. The estimated cointegrating vectors are found to be significantly different from the theoretically suggested vector [1, –1, C] by the LR (Likelihood Ratio) test.

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both measures are based on the residuals from reduced-form VECMs, and it is Yan and Zivot [2006] who rigorously address Lehmann's concern.

[12] We set the starting point at the first date of fiscal 2004. This is because prior to this date, it is quite difficult to get daily or even weekly data because of the low liquidity of the subordinated CDS and bond markets.

Given the cointegrating relationships above, we calculate (i) the GG measure and (ii) the Hasbrouck measure of price discovery, as reported in Table 7. These measures show a seemingly contradictory result. The Hasbrouck measure suggests that the CDS spread plays a role in price discovery that is almost equal to, or more dominant than, that of the bond spread, which is consistent with preceding studies such as Blanco, Brennan, and Marsh [2005].[13] On the other hand, the GG measure suggests a more dominant role of the bond spread in price discovery with an exception of SMBC where the CDS and bond spread play an almost equal role. The above result probably stems from (i) a high correlation between residuals in the VECM and (ii) the ambiguity of both measures' interpretation, as argued by Lehmann [2002] and Yan and Zivot [2006]. Regarding the second issue, if we follow the findings by Yang and Zivot [2006], we can clarify the interpretation. That is, a higher Hasbrouck measure and a lower GG measure for the CDS market is likely to indicate that the CDS spread tends to respond more significantly to permanent shocks than the bond spread. The CDS spread has a more dominant role in price discovery in this respect.

Here, we consider the issue of the time difference in quotes between the CDS and bond spreads mentioned in section 2-1. The CDS spread is the quote as of 5 p.m. in London, while the bond spread is as of 5 p.m. in Tokyo. Thus, the information set common to both spreads is the information between 5 p.m. in London (last day) and 5 p.m. in Tokyo. In normal circumstances, major credit events and/or related news for Japanese banks happen during Tokyo trading time, and thus this time difference does not seem to cause major trouble.[14]

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[13] If we set $\lambda_1$=0 for SMBC because this coefficient is not significant, then the Hasbrouck measure becomes 0.989 (Has 1), 1 (Has 2), and 0.995 (mid), respectively, which is more supportive for the CDS spread playing a more dominant role in price discovery.

[14] For instance, in investigating the determinants of the TIBOR-LIBOR spread, Covrig, Low, and Melvin [2004] compare today's TIBOR and the one-day lagged LIBOR to adjust for the time difference in fixing (TIBOR is fixed at 11 a.m. Tokyo time, and LIBOR is fixed at 11 a.m. London time). In this case, the information set common to both spreads becomes the information between 5 p.m. in Tokyo (last day) and 5 p.m. in London (last day). To check the robustness of our result, we recomputed the price discovery measures with the one-day lagged CDS spreads. The price discovery measures thus computed also support a more dominant role of the CDS spread in price discovery.

### 2-5 Volatility Spillovers between CDS and Bond Spreads

Finally, we explore the volatility spillovers between the CDS and bond markets. Many studies including French and Roll [1986] show that the variance of financial assets is closely related to new information flows. More specifically, French and Roll [1986] find that asset price returns are much more volatile during weekdays than during weekends, and this gap is caused by differences in the information flow.

We use the bivariate GARCH model to examine volatility spillovers between these two markets, whose basic structure can be written as:

$$\boldsymbol{\varepsilon}_t | \boldsymbol{\Omega}_{t-1} \sim N(0, \mathbf{H}_t), \tag{3}$$

where $\boldsymbol{\varepsilon}_t = (\varepsilon_{it}, \varepsilon_{2t})'$ is the vector of residuals from the VECM (or VAR) system (1), and $\boldsymbol{\Omega}_{t-1}$ is the matrix of conditional past information. Here, we assume that $\boldsymbol{\varepsilon}_t$ follows a Gaussian distribution.[15,16] Thus, we adopt a two-step estimation strategy: (i) estimating the VECM (VAR), and (ii) applying the bivariate GARCH model to the residuals derived from the VECM (VAR) system. The sample period and data frequency are the same as in the preceding price discovery analysis.

