Pricing Illiquidity of Corporate Bonds through Static and Dynamic Measures

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This paper studies the price impact of corporate bond illiquidity. Through dynamic panel estimation, price dispersion and resiliency, which have been used separately in extant studies, are simultaneously considered to price illiquidity. We find that the dynamic model, which has both measures, fits better than the static model that incorporates only price dispersion. We also confirm that the impact of the two measures systematically react to credit ratings of bonds. These results imply the importance of considering multiple measures to price illiquidity.

Keywords: Bond spreads, Illiquidity, Price dispersion, Persistency

JEL Classification: C23, C58, E44, G12

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I. Introduction

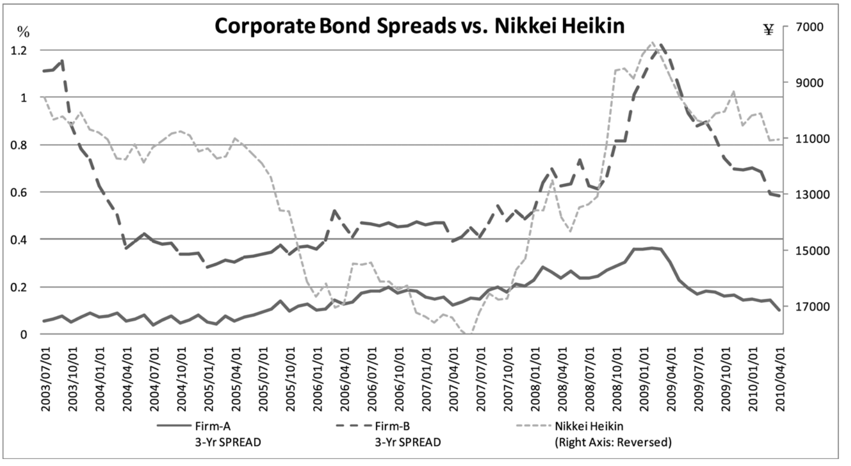
During previous financial crises and in the recent episode triggered by the Lehman shock, we have observed large time-series and cross-sectional variations in asset prices. These variations do not necessarily move well with each other and occasionally show rough dynamics (Figure 1). As demonstrated in a number of extant studies, classical asset pricing models that focus on credit risk factors cannot generally explain the reality of asset price variations (*e.g.*, Collin-Dufresne *et al.* 2001). With the presumption that the misalignment between model prediction and data occurs when liquidity of financial assets becomes low, the field of asset pricing has — over the last decade — focused on establishing appropriate proxies for liquidity factors.

Compared with extant studies that focus exclusively on a single illiquidity measure, the current paper studies the pricing impact of corporate bond illiquidity by simultaneously using static and dynamic measures of illiquidity. Among various illiquidity measures proposed in the literature, we consider price dispersion among multiple market makers (Houweling *et al.* 2005; Tychon and Vannetelbosch 2005; Kaguraoka 2010; Cici *et al.* 2011; Ou 2011), which we call GAP, as a static measure of illiquidity. GAP is the difference between the highest minus the lowest values of the simultaneously reported corporate bond spreads among multiple market makers. For the dynamic measure, we use price resiliency (*e.g.*, Kyle 1985; BIS 1999). A standard definition of resiliency is the rapid restoration of normal market prices. In this sense, low price resiliency is also quantified as the high persistency of prices. We study how static and dynamic measures affect the median spread of reported corporate bonds among multiple market makers by using a dynamic panel estimation framework.

In this paper, the asset price of corporate bonds is represented by "spread," which denotes the yield difference between similarly mature corporate bonds and Japanese Government Bonds (JGB). We particularly focus on the difference between linearly interpolated three-year corporate bond yield and government bond yield. Based on the presumption that government bonds are risk-free and are highly liquid (especially in relatively short-term maturities), we use the difference as a measure of credit and illiquidity risk premium, which is the dependent variable used in our empirical study.[1]

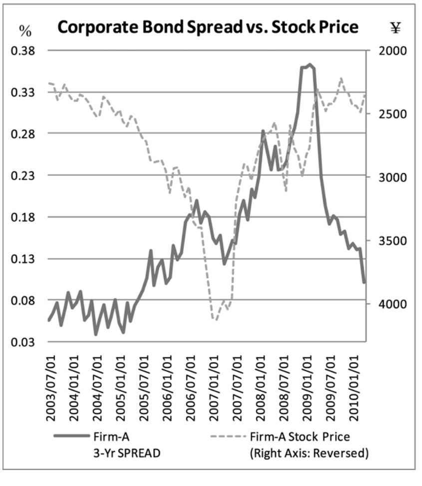
To motivate the study of corporate bonds, we briefly recall the illiquid

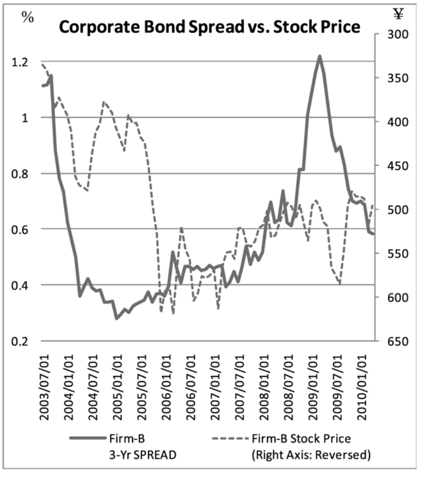
[1] We will discuss the validity of this presumption later.



Note: The bold solid line and bold dashed line show the three-year corporate bond spreads of two sample firms in our dataset. The fine dashed line accounts for the level of Nikkei 225 Index plotted on a reversed axis.

FIGURE 1(1)
CORPORATE BOND SPREADS AND STOCK INDEX





Note: Bold solid lines show the three-year corporate bond spreads of two sample firms in our dataset. The dashed lines account for the stock prices of those two firms plotted on a reversed axis.

FIGURE 1(2)
CORPORATE BOND SPREADS AND INDIVIDUAL STOCK PRICES

nature of corporate bond markets. Given that majority of investors of corporate bonds rely on the "buy and hold" type of investment strategy, the transaction volume is usually limited, and market making is very inactive. This is different from markets with frequent trading, such as stock and government bond markets.[2] Widely accepted liquidity proxies, such as bid-ask spreads, are not easily obtainable for corporate bonds. Consequently, the use of many major liquidity proxies referenced in other liquid markets becomes difficult.[3] Here, we choose price dispersion and resiliency to capture corporate bond illiquidity mainly because both are accessible even in illiquid markets.[4]

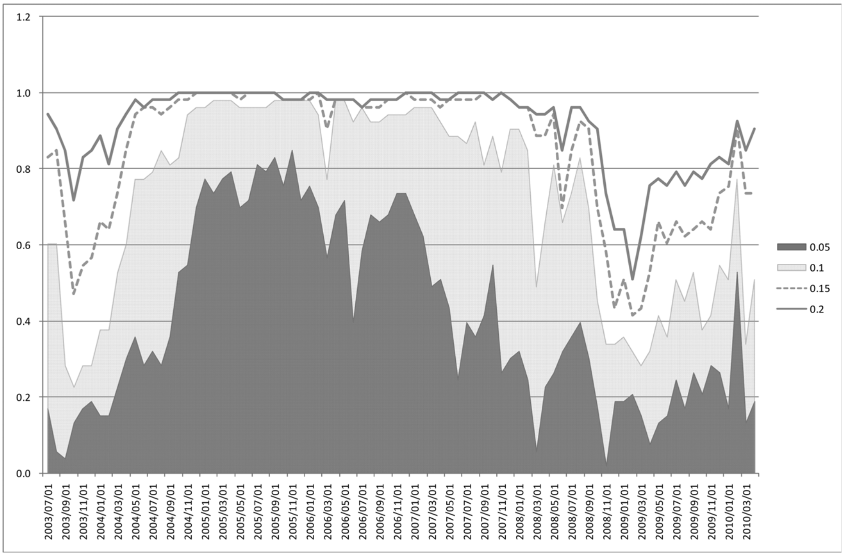
The Japanese corporate bond market, which is the focus of our study, is an appropriate environment for the study on price implication of bond illiquidity. This is mainly due to thick and highly liquid government bond markets, the existence of which allows us to reasonably employ the yield difference between corporate and government bonds as a measure of risk premium. Fortunately, there was no major change in tax treatment asymmetrically affects the return from corporate and government bond investments in Japan during our sample periods. In Friedman and Kuttner's (1993) discussion about the spread of U.S. commercial papers and treasury bills, the illiquidity risk of government bonds and their institutional feature (*e.g.*, tax treatment) can potentially affect the spread. In describing the characteristics of Japanese corporate and government bond markets, we can largely ignore these issues and study the price implication of credit and illiquidity factors in a relatively clean way.

To illustrate the behavior of price dispersion, Figure 2(1) reflects the share of firms that remain under a certain level of GAP (*e.g.*, equal to or smaller than 5, 10, 15, and 20 basis points) out of the total sample firms at a specific date. For example, the share of firms showing GAP that is smaller than 5 basis points is more than 80% in the second half of 2005. This was when the market environment was relatively good and the liqui-

[2] For U.S. stock and government bond markets, see information provided by Securities Industry and Financial Markets Association (SIFMA). SIFMA publishes "The SIFMA Fact Book," which is an annual data book on security markets.

[3] In fact, the Japan Security Dealers Association (JSDA), from which we obtained our data, has almost no data for bid-ask spreads although their tables include columns for recording them. Some exceptions to this lack of proxy references exist. One is Chen *et al.* (2007), who used quoted bid-ask spreads as well as percentage zeros and the LOT measure proposed by Lesmond *et al.* (1999), in their study of corporate bond prices.

[4] Another measure appropriate for illiquid markets is the latent liquidity measure proposed by Mahanti *et al.* (2008).



Note: Each line corresponds to the proportion of firms having GAP smaller than various percentages (*i.e.*, 0.05%, 0.1%, 0.15%, and 0.2%).

