Diversifying Risks in Bond Portfolios: A Cross-border Approach

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ABSTRACT

This study recalibrates corporate bond idiosyncratic risks in an international context. Applying a statistically powerful risk decomposition scheme, we show in this study that diversification is improved by the addition of a global risk benchmark. We build a long-run stationary yield spread decomposition scheme which provides better diversification effect. Both the global and domestic risk benchmarks are observable yield spreads and are free of measurement or availability issues. The idiosyncratic risk component is estimated as a fixed effect along with all the parameter estimates, rather than separately from an exogenous generating process. Our linear model is simple, yet it can be easily and promptly applied by practitioners.

Keywords: bond pricing; credit spread; systematic risk; diversification; global risk; heterogeneous panel; pooled mean group.

JEL Classification: C32, E4, E21, G13, G3

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1. Introduction

In the recent European sovereign debt crisis, corporate yield spreads are unusually high in some countries or regions, but not necessarily so elsewhere. It implies then opportunities for further cross-border diversification for fixed income portfolios. According to Bank of America, out of the overall 12 trillion dollars corporate bond market, 20% is held by ETF and mutual funds. Bond mutual funds invested $1.44 trillion in corporate bonds, up by almost $380 billion since the year of 2000. Although systematic and idiosyncratic risks of corporate bonds have been examined in many studies, global factors are rarely considered, possibly due to measurement difficulties. However, the increasingly stronger influence of global risks on corporate bond yields makes it more important than ever to identify idiosyncratic components in yield spreads so they that can be diversified away adequately in cross-border portfolios.

Yield spreads of fixed income instruments over riskless benchmarks are crucial in pricing and trading. The composition and forming process of yield spreads therefore determine if spreads are adequately assessed and practically applicable for practitioners to revise timely. Given that default risks, political or business cycle risks, as well as liquidity risks have been considered as three major corporate bond risk components in literature (see, among others, Dastidar & Phelps, 2011; Xie, Shi & Wu, 2008; Longstaff, Mithal & Neis, 2005; Chen, Lesmond & Wei, 2007; Block & Vaaler, 2004), signals used to proxy these components are sometimes difficult to observe or measure with precision, making it impractical to utilize those signals directly. In the work of Dastidar and Phelps (2011), liquidity risk is idiosyncratic. However, as international capital markets integrate, domestic economy cannot be the only major source of systematic risk. Dungey, Martin and Pagan (2000), whose factor model for government bonds incorporates both world as well as country risks, indicate that risks related to global economic or political developments need to be properly accounted for, beyond what has already been incorporated in sovereign yields. More importantly, yields or yield spreads for corporate bonds of different credit rating should have reflected the same systematic risks, which can be extracted from traded yields rather from other exogenous sources.

Decomposition is crucial in studying various risks reflected in yield spreads. Lin and Curtillet (2007) indicate that it is inappropriate to just analyze full credit spreads. Lerner and Wu (2005) also suggest that full spreads could be under- or over-estimated under different credit ratings. Wilson (1998) starts the research on credit spreads decomposition by studying systematic and idiosyncratic risks in the loss distribution. Duffee (1999) adopts a reduced-form model to decompose credit spreads, while Gatfaoui (2003) uses a structural model instead. Jarrow, Lando and Yu (2005) assume a perspective of investment portfolio and discuss how idiosyncratic risk can diversify risks in the portfolio. Churm and Panigirtzoglou (2007) incorporate the choice of default point in the calculation of spread decomposition[1] as an extension of Liu, Longstaff and Mandell (2006), where swap spreads are adopted as an estimation basis for idiosyncratic credit spreads. On the method of decomposition, this project will extend the two-factor spread decomposition scheme proposed in Sun, Lin and Nieh (2008) to a three-factor model as a fundamental structure of cross-border analysis.

This study proposes a model for decomposing corporate yield spreads in a context of international risk sharing where global, as well as domestic, systematic risks and idiosyncratic risks are considered at the same time. Our model employs indirect risk measurements rather than the commonly used direct default and liquidity risk measures. Specifically, the yield spread of a top-grade US corporate bond, which carries little credit and liquidity risks, is used as a proxy for global systematic risk for each single country covered in the study. While the yield spreads of domestic high-grade corporate bond in excess of the top-grade US corporate yield spreads, would serve as the domestic systematic risk proxy. Systematic credit and liquidity risks, which have been incorporated in these domestic high-grade corporate yield spreads, thus influence other individual corporate yield spreads indirectly. Indirect risk components are better than direct ones as they are produced by the same capital markets that price other individual corporate yields, so observing and measuring them is not a issue. Top-grade corporate bonds have better market liquidity than other issues hence they are also more stable and precise. With the aid of a statistically powerful risk decomposition scheme, we show in this study that diversification is improved after adding in an global risk benchmark.

In terms of the econometric treatment on yield data, *changes* had been used (.e.g., Wilson, 1998; Duffee, 1999; Dastidar & Phelps, 2011; Lee, Xie & Yau, 2011) to avoid partially problems arising from non-stationarity and autocorrelation in the *level* of credit spreads. But it is accompanied by fundamental drawbacks such as the loss of information, and being leptokurtic as indicated by Pedrosa and Roll (1998). Changes of yield spreads are also found to persist over time in Duffee (1998). Extending the credit spread decomposition model of Sun, Lin and Nieh (2008) and panel decomposition model of Lin and Sun (2007), we explore in this study a model containing both global and domestic systematic bond risks, in addition to idiosyncratic default and liquidity risks. The AutoRegressive Distributed Lag (ARDL) panel time series model of Pesaran, Shin and Smith (1999),

[1] This perspective is similar to Lin and Sun (2009), which is based on the model of Merton 1974) and analyzes nonlinear price changes of debt claims in the neighborhood of default point, whose direct contribution is to account for the differences in idiosyncratic credit spreads between investment and high-yield corporate bonds.

which emphasizes the long-run relations among economic variables, helps us building a long-run stationary yield spread decomposition scheme in the study.

We find from our analysis that, for all the countries, both the global and domestic systematic components are significant in constituting yield spreads of individual issues in each country. The inclusion of a global risk component provides more abundant and explicit information, which the traditional domestic model lacks, for the pricing and risk management practices of fixed income portfolios. The contributions of both the global and the domestic risk benchmarks are estimated with a statistically more powerful time series model in an econometrically long-run context. The idiosyncratic risk component is estimated as a fixed effect in our data panel along with all other parameter estimates, rather than being introduced separately from an exogenous generating process. As a result, parameter estimates from our yield decomposition model can be used to construct yield spreads directly, simply by employing market data. Our linear decomposition model may contain other econometric imperfections, but our estimates can be applied promptly and easily by practitioners.

Yield spread panels are often studied in regressions with fixed or random effects, in which homogeneity of parameters is imposed across all the group time series. While the long-run relationship can be predicted by economic theory, both the short-run dynamics and particularly the speed of adjustment to equilibrium mainly depend on group-specific factors. This study employs a panel estimation approach which allows heterogeneous short-run dynamics and how they revert to long-run equilibrium. Yet the approach constrains long-run equilibrium to be homogeneous across groups of corporate yield spreads. This modification of traditional methods proves to be consequential. For each country, the portfolio Value at Risk (VaR) measure on idiosyncratic risk falls significantly, which implies better cross-border diversification.

There are a growing number of studies using CDS spread, as they become more available, to measure systematic default risks (e.g., Blanco, Brennan & Marsh, 2005; Longstaff *et al.*, 2005). However, most of the CDS reference entities are financial institutions so it does not help identifying idiosyncratic risk component within corporate capital cost in general. Although prices of CDS are market-based measures, counterparty risk introduces significant measurement biases when financial market is distressed. The apparent higher volatility of CDS prices than corporate yields also make them less ideal as subjects for the study of credit risk decomposition.

