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Monetary Policy and Long-Term Interest Rates in Korea: A Decomposition Analysis*

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We fit an affine term structure model to Korean nominal treasury yields between 1999 and 2020 to identify four components of long-term interest rates: real short-term interest rate expectations, real term premia, inflation expectations, and inflation risk premia. We then examine how long-term interest rates and their components respond to changes in monetary policy. We find that long-term interest rates do react to monetary policy changes, but this responsiveness has weakened since the global financial crisis of 2008. The decline of the responsiveness is largely attributable to real term premia. We compare these patterns to those in the U.S. and discuss possible explanations of our findings.

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I. Introduction

The effectiveness of monetary policy, to a large extent, rests on the ability of central banks to influence the cost of credit to households and businesses. Long-term interest rates are more relevant cost-of-credit measures than short-term interest rates for many households and businesses. Thus, the responsiveness of long-term interest rates to central banks' target rates is a critical element of the interest rate channel of monetary policy transmission (Mishkin, 1996; Boivin, Kiley, and Mishkin, 2010). Summarizing the early literature for the U.S., Berument and

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Froyen (2006) state that there exists "considerable uncertainty about the response of longer-term interest rates to changes in the federal funds rate" and further note that "even the direction of the effect is a subject of disagreement."[1] The debate regarding the responsiveness of long-term interest rates to monetary policy target rates, and the effectiveness of conventional monetary policy in general, only intensified as target rates approached the lower bound of zero in response to the global financial crisis of 2008 (Mishkin, 2009; Bauer, 2017; Jannsen, Potjagailo, and Wolters, 2019; Eo and Kang, 2020). Recent literature focuses on the responses of the different long-term interest rate components and their time-variation. For example, Hanson and Stein (2015) suggest that the term premia component of long-term interest rates may respond positively to monetary policy. They reason that lower (higher) central bank target rates may induce investors to buy (sell) long-term bonds thereby lowering (raising) long-term interest rates. Bauer, Rudebusch, and Wu (2014) find that the inflation risk premium component of long-term interest rates is counter-cyclical, which may lead to a negative correlation between long-term and short-term rates since short-term rates tend to be pro-cyclical.

The focus of this paper is the relationship between monetary policy and long-term interest rates in Korea. A number of papers have examined this relationship. A partial list includes Yoo and Oh (2005), Eom, Lee, and Ji (2007), Song (2009), Kim (2011), Kim, Eom, and Jang (2014), Jang and Hahn (2014), Bank of Korea (2015), Kim and Ku (2019), and Kim (2019). Among these, Eom et al. (2007) and Kim and Ku (2019) emphasize a potentially negative relationship between monetary policy and long-term interest rates while the remaining authors support the traditional role of the interest rate channel. While these papers highlight various aspects of the monetary-policy-and-long-term-interest-rates relationship, none has explicitly identified and estimated the components of long-term interest rates. Our goal is to do just that. We estimate the components of long-term interest rates and examine how these components respond to monetary policy. We pay special attention to the change in the responsiveness after the global financial crisis and compare the pattern in Korea to that of the U.S.

Following Burasschi and Jiltov (2005), Ang, Bekaert, and Wei (2008), and Bekaert and Wang (2010), we begin by decomposing long-term interest rates into three components: (i) the real rate, (ii) inflation expectations, and (iii) an inflation risk premium. The inflation expectations component reflects market participants' expectation on the future path of inflation rates while the inflation risk premium describes the compensation that savers require to hold nominal assets instead of real assets. Following Kim and Wright (2005), Durham (2006), Abrahams, Adrian, Crump, Moench, and Yu (2016),[2] we further divide the real rate into (1) the

[1] See Moreno (2008) for a study of emerging markets.

[2] See, in particular, the 4-component decomposition of long-term interest rates shown in Table 1 of

expectation of the short-term real rate and (2) a real term premium. This step allows us to examine the effect of agents' expectations of the future state of the economy and monetary policy (reflected in the short-term-real-rate-expectations component) and the effect of liquidity and other risk (reflected in the real term premium component).[3] We therefore investigate a 4-component decomposition of long-term interest rates. Note that interest rates in this paper refer to the yields of default-free government bonds. That is, we ignore the default risk component of interest rates.

Our decomposition of long-term interest rates enables us to distinguish between important aspects of the interest rate channel of monetary policy transmission. For example, when the central bank lowers target rates, agents may believe that the central bank has become more pessimistic about the future course of the economy, believing it is likely to keep target rates lower in the future (lowering the real rate expectation component). Alternatively, agents may become less sensitive to liquidity risk (lowering the real term premium). A third possibility is that agents become more pessimistic of the future economic environment (decreasing the inflation expectations component). A final possibility is that agents become less sensitive to inflation risk, perhaps because of the increased concern for output risk (lowering the inflation risk premium). Identifying these components of long-term interest rates enables us to examine and quantify each of these possibilities.

We estimate the components of long-term interest rates by fitting an affine term structure model using the linear regression method of Adrian, Crump, and Moench (2013) to nominal treasury yields data. All asset prices in the model are expressed using pricing kernels and inflation rates. Once we identify the dynamics of these variables, we can determine the components of long-term interest rates from their first two moments.

After obtaining the estimates of the long-term interest rate components, we examine how the changes in long-term interest rates and their components relate to changes in short-term target rates. The key findings are: (i) Long-term interest rates respond positively to short-term target rates. (ii) Long-term interest rates' responsiveness is reduced in the period following the global financial crisis. (iii) This reduction is largely attributable to the change in the responsiveness of the real rate component--in particular, the real term premium. Our findings are robust across alternative specifications. We obtain comparable results whether we use the

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Kim and Wright (2005), Tables 1 and 2 of Durham (2006) and Figure 7 of Abrahams et al. (2016).

[3] In other words, separating out the term premium is necessary to evaluate the deviation from the expectation hypothesis (predicting that long-term interest rates are equal to the sum of the expected short-term rates over the horizon of the long-term bond). Evans and Marshall (1998), Cochrane and Piazessi (2008), and Wright (2008) compare the expectation component to the term premium component of nominal interest rates, whereas Hanson and Stein (2015) compare the two components of real interest rates.

term of 5 years, 10 years, or 15 years. The results do not differ significantly whether we use effective rates (1-day interbank-market rates) or target rates announced by the Bank of Korea. To further ensure the robustness of our results, we estimate the policy response function (the Taylor rule) and identify the surprise component from effective rates. We then repeat the analysis, replacing the changes in target rates with the surprise component. We obtain mostly comparable results.

We have repeated all our estimations with U.S. data. We find that the patterns in the Korean data differ from those in the U.S. data. In the U.S., unlike in Korea, the overall response of long-term interest rates changed from being negative to being positive after the crisis, strengthening the monetary policy transmission channel. We find this change in responsiveness is mainly attributable to real term premia, whose responsiveness also changed from negative to positive. This contrasts to the Korean case where the responsiveness of real term premia changed from positive to negative. We interpret this difference as indicating the significance of the U.S. Federal Reserve's large-scale-asset-purchases (LSAP) program in the post-crisis period. Several papers including those by Abrahams et al. (2016) and Bonis, Ihrig, and Wei (2017) document that the LSAP program lowered the real term premium component of U.S. long-term interest rates. During the implementation of the LSAP program, the Fed fund rate was at a historically low level. This implies a likely positive correlation between long-term interest rates and the Fed fund rate. In Korea, on the other hand, the post-crisis asset purchase program was relatively small and lasted a relatively short period. This helps to explain the difference between the U.S. and Korean patterns.

In sum, our analysis indicates that the interest rate channel of monetary policy transmission in Korea became less effective after the global financial crisis and that the real term premia component is largely responsible. The relatively moderate quantitative easing program in Korea relative to the U.S. likely contributed to this pattern. The main distinguishing feature of our paper is to document how different components of long-term interest rates respond to monetary policy in Korea. While there exist a number of papers that examine the response of long-term interest rates to monetary policy in Korea, ours is the first to focus on the specific components of long-term interest rates.

The rest of the paper is organized as follows. In Section 2, we describe the decomposition of long-term interest rates into four components. In Section 3, we describe the empirical implementation of the decomposition analysis and report the results based on Korean data. In Section 4, we analyze the response of long-term interest rates and their components to target rate changes. In Section 5, we repeat the analysis for the U.S. data. In Section 6, we examine potential explanations of our findings, and Section 7 concludes the paper. We have placed further details on data and estimation procedures as well as additional results in the Appendix.

II. Components of Long-Term Interest Rates

The well-known Fisher equation specifies nominal interest rates as the sum of real interest rates and inflation expectations. Yet this formula is true only if agents are indifferent to inflation risk. With inflation risk, nominal interest rates are the sum of real interest rates and inflation *compensation*, the sum of inflation expectations and an inflation risk premium. This led Burasschi and Jiltov (2005), Ang, Bekaert, and Wei (2008), and Bekaert and Wang (2010) to posit a 3-way decomposition of long-term interest rates into (i) the real rate, (ii) inflation expectations, and (iii) an inflation risk premium.

As we are investigating a *long-term* interest rate, we also consider the term structure of interest rates. Thus, it is useful to further split the real rate into the expectations of future short-term real rates and a real term premium. The former is influenced by agents' beliefs about the central bank's monetary stance and also the future course of the economy. The latter is influenced by agents' preference toward liquidity and other risk factors. The term premium can also be characterized as the deviation from the expectation hypothesis (See Evans and Marshall, 1998; Cochrane and Piazessi, 2008; Wright, 2008). This consideration led Kim and Wright (2005), Durham (2006), and Abrahams et al. (2016) to adopt a 4-way decomposition of long-term interest rates into (i) short-term real rate expectations, (ii) real term premia, (iii) inflation expectations, and (iv) inflation risk premia.

Below we present a concise but complete derivation of the decomposition formula.[4] If a nominal log pricing kernel m_{t+1} exists,[5] one-period nominal log interest rates i_t satisfy the following equation[6]:

$$e^{-i_t} = E_t[e^{m_{t+1}}] \tag{1}$$

The left-hand side indicates the price of a one-period riskless bond, and the right-hand side indicates the discounted present value of a \$1 payment to be made one period later where the "stochastic discount factor" is given by $e^{m_{t+1}}$. Similarly, one-period real interest rates r_t satisfy the following equation:

[4] Kim and Wright (2005) and Durham (2006) present a similar formula in a continuous-time framework. Abrahams et al. (2016) present a similar formula utilizing the risk-neutral pricing concept. We present the formula within a discrete time framework without the risk-neutral pricing concept for a formula that corresponds more closely to the computation implemented in this paper.