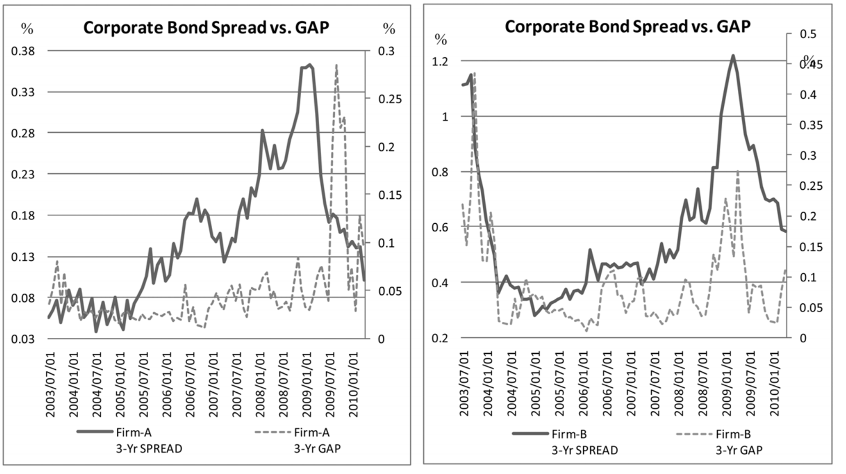
FIGURE 2(1)

DISTRIBUTION OF THE GAP MEASURE

dity was high. In contrast, shares were below 10% several periods after the Lehman shock. As discussed in the literature (*e.g.*, Houweling *et al.* 2005; Kaguraoka 2010), price dispersion can be generated by, for example, the difference in the prior of prices, the frictional trade environment, and/or by the slow diffusion of information. In any case, the static measure constructed from cross-sectional information naturally contains valid information about illiquidity.

As a concept, resiliency has been studied in the literature on market microstructure. Extant studies argue that illiquidity of corporate bonds may be inferred by checking whether or not reported quotes by market makers have persistency. If the market becomes highly illiquid and almost no transaction is carried out, a lesser portion of quotes depend on concurrent information and thus, stronger persistency is displayed. These discussions justify the usage of dynamic measure constructed from time-series information as a valid proxy for illiquidity.

While price dispersion and resiliency are certainly plausible as proxies for the illiquidity of corporate bonds, it is unclear whether or not they are mutually exclusive. In other words, the possibility of improving the preci-



Note: Bold solid lines show the three-year corporate bond spreads of two sample firms in our dataset. The dashed lines account for GAP of those two firms.

FIGURE 2(2)
INDIVIDUAL CORPORATE BOND SPREADS AND GAP MEASURES

sion of bond pricing by including multiple measures of illiquidity is still an open question. The purpose of our empirical analysis is to illustrate that the median level of reported spreads is correlated not only with standard covariates of bond spreads, but also with the dispersion of reported prices as well as with the level of past spreads. Through analysis, we aim to examine how static and dynamic measures of illiquidity complement each other.

We also attempt to revisit the well-known view called "flight-to-quality" (*e.g.*, Dick-Nielsen *et al.* 2012), which predicts that financial assets with low ratings rates can suffer more from the deterioration of liquidity. We are interested in finding to what extent our model with two illiquidity measures remain consistent with this view.

This paper is structured as follows. Section 2 briefly details the related literature. Section 3 formulates the hypotheses to be tested in subsequent sections. Section 4 illustrates our empirical framework. Sections 5 and 6 respectively present the data and the interpretation of the results. Section 7 concludes and suggests future research questions.

II. Related Literature

In asset pricing literature, two groups of studies are identified. The first group consists of works that perform classical empirical analysis for corporate bond spreads. The key implication shared among related literature is the insufficient ability of macro and/or micro credit factors to explain corporate bond spreads (*e.g.*, Jarrow *et al.* 2000; Edwin *et al.* 2001; Huang and Huang 2003; Eom *et al.* 2004). Collin-Dufresne *et al.* (2001) confirmed that credit risk factors cannot explain bond spreads even after incorporating various aggregate variables (*i.e.*, S&P 500 returns, slope of government bond yields, and government yields) into the estimation. This empirical finding is also discussed in another literature as the Credit Spread Puzzle. Investigations by Hull *et al.* (2004, 2005) show that the default probability implied by the corporate bond yield in the secondary market is more than ten times the default rate calculated from historical data. They also identify the significant gap between corporate bond spread and the premium of Credit Default Swap (CDS). Our motivation in introducing the viewpoint on liquidity premium, especially through simultaneous employment of static and dynamic measures, is to partially resolve the inconsistency between classical model prediction and observed data.

The results of the first group of works naturally motivate the second group, which perform empirical analysis by incorporating an illiquidity factor. Most of these studies first choose an appropriate proxy which is supposed to capture market-level liquidity. To evaluate the performance of various proxies, majority of extant studies (*e.g.*, Fleming 2003; Goldreich *et al.* 2005) in this direction use government bond yield, which is theoretically unaffected by the credit factor. With the control of established market-level liquidity proxies, several papers then began to incorporate an additional individual liquidity factor. Amihud (2002), for example, pioneers the measurement of individual illiquidity by introducing the ILLIQ measure, which is computed as the average daily variation in stock return.[5] Houweling *et al.* (2005) comprehensively examine how illiquidity factors work. Using European data, they apply a standard multi-factor model and separately establish the validity of several illiquidity measures.

[5] Precisely speaking, the ILLIQ measure is computed as the absolute price change divided by the trading volume of a given stock per day. Given that this ratio is actually a very noisy measure on any day, the average of all trading days in a month or in a year is derived to obtain a monthly or an annual liquidity estimate for the targeted stock.

Kaguraoka (2010) employ a price dispersion measure called "yield discrepancy" under a static panel estimation framework to measure the illiquidity of Japanese corporate bonds from 2002 to 2004. Our paper shares a very close motivation with the above-mentioned studies. The difference is that we use dynamic panel estimation techniques to explicitly incorporate additional liquidity proxies (*i.e.*, persistency) under a dynamic panel estimation framework. For the quantification of pricing implication of illiquidity risk for stocks, Acharya and Pedersen (2005) directly incorporate the notion of "transaction cost" to a standard CAPM structure and explicitly evaluate illiquidity risk.[6] Aside from the above papers based on a standard factor model specification, Bao *et al.* (2008) examine the price dynamics of corporate bonds and establish a proxy for corporate bond market liquidity. Our paper shares the belief that price dynamics (*e.g.*, persistency) has information related to market liquidity.[7]

Among empirical studies on the correlation between corporate bond spreads and cross-sectional price dispersion, the latter — which we use as a static measure for illiquidity — is also theoretically motivated. A theoretical model closest to our motivation is that proposed by Tychon and Vannetelbosch (2005) who develop a corporate bond valuation model that has both credit and liquidity/marketability risks generated by a matching friction between bondholders and potential investors in a secondary market. Tychon and Vannetelbosch (2005) assume that each bondholder can be matched with a limited number of potential buyers in an illiquid market, which may not result in the match-up with positive gain from the trade. Meanwhile, matching is guaranteed to always generate the positive gain in a liquid market.[8] The main idea is that the liquidity premium originates from the interaction between (i) the difference in matching frictions under illiquid and liquid market environments and (ii) the heterogeneity of investors' valuations. To summarize, as the heterogeneity of belief becomes larger, bondholders/investors place lower

[6] Although our paper does not have an explicit asset pricing formulation, the model structure is actually an extension of the model first proposed by Fama and French (1993). Thus, to some extent, we follow a traditional factor model/CAPM structure.

[7] Bao *et al.* (2008), referring to extant theoretical studies such as that of Roll (1984), hypothesize that the amount of price reversal or the negative of autocovariance of prices is associated with the illiquidity of corporate bonds.

[8] Therefore, matching in the liquid market is frequently to ensure that each bondholder can be matched with an investor who places higher evaluation for his bond. The authors assumed that the liquidity market always makes it possible to achieve profitable matching.

relative values (*i.e.*, higher spread) on bonds in the illiquid market than those in the liquid market as the heterogeneity of belief becomes larger. Bondholders and investors are less likely to find an appropriate counterpart for bond trading in the future when the matching friction is amplified by larger heterogeneity of bond evaluation.[9] The recently developed "differences of opinion" literature (Banerjee *et al.* 2009; Cao and Ou-Yang 2009; Banerjee and Kremer 2010) further pursues this direction. Authors of these studies construct a theoretical model that incorporates investors with heterogeneous beliefs about common public information and characterizes how such heterogeneous views are reflected in a market maker's opinion, in market prices (average opinion), and in trading volumes through the learning process.[10] From Houweling *et al.* (2005), we consider these discussions as theoretical underpinnings of our empirical study.

Studies on the Japanese corporate bond market are limited. First, Hongo and Oyama (2010) study the mechanism governing corporate bond spreads using a model without a liquidity factor. Second, Saito *et al.* (2001) examine the importance of a liquidity factor by featuring the liquidity demand described by Holmstrom and Tirole (2001). Shirasu and Yonezawa (2008) also challenge the same question using Japanese corporate bond market data. Third, Nakamura (2009) employ two methods to quantify the illiquidity risk on bond spreads. Differences include the incorporation of static and dynamic proxies for bond illiquidity and the estimation strategy.

III. Hypothesis Formulation

In this section, we mention several theoretical statements that motivate our empirical study. First, we review Tychon and Vannetelbosch (2005) who illustrate the connection between diverged price evaluation and corporate bond spreads. Second, we discuss the conditionality of the illiquidity premium on the firm-level risk. We also demonstrate a theoretical illustration of the relationship between resiliency and illiquidity, as well as its conditionality on firm-level credit risk. For all items, no attempt

[9] Consider an environment with no heterogeneous beliefs. Regardless of matching friction, no trade emerges and the heterogeneity among bondholders/investors does not matter.

[10] Easley and O'Hara (2010) construct a model that generates equilibrium quotes and the non-existence of trading at the quotes, which can be observed during the financial crisis period.

is made to provide an exhaustive survey or produce original theoretical models. We simply intend to establish a conceptual framework for reference in the empirical section.

As a first step of our empirical study, we consider a static panel estimation framework that incorporates standard covariates and GAP. To illustrate the inclusion of GAP into our empirical study, we recall the model used by Tychon and Vannetelbosch (2005). As described in the previous section, the framework demonstrates a linkage between the opinion difference and the bond spreads. The liquidity premium in the model of Tychon and Vannetelbosch (2005) originates from the interaction between the less frequent trading in illiquid markets and the opinion difference among investors. The model clearly illustrates the emergence of the liquidity premium when the economy has heterogeneous beliefs. We use this perspective to construct our first hypothesis, which aims to confirm the validity of our static GAP measure. We also use empirical results based on the static model to evaluate the performance of the dynamic panel estimation.