Our results help enhancing the performance of global fixed income portfolio diversification as we extend a domestic framework to a cross-border one. Secondly, the analysis of risk factors in international investment portfolio adds insights to the practice of pricing and risk management of international asset management, especially in effective cross-border and cross-segment management. A theoretical model for decomposition is introduced in Section 2. Section 3 gives an empirical decomposition scheme to fit our international bond data. Findings of empirical analysis are given in Section 4, followed by concluding remarks in Section 5.

2. Risk decomposition in a global context

To characterize systematic and idiosyncratic risks driving corporate yield spreads, we use a framework adapted from Duffie and Singleton (1999) and Liu, Longstaff and Mandell (2006). As our interest in this study is in the spreads among global top-grade, domestic high-grade and lower-grade corporate bonds, we include in our model only one default-free bond, two higher-grade and one lower-grade default able bond. First, assume there is a globally default-free zero-coupon bond maturing at T has at time t a value of

$$D(t,T) = E_Q\left[\exp\left(-\int_t^T r_s ds\right)\right], \tag{1}$$

where r_s is the short rate and E_Q is the expectation with respect to measure Q, the risk-neutral counterpart of the physical or objective measure P.

The value of a global top-grade defaultable bond will incorporate in addition a default intensity spread λ_s which is from a Poisson process with time varying parameter. At time t it can be expressed as

$$\boldsymbol{A_G(t,T) = E_Q}\left[\exp\left(-\int_t^T [\boldsymbol{r_s} + \lambda_s]\boldsymbol{ds}\right)\right]. \tag{2}$$

A domestic high-grade defaultable bond is also considered, which includes a liquid spread γ_s to compensate for the illiquidity compared with default-free bonds[2] and has the value of

$$\boldsymbol{A_D(t,T) = E_Q}\left[\exp\left(-\int_t^T [\boldsymbol{r_s} + \lambda_s + \gamma_s]\boldsymbol{ds}\right)\right], \tag{3}$$

which is identical to the expression in Liu *et al*. (2006), but the liquidity spread is imposed on a defaultable bond rather. For the purpose of differentiating the global systematic risk from the

[2] As our focus in the study is on the yield spreads of corporate issues, the modeling here is essentially a mix of the illiquid default-free bond and a defaultable bond as presented in Liu *et al*. (2006).

domestic systematic risks, we choose not to include liquidity spread in (2). (3) therefore in essence relates Liu *et al*. (2006) to Jarrow *et al*. (2003) with a reasonable theoretic support.

A lower-grade defaultable bond with similar structure is then modeled to have the value of

$$\boldsymbol{B}(\boldsymbol{t},\boldsymbol{T}) = \boldsymbol{E}_Q\left[\exp\left(-\int_t^T [\phi_1 \boldsymbol{r}_s + \phi_2\lambda_s + \phi_3\gamma_s]\boldsymbol{ds}\right)\right], \tag{4}$$

at time t. The three coefficients, ϕ_1, ϕ_2 and ϕ_3, are all positive and modeled in to reflect different sensitivity to the short rate, possible larger liquidity and default spreads. ϕ_1 could be considered as reflecting the agency effect argued by Leland and Toft (1996), should be greater than 1. So $\boldsymbol{B}(\boldsymbol{t},\boldsymbol{T})$ or its yield is expected to be more responsive, than $\boldsymbol{A_D}(\boldsymbol{t},\boldsymbol{T})$ or its yield, to the short rate. Similarly, ϕ_2 and ϕ_3 should both be greater than 1 as well, reflecting the fact that more risky bonds are more sensitive to changes in default intensity and market liquidity.

The dynamics of the three endogenous variables are characterized by a general affine model with four state variables which are Markovian under the equivalent martingale measure *Q* and square-root diffusions. The short rate is assumed to be driven by two state variables[3] to represent common shocks to the economy,

$$r_s = \delta_0 + X_1 + X_2, \tag{5}$$

where δ_0 is a constant. The liquidity spread in the domestic high grade defaultable bond is assumed to take the form of

$$\gamma_s = \delta_1 + X_3, \tag{6}$$

where δ_1 is also a constant and the state variable X_3 represents the premium required for the illiquid corporate issues, regardless of default risks. The default intensity is assumed such that

$$\lambda_s = \delta_2 + \tau r_s + X_4, \tag{7}$$

where δ_2 and τ are both constants and the latter stands for the sensitivity of default to the short rate. Structural models would predict τ to be negative.

The state variable vector $\boldsymbol{X} = (X_1, X_2, X_3, X_4)$, with general Gaussian processes under an affine

[3] The interpretation of factors X_1 and X_2, which come from the affine model of Duffie and Singleton (1997), can be found in Longstaff and Schwartz (1992) and Duffee (2002). In a continuous time context, the first factor is related to a long term mean of instantaneous rate while the second one to the instantaneous variance.

term-structure model, should be characterized by

$$dX = -\beta X dt + \Sigma dB^Q \, , \tag{8}$$

where β is a diagonal matrix and B^Q is a vector of independent standard Brownian motions under the risk-neutral measure of Q. Σ is a lower diagonal matrix containing covariances among the state variables, and it is assumed also that the covariance matrix $\Sigma\Sigma'$ is of full rank to allow correlations of state variables. Corresponding to this affine structure is the dynamics under the physical measure P,

$$dX = \kappa(\xi - X)dt + \Sigma dB^P \, , \tag{9}$$

where κ is also a diagonal matrix and ξ is a vector of long-term value of the state variables. The solutions to (1) through (4) can be solved under the risk-neutral dynamics (8). Generalizing the characterizations of (4) to bonds of other grades, we could consider X_1 and X_2 as globally common risks as their effects are proportional across all bonds.

On the physical measure P, the yield difference between the global top-grade bond, $A_G(t,T)$, and riskless bond, or the global systematic spread, can be seen from (8) and (9) as governed by

$$\boldsymbol{r_t} + \boldsymbol{\varpi r_t} + \eta_1^a(\boldsymbol{t})(\beta - \kappa)\boldsymbol{X_t} + \eta_1^b(\boldsymbol{t})\kappa\xi \, , \tag{10}$$

where $\eta_1^a(\boldsymbol{t})$ and $\eta_1^b(\boldsymbol{t})$ are functions of parameters. The first term in (10) is an instantaneous spread compensating for holding a risky bond which is less liquid than a riskless bond. The second term is also a short-run spread covering default related risk at current state, which is indirectly related to the interest rate. The third term is a long-run premium compensating for possible future default and liquidity related price changes. The last term is related to the risk-adjusted long-run level of bond yield spread.

Similarly on measure P, the yield spreads for the domestic high-grade bond or the domestic systematic spread should be

$$r_t + \varpi r_t + \eta_2^a(t)(\beta - \kappa)X_t + \eta_2^b(t)\kappa\xi \, , \tag{11}$$

which amounts to about 73 b.p. for a AAA 10-year bond based on US data according to Liu, *et al.* (2006). Similarly, the yield difference between the lower-grade and riskless bond can be derived as

$$\phi_1 r_t + \phi_2 \varpi r_t + \phi_3 \gamma_s + \eta_3^a(t)(\beta - \kappa)X_t + \eta_3^b(t)\kappa\xi \, , \tag{12}$$

The yield spread of the corporate bonds containing idiosyncratic risks should exhibit in the long run stronger responses to X_1 and X_2 contained in the interest rate due to agency risk. It should be more sensitive to interest rate-induced default risk in the short run. Both have been well documented by Sun, Lin and Nieh (2007) using US corporate indices.

3. Three-factor Credit Spread Decomposition

Instead of using the usually seen change-based short-run model, we decompose yield spread with a level-based long-run model which has better implications for cross-border diversification. Duffee（1998）and Xie, et al. (2008) both examine a three-factor reduced form model for corporate yield spreads, but the focus is on the idiosyncratic rather on the systematic risks. Xie, et al. (2008) indicate that findings of Duffee（1998）omit certain common factors in a firm's default risk, while arguing that macroeconomic variables, in addition to term structure and default intensity, affect corporate yield spreads. Our focus in this study lies instead on systematic risks to capture the unexplained variations in yield spreads. In addition to distinguishing systematic from idiosyncratic yield spreads proposed in literature, we further decompose systematic spreads into global and domestic components, so that global systematic risks can be flexibly priced in respective fixed income instruments. Under a framework of international market, long-run and common risk factors should be more important, particularly the part that cannot be effectively extracted under a single-country model.