[5] In a representative agent model, the pricing kernel represents the agent's marginal utility. In a more general setting, the existence of the pricing kernel is often motivated using the no-arbitrage principle.

[6] In our empirical implementation, one period corresponds to a month. We calculate interest rates, inflation rates, and pricing kernels in non-annualized terms.

$$e^{-r_t} = E_t[e^{\pi_{t+1}} e^{m_{t+1}}] \tag{2}$$

where π_{t+1} represents log inflation rates. The left-hand side of the equation indicates the price of a one-period real (indexed) bond, and the right-hand side indicates the discounted present value of an $\$ e^{\pi_{t+1}}$ payment to be made one period later.[7,8] Assuming joint normal distribution of m_{t+1} and π_{t+1}, Eqs. (1) and (2) can be rewritten as:

$$i_t = -E_t(m_{t+1}) - \frac{1}{2} V_t(m_{t+1}) \tag{3}$$

$$r_t = -E_t(m_{t+1} + \pi_{t+1}) - \frac{1}{2} V_t(m_{t+1} + \pi_{t+1}) \tag{4}$$

Combining the two,

$$i_t = r_t + E_t(\pi_{t+1}) + cov_t(m_{t+1} + \pi_{t+1}, \pi_{t+1}) - \frac{1}{2} V_t(\pi_{t+1}) \tag{5}$$

We may apply the same argument to n-period bonds instead of one-period bonds. Consider two n-period bonds: a nominal n-period zero-coupon bond and a real n-period zero-coupon bond. Writing the log yields on these two bonds as as $i_t^{(n)}$ and $r_t^{(n)}$, respectively, Eqs. (1) and (2) become[9]:

$$e^{-i_t^{(n)}} = E_t[e^{m_{t+1} + \cdots + m_{t+n}}] \tag{6}$$

$$e^{-r_t^{(n)}} = E_t[e^{\pi_{t+1} + \cdots + \pi_{t+n}} e^{m_{t+1} + \cdots + m_{t+n}}] \tag{7}$$

Then Eqs. (3) and (4) become:

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[7] $\$ e^{-r_t}$ is the price of a real bond that promises to pay $\$ e^{\pi_{t+1}}$ in the next period. Note that the real interest rate is determined by comparing the "real" future payment to the "real" current price. Taking the current period as the base period, the real future payment of this bond is $1 and the real current price is $\$ e^{-r_t}$ so that these numbers are consistent with the definition of the real interest rate. Confusion may arise unless one notes that the real current price equals the unadjusted price when we choose the current period as the base period.

[8] Korean inflation-indexed real bonds were first introduced in 2007 with new issues roughly every two years. As of November 2020, the aggregate principal of outstanding securities are approximately 9.3 trillion Korean wons (Korea Securities Depository, 2020). Trading of these securities is rather limited, making the computation of real interest rates from transaction prices impractical.

[9] In the empirical implementation, one period is to be taken as one month, and $i_t^{(n)}$ and $r_t^{(n)}$ are (non-annualized) log yields on n-month zero-coupon bonds.

$$i_t^{(n)} = -E_t(\sum_{j=1}^{n} m_{t+j}) - \frac{1}{2}V_t(\sum_{j=1}^{n} m_{t+j}) \tag{8}$$

$$r_t^{(n)} = -E_t(\sum_{j=1}^{n} m_{t+j} + \pi_{t+j}) - \frac{1}{2}V_t(\sum_{j=1}^{n} m_{t+j} + \pi_{t+j}) \tag{9}$$

Finally, Eq. (5) becomes:

$$\begin{aligned} i_t^{(n)} = r_t^{(n)} &+ E_t(\sum_{j=1}^{n} \pi_{t+j}) + cov_t(\sum_{j=1}^{n} m_{t+j} + \pi_{t+j}, \sum_{j=1}^{n} \pi_{t+j}) \\ &- \frac{1}{2}V_t(\sum_{j=1}^{n} \pi_{t+j}) \end{aligned} \tag{10}$$

We may further divide $r_t^{(n)}$ into the expectations component and the term premia component to obtain the main decomposition formula:

$$i_t^{(n)} = re_t^{(n)} + rtp_t^{(n)} + ie_t^{(n)} + irp_t^{(n)} \tag{11}$$

where four components--real interest rate expectations, real term premia, inflation expectations, and inflation risk premia--are defined as follows:

$$re_t^{(n)} \equiv E_t(\sum_{j=1}^{n} r_{t+j-1}) \tag{12}$$

$$rtp_t^{(n)} \equiv r_t^{(n)} - E_t(\sum_{j=1}^{n} r_{t+j-1}) \tag{13}$$

$$ie_t^{(n)} \equiv E_t(\sum_{j=1}^{n} \pi_{t+j}) \tag{14}$$

$$irp_t^{(n)} \equiv cov_t(\sum_{j=1}^{n} m_{t+j} + \pi_{t+j}, \sum_{j=1}^{n} \pi_{t+j}) - \frac{1}{2}V_t(\sum_{j=1}^{n} \pi_{t+j}) \tag{15}$$

Equation (13) shows that the real term premium is analogous to the nominal term premium, that is, the real term premium is the compensation given to a saver for buying a long-term real asset instead of buying and rolling a short-term real asset over the long-term asset's maturity.

Other authors have adopted variations of Eq. (11). Burasschi and Jiltov (2005), Ang, Bekaert, and Wei (2008), Bekaert and Wang (2010), and Grishenko and Huang (2013) study inflation risk premia with the framework of the three-component decomposition: $re_t^{(n)} + rtp_t^{(n)}$, $ie_t^{(n)}$, vs. $irp_t^{(n)}$. Evans and Marshall (1998), Cochrane and Piazzesi (2008), and Wright (2008) examine the expectations hypothesis by comparing the sum of two expectation components to the sum of two premia components, i.e. $re_t^{(n)} + ie_t^{(n)}$ vs. $rtp_t^{(n)} + irp_t^{(n)}$. Evans and Marshall (1998) also compare the sum of two real interest rate components to the sum of two inflation components, i.e., $re_t^{(n)} + rtp_t^{(n)}$ vs. $ie_t^{(n)} + irp_t^{(n)}$.

III. Decomposition Results

Below we describe the data, the empirical model and its estimation, and the decomposition results. We keep our discussion brief when we closely follow the existing literature. Further details are included in the Appendix.

3.1. Data

We have collected treasury yields, other interest rates, and inflation rates from the Bank of Korea web site.[10] We arrange all the time series into monthly series. For the estimation of the interest rate components, the sample period is from February 1999 to May 2020. Prior to February 1999, treasury yields data are available for three maturities only and thus are not adequate for estimation. Most of the analysis is performed for the period between May 1999 and May 2020 since the target rates of the Bank of Korea are available only from May 1999. The Bank maintained a money-supply-targeting regime until 1998 and announced its first target interest rates only in May 1999 (Bank of Korea, 2017). Between May 1999 and February 2008, the Bank used the 1-day interbank-market rate ("the call rate") as the reference rate. Since March 2008, the Bank has used the 7-day interbank-market rate ("the repo rate") as the reference rate.

3.2. Empirical Model and Estimation

For the estimation of the components of long-term interest rates, we adopt the affine term structure model of Adrian et al. (2013). This model expresses the term structure of interest rates via the moments of pricing kernels and inflation rates. Fitting this model to treasury yields data, we obtain the parameter estimates from which we can determine the moments of pricing kernels and inflation rates, which is what we need to calculate the components of long-term interest rates shown in Eqs. (12)-(15).

Other researchers have used similar estimation techniques in the study of Korean interest rate dynamics. Bank of Korea (2015) uses the same model to isolate the term premia component of long-term interest rates (but not inflation components). Song (2009) adopts an affine term structure model where factors are unobservable. Cho (2009) and Jang and Hahn (2014) present a macro-term-structure model which explicitly describes the effect of macro variables on the pricing kernel dynamics. Abrahams et al. (2016) adopted a similar approach for the U.S. data, but they utilize real (indexed) treasury yields as well as nominal treasury yields. Given the lack of active real treasury bond markets in Korea, we adopt a nominal-treasury-

[10] Further details are included in Appendix A.

only approach. Thus, we determine the distribution of inflation rates from realized inflation rates without supplementing them with breakeven inflation rates implied by real bond yields.

The affine term structure model of Adrian et al. (2013) consists of the following 5 equations:

$$X_{t+1} = \mu + \Phi X_t + v_{t+1}, \quad v_{t+1} \sim N(0,\Sigma) \tag{16}$$

$$i_t = \delta_0 + \delta_1' X_t \tag{17}$$

$$\lambda_t = \Sigma^{-\frac{1}{2}}(\lambda_0 + \lambda_1 X_t) \tag{18}$$

$$m_{t+1} = -i_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\Sigma^{-\frac{1}{2}}v_{t+1} \tag{19}$$

$$\pi_t = \omega_0 + \omega_1' X_t \tag{20}$$

Eq. (16) describes the dynamics of X_t, the vector of state variables. In this model, the dynamics of X_t generate time variation in the term structure of interest rates. Eq. (17) assumes that one-period interest rates, i_t, are linear in X_t (thus the name “affine”). Eq. (18) assumes the price of risk, λ_t, to be also linear in X_t. That λ_t is the price of risk is clear in Eq. (19), which shows that the conditional volatility of the nominal pricing kernel, $V_t(m_{t+1})$, is $\frac{1}{2}\lambda_t'\lambda_t$. Finally, Eq. (20) assumes inflation rates are also linear in X_t.[11]

The model has 9 parameters to be estimated: μ, Φ, Σ, δ_0, δ_1, λ_0, λ_1, ω_0, ω_1. We estimate these parameters by following the three-step estimation procedure of Adrian et al. (2013). We explain the procedure briefly as we introduce only minor variation of the procedure described by Adrian et al. (2013).[12] The input data consist of: (i) the principal components of the treasury yields of various maturities, assumed to be observed values of X_t, (ii) short-term interest rates, i_t, (iii) inflation rates, π_t, and (iv) excess returns on treasury bonds of various terms. For the principal component calculation, we use the yields of 1, 3, 6, ..., 120 months. We extracted three principal components after verifying that the eigenvalue associated with the third principal component is sufficiently different from zero. For the excess returns, we calculate them for the terms of 6, 12, 18, 24, 30, 36, 42, 48, 54, 60, 84, and 120 months. In the first step of the three-step estimation, we estimate Eqs. (16), (17), and (20) and obtain the estimates of μ, Φ, Σ, δ_0, δ_1, ω_0, and ω_1. In the second step, we regress excess returns on X_t and its residual components ($\hat{v}_t$). In the final step, we calculate λ_0 and λ_1 out of the second-step estimates.

