

THE BOND RISK PREMIA: ANALYTICAL EVIDENCE FROM TURKEY

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2024

THE BOND RISK PREMIA: ANALYTICAL EVIDENCE FROM TURKEY

Thesis submitted to the
Institute for Graduate Studies in Social Sciences
in partial fulfillment of the requirements for the degree of

Doctor of Philosophy
in
Management

by
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Boğaziçi University

2024

The Bond Risk Premia: Analytical Evidence from Turkey

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ABSTRACT

The Bond Risk Premia: Analytical Evidence from Turkey

The term structure of government bond yields provides valuable and predictive information. In the first part of this dissertation, I decompose the Turkish government nominal bond yields into their risk-neutral rate and term premium components, and then investigate the local and global drivers of each component separately. One key finding is that spillovers from the U.S. bond market to the Turkish local bond market mainly work through the channel from the U.S. term premia to the Turkish term premia, and solely during the pre-November 2017 period. Regarding the impacts of unconventional bond purchases by the Central Bank of Turkey (CBT), along with the bond purchases by local banks under the scope of the “Asset Ratio” (AR) regulation and the liraization strategy, on the yield components, the results reveal the operations of signaling and duration risk channels in lowering local risk-neutral rates and term premia, respectively. In the second part, I decompose the breakeven inflation (BEI) rates in Turkey. An important finding is that the local bond market does not price in the differential liquid premium between nominal and inflation-indexed bonds. Next, I explore the main determinants of the inflation risk premium component. Notably, changes in the Brent crude oil prices exert a strong positive impact on variations in the country’s inflation risk premia. Moreover, there is a statistically significant negative effect of the liraization strategy on inflation risk premia dynamics. Finally, I conduct an ex ante cost-benefit analysis of issuing nominal versus inflation-indexed bonds from the perspective of the Turkish Treasury.

ÖZET

Tahvil Risk Primleri: Türkiye’den Analitik Kanıtlar

Devlet tahvil getiri eğrisi değerli ve tahmine yönelik bilgiler sağlar. Bu tezin ilk bölümünde, Türk devlet tahvili getirilerini risk-nötr oran ve vade primi bileşenlerine ayırıyor ve ardından her bir bileşenin yerel ve küresel etkenlerini ayrı ayrı araştırıyorum. Temel bulgulardan biri, ABD tahvil piyasasından Türk yerel tahvil piyasasına etkileşimlerin esas olarak ABD vade primlerinden Türk vade primleri kanalıyla ve sadece Kasım 2017 öncesi dönemde gerçekleştiğidir. Türkiye Cumhuriyeti Merkez Bankası tarafından gerçekleştirilen geleneksel olmayan tahvil alımları ile yerel bankaların “Aktif Rasyosu” düzenlemesi ve liralasma stratejisi kapsamındaki tahvil alımlarının getiri bileşenleri üzerindeki etkilerine ilişkin sonuçlar, yerel risk-nötr oranları ve vade primlerini düşürmede sırasıyla sinyalizasyon ve durasyon riski kanallarının işleyişlerini ortaya koymaktadır. İkinci bölümde Türkiye’deki başabaş enflasyon oranlarını araştırıyorum. Önemli bir bulgu, yerel tahvil piyasasının nominal ve enflasyona endeksli tahviller arasındaki likidite prim farkını fiyatlamamasıdır. Sonrasında enflasyon risk priminin ana belirleyicilerini araştırıyorum. Özellikle Brent ham petrol fiyatlarındaki değişimler ülkenin enflasyon risk primlerindeki değişimler üzerinde güçlü bir pozitif yaratmaktadır. Ayrıca liralasma stratejisinin enflasyon risk primi dinamikleri üzerinde istatistiksel olarak önemli bir negatif etkisi bulunmaktadır. Son olarak, Türkiye Cumhuriyeti Hazine ve Maliye Bakanlığı perspektifinden nominal ve enflasyona endeksli tahvillerin ihraç edilmesine yönelik ön maliyet-fayda analizi yapıyorum.

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ACKNOWLEDGMENTS

I would like to express my heartfelt gratitude to my thesis advisor, Prof. Nesrin Okay Akman, for her guidance, support, and unwavering encouragement throughout my research journey. Her openness and responsiveness fostered an empowering environment that allowed me to navigate challenges with confidence.

I am also grateful to the members of my dissertation committee, Assoc. Prof. Asunur Cezar, Prof. Ayşe Dilara Altıok Yılmaz, Assoc. Prof. Yavuz Acar, and Prof. Fatih Yılmaz, for their insightful comments and constructive feedback.

Finally, I would like to thank my family members for their unconditional love, motivation, and support during this process and in my entire life.



Dedicated to my beloved family

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ABBREVIATIONS

ACM	Adrian, Crump, and Moench
ADF	Augmented Dickey-Fuller
AR	Asset Ratio
ATM	At-the-Money
ATSM	Affine Term Structure Model
BCFF	Blue Chip Financial Forecasts
BEI	Breakeven Inflation
BFGS	Broyden-Fletcher-Goldfarb-Shanno
BIS	Bank of International Settlements
BOE	Bank of England
BOJ	Bank of Japan
BP	Basis Point
BRSA	Banking Regulation and Supervision Agency
CBT	Central Bank of Turkey
CDS	Credit Default Swap
DE	Differential Evolution
ECB	European Central Bank
EH	Expectations Hypothesis
EMBI+	Emerging Markets Bond Index Plus
FED	Federal Reserve Bank
GDDS	Government Domestic Debt Securities
GFC	Global Financial Crisis
GRG	Generalized Reduced Gradient

HW	Hamilton and Wu
IRR	Internal Rate of Return
IV-GMM	Instrumental Variables Generalized Method of Moments
JSZ	Joslin, Singleton, and Zhu
L-BFGS-B	Limited BFGS with Boundary Constraints
LPH	Liquidity Preference Hypothesis
MAE	Mean Absolute Error
ML	Maximum Likelihood
MCSE	Minimum Chi-Square Estimation
NS	Nelson and Siegel
NSS	Nelson-Siegel-Svensson
OLS	Ordinary Least Squares
OMO	Open Market Operations
OMT	Outright Monetary Transactions
PC	Principal Component
PEH	Pure Expectations Hypothesis
QE	Quantitative Easing
RE	Rational Expectations
RMSE	Root Mean Squared Error
SMP	Securities Markets Programme
SPF	Survey of Professional Forecasters
UIF	Unemployment Insurance Fund
VAR	Vector Autoregressive

CHAPTER 1

DETERMINANTS OF TERM PREMIA AND RISK-NEUTRAL RATES

1.1 Introduction and main results

The term structure of interest rates, commonly known as the yield curve, maps the relationship between prevailing interest rates and their respective maturities. If these rates represent returns of zero-coupon bonds, they are defined as zero-coupon or spot rates, and the corresponding term structure is called the zero-coupon or spot yield curve. Since zero-coupon bonds are not available across the entire maturity spectrum, the spot yield curve is constructed from the prices of existing zero-coupon and fixed-coupon bonds.

The term structure of government bond yields, expressed in terms of zero-coupon rates, draws a lot of attention and interest from policymakers, academics, and financial market practitioners because it is a major pricing tool and contains an extremely valuable and predictive information. To begin with, it is a key input into valuation of financial securities, thus playing a vital role in the process of asset pricing. For instance, to determine fair value of an existing government bond or theoretical price of a new issue, it is ideal to use zero-coupon yield curve. Similarly, when pricing a corporate bond or inferring its implied spread, it acts as a benchmark. Another prime application area is that financial institutions conduct parallel and non-parallel shifts in the zero-coupon yield curve to identify their interest rate risk exposures, devise active trading strategies, or implement passive portfolio strategies (Ho, 1992; Nawalkha , Soto, & Beliaeva, 2005). Beyond these widespread uses, several value-at-risk and fixed income derivative valuation models also rely on the volatility estimates derived from the spot rates. Ultimately, the term structure of

interest rates is a key instrument to predict the market expectations of future short-term interest rates. To effectively disentangle and rationalize this crucial information, it is essential to comprehend the standard two-part decomposition of a long-term yield. Empirical studies reveal that the yield on a long-term bond is the sum of the average expected level of short-term interest rates over the maturity of the bond and a time-varying risk premium (Gürkaynak & Wright, 2012). The first component, known as the expectations component of interest rates or risk-neutral rate, provides information about market expectations with regards to monetary policy outlook, inflation and economic growth (Ceballos & Romero, 2016), while the second component, so-called the term premium, represents the compensation demanded by investors for the additional risks associated with holding a long-term bond instead of investing in a series of short-term bonds. Term premium is primarily a hedge against the risk that the future path of short-term interest rates over the life of the long-term bond does not evolve as expected (i.e., interest rate risk or duration risk). The deviations from expectations mainly originate from macroeconomic uncertainty about future inflation and growth. The term premium also reflects the compensations for risks associated with possibility of experiencing losses from credit events (i.e., sovereign credit risk), fluctuations in the market liquidity conditions (i.e., liquidity risk) and adverse movements in the exchange rate (i.e., currency risk). In this context, given that risk-averse investors require additional risk premiums to hold a long-term bond, uncovering the information related to the market expectations of future short-term interest rates from the yield curve necessitates a proper decomposition of zero-coupon bond yields.

The breakdown of bond yields into their risk-neutral rate and term premium components also serves two purposes. First, it allows for the assessment of the

relative contribution of each component to dynamic evolution of yield curves over time. Second, and more importantly, quantifying each component enables the exploration of their local and global drivers. In this regard, the main objective of this chapter is to decompose the Turkish government nominal bond yields into their risk-neutral rate and term premium components and then investigate the main determinants of each component separately across the sample period from February 2005 to June 2022 and two sub-periods. To the best of my knowledge, this study represents the first attempt to comprehensively analyze the key influencing factors of both yield components in the context of Turkey, using the same set of explanatory variables and considering different sub-periods.

The analysis starts with the construction of end-of-month zero-coupon yield curves over the sample period. For this purpose, I employ the Nelson-Siegel-Svensson (NSS) model and an optimization objective function in which each squared price error is weighted by the inverse of the square root of bond duration. This weighting scheme, proposed by Gimeno and Nave (2006), strikes a much better balance between the short-maturity and the long-maturity bonds in the optimization process, in contrast to the most commonly used weighting scheme that employs the inverse of bond duration as a weight for each price error. In estimating all six parameters of the NSS model, given that the optimization problem is non-convex and exhibits multiple local optima, I utilize not only four distinct non-linear optimization algorithms but also three different parameter initialization approaches to achieve a better quality of fit. The in-sample Root Mean Squared Error (RMSE) statistics for price errors in different maturity buckets indicate that my curve fitting methodology provides fairly accurate fits during the sample period.

To decompose the zero-coupon nominal government bond yields into risk-neutral rate and term premium components, I employ the model developed by Adrian, Crump, and Moench (ACM, 2013). This model uses a three-step linear regression approach and therefore allows for a simple and computationally fast estimation procedure. In the monthly application of the ACM model, I choose a five-factor specification, as supported by the model specification test results. The empirical results indicate that while yields are explained almost entirely by the first three principal components (PCs), expected excess bond returns are explained by all five PCs, with notable contributions from the higher-order PC4 and PC5. Moreover, the market price of risk associated with each factor is significantly priced in the five-factor model. The variance decomposition analysis of nominal yields in level form reveals that the term premium component is an important driver of nominal yields, accounting for at least 39.9% of the yield variation as of the 3-year maturity and becoming the dominant factor at the 5-year maturity. The same analysis in first-difference form suggests that term premium component emerges as the primary contributor to changes in nominal yields starting from the 3-year maturity.

Armed with the estimates of term premium and risk-neutral components of the Turkish government nominal bond yields, I then proceed to examine their determinants through a series of regressions. The analysis is conducted for the full sample period and two sub-periods to identify changes in both the statistical significance and magnitude of the effects of explanatory variables over time. I choose to split the sample in November 2017 as Turkey started to adopt unorthodox economic and financial policies. These include unconventional liquidity measures implemented by the CBT to tighten monetary conditions, the unconventional bond purchases by the CBT between April and July 2020 and post-April 2022, the

implementation of the AR regulation by the Banking Regulation and Supervision Agency (BRSA) between April and December 2020, the monetary easing cycle between September and December 2021 despite a surge in domestic inflation, the launch of a state-backed foreign currency-protected deposit scheme in December 2021, and changes in collateral requirements for banks under the liraization strategy in May and June 2022.

The selection of explanatory variables for the baseline model is based on both theoretical and empirical considerations. In order to examine the effects of sovereign credit risk, exchange rate risk, market liquidity risk, inflation uncertainty, international oil prices, and global risk perceptions on local yield components, I incorporate appropriate proxies for these factors into the model. Furthermore, to understand the international spillover effects from major advanced economies, particularly the United States, on the Turkish local bond market, I employ both the U.S. term premium and risk-neutral components as explanatory variables. Instead of relying solely on the U.S. bond yield, using its components allows for a more detailed analysis of spillover effects on both the Turkish term premium and risk-neutral components. Additionally, in an effort to assess the individual impact of the CBT's unconventional bond purchases, the BRSA's AR regulation, and the CBT's liraization strategy on each Turkish yield component during the post-November 2017 period, I construct time dummy variables corresponding to each of these policies. However, the simultaneous inclusion of these three dummies in a time-series regression model gives rise to multicollinearity issues due to their temporal overlap, resulting in imprecise estimates of individual effects. To address this challenge, I introduce the "Aggregate Policy" time dummy variable into the baseline model to capture their potential joint impact on each local yield component.

The regression results provide several important and novel empirical findings. In term premia regressions, the results indicate that there is a strong and statistically significant positive relationship between changes in the term premia and changes in the sovereign credit risk, as proxied by changes in the Turkish 5-year credit default swap (CDS) spread, across the full sample period and two sub-periods. Notably, the impact of changes in the sovereign CDS spread is more pronounced for longer maturities, as evidenced by a 100 basis points (bp) increase leading to an 82 bp increase in the 2-year term premium and a 109 bp increase in the 5-year term premium over the entire sample period. Focusing on the impact of changes in exchange rate risk, where exchange rate risk is represented by the implied volatility of the 6-month at-the-money (ATM) USD/TRY option contract, on changes in local term premia, the regression coefficients reveal a positive relationship for the full sample period and the first sub-period. Specifically, a 1% increase in the implied volatility is associated with around 5 bp and 9 bp increases in the term premia for the respective periods. However, the diminishing statistical significance in the second sub-period highlights the evolving dynamics in the contribution of exchange rate risk premium to term premium component. When analyzing the impact of changes in market liquidity conditions on changes in local term premia, I identify a statistically significant relationship only for the 5-year maturity in the pre-November 2017 period, and no such relationship is observed for any maturity in the post-November 2017 period. As for the spillover effects from the U.S. government bond market, the findings suggest that changes in the U.S. term premia have a statistically significant impact on changes in the Turkish term premia in the full sample period. Importantly, the magnitude of the impact declines with maturity. For instance, a 100 bp shock in the U.S. 2-year term premium leads to a 114 bp increase in the Turkish 2-year term

premium, while the same shock in the U.S. 5-year term premium prompts a 51 bp increase in the Turkish 5-year term premium. The comparison of two sub-periods shows that the significance of spillover effects from the U.S. term premia to local term premia over the entire sample period is mainly driven by the pre-November 2017 period. In the post-November 2017 period, characterized by Turkey's shift to unconventional economic and financial policies, this spillover channel is no longer functional. With regard to spillover effects from the U.S. risk-neutral rates to local term premia, this channel is relevant only for the 2-year maturity during the full sample and the first sub-periods. The regression results also suggest that changes in the VIX index have no statistically significant impact on changes on local term premia in any of the examined sample periods. Analyzing the effect of fluctuations in Brent crude oil prices on changes in local term premium component indicates that there is a positive and statistically significant relationship throughout the full sample period. The heightened sensitivity to oil price fluctuations is primarily associated to the second sub-period, which is marked by Turkey's experience with high and volatile inflation during the implementation of unconventional policy measures. Another notable finding is that the exchange rate and inflation surprise measures, which are used as proxies for domestic inflation uncertainty, do not contribute significantly to variations in local term premia over the considered time periods. Focusing on the impact of the "Aggregate Policy" in the post-November 2017 period, I observe a negative and statistically significant effect on local term premia across various maturities. Importantly, longer maturities exhibit a stronger response, as evidenced by the 2-year and 5-year local term premiums declining by 82 bp and 169 bp, respectively. This empirically supports the notion that the unconventional bond purchases by the CBT for its Open Market Operations (OMO) portfolio, along

with the bond purchases by local banks under the scope of AR regulation and within the context of the liraization strategy, jointly reduce term premia through the duration risk channel.

Turning to the risk-neutral rate regressions, the results over the entire sample period suggest that positive changes in the sovereign CDS spread prompt investors to revise their average expectations of future short-term interest rates upward, with the effect diminishing as the maturity increases. A 100 bp increase in the sovereign CDS spread elevates the 2-year and 5-year risk-neutral rates by 36 bp and 26 bp, respectively, in the full sample period. This relationship weakens during the adoption of unorthodox policies in the post-November 2017 period. In the context of exchange rate risk, implied volatility changes in the 6-month ATM option are positively and significantly associated with changes in risk-neutral rates across all examined periods. Given the strong exchange rate pass-through to domestic inflation in Turkey, this result is quite intuitive in the sense that fluctuations in market expectations about future exchange rate movements result in revisions in inflation expectations, which subsequently trigger changes in the expected path of future short-term interest rates. It is also evident that the sensitivity of changes in the risk-neutral rates to changes in the implied volatility decreases with maturity. A 1% increase in the implied volatility corresponds to an 11 bp increase in the 2-year risk-neutral rate and a 7 bp increase in the 5-year risk-neutral rate over the full-sample period. With regard to liquidity risk effects, changes in the employed market illiquidity measure, which is a price-space version of the illiquidity indicator proposed by Hu, Pan, and Wang (2013) and tends to spike during market crises, have a significant positive effect on changes in local risk-neutral rates in the full sample period and the first sub-period. When this indicator surges amid market turmoil, investors may preemptively anticipate that the

central bank will take actions to tighten the monetary conditions in order to ensure financial stability. This anticipation, in turn, leads to an increase in the average expected level of future short-term interest rates. Furthermore, the corresponding impact on risk-neutral rates declines with increasing maturity. Shifting to spillover effects from the U.S. term premia to the Turkish risk-neutral rates, the results for the full sample and pre-November 2017 periods suggest that changes in the U.S. term premia have a positive and statistically significant influence on changes in local risk-neutral rates for shorter maturities. For instance, across the full sample, a 100 bp rise in the U.S. 2-year term premium corresponds to an 89 bp increase in the local 2-year risk-neutral rate, while the same change in the U.S. 3-year term premium results in a 44 bp increase in the local 3-year risk-neutral rate. Nevertheless, in the post-November 2017 period, changes in the U.S. term premia do not significantly affect changes in local risk-neutral rates, mirroring their effects on changes in local term premia during the same period. Regarding the spillover effects from the U.S. risk-neutral rates to local risk-neutral rates, no statistical evidence is found to support the significance of this channel in any of the sample periods. The findings also demonstrate that there is no association between changes in the VIX index and changes in local risk-neutral rates across the considered maturities and sample periods. On the other hand, I document that fluctuations in the Brent crude oil prices have a strong positive impact on changes in the Turkish risk-neutral rates. This finding is well consistent with previous empirical findings that changes in oil prices have significant effect on agents' inflation expectations in Turkey. Considering that inflation expectations are a key component of risk-neutral rates, I conclude that the Brent crude oil prices influence local risk-neutral rates potentially through the inflation expectations channel. The regression results also suggest that the impact of

oil price shocks on changes in the risk-neutral rates decreases with increasing maturity, in line with theoretical predictions. Moving to the effects of surprise measures, inflation and exchange rate surprises have positive and statistically strong impacts on changes in local risk-neutral rates in the full sample and pre-November 2017 period. This indicates that market participants adjust their expectations of future short-term interest rates in the direction of these surprises over the specified periods. Regarding the impact of the “Aggregate Policy” on changes in local risk-neutral rates, the corresponding coefficients are negative and statistically significant. Importantly, the response of changes in the risk-neutral rates to the “Aggregate Policy” is more pronounced for shorter maturities, providing empirical evidence for the operation of signaling channel in the context of the central bank’s unconventional bond purchases. When the “Aggregate Policy” is in effect, the 2-year and 5-year risk-neutral rates decline by 77 bp and 53 bp, respectively, in the second sub period. A comparison of the responses of local term premia and local risk-neutral rates to the “Aggregate Policy” reveals that the impact is greater on local term premia than on local risk-neutral rates across various maturities, with the magnitude of the difference in impacts increasing with maturity.

In this study, to strike a balance between capturing the individual impact of unconventional bond purchases, the AR regulation and the liraization strategy, while still preventing multicollinearity issues due to their temporal overlap, I also employ an alternative model specification where the composite “Aggregate Policy” dummy is replaced with two other time dummies. The first combines the unconventional bond purchases and the AR regulation time dummies, while the second exclusively represents the liraization strategy. The regression results show a significant negative combined impact of the CBT’s unconventional bond purchases and local banks’

bond purchases under the AR regulation on local term premia across the 2-year to 4-year maturities in the second sub-period. While there is lack of detailed information on the maturity profile of the Turkish banks' government bond purchases under the mandate of the AR regulation, this result is consistent with the fact that the CBT's unconventional bond purchases in the second sub-period are mostly concentrated on bonds with 2- to 4-year maturities. Regarding the impact of the liraization strategy, it significantly reduces local term premia for maturities ranging from 3 to 5 years. This is in line with the CBT's June 2022 mandate, which required banks to maintain holdings between 3% and 10% of foreign currency deposits in local currency fixed-coupon government bonds with an original maturity of at least 5 years and a residual maturity of at least 4 years. Focusing on the pattern of coefficients associated with these two time dummy variables indicates that negative changes in local term premia increase with maturity. This provides further empirical evidence for the operation of duration risk channel. Turning to the results of the alternative specification for local risk-neutral rates, the coefficients representing the combined impact of the CBT's unconventional bond purchases and local banks' bond purchases under the AR regulation are negative and decline in magnitude with maturity. However, these coefficients have weaker statistical significance when compared to the coefficients associated with the "Aggregate Policy" variable in the baseline specification, implying only a marginal role of the signaling channel in reducing local bond yields. Lastly, the regression results suggest that the liraization strategy has no impact on changes in local risk-neutral rates, in line with theoretical expectations.

The rest of this chapter is organized as follows. Section 2 presents the literature review. Section 3 outlines the construction of zero-coupon yield curves. Section 4 details the ACM methodology and applies it to decompose bond yields into

their term premium and risk-neutral rate components. Section 5 focuses on the determinants of these yield components. Section 6 summarizes and concludes.

1.2 Literature review

The literature on the term structure of interest rates presents different theories to explain the shape of the yield curve. These theories basically attempt to answer why interest rates change with maturity. The starting point is the Pure Expectations Hypothesis (PEH), which plays a dominant role in the literature by establishing a link between short-term and long-term interest rates. According to the PEH, which is a risk-neutral approach, the expected return from holding a long-term bond until its maturity is equal to the expected return from rolling over a series of short-term bonds with a total maturity equal to that of the long-term bond, meaning that the yield of the long-term bond is the average of current and future expected short-term rates (Gürkaynak & Wright, 2012; Kim & Orphanides, 2007).¹ Equivalently, the PEH postulates that the expected excess return or term premium on the long-term bond over the short-term bonds is essentially zero, implying that any bond investor will be indifferent between purchasing a long-term bond or a sequence of short-term bonds covering the same time period (Rudebusch, 1998).

There also exists a less restrictive Expectations Hypothesis (EH), which is a weaker version of the PEH. The EH states that the term premium is constant over time and independent of bond maturities. A further extension from the EH brings us to the Liquidity Preference Hypothesis (LPH), which retains the assumption of time-invariant term premium component but allows it to monotonically increase with the maturity of the bond. In the context of the LPH, bond investors typically view long-

¹ Jensen's inequality term, which represents the convexity bias, is neglected. Campbell (1986) shows that its impact is empirically irrelevant, especially for maturities below ten years.

maturity bonds as riskier than short-maturity bonds and therefore demand an additional liquidity premium for holding longer maturities as opposed to shorter ones. Under the assumption of rational expectations (RE), the main testable implications of the PEH, EH and LPH using the regression analysis are identical, with the exception of the intercept term which represents the term premium component.² Because of their similarity, the PEH, EH and LPH are frequently referred to as the EH in the term structure terminology (Cuthbertson and Nitzsche, 2004). Within this framework, a significant feature that characterizes the EH is that, assuming a time-invariant term premium component, changes in the long-term interest rates are solely attributed to the shifts in market expectations of future short-term interest rates.

The EH holds an important place in the finance literature and has undergone extensive examinations through various tests and datasets. Much of the empirical studies, such as those of Fama and Bliss (1987), Campbell and Shiller (1991), Bekaert and Hodrick (2001), Cochrane and Piazzesi (2005, 2009) and Sarno, Thornton, and Valente (2007), suggest that different or modified versions of the EH struggle to explain the term structure of interest rates, as the methods used to test their implications often find little or no support for their validity. A commonly

² One well-known test of the EH is developed by Campbell and Shiller (1991). It is based on the idea that the long-term spot rate is equal to the sum of the average of current and expected future short-term interest rates and the term premium. By subtracting the current short-term interest rate from the both sides of this equation, one can easily show that the slope of the term structure (i.e., the term spread) forecasts the weighted average of expected changes in the future short-term interest rates, which is known as perfect foresight spread (PFS), plus the term premium. If the terms of the equation are rearranged to test the predictive power of the term spread, the regression equation takes the following form:

$$PFS_t^{(n,m)} = \alpha + \beta (\text{Term spread}_t^{(n,m)}) + \varepsilon_t$$
, where m and n are maturities of short-term and long-term interest rates, respectively. Under the null of PEH, EH and LPH with the RE assumption, the slope coefficient β in the regression should be equal to one. Regarding the intercept (α), which corresponds to the negative of the term premium component, it is expected to be zero for the PEH, a non-zero constant for the EH, and a different constant in a monotonically increasing manner with maturity of the bond for LPH.

proposed or standard finance explanation for empirical rejections is that the bond risk premium or term premium varies both with time-to-maturity and over time.³

Therefore, it is now widely accepted that the long-term interest rates can be decomposed into expectations of future short-term interest rates (i.e., risk-neutral rate) and time-varying term premium (Crump, Eusepi, & Moench, 2016). The latter component thus serves as a “catch-all” measure for other risk factors beyond the expectations component that influence long-term interest rates (Kim & Wright, 2005).