Based on the specification of (2)~(4) in the previous section, the yield spread of a corporate bond issued in a particular country can be modeled to reflect the influence of short rate, liquidity risk within the country, as well as the credit risk of the issuer. In practice, there should be a benchmark country where a top-grade corporate bond produces only liquidity risk and almost no credit risk, like A_G in (2). The short rate in this country would be considered as a riskless rate benchmark, and the yield spread implied in A_G in (2) as a global systematic risk benchmark. To the extent that the difference between local short rate and benchmark short rate is only a constant, after combining (10)~(12), we could consider the yield spread of the bond issued by firm j at time t in a non-benchmark country i, $\boldsymbol{SP}_{it}^{j}$, in a linear model like

$$\boldsymbol{SP}_{it}^{j} = \gamma_{0i}^{j} + \gamma_{1i}^{j}\boldsymbol{SP}_{t}^{G} + \gamma_{2i}^{j}\boldsymbol{SP}_{it}^{D} + \gamma_{3i}^{j}\boldsymbol{RP}_{it} + \gamma_{4i}^{j}\boldsymbol{TS}_{it} + \xi_{it}^{j},\ j=1,2,\ldots,M,\ i=1,2,\ldots,N,\ t=1,2,\ldots,T \qquad (13)$$

where $\boldsymbol{SP}_{t}^{G}$ is the global systematic risk benchmark and $\boldsymbol{SP}_{it}^{D}$ is the local systematic risk benchmark in the local country, as implied by A_D in (3). γ_{0i}^{j} is considered as the idiosyncratic spread and assumed, without loss of generality, to be invariable in time. Under the specification above, ξ_{it}^{j} would be a disturbance. $\boldsymbol{RP}_{it}$ is the domestic repo rate and $\boldsymbol{TS}_{it}$ would be the difference between domestic long and short rates, together serving as control factors beyond the three risk factors, the global systematic, the domestic systematic and the idiosyncratic factor.

A commonly used model for (13) is a pooling panel OLS regression on changes of $\boldsymbol{SP}_j$ with fixed or random effects (Duffee, 1998; Jacoby, Liao & Batten, 2009), but that would require coefficients of all the regressors to be the same across all firms. Besides, yield spreads and term structure parameters are autocorrelated. Disturbances in (13) maybe nonstationary as Morris, Neal, and Rolph (2000) has found. Taking simple changes of $\boldsymbol{SP}_j$ not only discards valuable information but also helps little due to possible higher order autocorrelations. So we would employ an ARDL version of (13) according to Pesaran and Smith (1995) in the following form, for a given country,

$$\boldsymbol{SP}_{jt} = \sum_{k=1}^{p} \lambda_{jk} \boldsymbol{SP}_{j,t-k} + \sum_{k=1}^{q} \boldsymbol{\delta}_{jk}^{'} \boldsymbol{X}_{j,t-k} + \mu_j + \varepsilon_{jt}, \; j=1,2,\ldots,M,\; t=1,2,\ldots,T, \tag{14}$$

where j denotes a certain firm, $\boldsymbol{X}_{jt}$=($\boldsymbol{SP}_t^{G}, \boldsymbol{SP}_t^{D}, \boldsymbol{RP}_t, \boldsymbol{TS}_t$)' , $\boldsymbol{\delta}_{jk}$=($\boldsymbol{\delta}_{jk}^{1}, \boldsymbol{\delta}_{jk}^{2}, \boldsymbol{\delta}_{jk}^{3}, \boldsymbol{\delta}_{jk}^{4}$)' , and ε_{jt} is the disturbance independently distributed across j and t with mean 0 and $\sigma_j^2 > 0$. μ_j is assumed to be the fixed effect for firm j in the panel ARDL model of (14), and can be considered as reflecting the idiosyncratic risk in this firm's corporate bond yield in the sense of decomposition argued by Sun, et all (2007) as well as Dastidar and Phelps (2011).

If the variables in (14) are processes of I(1) and cointegrated[4], then the error term should be of I(0) for all j. So we can reparameterize (14) into an error correction form like

$$\Delta \boldsymbol{SP}_{jt} = \phi_j (\boldsymbol{SP}_{j,t-1} - \boldsymbol{\theta}_j^{'} \boldsymbol{X}_{jt}) + \sum_{k=1}^{p-1} \lambda_{jk}^{*} \Delta \boldsymbol{SP}_{j,t-k} + \sum_{k=1}^{q-1} \boldsymbol{\delta}_{jk}^{'*} \Delta \boldsymbol{X}_{j,t-k} + \mu_j + \varepsilon_{jt}, \tag{15}$$

where ϕ_j=-(1- $\sum_{k=1}^{p} \lambda_{jk}$), $\boldsymbol{\theta}_j$= $\sum_{k=0}^{q} \delta_{jk}$ /(1- $\sum_{k=1}^{p} \lambda_{jk}$), $\lambda_{jk}^{*} = -\sum_{m=k+1}^{p} \lambda_{jm}$ and $\boldsymbol{\delta}_{jk}^{*} = -\sum_{m=k+1}^{q} \delta_{jm}$ according to Pesaran, et al. (1999). ϕ_j is the speed of error-correction on the process' deviation from its long term equilibrium, which is the expression $\boldsymbol{SP}_{j,t-1} - \boldsymbol{\theta}_j^{'} \boldsymbol{X}_{jt}$ in (15). If $\boldsymbol{SP}_j$ and ($\boldsymbol{SP}_t^{G}, \boldsymbol{SP}_t^{D}, \boldsymbol{RP}_t, \boldsymbol{TS}_t$) are

[4] Neal, Rolph, Dupoyet and Jiang (2012), among others, have argued that levels of the intermediate and long-term corporate as well as government bond yields are nonstationary while their changes appear to be stationary. Before that, Mehra (1994) and Campbell and Shiller (1987) have found similar results for long-term nominal interest rates.

cointegrated then ϕ_j should be significantly negative in order for for ε_t to revert to 0. The vector $\boldsymbol{\theta}_j$ characterizes the long run relation between $\boldsymbol{SP}_j$ and ($\boldsymbol{SP}_t^G$, $\boldsymbol{SP}_t^D$, $\boldsymbol{RP}_t$, $\boldsymbol{TS}_t$). Short run effects are reflected by ϕ_j, λ_{jk}^* and the vector $\boldsymbol{\delta}_{jk}^*$. The ARDL model retains the *level* terms of $\boldsymbol{SP}_{j,t-1}$ and $\boldsymbol{X}_{jt}$, and is therefore superior to models employing only *changes* of yield spreads and explanatory variables.

Pesaran and Smith (1995) show that a panel model like (15) can be estimated separately for each firm (j=1,2,…,M) first and then make inferences on the averages of coefficients from individual ARDL equations and standard errors of these averages. This approach, or the Mean Group (MG) estimation, is superior to a pooling panel model which has distinct fixed effects for each firm but common slope coefficients across all firms. The latter does not distinguish short-run effects from long-run ones, and also produces inconsistent results for a dynamic heterogeneous panel. The MG estimation is the first ARDL method used in this study for decomposing corporate yield spreads within a given country.

The second decomposition method is a Pooled Mean Group (PMG) model according to Pesaran, et al. (1999), which allows the intercept, short-run coefficients, and error variances to differ across groups, similar to the MG estimation method. The long-run coefficients under PMG are, however, constrained to be equal across groups like in a pooling model with fixed effects. So the second method requires the assumption of $\boldsymbol{\theta}_j=\boldsymbol{\theta}$, $\forall j$. To compare against the PMG method, we also include in our analysis a third method, which is the traditional pooling panel model with fixed effects, where both long- and short-run parameters are constrained to be equal across all firms within each country. To tell which model utilizes information better, tests according to Hausman (1978) is utilized.

