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International linkages of term structures: US and Korea Treasury bond yields

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ABSTRACT

This study analyzes the international linkages of term structures between the United States (US) and Korea using a Gaussian dynamic term structure model. The empirical analysis shows that the US level factor makes the most important contribution to the Korean term structure fluctuations. When two separate channels (i.e., policy and risk-compensation) are considered for propagating structural shocks into Korean bond yields, the policy channel dominates for the short-term rates, whereas for the long-term yield, the policy channel is dominant for the US level shock, but the risk channel is dominant for other shocks. These results are sensitive to whether or not small-sample bias is corrected. With a help of the bias correction, term premium estimates in the US and Korea exhibit substantial countercyclicality with respect to the US output gap. The US slope factor, which also has a significant impact on Korean long-term yields, is closely related to the global liquidity conditions for Korean bond returns.

1. Introduction

Many researchers have documented that global factors play an important role in the international comovement of bond yields (e.g., Diebold, Li, and Yue 2008, Dahlquist and Hasseltoft 2013, Jotikasthira, Le, and Lundblad 2015, Abbritti, Dell'Erba, Moreno, and Sola 2018). For example, according to Jotikasthira et al. (2015), the correlations of 10-year Treasury yields between the US and UK and between the US and Germany are all greater than 0.8. However, researchers have mostly studied the effects of global factors on term structures in advanced economies. It would be interesting to know whether the phenomenon of advanced economies also applies to emerging market economies.

Yun (2019) investigated the effect of global liquidity factors on the risk premia in Korea Treasury Bonds (KTBs). According to Yun (2019), the Korean financial market features highly volatile capital flows and nonreserve currency, and the KTB market is highly liquid with diverse bond maturities. The KTB market has been fully open since 1998 (immediately after the Asian financial crisis), and foreign investors have substantially influenced bond prices through both spot and futures markets.¹ As such, the Korean bond market is unique because of high market liquidity and a well-established market infrastructure even though it is an emerging market.

My aim with this study is to analyze the influence of structural shocks from certain global yield-curve factors on the term structure in Korea. To approximate the factors, I employ the US yield-curve factors because the US has the largest share in the global economy; therefore, US variables function sufficiently as proxies for global variables. For example, Abbritti et al. (2018) analyzed the effects of

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¹ The reader may refer to Yun (2019) for further details.

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the global yield-curve factors extracted from several countries' bond yields on the term structures in the selected individual countries and obtained similar results to those obtained using solely the US factors instead.

For this paper, I use a Gaussian dynamic term structure model (henceforth, GDTSM), which has been enormously popular in the finance literature since [Ang and Piazzesi \(2003\)](#) produced their work. GDTSMs exploit the predictability of future interest rates (from a vector autoregressive model) in a framework that rules out the possibility of pure arbitrage ([Gürkaynak and Wright, 2012](#)). Belonging to the affine class of models, GDTSMs provide tractable solutions for bond yields.

To specify a vector autoregression (VAR) implied by a GDTSM, I employ a two-block structure with US and Korean blocks. Each block consists of each country's yield-curve factors, specifically the first three principal components of the yield covariance matrix: level, slope, and curvature. The US block, which I use as a proxy to represent the global economy, is set exogenously to have a unilateral influence on the Korean block, and the GDTSM in this paper ultimately addresses six state variables.

I consider two separate channels for propagating structural shocks into Korean bond yields. As [Jotikasthira et al. \(2015\)](#) point out, international comovement of bond yields arises through either monetary policy (related to fundamental economic factors) or risk compensation (related to investor linkages). Through the policy channel, internal and external economic shocks affect current and expected future short-term rates as central banks react to changing economic environments. In contrast, through the risk compensation channel, investors can respond to these shocks by altering their required compensation for risk. In the term structure literature, bond yields have often been decomposed into expectation and term premia; in fact, the policy and risk compensation channels correspond to the expectation and term premia, respectively.

To analyze a GDTSM, I employ [Joslin, Singleton, and Zhu's \(2011\)](#) maximally flexible model: A GDTSM is transformed into a maximally flexible canonical representation via Jordan decomposition. Joslin et al.'s method requires that the pricing factors in a GDTSM be linear combinations of observed yields, and the principal components from bond yields, which I employ as the pricing factors in this paper, by construction, satisfy this condition; the maximally flexible models can be conveniently estimated by maximum likelihood estimation (MLE) with low computational cost. Meanwhile, [Ludvigson and Ng \(2009\)](#), [Duffee \(2011\)](#), and [Joslin, Priebsch, and Singleton \(2014\)](#), among others, argued that *unspanned* macro factors have predictive power for excess bond returns beyond the information contained in the yield-curve. In the current paper, I treat US yield-curve factors as unspanned for the term structure in Korea.

Another consideration in GDTSM estimation is that highly persistent bond yields can cause serious small-sample bias, which can engender too-fast mean reversion. Given biased estimates, the path of future short-term interest rates projected by a biased model becomes too smooth, and model-implied term premium estimates are also distorted. To avoid this problem, I perform bias corrections using [Kilian's \(1998\)](#) bootstrap-after-bootstrap method.

The results of the empirical analysis in this paper are as follows. First, the empirical results from impulse responses and variance decompositions show that the US level factor makes the largest and most persistent contribution to fluctuations in the term structure in Korea, generating a strong tendency of comovement between the US and Korean Treasury yields. Second, when the two separate channels (i.e., policy and risk compensation) are considered for propagating structural shocks into KTB yields, the policy channel dominates in the short-term yield. However, in the long-term yield, the dominant transmission channels differ depending on the type of yield-curve shock. For US level shocks, the policy channel dominates, and for other shocks, the risk channel dominates. The channel dominant for US level shocks is particularly sensitive to whether or not small-sample bias is corrected.

Third, the GDTSM-implied five- to ten-year forward term premia in the US and Korea show substantial countercyclicality with respect to the US output gap. The bias correction produces more countercyclicality in the term premium estimates. Finally, the US yield-curve factors are closely related to [Yun's \(2019\)](#) global liquidity factor (GLF). Yun investigated bond risk premia embedded in KTBs by extracting the GLF from a panel data set of 39 global liquidity variables, and showed that compared with a variety of competitors including macro factors, yield-curve factors, and the Cochrane–Piazzesi factor, the GLF was the only predictor that performed well across both in- and out-of-sample forecasts for one-year-ahead excess returns on the KTBs. In particular, the US slope is the most closely related to the global liquidity condition for KTB returns.

The remainder of this paper is structured as follows. In [Section 2](#), I summarize the related past studies, and in [Section 3](#), describe the GDTSM I used in the empirical analyses. In [Section 4](#), I explain the results of the empirical analyses. In [Section 5](#), I conduct robustness checks, and [Section 6](#) concludes the paper.

2. Related literature

In this paper, I analyze the influence of structural shocks identified from global yield-curve factors on the term structure in Korea. To proxy these global variables, I employ the US yield-curve factors, and I base my discussion and findings on the results from the following past studies on the international linkages of term structures.

For instance, [Diebold et al. \(2008\)](#) investigated the global yield-curve dynamics by extending the dynamic Nelson–Siegel model of [Diebold and Li \(2006\)](#), specifically analyzing the yield curves of four countries: the US, Germany, Japan, and the UK. The authors extracted global yield-curve factors (global level and slope) from these countries' bond yields and found that these global factors accounted for a significant part of the fluctuations in each country's interest rates.

Noting that fluctuations in bond yields are highly correlated across countries, [Jotikasthira et al. \(2015\)](#) analyzed the interactions among different countries' bond yields. They used a maximally flexible model with unspanned macro factors, as suggested by [Joslin et al. \(2014\)](#). For the model factors, Jotikasthira et al. employed macro factors and interest rates with a three-level hierarchical structure of the global, US, and local levels. They considered three countries, the US, the UK, and Germany, and both the policy and risk compensation channels, and their empirical analysis showed that the world inflation and US yield levels together explained over two-

thirds of the covariance of yields at all maturities. Further, these effects operated largely through the risk compensation channel for long-term bonds. I have based the model specification I used in the current paper largely on that of Jotikasthira et al. In particular, to specify a VAR implied by a GDTSM and to identify structural shocks for impulse response functions, I refer to their block structure and variable transformation scheme.

Abbritti et al. (2018) introduced unspanned global factors within a VAR framework using three principal components (global level, slope, and curvature) estimated from yield-curve data for eight countries: Australia, Canada, Germany, Japan, New Zealand, Switzerland, the UK, and the United States. They found that the global yield-curve factors accounted for more than 80% of term premia in the advanced economies and that local term structure factors were more relevant for the short-run dynamics of interest rates; noteworthy, the curvature of the global term structure had a large effect on the interest rate. In separate robustness checks, Abbritti et al. obtained similar results when they replaced their global factors with the US yield-curve factors.

The aforementioned studies all investigated the effects of global factors on advanced economies' term structures. With the current paper, I attempt to answer the question of to what extent global factors affect the term structure of an emerging market economy.

3. Model

3.1. Gaussian dynamic term structure model

In this section, I describe the GDTSM used throughout this paper. GDTSM is a discrete-time essentially affine model with homoskedastic disturbance terms but heteroskedastic stochastic discount factor (SDF). After Ang and Piazzesi (2003) pioneered the model, many term structure studies (to name a few, Cochrane and Piazzesi 2005, Joslin et al. 2011, Wright 2011, Joslin et al. 2014, and Jotikasthira et al. 2015) used GDTSMs to address many important macro-finance topics, such as the relationship between central bank monetary policy and the interest rate term structure.

If the price of an n -period zero-coupon bond at time t is denoted as $P_t^{(n)}$, then the corresponding yield can be defined as $y_t^{(n)} \equiv -\log(P_t^{(n)})/n$. By defining M_{t+1} as a one-period SDF, we can express the bond price as

$$P_t^{(n)} = E_t \left(M_{t+1} P_{t+1}^{(n-1)} \right). \tag{1}$$

We now assume that a $m \times 1$ state vector X_t follows a VAR process:

$$X_{t+1} = \mu + \Phi X_t + \Sigma \varepsilon_{t+1} \quad (\varepsilon_{t+1} \sim \text{i.i.d. } \mathcal{N}(0, I)), \tag{2}$$

this is the law of motion under the physical measure (or \mathcal{P} -measure). The time-invariant Σ in (2) indicates homoskedastic disturbance terms.

Next, the short-term (one-period) interest rate $y_t^{(1)}$ is specified as an affine function of X_t :

$$y_t^{(1)} = \delta_0 + \delta_1' X_t, \tag{3}$$

where δ_0 is a scalar and δ_1 is a $m \times 1$ vector.

Third, SDF is conditionally lognormal as in (4):

$$M_{t+1} = \exp \left(-y_t^{(1)} - \frac{1}{2} \lambda_t' \lambda_t - \lambda_t' \varepsilon_{t+1} \right), \tag{4}$$

where the market price of risks λ_t is specified as an affine function $\lambda_t \equiv \lambda_0 + \lambda_1 X_t$ (λ_0 is a $m \times 1$ vector and λ_1 is a $m \times m$ matrix.).

The above equations (1), (2), (3), and (4) together build a GDTSM. Given this specification, an n -period zero-coupon yield at time t also has an affine form:

$$y_t^{(n)} = -\frac{A_n}{n} - \frac{B_n'}{n} X_t, \tag{5}$$

where A_n is a scalar and B_n is a $m \times 1$ vector that satisfies the following recursions:

$$A_{n+1} = -\delta_0 + A_n + B_n' (\mu - \Sigma \lambda_0) + \frac{1}{2} B_n' \Sigma \Sigma' B_n, \tag{6}$$

$$B_{n+1} = (\Phi - \Sigma \lambda_1)' B_n - \delta_1,$$

starting from $A_1 = -\delta_0$ and $B_1 = -\delta_1$.

The bond price above is the same as if investors were risk-neutral (i.e., $\lambda_0 = 0$ and $\lambda_1 = 0$), but the state vector X_t followed a distorted VAR process:

$$X_{t+1} = \mu^* + \Phi^* X_t + \Sigma \varepsilon_{t+1}^*, \tag{7}$$

where $\mu^* \equiv \mu - \Sigma \lambda_0$ and $\Phi^* \equiv \Phi - \Sigma \lambda_1$.

Laws of motion (2) and (7) are known as the VAR processes under the physical and risk-neutral measures (or \mathcal{P} - and \mathcal{Q} -measures), respectively.