Once we obtain the estimates of model parameters, we calculate the components

[11] As noted earlier, in the empirical implementation, one period is to be taken as one month, and interest rates, inflation rates, and pricing kernels are non-annualized variables.

[12] We describe further details in Appendix B.

of long-term interest rates using Eqs. (12)-(15). Below we present necessary formulae with minimal discussion as they are minor variations of the equations of Adrian et al. (2013).[13] The real interest rate expectation component (Eq. (12)) is given by:

$$E_t(\textstyle\sum_{j=1}^{n} r_{t+j-1}) = n\left(\delta_0 - \omega_0 - \omega_1'\mu - \frac{1}{2}\omega_1'\Sigma\omega_1 + \omega_1\lambda_0\right)$$
$$+(\delta_1' - \omega_1'\Phi + \omega_1'\lambda_1)[(n-1)(I-\Phi)^{-1}\mu$$
$$-(\Phi-\Phi^n)(I-\Phi)^{-1}(I-\Phi)^{-1}\mu + (I-\Phi^n)(I-\Phi)^{-1}X_t] \tag{21}$$

For the real term premia component (Eq. (13)), note the formula for real interest rates:

$$r_t^{(n)} = A_t^{(n)} + B_t^{(n)\prime} X_t \tag{22}$$

where $A_r^{(n)}$ and $B_r^{(n)}$ are determined by the following recursive formula:

$$A_r^{(n)} = A_r^{(n-1)} + B_r^{(n-1)\prime}(\mu - \Sigma\omega_1 - \lambda_0) - \frac{1}{2}B_r^{(n-1)\prime}\Sigma B_r^{(n-1)} - \omega_0$$
$$-\omega_1'\mu + \delta_0 - \frac{1}{2}\omega_1'\Sigma\omega_1 + \omega_1'\lambda_0 \tag{23}$$
$$B_r^{(n)} = (\Phi - \lambda_1)' B_r^{(n-1)} - \Phi'\omega_1 + \delta_1 + \lambda_1'\omega_1 \tag{24}$$

The initial values are given by $A_r^{(1)} = -\omega_0 - \omega_1'\mu + \delta_0 - \frac{1}{2}\omega_1'\Sigma\omega_1 + \omega_1'\lambda_0$ and $B_r^{(1)} = -\Phi'\omega_1 + \delta_1 + \lambda_1'\omega_1$. The inflation expectations component (Eq. (14)) is given by:

$$E_t \textstyle\sum_{j=1}^{n} \pi_{t+j} = n\omega_0 + \omega_1'[n(I-\Phi)^{-1}\mu$$
$$-(\Phi-\Phi^{n+1})(I-\Phi)^{-1}(I-\Phi)^{-1}\mu + (\Phi-\Phi^{n+1})(I-\Phi)^{-1}X_t] \tag{25}$$

Finally, for the inflation risk premia component (Eq. (15)), note the formula for nominal interest rates:

$$i_t^{(n)} = A^{(n)} + B^{(n)\prime} X_t \tag{26}$$

$A^{(n)}$ and $B^{(n)}$ are determined by the following recursive formula:

[13] For completeness, we provide the derivation of Eqs. (21)-(28) in Appendix B.

$$A^{(n)} = A^{(n-1)} + B^{(n-1)\prime}(\mu - \lambda_0) - \frac{1}{2} B^{(n-1)\prime} \sum B^{(n-1)} + \delta_0 \tag{27}$$

$$B^{(n)} = (\Phi - \lambda_1)' B^{(n-1)} + \delta_1 \tag{28}$$

The initial values are given by $A^{(1)} = \delta_0$ and $B^{(1)} = \delta_1$.

3.3. Results

We present the decomposition results for the long-term rates with 10-year maturity in Figure 1.[14] We plot 10-year interest rates in Panel A and their components in Panel B.[15] (The plots also include monetary policy target rates, which we will discuss in the next section). During the sample period, the average 10-year interest rate was 4.4% per year. As for the components, inflation expectations was largest, averaging 1.9% per year, while the inflation risk premium was the smallest with an average 0.43% per year. The inflation risk premium is sometimes negative, suggesting that agents are afraid of a deflationary tendency. During a period of low inflation or deflation, income and inflation rates may be positively correlated, and so nominal fixed income assets provide a hedge against low income. (That is, when income is small and inflation rates are low or negative, nominal assets pay more than real assets). Thus, the required premium on nominal assets can be negative. Over the sample period, 10-year interest rates exhibit a clear downward trend though there also exist periods of rising interest rates such as 2013 and 2017. The four components exhibit similar trends but month-to-month fluctuation varies across the components, suggesting that the time variation in the components may be driven by multiple factors.

IV. Responses of Long-Term Interest Rates and Components to Monetary Policy

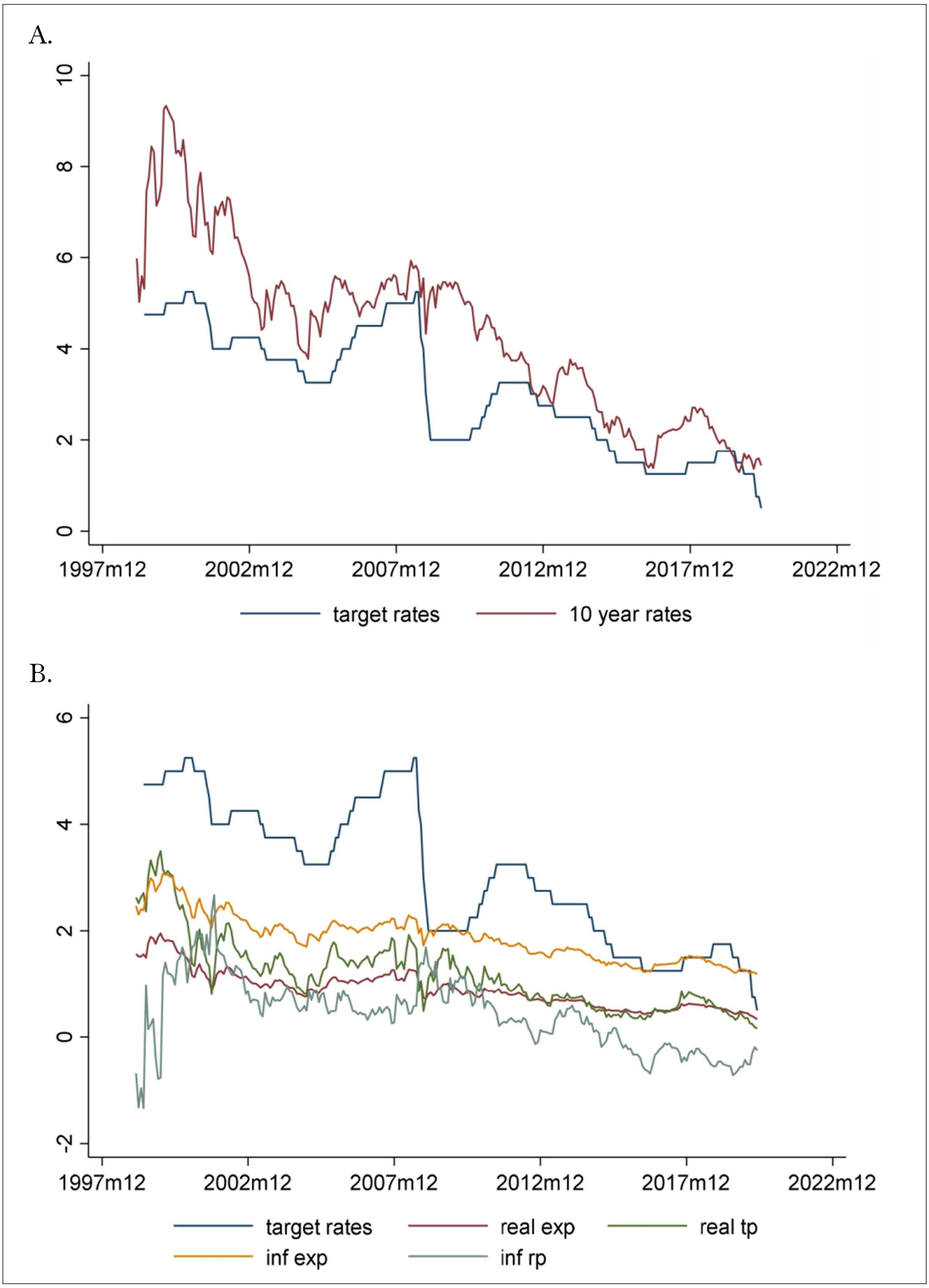
Having obtained the estimates of the components of long-term interest rates, we now turn to the responses of long-term interest rates and their components to the target rates of the Bank of Korea.

For intuition, Figure 1 shows target rates together with 10-year interest rates and the components. Over the long run, the 10-year interest rates move in the same direction as target rates, but over shorter periods the 10-year and target rates

[14] See Appendix C for the 3-way decomposition results where two real rate components are combined into a single component.

[15] In all tables and figures, interest rates and their components have been annualized for the ease of interpretation.

[Figure 1] Target Rates, 10-Year Rates, and Components



Notes: Target rates refer to the interest rate target set by the Bank of Korea. 10-year rates refer to the fitted values of 10-year treasury yields; the fitted values and the components are obtained from the affine term structure model of Adrian et al. (2013). The components identified are real interest rate expectations (real exp), real term premia (real tp), inflation expectations (inf exp), and inflation risk premia (inf rp).

sometimes move in opposite directions. This is particularly noticeable when target rates rose in 2010 and 2018. As noted earlier, month-to-month fluctuations vary across the components of 10-year interest rates, suggesting that the relationship to target rates is likely to be different for different components.

In the remainder of this section, we present two sets of analyses. The first is the single equation approach of Ellingsen and Soderstrom (2004) where the change in long-term interest rates and their components are regressed on the changes in target rates. We discuss the results for the maturity of 10 years in detail and report several additional calculations for robustness checks. In the second set of analyses, we take steps to reduce the potential endogeneity issue. We estimate the Taylor rule, identify policy surprises, and replace the changes in target rates with policy surprises.