Model in (15) estimates idiosyncratic spreads of each firm μ_j separately. Its accuracy depends on whether all the other coefficients are estimated correctly. Although traditional pooling panel estimation could allow the fixed effect μ_j to serve as an estimate for idiosyncratic spread, restricting all other coefficients to be the same would just result in inconsistent estimates of μ_j. The two main ARDL methods, MG and PMG, we employ both allow short-run coefficients and μ_j to differ across firms. So their estimates for (15) would produce more accurate idiosyncratic spreads than the traditional change-based panel model, and thus benefit practitioners more in diversification within or across borders.

4. Empirical Findings

Table 1
Summary Statistics of Investment-Grade Corporate Bond Spreads

	No. of Issues	Average Maturity	Average Rating[a]	AA Average Spreads (bp)	A Average Spreads (bp)	BBB Average Spreads (bp)
				3 to 7-year Maturities		
Canada	44	4.64	3.57	88.42	116.19	166.40
Germany	196	4.95	3.11	65.93	89.50	120.56
France	104	5.12	4.25	100.97	129.30	170.67
UK	119	5.56	4.90	122.16	163.76	206.52
US	221	5.03	4.14	97.76	126.95	167.01
				8 to 12-year Maturities		
Canada	29	8.87	3.98	143.38	179.54	220.07
Germany	147	9.25	3.42	110.89	141.49	186.43
France	110	10.18	4.73	176.14	213.07	263.63
UK	134	10.96	5.66	215.45	259.91	316.31
US	194	9.77	4.51	161.97	203.05	251.26

Value	1	2	3	4	5	6	7	8	9	10
Moody's	Aaa	Aa1	Aa2	Aa3	A1	A2	A3	Baa1	Baa2	Baa3
Standard & Poors	AAA	AA+	AA	AA-	A+	A	A-	BBB+	BBB	BBB-

Monthly investment-grade industrial corporate bond yields reported in this table are obtained from Bloomberg for the period between 2006 and 2011. Issues with floating coupon rates and embedded options are not included. Also, issues with unreasonably high or low prices are eliminated. Spreads for each issue in the corresponding maturity category are calculated against yields of average government bond with the closest matching maturity in the respective country.
[a] Rating scales are in the following chart.

For the estimation of (15) we use monthly pricing data of corporate bonds issued in Canada, Germany, France, UK and US from Bloomberg between January 2006 and June 2011. Only yields of investment-grade industrial coupon bonds with maturities between 3 and 12 years, and Standard & Poor credit ratings of AA, A or BBB, are collected and those with floating coupon rates and embedded options are not included. Unreasonably high or low prices are also discarded. Issues from other countries are not included as there are too few concurrent issues available to support the construction of spot yields. Table 1 shows that issues from France and UK have the longer maturities and lower credit ratings among the five countries. Government bond yields are obtained from Thompson Datastream for the same period. The variable $\boldsymbol{TS}_t$ in (15) is constructed by taking the monthly difference between yields of 10- and 2-year government bonds of each country. Each country's repo rates for $\boldsymbol{RP}_t$ in (15) are also from Bloomberg.

Nelson and Siegel (1987) method is used to compute for each country monthly spot yields of government bonds, as well as corporate bonds of each credit rating, with rounded maturities between 3 and 12 years. Although fixing the shape parameters could be problematic as argued by Annaert, Claes and Zhang (2013), we still adopt the ordinary approach due to the complexity of it remedy. Individual corporate spreads are calculated for each rating-maturity category and then combined and averaged into a short maturity group (3 to 7 years) as well as a long maturity group (8 to 12 years). Table 1 also shows that average yield spreads for the former group are about 100 to 120 basis points lower than the latter in a given rating class. The yield spreads of long maturity US issues reported in Table 1 are compatible with the average yield spread between Moody's seasoned Baa corporate bond portfolio and 30-year US Treasury bond, which amounts to 223 basis points. For the Moody AA portfolio spread is around 152 bps. Our spread estimates for the long-maturity category are higher probably because our spreads are based on spot yields and also Moody portfolio includes issues from utility and financial companies.

Treating US as the benchmark country, we use the average yield spreads of AA bonds in the corresponding maturity category from US as the global systematic risk benchmark $\boldsymbol{SP}_t^G$, and we use the average yield spreads of AA bonds in Canada, Germany, France or UK as the local systematic risk benchmark, $\boldsymbol{SP}_t^D$, for the respective country. We apply (15) on the spot yield spreads of the other four countries with the help of the ***xtpmg*** procedure provided in the *Stata* package, which is available only after 2007. Allowing heterogeneous short-run dynamics helps giving better statistical properties to long-run parameters, which are $\boldsymbol{\theta}_j$ under the MG method and $\boldsymbol{\theta}$ under the PMG method. For comparison, we add in a traditional panel fixed-effect model, which constrains ϕ_j, λ_{jk}^* and the vector $\boldsymbol{\delta}_{jk}^*$ to be the same across *j*.

For simplicity, we adopt the error correction form of an ARDL(1,1,1,1,1) version of (15), for all of the four maturity-rating categories[5], as follows,

$$\Delta \boldsymbol{SP}_{jt} = \phi_j (\boldsymbol{SP}_{j,t-1} - \boldsymbol{\theta}_j^{'} \boldsymbol{X}_{jt}) + \lambda_j^* \Delta \boldsymbol{SP}_{j,t-1} + \boldsymbol{\delta}_j^{'*} \Delta \boldsymbol{X}_{j,t-1} + \mu_j + \varepsilon_{jt}, \tag{16}$$

where $\phi_j = -(1-\lambda_j)$, $\boldsymbol{\theta}_j = \delta_j/(1-\lambda_j)$, $\lambda_j^* = -\lambda_j$ and $\boldsymbol{\delta}_j^* = -\boldsymbol{\delta}_j$. The traditional panel fixed-effect model constrains ϕ_j, λ_j^* and vector $\boldsymbol{\delta}_j^*$ to be the same across *j*, while the MG method loops through all

[5] According the Variable Addition Test (VAT) specified in Pesaran, et al. (2001), ARDL(2,2,2,1,1) should be chosen for the short maturity-rating A category, with lags selected based on Schwarz Bayesian Criterion. ARDL(1,2,2,2,1), ARDL(2,2,1,1,1) and ARDL(1,1,1,2,1) are the appropriate models according to VAT for the long maturity-rating A, short maturity-rating BBB and long maturity-rating BBB respectively. Analyzing the ARDL(1,1,1,1,1) model instead, however, affects mainly the short-run estimates. Long-run estimates, which are our focus, are only slightly different.

firms in each country and reports the unweighted average of θ_j, ϕ_j, λ_j^* and δ_j^*. The PMG method constrains θ_j to be equal to θ for all j but reports also the average of ϕ_j, λ_j^* and δ_j^*. We use the average yield spreads of AA bonds in the corresponding maturity category from US as the global systematic risk benchmark SP_t^G, and we use the average yield spreads of AA bonds in Canada, Germany, France or UK as the domestic systematic risk benchmark, SP_t^D, for each country.

Table 3 gives the results, for issues with short maturities and the rating of A, from the dynamic Fixed Effect (FE), Mean Group (MG) ARDL and Pooled Mean Group (PMG) ARDL estimations based on (16). Few of the long-run decomposition coefficients (θ_j) and the short-run ones (δ_j^*) from the dynamic FE model are significant, except for the long run coefficient for RP_t and the error correction coefficient. Most of the long-run decomposition coefficients from the PMG ARDL model are significant at the 1% level, while only half of the coefficients from the MG ARDL model are significant. The short-run decomposition coefficients are mostly insignificant. The error correction coefficients (ϕ_j) are, however, uniformly significant across all three models, with the PMG and MG models exhibiting stronger significance. Across the four countries studied, estimates for issues in Germany and UK appear to exhibit stronger statistical significance in general. Hausman tests results indicate that PMG model utilization information better than the MG and dynamic FE models.