As I have mentioned, I employ six state variables for GDTSM in this paper; here, I clarify their notations. The yield-curve factors (the first three principal components of the yield covariance matrix) for the US (denoted by US) and Korea (denoted by KR) are denoted by $P_t^{US} \equiv (l_t^{US}, s_t^{US}, c_t^{US})'$ and $P_t^{KR} \equiv (l_t^{KR}, s_t^{KR}, c_t^{KR})'$, respectively. As is well known in the literature, each element of the yield-curve vector represents the level, slope, and curvature. The resulting state-variable vector for the GDTSM can be written as $X_t \equiv (P_t^{US}, P_t^{KR})'$ (6-element vector).

To specify the \mathcal{P} -measure VAR in (2), based on Jotikasthira et al. (2015), I employ a two-block structure with the US and Korean blocks. I set the US block (corresponding to P_t^{US}), which is intended to represent the global economy, exogenously to have a unilateral influence on the Korean block (corresponding to P_t^{KR}). Hence, the coefficient matrix Φ in (2) is set to have the following block structure:

$$\Phi = \begin{pmatrix} \Phi_{11} & 0 \\ \Phi_{21} & \Phi_{22} \end{pmatrix} \quad (8)$$

$\begin{matrix} [3 \times 3] & [3 \times 3] \\ [3 \times 3] & [3 \times 3] \end{matrix}$

In addition, for the \mathcal{Q} -measure VAR in (7), I impose some restrictions because of the existence of *unspanned* factors. I discuss this in the next subsection.

3.2. Unspanned factors

Ludvigson and Ng (2009), Duffee (2011), and Joslin et al. (2014), among others, argue that some macroeconomic variables, such as output and inflation, have predictive power for excess bond returns beyond the yield curve information. These are referred to as unspanned factors in the related literature. In addition to the unspanned macro factors, Yun (2019) also documented the possibility of unspanned global liquidity factors in the KTBs.

Table 1 reports the regression results using one-year excess bond returns (for each of two- through five-year-maturity bonds) as dependent variables and as independent variables, combinations of KPC (three yield-curve factors in Korea), USPC (three yield-curve factors in the US), KM (two macro variables in Korea, year-on-year industrial production [IP] and consumer price index [CPI] growth), and USM (two macro variables in the US, also year-on-year IP and CPI growth).

As a preliminary analysis, I regressed the KTB yields on the independent variables in Table 1 (not reported here) and found that the KTB yield-curve factors (KPC) had already explained 99% cross-sectional variations in bond yields (measured as adj. R^2). All adj. R^2 exceeded 99% for the remaining combinations of the independent variables. This result implies that the US yield-curve factors or macro variables do not have additional explanatory power beyond the KTB factors in explaining cross-sectional variations in the KTB yields.

Now I turn to the regression results in Table 1. Different maturity produces similar results, so focus on the five-year bond. First, KPC shows adj. R^2 of 28.4% for one-year excess bond returns. When the US yield-curve factors (USPC) are used along with KPC, the predictive power significantly rises to 43.9%. This is higher than the predictive power (29.2%) when KM is used along with KPC. Adding either KM or USM to the combination of KPC and USPC slightly increases the explanatory power (adj. R^2 of 0.484 and 0.450, respectively). When all variables, that is, KPC, USPC, KM, and USM, are jointly used as independent variables, the predictive power is the highest at 54.5%.

The above results provided a basis for determining the number of factors for GDTSM in this paper. To model the term structure in Korea, I set only the US yield-curve factors as unspanned factors. Although Table 1 shows that the highest predictive power arises when all the variables are included as the predictive variables, including all these variables into the GDTSM produces a 10-factor model. However, given the relatively short sample period, 10 factors might be too many. Accordingly, I only considered the US yield-curve factors, which showed the most impressive contribution to predicting the excess bond returns in Table 1, as the unspanned factors. However, for the robustness check, I will later discuss the analysis results of the 10-factor model including the US and Korean macro variables.²

Because of the two-block structure, I model the US bond yields as a three-factor GDTSM with no unspanned factors. Thus, the US yield-curve factors in P_t^{US} are the only state-variables for both the \mathcal{P} - and \mathcal{Q} -measure VARs for the US term structure.

In contrast, the KTB yields are modeled as a six-factor GDTSM with the US yield-curve factors as unspanned factors. Thus, for the \mathcal{P} -measure VAR, the yield-curve factors for the US and Korea are its six state variables. As noted, the unspanned factors are important for predicting future excess bond returns, but they have little incremental contribution to explaining the current cross-section of bond yields beyond the yield-curve factors. Therefore, for the \mathcal{Q} -measure dynamics, which must fit cross-sectional KTB yield variations, are defined as the dynamics of P_t^{KR} only. In terms of the GDTSM specification, unspanned factors show up only in the VAR specification under the \mathcal{P} -measure VAR in (2). This restriction can work because changes in an unspanned factor affect both term premia and expected future short-term interest rates, but in opposite directions; thereby, the two effects cancel each other out, leaving the current interest rates unchanged (e.g., Wright 2011).

² The first draft of this paper was analyzed with a 10-factor GDTSM. I thank the referee for suggesting a more parsimonious six-factor model.

Table 1
Adjusted R^2 for the regressions for one-year excess returns on the KTBs.

Bond Maturity	KPC	KPC + USPC	KPC + KM	KPC + USPC + KM	KPC + USPC + USM	KPC + USPC + KM + USM
2 years	0.321	0.497	0.332	0.505	0.543	0.568
3 years	0.281	0.468	0.276	0.486	0.470	0.533
4 years	0.326	0.494	0.322	0.516	0.494	0.563
5 years	0.284	0.439	0.292	0.484	0.450	0.545

Note: This table reports the Adjusted R^2 obtained from regressing one-year excess returns on the two- through five-year KTBs on selected combinations among KPC (three principal components from KTB yields), USPC (three principal components from US bond yields), KM (year-on-year IP and CPI growth in Korea), and USM (year-on-year IP and CPI growth in the US). The sample period is 2001:1–2020:12.

3.3. Identifying structural shocks

To identify structural shocks in the impulse response analysis below, I employ [Jotikasthira et al.'s \(2015\)](#) variable transformation, which yields an observationally equivalent VAR representation that permits a better economic interpretation of structural shocks. I regress each of the KTB yield-curve factors I_t^{KR}, S_t^{KR} , and C_t^{KR} on P_t^{US} (with an intercept) and form a residual vector $P_t^{KR,e} \equiv (I_t^{KR,e}, S_t^{KR,e}, C_t^{KR,e})'$ by taking residuals from each regression. The resulting $P_t^{KR,e}$, can be interpreted as the Korea-specific local yield-curve factor. We now have $P_t^{US} \equiv (I_t^{US}, S_t^{US}, C_t^{US})'$ and $P_t^{KR,e} \equiv (I_t^{KR,e}, S_t^{KR,e}, C_t^{KR,e})'$ and define the transformed state-variable vector $X_t^e \equiv (P_t^{US}, P_t^{KR,e})'$. The \mathcal{P} -measure VAR process for X_t^e can be written as

$$X_t^e = \mu^e + \Phi^e X_{t-1}^e + \Sigma^e e_t^e, \tag{9}$$

where $e_t^e \sim \text{i.i.d. } \mathcal{N}(0, I)$ and Σ^e is a lower triangular matrix. To distinguish the parameters in VAR (9) from those in VAR (2), I use the superscript e for VAR (9). The two-block structure in (8) is still preserved for Φ^e , so the US state variables in X_t^e cannot be, by construction, influenced by any Korea-specific local structural shocks.

After fitting X_t^e to the VAR in (9), I identify structural shocks via Cholesky factorization, so Σ^e in (9) is lower triangular. I can now implement the usual VAR analysis: impulse response analysis and forecasting error variance decomposition.

The original state-variable vector $X_t \equiv (P_t^{US}, P_t^{KR})'$ and the transformed state-variable vector $X_t^e \equiv (P_t^{US}, P_t^{KR,e})'$ have a one-to-one mapping, because the i th element in X_t (denoted by $X_t^{(i)}$) is given by

$$X_t^{(i)} = c^{(i)} + T^{(i)'} X_t^e, \tag{10}$$

where each of $c^{(i)}$ (scalar) and $T^{(i)}$ (6×1) can be obtained by regressing $X_t^{(i)}$ onto the intercept and X_t^e . It is now possible to define the matrices C (6×1) and T (6×6) by stacking $c^{(i)}$ and $T^{(i)'}$, respectively, across $i = 1, \dots, 6$. This gives

$$X_t = C + T X_t^e. \tag{11}$$

With this transformation, it is now possible to easily switch between X_t and X_t^e . By plugging (11) into (2), transformation equations between the VARs (2) and (9) can also be derived as $\mu^e = T^{-1}(\mu - C + \Phi C)$ and $\Phi^e = T^{-1} \Phi T$. Moreover, Σ^e in (9) is a lower triangular matrix from a Cholesky decomposition of $T^{-1} \Sigma T^{-1}$.

3.4. Estimation method

To estimate the GDTSM, I first fit X_t^e to the aforementioned two-block VAR to obtain the \mathcal{P} -measure parameters in (9): μ^e, Φ^e , and Σ^e . Usual least squares produce the ML estimates.

However, as [Bauer et al. \(2012\)](#) highlighted, because bond yields are highly persistent, serious small-sample bias may lead to underestimating their persistence; model estimates in turn could imply too-fast mean reversion. As a result, the path of future short-term interest rates projected by a biased model becomes too smooth, and model-implied term premium estimates are also distorted to be too volatile. Accordingly, [Bauer et al. \(2012\)](#) recommended a bias correction in the \mathcal{P} -measure dynamics of GDTSM, and I here also perform bias correction to avoid a small-sample bias. In the following, the BC estimates denote bias-corrected estimates. To assess the effect of bias correction, I also provide the results of the OLS estimates, the estimates without bias correction.

As noted, I correct small-sample bias using bootstrap-after-bootstrap ([Kilian 1998](#)),³ where I adjust bias using the first round of bootstrap and obtain confidence intervals (CIs) or standard errors for impulse responses using the second round of bootstrap.

³ Alternatively, the indirect inference bias correction method [Bauer et al. \(2012\)](#) propose can also be used. [Abbritti et al. \(2018\)](#) argue that the difference between the two methods is not significant.

Meanwhile, in the process of adjusting bias to estimate either parameters or CIs, bias-corrected estimates can be pushed outside the stationary region. To prevent this, I also employ Kilian's stationary correction.⁴ After correcting small-sample bias via bootstrap-after-bootstrap, I apply the transformation equations in the previous subsection, which allows for recovering the parameters for the \mathcal{P} -measure VAR in (2).

The remaining task is to estimate the risk-neutral parameters for the \mathcal{C} -measure VAR in (7). To estimate the \mathcal{C} -measure parameters via MLE, I employ a maximally flexible model (Joslin et al. 2011). As noted earlier, Joslin et al. (2011) present an econometric methodology that can efficiently estimate GDTSMs when all state variables are observable. This efficiency is obtained because it becomes possible to separate the log-likelihood function into physical and risk-neutral components. Joslin et al. (2014) extend the method of Joslin et al. (2011) to term structure models with unspanned factors.

I consider the decomposition of the (observed) state variable vector X_t ($m \times 1$) into the (spanned) yield-curve vector X_t^S ($m_S \times 1$) and the *unspanned* component vector X_t^U ($m_U \times 1$ and $m_S + m_U = m$). In the present case, $X_t^S = P_t^{KR}$ with $X_t^U = P_t^{US}$ gives the term structure in Korea, and the US term structure is given by $X_t^S = P_t^{US}$ with no X_t^U . Note that for the \mathcal{C} -measure dynamics, X_t^S is the only relevant state variables.

The maximally flexible canonical GDTSM model (Joslin et al. 2011) is constructed with a latent state vector Z_t that follows a VAR with zero intercept and diagonal coefficient matrix $\Lambda^Q \equiv \text{diag}[\lambda_1^Q, \lambda_2^Q, \dots, \lambda_{m_S}^Q]$, and with short-term interest rates in the form $r_t = r_\infty^* + i'Z_t$ where i is a vector of ones. In this canonical form, r_∞^* is interpreted as the risk-neutral long-run mean of the short rate, and the diagonal elements in Λ^Q are the eigenvalues of the $m_S \times m_S$ block of Φ^* (corresponding to X_t^S) in (7). The risk-neutral part of the likelihood is maximized with respect to the risk-neutral parameters $\{r_\infty^*, \Lambda^Q\}$. Joslin et al. (2011) suggest a mapping from $\{r_\infty^*, \Lambda^Q\}$ to the original parameters for the original \mathcal{C} -measure VAR in (7).