4.1. Using Changes in Target Rates to Represent Monetary Policy

We follow Ellingsen and Soderstrom (2004) and regress the change in long-term interest rates and their components to changes in target rates. That is, we estimate the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t \tag{29}$$

where Y_t is $i_t^{(n)}$, $re_t^{(n)}$, $rtp_t^{(n)}$, $ie_t^{(n)}$, or $irp_t^{(n)}$, and Z_t is the target rate set by the Bank of Korea. We interpret β as indicating the responsiveness of Y_t to Z_t. Such interpretation may be criticized unless $Z_t - Z_{t-1}$ is truly exogenous. To address such criticism, we repeat our analysis after replacing $Z_t - Z_{t-1}$ with a proxy for policy surprises, as described in the next subsection.

An alternative to our single equation approach is the vector auto regression approach where other relevant macro variables are considered together with interest rates. Berument and Froyen (2009), in Table 1 of their paper, list eight studies that adopt a single equation approach and four adopting the vector auto regression approach. After comparing their results, Berument and Froyen (2009) suggest that the single equation approach may be more appropriate for recent periods. Among the papers that study this issue in Korea, Eom et al. (2007) adopt the single equation approach while Seo (2002), Kim (2011), Hwang (2013), and Kim (2019) adopt some variations of the vector auto regression approach.[16] One advantage of the single equation approach is that its estimates are simpler to provide intuitive understanding. Nonetheless, we carry out a vector autoregression analysis and report our findings in Appendix C. We confirm that the vector autoregression produces comparable results to the single equation approach.

[16] As an alternative to regression analysis, Park and Hur (2012) applied wavelet analysis to emphasize potentially different relationships at different frequencies.

We present the estimation results for the long-term rates with 10-year terms in Table 1.[17] Column (1) reports the estimates of coefficient β in Eq. (29) for 10-year interest rates and each of its components. All the coefficient estimates are at least marginally significant. The estimated coefficient on 10-year interest rates supports the traditional transmission mechanism of monetary policy. The 10-year rates move by about 25 basis points when target rates move by 1 percentage point; such response is somewhat larger than the U.S. estimates reported in the literature but still within a plausible range.[18] The responsiveness of inflation risk premia is negative while all the other components respond positively to the target rate changes.[19] The response of inflation compensation, i.e. the sum of inflation expectations and inflation risk premia, is negative corresponding to the pattern in the U.S. documented by Abrahams et al. (2016). The negative response of inflation risk premia is approximately offset by the positive response of real term premia, so the total response equals the sum of the responses of the two expectation components.

Columns (2) and (3) show coefficient estimates when the sample is split into two subperiods: the pre-global financial crisis period (May 1999 ~ February 2009) and the post-crisis period (March 2009 ~ May 2020). We split the sample around the month of February 2009 since the Bank of Korea lowered the target rate to 2.0% in that month and did not change the rate for the next 15 months until July 2010. Choosing any month between February 2009 and June 2010 does not alter the results substantially. It turns out that the traditional transmission channel of monetary policy became weaker after the crisis: The coefficient on 10-year rates declined from 0.35 to 0.008. Three of the four components--real rate expectations, the real term premium, and inflation expectations--contributed to this trend. In particular, the change in the responsiveness of the real term premium is largest: The coefficient declined from 0.341 to -0.023. The inflation risk premium is the only component whose responsiveness changed in the opposite direction: i.e., prior to the financial crisis it responded negatively to large rate changes but it became unresponsive after the financial crisis. In the last column of Table 1, we test the significance of the difference between the pre-crisis period and the post-crisis period. The decline in the responsiveness of real rate expectations and the real term premium are highly significant and the decline in the responsiveness of inflation

[17] See Appendix C for the 3-way decomposition results where two real rates components are combined into single component.

[18] We report the estimates for the U.S. in Table 3. For an earlier period, Cook and Hahn (1989) show that a one percentage point rise in the Fed target rate in the 1970s brought a 13 basis point increase in 10-year rates. Nilsen (1999) found somewhat lower responsiveness for the 1980s.

[19] The positive response of inflation expectations suggests that, when the central bank adopts expansionary policy (low target rates), agents do not believe that expansionary policy will create long-run inflation. The negative reaction of the inflation risk premium however shows that agents wish to have a greater compensation for the elevated uncertainty of inflation.

expectations is also somewhat significant.

[Table 1] Responses of 10-Year Rates and Components to Target Rates

	all period (1)	pre-crisis (2)	post-crisis (3)	(2) vs. (3)
changes in				
10-year rates	0.254*	0.350	0.008	-0.342
	[1.72]	[1.49]	[0.06]	[-1.03]
real rate expectations	0.135***	0.169***	0.042	-0.126**
	[5.40]	[4.37]	[1.50]	[-2.25]
real term premia	0.244***	0.341***	-0.023	-0.363**
	[3.44]	[3.13]	[-0.28]	[-2.28]
inflation expectations	0.102***	0.138***	0.006	-0.132*
	[3.17]	[2.75]	[0.17]	[-1.82]
inflation risk premia	-0.227*	-0.296	-0.017	0.279
	[-1.88]	[-1.53]	[-0.16]	[1.02]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the changes in target rates, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and Z_t is target rates. Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

To show the robustness of the results shown in Table 1, we have carried out several variations of the analysis. We have repeated the analysis for the 5-year and 15-year interest rates and found the results to be similar. We have also repeated the analysis using the effective inter-bank rates ("call rates") in place of the target rates and obtained very similar results.[20]

4.2. Using Rule-Based Policy Surprises

The changes in target rates, the right-hand side variable in Eq. (25), may be endogenous, making the interpretation of the estimation results potentially problematic. To remove potential endogeneity, we have estimated the Taylor rule for the monetary policy in Korea and identified the policy surprises component. The policy surprises are, at least within the framework of the Taylor rule, exogenous. We

[20] The first two variations are reported in Appendix C.

repeat the estimation of Eq. (29) using the policy surprises as the explanatory variable.

To estimate the Taylor rule for the Bank of Korea, we follow Kim and Ku (2019) and adopt the following specification[21]:

$$i_t^s = \beta_0 + \beta_1 \pi_{t+6} + \beta_2 x_t + \beta_3 i_{t-1}^s + \varepsilon_t \tag{30}$$

In the above, i_t^s is the one-day interbank rates at the end of month t, π_{t+6} is the year-on-year inflation rates at the end of month $t+6$, and x_t is a proxy for the GDP gap at the end of month t. Note that we estimate Eq. (30) from monthly data rather than from quarterly data since the number of quarterly observations is inadequate for our purpose. Estimating the Taylor rule from monthly data is not uncommon. See, for example, Gerdesmeir and Roffia (2003). Following the convention in the literature, we estimate GDP gap out of quarterly GDP by applying the Hodrick-Prescott filter; we then create monthly series by simply filling in each month with the latest value.[22] We have experimented with alternative proxies with the GDP gap, including the one based on monthly industrial production, and obtained similar results.

We use a GMM estimation procedure. We apply a two-step GMM procedure: in the first step, the weight matrix corresponds to the assumption of IID errors. For the second step, the weight matrix corresponds to the assumption of heteroskedastic errors. The instrument list consists of the past 6-month values of inflation rates, GDP gap, and 1-day interbank rates (but not the current-month values).

The residuals $\hat{\varepsilon}_t$ from the estimation of Eq. (30) are our measure of policy surprises. Figure 2 plots policy surprises together with target interest rate changes. While such policy surprises exhibit more variation than the target rate changes, the overall patterns are quite similar. The correlation coefficient between the two series is high at 0.76. Given the similarity between the two series, one can expect to obtain similar results when we replace the changes in target rates with policy surprises.

In Table 2, we report the responsiveness of 10-year rates and their components to policy surprises, i.e., we report the estimates of β in Eq. (29) after substituting policy surprises for $Z_t - Z_{t-1}$. Policy surprises are exogenous, at least within the framework of the Taylor rule. Thus, it is now less controversial to call β "responsiveness." The results shown in Table 2 are very similar to the results shown in Table 1. The responsiveness of 10-year rates declined. Regarding the responsiveness of components, that of two expectation components and the real

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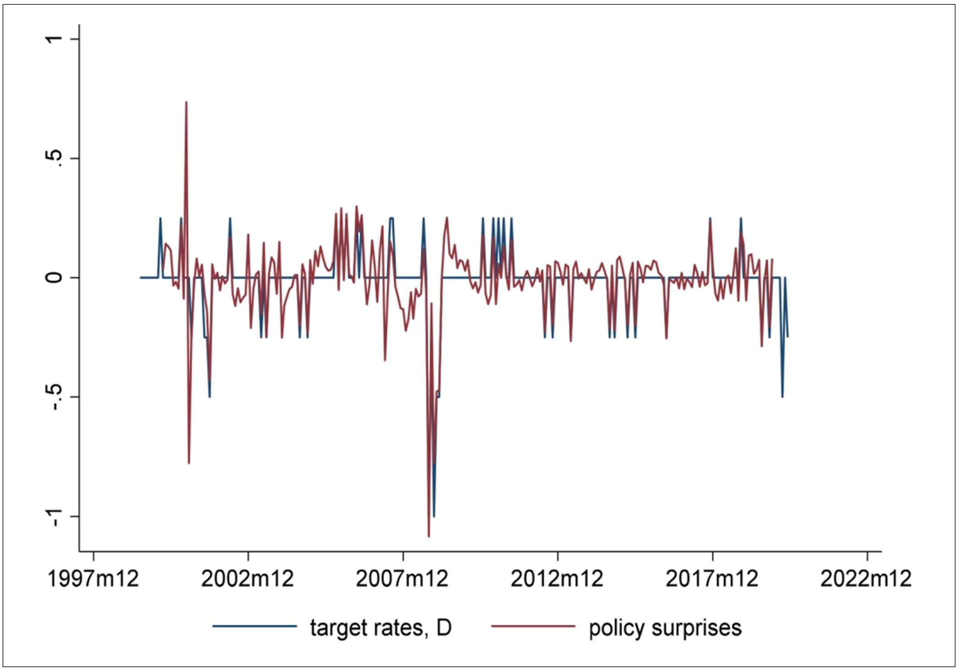
[21] Clarida, Gali, and Gertler (1998) derive a similar version of the Taylor rule by incorporating a central bank's interest rate smoothing objective into the standard Taylor rule. Cho (2020) estimate a variant of Taylor rule for Korea using a two-country panel framework.

[22] That is, we record quarterly values in the months of March, June, September, and December and for the remaining months, fill in with the latest available values.

term premium declined in the post-crisis period while that of the inflation risk premium moved in the opposite direction, i.e., its responsiveness changed from being negative to being zero.