Estimated coefficients for SP_t^G, the global systematic benchmark, are also uniformly more significant than SP_t^D, the local systematic benchmark across all four countries in Table 3. The lack of significance in estimated decomposition coefficients from the dynamic FE model suggests that its weaker statistical power stems from cross-panel constraining both the long- and short-run coefficient estimates to be the same across spread time series of all firms. The highly significant Hausman test result in comparing the dynamic FE against the MG method is consistent with the statement above, so is the fact that coefficient estimates from the latter model are in general more significant those from the former. Although the PMG model requires, for each country, all the long-run decomposition coefficients to be the same across individual corporate spread series, which causes the estimated standard deviations from the PMG method to be higher than those from the MG method, significance in long-run coefficients and Hausman tests between the two models are in favor of PMG over MG.

Table 2

Cross-border Yield Spread Decomposition with ARDL Error Correction Estimations,
Short maturities and credit rating A

	Canada		*Germany*		*France*		*UK*	
Dynamic FE Model								
Error correction (long-run)								
SP_t^G	0.2807	(0.1713)	0.2141	(0.1342)	0.3159	(0.1984)	0.3026**	(0.1185)
SP_t^D	0.1739	(0.1362)	0.1031	(0.0755)	0.1552	(0.1236)	0.1218	(0.1094)
RP_t	0.2404**	(0.0837)	0.1364**	(0.0556)	0.1835*	(0.0878)	0.3135**	(0.0992)
TS_t	0.1522	(0.0976)	0.1055	(0.0683)	0.2199	(0.1767)	0.2825*	(0.1410)
Short-run								
ϕ	-0.1236*	(0.0538)	-0.1481**	(0.0419)	-0.1148*	(0.0557)	-0.1669**	(0.0542)
ΔSP_t^G	0.0203	(0.1014)	-0.0317*	(0.0168)	0.0183	(0.1125)	-0.0545*	(0.0252)
ΔSP_t^D	0.4165	(0.9836)	-0.9006	(0.8815)	0.2455	(1.1356)	-1.3793	(1.0066)
ΔRP_t	-2.2980	(1.7361)	-1.5059*	(0.8213)	-3.0773	(1.9980)	-3.3837*	(1.6061)
ΔTS_t	-0.0654	(0.0558)	0.0325	(0.0539)	0.0271	(0.0845)	-0.0844*	(0.0416)
MG Model								
Error correction (long-run)								
SP_t^G	0.3171**	(0.1461)	0.2224**	(0.0846)	0.3479**	(0.1376)	0.3631**	(0.1064)
SP_t^D	0.1998	(0.1010)	0.1313*	(0.0625)	0.1928*	(0.0861)	0.1893*	(0.0887)
RP_t	0.2605**	(0.0481)	0.2139**	(0.0491)	0.2858**	(0.0776)	0.3007**	(0.0713)
TS_t	0.1627	(0.0902)	0.1129	(0.0611)	0.2434	(0.1483)	0.2922*	(0.1348)
Short-run								
ϕ	-0.2377**	(0.0401)	-0.2678**	(0.0338)	-0.2273**	(0.0446)	-0.2761**	(0.0385)
ΔSP_t^G	-0.0475*	(0.0221)	-0.0647**	(0.0188)	0.0019	(0.0449)	-0.0529**	(0.0164)
ΔSP_t^D	-0.2098	(0.2046)	-0.7293	(0.4756)	-0.1786	(0.5327)	-0.4489	(0.2855)
ΔRP_t	-2.7776*	(1.3592)	-1.6331**	(0.6695)	-3.3015*	(1.6234)	-3.1903**	(1.1220)
ΔTS_t	-0.0803	(0.0504)	-0.0442	(0.0303)	-0.0381	(0.0440)	-0.0767*	(0.0374)
PMG Model								
Error correction (long-run)								
SP_t^G	0.3628**	(0.1493)	0.2561**	(0.0521)	0.3733**	(0.1443)	0.4101**	(0.1246)
SP_t^D	0.2264	(0.1215)	0.1787**	(0.0649)	0.2512*	(0.1208)	0.2418*	(0.1025)
RP_t	0.2718**	(0.0527)	0.2220**	(0.0655)	0.2945**	(0.0790)	0.3252**	(0.0706)
TS_t	0.2196*	(0.0968)	0.1557**	(0.0679)	0.2756*	(0.1318)	0.2800**	(0.1157)
Short-run								
ϕ	-0.2686**	(0.0419)	-0.2709**	(0.0375)	-0.2554**	(0.0497)	-0.2888**	(0.0320)
ΔSP_t^G	-0.0431	(0.0249)	-0.0626**	(0.0201)	-0.0550*	(0.0276)	-0.0501**	(0.0188)
ΔSP_t^D	-0.3551	(0.8035)	-0.7559	(0.5213)	-0.1603	(0.6081)	-0.3445	(0.9294)
ΔRP_t	-2.2520	(1.5335)	-1.7893**	(0.6804)	-3.1314	(1.8525)	-3.2107**	(1.0049)
ΔTS_t	-0.0749	(0.0582)	-0.0410	(0.0311)	-0.0574	(0.0663)	-0.0713	(0.0353)

Hausman Tests

MG (unrestricted) over *Dynamic FE* (restricted)	$\chi^2(2)=11.37$	(p=0.0034)	*MG* is preferred over *FE*
MG (unrestricted) over PMG (restricted)	$\chi^2(2)=7.35$	(p=0.0253)	*PMG* is preferred over *MG*

For simplicity, we adopt the error correction form of the ARDL(1,1,1,1,1) version of (13) like

$$\Delta \boldsymbol{SP}_{jt} = \phi_j(\boldsymbol{SP}_{j,t-1} - \boldsymbol{\theta}_j^{'}\boldsymbol{X}_{jt}) + \lambda_j^{*}\Delta \boldsymbol{SP}_{j,t-1} + \boldsymbol{\delta}_j^{'*}\Delta \boldsymbol{X}_{j,t-1} + \mu_j + \varepsilon_{jt}\,,$$

where $\boldsymbol{SP}_{jt}$ is the yield spread of the bond issued by firm j at time t in a non-benchmark country, $\boldsymbol{X}_{jt}$=($\boldsymbol{SP}_t^G$, $\boldsymbol{SP}_t^D$, $\boldsymbol{RP}_t$, $\boldsymbol{TS}_t$)', $\boldsymbol{\delta}_{jk}$=(δ_{jk}^1, δ_{jk}^2, δ_{jk}^3, δ_{jk}^4)', $\boldsymbol{SP}_t^G$ is global spread benchmark and $\boldsymbol{SP}_t^D$ is the local spread benchmark in the local country as implied by A_D in (3). $\boldsymbol{RP}_t$ is the repo rate and $\boldsymbol{TS}_t$ would be the difference between long and short rates. ε_{jt} is the disturbance independently distributed across j and t with mean 0 and $\sigma_j^2 > 0$. μ_j is assumed to be the fixed effect for firm j in the panel ARDL model of (13), and can be considered as reflecting the idiosyncratic risk. Additionally, ϕ_j=-(1-λ_j), $\boldsymbol{\theta}_j$=δ_j/(1-λ_j), $\lambda_j^* = -\lambda_j$ and $\boldsymbol{\delta}_j^* = -\boldsymbol{\delta}_j$. The traditional panel fixed-effect model constrains ϕ_j, λ_j^* and vector $\boldsymbol{\delta}_j^*$ to be the same across j, while the Mean Group (MG) method loops through all firms in each country and reports the unweighted average of $\boldsymbol{\theta}_j$, ϕ_j, λ_j^* and $\boldsymbol{\delta}_j^*$. The Pooled Mean Group (PMG) method constrain $\boldsymbol{\theta}_j$ to be equal to $\boldsymbol{\theta}$ for all j but reports also the average of ϕ_j, λ_j^* and $\boldsymbol{\delta}_j^*$. we use the average yield spreads of AA bonds in the corresponding maturity category from US as the global systematic risk benchmark $\boldsymbol{SP}_t^G$, and we use the difference between average yield spreads of AA bonds in Canada, Germany, France or UK and $\boldsymbol{SP}_t^G$ as the local systematic risk benchmark, $\boldsymbol{SP}_t^D$, for the respective country.