I employ the estimate of Σ (obtained after the bias correction) as an initial value to maximize the risk-neutral part of the likelihood function with respect to $\{r_\infty^*, \Lambda^Q, \Sigma\}$. Using the mapping suggested by Joslin et al. (2011) gives the original \mathcal{C} -measure parameters (i.e., μ^*, Φ^*, δ_0 , and δ_1 in (3) and (7)).

4. Empirical analysis

4.1. Data

For this study, I have used the zero-coupon bond yield data on the KTBs (sample period: 2001:1–2020:12) provided by Korea Asset Pricing. Although long-term zero-coupon bonds are not issued in Korea, Korea Asset Pricing releases the continuously compounded zero-coupon yield curves estimated using the bootstrap method (Fama and Bliss 1987) using the coupon-bond price data traded in the market. I consider 12 bond maturities: 3, 6, 9, 12, 18, 24, 30, 36, 48, 60, 84, and 120 months.

For the US Treasury bond yields, I use the database of Gürkaynak et al. (2007) that is updated periodically on the homepage of the US Federal Reserve.⁵ I selected the same US Treasury bond maturities as those on KTBs. Fig. 1 illustrates the US (panel (a)) and Korean (panel (b)) bond yields at selected maturities over the sample period.

Over the latter half of the sample, the zero lower bound (ZLB) period, in which the US short-term interest rate approached zero, lasted for a long time. Although there was no episode similar to the ZLB in Korea, the interest rates fell constantly throughout the sample. Broadly, it appears that the bond yields of the US and Korea share a highly persistent common trend. Exceptionally, when the Fed began tapering after 2015, the short-term rates in both countries moved in the opposite direction for a few years (i.e., the US rate rose, and Korea's rate fell slightly).

4.2. Preliminary analysis

Table 2 reports the correlations of macro variables (year-on-year IP and CPI growth), bond yields, and one-year excess bond returns between the US and Korea. I estimated these correlations across three sample periods (2001:1–2020:12, 2001:1–2010:12, and 2011:1–2020:12). For the full sample (2001:1–2020:12), the magnitudes of correlations between the bond yields (0.58–0.90) are larger than those between the macro variables (around 0.5); in particular, the correlation for the 10-year yields even reaches around 0.90. However, the correlations between the short-term yields differ depending on the chosen sample period. To be specific, although the three-month yields show a correlation of 0.63 for the first-half sample (2001:1–2010:12), it changes to negative (– 0.37) for the latter half (2011:1–2020:12). Meanwhile, the 10-year yields continue to show large correlations across the two subsamples (0.64 and 0.61, respectively).

Table 2 also shows large correlations between one-year excess bond returns for the US and Korea; these are computed by subtracting one-year zero-coupon yield at the time of initial investment from a one-year change in the log price of a bond at a given maturity (either two or five-years). These correlations are all greater than 0.5 for each sample. Such large correlations between the bond excess returns imply the possibility of comovements between the risk premia embedded in bond yields.

As noted above, the first three principal components estimated from the bond yield data of the US and Korea are used as the yield-

⁴ For a detailed explanation, see Chapter 12 of Kilian and Lütkepohl (2017).

⁵ <https://www.federalreserve.gov/pubs/feds/2006>.

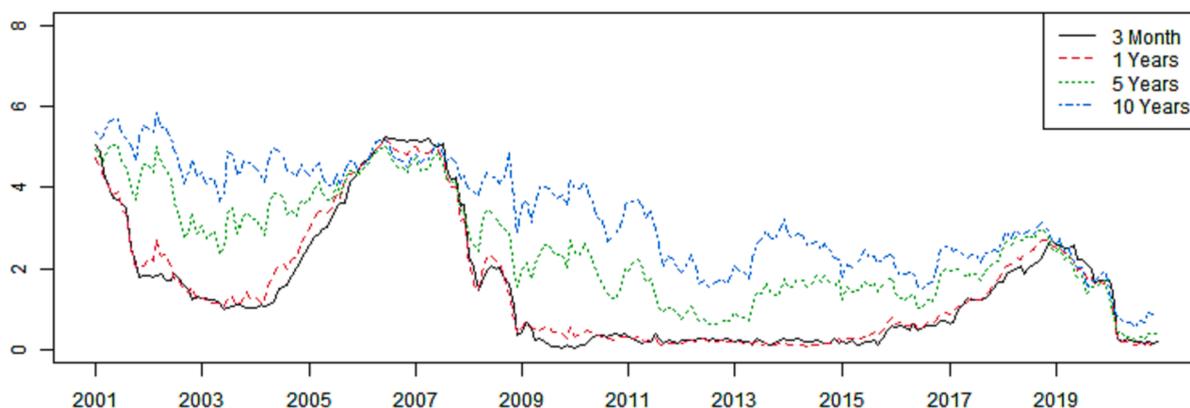
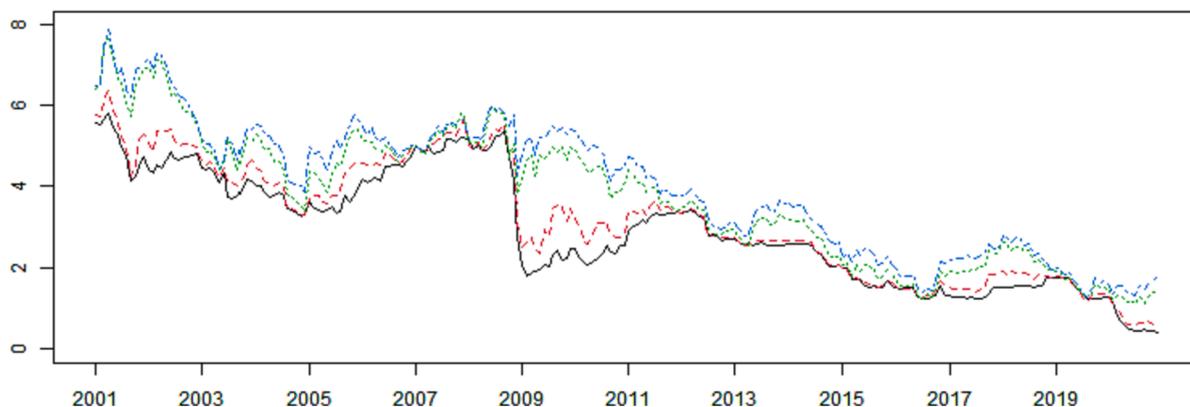
(a) US Treasury Bond Yields**(b) Korea Treasury Bond Yields**

Fig. 1. The zero-coupon Treasury bond yields across selected maturities for the US (panel A) and Korea (panel B) from 2001:1 to 2020:12.

Table 2

Correlations between selected macro variables and bond yields for the US and Korea.

Sample period	IP growth	CPI growth	Treasury bond yields				One-year-ahead excess bond returns	
			3 months	1 year	5 years	10 years	2 years	5 years
2001:1–2020:12	0.481	0.525	0.580	0.574	0.758	0.902	0.546	0.582
2001:1–2010:12	0.657	0.341	0.633	0.577	0.508	0.641	0.510	0.547
2011:1–2020:12	0.300	0.728	−0.373	−0.355	0.062	0.610	0.573	0.657

Note: This table reports the US–Korea correlations for year-on-year IP and CPI growth, Treasury bond yields, and one-year-ahead excess returns on the two- and five-year bonds. The sample periods considered are i) 2001:1–2020:12, ii) 2001:1–2010:12, and iii) 2011:1–2020:12.

curve factors in the GDTSM; specifically, they are usually referred to as the *level*, *slope*, and *curvature*. Several researchers, such as Mishkin (1990a, b), Barr and Campbell (1997), Dewachter and Lyrio (2006), and Rudebusch and Wu (2007), interpreted the *level* as expectations on long-term inflation, and Estrella and Mishkin (1998), Hamilton and Kim (2002), Ang et al. (2006), and Rudebusch and Wu (2008) linked the *slope* to business cycles and monetary policy. In a similar vein, Bekaert et al. (2010) found that as a result of variance decomposition of the US term structure under the New Keynesian framework, the inflation target shock dominates the variations in the level factor whereas monetary policy shocks dominate the variations in the slope and curvature.

In contrast to the level and slope factors, there is little consensus on the implications of the curvature. Dewachter and Lyrio (2006) document that the curvature effect is related to the current stance of monetary policy, that is, to real interest rate movements unrelated to standard macroeconomic conditions.

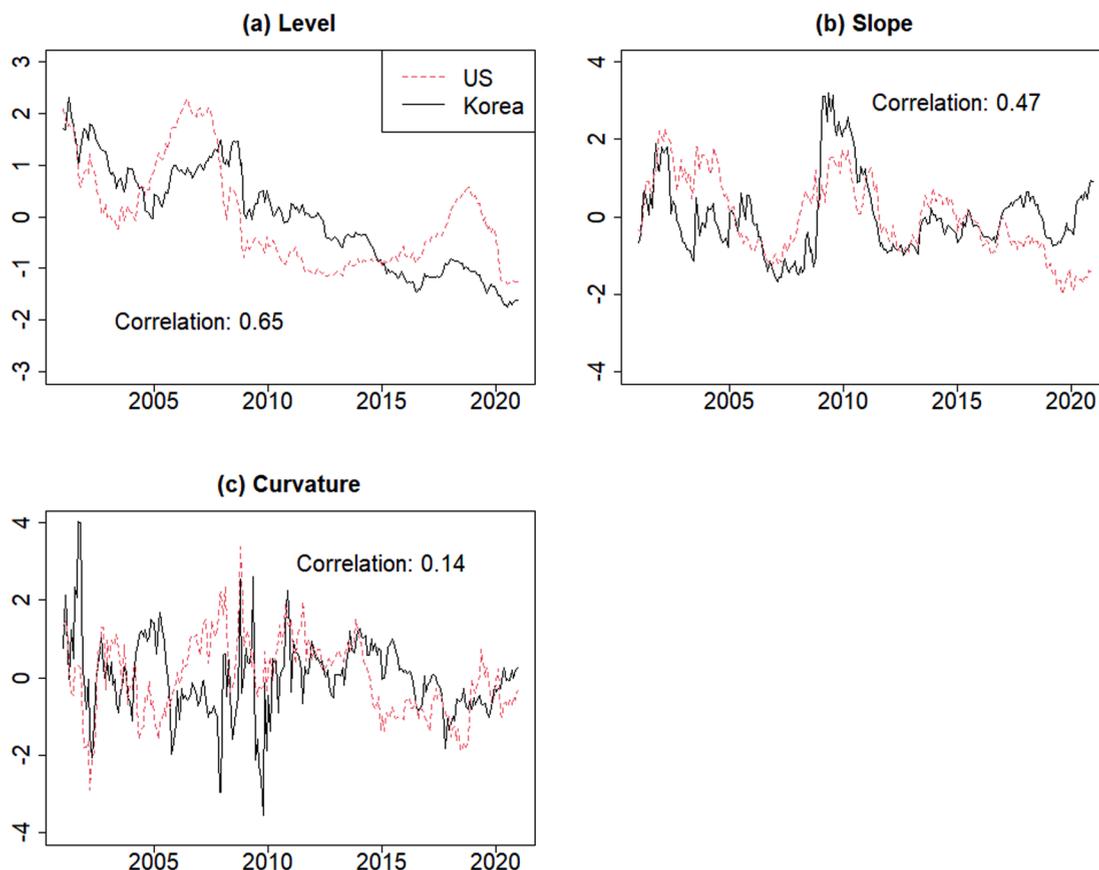


Fig. 2. Time-series of the yield-curve factors in the US and Korea. The factors are estimated as the first three principal components, that is, (a) level, (b) slope, and (c) curvature, from bond yield data for each country. Each figure also provides a correlation coefficient for the corresponding yield-curve factor.

The close linkage between the US and Korean Treasury yields is also clearly illustrated in Fig. 2, which shows the level, slope, and curvature factors over time along with the correlations for each factor. The level factors for the US and Korean Treasury yields (panel (a)) show a common downward trend during the sample period (with a correlation of 0.65). The slope factors shown in panel (b) show a correlation coefficient of about 0.47; here, an increase in the slope means a widening term spread. Finally, the curvatures have the smallest correlation coefficient among the yield-curve factors (about 0.14).

4.3. Impulse responses

Using the econometric method detailed in Section 3, I estimate the \mathcal{P} - and \mathcal{Q} -parameters for the GDTSM. As mentioned earlier, to correct the small-sample bias due to persistent bond yields, I employ Kilian's (1998) bootstrap-after-bootstrap method (i.e., the BC estimates). For comparison, I also provide the OLS estimates.