The time-varying risk premium obscures the source of variation in the long-term bond yields because its existence shows that the long-term interest rates depend not only on the expected future short-term interest rates, as posited by the EH, but also on the term premium priced by investors. In particular, the evidence of term premium is highly critical in the conduct of monetary policy. While central banks have a direct control on the short-term interest rates, they also aim to affect the long-term interest rates by influencing the market expectations of future path of short-term interest rates.⁴ This phenomenon is known as the leverage effect of central banks over the long-term interest rates (Gertler & Karadi, 2015). The use of monetary policy in this manner is completely in line with the EH. However, the time-varying term premium demanded by risk-averse investors complicates the task of central

³ Alternative explanations have also been offered for the failure of the EH. The first one is based on the econometric issues related to persistent variables (Bekaert, Hodrick, & Marshall, 1997; Mankiw & Shapiro, 1986), overlapping errors problem (Campbell & Shiller, 1991), small-sample inference known as the peso problem (Bekaert, Hodrick, & Marshall, 2001). The second potential reason for the empirical rejections is associated with the irrational behavior of investors (i.e., overreaction or underreaction to information), which contradicts with the RE assumption under which the validity of the EH is evaluated (Campbell & Shiller, 1991; Froot, 1989; Hardouvelis, 1994; Shiller, Campbell, Schoenholtz, & Weiss, 1983).

⁴ The transmission process of monetary policy is typically considered to run from the short-term interest rates to the long-term interest rates that impact interest-sensitive components of aggregate demand (Goodfriend, 1998). These components, such as investment and consumption, ultimately determine the levels of economic growth, employment, and inflation.

banks by driving a wedge or breaking the link between the long-term interest rates and the market expectations of futures short rates. Similarly, when central banks attempt to infer the interest rate expectations from the term structure to evaluate their own credibility and efficacy in implementing the monetary policy, it is essential to be aware of the presence of the term premium in the long-term interest rates. The extraction of expectations component could also be relevant for economists and investors while forecasting the path of monetary policy and key macroeconomic variables, such as inflation and output growth. In this vein, for an accurate reading of the market expectations embedded in the long-term interest rates, there is a need for a term structure model that decomposes the long-term interest rates into term premium and expectations components that are not directly observable.

The workhorse model employed by central bankers and investors is called a no-arbitrage affine term structure model (ATSM), in which observed yields are an affine (a constant plus linear) function of risk factors (or state variables), which themselves follow a first-order vector autoregressive (VAR) process. The classical single factor models of Vasicek (1977) and Cox, Ingersoll, and Ross (1985) are the pioneers of the ATSM.⁵ Later, Duffie and Kan (1996) introduce and characterize multi-factor ATSM framework and explain the yields in terms of unobservable or latent risk factors. This framework nests most of the previously developed term structure models in a unified approach.⁶ Additionally, Dai and Singleton (2000) classify ATSMs into subfamilies with respect to the total number of state variables (N) and the number of state variables that drive the conditional variances (m).⁷

⁵ The single factor in these studies is the instantaneous short-term interest rate. It drives the evolution of whole term structure.

⁶ The models of Vasicek (1977), Langetieg (1980), Cox et al. (1985), Chen and Scott (1992), and Longstaff and Schwartz (1992) are the special cases of the model proposed by Duffie and Kan (1996).

⁷ Dai and Singleton (2000) use $A_m(N)$ notation to represent a model that includes N state variables, m of which determine the conditional volatility of the state vector. With $0 \leq m \leq N$, they show that all N -

Most multi-factor ATSMs use the first three PCs of yields as latent risk factors to describe the dynamics of the term structure. Litterman and Scheinkman (1991) find that these factors are closely related to the level, steepness, and curvature of the yield curve and explain almost all of the cross-sectional variation in interest rates. Although such affine models consistently capture the evolution of interest rates in an accurate and parsimonious way, it is often difficult to give an economic interpretation to these factors or relate them to relevant macroeconomic variables (Pericoli & Taboga, 2008). For instance, these models are not capable of analyzing how the yields respond to macro shocks. To gain insight about economic forces that drive the yield curve, some studies estimate ATSMs that use only observable macroeconomic variables as underlying factors.⁸ However, compared to affine models with latent factors, this approach does not generate a close a fit to observed interest rates (Gürkaynak & Wright, 2012).

An alternative strategy is to incorporate macroeconomic variables into a latent factor model. In their seminal work, Ang and Piazzesi (2003) augment traditional three latent yield factors with two observable macro factors in an ATSM.⁹ Combining these factors within the same state space system establishes a joint macro-finance ATSM, allowing for the empirical analysis of the term structure of interest rates from both macroeconomics and finance perspectives. The study shows that the macro factors mainly explain the movements at the short and medium parts of the yield curve, while the latent yield factors still account for the bulk of the

factor admissible ATSMs can be uniquely classified into $N + 1$ non-nested subfamilies. Admissibility of ATSMs is established if the conditional covariances of m state variables are strictly positive. Moreover, for each of these non-nested subfamilies, there is a maximal model that nests all other models within this subfamily.

⁸ Examples of research not relying on latent factors in their affine models are Bernanke, Reinhart, and Sack (2004) and Smith and Taylor (2009).

⁹ These two macro factors are the measures of inflation and real economic activity. They are constructed as the first PCs of two sets of variables that are chosen to represent price changes and economic growth.

variation at the long end. They also present an empirical evidence that addition of macro factors to usual latent factors improves the forecasting performance of an affine model.

As the standard macro-finance term structure models assume that all bond yields are affine function of observable macroeconomic variables and latent yield factors, by inverting the linear mapping between bond yields and macro factors, one can express the macro factors as a linear combination of yields. Therefore, the macro-finance models imply that the yield curve fully spans the information in the macroeconomic variables (Bauer & Rudebusch, 2017). This theoretical argument is referred to as spanning hypothesis in the literature. It does not suggest that macroeconomic variables are unimportant for forecasting future interest rates and (excess) bond returns, but simply says that the yield curve subsumes all the relevant information in the macroeconomic variables, as well as other asset prices, volatilities, and survey expectations (Bauer & Hamilton, 2017). However, the empirical evidence about the spanning hypothesis is still inconclusive. Some studies provide evidences against the theoretical implications of macro spanning. The first evidence reveals that the regression of macro variables on latent yield factors produces low R^2 values, instead of values near one.¹⁰ The second evidence documents that the macro variables still have important predictive power in explaining the (excess) bond returns when controlling the information in the current yield curve.¹¹ On the other hand, Bauer and Rudebusch (2017) claim that even relatively minor yield measurement errors can make it impossible for spanning regressions to accurately detect the presence or absence of a spanning condition in macro-finance term structure models. Similarly, Bauer and Hamilton (2018) claim that small-sample

¹⁰ Examples include Duffee (2013) and Joslin, Priebsch, and Singleton (2014).

¹¹ See Cooper and Priestley (2009), Ludvigson and Ng (2009), and Cieslak and Povala (2015).

issues arising from the violation of strict exogeneity assumption and high persistence of regressors in the predictive regressions of previous studies raise questions about reliability of their results. To address these econometric issues, they propose an alternative parametric bootstrap procedure and then conclude that the evidences against about the spanning hypothesis in the widely cited previous studies are considerably weaker. More recently, Bekaert, Engstrom, and Ermolov (2021) reexamine the spanning hypothesis and rerun the predictive regressions of excess bond returns on some macro level and macro risk factors when conditioning on conventional yield curve factors.¹² Using the bootstrap methodology of Bauer and Hamilton (2018), they show that macro risk factors possess statistically and economically significant explanatory power for excess bond returns, unveiling a new evidence against the theoretical underpinnings of the spanning hypothesis.

The empirical estimation of ATSMs with observable and/or unobservable state variables primarily relies on maximum likelihood (ML) methods. However, there is an important computational challenge when applying ML techniques to estimate the affine model parameters. The likelihood function of these models is high-dimensional and exhibits a relatively flat surface with multiple local optima. Therefore, despite the same affine model specification, differences in the initial parameter values could lead to convergence to different local optima with inequivalent economic implications (Bauer, Rudebusch, & Wu, 2012; Kim & Orphanides, 2012). In response to these challenges, some researchers develop innovative approaches to lessen the burden of computational complexity and overcome convergence issues in the estimation of ATSMs. Joslin, Singleton, and Zhu (JSZ, 2011) suggest that if the yield pricing (or yield fitting) errors are assumed

¹² In this study, the macro risk factors represent the variables that drive the second and higher-order moments of the aggregate demand and supply shocks.

to be conditionally independent of their lagged values (i.e., serially uncorrelated) and also satisfy the consistency condition, then the affine model parameters can be separated into physical and risk-neutral subsets.¹³ They show that ML estimates of physical parameters can be recovered by simple ordinary least squares (OLS) estimator. Consequently, they propose a two-step estimation method. In the first step, the physical parameters that govern the physical distribution of pricing factors are estimated by OLS. In the second step, the estimated parameters of the first step are fixed at their OLS estimates, and the remaining parameters that govern the risk-neutral distribution of pricing factors are estimated by using conventional ML method. Alternatively, Hamilton and Wu (HW, 2012) develop another two-step estimation procedure for affine models. Similar to JSZ (2011), it also assumes that yield pricing errors are serially uncorrelated. The first step of the method involves the estimation of reduced-form parameters via OLS regressions. In the following step, they make use of a well-known minimum-chi-square estimation (MCSE) method and find the structural parameters of an affine model by minimizing the chi-square statistic, which measures quadratic form in the difference between OLS estimates of the reduced-form parameters from the first step and the reduced-form parameters implied by the structural model. HW (2012) demonstrate that, in backing out the structural parameters from the reduced-form estimates, MCSE method is asymptotically equivalent to ML estimation. Although it inherits all the asymptotic benefits of ML estimation, it avoids many computational difficulties encountered with ML methods. In another work, ACM (2013) introduce a three-step linear regression approach that completely eliminates the need for the numerical

¹³ In JSZ (2011), the pricing factors, which are denoted by P_t , are linear combinations (i.e., portfolios) of the zero-coupon bond yields with corresponding weight matrix W . Yield pricing (or measurement) errors are the difference between observed yields (y_t^o) and model-implied yields (y_t). The consistency condition is implied by $\mathbb{P}(W y_t^o = P_t | P_t) = 1$

optimization in the estimation of ATSM parameters, making the estimation process extremely fast. While all prior methods focus on using yields to estimate the model parameters, ACM (2013) utilize excess holding period returns in their estimation methodology.¹⁴ Moreover, they theoretically show that the assumption of serially uncorrelated yield pricing errors, as employed in JSZ (2011) and HW (2012), results in serially correlated return pricing errors, which in turn imply that the predictability in the excess holding period returns is not fully captured by the underlying pricing factors. From this perspective, ACM (2013) argue that the models of JSZ (2011) and HW (2012) are founded on an undesirable assumption. As the estimation framework of ACM (2013) is based on the regression of excess returns, there is no need to impose any distributional assumption on the yield pricing errors. Instead, ACM (2013) assume serially uncorrelated return pricing errors.¹⁵ In fact, consistent with the theoretical relationship between yield and return pricing errors, their empirical results reveal that the estimated yield pricing errors exhibit a significant degree of serial correlation whereas the corresponding return pricing errors show almost no serial correlation. ACM (2013) also employ a baseline specification that incorporates the first five PCs of zero-coupon yields as pricing factors. Although this five-factor model specification is selected on the basis of providing a better fit for the time series and cross section of excess bond returns, the associated model-implied yields also almost perfectly match the observed yields. Typically, the estimation of traditional ATSMs with more than three factors is computationally intensive under the likelihood-based estimation approaches as these models explicitly incorporate non-linear cross-equation restrictions imposed by the no-arbitrage assumption. As

¹⁴ In the ACM (2013) framework, the model is fitted to excess returns, rather than yields.

¹⁵ This assumption implies that the pricing factors are the single source of predictability in excess returns.

opposed to this, ACM (2013) do not enforce these cross-equation constraints (or bond pricing recursions) in their estimation procedure. This enables them to derive all model parameters by a simple regression-based approach in which there is no computational cost of including additional pricing factors.¹⁶ Due to its fast and straightforward implementation, the ACM method is a convenient choice for decomposing long-term bond yields, and therefore will be utilized later in this study.

The decomposition of bond yields permits the identification of key factors that influence the dynamics of their components. In this context, the determinants of term premium have received special attention in the literature, attributed to the crucial role of this component in explaining the movements in long-term bond yields. Although alternative ATSMs generally produce divergent quantitative estimates of the term premium due to differences in the number and type of state variables, as well as variations in parameter estimation techniques, the studies generally agree on its main determinants. Wright (2011) and Bauer, Rudebusch, and Wu (2014) document that the uncertainty about future inflation is an important driver of the term premium.¹⁷ As the inflation erodes the real value of nominal bond payments, investors require an additional compensation (so-called inflation risk premium) for the risk related to unanticipated changes in future inflation (Hördahl, 2008). Wright (2011) and Bauer et al. (2014) also find that uncertainty related to future growth, as proxied by dispersion of growth forecasts, has statistically and economically significant impact on term premium in most of their model specifications. Their results point out that higher macroeconomic uncertainty leads to an increase in the

¹⁶ ACM (2013) show that although the cross-equation restrictions are not imposed in their estimation procedure, the factor loadings estimated via linear regressions satisfy them to a high degree of precision.

¹⁷ Bauer et al. (2014) demonstrate that the key conclusion about the relationship between the inflation uncertainty and term premium in Wright (2011) remains intact even after correcting the small-sample bias problem in the estimated model parameters of his model.

term premium. In addition, the literature offers a wide range of evidence that term premium displays a pronounced countercyclical behavior. Consistent with the conventional wisdom that investors become more risk-averse and thus demand a higher risk premium for holding riskier assets during economic downturns, Bauer et al. (2014), Wu (2014), and Cohen, Hördahl, and Xia (2018) show that term premia tend to rise during recessions and fall during recoveries.¹⁸

Under traditional two-part decomposition of bond yields, the term premium component implicitly accounts for other risk premia in addition to inflation risk premium (Aguilar-Arguez, Diego-Fernández, Elizondo, & Roldán-Peña, 2022). This is particularly relevant for local currency government bonds issued by emerging market countries as investors charge additional risk premia for holding these riskier assets (Miyajima, Mohanty, & Yetman, 2014). One factor that contributes to the term premium is the credit risk premium. From a broader perspective, the credit risk encompasses the default risk, capital control risk, jurisdiction risk, currency convertibility risk and regulation risk (Du & Schreger, 2016).¹⁹ Therefore, the credit risk premium reflects the compensation for bearing these credit risks.²⁰ Another factor embedded in the term premium is the liquidity risk premium. Liquidity, a key

¹⁸ During financial crises, the “flight-to-quality” demand for government bonds could put downward pressure on term premium. However, Bauer et al. (2014) argue that if the financial crises coincide with economic recessions, it is highly unlikely that these safe-haven flows could completely offset countercyclical pattern of risk as the poor economic conditions during the recessions usually last longer than the financial market dislocations.

¹⁹ The currency convertibility risk is the risk in which a government imposes capital controls to prevent the repatriation of funds while avoiding a technical default (Du & Schreger, 2016). For the Euro area sovereign bonds, the conventional credit risk also involves redenomination risk. It is the risk that a sovereign could exit the euro and subsequently redenominate its debt into a new local currency at a depreciated exchange rate, causing losses for bondholders (Corradin, Grimm, & Schwaab, 2021; De Santis, 2019; Krishnamurthy, Nagel, & Vissing-Jorgensen, 2018).

²⁰ The credit spread (or credit risk compensation) reflects both the expected loss from credit events and corresponding risk premium. The latter relates to uncertainty that the realized loss could be larger than expected. While decomposing default-related credit spreads of 26 emerging market sovereigns, Remolona, Scatigna, and Wu (2007) show that the risk premia often make up a larger portion of the credit spreads than the expected losses. For simplicity, I do not separately identify these two components, and refer to their sum as the credit risk premium.

factor in asset pricing, can be defined as the ability to trade large quantities of an asset quickly and without adversely affecting its price (Chordia, Sarkar, & Subrahmanyam, 2004; Liu, 2006).²¹ Liquidity risk in the fixed income markets mainly arises from the potential challenge of unwinding a bond position prior to its maturity date. In this context, Amihud and Mendelson (2006) and Amihud, Mendelson, and Pedersen (2006) argue that investors are exposed to significant illiquidity costs when trading assets in the capital markets. In addition to direct trading costs (such as brokerage commissions, exchange fees or transaction taxes), the illiquidity costs include price-impact costs (i.e., price concessions made in return for immediate executions, or the cost of immediacy) and search and delay costs (i.e., opportunity cost incurred while searching for a counterparty in the market for a better price). Considering these illiquidity costs and uncertainty with regards to time-varying liquidity conditions, the risk-averse investors require a compensation for bearing the liquidity risk, leading to the liquidity risk premium.²² A vast literature documents the existence of the liquidity risk premium. It is even evident in nominal U.S. Treasuries (Fontaine & Garcia, 2012; Gürkaynak, Sack, & Wright, 2007), which are usually considered the most liquid securities in the global fixed-income markets.²³ Apart from these risk premia, the term premium derived from the two-part

²¹ Liquidity is a multidimensional concept. According to Sarr and Lybek (2002), it has five different dimensions, namely tightness (related to amount of transaction cost, such as bid-ask spreads), immediacy (associated with order execution speed), depth (linked to existence of wide range of orders above and below current trading price), breadth (connected to existence of both numerous and large orders in volume with minimal impact on price) and resiliency (linked to characteristic of a market to quickly trigger new order flows to correct order imbalances that drive prices away from what is justified by fundamentals). Liu (2006) summarizes the dimensions as trading cost, trading speed, trading quantity and price impact.

²² Similar to the compensation for credit risk (see footnote 20), the compensation for liquidity risk can be further decomposed into the expectation and risk-premium parts. Monfort and Renne (2013) demonstrate that the risk premium parts are substantially larger than the respective expectation parts when decomposing the liquidity-driven part of the spreads embedded in the Euro area sovereign bond yields. For simplicity, I do not separately identify these two components, and refer to their sum as the liquidity risk premium.

²³ The yield difference between on-the-run and off-the-run U.S. Treasury securities is the liquidity risk premium. Gürkaynak et al. (2007) measure it by comparing an on-the-run issue to a synthetic off-the-

yield decomposition also incorporates the currency risk premium. This premium represents the compensation for risks associated with unanticipated changes in the exchange rate (Domowitz, Glen, & Madhavan, 1998). Gadanecz, Miyajima, and Shu (2018) suggest that the increase in the uncertainty around the future path of exchange rate in emerging market countries raises local currency bond yields as global investors demand a larger yield compensation, specifically the currency risk premium, for holding these bonds.²⁴

Another frequently documented key factor that drives term premia is related to large-scale asset purchases by the central banks of developed economies, such as the Federal Reserve Bank (FED), the Bank of England (BOE), the European Central Bank (ECB), the Bank of Japan (BOJ), and some others. This unconventional monetary tool is commonly referred to as quantitative easing (QE). In the aftermath of the Global Financial Crisis (GFC) in 2008, once the conventional policy interest rates in major developed economies had reached to their effective lower bounds (i.e., the levels at which further rate cuts would have little or no effects), the central banks of these economies decided to implement further monetary easing by purchasing large quantities of government debt and other types of assets.²⁵ Such asset purchases

run security with the same coupon rate and maturity date. They show that the resulting liquidity premium measure for the 10-year Treasury note fluctuates around 10 bp in normal times and tends to move up during episodes of financial turmoil.

²⁴ The authors use the implied exchange rate volatility as a proxy for the exchange rate risk.

²⁵ The BOJ introduced its first QE in March 2001 and ended it in March 2006. The second episode of QE in Japan was initiated in October 2010. The FED, BOE and ECB announced their first QEs in December 2008, March 2009 and January 2015, respectively. Prior to QE, the ECB launched the “Covered Bond Purchase Programme” (CBPP) and the “Securities Markets Programme” (SMP). The CBPP, which was launched in two phases from July 2009 to June 2010 and November 2011 to October 2012, was designed primarily to ease financing conditions for credit institutions and enterprises, and thereby stimulate credit lending to the private sector. The SMP was adopted in May 2010 to ensure depth and liquidity in the dysfunctional market segments of the Euro area debt securities market and restore the monetary policy transmission mechanism. This purchase program was not in the form of QE since all related sovereign debt purchases were fully sterilized through the specific liquidity-absorbing operations. The SMP was replaced by the “Outright Monetary Transactions” (OMT) in September 2012. The OMT framework, which was stipulated to be unlimited bond buying program, included a strict conditionality that the ECB would only purchase sovereign bonds of a given country if it is covered under the “European Financial Stability Facility” (EFSF, a

are intended to lower the long-term bond yields, which in turn stimulate the economic activity. QE operates through a number of different potential channels. The first transmission channel is called the signaling channel (or expectations channel). It reflects that when a central bank announces a large-scale asset purchase program, the announcement provides a credible signal to investors that the central bank is concerned about current and future economic conditions (Bauer & Rudebusch, 2014) and therefore is willing to keep its policy rate at an exceptionally low level for an extended period of time. The signal is credible because if the central bank raises its short-term policy rate, it will be exposed to massive capital losses for the stock of purchased assets under the QE program (Krishnamurthy & Vissing-Jorgensen, 2011). In this context, the large-scale asset purchase announcement leads investors to revise down their expectations of the future short-term interest rates. This sequentially reduces the long-term bond yields through the expectations component (risk-neutral rate). The second transmission channel of QE is the portfolio balance channel, which acts on the term premium component of long-term bond yields. It is widely known as the key (or the most natural) transmission channel of QE and relies on the premise of reductions in the relative quantities of long-term government bonds held by private investors, who then try to rebalance their portfolios towards assets with similar characteristics (Joyce, Miles, Scott, & Vayanos, 2012). The portfolio balance channel mainly works through two subchannels. The first subchannel is the scarcity channel (or local supply channel), which is related to the local availability of fixed-income securities with a particular maturity. According to the preferred habitat theory, which provides the theoretical foundations of the local supply channel, there are preferred-

temporary bailout fund created in May 2010) or “European Stability Mechanism” (ESM, a permanent bailout fund adopted in March 2011 and officially set up in October 2012). Although the OMT scheme has never been used since its inception, it is widely credited for ending the sovereign debt crisis in Europe.

habitat investors with preferences for specific maturities, and the interest rate for a particular maturity is influenced by the demand of these investors and the supply of bonds with that maturity (Greenwood & Vayanos, 2010).²⁶ For instance, the long-term investors, such as pension funds and insurance companies, prefer long-term bonds to match their long-term liabilities, while short-term investors, such as money market mutual funds, prefer short-term government bonds and other cash-equivalent instruments to maintain a high level of liquidity in their portfolio (Li and Wei, 2012). Under the scarcity mechanism, when a central bank purchases a large quantity of long-term government bonds with a specific maturity, it creates a shortage of these bonds in the market. The sustained demand from preferred-habitat investors for this specific maturity then leads to an increase in the prices (or equivalently a reduction in the yields) of government bonds with this particular maturity or similar maturities (D’Amico, English, López-Salido, & Nelson, 2012).²⁷ Such a scarcity or local supply effect of QE is therefore concentrated on both the purchased government bonds and those at nearby maturities. It has no impact on the yields of government bonds that have maturities distinctly different from those of the purchased ones (Lemke & Werner, 2020; McLaren, Banerjee, & Latto, 2014). Under the maturity-specific bond scarcity channel of QE, the decline in the long-term government bond yields is attributed to the decline in the term premium component. The investors are willing to pay a convenience yield or scarcity premium for these long-term government bonds (Krishnamurthy & Vissing-Jorgensen, 2013). This translates into a lower term premium. The second subchannel through which the portfolio balance effect works is

²⁶ The preferred habitat theory was pioneered by the works of Culbertson (1957) and Modigliani and Sutch (1966).

²⁷ In this framework, as an additional requirement, the arbitrageurs are assumed to be risk-averse (e.g., capital constrained), which implies that they cannot fully exploit arbitrage opportunities across different maturities. Therefore, the demand from preferred-habitat investors will determine the bond yields (Vayanos & Vila, 2009).

the duration risk channel. It is associated with the removal of duration risk from the market.²⁸ The large-scale asset purchases by a central bank reduce a considerable amount of long-duration government bonds held by the private sector. Lower aggregate duration risk available to bond investors then increases their risk-bearing capacity, and thus they demand a lower risk premia to hold the remaining long-duration government bonds, implying a reduction in the term premia and hence the corresponding bond yields (Gagnon, Raskin, Remache, & Sack, 2010; Krishnamurthy & Vissing-Jorgensen, 2011, 2013).²⁹ Under the duration risk channel, when the central bank withdraws the duration from the market, the displaced investors with fewer assets restore their lost duration by purchasing other debt, particularly other long-duration government bonds. In this prospect, in contrast to scarcity channel, the duration risk channel operates in a way that the QE-related government bond purchases at a specific maturity range not only put downward pressure on the yields of the government bonds in that part of the yield curve, but also on the yields of government bonds located in different maturity segments (Lemke & Werner, 2020; McLaren, Banerjee, & Latto, 2014).³⁰ This channel could even affect the yields of other domestic fixed-income securities, such as corporate bonds, that are not directly targeted by QE operations (Gagnon et al., 2010; Joyce,

²⁸ The duration risk refers to exposure of long-term bonds to unexpected changes in the short-term policy interest rates. Long-maturity bonds are more exposed to the duration risk (or interest rate risk) than short-maturity bonds (Andrade, Breckenfelder, De Fiore, Karadi, & Tristani, 2016).

²⁹ The duration risk induces a duration risk premium, which is captured by the term premium component in the bond yields. According to the theoretical model of Vayanos and Vila (2009), the duration risk premium is a function of the duration and the market price of duration risk. The latter is also a function of the amount of duration risk held by the arbitrageurs and their risk aversion. Therefore, a central bank can reduce the aggregate duration risk in the fixed-income market by either increasing the amount of purchases for a given average maturity or by lengthening the average maturity without altering the total purchase amount. The duration risk channel emerges from the behavior of marginal bond market investors (arbitrageurs) and is not limited to models with preferred-habitat investors (Joyce & Tong, 2012).

³⁰ Compared to scarcity mechanism which operates through a narrow channel (i.e., effects limited to purchased bonds and those at nearby maturities), the duration risk channel has broad effects and reduces the term premium on all long-term government bonds (Krishnamurthy & Vissing-Jorgensen, 2013).