***Hypothesis* 1:** Median quoted spreads (*i.e.*, three-year SPREAD), which we regard as representative market prices, are positively correlated with absolute dispersions of market makers' quoted prices that are captured by the GAP measure in our setup, after controlling market and individual factors.

For (i) the market, we use the slope of Japanese government bond (JGB) yield curves as represented by the 10-year JGB yield minus the two-year JGB yield (*JGBSLOPE*), the 10-year JGB yield (*JGB10Y*), and the level of the Nikkei Average Stock Index (*NKY*); for (ii) the individual credit factors, we use the abnormal historical volatility of individual stocks and the credit ratings (*eHV* and *RATE_RI*);[11] and for (iii) the market liquidity factor, we use the three-month Tokyo inter-bank offered rate (Tibor) minus the three-month JGB yield, which represents the tightness of the money market (*T_JGBGAP*).

The inclusion of *JGBSLOPE*, *JGB10Y*, and *NKY* is based on the discussion made by Friedman and Kuttner (1993), who report that market-level credit risk factor affects the yield difference between various classes of security and government bonds. To check robustness, we also employ

[11] The unexpected historical volatility is estimated as the residual obtained from the regression of the individual historical volatility of stock prices on the historical volatility of the Nikkei average index.

the spread of three-month commercial papers (*i.e.*, CP) as an alternative proxy for market-level credit risk. This is computed as the difference between the average yield of commercial papers and the yield of JGB in the same maturity level (*i.e.*, three-month: *CP_JGB3M*). Considering high correlations with *CP_JGB3M*, we exclude *T_JGBGAP* in the robustness check. Note that *CP_JGB3M* is interpreted as the proxy for both market-level credit and liquidity factors, because it strongly reflects the tightness of the short-term money market.

To explicitly control the condition of foreign financial markets that may affect the attitudes of foreign investors, we employ the three-month dollar London inter-bank offered rate (*DLIBOR*). As discussed in extant studies (*e.g.,* Cook and Hahn 1989; Bernanke and Blinder 1992), market interest rates are largely affected by monetary policies. Thus, we also include the uncollateralized overnight call loan rate (*ON_AVERAGE*) to our model, which is the target of Japanese monetary policy.[12] We summarize the expected signs of each coefficient in the succeeding section.

For our second hypothesis, we consider the extant literature that emphasizes on systematic conditionality on credit ratings of coefficients associated with liquidity factors for stock returns (*e.g.*, Watanabe and Watanabe 2008). This also reflects the view called “flight-to-quality” (*e.g.*, Dick-Nielsen *et al.* 2012). A key premise behind this view is that investors rush into safer assets when they perceive a larger risk. In other words, extant theoretical discussions presume that the liquidity premium interacts with the credit-risk factor. The liquidity spiral story discussed by Brunnermeier (2009) provides an illustrative example for this. In the liquidity spiral story, liquidity in financial markets can suddenly evaporate due to the enhanced mechanism between two items. The first item is the easy conversion of assets into cash, while the second item represents the easy borrowing procedure for investors (see also Brunnermeier and Pedersen 2009). When a severe shock hits financial markets, it results in high difficulty for investors with weak funding availability to borrow and/or to recover from their debts. Thus, they may be forced to instantly sell their assets. The selling of assets can inevitably place burden on their borrowing ability brought about by the value depreciation of their collateralized assets. In this case, assets with higher ratings can be sold immediately, which worsens liquidity of the remaining

[12] Considering the high correlation between these additional variables and the macro-factors employed in the baseline model, we exclude some of the variables when we incorporated *DLIBOR* and *ON_AVERAGE*.

low-rated assets. This is the mechanism we theorize in this paper to formulate our second hypothesis.

***Hypothesis* 2:** The GAP measure's quantitative impact on bond spreads becomes larger (smaller) as credit ratings become worse (better).

This hypothesis can be tested by checking whether or not the coefficient associated with the GAP measure multiplied by the standard deviation of the GAP measure is different among the samples that were split based on credit ratings.

Third, we add the dynamic resiliency factor to the static model. As intensively discussed in various empirical studies (*e.g.*, BIS 1999), illiquidity in the financial market is "multi-faceted." However, extant empirical studies have mostly focused on a single index to measure and to price illiquidity. The main purpose of this paper is to improve the model fit by adding the dynamic resiliency factor, which is examined by testing the following hypothesis.

***Hypothesis* 3:** GAP and resiliency (*i.e.*, persistency on spreads) measures have statistically significant coefficients.

We also hypothesize that the conditionality considered in the second hypothesis is also applicable to the resiliency factor. We test this hypothesis by checking whether or not the coefficient of the lagged dependent variable varies with credit ratings.

***Hypothesis* 4:** The quantitative impact of the resiliency term on bonds with spreads becomes larger (smaller) as credit ratings become worse (better).

IV. Empirical Framework

We discuss the empirical framework in this section. Unlike the typical time-series estimation based on sorted hypothetical portfolios employed in the extant literature, we use panel estimation with a balanced panel data of monthly Japanese corporate bond spreads, which are detailed in the next section. The benefits of employing panel estimation are twofold. First, it can fully extract both the time-series and the cross-sectional properties of our firm-level data. Thus, by applying the panel estimation framework to this dataset, we can precisely examine potential determinants of individual corporate bond spreads. Second, our dynamic panel estimation enables us to establish empirical implications of our GAP measure as well as persistency in median quoted bond spreads.

Based on the literature on asset pricing, we start from a simplified version of a proposed multi-factor model, such as the model proposed by Fama and French (1993).

$$SP_{it}=\beta_1+\beta_2 F_t+\beta_3 L_t+\alpha_i+\varepsilon_{it} \tag{1}$$

Here, the dependent variable SP_{it} denotes the spread of firm i at time t, which is computed as the difference between the three-year JGB yield and the linearly interpolated corporate bond yield.[13] As an explanatory variable, F_t denotes various market credit indexes at time t (*i.e.*, stock market index, information related to the government bond yield curve, the spread of commercial paper, variable related to monetary policy, and variable related to foreign financial markets). In addition to these aggregate credit factors, we incorporate the market liquidity factor L_t represented by Tibor minus JGB spread, which indicates the tightness of short-term financial markets. As mentioned in the previous section, we use the average of commercial paper spreads as another proxy for market liquidity and for market credit index. We attempt to confirm that each coefficient has an expected sign implied by extant theoretical studies. According to standard panel estimations, α_i captures the firm-specific individual effect, which acquires either a fixed value (fixed-effect model) or an independent random variable (random-effect model) for each group with a zero mean and with the standard deviation σ_α. ε_{it} denotes the error term that is allowed to have some correlation within the same i by using the cluster-robust standard error.

Subsequently, we advance to the extended version of the model by incorporating f_{it}, which denotes the individual firm-specific credit risk factor of firm i at time t. We use credit ratings and historical abnormal volatility of stock returns as proxies for f_{it}. Credit ratings are defined as discrete numbers from 2 to 11. The larger number corresponding to a worse credit rating, which is reported by the Rating and Investment Information, Inc (R&I).[14] Historical abnormal volatility of stock returns

[13] In the next section, we will describe in detail the process of computing this number.

[14] If firms in our sample do not have their credit ratings from R&I but from other rating agencies, we transform these into hypothetical R&I ratings (*e.g.*, AAA=1, AA+=2, and so on). This transformation is done by referring to the

is computed as the regressing residual of the estimation of the 20-day historical volatility of individual stock prices on that of *NKY*.[15]

$$SP_{it} = \beta_1 + \beta_2 F_t + \beta_3 L_t + \beta_4 f_{it} + \alpha_i + \varepsilon_{it} \tag{2}$$

The model below adds the individual liquidity proxy l_{it} to the extended multi-factor model above.

$$SP_{it} = \beta_1 + \beta_2 F_t + \beta_3 L_t + \beta_4 f_{it} + \beta_5 l_{it} + \alpha_i + \varepsilon_{it} \tag{3}$$

The additional factor l_{it} is represented by GAP, which denotes the highest minus the lowest reported spreads among market makers at time t.[16] Note that we observe that most corporate bonds, including those issued by high credit utility companies (*e.g.*, companies who generate electricity), maintain a significant GAP in our sample period (see Figure 2(2)).[17] Our concern is whether, and if so how, model (3) is better than models (1) and (2), and whether or not the signs of estimated coefficients are qualitatively consistent with our predictions. The estimation is implemented by fixed-effect and random-effect estimations.

From a technical point of view, we may need to control the level of reported spreads to incorporate appropriately the GAP variable. One method is by simply including the level of the highest or the lowest quoted spread and the GAP itself. Another method is to construct the so-called "relative distance measure," similar with Houweling *et al.* (2005), which is calculated by GAP divided by an appropriate-level variable (*e.g.*, concurrent JGB yield). We employ the second method in a later section as a checking procedure for robustness of our results.

Finally, model (4) is used to implement dynamic panel estimation. Specifically, the lagged dependent variable SP_{it-1} is incorporated. As mentioned in the previous section, this formulation is motivated by the notion of resiliency. Supposedly, the coefficient γ assumes a value close to one, which implies that the median reported spread is persistent, if the market exhibits low resiliency. We then become interested in how model (4) is better than model (3).

companies holding both R&I ratings and ratings provided by other agencies.

[15] The estimated historical volatility is obtained by regressing individual historical volatility on that of the Nikkei stock average index (*i.e.*, NKY).

[16] We use one-day lagged GAP to avoid a simultaneity bias.