To gain insight into the magnitude of the small-sample bias, I compute the maximum eigenvalues for the VARs implied by the BC and OLS estimates. The largest eigenvalue in absolute value for the \mathcal{Q} -measure VAR (i.e., Φ^*) is estimated to be 0.998, and those for the \mathcal{P} -measure VARs (i.e., Φ) implied from the BC and OLS estimates are 0.999 and 0.975, respectively, where the former is very close to the unit root. In short, the difference between the BC and OLS estimates is significant, and therefore, bias correction could substantially affect the impulse response estimations.

Panel A of Table 3 reports the \mathcal{Q} -parameter estimates for the US and Korean GDTSM blocks; I used these estimates to fit the two countries' bond yield data. Panel B shows that the resulting yield pricing errors (measured by the root mean squared errors) for the US and Korean Treasury yields⁶ (3.77 bp and 5.26 bp for the US and Korea, respectively) are comparable with the errors from past studies, implying that the GDTSM fits the term structures well in both countries.

Using the \mathcal{P} - and \mathcal{Q} -parameter estimates, I estimate the impulse responses for the GDTSM-implied VAR (implied from each of the BC and OLS estimates). The online Appendix Section A describes how to derive the impulse responses in detail. As noted, the six state

⁶ Because there is little difference in the yield pricing errors between the BC and OLS estimates, Table 3 reports only the BC estimates.

Table 3
Risk-Neutral Parameter Estimation Results.

A. Risk-Neutral Parameter Estimates											
US					Korea						
λ_1^Q	λ_2^Q	λ_3^Q	r_∞^*	-	λ_1^Q	λ_2^Q	λ_3^Q	r_∞^*			
0.998 (0.000)	0.961 (0.007)	0.951 (0.008)	0.104 (0.012)		0.998 (0.000)	0.937 (0.022)	0.927 (0.025)	0.095 (0.007)			
B. Yield Pricing Errors (bp)											
US						Korea					
3 months	1 year	3 years	5 years	10 years	Total	3 months	1 year	3 years	5 years	10 years	Total
7.14	4.25	1.84	4.23	4.98	3.77	5.66	4.31	7.03	6.42	6.16	5.26

Note: This table reports the risk-neutral parameter estimates for the US and Korea (panel A) and the root mean squared errors for each country's yield pricing errors (in terms of a basis point) for the selected maturities (panel B). The numbers in parentheses in panel A are standard errors. The estimation sample period is 2001:1–2020:12.

variables for the VAR are the three US yield-curve factors (l_t^{US} , s_t^{US} , and c_t^{US}) and three KTB yield-curve factors (l_t^{KR} , s_t^{KR} , and c_t^{KR}). The lag order of the VAR is set to one, based on the Bayesian information criterion (BIC). As explained in Section 3, I identify the structural shocks for the IRF analysis through the VAR of the transformed state variables in (9). The six structural shocks in this IRF analysis will be considered to be from l_t^{US} , s_t^{US} , c_t^{US} , $l_t^{KR,e}$, $s_t^{KR,e}$, and $c_t^{KR,e}$. As a robustness check, I will also discuss the extended 10-factor GDTSM including the US and Korean macro variables in Section 4.8.1.

Given the GDTSM-implied VAR, Figs. 3 and 4 report the IRFs of the one- and ten-year KTB yields in response to each structural shock (one standard deviation). For each figure, the dotted lines represent the 90% (bootstrapped) CI, and along with the impulse response of a bond yield, the impulse responses of the expectation (policy channel) and term premium (risk compensation channel) components are also provided. To save space, the IRF results for the US bond yields are reported in the online Appendix Section B.

First, Fig. 3 illustrates the impulse responses of the one-year KTB yield, indicating the largest and most persistent effect for the US level shock. Because the US level shock reflects global long-term inflation expectations, the result of this impulse response can be seen as reflecting the global synchronization of inflation. Next, the US slope shock also has a large and persistent effect on this short-term interest rate. As mentioned earlier, the slope shock is often interpreted as a business cycle or monetary policy shock. As can be expected, the size and persistence of these US level and slope shocks are more prominent in the BC estimates than in the OLS estimates.

In contrast, impulse responses to Korean local shocks show similar results between the BC and OLS estimates. Among the Korean shocks, local level shock ($l_t^{KR,e}$) appears to have a sizeable positive influence on impact. However, in contrast to the US shock, this local shock temporarily moves the bond yield and is quickly dissipated.

Next, Fig. 4 presents the impulse responses of the 10-year KTB yield. The impacts of the US and Korean yield-factor shocks are all significant together, although their size and persistence differ. In particular, the US yield-curve shocks (l_t^{US} , s_t^{US} , and c_t^{US} shocks) all have statistically significant and persistent effects that appear in both the BS and OLS estimates. As in the case of the one-year yield, the persistence of these shocks is greater for the BC estimates than for the OLS estimates.

Interestingly, the curvature shock (i.e., the c_t^{US} shock) has a substantial effect on the 10-year yield. In fact, many authors (Fama and Bliss 1987, Campbell and Shiller 1991, Bauer 2018) emphasized the influence of the slope factor on the term premium and have tended to recognize the curvature factor as less important. In contrast, Abbritti et al. (2018) emphasize the importance of the curvature factor in the international linkage of term structures. They claim that the curvature has the highest contribution among the global yield-curve factors in determining the term premium. The analysis results in the current paper support Abbritti et al. (2018).⁷

In addition to the US yield-curve shocks that represent global yield-curve shocks, the local yield curve shocks (e.g., $l_t^{KR,e}$, $s_t^{KR,e}$, and $c_t^{KR,e}$ shocks) have a significant effect on the 10-year yield. However, unlike the US shocks, the impact of these local yield-curve shocks appears to have a large contemporaneous impact, but rapidly dissipates.

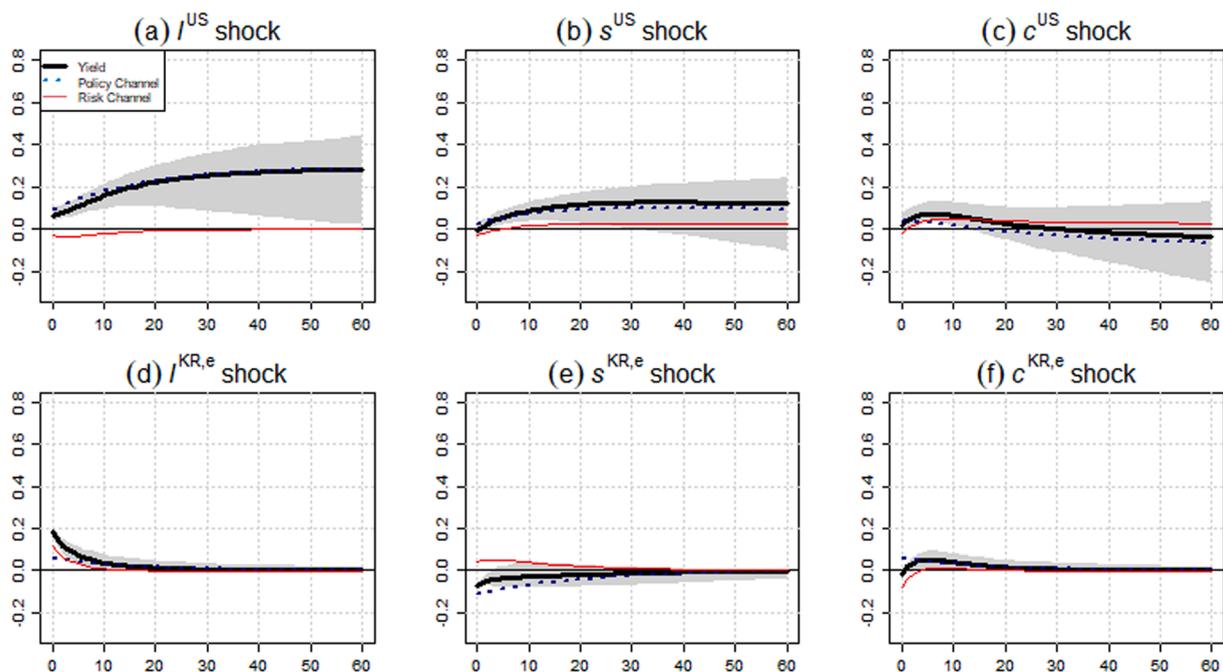
Figs. 3 and 4 also provide the impulse responses for the policy and risk compensation channels (that is, the expectation and term premium components, respectively).⁸ As illustrated by Fig. 3, for the one-year yield, structural shocks appear to be mainly transmitted to the bond yield through the policy channel. This result occurs for both BC and OLS estimates, and is consistent with the results of Jotikasthira et al. (2015).

Fig. 4 presents the 10-year yield results. Contrary to the case of the one-year yield, the BC and OLS estimates produce different results. There are also important departures from Jotikasthira et al.'s (2015) results. First, in the case of BC estimates, for the US level shock, which makes the most important contribution to the impulse responses of bond yields, the policy channel dominates; for all other shocks, the risk channel is dominant. However, in the OLS estimates, for all shocks, the risk channel seems to dominate the long-

⁷ According to my separate analysis (not reported here), in the impulse responses or forecasting error variance decomposition of the term premium for the 10-year KTB yield, the contribution of the US curvature factor is larger than that of the US slope. These results will be provided upon request.

⁸ See the online Appendix Section A for more details.

A. BC estimates



B. OLS estimates

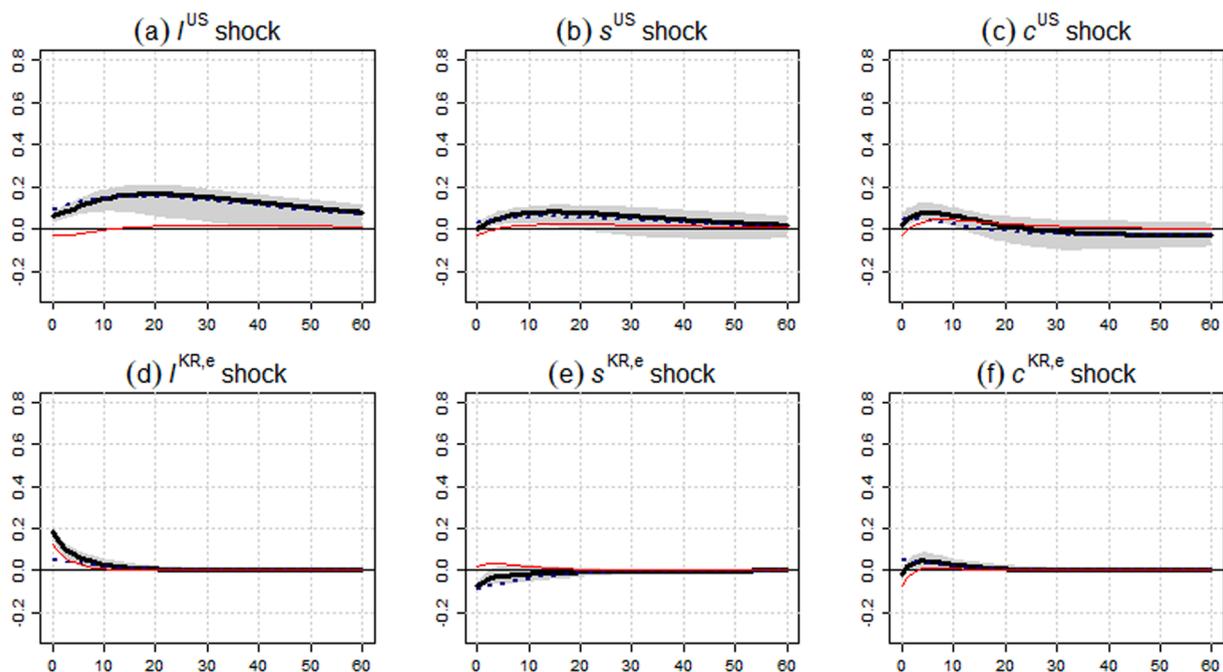
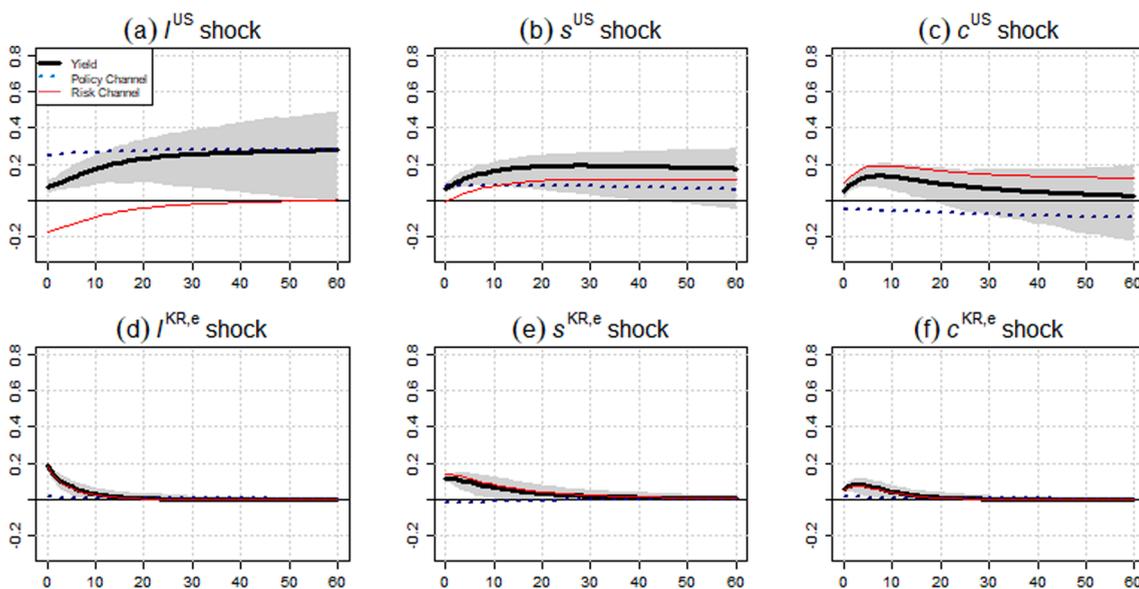