Lasaosa, Stevens, & Tong, 2011; Krishnamurthy & Vissing-Jorgensen, 2011).

Another primary channel through which the large-scale asset purchases decrease the government bond yields via the term premium component is the liquidity channel.

The presence of the central bank in the government bond market as a significant and committed buyer enhances the functioning of the market and the overall liquidity conditions, especially in stressed market conditions. The dealer and investor communities recognize that persistent central bank purchases make it less costly for them to liquidate their government bond holdings when needed, prompting them to take larger positions or increase their market-making activities in these securities (Gagnon et al., 2010). Therefore, by alleviating trading frictions in the market in a broad sense, the QE program depresses the liquidity risk premium embedded in the overall term premium component and thereby reduces government bond yields.

The purchases of government bonds of advanced economies by foreign official institutions, including foreign central banks and sovereign wealth funds, also lead to a decline in the term premium component of these bonds and their corresponding yields (Beltran, Kretchmer, Marquez, & Thomas, 2013). For instance, since the early 2000s, a handful of emerging market countries have invested a significant portion of their accumulated foreign exchange reserves in U.S. Treasury securities.³¹ These foreign official inflows resemble relative supply shocks that reduce the amount of bonds available to private investors (Sierra, 2014). From this

³¹ According to Csonto and Tovar (2017), foreign official purchases of U.S. Treasury securities between early 2000s and the end of 2014 occurred in two distinct phases. In the first phase (pre-GFC period), the purchases were primarily driven by emerging market countries' precautionary reserve accumulation to guard against future crises and investments by commodity exporters reaping significant benefits from the surge in the commodity prices. In the second phase (the post-GFC period), the huge capital inflows to emerging market countries during the QE programs led to a sharp rise in their foreign exchange reserves, part of which was then invested back into U.S. Treasury securities. Using the data from the Treasury International Capital system, Weiss (2022) reveals that foreign official holdings of U.S. Treasury securities almost levelled off, ranging from USD 4.11 to USD 4.57 trillion, from the end of 2014 to 2020.

perspective, when a foreign official agent purchases a long-term government bond, it should have a similar impact on the long-term yields as when a central bank purchases that bond under its QE program (Gagnon et al., 2010). Thus, akin to the portfolio rebalancing mechanism of QE, foreign official purchases of government bonds compress the term premium component through the portfolio balance effects and thereby lower the respective bond yields. Broadly, this effect can also arise from large-scale purchases by other significant market participants who are expected to hold bonds over time, often due to regulations.³²

In a highly integrated global financial system, besides the country-specific factors, various global factors such as global interest rates and global risk sentiment could also be important drivers of local currency government bond yields and their components. The *prima facie* evidence of existence of the global factors is that the interest rates in advanced and emerging market economies move closely together although these economies could be at different stages of the business cycle. This co-movement, which is especially more pronounced for long-term interest rates, reflects the impact of common global factors that drive international bond yields in the same direction (Hofmann & Takáts, 2015).³³ In this context, several empirical studies focusing on the interest rate linkages between the countries suggest that there are significant monetary spillover effects from the major advanced economies, particularly the United States, to other economies. For example, Hördahl, Sobrun, and Turner (2016) find that changes in the U.S. long-term bond yields have strong

³² Bernanke (2015) notes that regulatory changes can lower the U.S. term premiums by boosting demand for U.S. Treasury securities. For instance, new financial sector regulations require banks to hold more liquidity and securities dealers to post additional collateral for their derivatives transactions.

³³ The authors report that the average correlation of long-term government bond yields in emerging market economies and smaller advanced economies in their sample with those in the United States is 0.71 over the period 2000-2014. Similarly, Kolasa and Wesolowski (2020) show that the correlations between the first PC of 10-year term premia in small open economies and 10-year term premium in the United States are 0.73 and 0.94 in the pre-QE and post-QE periods, respectively.

spillover impacts on the domestic long-term bond yields of emerging and smaller advanced economies. By further decomposing the U.S. long-term interest rates, they also reveal that the U.S. term premium component is more influential than the expected future U.S. short-term rates in explaining variations in domestic long-term bond yields, especially in the emerging market economies. Albagli, Ceballos, Claro, and Romero (2019) contribute to the debate of interest rate spillovers by investigating the effect of U.S. monetary policy shocks on international bond yields and their components.³⁴ They show that the local long-term bond yields in advanced and emerging market economies react similarly and significantly to U.S. monetary policy shocks. However, the composition of corresponding yield changes differs between the two country groups. The action is concentrated on the risk-neutral component for advanced economies but predominantly on the term premium component for emerging market economies. In a separate study, Miyajima et al. (2014) assess and quantify the effects of international monetary spillovers on Asian bond markets. They report that the response of domestic long-term bond yields in Asia to a change in the U.S. long-term bond yield is substantial both before and after GFC, though the response is more persistent during the post-crisis period. More importantly, they demonstrate that the impact of the U.S. term premium on Asian domestic long-term bond yields is about twice as much as that of the U.S. long-term bond yield in both periods, indicating the key role of the U.S. term premium in influencing the local currency government bond markets in Asia. In a similar vein, Naudon and Yany (2016) point out that a shock to the U.S. term premium has a clear pass-through to domestic long-term bond yields and term premiums in emerging

³⁴ Albagli et al. (2019) identify the U.S. monetary policy shocks as changes in the U.S. 2-year Treasury yields observed within a narrow window centered around the Federal Open Market Committee meetings.

market economies, with a notably stronger and more persistent impact in Latin American countries. Another common factor that drives the global co-movement of interest rates is the shift in the global risk sentiment. The most widely used proxy for this factor among researchers is the VIX index. It measures the implied volatility of the S&P 500 index over the next 30 days and reflects both risk aversion and uncertainty in the financial markets (Bekaert, Hoerova, & Lo Duca, 2013). Obstfeld (2015) reports that changes in the VIX index have a material impact on the domestic long-term bond yields in both advanced and emerging market countries, especially for the latter group. A rise in the index raises the local long-term interest rates in these two country groups. Working on a different time period and country sample, Hofmann and Takáts (2015) also document that the VIX index affects both domestic short-term and long-term interest rates in advanced and emerging market economies, highlighting the role of global investor sentiment as a source of international spillovers. Additionally, Özbek and Talaslı (2020) show that the VIX index has significant positive effect on the term premium component of medium-term and long-term bond yields of emerging market countries in their study.

1.3 Construction of zero-coupon nominal yield curves

To decompose the Turkish government nominal bond yields into risk-neutral and term premium components using the estimation approach of ACM (2013), the first step involves the construction of the zero-coupon yield curves. For this purpose, I employ the Nelson-Siegel-Svensson (NSS) model, which was originally proposed by Nelson and Siegel (NS, 1987) and later extended by Svensson (1994). According to Gürkaynak et al. (2007), the NSS framework provides an appealing balance of flexibility to capture various yield curve shapes and parsimony to avoid overfitting. It

is widely used by central banks, such as the FED (Gürkaynak et al., 2007), the ECB (Coroneo, Nyholm, and Vidova-Koleva, 2008), and some others in the Euro area (Bank for International Settlements (BIS), 2005).

1.3.1 The NSS model

The NSS model assumes that the instantaneous forward rate n years ahead is characterized by the following functional form:³⁵

$$f_t(n; \beta) = \beta_0 + \beta_1 e^{\frac{-n}{\tau_1}} + \beta_2 \left(\frac{n}{\tau_1} \right) e^{\frac{-n}{\tau_1}} + \beta_3 \left(\frac{n}{\tau_2} \right) e^{\frac{-n}{\tau_2}}$$

The vector $\beta = (\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)$ is the model parameter set to be estimated at each time period t .³⁶ Under the NSS specification, the instantaneous forward rate begins at the level $\beta_0 + \beta_1$ (when $n = 0$) and eventually asymptotically converges to the constant level β_0 (when $n \rightarrow \infty$).³⁷ In between, forward rates can have two humps (or troughs).³⁸ The magnitude and direction of the first hump (or trough) is determined by the parameter β_2 while its location is specified by τ_1 . If β_2 is positive, a hump (inverted U-shape) occurs at τ_1 , whereas if it is negative, a trough (U-shape) occurs at τ_1 . In an analogous manner, β_3 determines the magnitude and direction of the second hump (if $\beta_3 > 0$) or trough (if $\beta_3 < 0$) at τ_2 . For nominal interest rates,

³⁵ If $f_t(n, m)$ is the interest rate of a forward contract that begins n years ahead and matures m years beyond the start date, then the instantaneous forward rate is $f_t(n, 0) = \lim_{m \rightarrow 0} f_t(n, m)$. Thus, for a given time period t , the instantaneous forward rate n years ahead can be interpreted as the marginal cost of borrowing for the infinitesimal time period after n (Bolder & Strélski, 1999).

³⁶ The NS model is a special case of the NSS model when $\beta_3 = 0$ or $\tau_1 = \tau_2$. Therefore, the NS model is nested within the NSS model.

³⁷ This implies that $-\beta_1$ can be regarded as the slope of the yield curve. If β_1 is positive, the curve has a negative slope, and vice versa.

³⁸ This is the key advantage of the NSS model compared to the NS model. Although the functional form of the NS model allows for only one hump (or trough), that of the NSS model allows for a second hump (or trough) on the forward rate curve. For complex yield curve shapes, this property makes the NSS model perform better than the NS model in terms of goodness of fit.

the parameters must also satisfy the conditions that $\beta_0 > 0$, $\beta_0 + \beta_1 > 0$, $\tau_1 > 0$, and $\tau_2 > 0$.³⁹

Since the NSS model is stated in terms of instantaneous forward rates, integrating the instantaneous forward rate function and dividing by n lead to an expression for continuously compounded zero-coupon yield over a maturity of n years at time t :⁴⁰

$$y_t(n; \beta) = \frac{1}{n} \int_0^n f_t(x; \beta) dx$$

$$y_t(n; \beta) = \beta_0 + \beta_1 \frac{1 - e^{-\frac{n}{\tau_1}}}{\frac{n}{\tau_1}} + \beta_2 \left(\frac{1 - e^{-\frac{n}{\tau_1}}}{\frac{n}{\tau_1}} - e^{-\frac{n}{\tau_1}} \right) + \beta_3 \left(\frac{1 - e^{-\frac{n}{\tau_2}}}{\frac{n}{\tau_2}} - e^{-\frac{n}{\tau_2}} \right)$$

Once the functional form of the zero-coupon rate is specified, it is straightforward to derive the discount function, denoted as $d_t(n; \beta)$, which is used to price any fixed income security with specific coupon and principal payments.

$$d_t(n; \beta) = e^{-y_t(n; \beta) \cdot n}$$

As the zero-coupon rate and discount factor functions are transformations of the initial instantaneous forward rate function, they are all closely related and hold the same relationship to the six model parameters (Bolder & Stréliski, 1999).

1.3.2 Formulation of the objective function

For each different parameter set $\beta = (\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)$, the NSS model yields a different instantaneous forward rate, zero-coupon rate, and discount factor functions,

³⁹ The first two conditions ensure that the two extremes (i.e., the starting value and the asymptote) of the instantaneous forward rate function are both positive. The last two conditions induce that the function has a finite limit for the very long end (Gimeno & Nave, 2006).

⁴⁰ The continuously compounded zero-coupon yield is the average of the instantaneous forward rates over the maturity n . Given that the instantaneous forward rate is interpreted as the marginal cost of borrowing (see footnote 35), the continuously compounded zero-coupon yield represents the average cost of borrowing over the investment period.

thereby producing a different model-implied yield or price for any outstanding fixed income security trading in a particular market. Thus, the yield curve estimation is essentially the process of finding the optimal parameter set β that is the most consistent with the actual bond yields or prices at a given time period. In this regard, the choice of an appropriate objective function for the optimization is highly critical in the yield curve fitting process. The literature offers two broad choices for the objective function. In the first choice, the optimal model parameters are estimated by minimizing the sum of squared deviations between the actual and model-implied yields to maturity. The corresponding objective function is as follows:

$$\min_{(\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)} \sum_{i=1}^{N_t} (y_t^{i,actual} - y_t^{i,model})^2$$

subject to: $\beta_0 > 0$, $\beta_0 + \beta_1 > 0$, $\tau_1 > 0$, and $\tau_2 > 0$,

where N_t is the number of bonds trading in period t .⁴¹ The optimization procedure for this objective function consists of two stages. The first stage involves finding the model-implied prices based on the NSS model parameter estimates. In the second stage, the model-implied yields to maturity are calculated from the corresponding model-implied prices and then compared with actual yields. Because the calculation of the model-implied yields to maturity from the model-implied prices requires an extra iterative process for all N_t bonds in each step of the optimization algorithm, the minimization of yield errors is computationally intensive and time-consuming (BIS, 2005; Wahlström, Paraschiv, & Schürle, 2022).

An alternative method of obtaining the model parameters could be to minimize the sum of squared deviations between the actual and model-implied prices. However, Svensson (1994) points out that this method could cause fairly

⁴¹ The actual and model-implied yields to maturity are both continuously compounded.

large yield errors for short-maturity bonds as their respective prices are very insensitive to yield changes. Since a change in the yield leads to a greater change in the price of long-maturity bonds than that of short-maturity bonds, the minimization of price errors results in the overfitting of the long-end of the yield curve at the expense of the short-end (BIS, 2005; Bolder & Stréliski, 1999;). One approach to overcome this problem is to use a weighting scheme for the price errors in the objective function. In this respect, the most common practice in the yield curve fitting literature is to weight each price error by the inverse of bond duration.⁴² Although this method gives more relevance to the price errors of short-maturity bonds, they tend to have higher weights in the optimization process (Çepni & Küçüksaraç, 2017; Gimeno & Nave, 2006).⁴³ Therefore, to ensure that long-maturity bonds also play an important role in the optimization process, Gimeno and Nave (2006) propose using the inverse of the square root of bond duration as the weighting scheme for each squared price error. Hence, the objective function takes the following form:

$$\min_{(\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)} \sum_{i=1}^{N_t} (P_t^{i,actual} - P_t^{i,model})^2 \frac{1}{\sqrt{D_t^i}}$$

subject to: $\beta_0 > 0$, $\beta_0 + \beta_1 > 0$, $\tau_1 > 0$, and $\tau_2 > 0$,

where N_t is the number of bonds trading in period t , $P_t^{i,actual}$ is actual price of bond i , $P_t^{i,model}$ is the model-implied price of bond i given the NSS model parameters, and D_t^i is the duration of bond i .

⁴² Weighting price errors by their inverse durations transforms them into yield fitting errors, to a first order approximation. Fitting the yield curve with inverse-duration-weighted price errors instead of yield errors is preferable since it drastically reduces the computation time required to find the optimum NSS model parameters (Gürkaynak, Sack, & Wright, 2010).

⁴³ For instance, based on this weighting scheme, the bonds at the very short end of the term structure have extremely high weights in the optimization process (Kanlı, Küçüksaraç, & Özel, 2013).

In penalizing the long-maturity bond prices in favor of the short-maturity bond prices, this objective function provides a more balanced approach in contrast to the one using inverse duration as a weight for each price error. Therefore, in this study, for the estimation of the zero-coupon nominal yield curves by the NSS model, I choose to minimize the sum of squared price errors weighted by the inverse of the square root of duration.

1.3.3 Optimization algorithms

The estimation of all six parameters of the NSS model is a non-linear optimization problem. This is due to the non-linearity imposed by the τ_1 and τ_2 parameters in the model's functional form. In addition to this, the optimization problem is non-convex and exhibits multiple local minima, which further complicates the parameter estimation process (Bolder & Stréliski, 1999; Manousopoulos & Michalopoulos, 2009; Wahlstrøm et al., 2022).

In this study, inspired by the study of Manousopoulos and Michalopoulos (2009), I use four alternative optimization algorithms from two different categories to address the yield curve estimation problem. However, my objective is not to compare the performance of these algorithms or determine the superior one across various dimensions. Instead, my focus is on identifying the optimal parameter set that minimizes the objective function under specific constraints for each month in the sample period. The optimization algorithms employed in this study are the Broyden-Fletcher-Goldfarb-Shanno (BFGS), limited memory BFGS with boundary constraints (L-BFGS-B; Zhu, Byrd, Lu, & Nocedal, 1997), Generalized Reduced Gradient (GRG; Lasdon, Waren, Jain, & Ratner, 1978) and Differential Evolution

(DE; Storn & Price, 1997).⁴⁴ The first three algorithms belong to the gradient-based class, which requires the numerical computation of the gradient of the objective function (Manousopoulos & Michalopoulos, 2009). They are widely used for solving a variety of non-linear optimization problems in different disciplines. As for the DE algorithm, it is classified as a population-based evolutionary algorithm and does not require any gradient information. It optimizes a problem by maintaining a population of candidate solutions and generating new candidate solutions through mutation and crossover operations, and then keeping the candidate solution with the best score or fitness (Storn & Price, 1997). The DE algorithm has gained popularity among scholars and practitioners in recent years due to its ease of implementation for addressing complex optimization issues.

1.3.4 Bounds and initial values for model parameters

In addition to the already introduced strictly positive lower bounds on β_0 , $\beta_0 + \beta_1$, τ_1 , and τ_2 , I impose further restrictions on the parameter set to increase the convergence speed and avoid extreme values that do not have a clear financial interpretation. These restrictions can also reduce potential parameter instability (Bekker & Bouwman, 2009). In fact, when the parameters are not constrained, they occasionally jump discretely from one time period to the next, often with a minor change in the underlying bond prices (Gürkaynak et al., 2007).⁴⁵

When applying these restrictions on the parameters, it is critical to ensure that the NSS model retains its flexibility to fit complex yield curve shapes. Therefore, to

⁴⁴ The BFGS algorithm is simultaneously outlined by Broyden (1970), Fletcher (1970), Goldfarb (1970), and Shanno (1970). L-BFGS is the limited memory version of the BFGS algorithm. It requires less memory than BFGS by discarding old gradients (Liu & Nocedal, 1989). L-BFGS-B is the extension of L-BFGS that incorporates bounds on the input variables.

⁴⁵ Gürkaynak et al. (2007) highlight that while these jumps in parameters can be significant, the estimation produces similar yield curve shapes through different combinations of parameters.

balance these considerations, I use the following lower and upper bounds in the estimation procedure.⁴⁶

$$0 < \beta_0 \leq 0.5, -0.5 < \beta_1 \leq 0.5, -1 \leq \beta_2 \leq 1, -1 \leq \beta_3 \leq 1, 0 < \tau_1 \leq 10, 0 < \tau_2 \leq 10$$

Similar to other applications that employ further parameter restrictions in the NSS framework, these bounds above are derived from a trial and error procedure and hence are not based on any economic theory.⁴⁷

In solving this non-convex optimization problem, as previously explained, the presence of multiple local minima poses significant challenges. Any attempt to locate the global minimum with certainty requires the test of all parameter combinations, which is practically impossible. For instance, with only four different initial values for each of the six NSS model parameters, $4^6 = 4,096$ different combinations need to be examined. Given this heavy computational burden, to avoid convergence to spurious local minima, it is necessary to have good initial parameter estimates before initializing the optimization algorithms.⁴⁸

In this regard, drawing inspiration from the study of Wahlström et al. (2022), I consider three different approaches to determine the initial parameter values. Since the DE algorithm generates the initial population randomly within the lower and upper bounds (see footnote 48), such parameter initialization approaches are only applicable to the BFGS, L-BFGS-B and GRG algorithms. In the first approach, I

⁴⁶ Since the BFGS algorithm does not handle bounds or constraints, I incorporate them into the objective function via a penalty function. The L-BFGS-B algorithm supports bounds but not constraints. To deal with the constraint condition on $\beta_0 + \beta_1$, I again adopt a penalty function approach.

⁴⁷ For instance, see Bolder and Strélski (1999) and De Lange, Ristad, and Westgaard (2022).

⁴⁸ In contrast to gradient-based optimization algorithms, the population-based DE algorithm does not require any specific initial values. It randomly generates the initial population of candidate solutions (the individuals) considering the lower and upper bounds of the model parameters. Each individual is a vector of six scalar parameters $\beta = (\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)$. To facilitate convergence to the global optima, the population size should be at least 10 times the number of parameters of the objective function (Price, Storn, & Lampinen, 2005). However, a large population size also slows down the convergence speed. Hence, to balance the convergence accuracy and speed of the DE algorithm, I use a population size of 90 individuals in this study.

utilize the parameter solution of the previous month. This is based on the assumption that the shape of the yield curve does not change significantly from the previous month.⁴⁹ In the second approach, I follow Manousopoulos and Michalopoulos (2009) methodology, which aligns with financial interpretation of the model parameters in the respect that β_0 and $\beta_0 + \beta_1$ correspond to very long-term and very short-term spot interest rates, respectively. The initial parameters are specified as follows:

$$\beta_0 = \frac{1}{m} \sum_{i=1}^m y_i, \quad \beta_1 = y_s - \beta_0, \quad \beta_2 = \beta_3 = 0, \quad \tau_1 = \tau_2 = 1$$

where m corresponds to the bonds with the longest maturities, either 2 or 3, and s is the bond with the shortest maturity.⁵⁰ Considering the limited number of long-term bonds traded in the Turkish secondary bond market, I choose m to be equal to 2. This means that β_0 is initially set to the average yield to maturity of the two longest maturity bonds available in the sample for a given estimation day. As for β_1 , its initial value is computed as the difference between the yield to maturity of the shortest maturity bond available in the sample and β_0 . In the third approach, I choose random initial values drawn from a uniform distribution within the pre-specified parameter bounds.⁵¹ This approach is similar to setting up the initial population of DE algorithm. By employing this method, I randomly generate 50 different parameter sets for the BFGS, L-BFGS-B and GRG algorithms.

Overall, with these proposed parameter initialization methods, each gradient-based optimization algorithm is run with 52 different sets of initial values. These include one set from the first approach, another set from the second approach, and 50

⁴⁹ This method is used by Gimeno and Nave (2006) and Wahlström et al. (2022) as one of the alternative methods for determining initial parameter values for daily yield curve estimation.

⁵⁰ The initial parameter values used by Manousopoulos and Michalopoulos (2009) and the Bundesbank are similar, except β_2 and β_3 . The Bundesbank set them equal to -1 (Nymand-Andersen, 2018).

⁵¹ This technique is also utilized by Gilli, Große, and Schumann (2010), Gilli and Schumann (2010), and Wahlström et al. (2022) while calibrating the NSS model.

random sets from the third approach. I also run DE algorithm five times, each starting with a different initial population of 90 randomly generated individuals.⁵² The best solution of each algorithm is then compared to determine the optimal NSS parameter set for a given month in the sample.

1.3.5 Data

To estimate the zero-coupon yield curve, I collect the secondary bond market data from the Borsa Istanbul (BIST). The curve fitting process is performed on the last full business day of each month from February 2005 to June 2022. I choose February 2005 as the starting period because Turkey launched its first 5-year fixed-coupon bond in local currency on February 15, 2005. From this point until the end of 2009, the upper bound of the bond maturity spectrum in the dataset was 5 years. It has been extended to 10 years with the debut auction of 10-year fixed-coupon bond on January 26, 2010.

After gathering the data, I determine the set of bonds to be used in the estimation. For this purpose, I apply a filtering process to exclude all floating-rate notes, inflation-indexed bonds, and corporate debt securities from the yield curve estimation, leaving only zero-coupon and fixed-coupon government bonds in the sample for each month. Moreover, if the trade and settlement dates of any bond transaction are different, that bond is also excluded. Finally, certain fixed-coupon bonds that remain in the sample might also be removed on an ad hoc basis. For instance, the primary dealer banks in Turkey are contractually obliged to enhance

⁵² I limit the number of runs to five for two reasons. First, the DE algorithm delivers similar results across repeated runs. Second, it requires significantly more CPU time than the gradient-based algorithms to solve this non-linear optimization problem.

liquidity in the secondary bond market.⁵³ Under this framework, on every trading day, they continuously quote bid and offer prices for the selected benchmark securities traded at the BIST Debt Securities Market.⁵⁴ Because of the greater liquidity provided by the primary dealers, these benchmark issues may trade at a premium relative to other securities in the secondary market. If any of these bonds is clearly disconnected from the rest of the curve, I also exclude it from the estimation.

1.3.6 Estimation results

Employing the above methodology, I estimate the end-of-month zero-coupon yield curves for Turkish government bonds over the period from February 2005 through June 2022. Table 1 presents the numerical results of the curve fitting procedure for selected dates using four different optimization algorithms and three different parameter initialization approaches. The results show that initiating any specific gradient-based optimization algorithm with different initial parameter sets may result in convergence to different local minima. For instance, setting initial parameter values for the GRG algorithm according to the first, second, and third approaches for January 31, 2012, yields the optimal objective function values of 0.1810, 0.1693, and 0.1203, respectively.⁵⁵ The results in Table 1 also suggest that different gradient-based algorithms with the same starting parameter values might generate very

⁵³ The primary dealership system in Turkey was first initiated in May 2000. Nevertheless, the program was terminated de facto with the emergence of the financial crisis in November 2000. It was reintroduced on September 2, 2002.

⁵⁴ Based on the latest updated primary dealership contract in 2021, each primary dealer must select seven benchmark securities from a list of 14 benchmark securities specified by the Ministry of Treasury and Finance to provide two-way quotes. These seven securities consist of four fixed-coupon bonds, one floating-rate note, one inflation-linked bond, and one TLREF-indexed bond. At least one of the four chosen fixed-coupon bonds must have a maturity of 5 years or more. Any bond with less than 182 days to maturity loses its benchmark status.

⁵⁵ As previously described, the third parameter initialization approach involves randomly selecting 50 initial values from the pre-specified parameter ranges. Each gradient-based algorithm undergoes 50 independent runs with these parameters, and I report the best solution for each algorithm in Table 1.