[17] In Japan, electricity, gas, and other utility companies have maintained high ratings due to institutional reasons.

$$SP_{it} = \gamma SP_{it-1} + \beta_1 + \beta_2 F_t + \beta_3 L_t + \beta_4 f_{it} + \beta_5 l_{it} + \alpha_i + \varepsilon_{it} \tag{4}$$

As Arellano (2003) carefully demonstrates, dynamic models with lagged dependent variables are suitable for the estimation of economic variables with adjustment costs and/or habit formation. In the current context, persistency represents the dependency of the bond evaluation of market makers on the spreads reported in the previous period due to insufficient information obtained from transactions. Estimation for this dynamic model is implemented through fixed-effect estimation, random-effect Generalized Least Squares (GLS) estimation, and random-effect Maximum Likelihood Estimation (MLE).[18]

A few points need to be discussed for model selection. First, we prefer MLE than that with other specifications. With the existence of individual effect (either fixed or random), OLS obviously cannot provide a consistent estimator. Moreover, fixed-effect and random-effect estimations with a lagged dependent variable also cannot provide a consistent estimator.[19] Although an Arellano-Bond GMM estimator that considers this problem is widely used for dynamic models, recent studies emphasize that the over-identification restriction test does not work on long time-series data. For example, based on the full instrument set, the Sargan-test is essentially never satisfied when T (and hence the number of moment conditions) becomes too large for a given value of N (Bowsher 2002). Since our sample contains a relatively large T (*i.e.*, 82), while N is just 52, GMM cannot be used.[20] In order to apply MLE where we can ignore the issue of correlation between residuals and lagged dependent variables to our dynamic model, we need to determine in advance the distribution for the initial dependent variable (*i.e.*, SP_{i0}). Fortunately, the large T of our sample alleviates this problem; thus, we can almost ignore the initial observation problem.[21]

[18] Although the results are not shown in this paper, we mention the employment of the Arellano-Bond GMM estimation (see Arellano and Bond 1991). The results are very close to those obtained by a random-effect MLE, unless otherwise noted.

[19] Nonetheless, we show the results of fixed and random effects in model (4) for comparison with other estimations.

[20] Blundell and Bond (1998) also point out that the instruments used in standard first-difference GMM estimators are less informative in two important cases: when the value of γ increases toward unity and when the relative variance of fixed effect increases. We observe that the long T problem is prominent in our analysis.

[21] The key idea is that the effect coming from the initial observation for the

Second, in relation to the first point, the bias of the estimator associated with the random effect is shown to be smaller as the time-invariant standard deviation σ_α also becomes smaller. Hsiao (2003) concisely shows the following expression for the estimated AR(1) coefficient:

$$\hat{\gamma} = \frac{\Sigma_{i=1}^{N}\Sigma_{t=1}^{T} SP_{it} SP_{it-1}}{\Sigma_{i=1}^{N}\Sigma_{t=1}^{T} SP_{it-1}^2} = \gamma + \frac{\Sigma_{i=1}^{N}\Sigma_{t=1}^{T} (\alpha_i + \varepsilon_{it}) SP_{it-1}}{\Sigma_{i=1}^{N}\Sigma_{t=1}^{T} SP_{it-1}^2} \tag{5}$$

where

$$\begin{aligned} &\text{plim}_{N\to\infty} \frac{1}{NT} \sum_{i=1}^{N} \sum_{t=1}^{T} (\alpha_i + \varepsilon_{it}) SP_{it-1} \\ &= \frac{1}{T} \frac{1-\gamma^T}{1-\gamma} \operatorname{cov}(SP_{i0}, \alpha_i) + \frac{1}{T} \frac{\sigma_\alpha^2}{1-\gamma} [(T-1) - T\gamma + \gamma^T] \end{aligned}$$

The last term at the right-hand side of the second expression becomes smaller as σ_α becomes smaller. We use this property to evaluate our estimation results in a later section.

V. Data

We obtained data from the Japan Securities Dealers Association (JSDA), which has reference data on Japanese corporate bond markets. The data on daily frequency consist of the highest, the lowest, mean, and median bond yields of all companies reported by selected market makers (*i.e.*, "member security firms"), after excluding outliers based on predetermined rules.[22] These member security firms report the yield of each company's existing bonds at different maturity levels by 4:30 pm every

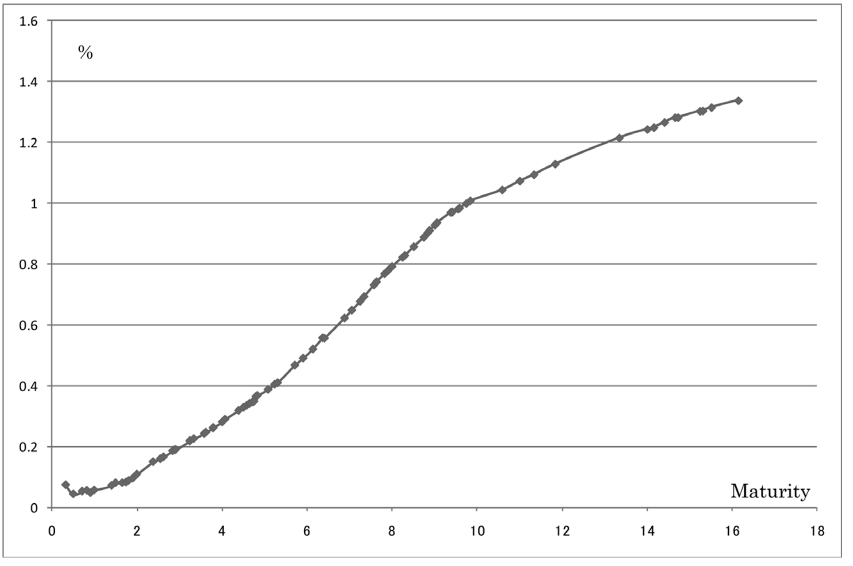
estimation of model parameters through MLE becomes almost negligible as the length of data becomes longer (Hsiao 1986). Meanwhile, the validity of assuming normal distribution remains the same.

[22] The selected market makers are as follows: SMBC Friend, Okasan, Credit Swiss, Cosmo, Goldman Sachs, Citigroup, Shinsei, JP Morgan, Daiwa SMBC, Deutsch, Tokai Tokyo, Nomura, BNP Paribas, Marusan, Mizuo, Mizuho Investors, Mitsubishi UFJ, Merrill Lynch Japan, Morgan Stanley, and UBS. If the number of reporting security firms is 15 to 20, the three security firms with the highest yields and the three security firms with the lowest yields are excluded. The criterion for exclusion is as follows: among 10 to 14 security firm reports, we determine two firms with highest yields and two firms with lowest yields; in addition, among 5 to 9 security firm reports, we exclude the highest and the lowest.

TABLE 1
SUMMARY STATISTICS

Variable		Definition	Unit	Expected Sign	Obs	Mean	Std.	Min	Max
SP_t	3-Yr SPREAD	3-year Corporate bond yield minus JGB yield	%		4173	0.40	0.45	-0.11	4.21
L_t	T_JGBGAP	3-month Tibor minus 3-month JGB yield	%	+	4264	0.19	0.14	-0.03	0.54
F_t	JGBSLOPE	10-year JGB yield minus 2-year JGB yield	%	− / +	4264	1.06	0.22	0.69	1.64
F_t	JGB10Y	10-year JGB yield	%	− / +	4264	1.49	0.20	0.95	1.93
L_t, F_t	CP_JGB3M	3-month commercial paper yield minus JGB yield	%	+	3952	0.15	0.21	-0.05	1.22
F_t	NKY	Level of Nikkei stock index	One thousand Yen	−	4264	12.77	2.92	7.57	18.14
F_t	ON_AVERAGE	Uncollateralized Overnight Call Rate (Target of Japanese Monetary Policy)	%	− / +	4264	0.18	0.20	0.00	0.52
F_t	DLIBOR3M	US Dollar Libor (London Interbank Offered Rate)	%	− / +	4264	2.91	1.84	0.25	5.62
f_{it}	eHV	20-day historical volatility of individual stock minus estimated historical volatility of each individual	%	+	4264	0.00	13.69	-76.03	83.84
f_{it}	RATE_RI	R&I credit ratings (the number increases as the credit risk increases)	Numeric measure from 2 to 11	+	4264	5.63	2.53	2	11
l_{it}	GAP3_1DLAG	Highest reported yield minus lowest reported yield in the previous day of 3-Yr SPREAD	%	+	4172	0.09	0.13	0.00	3.95
l_{it}	GAP3_1DLAG_Adj	GAP3_1DLAG divided by concurrent JGB yield	%	+	4172	0.25	0.34	0.01	8.51

Note: All the numbers are computed from the total sample. Expected sign indicates the sign of each coefficient, which is expected from theoretical discussion.



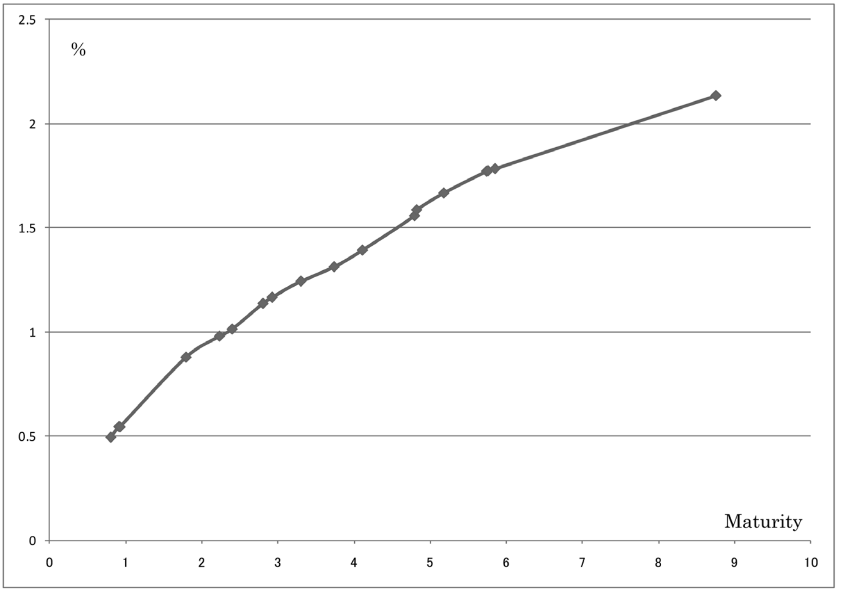
Note: Each dot corresponds to the reported yields from JSDA data. The three-year yield data used for computing the spread are obtained by interpolating those points linearly.