[Figure 2] Policy Surprises vs. Changes in Target Rates



Notes: Policy surprises are the residual obtained after estimating the following version of the Taylor rule:

$$i_t^s = \beta_0 + \beta_1\pi_{t+6} + \beta_2 x_t + \beta_3 i_{t-1}^s + \varepsilon_t$$

In the above, i_t^s is the one-day interbank rates at the end of month t, π_{t+6} is the year-on-year inflation rates at the end of month $t+6$, and x_t is a proxy for the GDP gap at the end of month t. The estimation is based on monthly data and via GMM. See the text for further details.

[Table 2] Responses of 10-Year Rates and Components to Policy Surprises

	all period (1)	pre-crisis (2)	post-crisis (3)	(2) vs. (3)
changes in				
10-year rates	0.215**	0.248	0.064	-0.185
	[2.01]	[1.58]	[0.41]	[-0.66]
real rate expectations	0.115***	0.134***	0.025	-0.109**
	[5.57]	[4.48]	[0.82]	[-2.03]
real term premia	0.211***	0.269***	-0.061	-0.330**
	[3.55]	[3.13]	[-0.69]	[-2.14]

inflation expectations	0.087***	0.105***	0.004	-0.101
	[3.28]	[2.69]	[0.11]	[-1.45]
inflation risk premia	-0.198***	-0.260***	0.095	0.355*
	[-2.82]	[-2.68]	[0.79]	[1.94]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the policy surprises, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and $Z_t - Z_{t-1}$ is policy surprises. Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

V. Comparison to the U.S.

To better assess the patterns reported in the previous section, we compare them to those of the U.S. We have repeated our analysis applying the exact same procedures described in the previous sections to the U.S. data. Using the yield data described by Gurkaynak, Sack, and Wright (2006), we have estimated the affine term structure model and calculated the four components of 10-year interest rates. Then we have examined the responsiveness of 10-year interest rates and the four components to the target rates set by the U.S. Federal Reserve. The U.S. Federal Reserve announces the upper and lower limits instead of a single target rate after the global financial crisis. For this period, we take the upper limit as the target rate in our calculation. We have also identified policy surprises after estimating the policy function and examined the responsiveness of 10-year interest rates and the four components to the policy surprises.[23]

Table 3 reports the responsiveness of 10-year rates and their components to target rates in the U.S. (comparable to Table 1), and Table 4 reports the responsiveness of the same variables to policy surprises (comparable to Table 2). In both tables, one can see that the overall response of the 10-year interest rates changed from being negative to being positive after the crisis. That is, the traditional transmission channel of monetary policy became stronger after the crisis. This is in contrast to the Korea case where the responsiveness of the 10-year rates declined in the post-crisis period and the traditional transmission channel became weaker. In both countries,

[23] See Appendix A for a further description of the data used in the analysis.

[Table 3] Responses of 10-Year Rates and Components to Target Rates in the U.S.

	all period (1)	pre-crisis (2)	post-crisis (3)	(2) vs. (3)
changes in				
10-year rates	0.052	-0.015	0.241*	0.256
	[0.64]	[-0.14]	[1.71]	[1.40]
real rate expectations	0.107***	0.119***	0.067**	-0.052
	[4.67]	[3.51]	[2.17]	[-1.00]
real term premia	-0.022	-0.065	0.103	0.167
	[-0.49]	[-1.12]	[1.23]	[1.60]
inflation expectations	0.013	0.008	0.023	0.016
	[0.78]	[0.32]	[1.00]	[0.43]
inflation risk premia	-0.046	-0.076*	0.048	0.125
	[-1.35]	[-1.77]	[0.81]	[1.63]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the changes in target rates, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and Z_t is target rates. Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table 4] Responses of 10-Year Rates and Components to Policy Surprises in the U.S.

	all period (1)	pre-crisis (2)	post-crisis (3)	(2) vs. (3)
changes in				
10-year rates	-0.140**	-0.166**	0.057	0.223
	[-2.10]	[-2.12]	[0.30]	[1.00]
real rate expectations	0.075***	0.082***	-0.023	-0.104
	[3.91]	[3.16]	[-0.55]	[-1.63]
real term premia	-0.117***	-0.135***	0.045	0.181
	[-3.09]	[-3.15]	[0.40]	[1.42]
inflation expectations	0.006	0.005	-0.007	-0.012
	[0.43]	[0.29]	[-0.22]	[-0.26]
inflation risk premia	-0.103***	-0.118***	0.041	0.159*
	[-3.744]	[-3.723]	[0.502]	[1.713]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the policy surprises, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and $Z_t - Z_{t-1}$ is policy surprises.

Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

the component that is most responsible for the change is the real term premia component. In the U.S. the responsiveness of real term premia changed from being negative to being positive whereas in Korea it moved in the opposite direction, from being positive to being negative.

In the U.S., the positive response of real term premia in the post-crisis period can be attributed to the LSAP program of the Federal Reserve. Between November 2008 and June 2012, the Federal Reserve initiated a number of LSAP programs and purchased assets worth 4.5 trillion dollars, amounting to the 30% of annual GDP.[24] The LSAP program, by design, put downward pressure on long-term yields; since the LSAP program tended to be accompanied by lower target rates, the outcome was a positive correlation between long-term yields and target rates. Gagnon, Raskin, Remache, and Sack (2011) and Abrahams et al. (2016) explain that the effect of LSAP program was mainly via real term premia though Krishnamurthy and Vissing-Jorgensen (2011) and Bauer and Rudebusch (2014) assign more significance to the role of (nominal and real) interest rate expectations.[25]

Contrary to the situation in the U.S., the quantitative easing program in Korea was of small scale and lasted for a relatively short period of time. The Bank of Korea injected 18.5 trillion wons (about 18.5 billion U.S. dollars) into the economy between September 2008 and February 2009; 18.5 trillion wons is about 1.5% of the GDP in 2009, a much smaller magnitude relative to the U.S. program.[26] Moreover, most of this money was withdrawn by the end of 2011. In early 2012, the Bank of Korea reported to have withdrawn more than 90% of the money supplied in the quantitative easing program.[27] Another difference with the U.S. program is that

[24] See Appendix A of Woodford (2012) for the list of the LSAP announcements.

[25] Krishnamurthy and Vissing-Jorgensen (2011) do not adopt a formal decomposition of long-term interest rates. Their empirical findings thus allow for alternative interpretations. Bauer and Rudebusch (2014) decompose long-term interest rates into an expectations and a term premium component but do not distinguish between real and inflation components, which also allows alternative interpretations. They suggest that 40-50% of long-term interest rate changes are attributable to the changes in the expectations component. If we combine the coefficients on the real and inflation components in Table 3, we find that these two expectation components together explain more than 40% of long-term interest rate changes. Thus, the Bauer and Rudebusch (2014) results are not irreconcilable with ours.

[26] See Bank of Korea (2017).

[27] See Bank of Korea (2012).

"maturity extension" was never an explicit goal of the Bank of Korea program. Only 1 trillion wons were allocated to the purchase of treasury bonds, and only a part of this amount was spent for the purchase of long-term treasury bonds. These differences in the quantitative easing program in Korea and the U.S. are likely to be a main reason for the different changes in the responsiveness of real term premia in the two countries.

VI. Reduced Responsiveness of Long-Term Interest Rates: Further Explanations

As we have discussed in Section 4, the response of long-term interest rates to monetary policy in Korea became weaker after the global financial crisis and three of the four components— real interest rate expectations, the real term premium, and inflation expectations--contributed to this change. The inflation risk premium was the only component that offset the others. In this section we consider potential explanations of these patterns.

Let us first consider the response of the two expectations components, real interest rate expectations and inflation expectations. The weakened responsiveness of these two components indicates that the changes in target rates do not reveal new information about the future state of the economy. This may be due to increased uncertainty making agents more pessimistic about the central bank's quality of private information about the future state of the economy.[28] Additionally, it may be due to the fact that target rates are near the lower bound in the later period. Since target rates are near the lower bound, the central bank becomes more reluctant to adjust target rates downward (to preserve the ability to do so in a more critical moment), which limits the information contents of target rates.[29]

For the real term premium, a combination of a weakened demand channel and a stronger risk channel are likely to be at play. Bauer (2017), Bonis et al. (2017), and Abrahams et al. (2016) describe the demand channel and suggest that, when the central bank buys long-term assets while cutting target rates, the greater demand for long-term assets cuts real term premia. Hanson and Stein (2005) propose a related logic that lower target rates reduce short-term interest rates and induce investors to buy more long-term bonds, which again cuts real term premia. Similarly, Bank of Korea (2015) suggests that greater liquidity from an expansionary monetary policy

[28] Note that our discussion does not preclude money neutrality. Even if money is neutral, the central bank may *know* more about what is likely to happen, in which case central banks' action still signals the future path of short-term rates.

[29] Jang and Kim (2017) present a detailed analysis of the formation of inflation expectations in Korea.

(i.e. lower target rates) cuts the premium for liquidity. Such a demand channel has apparently been weakened after the financial crisis.

Regarding the risk explanation, Bauer et al. (2014) view real risk premia as having a counter-cyclical nature. When an economy is doing poorly, an investor requires a greater premium for longer-term assets, increasing the risk premia. At the same time, the central bank is likely to lower target rates in response to the poor state of the economy. Thus, real term premia are negatively correlated with target rates. Foreign investors' behavior is likely to exacerbate this pattern. When the economy is doing poorly and the perceived risk level goes up, foreign investors are likely to sell long-term assets increasing real term premia. Such a risk channel may have been strengthened after the financial crisis, given the increased concern for risk in the government bond market in Korea.

The different responsiveness of real term premia in the U.S. after the financial crisis can be atrributed to a stronger demand channel and, possibly, a relatively weak risk channel. The LSAP of the Federal Reserve strengthened the demand channel. The Fed's massive purchases of long-term bonds shrank the amount of long-term bonds available for private investors to push up their prices and decrease the term premia. Aggressive Federal Reserve policy may have also weakened the risk channel effect, or perhaps weakened the risk channel effect vis-à-vis the effect of a strong demand channel.