There exist numerous parameterization methods for the conditional covariance matrix $\mathbf{H}_t$ in equation (3).[17] The specification that we adopt here is the BEKK[18] model proposed by Engle

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[15] We use the VECM residuals for SMBC, MIZUHO, and UFJ, and the VAR residuals for BTM, because we cannot find a significant cointegration vector for BTM.

[16] This treatment is just for securing the efficiency of the bivariate GARCH model estimation, which has 11 residual-covariance parameters to be estimated. This two-stage approach is asymptotically equivalent to a joint estimation of the VECM and GARCH, as shown in Tse [1999].

[17] We prefer the BEKK model to the so-called "diagonal VEC model" proposed by Bollerslev, Engle, and Wooldridge [1988] because the latter model does not guarantee positive definiteness of the conditional covariance matrix. For details of ARCH-type models, see Bollerslev [1986], Bollerslev, Chou, and Kroner [1992], and Bollerslev, Engle, and Nelson [1994]. For a survey of multivariate GARCH models in particular, see Bauwens, Laurent, and Rombouts [2006].

[18] BEKK is the acronym for Baba, Engle, Kraft, and Kroner [1989].

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:[19]

$$\mathbf{H}_t = \mathbf{C}'\mathbf{C} + \mathbf{A}'(\boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}'_{t-1})\mathbf{A} + \mathbf{B}'\mathbf{H}_{t-1}\mathbf{B}\,, \tag{4}$$

where $\mathbf{H}_t = \begin{bmatrix} h_{11,t} & h_{12,t} \\ h_{21,t} & h_{22,t} \end{bmatrix}$, $\mathbf{C} = \begin{bmatrix} c_{11} & c_{12} \\ 0 & c_{22} \end{bmatrix}$, $\mathbf{A} = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}$, and $\mathbf{B} = \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix}$.

We estimate the system described by equations (3) and (4) by maximizing the log-likelihood function:

$$\mathrm{L}(\mathbf{\Theta}) = -T\ln(2\pi) - \frac{1}{2}\sum_{t=1}^{T}\left(\ln\left|\mathbf{H}_t\right| + \boldsymbol{\varepsilon}'_t\mathbf{H}_t^{-1}\boldsymbol{\varepsilon}_t\right),$$

where $\mathbf{\Theta}$ is a $11 \times 1$ parameter vector.

In the above BEKK model, the off-diagonal parameters are of particular importance in terms of detecting volatility spillovers across markets. For instance, $a_{ij}$ measures the volatility spillovers in the form of the squared values of the shocks from the $i$ th market in the previous period to the $j$ th market in the current period. Similarly, $b_{ij}$ measures the volatility spillovers in the form of the conditional volatility of the $i$ th market in the previous period to the $j$ th market in the current period.

Table 8 reports the estimation result of the bivariate GARCH model. First, all of the diagonal ARCH and GARCH parameters except the GARCH term for SMBC's subordinated bond spread are found to be significant at least at the 10% level, implying that there is a high degree of persistence in the conditional standard deviations. Second, the off-diagonal parameter estimates suggest that volatility spillovers are more significant in the direction from the CDS spreads to the

[19] In practice, GARCH (1,1) suffices for model specification because it corresponds to ARCH($\infty$).

bond spreads than vice versa.[20] As is the case with the preceding price discovery analysis, the time difference in measuring the spreads in Tokyo and London favors the bond spreads. Nevertheless, volatility spillovers are relatively strongly detected in the direction from the CDS spread to the bond spread.