Fig. 3. Impulse responses of the one-year KTB yield implied from the GDTSM. This figure presents the impulse responses of bond yields to a one-standard deviation structural shock with a 90% confidence interval. Both BC (panel A) and OLS (panel B) estimates are provided. For each figure, along with the impulse responses of yields (denoted by Yield), those for the policy and risk compensation channels are also reported.

A. BC estimates



B. OLS estimates

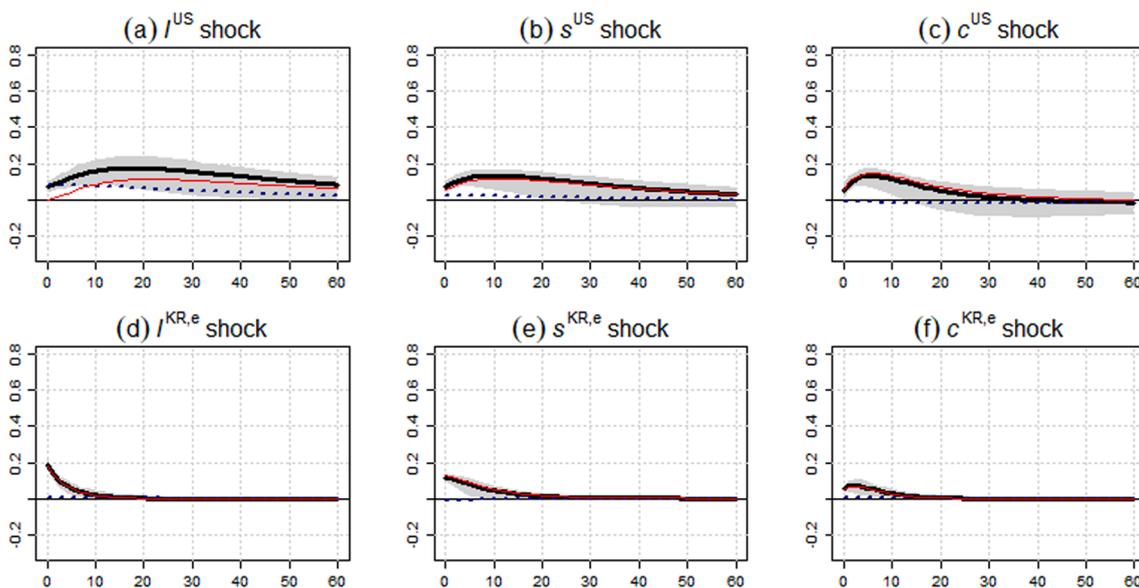


Fig. 4. Impulse responses of the 10-year KTB yield implied from the GDTSM. This figure presents the impulse responses of bond yields to a one-standard deviation structural shock with a 90% confidence interval. Both BC (panel A) and OLS (panel B) estimates are provided. For each figure, along with the impulse responses of yields (denoted by Yield), those for the policy and risk compensation channels are also reported.

term yield. The importance of policy channels in transmitting the US level shock into the 10-year yield is a different result from Jotikasthira et al. (2015); they found that the main transmission channel for long-term yields was the risk channel across all shocks, so they emphasized the risk compensation that global investors require as a critical channel for understanding the determinants of long-term yields across international markets. However, these results could vary depending on whether or not small-sample bias is corrected. As aforementioned, the US level factor represents the long-term expectation of global inflation; thus, it is an important economic fundamental to which a central bank will refer in conducting its monetary policy. Therefore, policy channels are considered to be important even in long-term interest rate fluctuations.

4.4. Forecasting error variance decomposition

To quantify the relative contributions of structural shocks in explaining fluctuations in bond yields, I discuss in this section the results of forecasting error variance decompositions for KTB yields. Table 4 reports those results by bond maturities (one and ten years), forecast horizons (12- and 60-months), and whether or not small-sample bias is corrected (either BC or OLS estimates). To save space, I provide the results for the US Treasury yields in the online Appendix Section B.

For each bond yield, forecasting error variances are also decomposed by the yield’s transmission channels (policy and risk compensation). Let $IRF_{i,h}^{(YLD)}$ denote the impulse response after h periods of a bond yield in response to the i th shock. Similarly, $IRF_{i,h}^{(EXP)}$ and $IRF_{i,h}^{(TP)}$ denote the impulse responses after h periods of the expectation (EXP) and term premium (TP), respectively, in response to the i th shock. As I showed above, $IRF_{i,h}^{(YLD)} = IRF_{i,h}^{(EXP)} + IRF_{i,h}^{(TP)}$. Therefore, the variance of forecasting errors of a bond yield up to period H , that is, $\sum_{h=0}^H [IRF_{i,h}^{(YLD)}]^2$, can be decomposed as

$$\sum_{h=0}^H [IRF_{i,h}^{(YLD)}]^2 = [IRF_{i,h}^{(YLD)} \times IRF_{i,h}^{(EXP)}]^2 + [IRF_{i,h}^{(YLD)} \times IRF_{i,h}^{(TP)}]^2 \tag{12}$$

The first and second components on the right-hand side in (12) can be regarded as the contributions of the policy and risk compensation channels, respectively, to the forecasting error variance of a bond yield.

Table 4 shows that similar to the results in Section 4.3, the relative contributions of structural shocks differ depending on forecast horizons. For the one-year KTB yield, the variance decomposition results are similar between the BC and OLS estimates. To be specific, at the 12-month horizon, the US level shocks (44.3% and 42.6% for BS and OLS, respectively) and Korean local level shocks (23.5% and 24.4% for BS and OLS, respectively) make sizeable contributions. However, at the 60-month horizon, the proportion of US level shocks is overwhelming (76.9% and 72.0% for BS and OLS, respectively). For this short-term yield, the policy channel nearly dominates, regardless of the bias correction. Exceptionally, for Korea’s local level shock, the risk channel is slightly dominant over the policy channel.

For the 10-year yield, nearly all the US and Korea yield-curve factor shocks make significant contributions at the 12-month horizon. For example, the contributions of the US shocks (I_t^{US} , s_t^{US} , and c_t^{US} , shocks) amount to 26.6%, 25.2% and 20.1% (based on the BC estimates), respectively, and the contributions of the Korea local level and slope shocks (I_t^{KR} and s_t^{KR} shocks) amount to 11.5% and 11.4%, respectively (based on the BC estimates); however, at the longer horizon, the US shocks become dominant. Similar to the case of the one-year yield, the US level shock (I_t^{US} shock) makes an overwhelmingly large contribution of 57.0% at the 60-month horizon. In addition, the US slope shock (s_t^{US} shock) also makes a sizeable contribution (31.8%). In contrast to the results from the 12-month horizon, at the 60-month horizon, the contributions of the Korea-specific local yield-curve shocks ($I_t^{KR,e}$, $s_t^{KR,e}$, and $c_t^{KR,e}$ shocks) add up to only around 5%.

I now turn to the variance decompositions for the policy and risk compensation channels (i.e., the expectation and term premium, respectively). As with the previous results, the policy channel is important for one-year maturity, consistent with the results of Jotikasthira et al. (2015).

However, the 10-year yield results in Table 4 tell a different story. Based on the BC estimates, the policy channel dominates for the US level shocks, and the risk channel dominates for the US slope and curvature shocks. These results depart from Jotikasthira et al. (2015), who found that across all shocks, the risk channel dominates for long-term bond yields. In Table 4, the OLS estimates produce results qualitatively similar to those by Jotikasthira et al. (2015): I again confirm that small-sample bias correction can have a significant effect on evaluating how differently the policy and risk channels contribute to the impulse responses of bond yields.

4.5. GDTSM-implied forward term premia

In this section, I discuss the term premia implied by the GDTSM. Following the past studies (e.g., Wright 2011, Bauer, Rudebusch, and Wu 2014, Abbritti et al. 2018, etc.), I focus on the term premia for the five- to ten-year forward rates.

As discussed earlier, the \mathcal{P} -measure dynamics for a GDTSM can be distorted by small-sample bias, which can result in misleading estimates of expected future short-term interest rates and of long-maturity term premia. To evaluate the influence of this bias, I consider both the BC and OLS term premium estimates below.

In Fig. 5, the five- to ten-year forward rate in Korea is decomposed into the expectation (i.e., risk-neutral forward rate) and term premium components based on each of the BC and OLS estimates (panels (a) and (b), respectively). Similar to past studies, the OLS expectation estimate appears much smoother than the BC counterpart because of faster mean reversion; as a result, the OLS term

Table 4
Forecasting error variance decomposition for KTB yields.

A. One-year maturity yield								
Horizon	Yield Comp.	l^{US}	s^{US}	c^{US}	$l^{KR,e}$	$s^{KR,e}$	$c^{KR,e}$	Total
(a) BC Estimates								
12 months	Overall	0.197	0.052	0.048	0.105	0.022	0.022	0.445
	(%)	(44.3)	(11.7)	(10.7)	(23.5)	(4.8)	(5.0)	(100.0)
	Policy channel	0.240	0.046	0.019	0.047	0.044	0.019	0.415
60 months	Risk channel	-0.043	0.006	0.029	0.058	-0.022	0.003	0.031
	Overall	3.341	0.759	0.080	0.109	0.030	0.028	4.347
	(%)	(76.9)	(17.5)	(1.8)	(2.5)	(0.7)	(0.6)	(100.0)
60 months	Policy channel	3.432	0.602	0.053	0.051	0.061	0.024	4.223
	Risk channel	-0.091	0.157	0.027	0.058	-0.031	0.004	0.124
	(b) OLS Estimates							
12 months	Overall	0.169	0.045	0.056	0.097	0.015	0.015	0.397
	(%)	(42.6)	(11.4)	(14.0)	(24.4)	(3.7)	(3.8)	(100.0)
	Policy channel	0.190	0.037	0.027	0.039	0.024	0.011	0.329
60 months	Risk channel	-0.021	0.008	0.028	0.058	-0.010	0.004	0.069
	Overall	1.038	0.190	0.083	0.099	0.016	0.017	1.441
	(%)	(72.0)	(13.1)	(5.8)	(6.8)	(1.1)	(1.2)	(100.0)
60 months	Policy channel	0.972	0.138	0.050	0.040	0.027	0.012	1.239
	Risk channel	0.065	0.051	0.033	0.059	-0.011	0.005	0.202
	B. Ten-year maturity yield							
Horizon	Yield Comp.	l^{US}	s^{US}	c^{US}	$l^{KR,e}$	$s^{KR,e}$	$c^{KR,e}$	Total
(a) BC Estimates								
12 months	Overall	0.240	0.228	0.181	0.104	0.103	0.048	0.903
	(%)	(26.6)	(25.2)	(20.1)	(11.5)	(11.4)	(5.3)	(100.0)
	Policy channel	0.445	0.135	-0.084	0.008	-0.021	0.006	0.489
60 months	Risk channel	-0.204	0.092	0.265	0.095	0.124	0.042	0.414
	Overall	3.359	1.873	0.378	0.104	0.123	0.051	5.888
	(%)	(57.0)	(31.8)	(6.4)	(1.8)	(2.1)	(0.9)	(100.0)
60 months	Policy channel	3.817	0.774	-0.302	0.009	-0.027	0.007	4.278
	Risk channel	-0.458	1.098	0.681	0.096	0.149	0.044	1.610
	(b) OLS Estimates							
12 months	Overall	0.213	0.183	0.164	0.095	0.077	0.034	0.766
	(%)	(27.8)	(23.9)	(21.4)	(12.4)	(10.1)	(4.5)	(100.0)
	Policy channel	0.122	0.029	-0.022	0.005	-0.008	0.003	0.129
60 months	Risk channel	0.091	0.154	0.186	0.090	0.086	0.031	0.637
	Overall	1.137	0.507	0.217	0.095	0.082	0.035	2.073
	(%)	(54.9)	(24.5)	(10.5)	(4.6)	(3.9)	(1.7)	(100.0)
60 months	Policy channel	0.426	0.061	-0.035	0.005	-0.009	0.003	0.451
	Risk channel	0.711	0.446	0.252	0.090	0.091	0.032	1.622

Note: This table reports the results of forecasting error variance decompositions of the one- and ten-year KTB yields for the main six-factor GDTSM. For both 12- and 60-month horizons, the forecasting error variances are computed for each of the overall, policy, and risk channels. The sample period is 2001:1–2020:12.

premium shows almost the same trend as the forward rate. In contrast, both the BC expectation and term premium components display much more variation.