Table 1. Estimation of the Optimal Parameter Set for Selected Dates

Date	Initial Parameter Approach	Opt. Algorithm	Obj. Funct. Value	β_0	β_1	β_2	β_3	τ_1	τ_2
Jan 31, 2012	First Approach	BFGS	0.1691	0.0920	0.0086	0.9277	-0.9463	0.8674	0.8525
		L-BFGS-B	0.1691	0.0920	0.0086	0.5778	-0.5964	0.8715	0.8478
		GRG	0.1810	0.0919	-0.0202	0.0189	0.0753	1.4076	0.1170
	Second Approach	BFGS	0.1691	0.0920	0.0086	0.8797	-0.8983	0.8679	0.8522
		L-BFGS-B	0.1694	0.0920	0.0086	-0.1047	0.0868	0.7891	0.9348
		GRG	0.1693	0.0920	0.0086	-0.1310	0.1130	0.8060	0.9210
	Third Approach	BFGS	0.1159	0.2256	-0.1260	-0.1114	-0.3629	1.0897	6.8468
		L-BFGS-B	0.1159	0.2219	-0.1224	-0.1095	-0.3532	1.0791	6.8204
		GRG	0.1203	0.3856	-0.2873	-0.1877	-0.7837	1.4426	7.8095
	Initial Population	DE	0.3397	0.1623	-0.0725	-0.9860	0.8749	9.4927	8.8945
Jan 30, 2015	First Approach	BFGS	0.4598	0.0001	0.0756	-0.8788	0.9996	3.5382	3.9917
		L-BFGS-B	0.4775	0.0424	0.0343	-0.5820	0.6182	2.7289	3.0505
		GRG	0.4632	0.0129	0.0630	-0.4947	0.5926	3.2269	3.8807
	Second Approach	BFGS	0.5624	0.0733	0.0062	-0.0213	-0.0114	1.0025	1.0024
		L-BFGS-B	0.5624	0.0733	0.0063	-0.0164	-0.0164	0.9859	1.0001
		GRG	0.5624	0.0733	0.0062	-0.0161	-0.0166	1.0039	0.9992
	Third Approach	BFGS	0.4598	0.0001	0.0756	1.0000	-0.8884	3.9907	3.5396
		L-BFGS-B	0.4598	0.0001	0.0756	-0.8793	1.0000	3.5382	3.9915
		GRG	0.4600	0.0010	0.0747	1.0000	-0.8900	3.9736	3.5282
	Initial Population	DE	0.4600	0.0001	0.0756	-0.3276	0.4555	3.2372	4.3344
Jan 31, 2018	First Approach	BFGS	0.4683	0.1258	-0.0073	0.9851	-1.0000	1.7480	1.8610
		L-BFGS-B	0.4683	0.1258	-0.0073	0.9590	-0.9739	1.7466	1.8628
		GRG	0.4836	0.1203	-0.0016	0.9983	-1.0000	1.6007	1.6873
	Second Approach	BFGS	1.0293	0.1065	0.0242	0.9992	-0.9774	0.4108	0.3933
		L-BFGS-B	1.0316	0.1065	0.0238	-0.2282	0.2533	0.3672	0.4381
		GRG	1.0298	0.1065	0.0243	-0.4807	0.5039	0.3834	0.4185
	Third Approach	BFGS	0.4683	0.1258	-0.0073	-1.0000	0.9856	1.8610	1.7480
		L-BFGS-B	0.4683	0.1258	-0.0073	-1.0000	0.9855	1.8616	1.7485
		GRG	0.4684	0.1257	-0.0072	-0.8763	0.8620	1.8669	1.7381
	Initial Population	DE	0.4991	0.1661	-0.0394	-0.0419	-0.1720	0.3606	4.4901
Sep 30, 2021	First Approach	BFGS	0.5735	0.1754	-0.0282	1.0000	-0.9812	0.3516	0.3667
		L-BFGS-B	0.6450	0.1754	-0.0008	-0.0354	-0.0284	0.0162	0.6411
		GRG	0.2204	0.1152	0.0402	0.0710	0.1736	0.3632	3.9236
	Second Approach	BFGS	0.2696	0.1421	0.0308	0.7588	-0.7120	2.1348	1.9389
		L-BFGS-B	0.2696	0.1420	0.0309	1.0000	-0.9525	2.1139	1.9658
		GRG	0.6508	0.1753	-0.0010	-0.0089	-0.0197	0.6166	0.6167
	Third Approach	BFGS	0.2204	0.1158	0.0395	0.0710	0.1720	0.3599	3.9181
		L-BFGS-B	0.2204	0.1158	0.0395	0.0710	0.1720	0.3599	3.9183
		GRG	0.2204	0.1158	0.0394	0.0710	0.1720	0.3597	3.9177
	Initial Population	DE	0.2204	0.1154	0.0402	0.0707	0.1732	0.3634	3.9259

This table displays the numerical results of the curve fitting procedure for the selected dates using four different optimization algorithms and three different parameter initialization approaches. The algorithms and parameter initialization approaches are described in Section 1.3.3 and 1.3.4, respectively. Note that in the third approach, each gradient-based optimization algorithm is run 50 times, and their best solution is reported. As for the DE algorithm, it is run five times to identify the corresponding optimal NSS parameter set.

different solutions. As an example, running BFGS, L-BFGS-B, and GRG algorithms under the first parameter initialization approach for September 30, 2021, leads to optimal objective function values of 0.5735, 0.6450, and 0.2204, respectively. Finally, in line with literature, I observe that the DE algorithm which starts the optimization process with a randomly generated initial population (see footnote 48) does not guarantee a global optimal solution and can easily become trapped in local minima.

These evidences indicate that finding the best possible solution for this non-convex curve fitting problem necessitates the implementation of various optimization algorithms and parameter initialization techniques. The diversity of algorithms and the choice of initial parameter values in the optimization procedure both have a significant impact on the quality of the fit achieved in the estimation of spot yield curve.

Table 2 gives the summary statistics for the end-of-month zero-coupon yields at various fixed maturities. Since the maturity spectrum of the bonds has been extended from 5 to 10 years in January 2010, I split the sample period into two separate periods. Table 2a presents the descriptive statistics for the period between February 2005 and December 2009. For this period, the average yield curve exhibits a hump-shaped pattern with a peak around 2 years. Furthermore, the volatility of yields decreases almost monotonically with maturity. Over this time interval, the yields reached their maximum at the height of the GFC in October 2008, while their minimums were observed in late 2009, which coincided with the post-crisis period of easy monetary conditions around the world.⁵⁶ Table 2b provides the same descriptive statistics for the period from January 2010 to June 2022. In this period, the average

⁵⁶ The CBT cut the lower bound of the interest rate corridor, which served as the policy rate prior to May 2010, by 1025 bp from 16.75% in October 2008 to 6.5% in November 2009.

yield curve is again hump-shaped with a peak around 2-2.5 years. The yield volatility rises for maturities up to 1.5 years and then falls as maturity increases. The interest rates reached their lowest levels in April 2013, shortly before the FED announcement in May 2013 of its intention to taper QE purchases. As for the maximums, yields for maturities between 1 and 5 years hit their highest levels during the local currency crisis in August 2018, while yields for maturities less than 1 year hit their record highs after the CBT raised its policy rate by 625 bp from 17.75% to 24% in September 2018. The peak for longer maturities from 5 to 10 years took place in March 2022 in the wake of the conflict between Russia and Ukraine.

Table 2. Summary Statistics for Zero-Coupon Nominal Yields

(a) Between February 2005 and December 2009								
Maturity (months)	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
6	0.1475	0.0334	0.0722	0.1291	0.1538	0.1747	0.2078	59
9	0.1506	0.0341	0.0739	0.1302	0.1551	0.1771	0.2151	59
12	0.1526	0.0340	0.0764	0.1323	0.1560	0.1782	0.2199	59
18	0.1547	0.0329	0.0825	0.1306	0.1579	0.1793	0.2268	59
24	0.1554	0.0315	0.0882	0.1280	0.1573	0.1779	0.2319	59
30	0.1551	0.0304	0.0924	0.1288	0.1590	0.1756	0.2350	59
36	0.1544	0.0295	0.0947	0.1311	0.1593	0.1751	0.2359	59
48	0.1524	0.0283	0.0962	0.1325	0.1564	0.1726	0.2318	59
60	0.1506	0.0279	0.0959	0.1317	0.1547	0.1692	0.2219	59
(b) Between January 2010 and June 2022								
Maturity (months)	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
6	0.1114	0.0437	0.0500	0.0797	0.0980	0.1319	0.2340	150
9	0.1128	0.0441	0.0502	0.0804	0.0972	0.1332	0.2330	150
12	0.1137	0.0446	0.0504	0.0816	0.0979	0.1339	0.2329	150
18	0.1146	0.0451	0.0511	0.0838	0.0980	0.1326	0.2460	150
24	0.1149	0.0449	0.0520	0.0862	0.0982	0.1324	0.2554	150
30	0.1149	0.0442	0.0529	0.0864	0.0997	0.1310	0.2588	150
36	0.1147	0.0433	0.0540	0.0871	0.0993	0.1291	0.2580	150
48	0.1142	0.0412	0.0561	0.0878	0.0999	0.1266	0.2486	150
60	0.1136	0.0393	0.0581	0.0882	0.1005	0.1263	0.2348	150
72	0.1131	0.0377	0.0597	0.0884	0.0995	0.1264	0.2317	150
84	0.1126	0.0363	0.0610	0.0890	0.0990	0.1254	0.2304	150
96	0.1124	0.0352	0.0620	0.0892	0.0990	0.1258	0.2278	150
108	0.1122	0.0344	0.0626	0.0891	0.0991	0.1271	0.2241	150
120	0.1122	0.0338	0.0629	0.0896	0.0992	0.1275	0.2194	150

Figure 1 displays time-series plots of end-of-month zero-coupon yields for a set of fixed maturities. The key trends, minimums and maximums are clearly visible in these plots.

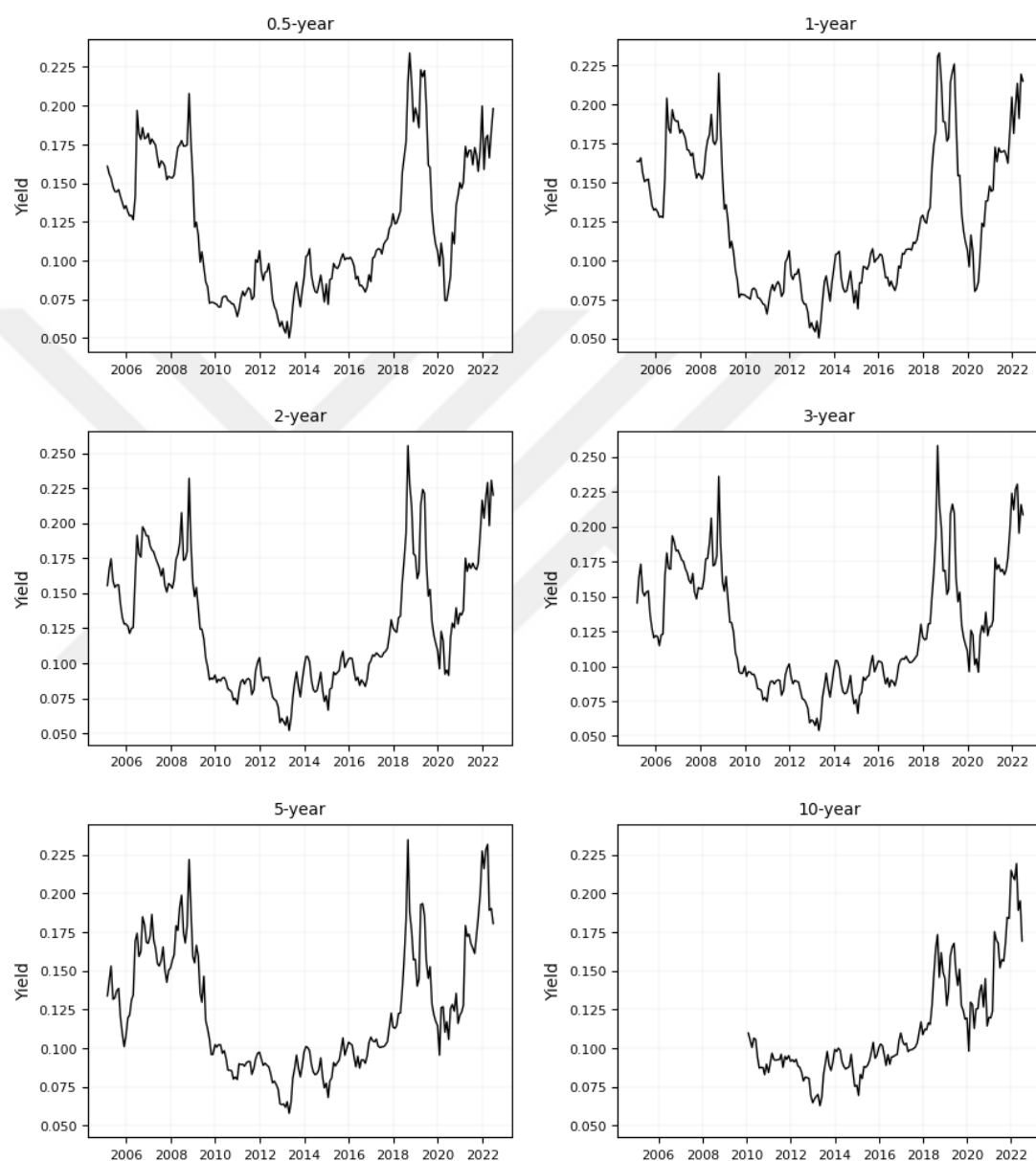


Figure 1. Zero-coupon nominal yields

Figure 2 illustrates a three-dimensional plot of the zero-coupon yield curve data. It reveals not only the cross-sectional variation of the yields across different

maturities but also the dynamic evolution of the yield curve throughout the sample period.

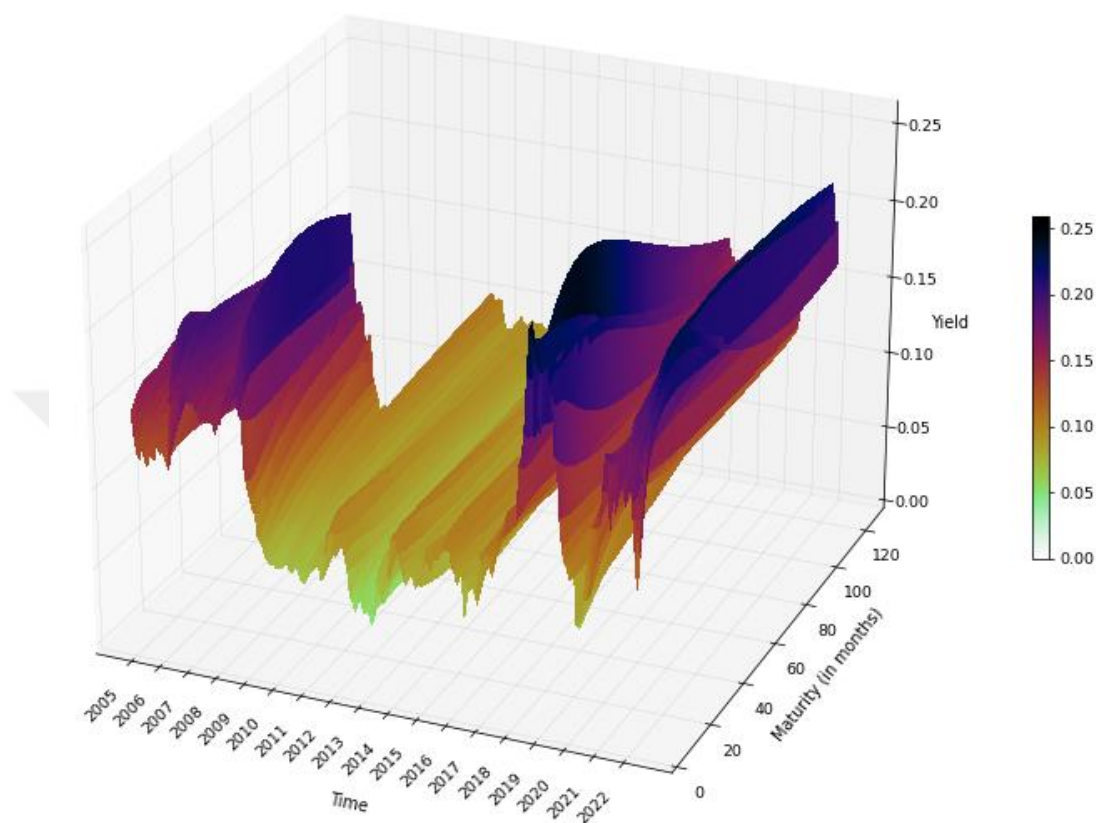


Figure 2. Zero-coupon nominal yield surface

Figure 3 demonstrates the plot the estimated zero-coupon yield curves along with the continuously compounded actual and fitted (model-implied) yields to maturity for the specific dates considered in Table 1.⁵⁷ It is apparent that the estimation procedure is able to capture various yield curve shapes effectively. More importantly, the overall fit is visually quite good.

⁵⁷ As also stated by Svensson (1994) and Söderlind and Svensson (1997), although continuously compounded fitted yields to maturity for zero-coupon bonds are directly on the zero-coupon yield curve, this is typically not the case for coupon bonds because yields to maturity on them are generally different from those on zero-coupon bonds of the same maturity.

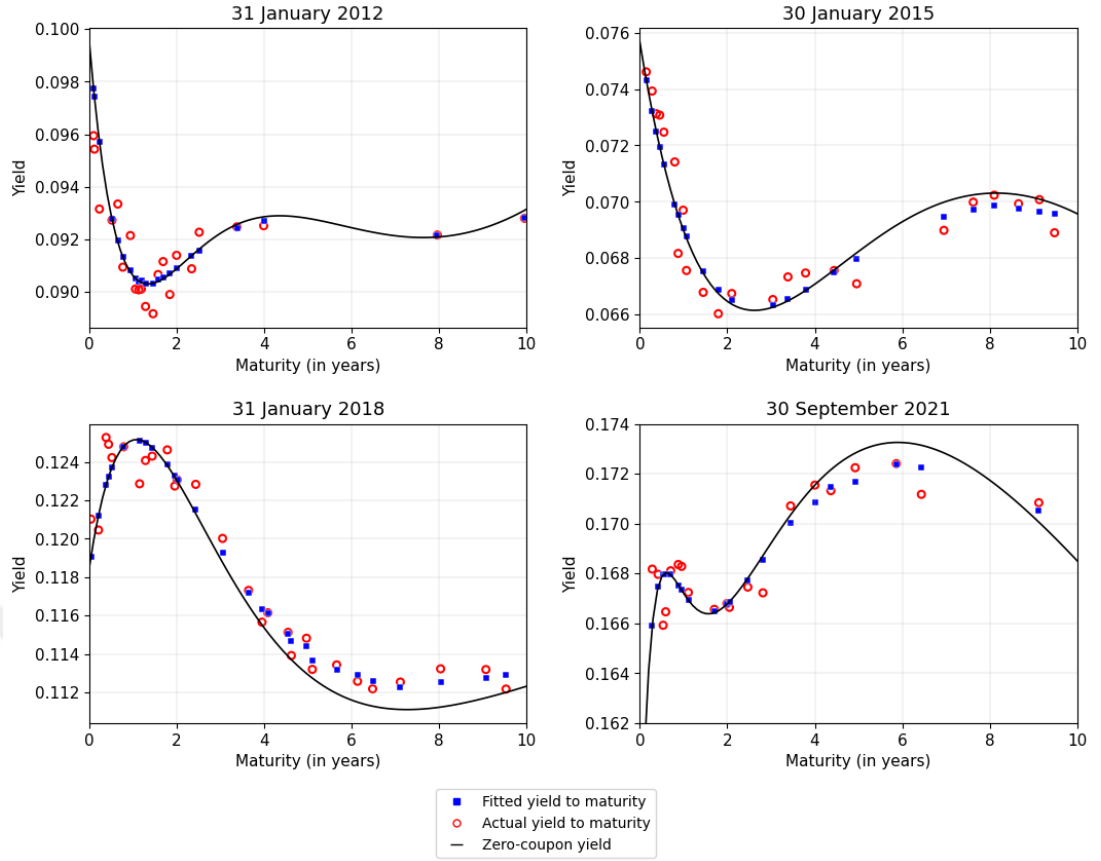


Figure 3. Yield curve estimation for selected dates

To assess in-sample fit over these dates, I report the RMSE and Mean Absolute Error (MAE) measures for both price and yield to maturity errors in Table 3.⁵⁸ The results show that the RMSE and MAE values for the price errors range from 0.078 to 0.201 and 0.061 to 0.152, respectively, while those for yield to maturity errors span from 7.7 bp to 10 bp and 6.7 bp to 7.3 bp.⁵⁹ These low error measures

⁵⁸ The error is defined as the difference between the actual and fitted value. The RMSE and MAE measures for the price error are calculated as follows:

$$RMSE = \sqrt{\frac{1}{N_t} \sum_{i=1}^{N_t} (P_t^{i,actual} - P_t^{i,model})^2}$$
 and $MAE = \frac{1}{N_t} \sum_{i=1}^{N_t} |P_t^{i,actual} - P_t^{i,model}|$, where N_t is the number of bonds trading in period t , $P_t^{i,actual}$ is actual price of bond i , $P_t^{i,model}$ is the fitted (model-implied) price of bond i . The RMSE and MAE measures for the yield to maturity error are calculated by replacing price values with continuously compounded yield to maturity values in the respective formulas.

⁵⁹ The RMSE values are generally higher than the corresponding MAE values because the RMSE incorporates the squared errors, which penalize larger errors more severely than smaller ones, whereas the MAE treats all errors equally.

provide the empirical evidence of the satisfactory performance of the estimation process and also verify the visually good quality of the curve fitting observed for each specified day in Figure 3.

Table 3. RMSE and MAE Measures for Price and Yield to Maturity Errors

Date	Price Errors		Yield to Maturity Errors	
	RMSE	MAE	RMSE (bp)	MAE (bp)
31 January 2012	0.078	0.061	10.0	7.3
30 January 2015	0.201	0.152	7.7	6.7
31 January 2018	0.173	0.135	9.6	7.2
30 September 2021	0.134	0.095	9.4	7.2

The RMSE and MAE values for price errors are expressed in TRY for bonds with a face value of 100 TRY while those for yield to maturity errors are expressed in bp.

Table 4 presents an overview of in-sample RMSE statistics for price errors in different maturity buckets and for all maturities.⁶⁰ Similar to the approach taken in Table 2, to account for the extension of the maturity spectrum to 10 years in January 2010, the sample period in Table 4 is divided into two separate periods. According to Table 4a, the estimation process produces fairly low pricing errors across all maturity segments for the period between February 2005 and December 2009. Notably, the mean RMSE value for the maturity segment between 2 and 5 years is lower than that for the maturity segment between 1 and 2 years, suggesting greater accuracy of the curve fitting process for the former segment. Table 4b demonstrates the same statistics for the period from January 2010 to June 2022. Similar to the first time interval, the price errors for all maturity buckets in this period imply reasonably accurate fit. The mean RMSE values in the second time period rise with increasing maturity, indicating a slightly less precise fit for longer-dated bonds compared to

⁶⁰ Calculating the RMSE for the yield to maturity errors is computationally intensive due to the extra iterative process required to derive the model-implied yields to maturity from the model-implied prices. Hence, for the entire sample period, I only report RMSE statistics for the price errors.

shorter-dated bonds. Overall, the findings suggest that the in-sample accuracy of the yield curve fit varies over time and across different maturity buckets.