FIGURE 3(1)
INTERPOLATED YIELD CURVE FOR FIRM-A (JULY 31, 2003)

business day and then determine the price of 500 million yen worth of bonds as of 3:00 pm on the same day. JSDA collects the data and releases four reference data mentioned above on their website by 5:30 pm every business day.[23]

We only consider the data of listed firms that maintain a certain number of issued bonds, enabling us to calculate the spread for the three-year tenor. Note that most existing works incorporate maturity level as an explanatory variable without adjusting yields by maturity. From the standard practical viewpoint, assuming a linear relationship between spread and maturity is too restrictive. We usually observe a non-linear yield curve in reality. For a given company, we draw its yield curve and calculate its three-year yield by linear interpolation. We apply this procedure to construct the yield corresponding to the highest, the lowest, mean, and median reported values. Figures 3(1) and 3(2) exhibit the

[23] If the gap between the highest and the lowest yields is more than 500 bp, the data are not released.



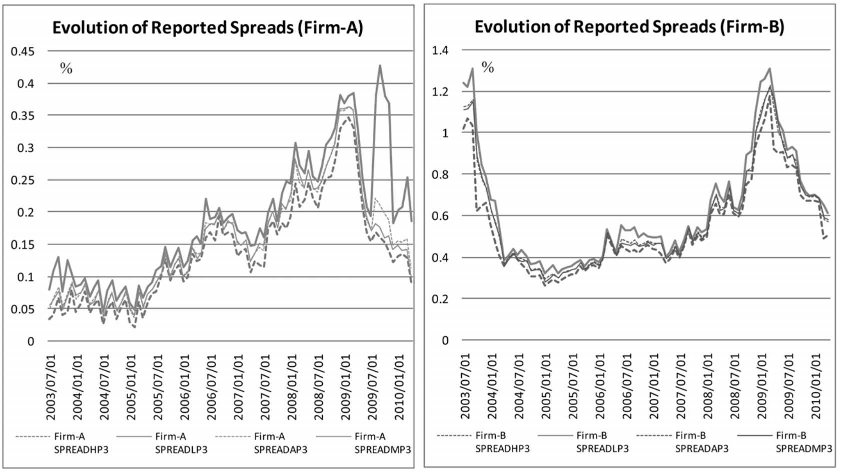
Note: Each dot corresponds to the reported yields from JSDA data. The three-year yield data used for computing the spread are obtained by interpolating those points linearly.

FIGURE 3(2)
INTERPOLATED YIELD CURVE FOR FIRM-B (JULY 31, 2003)

resulting median yield curves for two firms in our sample, respectively. Our targeted spreads corresponding to three-year maturity levels are then extracted from the median yield curve and from the concurrent JGB yields.[24] With this operation, we do not need to control maturity levels as explanatory variables. With the shape of the time-series data constructed and shown in Figure 4, we believe this interpolation does not generate any critical problems.

From the data, we use the median spread as our dependent variable. This is computed as the difference between the linearly interpolated median of the three-year corporate bond yield and that of the three-year Japanese government yield. We use the difference between the highest

[24] Estimation results based on five-year spreads are provided upon request. Constructing 10-year spreads is apparently difficult due to the insufficient amount of long-term corporate bond data. Meanwhile, we do not have one-year spreads in our analysis simply because bonds that are very close to maturity tend to exhibit irregular dynamics.



Note: SPREADHP3 and SPREADLP3 refer to the spreads corresponding to highest price (hence lowest spread) and lowest price (hence highest spread), respectively. SPREADAP3 and SPREADMP3 refer to the average and median reported spreads, respectively. These data are computed from the interpolated yield curve and the concurrent JGB yield.

FIGURE 4

INTERPOLATED SPREADS (HIGHEST, LOWEST, AVERAGE, MEDIAN)

and the lowest yields (GAP) as one of two illiquidity factors. The median is more ideal than the mean because it is not directly affected by the highest and lowest spreads. Considering that the data are significantly persistent, we also use data with monthly frequency. Table 1 shows the summarized statistics of our explanatory variables and their definitions, while Table 2 summarizes the correlation.[25]

For the market-level liquidity proxy (*T_JGBGAP*) and the individual credit factors (*eHV* and *RATE_RI*), we expect a positive correlation with the median reported spread. For the macro factors and for the monetary policy-related variables (*i.e.*, *JGBSLOPE*, *JGB10Y*, *ON_AVERAGE*, and *DLIBOR3M*), existing literature provides mixed implications. For example, the steeper slope of the government yield curve can be interpreted as a precursor of better economic environment as well as a risk in elevating future interest rates (*e.g.*, Fama and French 1993). Similarly, the lower uncollateralized overnight call loan rate can be a proxy of liquid short-

[25] As a reference for the estimation based on the sample splits, the summarized statistics and the correlation tables are provided upon request.

TABLE 2

CORRELATION TABLE FOR ALL SAMPLES

	3-Yr SPREAD	3-Yr SPREAD (Lagged)	T_ JGBGAP	JGBSLOPE	JGB10Y	CP_ JGB3M	NKY	ON_ AVERAGE	DLIBOR3M	eHV	RATE_ RI	GAP3_ 1DLAG	GAP3_ 1DLAG_ Adj
3-Yr SPREAD	1.00												
3-Yr SPREAD (Lagged)	0.97	1.00											
T_JGBGAP	0.43	0.41	1.00										
JGBSLOPE	-0.17	-0.14	-0.48	1.00									
JGB10Y	-0.16	-0.15	-0.38	-0.05	1.00								
CP_JGB3M	0.29	0.23	0.73	-0.45	-0.27	1.00							
NKY	-0.26	-0.25	-0.45	-0.33	0.71	-0.35	1.00						
ON_AVERAGE	0.06	0.04	0.39	-0.86	0.26	0.30	0.43	1.00					
DLIBOR3M	-0.26	-0.27	-0.41	-0.39	0.61	-0.15	0.89	0.41	1.00				
eHV	0.31	0.30	0.10	-0.07	-0.03	0.09	-0.02	0.02	-0.02	1.00			
RATE_RI	0.43	0.43	-0.04	0.04	0.01	-0.03	0.01	-0.04	0.01	0.33	1.00		
GAP3_1DLAG	0.70	0.65	0.32	-0.07	-0.19	0.26	-0.31	-0.03	-0.31	0.26	0.25	1.00	
GAP3_1DLAG_Adj	0.60	0.57	0.19	0.15	-0.37	0.11	-0.44	-0.25	-0.47	0.21	0.26	0.90	1.00

Note: Number of observations is 3823.

term financial markets as well as a result of lower policy rate corresponding to a poor economic environment. Thus, we prefer to treat these variables as control variables without initially assigning specific expected signs. For the rest of the macro factors (*i.e.*, *NKY* and *CP_JGB3M*), we expect negative and positive correlations with a median reported spread, respectively. The adjusted GAP measure obtained by dividing the spreads by concurrent JGB yields is used for the robustness check of our estimation results. This represents the micro liquidity factor with an expected positive coefficient.

From almost 120 firms in our original sample, the number of groups (*i.e.*, firms) is reduced to 52 because of the maturity level control explained above. However, our sample still contains a large variation in credit ratings and in the GAP measure. The sample period is 82 months, spanning July 2003 to April 2010.[26] Although the original data cover a longer time period, we focus on the period when the GAP measure is available.

VI. Estimation Results

In this section, we implement the estimations proposed in the previous section. All results from the static panel estimations are summarized in Table 3. With the target variable (*i.e.*, GAP) defined in a firm level, we allow the correlation of observations within the same group i (*i.e.*, firm). As a standard treatment in panel studies, we use the firm-level cluster-robust standard error to evaluate estimated coefficients.

First, the models without the GAP measure, which are introduced in (1) and (2), are estimated by the random-effect model with the cluster-robust standard error.[27] Then, the models with the GAP measure, which are introduced in (3), are estimated by the fixed-effect model, the random-effect model, and MLE.[28] After the standard model specification tests, we confirm that the fixed-effect model is more suitable than pooling OLS (see F-test results in Table 3) and that the random-effect model is more suitable than the pooling OLS model (see results of the Breusch

[26] Due to the limitation of data availability of *CP_JGB3M*, the sample period is shortened from July 2003-April 2010 to July 2003-October 2009 when we incorporate *CP_JGB3M* in our model.

[27] Estimation results based on the fixed-effect are very close to the numbers in Table 3.

[28] For the test of model specification, the fixed-effect and random-effect models are estimated without the use of the cluster-robust standard errors.

TABLE 3

STATIC MODEL

	Model (1) RE		Model (2) RE		Model (3) FE		Model (3) RE		Model (3) MLE	
3-Yr SPREAD	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Std.
T_JGBGAP	1.0830	0.2073 ***	1.1110	0.1995 ***	0.9242	0.1596 ***	0.9153	0.1593 ***	0.9171	0.0460 ***
JGBSLOPE (10Y-2Y)	-0.1019	0.0326 ***	-0.1244	0.0357 ***	-0.0155	0.0503	-0.0104	0.0498	-0.0116	0.0284
JGB10Y	0.2179	0.0255 ***	0.2424	0.0279 ***	0.1010	0.0329 ***	0.0972	0.0340 ***	0.0983	0.0314 ***
NKY	-0.0277	0.0061 ***	-0.0289	0.0056 ***	-0.0030	0.0056	-0.0025	0.0059	-0.0027	0.0029
e_HV			0.0043	0.0010 ***	0.0022	0.0007 ***	0.0022	0.0007 ***	0.0022	0.0004 ***
RATE_RI			0.0929	0.0261 ***	0.0757	0.0329 **	0.0629	0.0140 ***	0.0655	0.0053 ***
GAP3_1DLAG					1.5840	0.2140 ***	1.6055	0.2151 ***	1.5978	0.0371 ***
_cons	0.3303	0.0768 ***	-0.1949	0.1905	-0.4480	0.2331 *	-0.3823	0.1445 ***	-0.3944	0.0655 ***
R-sq: # Obs	4173		4173		4172		4172		4172	
# Group	52		52		52		52		52	
within	0.2683		0.3109		0.5223		0.5219			
between	0.0384		0.6120		0.7518		0.7704			
overall	0.1802		0.4090		0.5997		0.6083			
					F-test that all u_i=0 F(51, 4113)=19.13 Prob>F=0.0000		Breusch and Pagan test for random effects chi2(1)=4631.59 for random effects Prob>chi2 = 0.0000		sigma_alpha=0.1244 (Std.=0.0131) sigma_e=0.2513 (Std.=0.0028) rho: AR(1) on e=0.1967 (Std.=0.0335)	

Note: ***:1%, **:5%, *:10%. Models (1) and (2) are estimated by a random-effect model with cluster-robust (firm-level) standard errors. Model (3) on FE (fixed-effect) and RE (random-effect) models are also estimated with cluster-robust standard error while the model specification tests are implemented by the models without the cluster-robust standard error. Models fitted on these data fail to meet the asymptotic assumptions of the Hausmann test.