## 3. Concluding Remarks

We have empirically investigated the determinants of credit spreads for the Japanese mega-banks using both subordinated CDS and bond spreads, paying particular attention to detecting the primary role of price discovery for bank credit risk between these credit spreads. The major findings are summarized as follows. First, both spreads are significantly cointegrated for three of the four mega-banks, and price discovery measures based on the VECM suggest that the CDS spread plays a more dominant role than the bond spread in price discovery for Japanese banks' credit risk. Second, both CDS and bond spreads significantly react to the Japanese sovereign CDS spread, but only the CDS spread significantly reacts to other financial market variables including its own volatility and equity returns. Thus, while the Japanese sovereign CDS spread is likely to be one of the common factors, the volatility of the CDS spread and equity returns tend to cause a temporary break in the long-run relationship between credit spreads. Third, both credit spreads are priced in response to changes in fundamental accounting variables such as the capital ratio and the nonperforming loan ratio. These accounting variables are also likely to constitute the common factors that are behind the cointegration relationship. Finally, significant volatility spillovers are detected from the CDS to bond spreads. This result implies that new information flows more in this direction than vice versa.

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[20] Note here that in assessing the volatility spillovers, not the sign but only the significance level of the parameters is important because only squared ARCH and GARCH terms enter into the paths of volatility spillovers.

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## Table 1: Definition of Financial Market Variables

| | Definition | Source |
|---|---|---|
| Bond volatility | 10-day historical volatility of subordinated bond spread | Bloomberg |
| CDS volatility | 10-day historical volatility of subordinated CDS spread | Bloomberg |
| Equity return | Change rate of stock price for each bank (%) | Bloomberg |
| Japan CDS | 5-year Japanese sovereign CDS spread (bp) | Bloomberg |
| 10-year JGB | 10-year Japanese government bond yield (%) | Bloomberg |
| Yield curve slope | 10-year JGB yield minus 2-year JGB yield (% point) | Bloomberg |

## Table 2: Unit Root Tests

Sample Period (weekly) : April 2, 2004 to December 30, 2005 (Number of Observations: 92)

| | | Augmented Dickey–Fuller Test | | Phillips–Perron Test | |
|---|---|---|---|---|---|
| | | Level | 1st difference | Level | 1st difference |
| BTM | CDS spread | –1.881 | –5.233*** | –1.921 | –8.981*** |
| | Bond spread | –1.736 | –9.059*** | –2.091 | –9.104*** |
| | CDS volatility | –1.881 | –5.233*** | –1.921 | –8.981*** |
| | Bond volatility | –4.142*** | –11.399*** | –5.555*** | –19.987*** |
| | Equity return | –11.672*** | –8.884*** | –11.676*** | –57.418*** |
| SMBC | CDS spread | –1.347 | –7.682*** | –1.361 | –8.788*** |
| | Bond spread | –1.926 | –10.015*** | –1.908 | –10.129*** |
| | CDS volatility | –1.347 | –7.682*** | –1.361 | –8.788*** |
| | Bond volatility | –7.268*** | –10.506*** | –6.179*** | –14.104*** |
| | Equity return | –9.453*** | –9.247*** | –9.453*** | –84.301*** |
| MIZUHO | CDS spread | –1.741 | –12.105*** | –1.651 | –11.985*** |
| | Bond spread | –1.735 | –8.444*** | –1.717 | –8.553*** |
| | CDS volatility | –1.741 | –12.105*** | –1.651 | –11.985*** |
| | Bond volatility | –5.992*** | –10.048*** | –5.992*** | –18.411*** |
| | Equity return | –9.998*** | –8.415*** | –9.989*** | –34.626*** |
| UFJ | CDS spread | –1.629 | –11.540*** | –1.538 | –11.416*** |
| | Bond spread | –1.473 | –8.709*** | –1.694 | –10.028*** |
| | CDS volatility | –1.629 | –11.540*** | –1.538 | –11.416*** |
| | Bond volatility | –5.244*** | –9.684*** | –5.349*** | –13.713*** |
| | Equity return | –9.743*** | –8.304*** | –9.744*** | –54.797*** |
| Japan CDS | | –1.331 | –8.794*** | –1.687 | –8.810*** |
| 10-year JGB | | –1.818 | –8.482*** | –1.923 | –8.424*** |
| Yield curve slope | | –1.690 | –9.002*** | –1.690 | –8.985*** |

*Notes*

1. ** and *** denote significance at the 5% and 1% level, respectively. Test statistics are based on the specification including a constant term.