Table 4. RMSE Statistics for the Price Errors in Different Maturity Buckets

(a) Between February 2005 and December 2009							
Maturity Bucket	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.
Btw. 0 and 1 year	0.072	0.042	0.019	0.046	0.063	0.090	0.227
Btw. 1 and 2 years	0.163	0.088	0.037	0.116	0.146	0.181	0.487
Btw. 2 and 5 years	0.124	0.093	0.005	0.047	0.095	0.186	0.375
All	0.112	0.046	0.033	0.084	0.097	0.131	0.279

(b) Between January 2010 and June 2022							
Maturity Bucket	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.
Btw. 0 and 1 year	0.084	0.067	0.010	0.041	0.066	0.097	0.417
Btw. 1 and 2 years	0.142	0.096	0.001	0.073	0.126	0.175	0.518
Btw. 2 and 5 years	0.221	0.146	0.025	0.129	0.187	0.255	1.030
Btw. 5 and 10 years	0.221	0.197	0.001	0.091	0.182	0.290	1.332
All	0.187	0.117	0.042	0.112	0.155	0.235	0.764

Figure 4 depicts the temporal evolution of the RMSE for the price errors across the specified maturity buckets and for all maturities. It is observed that there is a positive correlation between the occurrence of occasional spikes in RMSE values and periods of financial crises, indicating heightened level of pricing inaccuracies during such times. According to Hu et al. (2013), the scarcity of arbitrage capital during market crises compels market participants to reduce or completely abandon their relative value trades, which results in yields to fluctuate more freely in their own habitats and ultimately leads to substantial price deviations from the fundamental values. In this context, the misalignments in the Turkish local yield curve, reflected by pronounced deviations above the RMSE means, have been observed during several episodes within my sample period. These include the emerging markets sell-off in May and June 2006, the GFC in October 2008, the taper tantrum in May 2013, the local currency crisis in August 2018, the liquidity squeeze

in the offshore swap market for Turkish lira prior to local elections in March 2019, and the renewed currency crisis in December 2021.⁶¹

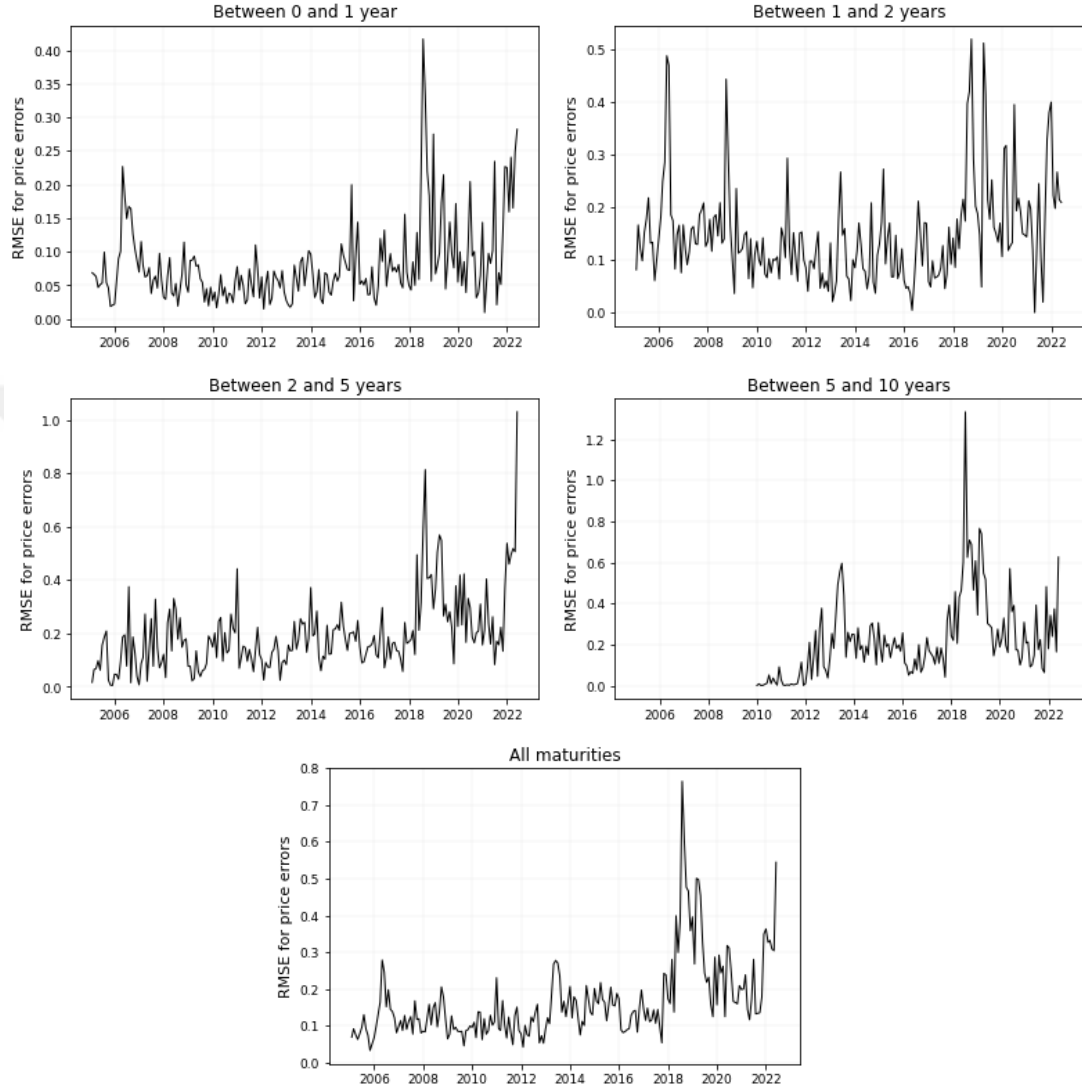


Figure 4. Temporal evolution of the RMSE for the price errors across different maturity buckets and for all maturities

Finally, to evaluate the in-sample performance of my estimation methodology, I compare my RMSE statistics with those reported by Çepni and Küçüksaraç (2017), who also employ the NSS model and the same secondary market

⁶¹ The RMSE for 2-5 year maturity bucket in June 2022 surged due to the CBT's regulatory changes.

data from the BIST to estimate the zero-coupon yield curve for Turkish government local currency bonds. Given that their study examines the period between January 2011 and May 2016, I recalculate my RMSE statistics for the same time period in order to facilitate the comparison. Table 5 presents these recalculated statistics. Although a direct comparison is still not possible because of the difference in data frequencies, I obtain very similar RMSE mean and standard deviation values for the first two maturity buckets.⁶² This similarity, particularly in maturity bucket between 0 and 1 year, is noteworthy given that, while the Çepni and Küçüksaraç (2017) assigns more weight to the short-maturity bonds by weighting every price error by the inverse duration in their objective function, I use the inverse of the square root of bond duration as the weighting scheme for each squared price error to ensure a balance between the short-maturity and long-maturity bonds in the optimization process (see section 1.3.2). Although I give less weight to short maturity bonds in the objective function, the observed similarity of in-sample performance for the short maturity buckets is potentially explained by the fact that I employ different optimization algorithms and parameter initialization approaches in the whole estimation framework. For the maturity buckets between 2 and 5 years and between 5 and 10 years, the RMSE values (see footnote 62 and Table 5) indicate a superior in-sample performance of my overall curve fitting methodology. The substantial improvement in the goodness of fit for these maturity buckets is attributable to both my weighting scheme, which gives less relevance to short-term bonds, and the

⁶² The study by Çepni and Küçüksaraç (2017) is based on the daily data. They report the RMSE statistics of the “no filter” scenario in Table 2 of their study. The means for maturity buckets between 0 and 1 year, between 1 and 2 years, between 2 and 5 years, and between 5 and 10 years are 0.056, 0.114, 0.211, and 0.334, respectively, with corresponding standard deviations of 0.036, 0.068, 0.133, and 0.330.

utilization of various optimization algorithms and parameter initialization techniques.⁶³

Table 5. RMSE Statistics for In-Sample Performance Comparison

Between January 2011 and May 2016							
Maturity Bucket	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.
Btw. 0 and 1 year	0.060	0.032	0.015	0.036	0.057	0.074	0.200
Btw. 1 and 2 years	0.104	0.060	0.005	0.061	0.093	0.143	0.293
Btw. 2 and 5 years	0.160	0.077	0.025	0.106	0.141	0.205	0.442
Btw. 5 and 10 years	0.170	0.134	0.001	0.068	0.167	0.249	0.595

1.4 Decomposition of yields

In this section, I decompose the Turkish government local currency nominal bond yields into their risk-neutral and term premium components using the methodology outlined in ACM (2013).

1.4.1 The ACM model

In Gaussian ATSMs, it is assumed that the state variables (also referred to as risk factors or pricing factors) X_t evolve under the physical measure according to the following first order VAR model:

$$X_{t+1} = \mu + \Phi X_t + v_{t+1} \quad (1)$$

where the shocks v_{t+1} are conditionally Gaussian with zero mean and a conditional variance-covariance matrix Σ , denoted as $v_{t+1} \sim N(0, \Sigma)$.

Letting $P_t^{(n)}$ indicate the price of a zero-coupon bond with maturity n at time t , the assumption of no-arbitrage implies that there exists a pricing kernel or stochastic discount factor M_t such that:

⁶³ I still acknowledge that a part of the difference in RMSE values may be attributable to differences in data frequencies.

$$P_t^{(n)} = E_t[M_{t+1}P_{t+1}^{(n-1)}] \quad (2)$$

The pricing kernel M_t is assumed to be in the following exponentially affine functional form:

$$M_{t+1} = \exp(-r_t - \frac{1}{2}\lambda_t' \lambda_t - \lambda_t' \Sigma^{-1/2} v_{t+1}) \quad (3)$$

where $r_t = -\ln P_t^{(1)}$ denotes the continuously compounded risk-free rate, and λ_t represents the market prices of risk. Both r_t and λ_t are further assumed to be affine functions of the underlying state variables:

$$r_t = \delta_0 + \delta_1 X_t \quad (4)$$

$$\lambda_t = \Sigma^{-1/2}(\lambda_0 + \lambda_1 X_t) \quad (5)$$

The log excess one-period holding return of a zero-coupon bond maturing in n periods is defined as follows:

$$rx_{t+1}^{(n-1)} = \ln P_{t+1}^{(n-1)} - \ln P_t^{(n)} - r_t \quad (6)$$

ACM (2013) show that using equation (3) and equation (6) in equation (2) leads to:

$$1 = E_t \left[\exp \left(rx_{t+1}^{(n-1)} - \frac{1}{2} \lambda_t' \lambda_t - \lambda_t' \Sigma^{-1/2} v_{t+1} \right) \right] \quad (7)$$

Then, assuming that $\{rx_{t+1}^{(n-1)}, v_{t+1}\}$ are jointly normally distributed, ACM (2013)

find that expected excess return follows:

$$E_t[rx_{t+1}^{(n-1)}] = Cov_t[rx_{t+1}^{(n-1)}, v_{t+1}' \Sigma^{-1/2} \lambda_t] - \frac{1}{2} Var_t[rx_{t+1}^{(n-1)}] \quad (8)$$

Denoting $\beta_t^{(n-1)'} = Cov_t[rx_{t+1}^{(n-1)}, v_{t+1}'] \Sigma^{-1}$ and using the equation (5), they rewrite

the equation (8) as the following expression:

$$E_t[rx_{t+1}^{(n-1)}] = \beta_t^{(n-1)'} (\lambda_0 + \lambda_1 X_t) - \frac{1}{2} Var_t[rx_{t+1}^{(n-1)}] \quad (9)$$

Using this relationship, ACM (2013) decompose the unexpected excess return into a component that is correlated with v_{t+1} and another component that is conditionally orthogonal:

$$rx_{t+1}^{(n-1)} - E_t[rx_{t+1}^{(n-1)}] = \beta_t^{(n-1)'} v_{t+1}' + e_{t+1}^{n-1} \quad (10)$$

where e_{t+1}^{n-1} is assumed to be conditionally independently and identically distributed (i.i.d.) with zero mean and variance σ^2 . Assuming $\beta_t = \beta$ for all t , ACM (2013) subsequently derive the return generating process for log excess holding period returns, which can be expressed as:

$$rx_{t+1}^{(n-1)} = \beta^{(n-1)'} (\lambda_0 + \lambda_1 X_t) - \frac{1}{2} (\beta^{(n-1)'} \Sigma \beta^{(n-1)} + \sigma^2) + \beta^{(n-1)'} v_{t+1} + e_{t+1}^{(n-1)} \quad (11)$$

The four terms on the right hand side of equation (11) account for expected return, convexity adjustment, priced return innovation and return pricing error, respectively. Letting K be the number of pricing factors, N be the number of maturities for which excess returns are calculated, and T be the number of time periods, equation (11) can be rewritten in a stacked form across maturities and time periods:

$$rx = \beta' (\lambda_0 \iota_T' + \lambda_1 X_-) - \frac{1}{2} (B^* \text{vec}(\Sigma) + \sigma^2 \iota_N) \iota_T' + \beta' V + E \quad (12)$$

where rx is an $N \times T$ matrix of excess holding period returns, $\beta = [\beta^{(1)} \beta^{(2)} \dots \beta^{(N)}]$ is a $K \times N$ matrix of factor loadings, ι_T and ι_N are a $T \times 1$ and $N \times 1$ vectors of ones, $X_- = [X_0 \ X_1 \ \dots \ X_{T-1}]$ is a $K \times T$ matrix of lagged pricing factors, $B^* = [\text{vec}(\beta^{(1)} \beta^{(1)'}) \dots \text{vec}(\beta^{(N)} \beta^{(N)'})]'$ is an $N \times K^2$ matrix, V is a $K \times T$ matrix of VAR innovations, and E is an $N \times T$ matrix of return pricing errors.

Following the derivation of the return generating process for excess holding period returns, as presented in equation (12), ACM (2013) propose a three-step regression-based approach to estimate the model parameters. In the first step, they

estimate equation (1) via OLS. Stacking the estimates of the factor innovations $\hat{v}_{t+1}$ into the matrix $\hat{V}$, they develop an estimator of the state variable variance-covariance matrix $\hat{\Sigma} = \hat{V}\hat{V}'/T$. In the second step, they perform a regression of excess returns on a constant, lagged pricing factors and contemporaneous pricing factor innovations according to the reduced-form of equation (12):

$$rx = a\iota_T' + \beta'\hat{V} + cX_- + E \quad (13)$$

Gathering the regressors into the $(2K + 1) \times T$ matrix $\tilde{Z} = [\iota_T' \hat{V}' X_-']'$, the OLS estimates of a , β' and c are obtained as follows:

$$[\hat{a} \hat{\beta}' \hat{c}] = rx\tilde{Z}'(\tilde{Z}\tilde{Z}')^{-1} \quad (14)$$

The residuals of this regression are also stacked into the $N \times T$ matrix $\hat{E}$, which is then utilized to estimate $\hat{\sigma}^2 = \text{trace}(\hat{E}\hat{E}')/NT$. The second step concludes with the construction of matrix $\hat{B}^*$ using $\hat{\beta}$. In the third step, the market price of risk parameters λ_0 and λ_1 are estimated via the cross-sectional regressions. Noting from equations (12) and (13) that $a = \beta'\lambda_0 - \frac{1}{2}(B^*\text{vec}(\Sigma) + \sigma^2\iota_N)$ and $c = \beta'\lambda_1$, the estimates of the market price of risk parameters are:

$$\hat{\lambda}_0 = (\hat{\beta}\hat{\beta}')^{-1}\hat{\beta}\left(\hat{a} + \frac{1}{2}(\hat{B}^*\text{vec}(\hat{\Sigma}) + \hat{\sigma}^2\iota_N)\right) \quad (15)$$

$$\hat{\lambda}_1 = (\hat{\beta}\hat{\beta}')^{-1}\hat{\beta}\hat{c} \quad (16)$$

Denoting a $K \times (K + 1)$ matrix Λ as equal to $[\lambda_0 \lambda_1]$, the market price of risk estimator can be written in a single expression as:

$$\hat{\Lambda} = (\hat{\beta}\hat{\beta}')^{-1}\hat{\beta}\left[rx + \frac{1}{2}\hat{B}^*\text{vec}(\hat{\Sigma})\iota_T' + \frac{1}{2}\hat{\sigma}^2\iota_N\iota_T'\right]M_{\hat{V}}Z_-'(Z_-M_{\hat{V}}Z_-')^{-1} \quad (17)$$

where $Z_- = [\iota_T' X_-']'$ is a $(K + 1) \times T$ matrix, and $M_{\hat{V}} = I_T - \hat{V}'(\hat{V}\hat{V}')^{-1}\hat{V}$ is a $T \times T$ matrix. The equivalence of equation (17) with equation (15) and equation (16) can be inferred by considering that $\iota_T'M_{\hat{V}}Z_-'(Z_-M_{\hat{V}}Z_-')^{-1} = \varrho_1'$, where ϱ_1 is a $(K +$

1) $\times 1$ vector with first element equal to one and zeros elsewhere and since $[\hat{a} \ \hat{c}] = rxM_{\hat{v}}Z'_{-}(Z_{-}M_{\hat{v}}Z'_{-})^{-1}$. ACM (2013) show that β and Λ have a joint limiting distribution:⁶⁴

$$\sqrt{T} \begin{bmatrix} \text{vec}(\hat{\beta} - \beta) \\ \text{vec}(\hat{\Lambda} - \Lambda) \end{bmatrix} \xrightarrow{d} \mathcal{N} \left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{bmatrix} \mathcal{V}_{\beta} & \mathcal{C}'_{\Lambda, \beta} \\ \mathcal{C}_{\Lambda, \beta} & \mathcal{V}_{\Lambda} \end{bmatrix} \right) \quad (18)$$

with

$$\mathcal{V}_{\beta} = \sigma^2(I \otimes \Sigma^{-1}) \quad (19)$$

The joint limiting distribution given in equation (18) is employed for conducting inference on the significance of a given factor as an explanatory variable for excess bond returns, as well as its role in pricing the various components of interest rate risk.

With this three-step linear regression approach, ACM (2013) estimate the model parameters $(\Phi, \Sigma, \sigma, \beta, \lambda_0, \lambda_1)$. They further estimate the short rate parameters δ_0 and δ_1 by performing an OLS regression of the short rate on a constant and the pricing factors. Next, using all the estimated parameters, ACM (2013) generate the zero coupon yield curve. Based on the previous assumptions, it can be shown that the zero-coupon bond prices are exponentially affine in the state variables:

$$\ln P_t^{(n)} = A_n + B_n' X_t + u_t^{(n)} \quad (20)$$

Through the substitution of equation (20) into equation (6), the excess return can be written as:

$$rx_{t+1}^{(n-1)} = A_{n-1} + B_{n-1}' X_{t+1} + u_{t+1}^{(n-1)} - A_n - B_n' X_t - u_t^{(n)} + A_1 + B_1' X_t + u_t^{(1)} \quad (21)$$

By equating this expression for excess returns with the return generating expression stated in equation (11), they get:

⁶⁴ The analytical expressions for $\mathcal{V}_{\Lambda}$ and $\mathcal{C}_{\Lambda, \beta}$ are provided in the appendix of ACM (2013).

$$A_{n-1} + B'_{n-1}(\mu + \Phi X_t + v_{t+1}) + u_{t+1}^{(n-1)} - A_n - B'_n X_t - u_t^{(n)} + A_1 + B'_1 X_t + u_t^{(1)} = \beta^{(n-1)'}(\lambda_0 + \lambda_1 X_t + v_{t+1}) - \frac{1}{2}(\beta^{(n-1)'} \Sigma \beta^{(n-1)} + \sigma^2) + e_{t+1}^{(n-1)} \quad (22)$$

Letting $A_1 = -\delta_0$ and $B_1 = -\delta_1$, and matching the coefficients, ACM (2013) find that the coefficients in equation (20) are determined by the following system of recursive linear restrictions:

$$A_n = A_{n-1} + B'_{n-1}(\mu - \lambda_0) + \frac{1}{2}(B'_{n-1} \Sigma B_{n-1} + \sigma^2) - \delta_0 \quad (23)$$

$$B'_n = B'_{n-1}(\Phi - \lambda_1) - \delta'_1 \quad (24)$$

$$A_0 = 0, B'_0 = 0 \quad (25)$$

Once the bond pricing parameters A_n and B_n are obtained using equation (23) and equation (24), the model-implied zero-coupon yield for maturity n at time t is calculated as follows:

$$\hat{y}_t^{(n)} = -\frac{1}{n}(A_n + B'_n X_t) \quad (26)$$

Setting the market price of risk parameters λ_0 and λ_1 to zero in equation (23) and equation (24) produces the risk-neutral bond pricing parameters A_n^{RF} and B_n^{RF} . Based on these parameters, the risk-neutral yield, which represents the time t expectation of average future short-term interest rates over the next n periods, can be computed as:

$$RN_t^{(n)} = -\frac{1}{n}(A_n^{RF} + B_n^{RF'} X_t) \quad (27)$$

Finally, the term premium is the difference between the model-implied yield and the risk-neutral yield:

$$TP_t^{(n)} = \hat{y}_t^{(n)} - RN_t^{(n)} \quad (28)$$

1.4.2 Empirical results

By employing the ACM model, I can now decompose the local currency nominal interest rates of Turkish government into their corresponding risk-neutral and term

premium components. The analysis relies on utilization of zero coupon yield curve data derived from the fitted NSS curves in Section 1.3 for the period from February 2005 to June 2022. However, as previously mentioned, between February 2005 and December 2009, the upper bound of the bond maturity spectrum in Turkey was 5 years (see section 1.3.5). Consequently, attempting to estimate the cross-section of zero-coupon yields for maturities ranging from 5 to 10 years within this frame using the corresponding end-of-month NSS curve parameters would lead to misleading results. To address this inherent limitation and maintain the scope of the study over the entire sample period from February 2005 to June 2022, I restrict the yield decomposition analysis to maturities up to 5 years.

To implement ACM methodology, I initially back out the cross-section of zero-coupon yields for maturities $n = 1, 2, \dots, 60$ months using the end-of-month NSS curve parameters. The calculation is performed on a monthly basis throughout the sample period, resulting in a cross-section of 60 maturities and a total of $T = 209$ observations. In line with the baseline specification of the ACM model, I then extract the first 5 PCs from these end-of-month zero-coupon yields. The extracted PCs serve as pricing factors (X_t) within the ACM framework. Table 6 shows the proportion of the total variance of the yields accounted for by each PC. The findings indicate that the first three PCs collectively explain 99.94% of variation in yields, which is consistent with previous empirical research showing that these three factors capture nearly all of the cross-sectional variation in the term structure of interest rates.

Following the identification of the pricing factors, equation (1) is estimated as a first step in the ACM procedure. The specified first-order VAR model decomposes X_{t+1} into a predictable component ($\hat{\mu} + \hat{\Phi}X_t$), and an estimate of the factor innovation $\hat{v}_{t+1}$. Since the yield PCs are extracted from demeaned monthly yields, μ

Table 6. Proportion of the Total Variance of Yields Explained by PCs (%)

Maturity	PCs				
	PC1	PC2	PC3	PC4	PC5
6-month	93.579	5.972	0.432	0.000	0.015
1-year	97.576	2.289	0.032	0.100	0.002
2-year	99.691	0.040	0.258	0.004	0.007
3-year	99.436	0.465	0.043	0.053	0.003
4-year	98.184	1.740	0.069	0.001	0.005
5-year	95.980	3.190	0.601	0.194	0.032
All	97.954	1.766	0.221	0.051	0.007

is set to zero in the estimation. This VAR model allows for the estimation of $\hat{\Phi}$ and $\hat{v}_{t+1}$. The factor innovations $\hat{v}$ are then used to construct the variance-covariance matrix $\hat{\Sigma}$. In the second step, taking the risk free rate as $n = 1$ month yield, I compute log excess holding period returns for zero-coupon bonds with maturities $n = 6, 12, 18, \dots, 60$ months directly based on equation (6), giving a cross-section of $N = 10$ maturities for which there are a total of $T = 209$ observations. Next, these excess bond returns are regressed on a constant, lagged pricing factors and contemporaneous pricing factor innovations in accordance with equation (13) to obtain the estimates of σ and β . In the third step, I estimate of the market price of risk parameters λ_0 and λ_1 by using equations (15) and (16).

Once all the model parameters $(\Phi, \Sigma, \sigma, \beta, \lambda_0, \lambda_1)$ are derived using the three-step estimation approach, I proceed with estimating the short rate parameters δ_0 and δ_1 by regressing one-month zero-coupon yield on a constant and the pricing factors. I thereafter feed all parameters into equations (23) and (24) to obtain the recursive pricing parameters A_n and B_n , which are subsequently employed to compute the model-implied yields according to equation (26). Setting λ_0 and λ_1 to zero in the recursion equations (23) and (24) generates the risk-neutral bond pricing parameters A_n^{RF} and B_n^{RF} , which in turn are utilized, as indicated in equation (27), to calculate risk-neutral yields. Ultimately, by applying equation (28), I compute the term

premium component as the difference between the model-implied yield and the risk-neutral yield.

The ACM framework offers a direct test to explore the presence of unspanned or redundant pricing factors that do not explain the cross sectional variation of excess bond returns. In this respect, ACM (2013) use the Wald test as a model specification test with a null hypothesis that the exposures of bond excess returns to a given pricing factor in the model are jointly equal to zero. Then, letting β_i be i th column of β' , the Wald statistic under the null hypothesis that $\beta_i = 0_{N \times 1}$ is as follows:

$$W_{\beta^i} = \hat{\beta}_i' \hat{\mathcal{V}}_{\beta_i}^{-1} \hat{\beta}_i \stackrel{a}{\sim} \chi^2(N) \quad (29)$$

where $\hat{\mathcal{V}}_{\beta_i}$ is an $N \times N$ diagonal matrix that contains the asymptotic variances of $\hat{\beta}_i$ coefficient estimates. To test for the number of pricing factors necessary to explain excess bond returns, I check whether the K th column of β' within a K factor model is equal to zero. Table 7 reports the Wald test statistics for $K = 3, 4, 5$ and 6 factors and their associated p-values. The test results support the use of a five-factor model as the preferred model specification.⁶⁵

Table 7. Identification Test for the Number of Pricing Factors

	Number of Factors			
	K = 3	K = 4	K = 5	K = 6
Wald Statistic	3013.820	5546.864	6254.903	3379.871
p-value	0.000	0.000	0.000	0.000

The Wald test statistic follows $\chi^2(N = 10)$ distribution.

⁶⁵ If the Wald test is applied to the first four columns of β' in the five-factor specification, the null hypothesis is rejected for each pricing factor too. The Wald test statistics for PC1, PC2, PC3, and PC4 are 6131939.813, 440299.583, 105933.589, and 38477.389, respectively, with all corresponding p-values below 0.0001. Hence, the overall test results, including the result for PC5, imply that the cross-sectional variation of excess bond returns is largely captured by the first five PCs extracted from nominal zero-coupon yields.

In addition to the Wald test statistics, I also compute in-sample RMSE statistics by comparing the observed and model-implied yields for bonds with maturities of 6, 12, 24, 36, 48 and 60 months, as well as all maturities, across the models that incorporate $K = 3, 4, 5$, and 6 pricing factors.⁶⁶ According to Table 8, the five-factor model exhibits the smallest yield fitting errors, with overall RMSE value of 9 bp. Furthermore, consistent with the ACM model's use of excess returns instead of yields for estimating the model parameters, the RMSE values in Table 8 reveal that increasing the number of pricing factors does not necessarily enhance the yield fitting performance of the model across different maturities. For instance, moving from five to six factors results in larger RMSE of pricing errors for all maturities, except the 6-month maturity.

Table 8. Comparison of RMSE Across Models With Different Numbers of Factors

Maturity (months)	RMSE for Yields (bp)			
	K = 3	K = 4	K = 5	K = 6
6	19.11	17.03	12.44	11.76
12	23.54	14.92	5.70	6.89
24	10.49	11.35	5.48	6.83
36	12.79	15.39	5.72	7.92
48	5.68	21.51	6.41	11.33
60	20.66	25.10	8.25	17.17
All	20.31	19.40	8.99	10.58

Table 9 presents the summary statistics of the yield and return pricing errors of the five-factor model. The results indicate that the average yield pricing errors across the maturity spectrum are very small, not exceeding 4 bp in absolute value. Their variability is also notably low, with standard deviations ranging from 5.5 and 7.3 bp, with the exception of the 6-month maturity. Furthermore, in line with the

⁶⁶ Malik and Meldrum (2016) adopt the use of in-sample RMSE statistics as means to test the number of pricing factors.

findings discussed in ACM (2013) with regards to the relationship between yield and return pricing errors, I observe a significant presence of serial correlation in the yield pricing errors while the return pricing errors demonstrate little autocorrelation.⁶⁷ This suggests that the pricing factors in the five-factor model sufficiently account for the predictability in excess holding period returns on nominal bonds.

Table 9. Fit Diagnostics of Five-Factor Model

(a) Yield pricing errors $\hat{u}$ (bp)						
	n = 6	n = 12	n = 24	n = 36	n = 48	n = 60
Mean	3.246	-0.738	0.059	-0.494	1.155	3.835
Std. Dev.	12.038	5.662	5.491	5.715	6.318	7.324
Skewness	1.203	0.809	0.463	0.495	-0.792	1.983
Kurtosis	7.928	4.815	5.567	5.365	5.927	11.575
$\rho(1)$	0.441	0.758	0.672	0.702	0.616	0.468
$\rho(6)$	0.060	0.310	0.346	0.258	0.146	0.110
(b) Return pricing errors $\hat{e}$						
	n = 6	n = 12	n = 24	n = 36	n = 48	n = 60
Mean	-0.010	-0.043	0.013	0.002	0.162	0.151
Std. Dev.	0.933	0.592	1.251	1.662	2.616	4.498
Skewness	0.182	-0.419	0.344	-0.470	-0.211	-0.186
Kurtosis	6.556	6.160	6.307	10.676	7.902	13.559
$\rho(1)$	-0.316	0.072	-0.145	-0.197	-0.258	-0.373
$\rho(6)$	-0.047	-0.160	0.010	0.088	0.018	0.068

$\rho(1)$ and $\rho(6)$ denote autocorrelation coefficients of order one and six.