TABLE 4
DYNAMIC MODEL

	Model (4) FE		Model (4) RE		Model (4) AH MLE	
3-Yr SPREAD	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Std.
T_JGBGAP	0.0410	0.0310	0.0280	0.0298	0.0365	0.0183 **
JGBSLOPE (10Y-2Y)	-0.0747	0.0146 ***	-0.0801	0.0143 ***	-0.0788	0.0106 ***
JGB10Y	0.0211	0.0104 **	0.0197	0.0097 **	0.0209	0.0116 *
NKY	-0.0023	0.0021	-0.0022	0.0019	-0.0023	0.0011 **
e_HV	0.0004	0.0002 *	0.0003	0.0002 *	0.0003	0.0001 ***
RATE_RI	-0.0049	0.0049	0.0030	0.0012 ***	0.0032	0.0009 ***
GAP3_1DLAG	0.4036	0.1330 ***	0.3852	0.1274 ***	0.3933	0.0151 ***
3-Yr SPREAD (Lagged)	0.8679	0.0348 ***	0.8865	0.0334 ***	0.8779	0.0054 ***
_cons	0.1117	0.0311 ***	0.0710	0.0344 **	0.0690	0.0232 ***
R-sq: # Obs	4116		4116		4116	
# Group	52		52		52	
within	0.9384		0.9382			
between	0.9973		0.9980			
overall	0.9566		0.9584			
	F test that all u_i=0 F(51, 4056)=2.27 Prob>F=0.0000		Breusch and Pagan test for random effects chi2(1)=11.29 for random effects Prob>chi2=0.0008		sigma_alpha=0.0102 (Std.=0.0023) sigma_e=0.0905 (Std.=0.0010) rho: AR(1) on e=0.0125 (Std.=0.0055)	

Note: ***:1%,**:5%,*:10%. Model (4) is estimated by a fixed-effect model (FE), a random-effect model (RE), and random-effect MLE (AH MLE). Models (4) (other than AH MLE) is estimated with cluster-robust (firm-level) standard errors. The model specification tests are implemented by the models without the cluster-robust standard error. Models fitted on these data fail to meet the asymptotic assumptions of the Hausmann test.

and Pagan test in Table 3). We cannot, however, determine which model is better since the models fitted to our data fail to meet asymptotic assumptions of the Hausman test. With the estimated results of the fixed-effect and the random-effect models as reasonably similar, we use both as our main models. In these two models, results of sensitivity analysis of each coefficient with respect to credit ratings are further demonstrated in Table 5.

The estimated coefficients in our static panel estimations are as follows. First, higher Tibor-JGB spreads (*T_JGBGAP*), which represent tightness in the money market, contribute to a higher median spread. This is consistent with our prediction that the spreads are positively correlated with

TABLE 5

STATIC AND DYNAMIC MODELS WITH SAMPLE SPLIT

	FE Model (3) Hi-Rate		FE Model (3) Low-Rate		RE Model (3) Hi-Rate		RE Model (3) Low-Rate		Model (4) AH MLE Hi-Rate		Model (4) AH MLE Low-Rate	
3-Yr SPREAD	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Std.	Coef.	Std.
T_JGBGAP	0.3652	0.0648 ***	1.4586	0.2961 ***	0.3993	0.0601 ***	1.4742	0.2886 ***	0.0000	0.0246	0.1036	0.0258 ***
JGBSLOPE (10Y-2Y)	-0.1437	0.0490 ***	0.0795	0.0810	-0.1298	0.0432 ***	0.0770	0.0788	-0.0733	0.0142 ***	-0.1048	0.0145 ***
JGB10Y	0.1153	0.0642 *	0.1075	0.0464 **	0.1106	0.0599 *	0.1083	0.0447 **	0.0041	0.0157	0.0476	0.0157 ***
NKY	-0.0015	0.0084	-0.0111	0.0057 *	0.0011	0.0091	-0.0103	0.0058 *	0.0013	0.0014	-0.0050	0.0014 ***
e_HV	0.0025	0.0008 ***	0.0011	0.0009	0.0021	0.0008 ***	0.0016	0.0010	0.0002	0.0002	0.0007	0.0002 ***
RATE_RI	0.1774	0.0237 ***	0.0718	0.0311 **	0.1043	0.0284 ***	0.0816	0.0225 ***	0.0051	0.0024 **	0.0082	0.0020 ***
GAP3_1DLAG	1.2941	0.3753 ***	1.4695	0.2328 ***	1.3760	0.3883 ***	1.4895	0.2318 ***	0.8399	0.0318 ***	0.2728	0.0164 ***
3-Yr SPREAD (Lagged)									0.7340	0.0136 ***	0.8985	0.0059 ***
_cons	-0.5357	0.1444 ***	-0.5546	0.2834 *	-0.3336	0.1947 *	-0.6427	0.2215 ***	0.0442	0.0312	0.0442	0.0347
R-sq: # Obs	2029		2143		2029		2143		2003		2113	
# Group	25		27		25		27		25		27	
within	0.6329		0.5677		0.6126		0.5674					
between	0.6590		0.7460		0.7022		0.7353					
overall	0.5425		0.5856		0.6110		0.5888					
									sigma_alpha=0.0122 (Std.=0.0033) sigma_e=0.0846 (Std.=0.0013) rho: AR(1) on e=0.0204 (Std.=0.0110)		sigma_alpha=0.0118 (Std.=0.0029) sigma_e=0.0876 (Std.=0.0014) rho: AR(1) on e=0.0177 (Std.=0.0087)	

Note: ***:1%, **:5%, *:10%. All the models (3) are estimated by either a fixed-effect model (FE) or a random-effect model (RE) with cluster-robust (firm-level) standard errors. Models (4) is estimated by random-effect MLE (AH MLE). The sample is Hi-Rate for the firms whose initial credit ratings at July 2003 are better than 5, and Low-Rate otherwise.

the market-level liquidity index. Second, steeper JGB slope (*JGBSLOPE*) is correlated with lower corporate bond spreads in models (1) and (2).[29] This implies that a steeper slope can be interpreted as the precursor of a better future market condition. Third, the higher yield of JGB (*JGB10Y*) with controlling the slope has positive impact on the corporate bond yield. This result has several interpretations. One perception is that the higher yield of government bonds is a sign of worse market conditions for corporate bonds. For example, a higher interest rate may negatively affect firms with large borrowing. In order to see how the market-level credit factor more explicitly affects the corporate bond spread, we subsequently use the spreads of commercial paper as an alternative factor. Fourth, the higher level of stock price index (*NKY*) is correlated with lower corporate bond spreads in models (1) and (2).[30] This result is consistent with the view that lower market-level credit risk substituted by the higher stock price is associated with low corporate bond spreads. Fifth, the abnormal individual stock price volatility (*eHV*) has a positive effect on corporate bond spreads. Similarly, the credit ratings of each company, as provided by R&I (*RATE_RI*) that covers the largest number of Japanese companies, have the same implication as *eHV*.

As the most important result, we establish a strong positive correlation between GAP and median spreads. We also find that the influence of GAP depends negatively on the credit ratings (*i.e.*, as credit ratings worsen, GAP coefficients multiplied by the standard deviation of GAP become larger).[31] Consequently, the impact is 0.09 (coefficient 1.2941 $\times$ standard error 0.07) in the fixed-effect estimation. Meanwhile, the credit rating at the beginning of the sample period is greater than or equal to 4 (*i.e.*, AA-) or is approximately double at 0.18 (coefficient 1.4695 $\times$ standard error 0.12).[32]

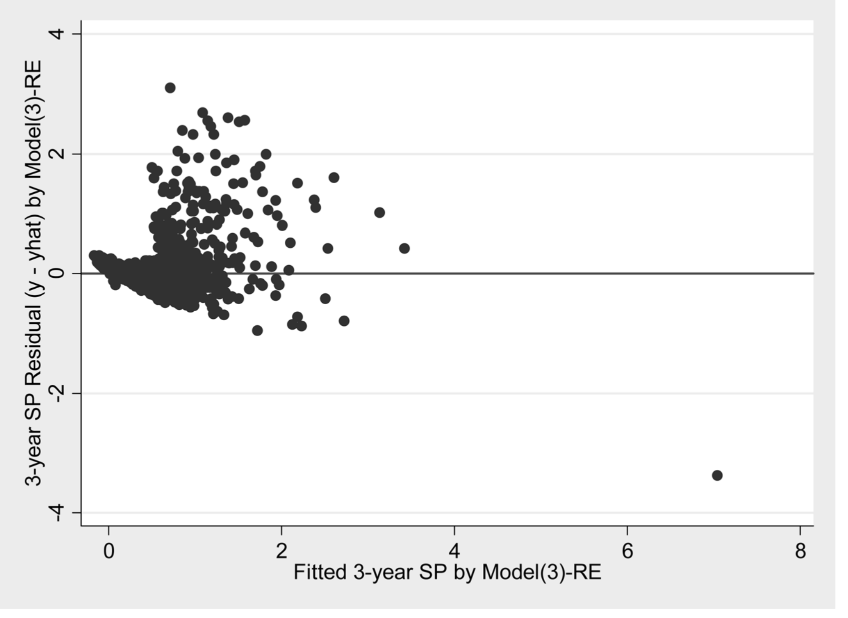
Although we successfully improved the model by adding the GAP measure in the static panel estimation, the residual plot of the random-effect model in Figure 5 shows obvious heteroskedasticity. The first feature of

[29] Although results from model (3) do not support this, we can confirm this in the results from the dynamic model presented in the succeeding section.