Next, Fig. 6 displays the BC and OLS term premium estimates for the five- to ten-year forward rates in the US and Korea; in each panel, the shading denotes National Bureau of Economic Research-dated recessions. Regardless of the bias correction, the term premia for both countries show very high correlations of around 0.90. Both figures imply countercyclical term premia with respect to the US business cycle. Among the two, the countercyclicality of the BC term premia appears more evident than that of the OLS term premia. In particular, the rise in the BC term premia is more pronounced around the Lehman bankruptcy in 2007.

To further analyze the relationship between the bias correction and countercyclicality of term premia, I run regressions of the term premia on monthly output gaps (Table 5); I employ the detrended log of IP index for the gaps. I use both the quadratic (denoted by Quadratic) and Hodrick and Prescott (1997) filtered (denoted by HP) trends to remove the trend from the log of IP. In addition, I consider year-on-year IP growth as an independent variable in the regressions. The numbers in parentheses in Table 5 are the heteroskedasticity and autocorrelation consistent (HAC) standard errors based on Andrews (1991).

First, I discuss the regression results for the US term premium estimates. Table 5 shows that the regressions of the BC term premium in the US on the US output gap measures (either Quadratic or HP) produce statistically significant negative coefficients, implying a significant countercyclicality of the US term premium. When the IP growth is used as an independent variable, the regression coefficient is still negative, although not significant. However, when the OLS term premium is used instead, the regression coefficients are no more statistically significant across all cases. When the term premium is regressed on IP growth, the regression coefficient even becomes positive. Thus, the bias correction makes a more marked countercyclicality of the term premium. The countercyclical term premia have been documented by Campbell and Cochrane (1999), Wachter (2006), Cochrane and Piazzesi (2005), and Ludvigson and

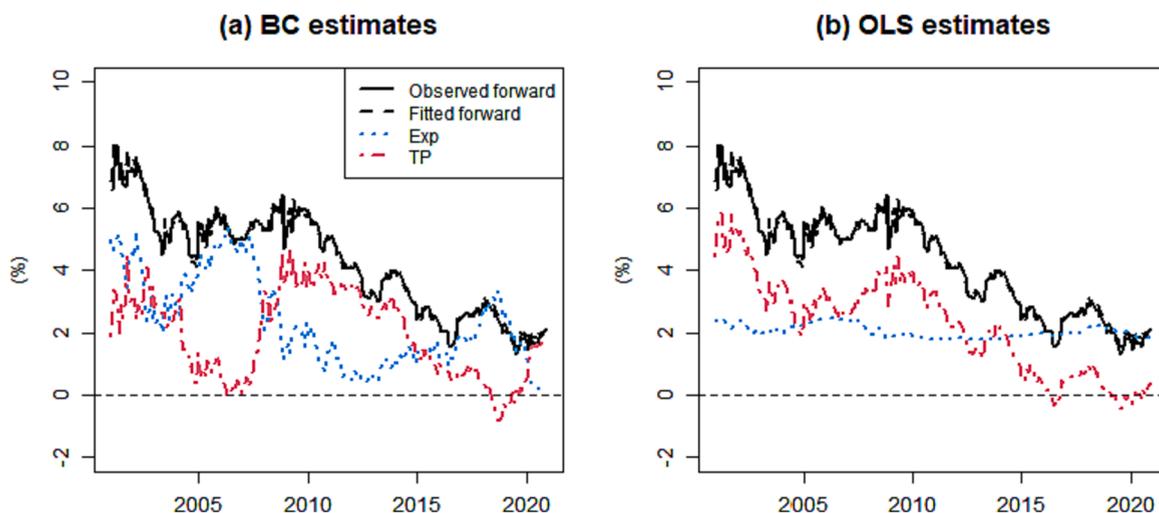


Fig. 5. The GDTSM-implied forward rate, expectation, and term premium for KTB. This figure presents the five- to ten-year forward rates (actual and fitted ones), its expectation (risk-neutral forward rate), and forward term premium components. Both BC (panel A) and OLS (panel B) estimates are provided.

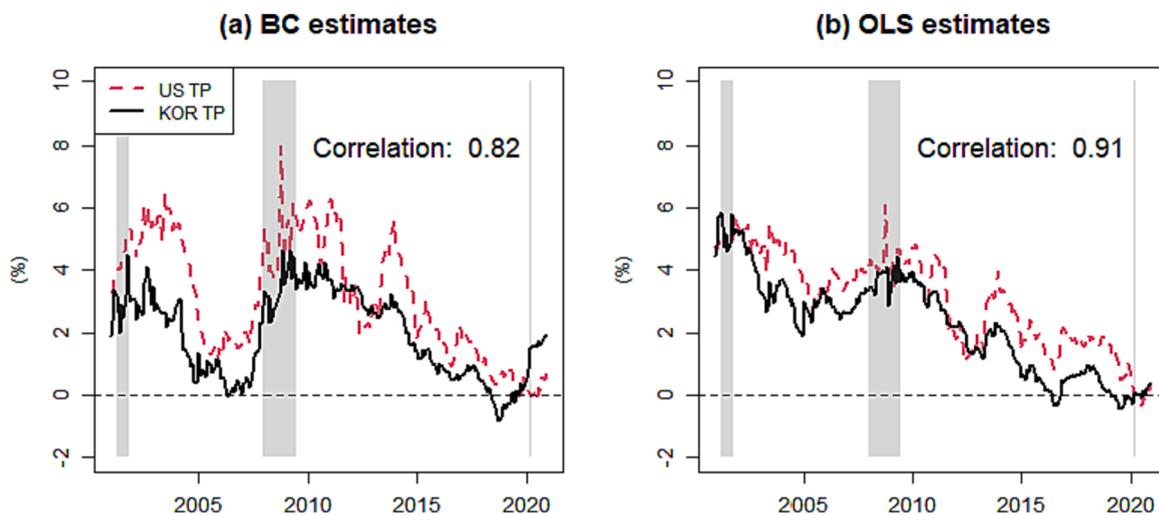


Fig. 6. The GDTSM-implied forward term premia. This figure presents the US and Korean term premia for the five- to ten-year forward rate. Both BC (panel A) and OLS (panel B) estimates are provided. Shading denotes NBER-dated recessions.

Ng (2009). Similar to the results of this paper, Bauer et al. (2014) and Abbritti et al. (2018) also found that bias correction makes the countercyclicality of term premia more pronounced.

Why does the bias correction make more countercyclical term premia? The BC estimates imply much slower mean reversion of short-term interest rates than the OLS estimates. For instance, during a recession, a central bank normally cuts its short-term policy rate, and the slow mean reversion in the BC estimates makes their expectation component (risk-neutral forward rate) lower than that in the OLS estimates. Because the term premium is obtained by subtracting the expectation component from the corresponding bond yield, the BC term premia are higher in recessions than the OLS term premia. Conversely, in an economic boom, for the same reason, the BC estimates produce lower term premia than the OLS estimates. As such, fluctuations in the BC term premia are more volatile with respect to the business cycle than the OLS estimates.

Here, I turn to the forward term premium estimates for Korea in Table 5. The results of regressing the term premia in Korea on the US output gap (either quadratic detrending or HP filtered) are similar to those for the US; only the BC estimates produce the countercyclicality of the KTB forward term premium with respect to the US business cycle. The bias correction plays an important role in making a pronounced countercyclical term premium in Korea. In contrast, when the term premium for Korea is regressed on the domestic IP gap variables, the regression coefficients are *not* significant for either BC or OLS estimates.

In sum, the BC term premium for the US shows marked countercyclicality with the US output gap. However, the BC term premium

Table 5
Countercyclicity of the forward term premium estimates in the US and Korea.

		IP gap of US			IP gap of Korea		
		Quadratic	HP	IP Growth	Quadratic	HP	IP Growth
US	BC	-0.794** (0.346)	-0.588* (0.367)	-0.155 (0.391)	-	-	-
	OLS	-0.148 (0.293)	-0.100 (0.287)	0.019 (0.320)	-	-	-
Korea	BC	-0.819*** (0.204)	-0.564*** (0.221)	-0.317 (0.253)	0.119 (0.337)	-0.091 (0.259)	0.124 (0.239)
	OLS	-0.316 (0.301)	-0.183 (0.291)	-0.149 (0.280)	0.101 (0.257)	0.012 (0.213)	0.494 (0.304)

Note: This table reports the estimation results from regressing each of the five- to-ten-year forward term premium estimates (either BC or OLS) on the IP gap for each of the US and Korea. The IP gaps are computed using i) the deviation of log IP from quadratic trend (denoted by Quadratic), ii) the deviation of the log of IP from Hodrick–Prescott (1997) trend (denoted by HP), or iii) year-on-year IP growth (IP Growth). All independent variables are standardized to have zero mean and unit variance. The numbers in parentheses are the heteroskedasticity and autocorrelation consistent (HAC) standard errors based on Andrews (1991). *, **, and *** indicate the significance at the 10%, 5%, and 1% levels, respectively. The sample period is from 2001:1–2020:12.

for Korea has little countercyclical relationship with the domestic business cycle, but substantial countercyclical relationship with the US business cycle. As the globalization of the international capital market progressed, the US and Korea's term premia as well as bond yields move very closely to each other, which explain the close countercyclical relationship between the KTB term premia and the US business cycle.

The remaining question is why the term premia for Korea are so weakly connected to the domestic business cycle. In the US, the 10-year Treasury yield moves very closely with the mortgage rate and has a substantial influence on the real economy, such as the housing market. In Korea, long-term interest rates are *not* as closely related to the real economy as in the US. To be specific, in Korea, most mortgage loans are floating rate loans, so short-term rates have more influence on the real economy than long-term rates. This fact helps to explain why the KTB term premia embedded in long-term rates are not significantly countercyclical with respect to domestic economic variables.

4.6. US Yield-Curve factors and global liquidity conditions

As noted earlier, Yun (2019) estimates a single GLF from 39 global liquidity-related variables⁹ by exploiting their relationship with one-year excess return on the KTBs. In his empirical analysis, where the Cochrane and Piazzesi (2005) factor, macro factors, and yield-curve factors are competing predictors, Yun finds that this GLF is the only significant predictor in the out-of-sample prediction for excess returns on the KTBs. As reported in Table 1, the US macro variables and yield-curve factors significantly predict one-year excess returns on the KTBs. This finding indicates a potential linkage between the GLF and US variables.

Table 6 reports the estimation results from regressing the GLF on the US variables: year-on-year IP growth, year-on-year CPI growth, the log of VIX,¹⁰ and yield-curve factors. To facilitate the interpretation of the results, I standardize all variables with zero mean and unit variance. The numbers in parentheses indicate the HAC standard errors (Andrews 1991).