Figure 5 shows the time series fit of the model. The top two panels, which display the observed and model-implied yields for the 2-year and 5-year bonds, reveal that the five-factor model provides an extremely accurate fit. This is evident as the solid blue lines representing the observed yields and the dashed green lines

⁶⁷ Matching terms in equation (22) also implies that $u_{t+1}^{(n-1)} - u_t^{(n)} + u_t^{(1)} = e_{t+1}^{(n-1)}$. This equivalence in turn suggests that if the yield pricing errors are i.i.d., then the return pricing errors exhibit serial correlation. The likelihood-based estimation methods for ATSMs, such as those employed by JSZ (2011) and HW (2012), assume serially uncorrelated yield pricing errors. This assumption is made in order to mitigate the computational complexity of the estimation process. However, by assuming serially uncorrelated yield pricing errors, these methods implicitly assume the presence of serially correlated return pricing errors, which consequently imply that excess return predictability is not captured by the pricing factors. In contrast, ACM (2013) take a different approach and do not impose any distributional assumption on the yield pricing errors.

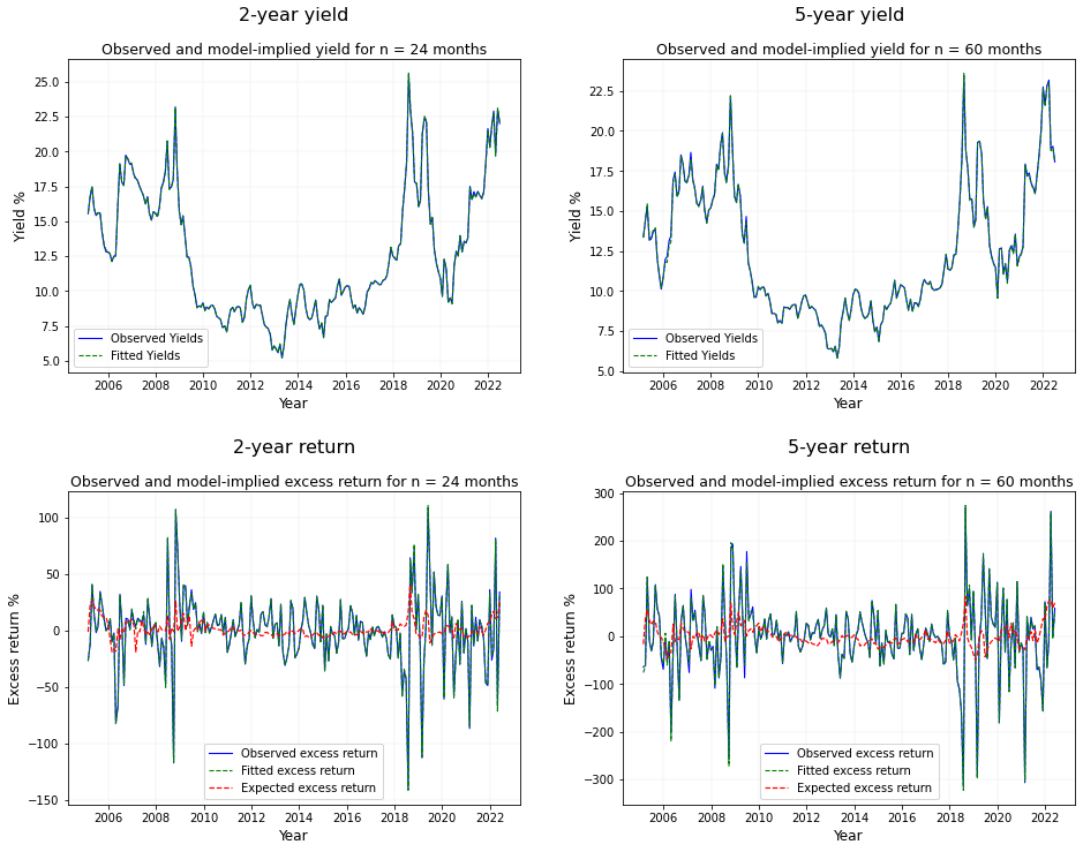


Figure 5. Observed and model-implied yields and excess returns

depicting the model-implied yields are nearly visually indistinguishable. The ACM model's ability to precisely capture the term structure is noteworthy, particularly considering its calibration on excess returns rather than yields. All model parameters are estimated via linear regressions without imposing the cross-equation recursions specified in equations (23) and (24), and these equations are utilized only ex post to compute the model-implied yields. The bottom two panels in Figure 5 illustrate the observed and model-implied excess one-month holding period returns on the same two maturities.⁶⁸ A remarkably tight fit implies that the five-factor model is able to closely replicate the dynamics of the observed excess bond returns. Additionally, the

⁶⁸ The model-implied excess returns are computed by the following expression:

$$\hat{B}'_{n-1}(\hat{\lambda}_0 + \hat{\lambda}_1 X_t) - \frac{1}{2}(\hat{B}'_{n-1}\hat{\Sigma}\hat{B}_{n-1} + \hat{\sigma}^2) + \hat{B}'_{n-1}\hat{v}_{t+1}$$

dashed red lines displayed in the bottom panels indicate the model-implied expected excess return component, which reflects the risk premium that bond investors demand for holding an n -month bond for one month.⁶⁹ It is clear that the compensation for risk exhibits significant temporal variability.

Figure 6 displays both the cross-sectional fit and interpretation of the pricing factors in the five-factor model. The top two panels plot unconditional means and standard deviations of observed and model-implied yields across the maturities. It is evident that the five-factor specification provides a very good fit for both moments.

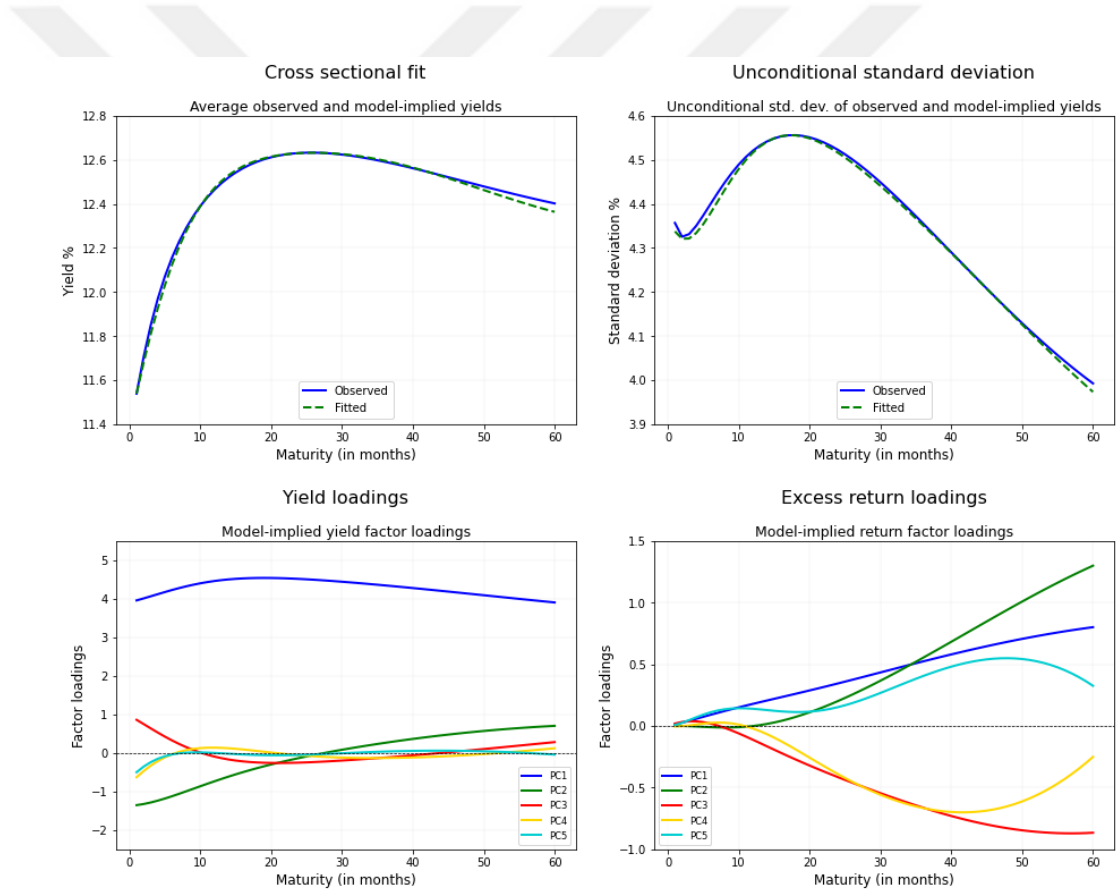


Figure 6. Cross-sectional diagnostics of the five-factor model

⁶⁹ The model-implied expected excess returns are calculated as $\hat{B}'_{n-1}(\hat{\lambda}_0 + \hat{\lambda}_1 X_t) - \frac{1}{2}(\hat{B}'_{n-1} \hat{\Sigma} \hat{B}_{n-1} + \hat{\sigma}^2)$.

The bottom left panel of Figure 6 shows a plot of the estimated yield loadings $b_n = -\frac{1}{n}B_n$ (see equation 26). These coefficients are interpreted as the response of the n -month yield to a contemporaneous shock to the respective pricing factor. The yield loadings on the first three PCs in the chart are consistent with common characterization of these components in the literature. The first PC effectively represents the level factor as a shock to this component moves all yields by a similar amount, ultimately leading to a parallel shift in the term structure. The second PC has negative loadings for the short-end and positive loadings for the long-end of the yield curve. Consequently, a shock to this component causes the yield curve to flatten (negative change) or steepen (positive change). In that sense, it can be regarded as the slope factor. The third PC demonstrates U-shaped factor loadings, altering the curvature of the term structure. Thus, it naturally represents the curvature factor. The same chart also illustrates that the yield loadings on the PC4 and PC5 are extremely small, highlighting their negligible explanatory power for the yield curve variation. However, a different picture emerges when examining the bottom right panel of Figure 6, which portrays expected excess return loadings $B'_n\lambda_1$ on all pricing factors.⁷⁰ In line with ACM (2013) findings, the chart reveals that the PC4 and PC5 are important drivers of expected excess returns. Moreover, expected excess returns have positive loadings on the PC2 with respective coefficients almost linearly increasing across the maturity spectrum. This result is in accordance with earlier studies, such as Campbell and Shiller (1991), which underscore the predictive ability of yield spreads for excess bond returns. The bottom right plot also shows that contemporaneous shocks to the PC2 and PC3 separately have a greater impact on

⁷⁰ The corresponding coefficients provide an estimation of how the expected one-month excess holding return on an n -month bond responds to a contemporaneous shock to the respective pricing factor.

expected excess return than a contemporaneous shock to the PC1 for maturities beyond the short end. For example, a one standard deviation increase in the PC2 and PC3 changes the annualized expected excess returns on the 5-year bond by 15.6% and -10.4%, respectively, whereas the corresponding change is 9.6% for the PC1. Taken together, the results indicate that while yields are predominantly explained by the first three PCs, expected excess returns are explained by all the PCs, including the higher-order PC4 and PC5.

To explore the role of each factor in pricing the various components of interest rate risk in the five factor model, I report the estimates of the market price of risk parameters and the corresponding t -statistics in Table 10. First, although the price of level risk has a negative constant component (the top element of λ_0 vector), which would normally suggest that a typical investor would expect a positive excess return for holding the level portfolio, the observed effect is not statistically significant.⁷¹ The first row of the table also provides evidence that the level risk, as measured by the exposure to the first PC, is significantly determined solely by the first PC itself. The negative sign of the loading of level risk on the level factor implies that higher level component is associated with higher expected excess returns. As indicated by the second row, the slope risk, which is measured by the exposure to the second PC, is driven by the slope factor itself, the curvature factor, as well as the fourth PC. In terms of the curvature risk, it is affected by all the pricing factors, except the level factor. Importantly, the results underscore highly important roles of the fourth and fifth PCs on the expected excess returns. While the fourth PC is a significant determinant of market prices of slope risk, curvature risk and risk of

⁷¹ Excess return betas are represented as the negative multiples of the yield loadings $b_n = -\frac{1}{n}B_n$. Also see footnote 69 for the calculation of the model-implied expected excess returns.

Table 10. Market Prices of Risk

Factor	λ_0	$\lambda_{1.1}$	$\lambda_{1.2}$	$\lambda_{1.3}$	$\lambda_{1.4}$	$\lambda_{1.5}$	W_Λ	W_{λ_1}
PC1	-0.018	-0.040	-0.032	0.038	0.027	-0.029	14.860	14.120
(t-statistic)	(-0.904)	(-2.022)	(-1.623)	(1.876)	(1.353)	(-1.453)	(0.021)	(0.015)
PC2	0.074	0.003	-0.168	0.116	0.089	0.062	38.526	34.749
(t-statistic)	(1.899)	(0.077)	(-4.304)	(2.877)	(2.288)	(1.596)	(0.000)	(0.000)
PC3	-0.111	-0.028	-0.100	-0.214	-0.240	-0.152	55.488	50.468
(t-statistic)	(-2.172)	(-0.548)	(-1.956)	(-4.062)	(-4.699)	(-2.949)	(0.000)	(0.000)
PC4	0.033	0.004	0.081	0.034	-0.412	0.156	50.568	50.264
(t-statistic)	(0.532)	(0.060)	(1.286)	(0.526)	(-6.505)	(2.437)	(0.000)	(0.000)
PC5	-0.005	-0.068	-0.049	0.056	-0.020	-0.648	75.021	75.021
(t-statistic)	(-0.073)	(-0.926)	(-0.667)	(0.732)	(-0.271)	(-8.543)	(0.000)	(0.000)

This table provides the estimates of the market price of risk parameters λ_0 and λ_1 for the five-factor model. Associated t -statistics are reported in parenthesis. Wald statistics for tests of the rows of Λ and of λ_1 being different from zero are presented along each row, with the corresponding p -values in parentheses below.

the fourth component itself, the movements of the fifth PC influence the market prices of curvature risk, risk of the fourth component and risk of the fifth component itself. Following ACM (2013), I can summarize the pricing implications of the five-factor model by testing the null hypothesis that different rows of $\Lambda = [\lambda_0 \ \lambda_1]$ and λ_1 are jointly different from zero. In particular, letting $\lambda'_{i.}$ be the i th row of $\Lambda = [\lambda_0 \ \lambda_1]$, the Wald statistic for the null hypothesis that $\lambda'_{i.} = 0_{1 \times (K+1)}$ is given by:

$$W_{\Lambda_{i.}} = \hat{\lambda}'_{i.} \hat{\mathcal{V}}_{\lambda_{i.}}^{-1} \hat{\lambda}_{i.} \stackrel{a}{\sim} \chi^2(K+1) \quad (30)$$

where $\hat{\mathcal{V}}_{\lambda_{i.}}$ is a $(K+1) \times (K+1)$ matrix that contains the estimated variances of the $\hat{\lambda}_{i.}$ coefficient estimates. This test allows for the assessment of whether a given risk factor is priced in the cross-section of excess returns. In a similar manner, I can determine whether the price of risk of a given factor displays significant time variation by testing if the corresponding row of λ_1 is jointly equal to zero. Letting $\lambda'_{1i.}$ be the i th row of λ_1 , the Wald statistic for the null hypothesis that $\lambda'_{1i.} = 0_{1 \times K}$ is defined as:

$$W_{\lambda_{1i.}} = \hat{\lambda}'_{1i.} \hat{\mathcal{V}}_{\lambda_{1i.}}^{-1} \hat{\lambda}_{1i.} \stackrel{a}{\sim} \chi^2(K) \quad (31)$$

The last two columns of Table 10 present these Wald statistics and associated p -values for the rows of Λ and λ_1 , respectively. The first test results indicate that all factors are strongly priced in the five-factor specification. The second test results also reveal significant temporal variation in the risk prices of all five factors.

I can also analyze the contribution of risk prices of all five factors to the variability of the pricing kernel. Recalling $\ln M_{t+1} = -r_t - \frac{1}{2} \lambda_t' \lambda_t - \lambda_t' \Sigma^{-1/2} v_{t+1}$ (see equation 3), the conditional volatility of the pricing kernel can be decomposed into the contributions from each price of risk as follows:

$$\text{Var}_t(\ln M_{t+1}) = \lambda_t' \lambda_t = \sum_{j=1}^K \lambda_{jt}^2 \quad (32)$$

Figure 7 shows that risk prices of all five factors contribute to the volatility of the pricing kernel, consistent with the findings reported in ACM (2013) for the U.S. market.

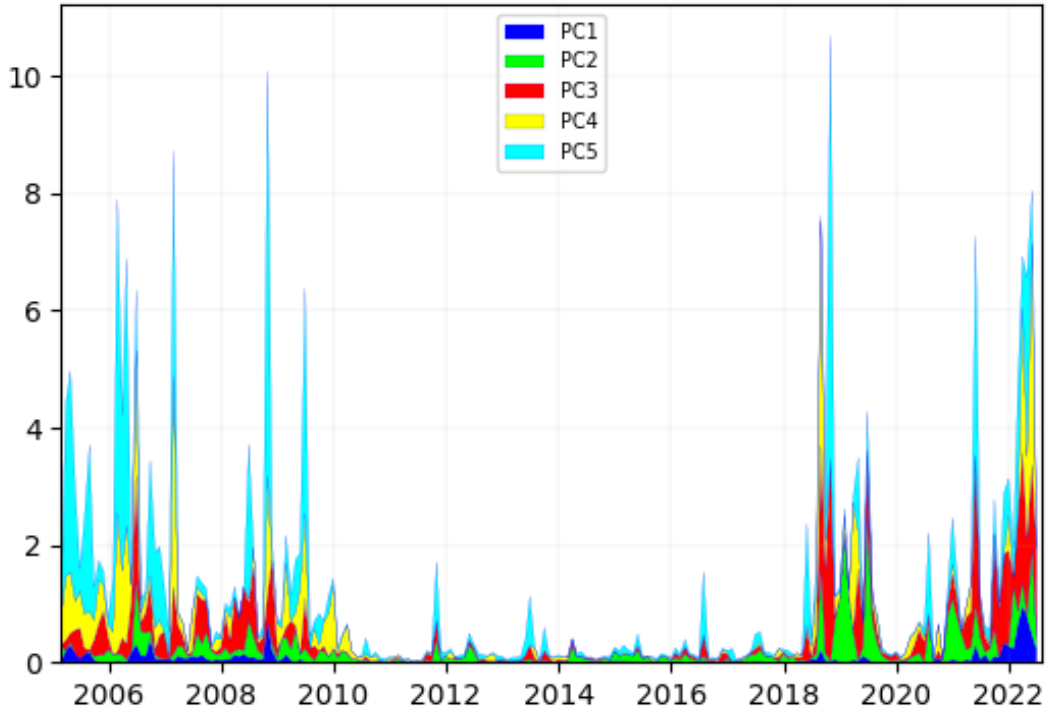


Figure 7. Decomposition of pricing kernel variance in the five-factor model

Next, I derive the model-implied maximal Sharpe ratio to evaluate the validity of the stochastic discount factor. From equation (8), the conditional Sharpe ratio for the log excess returns in the dynamic term structure models is computed as:

$$\frac{E_t[rx_{t+1}^{(n-1)}] + \frac{1}{2} Var_t[rx_{t+1}^{(n-1)}]}{\sqrt{Var_t[rx_{t+1}^{(n-1)}]}} = \frac{Cov_t[rx_{t+1}^{(n-1)}, \ln M_{t+1}]}{\sqrt{Var_t[rx_{t+1}^{(n-1)}]}} \quad (33)$$

The Sharpe ratio in equation (33) is maximized for a hypothetical return to a portfolio of bonds that spans the payoff of the log pricing kernel M_t . Then, the model-implied maximal Sharpe ratio based on log excess returns is equal to:

$$\sqrt{Var_t[\ln M_{t+1}]} = \sqrt{\lambda_t' \lambda_t} \quad (34)$$

Duffee (2011) argues that the fully-flexible four- and five-factor Gaussian term structure models produce quite high conditional maximal Sharpe ratios because of overfitting. However, in line with the results of ACM (2013), I find very reasonable maximal Sharpe ratios for my baseline five-factor specification. Figure 8 displays that the maximal Sharpe ratio peaks at 3.27, while the sample average stands at 0.88, which is below 1.⁷²

The recursive pricing formulas in equation (23) and equation (24) are derived by equating the return generating process in equation (11) with the definition of excess returns in equation (6) based on the assumption that log bond prices and therefore yields are linear in the states variables X . Matching terms on left-hand side and right-hand side of equation (22) implies that the recursive pricing parameters B_n , which are functions of the estimated parameters $\Phi, \Sigma, \sigma, \lambda_0, \lambda_1, \delta_0$ and δ_1 , need to be equal to the coefficients $\beta^{(n)}$ from the time series regressions of log excess holding period returns on lagged pricing factors and the contemporaneous pricing factor

⁷² The maximal Sharpe ratio implied by the five-factor model in ACM (2013) has a sample average of 0.67 and reaches its peak value at 2.30.

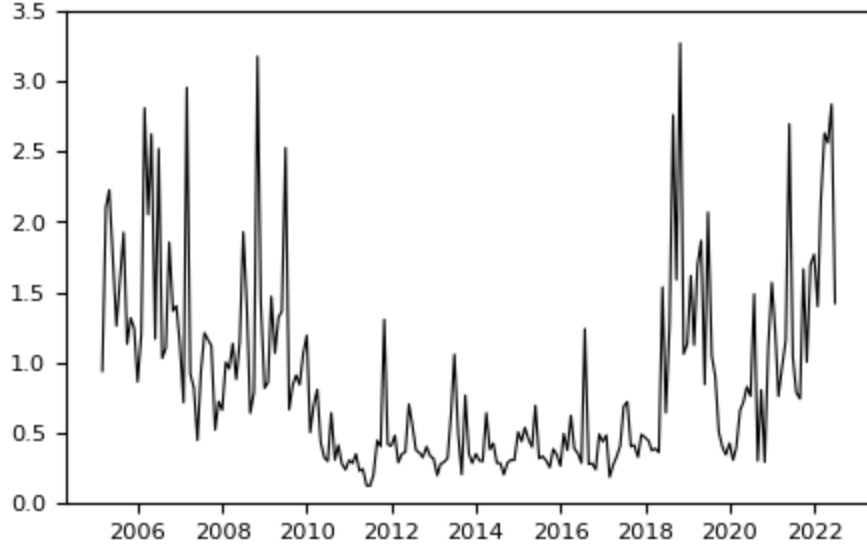


Figure 8. Maximal Sharpe ratio implied by the five-factor specification

innovations v for all maturities n . Nevertheless, this equality restriction is not enforced in the estimation. As suggested by ACM (2013), the difference between $\hat{B}_n$ and $\hat{\beta}^{(n)}$ can be considered as an additional diagnostic tool to assess the ability of the affine model to capture the dynamics of the excess returns. Figure 9 illustrates that these two sets of coefficients exhibit a high degree of similarity across all maturities for the first three factors. Although there are slight dissimilarities between the coefficients for the fourth and fifth factors, the differences are economically small. In conclusion, the five-factor specification is sufficiently flexible to replicate the dynamics of the yield curve and excess returns.