Finally, the change in the responsiveness of inflation risk premia (from a negative response to a near-zero response) suggests that the economy is making a transition from a non-deflationary environment to a deflationary environment. In a non-deflationary period, real income and inflation rates tend to be negatively correlated.[30] In this case, when low target rates signal a high likelihood of recession, agents perceive real assets to be a good hedge against low income and so demand a premium on nominal assets, leading to large inflation risk premia. Thus, the inflation risk premium responds negatively to target rates. This consideration led Abrahams et al. (2016) to characterize inflation risk premia as a counter-cyclical variable. The opposite response is likely in a deflationary environment as noted by Bauer (2017). When deflation fears are high, real income and inflation rates tend to be positively correlated.[31] In this case, when low target rates signal a high likelihood of recession, agents perceive nominal assets to be a good hedge against low income (since nominal assets pay more than real assets in times of low income and low inflation) and are willing to accept low or even negative inflation risk premia on nominal assets. Thus, inflation risk premia respond positively to target rates. Given

[30] Note that the real pricing kernel is essentially marginal utility and that it moves in the opposite direction to real income. Thus, a negative correlation between real income and inflation rates leads to a positive correlation between the real pricing kernel $\sum_{j=1}^{n} m_{t+j} + \pi_{t+j}$ and inflation rates $\sum_{j=1}^{n} \pi_{t+j}$. By Eq. (15), this implies that $irp_t^{(n)}$ is positive.

[31] In this case, $irp_t^{(n)}$ is negative. See the previous footnote.

this association between deflation fears and a positive responsiveness of inflation risk premia, we may interpret the close-to-zero response of inflation risk premia as suggesting that the economy may be moving toward a deflationary environment.

Alternatively, the change in the responsiveness of inflation risk premia may reflect the anchoring of inflation expectations. The Bank of Korea has pursued inflation targeting since 1998 which by design anchors agents' inflation expectations to target inflation rates (Ball and Sheridan, 2004; Capistran and Ramos-Francia, 2010; Brito, Swallow, and Gruss, 2018). If the Bank of Korea has successfully anchored inflation expectations to the target inflation level, the variation in inflation expectations and thus the inflation risk premium decreases (in absolute terms). This interpretation may be more applicable to the post-crisis period. As shown in Figure 1, inflation risk premia in the post-crisis period are near zero while not significantly negative.

VII. Conclusion

We follow Kim and Wright (2005), Durham (2006), Abrahams et al. (2016), and Bauer (2017) and make a sensible and easily comprehended decomposition of nominal long-term interest rates into four components: real short-term interest rate expectations, real term premia, inflation expectations and inflation risk premia. The decomposition allows us to examine how the long-term government bond markets have responded to Korean central bank target changes.

Our "all period" results in Table 1 shows a weak positive but significant response of long-term interest rates to target rates in Korea, leading us to conclude that, overall, the interest rate channel of monetary policy transmission remains relevant in Korea. Concerning the components, we find that a Bank of Korea monetary easing (a lower target rate) induces lower real rate expectations, lower real term premia, and lower inflation expected over the maturity. Inflation risk premia, on the other hand, show a significant rise in this case. Apparently, savers wish to be compensated for the perceived higher uncertainty in inflation. Using Taylor rule policy surprises instead of target rate changes brings approximately the same coefficient estimates on all dependent variables, the long-term rates as well as the components.

We do observe that the responsiveness of long-term interest rates to monetary policy declined after the global financial crisis. This change is understandable given that target rates are at a historically low level. The Bank of Korea is likely to exhibit precautionary behavior, delaying rate cuts to preserve the ability to lower the rates in the future, which in turn limits the information content of target rates. Among the four components of long-term interest rates, the real term premia component is the largest contributor to the decline in the responsiveness of long-term interest rates. This component responded positively to target rates in the pre-crisis period, but its response became negative in the post-crisis-period. Increased risk premia and weakening of investors' portfolio rebalancing are the two likely causes.

One striking difference compared to the U.S. case is that the responsiveness of real term premia in Korea declined after the crisis. The LSAP program in the U.S. is credited with reducing the real term premia component of long-term interest rates. Given the relatively small quantitative easing program adopted by the Bank of Korea, the different changes in the responsiveness of real term premia is not surprising. To put it another way, our analysis suggests that not adopting an aggressive quantitative easing policy contributed to the declining effectiveness of monetary policy in Korea.

Appendix

A. Data: Further Details

[Table A1] Time Series Used in This Study

Series	Beg. month	End. month	Source
Series used in ATSM estimation			
91D CD rates	1999m2	2020m5	ecos.bok.or.kr
6M KORIBOR	2006m2	2020m5	ecos.bok.or.kr
Yields of 1Y monetary stabilization bonds	1999m2	2020m5	ecos.bok.or.kr
Yields of 2Y monetary stabilization bonds	1999m2	2020m5	ecos.bok.or.kr
Yields of 3Y treasury bonds	1999m2	2020m5	ecos.bok.or.kr
Yields of 5Y treasury bonds	2000m1	2020m5	ecos.bok.or.kr
Yields of 10Y treasury bonds	2000m12	2020m5	ecos.bok.or.kr
Yields of 20Y treasury bonds	2006m1	2020m5	ecos.bok.or.kr
Yields of 30Y treasury bonds	2012m9	2020m5	ecos.bok.or.kr
Consumer price index	1999m2	2020m5	ecos.bok.or.kr
Target rates			
Bank of Korea reference rates	1999m5	2020m5	ecos.bok.or.kr
Additional series used in policy function estimation			
1D call rates	1999m5	2020m5	ecos.bok.or.kr
Industrial production index, s.a.	2000m1	2020m5	ecos.bok.or.kr
Real GDP, s.a. (quarterly)	1999m9	2020m5	ecos.bok.or.kr
Series used in U.S. ATSM estimation			
Nelson-Siegel-Svensson parameters	1999m2	2020m5	See Gurkaynak et al. (2006).
Consumer price index	1999m2	2020m5	fred.stlouisfed.org
U.S. target rates			
Fed fund rates, upper limit	1999m5	2020m5	fred.stlouisfed.org
Additional series used in U.S. policy function estimation			
Fed fund rates, effective	1999m5	2020m5	fred.stlouisfed.org
Industrial production index, s.a.	1999m9	2020m5	fred.stlouisfed.org
Real GDP, s.a. (quarterly)	1999m9	2020m5	fred.stlouisfed.org

B. Estimation of Eqs. (16)-(20): Further Details

Given the limited range of bond maturities published by the Bank of Korea, we first fit the daily Nelson-Siegel curve using the available bond yields and determine the bond yields of the other maturities. The Nelson-Siegel curve for day t can be written as:

$$i_t^{(n)}\frac{12}{n}=\beta_{t,0}+\beta_{t,1}\left(1-e^{-\frac{n}{12\tau_t}}\right)\frac{12\tau_t}{n}+\beta_{t,2}\left[\left(1-e^{-\frac{n}{12\tau_t}}\right)\frac{12\tau_t}{n}-e^{-\frac{n}{12\tau_t}}\right] \quad \text{(B1)}$$

where $i_t^{(n)}$ is the non-annualized, log yield of the zero bond of n-month maturity and $\beta_{t,0}$, $\beta_{t,1}$, $\beta_{t,2}$ and τ_t are the parmaters to be estimated. An alternative to the Nelson-Siegel curve is the Nelson-Siegel-Svensson curve which contains two extra parameters. Including two extra parameters does not alter the analysis much, so we chose the simpler specification of the Nelson-Siegel curve. We use the nonlinear least square method to estimate these parameters. Note that the yields published by the Bank of Korea for $n \geq 12$ are par yields rather than zero yields. If $P_t^{(n)}$ is the par yield in percent, then it can be expressed in terms of log zero yields as follows:

$$\frac{P_t}{200} = \frac{(1 - e^{-i_t^{(n)}})}{e^{-i_t^{(6)}} + e^{-i_t^{(12)}} + \cdots + e^{-i_t^{(n)}}} \tag{B2}$$

The above expression can be derived from the definition of the par bond: the present value of a par bond paying semi-annual coupons of $\$P_t$ and the final payment of \$1 equals \$1. Utilizing the above expression, we can determine $\beta_{t,0}$, $\beta_{t,1}$, $\beta_{t,2}$ and τ_t from any combination of zero yields and par yields as long as the degree of freedom is nonnegative.

While the Nelson-Siegel curves can be estimated for each day in our sample, we only use the end-of-month estimates. We create the monthly time-series of yields for the maturities of 1, 3, 6, ..., 120 months. From these yields, we extract three principal components. We also calculate excess returns for the terms of 6, 12, 18, 24, 30, 36, 42, 48, 54, 60, 84, and 120 months. The excess return for the term of n months is defined as:

$$xr_t^{(n)} = -i_t^{(n)} + i_{t-1}^{(n+1)} - i_{t-1} \tag{B3}$$

where i_{t-1} is the log yield for one-month maturity.

Before we describe the three-step estimation procedure, we first verify the formulae shown in Eqs. (21)-(28). We need to utilize these formulae to relate the excess returns to the model parameters. To verify Eq. (21), substitute the expressions in Eqs. (19) and (20) for m_{t+1} and π_{t+1} in Eq. (4) and then further substitute the expressions in Eqs. (16)-(18) for X_{t+1}, i_t, and λ_t in the resulting formula. Then we obtain the following expression for r_t:

$$r_t = \delta_0 + \delta_1' X_t + \omega_0 + \omega_1'(\mu + \Phi X_t) - \frac{1}{2}\omega_1' \Sigma \omega_1 + \omega_1'(\lambda_0 + \lambda_1 X_t) \tag{B4}$$

From the above, Eq. (21) can be easily verified.

To verify the recursive formulae of Eqs. (22)-(24), note that the time-t price of

n-period real bond equals the present value of the time-($t+1$) price of ($n-1$)-period real bond:

$$e^{-r_t^{(n)}} = E_t(e^{m_{t+1}} e^{\pi_{t+1}} e^{-r_{t+1}^{(n-1)}}) \tag{B5}$$

In the above, the prices of real bonds are normalized by dividing them by the current level of the price index. Given that the joint distribution of the variables is normal, the above equation implies:

$$-r_t^{(n)} = E_t(m_{t+1} + \pi_{t+1} - r_{t+1}^{(n-1)}) + \frac{1}{2}V_t(m_{t+1} + \pi_{t+1} - r_{t+1}^{(n-1)}) \tag{B6}$$

Substituting $A_r^{(n)} + B_r^{(n)\prime} X_t$ and $A_r^{(n-1)} + B_r^{(n-1)\prime} X_{t+1}$ for $r_t^{(n)}$ and $r_{t+1}^{(n-1)}$ in the above equation, respectively, we can verify Eqs. (22)-(24).

To verify Eq. (25), we may simply substitute the expression in (20) for for π_t in the left-hand side of Eq. (25).