* Significant at the 5% level.

** Significant at the 1% level.

Results in Table 2 also exemplify the advantage of applying an ARDL model in a hetergeneous panel. As level of terms retain more information than the difference terms of yield spreads, stronger significance exhibited by the long-run decomposition coefficients than the short-run ones demonstrates that an ARDL model works better in studying corporate yield spreads, possibly due to the information provided by level terms of lagged dependent variable as well as the level term of current independent variables. Based on the average yield and interest rate data within our data period, the PMG analysis in Table 3 predicts that the average long-run Canadian rating A short maturity corporate yield spread to amount to roughly 152 bps, only 13 bps below the observed average, while for UK that difference is about 10 bps. Through properly estimated long-run decomposition coefficients and μ_j, the fixed effect or the proxy for idiosyncratic risk in individual corporate spreads, our analysis would substantially help managing risks of holding corporate bond portfolios in a long period of time.

Table 3, 4 and 5 give results from the same procedures for the categories of long-maturity with rating BBB, short-maturity with rating A, as well as long-maturity with rating BBB. Uniformly significant error correction coefficients suggest apparent cointegration relationships exist among yield spreads and the four independent variables. Both the long- and shor-run coefficients go up in magnitude and the extent of significances is stronger with longer maturities and lower bond ratings[6]. Similar to the pattern in Table 3, across all the maturity-rating categories and countries, PMG model produces the largest coefficients and dynamic FE the smallest. The pattern of standard deviations is just the opposite. Hausman test results reported in each of the three tables also suggest the PMG procedure is superior to the MG and dynamic FE ones. In general, reconstructed yield spread estimates from coefficients given by the PMG model are slightly lower than the observed figures shown in Table 1, possibly due to apparent down-trend of yield spreads within the data period. It is also worth noting that the responses of yields to repo are in general stronger for issues with lower credit rating, validating the notion, brought up initially in Section 2, that the direct influence of short rate on yield spread should increase with credit risks.

[6] This is also consistent with findings in Lin and Sun (2009), which are based on US data and predict that yield spreads of bonds with lower credit rating would be more responsive to systematic risks.

Table 3

Cross-border Yield Spread Decomposition with ARDL Error Correction Estimations,
Short maturities and credit rating BBB

	Canada		***Germany***		***France***		***UK***	
Dynamic FE Model								
Error correction (long-run)								
SP_t^G	0.2887	(0.3921)	0.2242*	(0.1042)	0.3056	(0.3559)	0.3350**	(0.1623)
SP_t^D	0.2125	(0.1406)	0.1671	(0.1293)	0.2271	(0.1642)	0.2489	(0.2094)
RP_t	0.2514	(0.1308)	0.1770**	(0.0758)	0.2692	(0.1425)	0.3953**	(0.1218)
TS_t	0.2175	(0.1556)	0.1515	(0.1640)	0.3004	(0.2083)	0.3167	(0.2687)
Short-run								
ϕ	-0.1839*	(0.0964)	-0.1682**	(0.0521)	-0.1305*	(0.0585)	-0.1895**	(0.0638)
ΔSP_t^G	0.0287	(0.1495)	0.0198	(0.0344)	0.0290	(0.0955)	0.0808	(1.1563)
ΔSP_t^D	0.2332	(0.6697)	-0.9234	(0.9705)	0.0276	(0.6529)	-1.0005	(1.2560)
ΔRP_t	-1.8239	(1.9069)	-1.6434	(1.3371)	-1.1239	(1.4101)	-1.1191	(1.3987)
ΔTS_t	0.0468	(0.0721)	0.0500	(0.0658)	0.0460	(1.3037)	0.0119	(0.0884)
MG Model								
Error correction (long-run)								
SP_t^G	0.3049**	(0.1086)	0.2433**	(0.0846)	0.3274**	(0.1178)	0.3454**	(0.1369)
SP_t^D	0.2390*	(0.1115)	0.1835**	(0.0425)	0.2638*	(0.1235)	0.2993	(0.1661)
RP_t	0.3323**	(0.0774)	0.2571**	(0.0551)	0.3558**	(0.0848)	0.4007**	(0.0992)
TS_t	0.2921*	(0.1372)	0.1242*	(0.0613)	0.3160*	(0.1423)	0.3340*	(0.1582)
Short-run								
ϕ	-0.2854**	(0.0593)	-0.3036**	(0.0451)	-0.2518**	(0.0604)	-0.3157**	(0.0522)
ΔSP_t^G	-0.0535*	(0.0269)	-0.0777**	(0.0232)	-0.0733*	(0.0349)	-0.0529**	(0.0164)
ΔSP_t^D	-0.4198	(0.2834)	-0.7548	(0.4234)	-0.5985	(0.5610)	-0.8758	(0.6949)
ΔRP_t	-3.2528*	(1.5678)	-2.2480*	(1.1294)	-3.6420*	(1.7881)	-2.5596**	(1.1027)
ΔTS_t	-0.0838	(0.0543)	-0.0639	(0.0466)	-0.1255	(0.0664)	-0.1201*	(0.0561)
PMG Model								
Error correction (long-run)								
SP_t^G	0.3287**	(0.0959)	0.2834**	(0.0669)	0.3635**	(0.1031)	0.3834**	(0.1156)
SP_t^D	0.2791**	(0.0982)	0.2206**	(0.0404)	0.3017**	(0.1093)	0.3208**	(0.1200)
RP_t	0.3494**	(0.0695)	0.2689**	(0.0425)	0.3740**	(0.0704)	0.4203**	(0.0775)
TS_t	0.2267**	(0.0736)	0.1488**	(0.0510)	0.2838**	(0.1050)	0.3151**	(0.1117)
Short-run								
ϕ	-0.2994**	(0.0501)	-0.3237**	(0.0404)	-0.2994**	(0.0385)	-0.3753**	(0.0480)
ΔSP_t^G	-0.0510*	(0.0252)	-0.0714**	(0.0230)	-0.0677*	(0.0325)	-0.0488**	(0.0156)
ΔSP_t^D	-0.3915	(0.1994)	-0.7878	(0.4101)	-0.4065	(0.6081)	-0.7932	(0.5825)
ΔRP_t	-3.0511*	(1.5492)	-1.9676**	(0.8180)	-3.4298*	(1.7332)	-2.7685**	(1.0032)
ΔTS_t	-0.0831	(0.0515)	-0.0609	(0.0383)	-0.0899	(0.0602)	-0.0774	(0.0498)

Hausman Tests

MG (unrestricted) over *Dynamic FE* (restricted)	$\chi^2(2)$=10.86	(p=0.0044)	*MG* is preferred over *FE*
MG (unrestricted) over PMG (restricted)	$\chi^2(2)$=6.21	(*p*=0.0448)	*PMG* is preferred over *MG*

* Significant at the 5% level.
** Significant at the 1% level.