Finally, I present the results the decomposition of nominal yields into risk-neutral and term premium components. Table 11 provides the summary statistics for the estimates of these two components across various fixed maturities. It is evident that while the average risk-neutral yield component has a downward sloping pattern, the average term premium component exhibits a hump-shaped pattern with a peak around 3 years. The volatilities of these components demonstrate a more contrasting

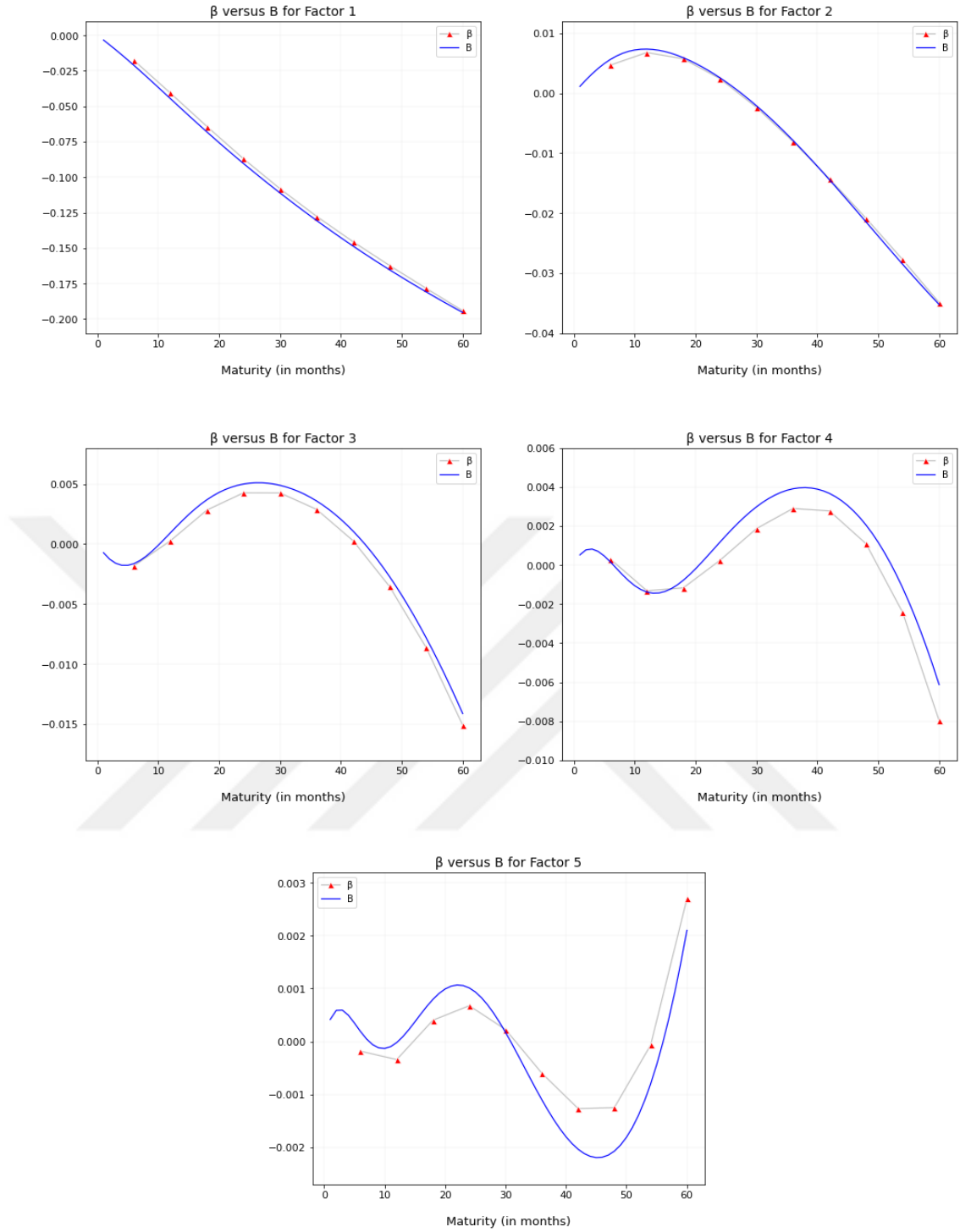


Figure 9. Regression coefficients $\beta^{(n)}$ versus recursive pricing parameters B_n

picture. Although the volatility of risk-neutral yields decreases with maturity, that of term premium increases with maturity. Figure 10 illustrates the temporal evolution of both risk-neutral and term premium components, manifesting substantial fluctuations

Table 11. Summary Statistics for Risk-Neutral and Term Premium Components

(a) Risk-Neutral Yield								
Maturity	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
6-month	0.1153	0.0397	0.0525	0.0805	0.1024	0.1497	0.2237	209
1-year	0.1152	0.0365	0.0562	0.0832	0.1027	0.1483	0.2071	209
2-year	0.1148	0.0311	0.0638	0.0877	0.1047	0.1437	0.1884	209
3-year	0.1143	0.0268	0.0702	0.0909	0.1057	0.1394	0.1766	209
4-year	0.1139	0.0233	0.0754	0.0934	0.1062	0.1357	0.1678	209
5-year	0.1134	0.0205	0.0796	0.0954	0.1066	0.1326	0.1608	209

(b) Term Premium								
Maturity	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
6-month	0.0059	0.0051	-0.0021	0.0020	0.0045	0.0095	0.0211	209
1-year	0.0095	0.0092	-0.0054	0.0020	0.0061	0.0168	0.0360	209
2-year	0.0115	0.0152	-0.0116	-0.0003	0.0069	0.0209	0.0693	209
3-year	0.0116	0.0183	-0.0160	-0.0021	0.0065	0.0242	0.0824	209
4-year	0.0110	0.0200	-0.0194	-0.0038	0.0048	0.0245	0.0825	209
5-year	0.0102	0.0211	-0.0216	-0.0056	0.0037	0.0235	0.0792	209

The sample period is from February 2005 to June 2022.

over the entire sample period. To assess their roles in the dynamics of nominal yields, I conduct a variance decomposition analysis. Using equation (28), the variance of nominal yield can be written as:

$$Var[\hat{y}_t^{(n)}] = Cov[\hat{y}_t^{(n)}, \hat{y}_t^{(n)}] = Cov[\hat{y}_t^{(n)}, RN_t^{(n)}] + Cov[\hat{y}_t^{(n)}, TP_t^{(n)}] \quad (35)$$

Dividing each side of equation (35) by $Var[\hat{y}_t^{(n)}]$ gives the following expression:

$$1 = \frac{Cov[\hat{y}_t^{(n)}, RN_t^{(n)}]}{Var[\hat{y}_t^{(n)}]} + \frac{Cov[\hat{y}_t^{(n)}, TP_t^{(n)}]}{Var[\hat{y}_t^{(n)}]} \quad (36)$$

Based on equation (36), Table 12 reports the variance decompositions of nominal yields in both their level and first-difference forms. The results in Table 12a indicate that risk-neutral rate plays a substantial role in explaining the level form of nominal yield, but its importance diminishes with maturity. While the term premium becomes the dominant factor solely at the 5-year maturity horizon, it remains a significant contributor, explaining at least 39.9% of the variation as of 3-year maturity. Since nominal yields display a high degree of persistence, I also examine the variance

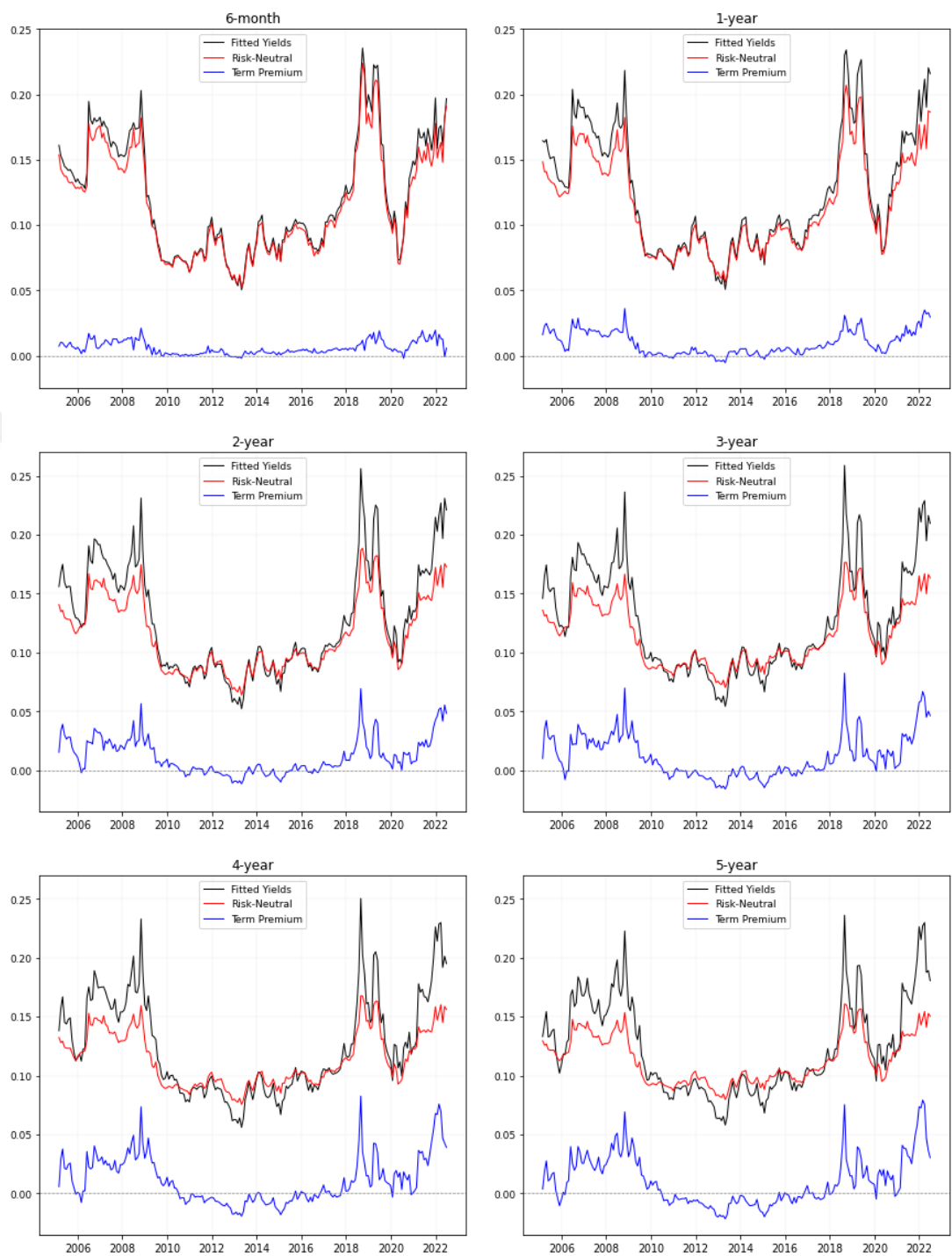


Figure 10. Nominal yield decomposition

decompositions of their monthly changes.⁷³ Table 12b reveals that when yields and their components are expressed in first differences, the contribution of term premiums to the variation of changes in yields increases across all maturities compared to the level case. In addition, the term premium component becomes the leading contributor starting from the 3-year maturity.

Table 12. Variance Decomposition of Nominal Yields

(a) Levels						
	6-month	1-year	2-year	3-year	4-year	5-year
Risk-Neutral	0.902	0.804	0.680	0.601	0.541	0.493
Term Premium	0.098	0.196	0.320	0.399	0.459	0.507
(b) First-Differences						
	6-month	1-year	2-year	3-year	4-year	5-year
Risk-Neutral	0.877	0.758	0.534	0.413	0.357	0.333
Term Premium	0.123	0.242	0.466	0.587	0.643	0.667

The variance decomposition analysis relies on equation (36).

1.4.3 Context for the evolution of risk-neutral and term premium components

In this sub-section, I delve into an analysis of the fluctuations observed in the risk-neutral and term premium components by providing a broad economic and monetary context to their swings.

As evident in Figure 10, the risk-neutral components of yields were on the rise from the beginning of 2008. In the first quarter of 2008, the surge in the food and energy prices along with the growing uncertainties in the global economy worsened the medium-term inflation expectations and heightened the upside risk on Turkish inflation figures. As a response, under its price stability mandate, the CBT decided to halt the rate cuts in March 2008 and signaled on its April 2008 meeting the adoption of a clear and immediate tightening bias to mitigate the potential second-round

⁷³ Crump, Eusepi, and Moench (2016) and Naudon and Yany (2016) also perform variance decompositions in both the level and first-difference forms of yields.

effects of supply-side shocks.⁷⁴ Subsequently, the CBT delivered the measured rate hikes in the policy rate during the May, June, and July 2008 meetings, incrementally raising the policy rate by 50 bp in each consecutive meeting.⁷⁵ Although inflation expectations moderated in August and September 2008 due to lower food and energy prices and previous monetary tightening, the sharp depreciation of the Turkish lira during the peak of the global financial turmoil in October 2008 had a significant impact on the medium term inflation outlook. Against this backdrop, the risk-neutral rates for different maturities surged to their highest levels of the year, reflecting investors' upward revisions in the expected path of future short-term interest rates.

In November 2008, the repercussions of the GFC on the Turkish economy started to become more evident. The CBT acknowledged that severe challenges in the international credit markets and the global economy would continue to impede both domestic and external demand for a prolonged period, thereby limiting the pass-through from the exchange rate depreciation to domestic prices. The CBT also envisaged that the significant fall in oil and other commodity prices since September 2008 would support disinflationary trend. In the light of these factors, anticipating a more rapid decline in inflation compared to its previous projections, the CBT initiated a policy rate cut cycle by lowering the overnight borrowing rate by 50 bp in the November 2008 meeting. Within a year, the cumulative policy rate cut amounted to 1025 bp, bringing the policy rate down to 6.5% by November 2009. During this period, the risk-neutral components of local bond yields witnessed sharp declines as the market participants revised downward their expectations of future short-term interest rates, mainly consistent with the inflation projections of the CBT. Within this

⁷⁴ After implementing an implicit inflation-targeting regime between 2002 and 2005, the CBT shifted to an explicit inflation-targeting framework at the beginning of 2006.

⁷⁵ The overnight borrowing rate, which served as the policy rate prior to May 2010, was raised from 15.25% to 16.75% in this period.

time frame, 2-year and 5-year risk-neutral rates dropped by 933 bp and 619 bp, respectively.

After demonstrating a relatively stable pattern through 2010 and in the first half of 2011, the risk-neutral yields underwent a temporary surge between August and December 2011. In fact, starting from August 2011, the deterioration in the global risk appetite due to deepening sovereign debt problems in the Euro area triggered significant capital outflows from emerging market countries, including Turkey. In October 2011, the central bank adopted a monetary tightening stance as the excessive depreciation in Turkish lira, coupled with the hikes in administered prices and base effects from unprocessed food prices, began to exert a negative impact on the medium-term inflation outlook. Consequently, in accordance with the new monetary policy framework implemented as of November 2010, which aimed to uphold both price stability and financial stability, the CBT widened the interest rate corridor upwards by raising the overnight lending rate from 9% to 12.5% in October 2011.⁷⁶ Moreover, in the period from December 29, 2011, to January 9, 2012, making use of the flexibility offered by the interest rate corridor, the central bank also implemented additional monetary tightening by temporarily suspending market funding through its usual one-week repo auctions at the prevailing policy rate of 5.75%, resulting in a significant increase in the average cost of funding and

⁷⁶ In November 2010, the CBT modified its existing inflation targeting framework, which had been in place since 2006, by incorporating the preservation of financial stability as a supplementary objective alongside its primary goal of achieving and maintaining price stability. During the period from November 2010 to August 2011, the CBT focused on the implementing the macroprudential tightening measures to bolster the resilience of the economy against vulnerabilities associated with external balance, credit expansion and capital flows, while simultaneously safeguarding price stability. To effectively address both the inflation and macro-financial risks, the new policy necessitated the combined and active use of interest rate corridor between overnight borrowing and lending rates, the reserve requirement ratios, and one-week repo auction rate (the policy rate as of May 2010). For more details, see Başcı and Kara (2011) and Kara (2013).

expectations of higher future short-term interest rates.⁷⁷ Between March 22, 2012 and June 4, 2012, five rounds of the additional monetary tightening were delivered by the central bank to mitigate the impact of temporary factors, such as the increase in oil and other commodity prices, on inflation expectations.⁷⁸

However, as of early June 2012, the CBT shifted to a more accommodative stance due to disinflationary impact of aggregate demand conditions, along with an increase in global risk appetite led by the adoption of new expansionary monetary policies by the central banks of advanced economies.⁷⁹ Accordingly, the CBT increased its liquidity injection into the market and gradually reduced its average funding rate in the following months.⁸⁰ During this period, the risk-neutral rates followed a downward trajectory, indicating a revision by market participants of their expectations for future short-term interest rates. In September 2012, foreseeing a faster decline in inflation and expecting a surge in short-term capital inflows towards emerging market economies in response to the monetary policy easing by the major central banks, the CBT decided to narrow the interest rate corridor by cutting the overnight lending rate from 11.5% to 10%. In the last quarter of 2012, with the

⁷⁷ During these “exceptional days” of additional monetary tightening, the CBT opted for intraday one-week repo auctions under the standard auction method, instead of holding its regular one-week repo auctions via the quantity auction method at the policy rate. Throughout the specified time frame, the average funding rate in intraday one-week repo auctions exhibited fluctuations, with a minimum of 10.63% on December 29, 2011, a peak of 12.02% on January 2, 2012, and final rate of 11.96% on January 9, 2012.

⁷⁸ The duration of the additional monetary tightening episodes between March and June 2012 spanned from 3 to 6 days. During these episodes, the average funding rates in intraday one-week repo auctions ranged from 10.51% (April 17, 2012) to 10.92% (April 12, 2012), significantly exceeding the CBT’s policy rate of 5.75% applied in regular one-week repo auctions.

⁷⁹ From early June to mid-September, the central banks of advanced economies, including the FED, the ECB and the BOE, implemented several additional monetary easing measures. Specifically, on June 20, 2012, the FED announced the extension of Operation Twist for an additional six months. Then, on July 5, 2012, the ECB reduced its main refinancing rate by 25 bp to 0.75%, while on the same date, the BOE increased its QE program. Moreover, on September 6, 2012, the ECB unveiled its new bond buying program known as OMT (see footnote 25), and shortly after, on September 13, 2012, the FED launched its third round of QE.

⁸⁰ The CBT’s average funding rate declined from 9.02% on June 29 to 7.63% on July 31, further reducing to 6.33% on August 31, and reaching 6.15% on September 17, 2012, one day before its September Monetary Policy Committee meeting.

acceleration of capital inflows to Turkey, the CBT continued to lower short-term interest rates to contain macro financial risks, while implementing a moderate tightening via reserve requirements. In this period, the medium-term inflation expectations also trended lower. In December 2012, the 24-month-ahead inflation expectations fell below 6% for the first time since April 2008. The capital inflows, while remaining mostly strong in the first quarter of 2013, reaccelerated following the BOJ's announcement of monetary easing policy in early April.⁸¹ Anticipating a further acceleration of capital inflows in the near future, the CBT maintained its stance of lowering short-term interest rates while simultaneously tightening through reserve requirement ratios. Against this backdrop, as depicted in Figure 10, the risk-neutral yields reached their historical low levels in April 2013, shortly before the FED's announcement in May 2013 of its intention to gradually wind down the QE program.

The global bond markets perceived the taper talk as a signal of earlier than expected rise in short-term policy rates in the U.S. The sudden shift in the expected path of U.S. short-term interest rates then prompted an increase in the U.S. long-term bond yields, which subsequently diminished the attractiveness of local currency emerging market debt and triggered sizeable capital outflows from these economies. To manage the challenges posed by currency depreciation and its potential inflationary implications, several emerging market countries responded by raising their short-term policy rates. The higher path of short-term policy rates in turn translated into higher long-term bond yields (Klemm, Meier, & Sosa, 2014). That said, during the taper talk period, the CBT kept the one-week repo rate (the policy

⁸¹ On April 04, 2013, the BOJ adopted an aggressive monetary easing policy known as “quantitative and qualitative monetary easing”, which aimed to combat deflation and foster economic growth in Japan. It included a substantial increase in the BOJ's asset purchase program with the aim of doubling the monetary base over a 2-year period.

rate) constant at 4.5%. Instead, in line with new policy framework adopted in November 2010, in order to address the negative impacts of excessive exchange rate volatility on price and financial stability, the CBT widened the interest rate corridor by raising the overnight lending rate from 6.5% to 7.25% on July 23, 2013, and further increased it to 7.75% on August 20, 2013. In parallel, given potential upside risks to pricing behavior, the CBT also implemented additional monetary tightening measures by adjusting the liquidity provided to the market at the policy rate and conducting unsterilized foreign exchange sale auctions. In the period from the end of April to end of August 2013, the risk-neutral components of 2-year and 5-year maturity local bonds in Turkey experienced substantial increases of 272 bp and 181 bp, respectively, driven by the upward shift in investors' expected path of future short-term interest rates amid the CBT's tight and cautious monetary policy stance.

After a period of relative stability, the risk-neutral rates for Turkey displayed a notable change in its trajectory following the U.S. presidential election in November 2016, which triggered a significant repricing of assets, particularly in emerging market economies. The expectation that the new U.S. administration would adopt expansionary fiscal policies, which in turn could prompt the FED to accelerate interest rates hikes, led to a strong rally in the U.S. dollar and a substantial surge in the U.S. interest rates. In December 2016, the FED raised the interest rates and signaled three more hikes in 2017, boosting the U.S. dollar further to a 14-year high against a basket of major currencies. The U.S. bond yields also rose, reaching seven-year highs for 2-year notes and 5.5-year highs for 5-year notes. During this period, soaring U.S. yields attracted the capital away from emerging market economies, including Turkey. In the aftermath of the U.S. presidential election, while the CBT immediately raised its one-week repo rate and overnight lending rate by 50 bp and 25

bp to 8.00% and 8.50%, respectively, at its November meeting, given subsequent further depreciation in the Turkish lira and worsening medium-term inflation outlook, it started to implement a set of liquidity tightening measures in mid-January 2017.⁸² This was later followed by a more aggressive monetary tightening through hikes in marginal funding rate and late liquidity window lending rate by 75 bp and 100 bp to 9.25% and 11%, respectively, in late January 2017. Considering the lagged spillover effects of exchange rate depreciation on inflation, the CBT not only raised the share of the late liquidity window funding within the overall funding but also increased the late liquidity window lending rate further in the March and April meetings to 11.75% and 12.25%, respectively.⁸³ Amid the sustained tight monetary policy stance and increasing average cost of funding, the risk-neutral rates shifted upward, with 2-year and 5-year tenures rising by 172 bp and 104 bp, respectively, in the period from end-September 2016 to end-April 2017.

In the subsequent months, a more favorable outlook in the global economic activity, combined with moderate and predictable monetary policy normalizations by the central banks of advanced economies due to low inflation rates, supported the global risk appetite, which in turn drove strong portfolio inflows to emerging market economies.⁸⁴ However, the Turkish lira started to decouple from its emerging market peers as of October 2017 due to geopolitical tensions and deterioration in the medium-term inflation outlook. The consumer inflation reached its highest level in

⁸² As of January 12, 2017, one-week repo auctions were suspended. In addition to that, starting from January 16, 2017, the CBT introduced a limit on the amount of funding provided through BIST Repo and Reverse Repo Market, which forced banks to meet a part of their liquidity needs through the late liquidity window facility at 10%. As a result of all these measures, there was a notable increase in the CBT's average funding rate, reaching 9.12% on January 20, 2017.

⁸³ The CBT's average funding rate at the end of April was 11.80%.

⁸⁴ Despite the FED announcing in its September 2017 meeting that it would begin to unwind its balance sheet as of October and the ECB deciding in its October meeting to halve its bond purchases starting from January 2018, portfolio inflows into emerging market remained strong, albeit at a slower pace than in the second and third quarters.

the past nine years at 11.9% in October, mainly driven by the depreciation of Turkish lira against the currency basket, rising global commodity prices, and strong domestic economic activity on the back of the Credit Guarantee Fund loans and macroprudential incentives.⁸⁵ Importantly, the core inflation measured by the C index surged to 11.82%, marking its highest level since January 2004 during the same month. Given the heightened exchange rate volatility after multi-year high inflation figures, starting from November 22, 2017, the CBT reduced the banks' overnight borrowing limits in the Interbank Money Market to zero, leading to the CBT's funding being solely provided via the late liquidity window facility. Following this unorthodox liquidity move, the CBT also raised its late liquidity window lending rate by 50 bp to 12.75% during its December meeting, while keeping the one-week repo rate, overnight borrowing and lending rates all unchanged. The upbeat US job data in early February 2018, which fueled inflation worries and triggered speculations of more aggressive monetary tightening by the FED, along with concerns about U.S. foreign trade protectionism, led to a substantial deterioration in the global risk appetite and resulted in depreciation of emerging market currencies. In this period, the Turkish lira continued to underperform its emerging market peers, mainly due to worsening inflation expectations, widening current account deficit and escalating geopolitical risks. In response to the Turkish lira hitting a series of record lows, the CBT lifted its late liquidity window lending rate by 75 bp to 13.50% in April 2018, followed by another significant hike of 300 bp in an extraordinary monetary policy meeting in May 2018. Despite these substantial hikes, the currency continued to plunge, prompting the CBT to finalize its

⁸⁵ The October 2017 inflation figures were announced on November 3rd, 2017. In Turkey, the inflation figures for a given month are typically released on the 3rd of the following month. If the 3rd falls on a holiday, the release is delayed to the next business day.

longstanding monetary policy simplification process. Accordingly, on May 28, 2018, the CBT announced that the one-week repo rate would be set at the prevailing funding rate of 16.5% and reconsidered as the main policy rate, effective from June 1, 2018.⁸⁶ Additionally, the overnight borrowing and lending rates would be determined at 150 bp below and above the one-week repo rate, respectively. In its June meeting, the central bank further increased the one-week repo rate by 125 bp to 17.75% in response to the uptick in inflation figures for May. Surprisingly, despite consumer inflation reaching a 14-year high of 15.39% in June, the central bank chose to keep rates unchanged in July as it considered the need to monitor the deceleration in domestic economic activity and the lagged effects of previous monetary policy decisions. Growing political tension between Turkey and the U.S. as of late July 2018, coupled with already large current account deficit and soaring inflation, resulted in an exchange rate crisis in early August, during which the Turkish lira faced a sharp depreciation.⁸⁷ As an initial step towards tightening the monetary policy, the CBT decided not to open its regular one-week repo auction on August 13, 2018, and instead opted to provide funding to the market at an overnight lending rate of 19.25%. Due to continued concerns about the inflation outlook, the CBT implemented a strong monetary tightening, raising its one-week repo rate from 17.75% to 24% on September 13, 2018. It also announced the resumption of one-week repo auctions as of September 14, 2018, after their suspension since August 13, 2018. In the period from the end of September 2017 to the end of September 2018, during which Turkey exhibited a notable negative divergence from its peers, there was a substantial readjustment in the risk-neutral component of yields, given the

⁸⁶ As of June 1, 2018, the CBT resumed its repo funding, which had been paused since January 2017 (see footnote 82).

⁸⁷ During early trading hours on August 13, 2018, the Turkish lira depreciated to 7.24 against the U.S. dollar, compared to its value of 4.91 against the U.S. dollar at the end of July 2018.

CBT's tight monetary policy stance and a significant deterioration in the inflation outlook. Within this time frame, the 2-year and 5-year risk-neutral rates widened by 813 bp and 523 bp, respectively.⁸⁸

While the CBT's tight monetary stance was maintained in the last quarter of 2018 and first quarter of 2019 due to persistent risks to price stability, the Turkish lira experienced a steep decline shortly before the local elections at the end of March 2019. In response to heightened volatility in the financial markets, the CBT took a step of suspending one-week repo auctions on March 25, 2019 and chose to provide the market funding at an overnight lending rate of 25.5%. Concurrently, Turkish banks began to restrict their lira offshore swap market transactions well below their regulatory limits. This action triggered a substantial liquidity squeeze for the Turkish lira in the offshore swap market, pushing the offshore overnight swap rate in London from 22% to over 1200% within two days. It also forced some foreign investors to divest parts of their Turkish assets, including local currency government bonds, to generate the necessary Turkish lira funds to close out their swap positions.⁸⁹ Although the CBT reinstated its one-week repo auctions on April 08, 2019, it halted them again from May 9 to 20, 2019, following the Turkish lira hitting its weakest level after the currency crisis in 2018. Amid higher exchange rate volatility in this period, the risk-neutral rates temporarily spiked up.

Nevertheless, owing to the downward trajectory in unprocessed food inflation and commodity prices, coupled with sluggish domestic demand, the medium-term inflation outlook began to improve as of May 2019. Against this backdrop, the CBT

⁸⁸ The risk-neutral yields for maturities above 3 years hit their sample highs in August 2018, while those less than 3 years reached their sample highs subsequent to the CBT's policy rate hike in September 2018.