2. The number of lags is chosen by the Schwarz Criterion. The bandwidth for the Phillips–Perron test is based on the Newey–West bandwidth.

3. For the definition of variables, see Table 1.

## Table 3: Credit Spreads and Financial Market Variables

Sample Period (weekly) : April 2, 2004 to December 30, 2005 (Number of Observations: 92)

| | Subordinated CDS Spreads | | Subordinated Bond Spreads | |
|---|---|---|---|---|
| Bond volatility | | | 0.079 | 0.079 |
| | | | [0.062] | [0.062] |
| CDS volatility | 0.549*** | 0.567*** | | |
| | [0.139] | [0.139] | | |
| Japan CDS | 3.606*** | 3.584*** | 0.726*** | 0.731*** |
| | [0.532] | [0.526] | [0.210] | [0.209] |
| Equity return | –12.643** | –13.405*** | –1.628 | –1.467 |
| | [5.090] | [5.010] | [1.974] | [1.830] |
| 10-year JGB yield | –3.901 | | 0.598 | |
| | [2.586] | | [1.140] | |
| Yield curve slope | | –4.439 | | 0.161 |
| | | [3.315] | | [1.389] |
| Dummy | –37.686*** | –37.709*** | –20.367*** | –20.384*** |
| | [1.544] | [1.557] | [0.536] | [0.537] |
| Constant | –0.018 | –0.022 | –0.169** | –0.169** |
| | [0.177] | [0.177] | [0.083] | [0.084] |
| Num. of Obs | 364 | 364 | 364 | 364 |
| Method | Pooled | Pooled | Pooled | Pooled |
| R-squared | 0.356 | 0.355 | 0.538 | 0.537 |

*Notes*

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

2. Figures in square brackets are robust standard errors corrected for heteroskedasticity using White's method.

3. For the definition of variables, see Table 1. Dummy is equal to one for UFJ on July 16, 2004, when the merger with the BTM was reported. All the variables except for Bond volatility and Equity return are first-differenced because they are found to be I(1). See Table 2 for the results of the unit-root tests.

4. Method refers to the methodology chosen from the alternative estimation methodologies: pooled model, fixed-effects model, and random-effects model.

**Table 4: Definition of Accounting Variables**

| Accounting Variables | Definition | Source |
|---|---|---|
| Capital ratio | Shareholders' equity / total assets (%) | Financial statements |
| Adjusted capital ratio | Adjusted capital / total assets (%)<br>Adjusted capital = shareholders' equity – (loans to be written off – loan loss reserves) – deferred tax assets + (profit from revaluation of securities – equivalent of deferred tax liabilities + profit from revaluation of derivatives – difference in land price because of revaluation) | Financial statements |
| Nonperforming loan ratio | Nonperforming and past due loans divided by loans and bills discounted (%) | Financial statements |
| Size | Natural logarithm of total deposit outstanding | Financial statements |
| Lending growth | Growth rate of loans and bills discounted during semiannual period (%) | Financial statements |