Table 4
Cross-border Yield Spread Decomposition with ARDL Error Correction Estimations,
Long maturities and credit rating A

	Canada		Germany		France		UK	
Dynamic FE Model								
Error correction (long-run)								
SP_t^G	0.3796*	(0.1793)	0.2169**	(0.0908)	0.3878*	(0.1827)	0.4154*	(0.2110)
SP_t^D	0.2550	(0.1599)	0.1894	(0.1009)	0.2029	(0.1684)	0.3038	(0.1889)
RP_t	0.2778**	(0.1047)	0.1964**	(0.0719)	0.2957**	(0.1068)	0.3343**	(0.1201)
TS_t	0.3078	(0.1802)	0.1605	(0.1011)	0.3158	(0.1970)	0.3539	(0.2123)
Short-run								
ϕ	-0.1695**	(0.0828)	-0.1553**	(0.0517)	-0.1284**	(0.0523)	-0.1774**	(0.0567)
ΔSP_t^G	-0.0392	(0.1115)	0.0198	(0.0344)	0.0290	(0.0955)	0.0808	(1.1563)
ΔSP_t^D	0.2567	(0.5883)	-0.9234	(0.9705)	0.0276	(0.6529)	-1.0005	(1.2560)
ΔRP_t	-2.2376	(1.3506)	-1.6434	(1.3371)	-1.1239	(1.4101)	-1.1191	(1.3987)
ΔTS_t	0.0689	(0.0678)	0.0500	(0.0658)	0.0460	(1.3037)	0.0119	(0.0884)
MG Model								
Error correction (long-run)								
SP_t^G	0.3927**	(0.1675)	0.2253**	(0.0704)	0.3968**	(0.1774)	0.4292**	(0.1998)
SP_t^D	0.2835**	(0.1363)	0.2189**	(0.0921)	0.3086**	(0.1410)	0.3168*	(0.1544)
RP_t	0.3769**	(0.0902)	0.2123**	(0.0570)	0.3629**	(0.0928)	0.3705**	(0.1092)
TS_t	0.3130	(0.1583)	0.1823*	(0.0896)	0.3204	(0.1632)	0.3689*	(0.1711)
Short-run								
ϕ	-0.2620**	(0.0565)	-0.2744**	(0.0409)	-0.2485**	(0.0546)	-0.3011*	(0.0473)
ΔSP_t^G	-0.0621**	(0.0249)	-0.0824**	(0.0222)	-0.0841**	(0.0286)	-0.0793**	(0.0147)
ΔSP_t^D	-0.4776	(0.2613)	-0.8135	(0.4202)	-0.6502	(0.4568)	-0.8086	(0.5097)
ΔRP_t	-3.8814**	(1.4485)	-2.5371**	(1.0076)	-3.9749**	(1.5135)	-2.7419**	(1.0203)
ΔTS_t	-0.0903	(0.0511)	-0.0712	(0.0413)	-0.1302	(0.0671)	-0.1310	(0.0636)
PMG Model								
Error correction (long-run)								
SP_t^G	0.4214**	(0.1559)	0.3107**	(0.0592)	0.4235**	(0.1610)	0.4454**	(0.1635)
SP_t^D	0.3033*	(0.1243)	0.2293**	(0.0840)	0.3252**	(0.1307)	0.3438**	(0.1349)
RP_t	0.4120**	(0.0700)	0.2976**	(0.0463)	0.4198**	(0.0729)	0.4335**	(0.0917)
TS_t	0.3565**	(0.1329)	0.2367**	(0.0711)	0.3623**	(0.1417)	0.3794**	(0.1668)
Short-run								
ϕ	-0.2828**	(0.0533)	-0.3110**	(0.0387)	-0.2754**	(0.0332)	-0.3555**	(0.0426)
ΔSP_t^G	-0.0767**	(0.0199)	-0.0887**	(0.0230)	-0.0885**	(0.0251)	-0.0861**	(0.0127)
ΔSP_t^D	-0.4898*	(0.2207)	-0.8381*	(0.4006)	-0.7047	(0.4250)	-0.8889*	(0.4843)
ΔRP_t	-3.9624**	(1.2321)	-2.7885**	(0.7354)	-4.0095**	(1.4039)	-3.2473**	(0.0844)
ΔTS_t	-0.1014*	(0.0495)	-0.0733*	(0.0321)	-0.1357*	(0.0590)	-0.1423*	(0.0619)

Hausman Tests

MG (unrestricted) over *Dynamic FE* (restricted)	$\chi^2(2)=11.09$	(p=0.0039)	*MG* is preferred over *FE*
MG (unrestricted) over PMG (restricted)	$\chi^2(2)=6.44$	(p=0.0399)	*PMG* is preferred over *MG*

* Significant at the 5% level.
** Significant at the 1% level.

Table 5

Cross-border Yield Spread Decomposition with ARDL Error Correction Estimations,
Long maturities and credit rating BBB

	Canada		*Germany*		*France*		*UK*	
Dynamic FE Model								
Error correction (long-run)								
SP_t^G	0.3610*	(0.1761)	0.2743*	(0.0915)	0.3702*	(0.1810)	0.4132**	(0.1234)
SP_t^D	0.3134	(0.1847)	0.2667	(0.1396)	0.3273	(0.1997)	0.3218	(0.2021)
RP_t	0.3887*	(0.1520)	0.2929**	(0.0787)	0.3914*	(0.1518)	0.4133**	(0.1349)
TS_t	0.3946	(0.2173)	0.2989*	(0.1421)	0.4065	(0.2299)	0.4201	(0.2547)
Short-run								
ϕ	-0.2071**	(0.0915)	-0.1920**	(0.0709)	-0.1556**	(0.0863)	-0.1895**	(0.0661)
ΔSP_t^G	-0.0411	(0.1346)	0.0048	(0.0344)	0.0076	(0.0955)	-0.0639	(0.0624)
ΔSP_t^D	0.1024	(0.6004)	-1.1453	(1.2232)	-1.3897	(0.8055)	-1.0005	(1.4981)
ΔRP_t	-2.4435	(1.3883)	-1.8896	(1.5904)	-1.4465	(1.6274)	-1.1191	(1.5734)
ΔTS_t	0.0113	(0.0772)	0.0038	(0.0745)	0.0854	(1.5298)	-0.1033	(0.0721)
MG Model								
Error correction (long-run)								
SP_t^G	0.4032**	(0.1158)	0.3099**	(0.0827)	0.4209**	(0.1025)	0.4665**	(0.0842)
SP_t^D	0.3682**	(0.1504)	0.3160**	(0.1053)	0.3741**	(0.1621)	0.3817**	(0.1698)
RP_t	0.4306**	(0.0883)	0.3215**	(0.0692)	0.4322**	(0.1033)	0.4464**	(0.0756)
TS_t	0.4367*	(0.1901)	0.3324**	(0.1235)	0.4371*	(0.1978)	0.4506**	(0.2110)
Short-run								
ϕ	-0.2881**	(0.0849)	-0.2912**	(0.0675)	-0.2769**	(0.0721)	-0.3305*	(0.0539)
ΔSP_t^G	-0.0731**	(0.0277)	-0.0893**	(0.0319)	-0.0841**	(0.0286)	-0.0869**	(0.0201)
ΔSP_t^D	-0.5104	(0.2900)	-0.9494	(0.4445)	-0.6502	(0.4568)	-0.8818	(0.7360)
ΔRP_t	-4.1035**	(1.5885)	-2.7344**	(1.1769)	-3.9749**	(1.5135)	-2.8323**	(1.1783)
ΔTS_t	-0.1097	(0.0644)	-0.0921	(0.0525)	-0.1302	(0.0671)	-0.1388	(0.0684)
PMG Model								
Error correction (long-run)								
SP_t^G	0.4466**	(0.0914)	0.3576**	(0.0710)	0.4663**	(0.0933)	0.4960**	(0.0728)
SP_t^D	0.3830*	(0.1102)	0.3421**	(0.0923)	0.4102**	(0.1267)	0.4273**	(0.1413)
RP_t	0.4653**	(0.0668)	0.3693**	(0.0505)	0.4723**	(0.0914)	0.4857**	(0.0680)
TS_t	0.4714**	(0.1194)	0.3708**	(0.0873)	0.4786*	(0.1242)	0.4951**	(0.1312)
Short-run								
ϕ	-0.3008**	(0.0801)	-0.3354**	(0.0502)	-0.2995**	(0.0665)	-0.3764**	(0.0498)
ΔSP_t^G	-0.0840**	(0.0229)	-0.0915**	(0.0289)	-0.0885**	(0.0251)	-0.0928**	(0.0175)
ΔSP_t^D	-0.5457*	(0.2621)	-0.9648*	(0.4213)	-0.7047	(0.4250)	-0.9190*	(0.5679)
ΔRP_t	-4.3478**	(1.3796)	-3.1308**	(0.8405)	-4.0095**	(1.4039)	-3.1415**	(1.1062)
ΔTS_t	-0.1241*	(0.0550)	-0.1039**	(0.0440)	-0.1357*	(0.0590)	-0.1267*	(0.0640)
Hausman Tests								
MG (unrestricted) over *Dynamic FE* (restricted)			$\chi^2(2)=9.15$	(p=0.0103)	*MG* is preferred over *FE*			
MG (unrestricted) over PMG (restricted)			$\chi^2(2)=5.79$	(p=0.0553)	*PMG* is preferred over *MG*			

* Significant at the 5% level.
** Significant at the 1% level.