The regression results show that when the US macro variables (year-on-year IP and CPI growth) are independent variables (Eq. (1)), the adj. R^2 is only 5%. When the log VIX is a single independent variable (Eq. (2)), the adj. R^2 reaches 32%. Eqs. (3)–(5) in Table 6 report the results for each of the US yield-curve factors. Interestingly, when the US slope factor is used as a single regressor (Eq. (4)), the adj. R^2 is as high as 65%. Its positive coefficient indicates that an increase in the US slope (a widening term spread) corresponds to an increase in the GLF. As shown in Eqs. (3) and (5) in Table 6, the remaining US yield-curve factors explain only about 2% of the variations in the GLF. When all US yield-curve factors are used as independent variables (Eq. (6)), the adj. R^2 reaches 74%. Finally, Eq. (7) in Table 6 shows that the variables considered jointly explain 76% of variations in the GLF. Therefore, given the US yield-curve factors, the incremental contribution from the US macro variables and VIX is only marginal. In particular, the log of VIX, which has fitted well to the univariate regression (Eq. (2)), becomes insignificant. This implies that VIX has no additional information regarding the GLF beyond the US yield-curve factors.

In short, the US yield-curve factors have a close relationship with the GLF suggested by Yun (2019). In particular, the US slope is the most important representative of the global liquidity condition relevant to KTB returns. For better understanding, Fig. 7 compares the GLF with the US slope factor and shows that the two variables share a common trend with a correlation coefficient of 0.81.

The bond return predictability of the slope factor (or the interest rate differential between long- and short-term rates) has often been documented since Fama and Bliss (1987) and Campbell and Shiller (1991). Recently, by estimating GDTSMs under the

⁹ They are nine monetary aggregates, fourteen interest rates, three real effective exchange rates, five stock indexes, five stock realized volatility, and three options-implied stock index volatilities. These variables are for the Euro, Japan, UK, and US.

¹⁰ The VIX index is the option-implied stock market volatility for the S&P500 index and is computed from a panel of options prices from Chicago Board Options Exchange (CBOE).

Table 6
The regressions of the GLF on the US macro variables, VIX and the US yield-curve factors.

Equation	US IP growth	US inflation	log VIX	US level	US slope	US curvature	Adj.R ²
(1)	-0.272** (0.129)	0.100 (0.178)					0.051
(2)			0.564*** (0.110)				0.315
(3)				-0.174 (0.282)			0.025
(4)					0.806*** (0.105)		0.648
(5)						0.165 (0.226)	0.022
(6)				-0.079 (0.107)	0.843*** (0.110)	0.294*** (0.087)	0.738
(7)	-0.037 (0.054)	0.174** (0.076)	0.095 (0.108)	-0.156 (0.116)	0.805*** (0.111)	0.212* (0.113)	0.757

Note: This table reports the estimation results from regressing the global liquidity factor (GLF) from Yun (2019) on selected combinations of the US variables: the IP and CPI growth, log VIX, and yield-curve factors (level, slope, and curvature). All independent variables are standardized to have zero mean and unit variance. The numbers in parentheses are the heteroskedasticity and autocorrelation consistent (HAC) standard errors based on Andrews (1991). *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The sample period is 2001:1–2017:12, the same as in Yun (2019).

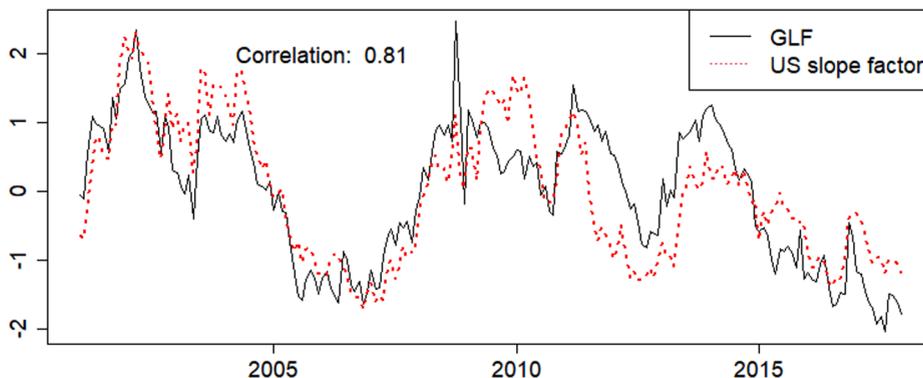


Fig. 7. Time-series of global liquidity factor (GLF) from Yun (2019) and the US slope factor (the second principal component from the US bond yields).

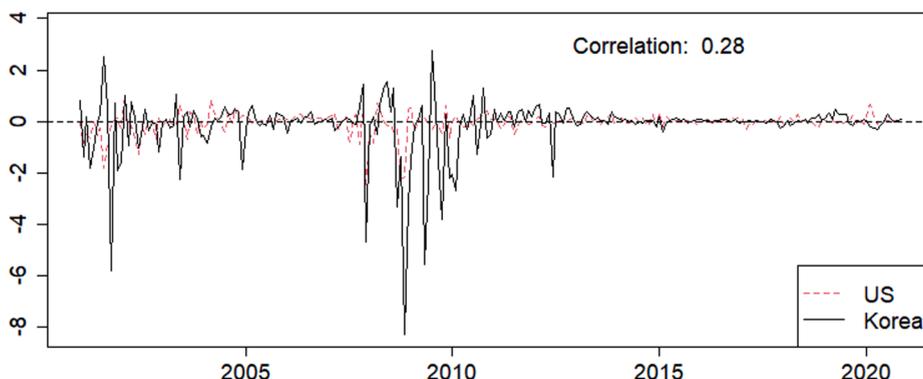


Fig. 8. The GDTSM-implied stochastic discount factors (SDFs) for the US and Korea.

restrictions on risk prices, [Bauer \(2018\)](#) concludes that only changes in the slope factor affect term premia. All these studies are about the US bond return predictability of the US interest rate differential. To extend these studies, I relate the US slope factor to the return predictability of a foreign bond in the current paper.

Why is GLF closely associated with the US slope factor? According to [Yun \(2019\)](#), GLF has a significant positive correlation with the VIX, and as documented by [Bruno and Shin \(2015a, b\)](#), VIX has a close relationship with the leverage of global banks, which is the determinant of the transmission of financial conditions across borders through banking sector capital flow. Thus, Bruno and Shin consider VIX an important global liquidity variable. When the VIX rises as economic uncertainty rises (for example, because of a financial crisis), the US Fed tends to respond by lowering its policy rate, leading to an increase in the US slope factor. These interactions among the VIX, the leverage of global banks, and the policy rates seem to give rise to a positive relationship between the GLF and the US slope factor.

4.7. Model implications for foreign exchange rates

Under the assumption of complete markets, the Korean won's (KRW) appreciation rate is determined by the difference in the log SDFs for the US and Korea ([Backus et al., 2001](#)):

$$\Delta s_{t+1} = m_{t+1}^{KR} - m_{t+1}^{US} \tag{13}$$

where m_{t+1}^{US} and m_{t+1}^{KR} indicate log SDFs for the US and Korea, respectively. Following Eq. (4), each GDTSM-implied log SDF is defined as

$$m_{t+1}^i = -y_t^{(1),i} - \frac{1}{2}\lambda_t^i \lambda_t^i - \lambda_t^i \varepsilon_{t+1}^i \quad \text{for } i = \text{US, KR} \tag{14}$$

where ε_{t+1}^i indicates each country's yield-curve shock vector (3-elements) for the VAR in (2).

Using the log SDFs estimated from the US and Korea bond yields, the GDTSM-implied KRW appreciation rate in Eq. (13) can be computed. [Fig. 8](#) illustrates the GDTSM-implied log SDFs for the US and Korea over time. The difference between these SDFs is the GDTSM-implied KRW appreciation rate under the complete market assumption. This figure shows very volatile SDFs around 2008 and among the two, Korea's fluctuates more. Needless to say, this is the period of the global financial crisis that occurred after the Lehman bankruptcy in 2007. As [Fig. 2](#) showed, the yield-curve factors in both countries vary greatly during this period. [Fig. 8](#) also displays that during this period, a sharp decrease in the SDFs (i.e., a decrease in marginal utility growth) occurred. Generally, the consumption-based asset pricing models tell us that an economic agent's marginal utility rises with a sluggish economy, but the opposite situation occurs in the SDFs implied by the bond markets. These results indicate a departure of the SDFs implied by the bond markets from those implied by the consumption-based asset pricing models.

Over the entire sample period, the annualized standard deviation of the KRW appreciation rate is computed as 10.6%. However, that of the GDTSM-implied appreciation rate reaches 365.7%! This arises because the global financial crisis exerted undue influence on the magnitude of standard deviation, but when this crisis period is excluded, the standard deviation of the actual appreciation rate is 8.1%. Notably, though, the standard deviation of the GDTSM-implied appreciation rate is still high at 250.2%.

This huge difference between the volatilities of actual and SDF-implied foreign exchange rates is not limited to this study. For example, according to [Brandt et al. \(2006\)](#), where the US is set as the home country and the UK, Germany, and Japan are the foreign countries, when a reasonable degree of risk sharing is assumed for country-pairs, the actual exchange rate volatility is about 10% per year, but the volatility of the SDF-implied exchange rate is greater than 100%. [Brandt et al. \(2006\)](#) estimated the SDF from the stock, foreign exchange, and money markets, whereas in the current paper, I estimate the SDF-implied exchange rate only from the bond market; in turn, I obtain greater volatility than Brandt et al.

[Table 7](#) presents the OLS estimation results from regressing the KRW appreciation rate (against USD) Δs_{t+1} on the SDF differential $m_{t+1}^{KR} - m_{t+1}^{US} = -(y_t^{(1),KR} - y_t^{(1),US}) - \frac{1}{2}(\lambda_t^{KR'} \lambda_t^{KR} - \lambda_t^{US'} \lambda_t^{US}) - (\lambda_t^{KR'} \varepsilon_t^{KR} - \lambda_t^{US'} \varepsilon_t^{US})$ (denoted by ΔSDF), or the three components of ΔSDF ($\Delta SDF(1)$, $\Delta SDF(2)$, and $\Delta SDF(3)$). Furthermore, excess returns on the CRSP value-weighted index and the Korean Composite Stock Price Index (KOSPI) are also considered regressors. The numbers in parentheses are HAC standard errors based on [Andrews \(1991\)](#). All independent variables are standardized to mean 0 and standard deviation 1 to facilitate comparison between regression coefficients.

The estimation result for Eq. (1) in [Table 7](#) shows that when the realized appreciation rate is regressed on the GDTSM-implied appreciation rate, the adj. R^2 is very low at 1.3%, indicating a huge deviation from the complete markets assumption of [Backus et al. \(2001\)](#). Next, in Eq. (2), I try to improve the fitting by using the individual components of the SDF-implied appreciation rate, but the goodness of fit is still only 2.3%. More problematically, some statistical coefficients show negative signs, unlike the model prediction.

[Yung \(2021\)](#) also performed similar regressions using the exchange rates of seven developed countries against USD. Her estimations produced R^2 ranging from 0 to 15%. She seems to have obtained higher R^2 than the current paper because she estimated her GDTSMs by exploiting information from both bond yields and exchange rates.

[Jotikasthira et al. \(2015\)](#) performed a similar analysis. To check market completeness, they regressed monthly currency appreciation rates on the yield curve factors and macro variables they employed as the factors of their term structure model. Similar to the result from the current paper, they obtained a small R^2 (6.7% and 10.3% for US/UK and US/Germany, respectively) and concluded that market completeness is difficult to rationalize. However, unlike in the current paper, they did not directly use the model-implied SDF in regression analysis; instead, they indirectly examined market completeness using their model factors as independent variables.

Table 7

The regressions of the KRW appreciation rate on the difference in the GDTSM-implied log SDFs.

Equation	ΔSDF	$\Delta SDF(1)$	$\Delta SDF(2)$	$\Delta SDF(3)$	CRSP	KOSPI	Adj.R ²
(1)	-0.133 (0.089)						0.013
(2)		0.102 (0.062)	-0.159** (0.072)	-0.114 (0.097)			0.023
(3)		0.071 (0.047)	-0.146** (0.066)	-0.076 (0.094)	0.453*** (0.130)		0.225
(4)		0.066 (0.048)	-0.146** (0.065)	-0.061 (0.103)	0.351** (0.152)	0.150* (0.084)	0.232

Note: This table reports the estimation results from regressing the KRW (Korean Won) appreciation rate (against USD) Δs_{t+1} on selected combinations of the variables: the SDF differential $\Delta SDF \equiv m_{t+1}^{KR} - m_{t+1}^{US} = -\left(y_t^{(1),KR} - y_t^{(1),US}\right) - \frac{1}{2}\left(\lambda_t^{KR'} \lambda_t^{KR} - \lambda_t^{US'} \lambda_t^{US}\right) - \left(\lambda_t^{KR'} \varepsilon_t^{KR} - \lambda_t^{US'} \varepsilon_t^{US}\right)$, each of its three components ($\Delta SDF(1) \equiv -\left(y_t^{(1),KR} - y_t^{(1),US}\right)$, $\Delta SDF(2) \equiv -\frac{1}{2}\left(\lambda_t^{KR'} \lambda_t^{KR} - \lambda_t^{US'} \lambda_t^{US}\right)$, and $\Delta SDF(3) \equiv -\left(\lambda_t^{KR'} \varepsilon_t^{KR} - \lambda_t^{US'} \varepsilon_t^{US}\right)$), excess return on the CRSP value-weighted index, and excess return on the KOSPI (Korean Stock Price Index). All independent variables are standardized to have zero mean and unit variance. The numbers in parentheses are the heteroskedasticity and autocorrelation consistent (HAC) standard errors based on Andrews (1991). *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The sample period is 2001:1–2020:12.