## Table 5: Credit Spreads and Accounting Variables

Sample Period (semiannual): 1st half of FY 2002 to 1st half of FY 2005

| | Subordinated CDS Spread | | | Subordinated Bond Spread | | |
|---|---|---|---|---|---|---|
| Capital ratio | –43.180** | | | –7.239 | | |
| | [16.217] | | | [9.713] | | |
| Adjusted capital ratio | | –19.905*** | | | –9.213** | |
| | | [6.806] | | | [3.796] | |
| Nonperforming loan ratio | | | 12.086*** | | | 5.415*** |
| | | | [4.465] | | | [1.784] |
| Size | –53.682* | –79.722** | –73.915*** | –284.706*** | –27.724* | –181.210** |
| | [28.738] | [30.941] | [25.456] | [91.307] | [15.991] | [67.610] |
| Lending growth | –0.512 | –0.568 | –1.343 | –0.572 | –0.467 | –0.423 |
| | [1.442] | [1.099] | [0.943] | [0.835] | [0.556] | [0.399] |
| Constant | 1155.499** | 1515.504** | 1313.513*** | 5116.67*** | 555.086* | 3231.525** |
| | [548.343] | [561.813] | [442.515] | [1614.01] | [290.521] | [1204.26] |
| Num. of Obs. | 24 | 24 | 24 | 23 | 23 | 23 |
| Method | Pooled | Pooled | Pooled | Fixed-effects | Pooled | Fixed-effects |
| R–squared | 0.425 | 0.504 | 0.555 | 0.539 | 0.442 | 0.744 |

*Notes*

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

2. Figures in square brackets are robust standard errors corrected for heteroskedasticity using White's method.

3. For the definition of variables, see Table 4.

4. Method refers to the methodology chosen from the alternative estimation methodologies: pooled model, fixed-effects model, and random-effects model.

## Table 6: Unit Root Tests

Sample Period (daily): April 1, 2004 to December 30, 2005 (Number of Observations: 430)

| | | Augmented Dickey–Fuller Test | | Phillips–Perron Test | |
|---|---|---|---|---|---|
| | | Level | 1st difference | Level | 1st difference |
| BTM | CDS | –2.043 | –22.019*** | –1.984 | –22.019*** |
| | Bond | –1.890 | –29.800*** | –2.018 | –31.016*** |
| SMBC | CDS | –1.577 | –18.108*** | –1.628 | –18.238*** |
| | Bond | –1.424 | –19.095*** | –1.394 | –26.203*** |
| MIZUHO | CDS | –2.119 | –28.248*** | –2.090 | –29.264*** |
| | Bond | –1.573 | –28.154*** | –1.528 | –27.238*** |
| UFJ | CDS | –1.912 | –25.567*** | –1.890 | –26.262*** |
| | Bond | –1.424 | –19.649*** | –1.468 | –19.650*** |

*Notes*

1. *** denotes significance at the 1% level. Test statistics are based on the specification including a constant term.

2. The number of lags is chosen by the Schwarz Criterion. The bandwidth for the Phillips–Perron test is based on the Newey–West bandwidth.

## Table 7: Estimation Result of Price Discovery Measures

### (i) Johansen Cointegration Test

Sample Period (daily): April 1, 2004 to December 30, 2005 (Number of Observations: 430)

| | Number of Cointegrating Vectors | | | |
|---|---|---|---|---|
| | Trace Test | | Maximum Eigenvalue Test | |
| | None | At Most 1 | None | At Most 1 |
| BTM | 13.051 | 2.532 | 10.519 | 2.532 |
| SMBC | 18.810* | 3.119 | 15.692* | 3.119 |
| MIZUHO | 25.928*** | 7.122 | 18.806** | 7.122 |
| UFJ | 25.225*** | 4.087 | 21.139*** | 4.087 |

### (ii) Cointegrating Vector

| | Estimated Cointegrating Vector | Restriction on Vector |
|---|---|---|
| | [CDS, Bond, Constant] | [1, –1, C] |
| BTM | n.a | n.a. |
| SMBC | [1, –1.913***, 19.682** ] | 7.482*** |
| MIZUHO | [1, –2.513***, 44.899***] | 10.337*** |
| UFJ | [1, –1.712***, 24.591***] | 11.590*** |