[30] Although the results from model (3) do not support this, we can confirm this in the results in many cases to be mentioned later.

[31] As we have explained above, we transform each rating into numbers. Detailed computation results are provided upon request.

[32] We obtain a similar implication from the estimation of the random-effect model. Detailed summary statistics for each sub-sample are provided upon request.



Note: Fitted three-year spreads are plotted on the horizontal axis while the residual is on the vertical axis. The upper (lower) panel is the result of fixed-effect (random-effect) estimation. TIME refers to 82 monthly periods, from July 2003 to April 2010.

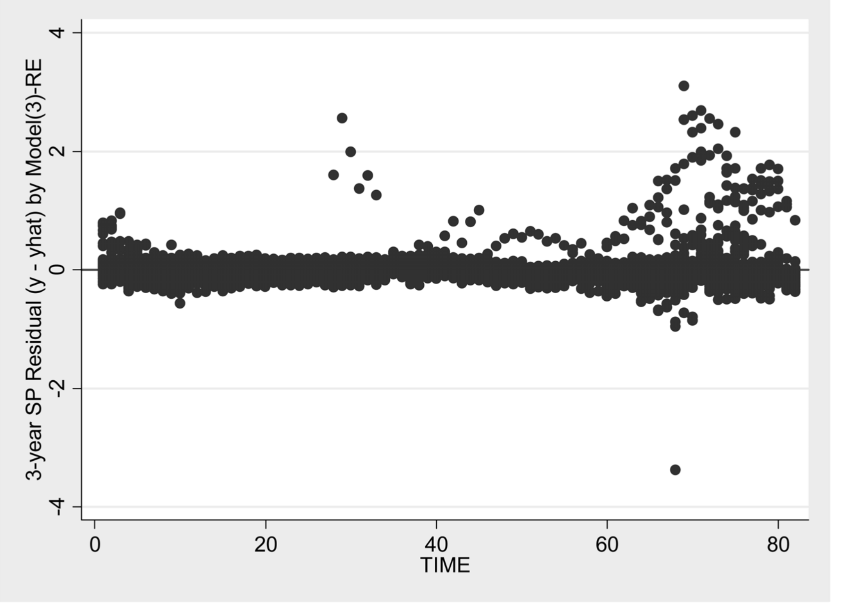
FIGURE 5(1)

RESIDUAL PLOT OF STATIC PANEL ESTIMATION (RANDOM-EFFECT)

the residual plot is the lower part of the scatter plot (*i.e.*, the spreads of firms with good credit and/or those in a good market environment period), which forms the down-sloped concentration. The second feature (*i.e.*, the spreads of firms with weak credit or those in a bad environment period) is in the upper scattered portion. With some potentially omitted variables that capture the behavior of the second feature, the static model generates a number of very large positive residuals in the higher spread range. These outliers attract the regression line upward and render the over-estimation of the first feature.

Under the presumption that the lagged dependent variable is one plausible omitted variable, the dynamic models proposed in (4) are estimated by fixed-effect and random-effect GLS and MLE.[33] The results summar-

[33] We also estimate Arellano-Bond GMM, and the results are very close to those from MLE, unless otherwise noted.

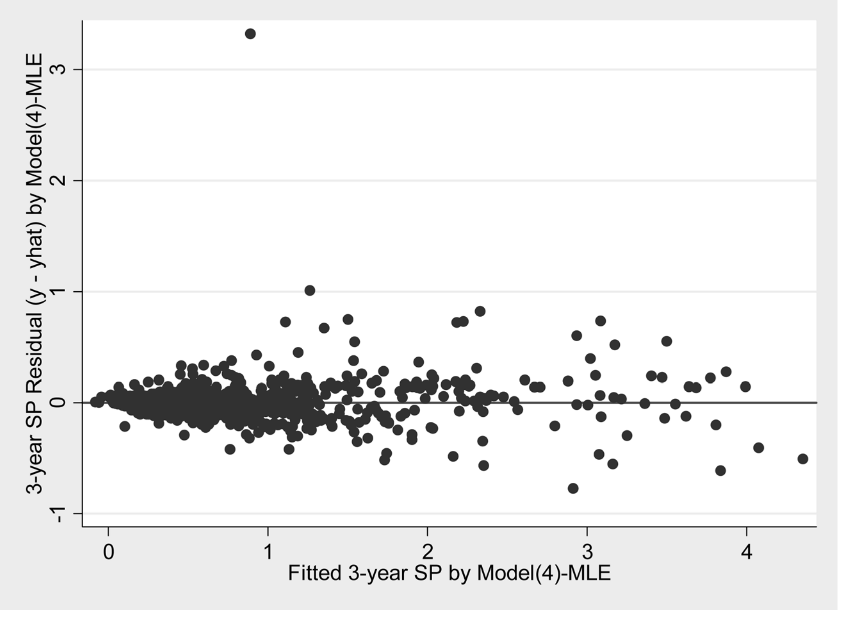


Note: The horizontal axis corresponds to TIME, while the residual is plotted on the vertical axis. The upper (lower) panel is the result of fixed-effect (random-effect) static panel estimation. TIME refers to 82 monthly periods, from July 2003 to April 2010.

FIGURE 5(2)
RESIDUAL PLOT OF STATIC PANEL ESTIMATION (RANDOM-EFFECT)

ized in Table 4 show that all estimation methods result in almost the same coefficients. Upon the examination of the scattered plots (Figure 6) corresponding to MLE estimations, the emerging problem from the omitted variables seems to be resolved satisfactorily. We can also confirm that the dynamic panel still provides qualitatively similar results, such as the coefficient of GAP, compared with those in the static model. The estimate of the AR(1) coefficient ranges from 0.86 to 0.89, which indicates a very strong auto-correlation. Hence, the conjectured persistency of the reported bond spreads is successfully confirmed through our dynamic panel estimation. Moreover, the results show that simultaneous inclusions of static and dynamic measures significantly improve the model fit.

Another important point is that the AR(1) coefficient also has apparent conditionality on credit ratings and/or the business cycle. Table 5 summarizes the spectrum of the AR(1) coefficient for different levels of credit



Note: Fitted three-year spreads are plotted on the horizontal axis while the residual is on the vertical axis. The panel shows the results based on MLE.

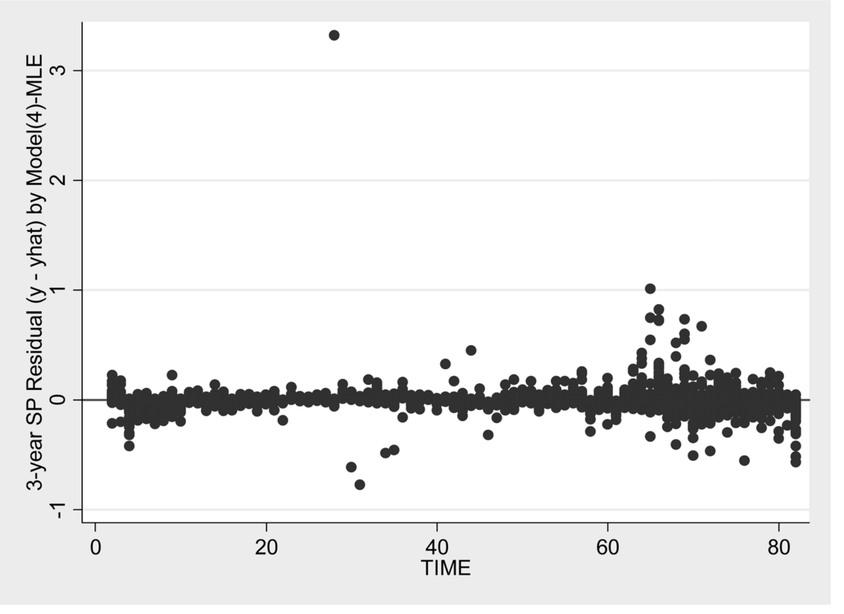
FIGURE 6(1)
RESIDUAL PLOT OF DYNAMIC PANEL ESTIMATION (MLE)

ratings. The purpose of this additional analysis is to compare AR(1) coefficients with respect to credit risk.[34] The negative conditionality of AR(1) coefficients on credit ratings is clear. The persistency of the median spread becomes stronger as credit ratings worsen, as established by Dick-Nielsen *et al.* (2012).

Furthermore, σ_α in MLE for the dynamic model is substantially smaller than that for the static model. This is consistent with the fact that estimated coefficients are similar among various estimation methods. If the individual characteristics are properly controlled, it is natural to have very small variations in the estimators of individual effects.

In order to check the robustness of the result, Table 6 reports the revised regression results in (3) and (4) under an adjusted GAP measure constructed by dividing GAP by the concurrent JGB yield. The results

[34] Recall that, in order to split the sample, we use (i) whether the credit rating at the beginning of the sample period is better than or equal to 4 (*i.e.*, AA-), or (ii) worse than or equal to 5 (A+).



Note: The horizontal axis corresponds to TIME while the residual is plotted on the vertical axis. The panel shows the results based on MLE. TIME refers to 82 monthly periods, from July 2003 to April 2010.

FIGURE 6(2)

RESIDUAL PLOT OF DYNAMIC PANEL ESTIMATION (MLE)

obtained in previous estimations are confirmed in this additional estimation.

Table 7 reveals the results of the static random-effect estimation and of the dynamic MLE with the split GAP measures, which consist of (i) the difference between the highest minus the median spreads and (ii) the difference between the median and the lowest spreads. The purpose of this exercise is to observe the evolution of reported spread distributions in correlation with the median spreads. As the table shows, both upward [*i.e.*, (i) mentioned above] and downward [*i.e.*, (ii) mentioned above] divergences have positive coefficients. This implies that the opinion difference in either direction has specific information correlated with the median evaluation.