Next, Eqs. (3) and (4) in Table 7 consider excess returns on the broad stock indices in the US and Korea as additional independent variables in the regressions. They are potential candidates for the main pricing factors in equity-based asset pricing models. First, in Eq. (3), adding excess returns on the CRSP value-weighted market index as an independent variable, the adj. R² rises significantly to 22.5%. However, in Eq. (4), excess returns on the KOSPI, which is a representative broad stock market index in Korea, leads to only a marginal increase in adj. R².

It is not surprising given the characteristics of Korean capital markets that the equity market is more informative about the KRW exchange rate than the bond market. In Korea, the amount of foreign equity investment is about three times greater than the amount of foreign bond investment.¹¹ Furthermore, in foreign bond investment, the share of the public sector (central bank, sovereign wealth fund, international organizations) is higher than that of the private sector (commercial banks, investment companies, securities companies), implying a large proportion of long-term investments in bonds. Given these characteristics, the exchange rate will be more influenced by the determinants of foreign equity investment.

Considering the above, it seems difficult to obtain useful implications for the exchange rate through GDTSMs using only information on the bond market; introducing stock market information should allow for better explaining the exchange rate as it approaches the assumption of market completeness. However, as seen in the case of Brandt et al. (2006), there is still a concern in this case that the exchange rate will appear too smooth compared with the volatility of the SDF differential.

4.8. Robustness check

In this subsection, I perform the robustness check in three different directions. I provide the detailed empirical analysis in the online Appendix Sections C, D, and E; and summarize the analysis results here.

4.8.1. 10-factor GDTSM

The first robustness check extends the main six-factor GDTSM, which considers only yield-curve factors, to the 10-factor GDTSM where I employ the US and Korea macro variables (year-on-year IP and CPI growth) as additional model factors. Similar to the six-factor model, I also employ a two-block structure with US and Korean blocks. Each block consists of two macro variables (IP and CPI growth) and three yield-curve factors. The US block is set exogenously to have a unilateral influence on the Korean block.

To identify structural shocks in the impulse response analysis, I employ the variable transformation scheme of Jotikasthira et al. (2015). This scheme gives a better economic interpretation of structural shocks. I provide a detailed description of structural shock identification in the online Appendix Section C.

Similar to the six-factor model above, I analyze impulse responses, forecasting error variance decomposition, and term premium analysis for the 10-factor GDTSM; these results show that the yield-curve shocks have a much greater effect on the bond yields than the macro shocks. Accordingly, the results of the 10-factor model are broadly very similar to those of the six-factor model. Importantly, the US level factor makes the most important contribution to fluctuations in the term structure in Korea. For the one-year yield, the US level factor overwhelmingly contributes, and for the 10-year yield, the US slope and curvature as well as the US level all contribute significantly to the bond yield fluctuations.

Furthermore, as is the case with the six-factor model, when the two separate channels are considered, the policy channel dominates the one-year yield. As for the 10-year yield, the policy channel is dominant for the US level shock, but the risk channel is dominant for

¹¹ As of the end of January 2023, the balance of foreigners' bond investment is KRW 222 trillion, and the balance of foreign stock investment is KRW 636 trillion.

other shocks. These results for the long-term bond are sensitive to whether or not bias is corrected.

The results of the GDTSM-implied five- to ten-year forward term premia for the US and Korea are also very similar to those of the six-factor model. With the help of bias correction, they show a more marked countercyclicality to the US output gap.

4.8.2. Subsample analysis

The sample period in this paper is from 2001:1 to 2020:12. The second half of this sample includes some extraordinary episodes, such as the ZLB and the COVID-19 pandemic, which raises concerns about outliers or structural breaks in model estimation. To avoid such problems, some recently published studies (e.g., [Abbritti et al. 2018](#), [Bauer 2018](#), [Bauer and Rudebusch 2020](#)) intentionally excluded the ZLB period from their sample periods. As shown in [Fig. 1](#), no episode such as ZLB has occurred in Korea, unlike the US.

To check the robustness, I performed the same analysis using the first half of the entire sample (i.e., during 2001:1–2010:12). The online Appendix Section D provides the analysis results for impulse responses, forecasting error variance decomposition, and term premium analysis. Briefly, this subsample analysis indicates that the main results of this paper are largely unaffected by the sample selection. In particular, whether the short- or long-term yields, the US level factor, l_t^{US} , is the most important factor in explaining fluctuations in the term structure in Korea by a considerable margin. In addition, for the 10-year yield, the US slope factor is also important. Furthermore, for the one-year yield, the policy channel dominates, whereas for the 10-year yield, the policy channel dominates for the US level shock, but the risk channel dominates for the remaining shocks.

For the 10-year yield, there occurs a notable deviation from the full-sample results. It turns out that for the subsample, the contributions of the Korean local factors are comparable with those of the US factors. In fact, the Korean slope factor makes the largest contribution to fluctuations in the 10-year yield. This could be because foreign investment in the KTBs began in earnest only after the mid-2000 s, so that over the first half of the sample, the influence of foreign bond investment on the KTB market was relatively small.

4.8.3. Global Yield-Curve factors

In this paper, I set the US yield-curve factors to represent global yield-curve factors and analyzed the impacts of these factors on the term structure in Korea. In this section, I check how the analysis results change when I directly use the principal components estimated from the bond yields of various countries.

I use [Yung's \(2021\)](#) data set to extract the global principal components. The data are three-month, six-month, and one- through ten-year government bond yields (a total of 12 yields) for eight countries: Australia, Canada, Japan, Norway, Sweden, Switzerland, the UK, and the US (total of $8 \times 12 = 96$ yields). Yung's sample period is 2001:1–2019:5 (221 months), which is slightly shorter than the sample period in the current paper.

The correlation coefficients between the level, slope, and curvature factors between the US and global bond yields are 0.70, 0.25, and 0.24, respectively (which I present in the online Appendix Section E with a time-series plot). Notably, the correlation coefficient between the level factors is very high, indicating the tendency of international synchronization of inflation.

The results of impulse response and variance decomposition (See online Appendix Section E) are similar to the main analysis results in this paper. As can be expected, the global *level* factor makes the most important contribution to fluctuations in the term structure in Korea.

In the analysis of the global factor model, it is noteworthy that the persistence of the \mathcal{P} -measure dynamics under the OLS estimates (to be specific, the absolute maximum eigenvalue is 0.998) is already very large, so there is almost no difference between the BC and OLS estimates.

Therefore, when the forward term premia are estimated using the global yield-curve factor model, there is little difference between the BC and OLS term premia, and the trends are similar to the trend of the term premium for Korea based on the BC estimates of the main six-factor model.

Based on these results, it seems that the US yield-curve factors in the main analysis properly represent global principal components.

5. Concluding remarks

In this paper, I investigated the international linkages of the term structures between the US and Korea using a GDTSM, in which its state variables are the yield-curve factors in the US and Korea; I employed the US factors to approximate the global factors that affect the term structure in Korea. The GDTSM-implied VAR was specified so that the US block exogenously had a unilateral effect on the Korean block. To estimate the GDTSM, I used a maximally flexible model suggested by [Joslin et al. \(2011\)](#), removing small-sample bias via [Kilian's \(1998\)](#) bootstrap-after-bootstrap.

The main empirical findings in this paper are as follows. First, the results from impulse responses and variance decompositions indicate that the US *level* factor makes the largest contribution to fluctuations in the term structure for Korea, generating a strong tendency of comovement between the bond yields in the two countries. Because the level factor reflects long-term inflation expectations, this result implies global synchronization of inflation trends. The relative importance of structural shocks differs across yield maturities and forecast horizons: the US level factor overwhelmingly contributes to the fluctuations in the short-term rate in Korea, but all the US yield-curve factors jointly contribute significantly to the fluctuations in the long-term rate in Korea. In contrast, in response to the Korea-specific *local* yield-curve shocks, such as the local level shocks, the bond yields rise significantly on impact; however, in contrast to the US shocks, the effects of the shocks soon dissipate.

Second, I considered two separate channels (i.e., policy and risk-compensation) for propagating structural shocks into the Treasury yields in Korea. The policy channel dominates the short-term yield. In contrast, for the long-term yield, the dominant transmission

channels differ depending on the types of yield-curve shock: for US level shock, the policy channel is dominant, and for other shocks, the risk channel is dominant. This result confirms that the US level factor, which represents the long-term expectation of global inflation, is an important economic fundamental that a central bank will refer to in conducting its monetary policy. The result of which channel is dominant for the US level shock is particularly sensitive to whether or not small-sample bias is corrected. For example, the OLS estimates (without bias correction) produce the dominance of the risk channel across all the shocks for the long-term bond, which are what Jotikasthira et al. (2015) observed.

Third, the five- to ten-year forward term premium estimates for the US and Korea showed countercyclicality with respect to the US output gap; the bias correction produced a more pronounced countercyclicality, consistent with the results of Bauer et al. (2014) and Abbrittie et al. (2018); they found that the bias correction made term premia more countercyclical. Interestingly, Korea's term premium estimates are *not* as significantly countercyclical with respect to the domestic output gap as they are with respect to the US output gap; this is because the long-term interest rates in Korea are not as closely associated with the domestic real economy as in the US. Unlike in the US, for example, a significant portion of mortgage loans in Korea are floating-rate loans. In contrast, because the US and Korean Treasury yields more closely follow each other as the globalization of the capital market progresses, the term premia for Korea have a rather close relationship with the US economy.

Finally, the US yield-curve factors are closely related to the GLF suggested by Yun (2019). In particular, the US slope is the factor most closely related to the global liquidity condition for KTB returns. According to Yun (2019), GLF highly correlates with economic uncertainty variables such as the VIX. In the global risk-taking channel literature (e.g., Bruno and Shin 2015a, b), the VIX is recognized as an important global liquidity variable because it is highly related to the leverage of global banks. As experienced in the global financial crisis, whenever economic uncertainty is high, the US Fed tends to respond by lowering the policy rate (increasing the US slope factor), and this produces a positive relationship between the GLF and the US slope factor.

Another important topic for the term structures in the US and Korea is the implications of the GDTSM for the exchange rates. As documented by Backus et al. (2001), the difference in the nominal SDFs of two countries determines the change in the exchange rate (see Eq. (13)). However, the key assumption for this connection is market completeness for bond markets. In the current paper, the regression of the KRW appreciation rate on the GDTSM-implied SDF differential showed a very bad fit, making it difficult to accept the complete market assumption. The CRSP excess return, which can be an important risk factor for equity markets, markedly improved the regression fit. This is not surprising because in Korea, the amount of foreign equity investment is about three times greater than the amount of foreign bond investment. Thus, it is difficult to obtain useful implications for the exchange rate through GDTSMs using only information on the bond market.

One alternative approach that does not require the complete market assumption is the present value approach (Campbell and Clarida 1987, Engel 2016, Balduzzi and Chiang 2020, Dahlquist and Pénasse 2022). In this approach, variations in real exchange rates are decomposed into cash flow and discount rate components. Because the interest rate differential between two countries is an important determinant of the real exchange rate, there is room for applying GDTSMs to the present value model for real exchange rates. This would be an interesting future research topic that could build on the current paper, and I am in fact pursuing it in my current research work.

CRedit authorship contribution statement

Jaeho Yun: Conceptualization, Methodology, Formal analysis, Investigation, Data curation, Writing – original draft, Visualization, Investigation, Writing – review & editing, Funding acquisition.

Declaration of Competing Interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Data availability

Data will be made available on request.

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Appendix A. Supplementary material

Supplementary data to this article (online Appendix) can be found online at <https://doi.org/10.1016/j.jimonfin.2023.102924>.

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