To verify the recursive formulae of Eqs. (26)-(28), note that the time-t price of n-period bond equals the present value of the time-($t+1$) price of ($n-1$)-period bond:

$$e^{-i_t^{(n)}} = E_t(e^{m_{t+1}} e^{-i_{t+1}^{(n-1)}}) \tag{B7}$$

Given the normal distribution of the variables, the above equation implies:

$$-i_t^{(n)} = E_t(m_{t+1} - i_{t+1}^{(n-1)}) + \frac{1}{2}V_t(m_{t+1} - i_{t+1}^{(n-1)}) \tag{B8}$$

Substituting $A^{(n)} + B^{(n)\prime} X_t$ and $A^{(n-1)} + B^{(n-1)\prime} X_{t+1}$ for $i_t^{(n)}$ and $i_{t+1}^{(n-1)}$ in the above equation, respectively, we can verify Eqs. (26)-(28).

We now describe the three-step estimation procedure. In the first step, we estimate Eqs. (16), (17), and (20) and obtain the estimates of μ, Φ, Σ, δ_0, δ_1, ω_0, and ω_1. Eq. (16) is estimated by equation-by-equation OLS. Eqs. (17) and (20) are also estimated by OLS.

In the second step, we regress excess returns on X_t and its residual compoents (v_t). That is, for each n, $n = 6, 12, \cdots, 60, 84$, we estimate the following equation:

$$xr_t^{(n)} = a_n + b_n' X_t + c_n \hat{v}_t + error \tag{B9}$$

Once again, the estimation is by OLS.

In the final step, we calculate λ_0 and λ_1 out of the second-step estimates. Substituting the recursive formulae for $i_t^{(n)}$, $i_{t-1}^{(n+1)}$, and t_{t-1} in (B3), we can verify:

$$xr_t^{(n)} = -B^{(n)\prime}\lambda_0 - \frac{1}{2}B^{(n)\prime}\Sigma B^{(n)} - B^{(n)\prime}\lambda_1 X_{t-1} - B^{(n)\prime}v_t \tag{B10}$$

Comparing (B10) to (B9), we see that $\hat{c}_n$ is a good estimate of $B^{(n)}$ and that the coefficient estimate in the regressing of $\hat{b}_n$ on $\hat{c}_n$ is a good estimate of λ_1. Moreover, the coefficient estimate in the regression of $\hat{a}_n$ on on $\hat{c}_n$ is a good estimate of λ_0.

C. Further Results

Three sets of analyses are reported in this appendix.

First, we present results based on the 3-way decomposition rather than the 4-way decomposition of long-term interest rates of the main text. Figure C1 plots the times series of components (in the format of Figure 1 of the main text). Table C1 reports the responses to target rate changes (in the format of Table 1 of the main text).

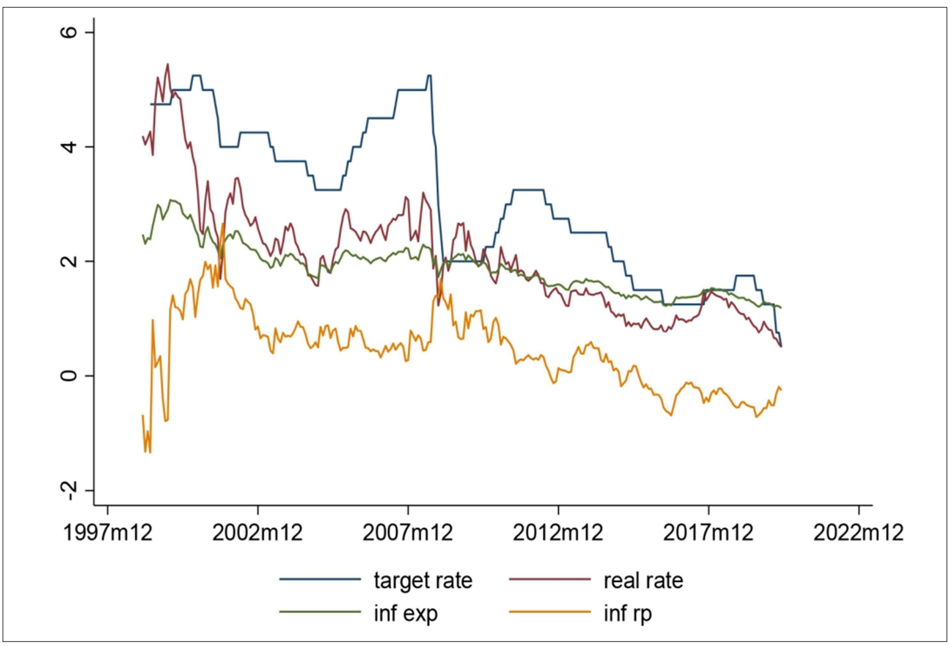
Second, we report on additional variations of the main regression analysis reported in Table 1 of the main text. In Table C2, the long-term interest rates are 5-year rates rather than 10-year rates. In Table C3, the long-term interest rates are 15-year rates rather than 10-year rates. In Table C4, the right-hand-side variable is effective rates rather than target rates. As mentioned in the main text, we obtain similar results from these variations.

The remainder of this appendix concerns a vector autoregression analysis, which is an alternative to the single equation analysis presented in Section 4.1 of the main text. We start with augmented Dickey-Fuller tests to establish that the time-series we deal with are integrated of a right order. Table C5 reports that all our time series are integrated of order 1 according to augmented Dickey-Fuller tests.

We then estimate vector error correction models. Model specifications are shown in Tables C6-C10. Our vector error correction models include 4 variables, three of which are the industrial production index, the consumer price index, and the target rate. The 4th variable is one of 10-year rates, real expectations, real term premia, inflation expectations, and inflation risk premia. We have determined the lag length and the number of cointegrating equations for the first model (where the 4th variable is 10-year rates) and applied the same numbers for the remaining models. The lag length of 1 has been determined via the Schwarz-Bayesian information criterion, and the number of cointegrating equations (which is 2) has been determined via the Johansen procedure. The vector error correction models have been estimated by the maximum likelihood method.

After estimating the vector error correction models, we have calculated the impulse response function. The response of 10-year rates and their components to the orthogonalized target rate shock is reported in Table C11. For the ease of interpretation, we have normalized the response by dividing by the response of target rate to the orthogonalized target rate shock. Thus, the response of 0.231 reported at the top of column (1) can be interpreted as follows: when the target rate moves by 1 percent point, the 10-year rate moves by 0.231 percent point. Given this normalization, the numbers shown in Table C11 can be easily compared to the numbers in Table 1 of the main text. One can verify that the vector autoregression analysis produces very similar results to those of the single equation analysis. For example, the all-period response of 10-year rates in Table 1 is 0.254, and the corresponding number in Table C11 is 0.231.

[Figure C1] Target Rates and Components: The 3-way Decomposition



Notes: Target rates refer to the interest rate target set by the Bank of Korea. The components (of 10-year interest rates) are obtained from the affine term structure model of Adrian et al. (2013). The components identified are real interest rate (real rate), inflation expectations (inf exp), and inflation risk premia (inf rp).

[Table C1] Responses of 10-Year Rates and Components to Target Rates: The 3-way Decomposition

	all period (1)	pre-GFC (2)	post-GFC (3)	(3)-(2)
changes in				
10-year rates	0.254*	0.350	0.008	-0.342
	[1.72]	[1.49]	[0.06]	[-1.02]
real rates	0.379***	0.509***	0.020	-0.490**
	[3.97]	[3.48]	[0.18]	[-2.28]
inflation expectations	0.102***	0.138***	0.006	-0.132*
	[3.17]	[2.75]	[0.17]	[-1.82]
inflation risk premia	-0.227*	-0.296	-0.017	0.279
	[-1.88]	[-1.53]	[-0.15]	[1.02]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the changes in target rates, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and Z_t is target rates. Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C2] Responses of 5-Year Rates and Components to Target Rates

	all period (1)	pre-crisis (2)	post-crisis (3)	(1) vs. (3)
changes in				
5-year rates	0.472***	0.615***	0.088	-0.527*
	[3.59]	[2.99]	[0.62]	[-1.78]
real rate expectations	0.220***	0.269***	0.084**	-0.185**
	[5.96]	[4.70]	[2.07]	[-2.22]
real term premia	0.377***	0.501***	0.029	-0.472**
	[3.83]	[3.28]	[0.26]	[-2.14]
inflation expectations	0.138***	0.185***	0.012	-0.174*
	[3.11]	[2.67]	[0.25]	[-1.73]
inflation risk premia	-0.263**	-0.340**	-0.037	0.303
	[-2.43]	[-1.96]	[-0.36]	[1.24]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the changes in target rates, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and Z_t is target rates. Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C3] Responses of 15-Year Rates and Components to Target Rates

	all period (1)	pre-crisis (2)	post-crisis (3)	(1) vs. (3)
changes in				
15-year rates	0.048***	0.053***	0.033***	-0.020
	[8.66]	[6.00]	[6.37]	[-1.60]
real rate expectations	0.039***	0.043***	0.027***	-0.016
	[6.60]	[4.49]	[5.29]	[-1.18]
real term premia	0.017***	0.019***	0.011***	-0.008
	[5.57]	[3.86]	[4.05]	[-1.16]
inflation expectations	0.007	0.009	0.001	-0.008
	[1.56]	[1.30]	[0.28]	[-0.80]
inflation risk premia	-0.015***	-0.018**	-0.006	0.012
	[-3.41]	[-2.54]	[-1.52]	[1.18]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the changes in target rates, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and Z_t is target rates. Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C4] Responses of 10-Year Rates and Components to Effective Rates

	all period (1)	pre-crisis (2)	post-crisis (3)	(1) vs. (3)
changes in				
10-year rates	0.237*	0.308	-0.011	-0.319
	[1.88]	[1.59]	[-0.08]	[-1.03]
real rate expectations	0.106***	0.126***	0.033	-0.093*
	[5.04]	[4.01]	[1.24]	[-1.80]
real term premia	0.171***	0.228***	-0.044	-0.272*
	[2.86]	[2.58]	[-0.57]	[-1.86]
inflation expectations	0.082***	0.105**	-0.002	-0.107
	[3.01]	[2.57]	[-0.06]	[-1.61]

inflation risk premia	-0.123	-0.151	0.002	0.153
	[-1.19]	[-0.94]	[0.02]	[0.60]