To demonstrate the crucial implication of potential diversification benefit from our PMG ARDL estimation method, we take, μ_j, the estimated fixed effect or proxy for idiosyncratic component in (16) and compare it against a counterpart from the following model,

$$\Delta SP_{jt} = \alpha_j + \beta_j \Delta SP_t^D + \gamma_{1j} \Delta RP_t + \gamma_{2j} \Delta TS_t + \eta_j + \nu_{jt}, \quad (17)$$

where η_j would be the alternative idiosyncratic component and ν_{jt} is the disturbance term. (17) emulates the commonly adopted change-based *domestic* yield spread decomposition model, like

Duffee (1998).

Table 6

VaR Analysis of Corporate Bond Portfolios,

Cross-border PMG ARDL approach versus traditional domestic approach

	Canada		***Germany***		***France***		***UK***	
	Cross-Border	Domestic	Cross-Border	Domestic	Cross-Border	Domestic	Cross-Border	Domestic
Short Maturities, A								
Portfolio 1% VaR[a]	-119.35	-139.97	-103.29	-111.94	-128.68	-146.89	-132.68	-145.80
S.D. of individual VaR's[a]	30.83	45.11	29.45	43.26	41.71	59.64	32.63	47.36
Paired *t*-tests	$t_{d.f.:19}$=-1.83	p=0.0129	$t_{d.f.:78}$=-2.14	p=0.0176	$t_{d.f.:44}$=-2.43[b]	p=0.0096	$t_{d.f.:49}$=-2.34	p=0.0116
Short Maturities, BBB								
Portfolio 1% VaR[a]	-131.21	-154.14	-109.38	-119.15	-134.77	-153.32	-134.64	-148.45
S.D. of individual VaR's[a]	35.27	50.51	32.91	48.55	43.85	61.41	34.69	50.08
Paired *t*-tests	$t_{d.f.:21}$=-2.50[b]	p=0.0104	$t_{d.f.:89}$=-2.31	p=0.0116	$t_{d.f.:52}$=-2.59[b]	p=0.0062	$t_{d.f.:51}$=-2.37	p=0.0107
Long Maturities, A								
Portfolio 1% VaR[a]	-135.46	-163.38	-113.81	-125.74	-136.22	-156.89	-136.54	-151.07
S.D. of individual VaR's[a]	34.09	50.26	33.53	49.59	44.63	63.19	40.81	55.11
Paired *t*-tests	$t_{d.f.:12}$=-2.34	p=0.0187	$t_{d.f.:65}$=-2.36	p=0.0106	$t_{d.f.:52}$=-2.82[b]	p=0.0034	$t_{d.f.:59}$=-2.37[b]	p=0.0104
Long Maturities, BBB								
Portfolio 1% VaR[a]	-138.06	-170.86	-118.42	-132.34	-140.13	-164.18	-139.96	-157.48
S.D. of individual VaR's[a]	30.83	45.11	35.31	53.76	46.61	66.24	43.67	58.02
Paired *t*-tests	$t_{d.f.:13}$=-2.69[b]	p=0.0090	$t_{d.f.:68}$=-2.63[b]	p=0.0053	$t_{d.f.:53}$=-3.17[b]	p=0.0013	$t_{d.f.:63}$=-2.79	p=0.0035

We calculate μ_j according to (14) based on parameter estimates from Table 2 through 5 for all firms in each country, as the proxies for idiosyncratic component of our corporate yield decomposition. Then we construct a traditional domestic approach counterpart in an ordinary panel OLS model, for each country, like,

$$\Delta \boldsymbol{SP}_{jt} = \alpha_j + \beta_j \Delta \boldsymbol{SP}_t^{D} + \gamma_{1j} \Delta \boldsymbol{RP}_t + \gamma_{2j} \Delta \boldsymbol{TS}_t + \eta_j + \nu_{jt} \text{ ,}$$

where η_j is corresponding idiosyncratic component from the alternative model. To compute VaR estimates, we rank μ_j and η_j derived from the two models for all the firms in a given country. The bottom values of μ and η for each firm are identified as our approximated historically simulated 1% VaR (quantile) estimates for the two models respectively. For each of the four countries, equally weighted portfolios are constructed separately for short and long maturities, as well as for ratings A and BBB. Each country's portfolio VaR is the average of all the individual firm VaR's.

a Numbers are in basis points.

b Significant at the 1% level.

In each of the four countries studied, historically simulated 1% one-tailed Value-at-Risk (VaR) estimates from ranked individual PMG-produced μ_j are identified for every single issue within a given maturity-rating category in the country. As there are only at most 66 observations for any issue, the smallest μ_j is selected as a proxy for the VaR estimate. A similar procedure is carried out on ranked η_j, and VaR estimates are obtained accordingly. For each of the four countries, equally weighted portfolios are constructed separately for short and long maturities, as well as for ratings A and BBB. The average of all individual firm's VaR's in each country would be adopted, in the spirit of Venkatesh (2003), as portfolio VaR of that country. Paired *t*-test results are given in Table 6 for each country and each maturity-rating category respectively.

Overall, the down side VaR estimates for the PMG model average at -132.93 b.p., while the average for the alternative model is -155.73 b.p.. The results of paired *t* tests are barely significant at the 1% level, except for France, within the category of short maturity and rating A. Lower and more significant *p* values appear as we move to longer maturity and lower credit rating, across all countries. The VaR analysis of bond portfolios in Table 6 indicates that the benefit of diversifying idiosyncratic risks produced by our PMG ARDL procedure is substantially greater than an alternatively constructed change-based *domestic* panel OLS procedure. Furthermore, combining all the VaR estimates across all four countries for a given maturity-rating category yields *t*-statistics more than twice as large, suggesting potential existence of further cross-border diversification benefits very much needed by managers of international bond portfolios.

V. Conclusion

This study recalibrates corporate bond idiosyncratic risks in the context of international portfolio diversification. Based on the ideas of Dastidar and Phelps (2011), Xie, et al. (2008), Churm and Panigirtzoglou (2007) and Venkatesh (2003), we extend the model of Sun, et al. (2007) to a cross-border context. The empirical framework of Pesaran, et al. (1999) is used to process the cross-border heterogeneous panels. By introducing a statistically powerful risk decomposition scheme, we show in this study that diversification is improved as both global and domestic risk benchmarks are utilized. Not only fixed income portfolio management, but also the pricing of traditional and innovative financial instruments can benefit from the scheme proposed in this very study.

In addition to domestic factors like interest rate and default risks, we include a global risk benchmark to be consistent with the thoroughly integrated international capital market. The ARDL panel time series model of Pesaran, et al. (1999), which emphasizes long-run relations among economic variables, helps us building a long-run stationary yield spread decomposition scheme in our study. We have employed top-grade yield spreads of US corporate bonds as the global risk benchmark, but many other instruments can serve the same purpose. The inclusion of a global risk component provides more abundant and explicit information, which the traditional domestic model lacks, for pricing and risk management practices of fixed income portfolios. The global and domestic risk benchmarks are observable yield spreads of traded top-grade bonds, so there are few measurement or availability issues involved. The idiosyncratic risk component is estimated as a fixed effect in a data panel along with all the parameter estimates, rather than being introduced separately from a exogenous generating process. Our linear model may contain other econometric imperfections, but our estimates can be applied promptly and easily by practitioners.

The idiosyncratic component of yield spread has been estimated from three different models. Hauseman tests show that the PMG ARDL method is the best in utilizing available information. The VaR analysis verifies that the idiosyncratic risks generated under this procedure have substantially better diversification implication than an alternatively constructed change-based domestic panel OLS procedure. So the results of our study not only extend a purely domestic fixed income model to a cross-border one, but they also help enhancing the diversification capability of international fixed income portfolios.

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