### (iii) Contribution to Price Discovery

| | Estimated Adjustment Coefficients | | Hasbrouck | | | Gonzalo/ Granger |
|---|---|---|---|---|---|---|
| | $\lambda_1$ | $\lambda_2$ | Lower | Upper | Mid | |
| BTM | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. |
| SMBC | –0.015 | 0.017*** | 0.814 | 0.887 | 0.850 | 0.531 |
| MIZUHO | –0.066*** | 0.013*** | 0.455 | 0.517 | 0.486 | 0.166 |
| UFJ | –0.050** | 0.026*** | 0.401 | 0.816 | 0.608 | 0.342 |

*Notes*

1. *, **, and *** denote significance at the 10%, 5%, and 1% level, respectively. Test statistics are based on the specification including a constant term. The restriction test on the cointegrating vector is conducted by the LR (Likelihood Ratio) test.

2. The number of lags is chosen by the Schwarz Criterion.

## Table 8: Estimation Result of Bivariate GARCH Model

Sample Period (daily): April 1, 2004 to December 30, 2005 (Number of Observations: 430)

| | BTM | SMBC | MIZUHO | UFJ |
|---|---|---|---|---|
| ARCH | | | | |
| a11 | 0.468*** | 0.524*** | 0.529*** | 0.344*** |
| a12 | –0.027 | 0.055*** | –0.018** | 0.047*** |
| a21 | –0.014 | –0.044 | 0.037 | –0.023* |
| a22 | 0.127* | 0.137** | 0.171*** | 0.095* |
| GARCH | | | | |
| b11 | 0.871*** | 0.897*** | 0.881*** | 0.911*** |
| b12 | 0.022* | –0.005 | 0.012*** | –0.022*** |
| b21 | 0.205 | –0.221 | –0.107* | 0.115*** |
| b22 | 0.914*** | 0.410 | 0.955*** | 1.004*** |

*Notes*

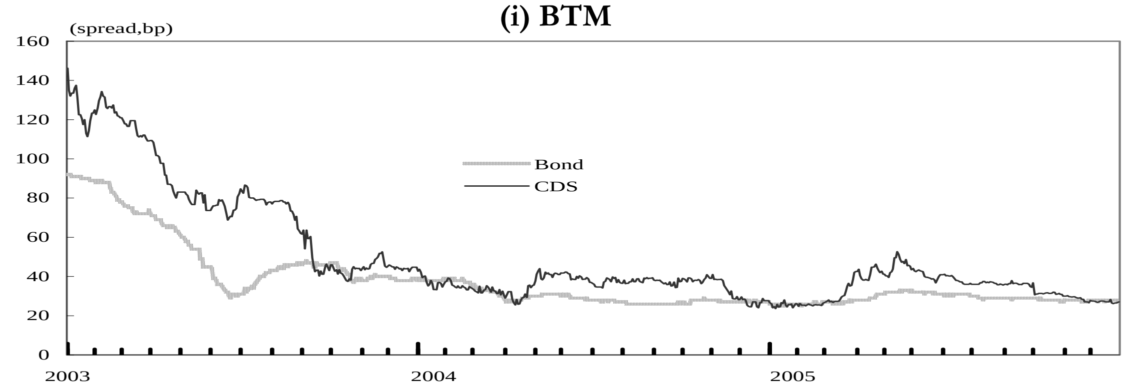
1. $a_{ij}$ and $b_{ij}$ measure the volatility spillovers from the *i*-th to *j*-th spreads (1: CDS, 2: Bond).