As indicated by additional robustness checks, we replace some market-risk factors (*i.e.*, *T_JGBGAP*, *JGBSLOPE*, and *JGB10Y*) with the average commercial paper spread (*CP_JGB3M*).[35] As revealed in the first three

TABLE 6

ROBUSTNESS CHECK (ADJUSTED GAP: RELATIVE DISTANCE MEASURE)

	FE Model (3) (Adjusted GAP)		RE Model (3) (Adjusted GAP)		Model (4) AH MLE (Adjusted GAP)	
3-Yr SPREAD	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Std.
T_JGBGAP	1.0606	0.1591 ***	1.0495	0.1586 ***	0.0324	0.0194 *
JGBSLOPE (10Y-2Y)	-0.1851	0.0317 ***	-0.1802	0.0294 ***	-0.1194	0.0111 ***
JGB10Y	0.4116	0.0362 ***	0.4108	0.0361 ***	0.0744	0.0124 ***
NKY	-0.0112	0.0046 **	-0.0109	0.0049 **	-0.0053	0.0011 ***
e_HV	0.0028	0.0007 ***	0.0028	0.0007 ***	0.0005	0.0001 ***
RATE_RI	0.0864	0.0393 **	0.0678	0.0178 ***	0.0028	0.0009 ***
GAP3_1DLAG_Adj	0.5558	0.0659 ***	0.5626	0.0664 ***	0.0937	0.0061 ***
3-Yr SPREAD (Lagged)					0.9058	0.0058 ***
_cons	-0.7013	0.2901 **	-0.6050	0.1883 ***	0.0766	0.0246 ***
R-sq: # Obs	4172		4172		4116	
# Group	52		52		52	
within	0.4898		0.4891			
between	0.7078		0.7291			
overall	0.5569		0.5715			
					sigma_alpha=0.0088 (Std.=0.0025) sigma_e=0.0952 (Std.=0.0011) rho: AR(1) on e=0.0085 (Std.=0.0048)	

Note: ***:1%,**:5%,*:10%. Models (3) (FE and RE) is estimated with cluster-robust (firm-level) standard errors.

columns of Table 8, we confirm that *CP_JGB3M* works consistently as a market-risk factor, while GAP and the persistency measure are robust against this alternative variable choice.[36] The results also imply that the lower short-term rate (*ON_AVERAGE*) is correlated with wider corporate bond spreads. This reflects the employment of monetary easing policy in the case of higher macro credit risk. Corporate bond spreads also become wider when a larger risk is perceived in the foreign financial market (*i.e.*, higher *DLIBOR3M*).[37] Although foreign investors are not ne-

[35] Due to high correlation among variables, we are not including some of the variables in our baseline estimation.

[36] We can infer that our result associated with GAP and the persistency term is not suffering from the potential omitted variable bias since various alterations of explanatory variables do not affect results.

TABLE 7
ROBUSTNESS CHECK
(SPLIT GAP: HI-MEDIAN SPREAD AND LOW-MEDIAN SPREAD)

	Model (3) FE GAP SPLIT		Model (3) RE GAP SPLIT		Model (4) AH MLE GAP SPLIT	
3-Yr SPREAD	Coef.	Cluster-Robust Std.	Coef.	Cluster-Robust Std.	Coef.	Std.
T_JGBGAP	0.9367	0.1563 ***	0.9294	0.1566 ***	0.0437	0.0183 **
JGBSLOPE (10Y-2Y)	0.0286	0.0533	0.0333	0.0521	-0.0733	0.0107 ***
JGB10Y	0.0748	0.0274 ***	0.0711	0.0281 **	0.0187	0.0116
NKY	0.0008	0.0051	0.0012	0.0054	-0.0018	0.0011 *
e_HV	0.0021	0.0007 ***	0.0021	0.0007 ***	0.0004	0.0001 ***
RATE_RI	0.0697	0.0327 **	0.0592	0.0138 ***	0.0031	0.0009 ***
GAP3_1DLAG_UP	0.8943	0.2514 ***	0.9072	0.2573 ***	0.3228	0.0221 ***
GAP3_1DLAG_DOWN	2.6067	0.2825 ***	2.6368	0.2728 ***	0.5149	0.0317 ***
3-Yr SPREAD (Lagged)					0.8727	0.0055 ***
_cons	-0.4794	0.2317 **	-0.4261	0.1449 ***	0.0606	0.2324 ***
R-sq: # Obs	4172		4172		4116	
# Group	52		52		52	
within	0.5458		0.5455			
between	0.7645		0.7807			
overall	0.6207		0.6273			
					sigma_alpha=0.0107 (Std.=0.0023) sigma_e=0.0902 (Std.=0.0010) rho: AR(1) on e=0.0138 (Std.=0.0058)	

Note: ***:1%,**:5%,*:10%. All the models (3) are estimated by either a fixed-effect model (FE) or a random-effect model (RE) with cluster-robust (firm-level) standard errors. Models (4) is estimated by random-effect MLE (AH MLE). Models (3) is estimated with cluster-robust standard error (firm-level). GAP3_1DLAG_UP and GAP3_1DLAG_DOWN denote the differences between (i) the highest reported spread minus median reported spread and (ii) the median reported spread minus the lowest reported spread, respectively. Both variables are one-day lagged.

cessarily major players in Japanese corporate bond markets, this finding implies the potential connection between the condition of foreign financial markets and that of the Japanese corporate bond spreads.[38]

[37] Due to the high correlation between NKY and DLIBOR3M, we exclude NKY when we incorporate DLIBOR3M in our estimation.

[38] According to the Bank of Japan, the outstanding amount of corporate bonds held by foreign investors is only 2.1 trillion yen out of 76.7 trillion yen of the total outstanding amount as of the end of September 2011.

TABLE 8

ROBUSTNESS CHECK (ALTERNATIVE MACRO FACTOR, TIME-EFFECT)

	Model (4) AH MLE (Alternative-1)		Model (4) AH MLE (Alternative-2)		Model (4) AH MLE (Alternative-3)		Model (6) AH MLE (w/ Monthly-dummy)		Model (6) AH MLE (w/ Year-dummy)	
3-Yr SPREAD	Coef.	Std.	Coef.	Std.	Coef.	Std.	Coef.	Std.	Coef.	Std.
T_JGBGAP							-0.7920	1.3141	0.2604	0.0321 ***
JGBSLOPE (10Y-2Y)							4.4264	0.1563 ***	-0.0669	0.0190 ***
JGB10Y							-2.9633	0.4769 ***	0.0367	0.0154 **
CP_JGB3M	0.1441	0.0074 ***	0.1525	0.0086 ***	0.1399	0.0076 ***				
NKY	0.0032	0.0005 ***	0.0040	0.0007 ***			-0.0949	0.0607	-0.0058	0.0015 ***
ON_AVERAGE			-0.0170	0.0090 *	-0.0165	0.0081 **				
DLIBOR3M					0.0081	0.0010 ***				
e_HV	0.0002	0.0001 *	0.0002	0.0001 *	0.0002	0.0001 *	0.0002	0.0001	0.0003	0.0001 ***
RATE_RI	0.0029	0.0008 ***	0.0027	0.0008 ***	0.0024	0.0008 ***	0.0018	0.0008 **	0.0018	0.0008 **
GAP3_1DLAG	0.4018	0.0152 ***	0.4010	0.0152 ***	0.4050	0.0152 ***	0.3874	0.0144 ***	0.3938	0.0148 ***
3-Yr SPREAD (Lagged)	0.8814	0.0053 ***	0.8828	0.0053 ***	0.8871	0.0054 ***	0.8973	0.0052 ***	0.8939	0.0055 ***
_cons	-0.0674	0.0091 ***	-0.0753	0.0100 ***	-0.0467	0.0058 ***	0.0203	0.4179	-0.0336	0.0286
R-sq: # Obs	3823		3823		3823		4116		4116	
# Group	52		52		52		52		52	
	sigma_alpha=0.0096 (Std.=0.0023) sigma_e=0.0880 (Std.=0.0010) rho: AR(1) on e=0.0117 (Std.=0.0056)		sigma_alpha=0.0094 (Std.=0.0023) sigma_e=0.0879 (Std.=0.0010) rho: AR(1) on e=0.0113 (Std.=0.0055)		sigma_alpha=0.0087 (Std.=0.0024) sigma_e=0.0876 (Std.=0.0010) rho: AR(1) on e=0.0098 (Std.=0.0053)		sigma_alpha=0.0080 (Std.=0.0022) sigma_e=0.0827 (Std.=0.0009) rho: AR(1) on e=0.0093 (Std.=0.0050)		sigma_alpha=0.0080 (Std.=0.0023) sigma_e=0.0880 (Std.=0.0010) rho: AR(1) on e=0.0081 (Std.=0.0047)	

Note: ***:1%, **:5%, *:10%. All the models are estimated by random-effect MLE (AH MLE). Model (6) is estimated with considering either monthly- or year-level unobservable time-effect.

$$SP_{it}=\gamma SP_{it-1}+\beta_1+\beta_2 F_t+\beta_3 L_t+\beta_4 f_{it}+\beta_5 l_{it}+\alpha_i+\eta_t+\varepsilon_{it} \tag{6}$$

Finally, we consider the unobservable time-effect η_t by including either monthly or annual time dummies into our model in our baseline estimation, as shown in (6). The last two columns in Table 8 display the model that has monthly or yearly frequency time-dummy variables, based on the model by Baltagi (2008). All results associated with GAP and the persistency term remain same. This further confirms the robustness of our empirical results.

VII. Concluding Remarks

This paper studies the price impact of corporate bond illiquidity. We employ the dynamic panel estimation to simultaneously examine the static price dispersion measured by GAP and the dynamic resiliency factor substituted by the lagged dependent variable. Unlike extant studies that treat these as separate objects, we study both price dispersion and resiliency in a unified framework and observe that the inclusion of both measures significantly improves the model fit. Their price impacts are also confirmed to systematically respond to credit ratings of bonds. These results imply that the incorporation of multiple measures of illiquidity improves the precision of corporate bond pricing.

To conclude, we list several future research questions. First, our estimated results can be used to motivate theoretical models. A model with a market maker that continuously revises its evaluation about bond spreads while still exhibiting heterogeneous valuations can be created.[39] This behavior stems, for example, from heterogeneous funding availabilities of clients under limited arbitrage opportunity. Second, the study on the determination of GAP can serve as informative research objects. Developing well-functioning corporate bond markets has been one of the most important policy issues, for example, in Asian financial markets (Kang 2007). We believe the results obtained in this paper and those from the extension proposed above provide further understanding of corporate bond pricing, which can certainly contribute to the development of corporate bond markets.

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[39] Feldhütter (2012) partly considers this environment.

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