Notes: Columns (1), (2), and (3) report the responsiveness of the variables listed in the left-most column to the changes in target rates, i.e. the estimate of β in the following equation:

$$Y_t - Y_{t-1} = \alpha + \beta(Z_t - Z_{t-1}) + \varepsilon_t$$

where Y_t is the variable listed in the left-most column and Z_t is effective rates. Numbers inside square brackets are t statistics. For column (1), the estimation is based on the all period (May 1999 ~ May 2020); for column (2), the pre-crisis period (May 1999 ~ Feb 2009); for column (3), the post-crisis period (March 2009 ~ May 2020). The last column reports the test of the null hypothesis that the estimates shown in columns (2) and (3) are identical. The difference is reported first and then the t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C5] Augmented Dickey-Fuller Test

	Null hypothesis	Test	Result
IP	The level is a unit root process	ADF, trend	Not reject at 10%
	The 1st difference is a unit root process	ADF	Reject at 1%
CPI	The level is a unit root process	ADF, trend	Not reject at 10%
	The 1st difference is a unit root process	ADF	Reject at 1%
TR	The level is a unit root process	ADF	Not reject at 10%
	The 1st difference is a unit root process	ADF	Reject at 1%
IR	The level is a unit root process	ADF	Not reject at 10%
	The 1st difference is a unit root process	ADF	Reject at 1%
RE	The level is a unit root process	ADF	**Reject at 5%, not reject at 1%**
		DFGLS	Not reject at 10%
	The 1st difference is a unit root process	DFGLS	Reject at 1%
RTP	The level is a unit root process	ADF	**Reject at 1%**
		DFGLS	Not reject at 10%
	The 1st difference is a unit root process	DFGLS	Reject at 1%
IE	The level is a unit root process	ADF	Not reject at 10%
	The 1st difference is a unit root process	ADF	Reject at 1%
IRP	The level is a unit root process	ADF	Not reject at 10%
	The 1st difference is a unit root process	ADF	Reject at 1%

Notes: IP is the industrial production index, CPI is the consumer price index, TR is target rates, IR is 10-year rates, RE is real rate expectations, RTP is real term premia, IE is inflation expectations, and IRP is inflation risk premia. Further description of each series is presented in Table A1 of Appendix A. In the test column, ADF refers to the augmented Dickey-Fuller test without a trend term, "ADF, trend" refers to the augmented Dickey-Fuller test with a trend term, and DFGLS refers to the Dickey-Fuller GLS test.

[Table C6] VECM Estimates: Including 10-year interest rates

A. Main equation

	D.IP	D.CPI	D.TR	D.IR
CE1	-0.135***	0.022**	2.069***	0.615
	[-3.72]	[2.14]	[5.59]	[0.83]
CE2	0.184***	-0.027*	0.615	-1.081
	[3.80]	[-1.93]	[0.83]	[-1.10]
constant	0.001	0.002***	<0.001	<0.001
	[0.55]	[6.70]	[<0.01]	[<0.01]

B. Cointegrating equation

	CE1	CE2
IP	1	0
CPI	0	1
TR	-0.099	-0.067
	[-1.12]	[-1.01]
IR	0.286***	0.213***
	[4.70]	[4.62]
constant	-5.26	-5.092

Notes: IP is the industrial production index, CPI is the consumer price index, TR is target rates, IR is 10-year rates, RE is real rate expectations, RTP is real term premia, IE is inflation expectations, and IRP is inflation risk premia. Further description of each series is presented in Table A1 of Appendix A. D is a difference operator. That is, D.x is x(t) - x(t-1). CE1 and CE2 refer to the first and second cointegrating equations, respectively. The t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C7] VECM Estimates: Including real rate expectations

A. Main equation

	D.IP	D.CPI	D.TR	D.RE
CE1	-0.138***	0.02**	1.903***	0.473***
	[-3.74]	[2.02]	[5.07]	[3.19]
CE2	0.192***	-0.028*	-2.456***	-0.737***
	[3.74]	[-1.89]	[-4.70]	[-3.56]
constant	0.001	0.002***	<0.001	<0.001
	[1.18]	[8.02]	[<0.01]	[0.03]

B. Cointegrating equation

	CE1	CE2
IP	1	0
CPI	0	1

TR	-0.201***	-0.141***
	[-3.29]	[-3.13]
RE	1.762***	1.296***
	[6.96]	[6.96]
constant	-5.36	-5.156

Notes: IP is the industrial production index, CPI is the consumer price index, TR is target rates, IR is 10-year rates, RE is real rate expectations, RTP is real term premia, IE is inflation expectations, and IRP is inflation risk premia. Further description of each series is presented in Table A1 of Appendix A. D is a difference operator. That is, D.x is x(t) - x(t-1). CE1 and CE2 refer to the first and second cointegrating equations, respectively. The t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C8] VECM Estimates: Including real term premia

A. Main equation

	D.IP	D.CPI	D.TR	D.RTP
CE1	-0.149***	0.022**	1.724***	0.987**
	[-3.95]	[1.99]	[4.52]	[2.35]
CE2	0.21***	-0.028*	-2.159***	-1.662***
	[3.98]	[-1.86]	[-4.03]	[-2.82]
constant	0.001	0.002***	<0.001	<0.001
	[1.01]	[8.06]	[<0.01]	[<0.01]

B. Cointegrating equation

	CE1	CE2
IP	1	0
CPI	0	1
TR	-0.089**	-0.059**
	[-2.39]	[-2.16]
RTP	0.701***	0.517***
	[8.50]	[8.54]
constant	-4.96	-4.861

Notes: IP is the industrial production index, CPI is the consumer price index, TR is target rates, IR is 10-year rates, RE is real rate expectations, RTP is real term premia, IE is inflation expectations, and IRP is inflation risk premia. Further description of each series is presented in Table A1 of Appendix A. D is a difference operator. That is, D.x is x(t) - x(t-1). CE1 and CE2 refer to the first and second cointegrating equations, respectively. The t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C9] VECM Estimates: Including inflation expectations

A. Main equation

	D.IP	D.CPI	D.TR	D.IE
CE1	-0.137***	0.02**	1.99***	0.292
	[-3.76]	[1.97]	[5.39]	[1.59]
CE2	0.19***	-0.025*	-2.542***	-0.509**
	[3.83]	[-1.73]	[-5.02]	[-2.02]
constant	0.001	0.002***	<0.001	<0.001***
	[0.79]	[7.42]	[<0.01]	[6.15]

B. Cointegrating equation

	CE1	CE2
IP	1	0
CPI	0	1
TR	-0.085	-0.056
	[-1.60]	[-1.39]
IE	1.022***	0.757***
	[6.13]	[6.01]
constant	-6.034	-5.661

Notes: IP is the industrial production index, CPI is the consumer price index, TR is target rates, IR is 10-year rates, RE is real rate expectations, RTP is real term premia, IE is inflation expectations, and IRP is inflation risk premia. Further description of each series is presented in Table A1 of Appendix A. D is a difference operator. That is, D.x is x(t) - x(t-1). CE1 and CE2 refer to the first and second cointegrating equations, respectively. The t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C10] VECM Estimates: Including inflation risk premia

A. Main equation

	D.IP	D.CPI	D.TR	D.IRP
CE1	-0.138***	0.031***	2.013***	-0.933*
	[-3.58]	[2.82]	[5.08]	[-1.83]
CE2	0.185***	-0.039***	-2.729***	0.91
	[3.66]	[-2.71]	[-5.25]	[1.36]
constant	0.001	0.002***	<0.001	<0.001
	[1.61]	[8.07]	[0.01]	[<0.01]

B. Cointegrating equation

	CE1	CE2
IP	1	0
CPI	0	1

TR	-0.043	-0.026
	[-1.01]	[-0.82]
IRP	0.505***	0.373***
	[6.27]	[6.15]
constant	-4.514	-4.534

Notes: IP is the industrial production index, CPI is the consumer price index, TR is target rates, IR is 10-year rates, RE is real rate expectations, RTP is real term premia, IE is inflation expectations, and IRP is inflation risk premia. Further description of each series is presented in Table A1 of Appendix A. D is a difference operator. That is, D.x is x(t) - x(t-1). CE1 and CE2 refer to the first and second cointegrating equations, respectively. The t statistic is reported inside square brackets. ***, **, * represent significance at the 1%, 5% and 10% levels, respectively.

[Table C11] Summary of Impulse Response Analysis

		all period (1)	pre-GFC (2)	post-GFC (3)	(3) - (2)
changes in					
10-year rates	at t	0.231	0.366	0.054	-0.312
	at $t+3$	0.231	0.229	0.098	-0.131
	at $t+6$	0.231	0.137	0.130	-0.007
real rates expectations	at t	0.147	0.192	0.043	-0.149
	at $t+3$	0.124	0.126	0.065	-0.061
	at $t+6$	0.116	0.086	0.065	-0.021
real term premia	at t	0.326	0.474	0.000	-0.474
	at $t+3$	0.217	0.224	0.033	-0.191
	at $t+6$	0.171	0.112	0.043	-0.068
inflation expectations	at t	0.109	0.163	0.011	-0.153
	at $t+3$	0.093	0.092	0.022	-0.070
	at $t+6$	0.085	0.046	0.033	-0.013
inflation risk premia	at t	-0.282	-0.480	0.086	0.566
	at $t+3$	-0.145	-0.079	0.108	0.186
	at $t+6$	-0.069	0.059	0.118	0.059

Notes: The table below reports the response of the variable shown in the first column to the orthogonalized target rate shock. Numbers have been normalized by dividing by the response of the target rate to the orthogonalized target rate shock. For example, the first number in column (1), 0.231, is the response of the 10-year rate at t to the orthogonalized target rate shock to the response of target rate to the same shock.

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한국의 통화정책과 장기이자율: 분해분석*

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논문초록

본 논문에서는 1999년부터 2020년까지 기간 한국의 명목 국채수익률 자료를 활용하여 선형기간구조모형을 추정하고 장기이자율을 네 가지의 구성요소(실질단기이자율 기대, 실질기간 프리미엄, 인플레이션 기대, 인플레이션위험 프리미엄)로 분해하였다. 분해 결과를 바탕으로 장기이자율과 구성요소가 통화정책의 변화에 어떻게 반응하는가를 검토하였다. 검토결과 장기이자율은 통화정책의 변화에 반응하지만 반응정도는 2008년 글로벌금융위기 이후 약화되었음을 확인하였다. 반응정도의 약화는 상당부분 실질기간프리미엄으로 인한 것이다. 이러한 패턴을 미국 데이터에서 발견되는 패턴과 비교하였고 패턴을 어떻게 설명할 수 있을지 논하였다.

핵심 주제어: 통화정책, 장기이자율, 실질기간 프리미엄, 인플레이션 기대, 인플레이션위험 프리미엄

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