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

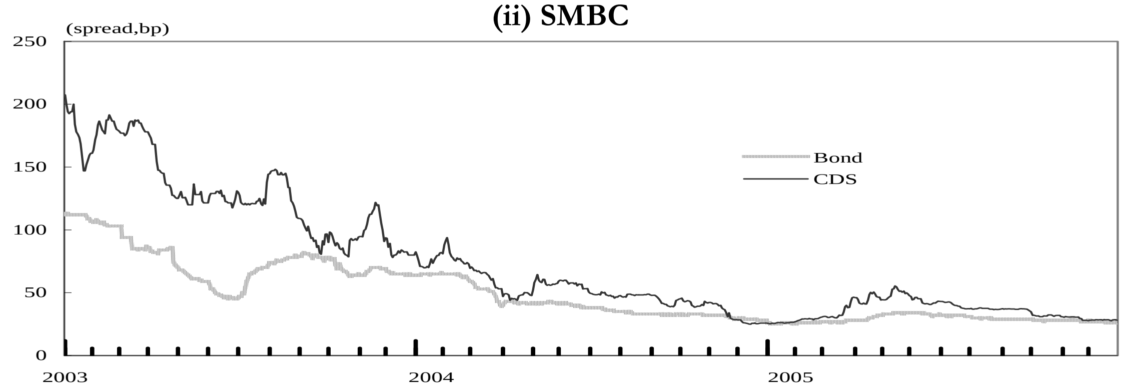
3. Estimation results of constant terms are omitted because of the limitation of space.

## Figure 1: Subordinated CDS and Bond Spreads

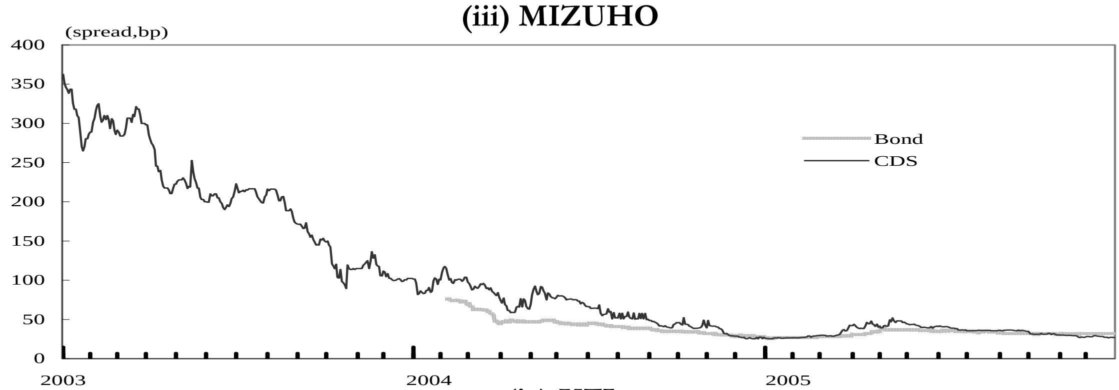
### (i) BTM



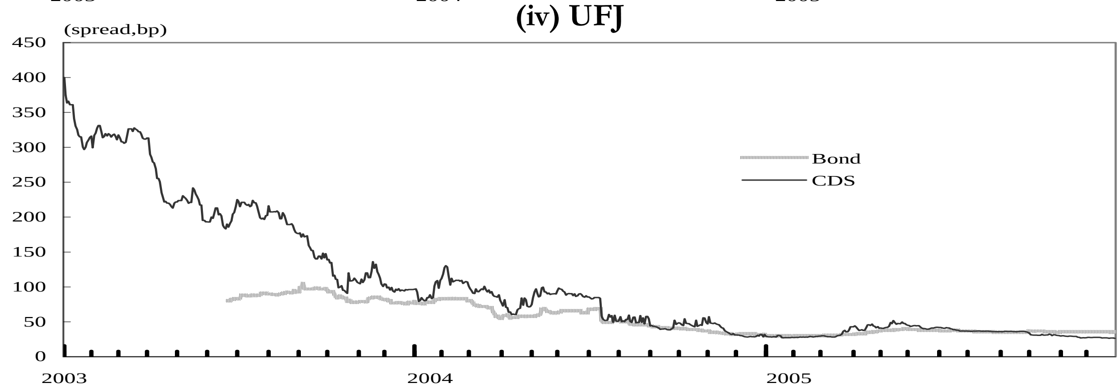

### (ii) SMBC




### (iii) MIZUHO




### (iv) UFJ




*Source*: Bloomberg.