⁸⁹ According to official statistics provided by the Republic of Turkey Ministry of Finance and Treasury, the share of domestic debt stock held by foreign investors dropped from 13.35% in February 2019 to 11.64% in March 2019.

decided to lower its one-week repo rate from 24% to 19.75% in its July meeting. With the disinflation process gaining momentum, the CBT proceeded with further rate cuts in September, October, December 2019, and January, February 2020, bringing the one-week repo rate down to 10.75%.⁹⁰ As the coronavirus outbreak started to spread across the world at the end of February 2020, it began to exert immense pressure on the global economic activity, compelling central banks and governments to enact expansionary monetary, credit, and fiscal policy measures.⁹¹ Considering the multifaceted adverse impacts of the pandemic on the Turkish economy, the CBT lowered the one-week repo rate by 250 bp in total in March, April and May 2020 to 8.25%. In addition to the policy rate cuts, the CBT offered banks targeted additional liquidity facilities to ensure a continuous credit flow to the corporate sector. These included the provision of Turkish lira liquidity through repo auctions with maturities up to 91 days at an interest rate 150 bp below the one-week repo rate, as well as Turkish lira currency swap auctions with maturities of six-month and one-year at interest rates 125 bp and 100 bp below the one-week repo rate,

⁹⁰ The adoption of expansionary monetary policies by the central banks of advanced economies due to downside risks to growth outlook and inflation also supported the CBT's move toward implementing a more accommodative monetary policy. The FED lowered the interest rates by 25 bp on July 31, 2019, for the first time since the GFC in 2008. On September 12, 2019, the ECB cut its deposit rate by 10 bp to -0.50% and decided to restart its asset purchase program with a monthly pace of EUR 20 billion as from the beginning of November 2019. With the aim of bolstering economic growth and fostering inflation expectations, the FED continued to reduce the interest rates by 25 bp on both September 18, 2019, and October 30, 2019, ultimately bringing the target range for federal funds rate to 1.5%-1.75%.

⁹¹ Regarding the monetary policy actions implemented by the advanced economies in March 2020, a series of measures were undertaken to mitigate the economic impact of the coronavirus pandemic. During emergency meetings on March 3 and March 15, 2020, the FED lowered the federal funds rate by 50 bp to a range of 1.00%-1.25% and by 100 bp to a range of 0%-0.25%, respectively. At its March 15 meeting, it also unveiled plans to purchase at least USD 500 billion worth of U.S. Treasury securities and at least USD 200 billion worth of agency mortgage-backed securities. Subsequently, on March 23, 2020, the FED pledged unlimited purchase of U.S. Treasuries and agency mortgage-backed securities to support the markets. Amid these developments, on March 16, 2020, the BOJ committed to doubling its annual purchases of Japanese stock exchange traded funds to 12 trillion yen. Moreover, on March 18, 2020, the ECB launched a 750 billion euro emergency bond purchase program. After initially reducing the interest rates from 0.75% to 0.25% during an emergency meeting on March 11, 2020, the BOE further lowered the interest rates to 0.1% and disclosed a front-loaded bond purchase plan of 200 billion pounds on March 19, 2020.

respectively.⁹² Furthermore, to strengthen the effectiveness of the monetary transmission mechanism by enhancing the depth and liquidity of the local currency government bond market, preventing unhealthy price formation, and offering banks greater flexibility in managing their liquidity, the CBT took a series of measures. On March 31, the central bank announced that the outright bond purchase operations under the OMO portfolio could be conducted in a front-loaded manner, with the possibility of adjusting the related purchase limits according to prevailing market conditions. Moreover, for a temporary period, it allowed primary dealer banks to sell their local currency government bonds that they had acquired from the Unemployment Insurance Fund (UIF) to the CBT. Subsequently, on April 17, the CBT decided to double the cap on the OMO portfolio size for 2020 to 10% of its total assets.⁹³ On April 18, 2020, the BRSA also introduced a new and unconventional AR regulation, pushing banks to either expand their loan portfolios, increase holdings of government bonds, or participate in swap transactions with the CBT (Kara, 2021).⁹⁴ Against this backdrop, in April 2020, the substantial decline in global commodity prices, combined with worsening aggregate local demand conditions, had a favorable impact on the domestic inflation outlook, causing risk-neutral rates to hit their lowest levels since September 2016.

⁹² The upper limit of funds that an eligible bank could receive from these new liquidity schemes was linked to the amount of credit that this bank had previously extended or would extend to the corporate sector.

⁹³ Between April 1 and July 27, 2020, the CBT purchased government domestic debt securities (GDDS) amounting to TRY 67.78 billion (nominal equivalent of TRY 57.09 billion), including indirect GDDS purchases of TRY 20.67 billion (nominal equivalent of TRY 12.2 billion) from the UIF in April, which were exempt from the 10% upper limit. The weighted average maturity of the CBT's purchases in April-July period was 3.44 years, while purchases from the UIF in April had a weighted average maturity of 3.09 years, all based on net purchase amounts. Kara (2021) regards these outright bond purchases as the CBT's first QE attempt in decades.

⁹⁴ The AR was calculated according to the following formula:

$$AR = (Loans + Securities * 0.75 + CBT Swaps * 0.5) / (TRY Deposits + FX Deposits * 1.25)$$

The monthly average of the weekly calculated AR was required to be at least 100% for deposit banks and 80% for participation banks.

However, as of May 2020, the strong domestic credit expansion, mainly fueled by the state-owned banks, restrained disinflationary impact of demand conditions. The recovery in global commodity prices and depreciation of the Turkish lira also induced supply-side cost pressures in the domestic economy. These developments affected the medium-term inflation expectations adversely. Despite this, the CBT refrained from conventional policy rate hikes and instead preferred to adopt monetary tightening through liquidity measures to restore the disinflation process. In this respect, the CBT suspended one-week repo auctions as of August 7, 2020, forcing banks to borrow at an overnight lending rate of 9.75%. Additionally, it reduced the liquidity limits offered to primary dealers by half and eventually eliminated these limits, followed by halving the borrowing limits of banks at the CBT Interbank Money Market for overnight transactions. Separately, the traditional repo auctions with variable rates were initiated on August 13, 2020. As a result of these unconventional tightening steps, the CBT's weighted average cost of funding increased to 10.15% by the end of August, in comparison to a low of 7.34% in mid-July. Because of the further weakness in the Turkish lira from the beginning of September, the CBT decided to hike its main policy rate by 200 bp to 10.25% in late September in an effort to contain inflation expectations and address potential risks to the inflation outlook. Nonetheless, the central bank surprisingly kept the policy rate unchanged in October and only raised the late liquidity window lending rate from 13.25% to 14.75%, signaling its intention to employ liquidity measures for additional monetary tightening. Following this decision, the Turkish lira continued to slide, mainly due to ex post negative real interest rates and worries about the central bank's dwindling foreign currency reserves. After a 30% plunge in the Turkish lira's value against the USD from the start of 2020, with the appointment of a new central bank

governor, the CBT announced on November 9 that it would decisively use all policy tools for its primary objective of price stability. Accordingly, the central bank began to implement a strong monetary tightening and increased the one-week repo rate by 475 bp to 15% in November. It also started to provide funding through one-week quantity repo auctions as of November 20. On November 24, the BRSA also decided to terminate its AR policy as of December 31. To mitigate potential risks to the inflation outlook, the CBT raised the one-week repo rate by an additional 200 bp to 17% at its December meeting. Consistent with the deterioration in inflation expectations between April and December 2020, the risk-neutral rates moved higher in this period. Specifically, the 2-year and 5-year risk-neutral rates surged by 428 bp and 266 bp from their April 2020 lows, respectively.

While the CBT kept its main policy rate steady in January and February 2021 after reversing the ultra-loose policy implemented in the early stages of the coronavirus, it adopted a hawkish stance by pledging to uphold a tight monetary policy stance for an extended period until a sustained decline in inflation occurred, and by promising additional front-loaded monetary tightening if necessary. Since the acceleration in retail loan growth, surge in global commodity prices, and weaker Turkish lira as of February posed upside risks to the inflation outlook, the CBT hiked the one-week repo rate by 200 bp to 19% in March. Shortly after the CBT's March meeting, the central bank governor was once again replaced. In April 2021, the CBT kept the main policy rate at 19%, but notably omitted a prior pledge to tighten monetary policy further if required. From May to August 2021, in contrast to soaring inflation and inflation expectations, the CBT kept rates on hold, and unexpectedly cut the policy rate from 19% to 18% in September 2021, citing the slowing effect of previous monetary tightening cycle on credit and domestic demand and attributing

the increase in inflation mainly to transitory supply-side factors. Based on the same reasoning, the CBT continued to cut the policy rate, reducing it to 16% in October 2021. Despite accelerating inflation, a series of record lows in the value of the Turkish lira, and a waning global risk appetite due to the FED's announcement in early November that it would begin tapering its monthly bond purchases as of November, the CBT further lowered the policy rate by 100 bp in both November and December 2021.⁹⁵ In 2021, amid significant deterioration in inflation expectations and rising exchange rate volatility, the risk-neutral rates for different maturities displayed an upward trajectory, reflecting the market participants' expectations of higher future short-term interest rates.

While the Turkish lira exhibited a stable course against major currencies throughout January and into February 2022, on the back of a state-backed foreign currency-protected deposit scheme (abbreviated as KKM in Turkish) introduced on December 20, 2021, and foreign currency interventions, it started to weaken again as of late February, mainly owing to the conflict between Russia and Ukraine, which subsequently led to the sharp increases in global energy, food, and agricultural commodity prices, exacerbating already high domestic inflationary pressures. Although the global monetary tightening process, in the form of rate hikes or reduced asset purchases, accelerated in March 2022 to address the multi-decade high inflation worldwide, the CBT decided to leave its policy rate unchanged at 14% in March.⁹⁶

⁹⁵ The Turkish lira, which closed at 8.64 against the USD on the day preceding the first rate cut in September 2021, traded as high as 18.36 on December 20, 2021, indicating that it lost nearly 53% of its value against the USD during this time frame. The annual inflation surged from 19.58% in September to 36.08% in December, surpassing the CBT's medium-term target of 5% by more than sevenfold.

⁹⁶ Following a surge in inflation to a 30-year high of 5.1% in January 2022, the Bank of Canada increased its benchmark interest rate from 0.25% to 0.50% on March 2, its first hike since October 2018. On March 10, the ECB decided to move faster in reducing its monthly asset purchases as of May and announced that it could end its net asset purchases in the third quarter, depending on the pace of inflation. On March 16, with US inflation hitting a 40-year high at 7.9% in February and considering further upside risks to inflation from the Russia-Ukraine conflict, the FED raised the

Similarly, despite revising its year-end 2022 inflation forecast to 42.8% in April from 23.2% in January, the CBT opted to retain its policy rate at the same level throughout April, May, and June 2022, thereby keeping the real rates in deeply negative territory.⁹⁷ Given the CBT's clear intention to keep its policy rate unchanged in the near term, the risk-neutral rates did not show any significant changes during the January-June 2022 period.

Turning to the evolution of term premium components of yields, as depicted in Figure 10, they were positive and exhibited an upward trend in the early phase of the analysis period, with a notable surge during the peak of the GFC in October 2008, which resulted in term premia for maturities below one year reaching their highest levels during the sample period. In the aftermath of the GFC, the implementation of QE programs by the FED and other central banks in certain developed economies had a pronounced impact on lowering bond yields within those economies. This led to substantial capital inflows into emerging market countries as global investors rebalanced their portfolios towards higher yielding emerging market assets (Chari, Stedman, & Lundblad, 2017). Turkey also benefited from this “search for yield” environment.⁹⁸ During the post-GFC period, the term premia for different yield maturities displayed a downward trend and hit their sample lows in April 2013, before the FED's surprise tapering signal in May 2013. Importantly, Figure 10 clearly illustrates that the term premia in Turkey resided in the negative territory prior to the taper talk, implying that investors were willing to accept lower returns for

target range of federal funds rate by 25 bp to 0.25%-0.5%. On March 17, anticipating a further rise in annual inflation to around 8% in the second quarter, the BOE hiked its key interest rate by 25 bp to 0.75%, marking the third increase since December 2021.

⁹⁷ In March, April, May, and June 2022, Turkey witnessed annual inflation rates of 61.14%, 69.97%, 73.5%, and 78.62%, respectively, each marking the highest levels in two decades.

⁹⁸ To put this in context, according to official statistics provided by the Republic of Turkey Ministry of Finance and Treasury, the share of domestic debt stock held by foreign investors jumped from 8.81% in June 2009 to 25.67% in April 2013.

holding longer-term bonds as compared to holding and rolling over a series of short-term bonds over the same investment horizon.

The taper talk triggered a financial market turmoil in emerging market economies, leading to sharp depreciations in their currencies and substantial rises in their local currency government bond yields, with the worst effects felt by the “fragile five” countries that included Brazil, India, Indonesia, South Africa, and Turkey.⁹⁹ Confronted with heightened uncertainty across various risk factors, investors demanded a higher risk premium in their bond investments during this period. Specifically, between the end of April and August 2013, the term premium components of 2-year and 5-year maturity Turkish Treasury yields surged by 145 bp and 193 bp, respectively.

After a strong upward adjustment in Turkish term premia as of late May 2013, the trend reversed in 2014. This change coincided with measures taken to address the global economic slowdown, particularly in the Euro area and Japan. In response to mounting deflationary and recessionary risks, both the ECB and BOJ pursued aggressive monetary easing policies. In fact, following the introduction of a negative deposit facility rate in June 2014, the ECB decided in September to start purchasing asset-backed securities (ABS) and covered bonds as of October 2014. In October 2014, the BOJ announced an expansion of its quantitative and qualitative monetary easing program by accelerating the purchases of Japanese government bonds to annual pace of 80 trillion yen from 50 trillion yen, and also decided to triple

⁹⁹ In August 2013, Morgan Stanley coined the term “fragile five” for a group of emerging market countries, including Brazil, India, Indonesia, South Africa and Turkey, characterized by significant current account deficits, reliance on foreign capital, and relatively low foreign exchange reserves, rendering them vulnerable to shifts in investor sentiment. Mishra, Moriyama, N’Diaye, and Nguyen (2014) report that, on average, bond yields in these five countries rose by 2.5%, while their exchange rates experienced a depreciation of 13.5% from the onset of the tapering talk on May 22 until the end of August in 2013.

its annual purchases of exchange-traded funds to 3 trillion yen and real estate investment trusts to 90 billion yen. Amid this more accommodative monetary environment, the global risk appetite increased, leading to a recovery in the capital inflows to emerging market economies. In late January 2015, the ECB took another significant step of expanding its asset purchase program to include bonds issued by the Euro area central governments and supranational agencies.¹⁰⁰ Meanwhile, the significant fall in oil and other commodity prices also improved the inflation outlook and current account deficit in Turkey, resulting in a favorable and notable shift in risk assessments regarding the country.¹⁰¹ Against this backdrop, in January 2015, the term premia embedded in the Turkish Treasury bond yields declined to their lowest levels since April 2013.

In the second and third quarter of 2015, the global risk appetite deteriorated, primarily driven by the deceleration across emerging market economies, especially in China, concerns surrounding Greek debt structuring, and lack of clarity regarding the FED's first policy rate hike. Consequently, significant volatility was observed in the global financial markets, and emerging market economies faced substantial capital outflows during this time frame. Similar to other emerging market countries, Turkey experienced a deterioration in risk sentiment, causing term premia to revert to post-tapering era levels at the end of September 2015.

In the last quarter of 2015, owing to persistent economic slowdown in emerging markets, capital flows to these economies continued to remain weak. However, the FED's policy rate hike in December did not result in any material or

¹⁰⁰ The ECB revealed that its new QE program, along with its existing ABS and covered bond purchases, would inject 60 billion euros a month into the eurozone economy from March 2015 until at least the end of September 2016.

¹⁰¹ The price of the Intercontinental Exchange Brent crude oil futures declined from 112.36 USD at the end of June 2014 to 94.67 USD at the end of September 2014, and further dropped to 52.99 USD at the end of January 2015, which marked its lowest month-end level since April 2009.

additional adverse impact on portfolio flows to emerging markets as the post-decision statement and following communications by the FED conveyed a message of gradual monetary normalization in the forthcoming period. In the first quarter of 2016, amid concerns over global growth, the ECB and BOJ unveiled new monetary easing measures.¹⁰² Moreover, in March 2016, given uncertain global economic conditions and absence of inflation pressures, the FED halved its outlook for the U.S. interest rate increases, projecting only two quarter-point hikes by the year-end, down from the initial forecast of four quarter-point hikes back in December 2015. On the back of these developments, capital inflows to emerging market economies started to recover. Subsequently, in the period from September 2015 to June 2016, the term premia for maturities ranging from 2 to 5 years in Turkey decreased by around 90 bp.

Nevertheless, after a fairly stable pattern in the third quarter, they spiked up sharply following the U.S. presidential election in November 2016. Amid growing uncertainties and heightened volatility, driven by the anticipation of the FED adopting an aggressive monetary tightening stance in response to the expected expansionary fiscal policies from the new U.S. administration, the term premia increased in both U.S. and emerging market economies, including Turkey, during November and December 2016. Given the tighter monetary policy measures by the CBT, coupled with mild and predictable monetary policy normalization by the central banks of advanced economies, particularly the FED, the term premia

¹⁰² In January 2016, the BOJ announced that it would apply a negative interest rate of -0.10% on a portion of excess reserves held at the central bank. Subsequently, in March 2016, the ECB decided to lower its deposit facility rate by 0.10% to -0.40%, following the rate cut in December 2015. Alongside this rate adjustment, the ECB decided to expand its asset purchase program, set to commence in April 2016, and include the investment grade euro-denominated corporate bonds in its regular purchases. Additionally, in the same monetary policy meeting, the ECB introduced a new series of four targeted longer-term refinancing operations (T-LTROs) with a four-year maturity, scheduled to start in June 2016.

embedded in the Turkish government local currency nominal bond yields steadily declined in the first three quarters of 2017.

However, starting from October 2017, Turkey began to diverge from its peers, mainly due to elevated geopolitical risks and a worsening inflation outlook. Despite a nine-year high headline inflation figure in October 2017, the CBT abstained from a conventional policy rate hike but instead took unconventional liquidity measures to tighten the monetary policy in November 2017. While providing the market funding entirely through the late liquidity window facility as of November 22, 2017, the central bank raised the late liquidity lending rate by a cumulative 425 bp to 16.5% between December 2017 and May 2018 to curb inflationary risks. Throughout this period, the CBT continued to keep the one-week repo rate constant at 8%. As the final hike in late liquidity window lending rate on May 23 did not succeed in halting the depreciation of Turkish lira, the CBT announced on May 28 that it would finalize the simplification process of the operational framework of its monetary policy and use the one-week repo rate once again as both the main policy rate and primary funding tool, to be effective from June 1, 2018 with an interest rate set at 16.5%. Following the May inflation figures, the CBT opted for further tightening and raised the one-week repo rate to 17.75% in its June meeting. However, in spite of a 14-year high headline inflation figure in June, the central bank unexpectedly left the policy rate on hold, which led to a widening of the sovereign CDS spread, a depreciation of the Turkish lira, and a sharp sell-off in local currency government bonds. The ongoing negative sentiment was amplified by the political tension between Turkey and the U.S. as of late July 2018, escalating into a full-fledged currency crisis in early August 2018 (see footnote 87). Against the backdrop of massive macroeconomic and financial uncertainty, as

displayed in Figure 10, the term premia in Turkish local currency government bond yields surged dramatically, reaching their sample highs for maturities from 2-year to 4-year. During the period from the end of September 2017 to the end of August 2018, the 2-year and 5-year term premiums increased by 653 bp and 791 bp, respectively.

The CBT's strong monetary tightening in September, which involved raising the one-week repo rate from 17.75% to 24%, and diminishing geopolitical risks led to an appreciation of the Turkish lira, a drop in the exchange rate volatility, and a decline in the sovereign CDS spread during the final quarter of 2018 and in January 2019. Amid a favorable shift in risk perceptions, the term premia retraced to levels last observed in October 2017 at the end of January 2019.

However, in early March 2019, a new political rift between Turkey and the U.S. began to exert pressure on the value of the Turkish lira. As the pace of depreciation notably accelerated on March 22, nine days ahead of the nationwide municipal elections, the CBT promptly responded by suspending one-week repo auctions as of March 25 for an indefinite period of to squeeze the Turkish lira liquidity in the market. On March 25, Turkish banks also began to restrict their offshore swap market transactions in the London market well below their regulatory limits, which pushed the overnight offshore swap rate to over 1200% on March 27. Although the massive squeeze on Turkish lira liquidity briefly strengthened the currency, it also forced some foreign investors to sell their Turkish government local currency bonds in order to secure funds needed to unwind their swap positions (see footnote 89). While Turkish banks started to provide Turkish lira liquidity to the offshore market on March 28, resulting in the overnight offshore swap rate falling sharply to 35%, the Turkish lira resumed its depreciation, accompanied by a rise in

yields on Turkish government bonds, on the last two business days of the month. In light of significant uncertainty and volatility witnessed in the financial markets, the term premium components of 2-year and 5-year Turkish government bonds rose by 233 bp and 328 bp, respectively, during March 2019.

Amid internal and external political tensions in both April and May 2019, the Turkish lira continued to experience persistent pressure. Nevertheless, in the rest of the year and January 2020, the exchange rate volatility and sovereign credit risk premium mostly exhibited a downward trend on the back of positive developments in inflation prospects, the current account balance, and global financial conditions.¹⁰³ In this environment, the term premia for maturities beyond 2 years fell back into negative territory in January 2020.

In late February 2020, the spread of coronavirus pandemic triggered a global economic crisis, leading to massive capital outflows from emerging market countries during March.¹⁰⁴ In the midst of heightened risk aversion, similar to its emerging market peers, the Turkish lira also depreciated sharply, coupled with a significant increase in the Turkey's CDS spread, in March. Given the abrupt change in the risk sentiment caused by uncertainty regarding the pandemic's severity, duration and its repercussions on the Turkish economy, the 2-year and 5-year term premiums were higher by 114 bp and 249 bp, respectively, by the end of March as compared to end of January. While the CBT's announcements regarding the purchase of local currency government bonds for its OMO portfolio on March 31 and April 17, along

¹⁰³ August 2019 was an exception to overall trend described above. During this month, escalating trade tensions between China and the U.S. negatively impacted global risk appetite, resulting in portfolio outflows from emerging market economies, including Turkey. Additionally, geopolitical developments further fueled the risk-off sentiment in Turkish financial markets.

¹⁰⁴ According to the Institute of International Finance, global investors pulled a record 83.3 billion USD from emerging market countries in March 2020, surpassing the outflow observed during the GFC. Debt outflows accounted for 31 billion USD of total, marking the second-largest monthly outflow on record since the peak of the GFC in October 2008.

with the BRSA's AR regulation on April 18, helped lower local currency government bond yields presumably and largely through the term premium component in the latter part of April, the Turkish lira remained under pressure and reached 7.00 against the USD on April 28, hitting its lowest level since the height of the currency crisis in August 2018.¹⁰⁵ Despite a recovery in risk sentiment across emerging market countries during May and June 2020, mainly on the back of renewed interest in carry trades and expectations of a swift economic recovery with new stimulus measures, the Turkish lira underperformed most of its peers, mainly due to concerns about the country's diminishing foreign currency reserves and mounting current account deficit. In the third quarter of 2020, the rebound in the domestic economic activity, driven by high loan growth, together with further depreciation of the Turkish lira, deteriorated the medium-term inflation outlook, which in turn kept the term premia at elevated levels.

Although the CBT surprisingly hiked the policy rate by 200 bp to 10.25% in late September 2020, its first hike since September 2018, the central bank unexpectedly left it unchanged in October despite a new spike in the Turkish lira due to geopolitical tensions. Amid these developments, the decline in the term premia following the rate hike in September was reversed in October, primarily because of the uncertainty surrounding the economic and monetary policy outlook. Nevertheless, with the appointment of a new central bank governor, the CBT made a significant policy move by raising interest rates by 475 bp in November, followed by another 200 bp hike in December.¹⁰⁶ These decisive actions bolstered the risk

¹⁰⁵ The Turkish banks increased their interest for local currency government bonds following the introduction of the BRSA's AR regulation. Refer to footnote 94 for the details of the AR formula.

¹⁰⁶ Five days after the CBT's November 2020 rate hike, the BRSA decided to terminate the AR regulation, which had been spurring the credit growth and contributing to adverse inflationary effects as well as financial stability risks since May. The termination was effective as of the end of December 2020, following a series of announced revisions in August, September, and October 2020, all aimed at

appetite for Turkish assets as investors hoped that Turkey would adopt more orthodox policies to address its double-digit inflation problem. Against this backdrop, term premia experienced a rapid compression.

In January and February 2021, the CBT kept the main policy rate unchanged with a hawkish commitment that it would decisively maintain tight monetary conditions for an extended period and deliver an additional monetary tightening if necessary. In line with the promise of sustaining its tight policy, in response to the deterioration in the medium-term inflation outlook after February, the CBT hiked the policy rate by 200 bp to 19% in March 2021. However, two days after the CBT's March meeting, the central bank governor was once again replaced, triggering a drastic sell-off in Turkish assets. Due to uncertainty regarding domestic monetary policy and increased risk aversion, the term premium component in 5-year bond yields rose by 362 bp in March. In its April meeting, the CBT decided to keep the policy rate unchanged, but it abandoned its previous commitment to enact further monetary tightening if needed. Despite a surge in headline inflation and medium-term inflation expectations in the following months, the Turkish lira and sovereign risk premium demonstrated a relatively better performance in July and August 2021, primarily owing to diminishing concerns about an immediate interest rate cut by the CBT, as it communicated that the prevailing tight monetary stance would be maintained until a permanent fall in inflation became evident. Amid this favorable backdrop, the 5-year term premium in August was 157 bp lower compared to its level in March.

However, even though the headline inflation figure reached 19.25% in August, the CBT unexpectedly cut the policy rate by 100 bp to 18% in September

easing the regulation during the process of policy normalization after the peak of the coronavirus pandemic.

2021, effectively concluding the tightening cycle initiated in September 2020. This was followed by a larger-than-expected 200 bp rate cut in October, along with two additional 100 bp cuts in November and December 2021, each causing the Turkish lira to hit record lows against major currencies. While the introduction of a state-backed foreign currency-protected deposit scheme (KKM) on December 20, 2021, partially alleviated the pressure on the Turkish lira, concerns about rising inflation, the central bank's loose monetary policy, and its diminished net foreign currency reserves continued to weigh on market sentiment.¹⁰⁷ In the context of increasing macroeconomic, financial, and monetary policy uncertainties from August to December 2021, the term premia displayed an upward trend, with the 2-year and 5-year term premiums increasing by 228 bp and 485 bp, respectively.

This upward trend further extended to late February 2022 because of the conflict between Russia and Ukraine, adding to the already high uncertainties.¹⁰⁸ After hovering slightly below the February levels in March, the term premia for all maturities declined in April, primarily attributable to the CBT's purchases of TRY-denominated fixed-coupon GDDS for its OMO portfolio from the primary dealer banks via auctions as of April 08, 2022.¹⁰⁹ Notably, this decrease occurred despite the significant increase in the country's sovereign CDS spread amid heightened

¹⁰⁷ After reaching a record low of 18.36 against the USD on December 20, the Turkish lira strengthened sharply and closed the day at 13.50, following the introduction of a state-backed foreign currency-protected deposit scheme. Despite extending its gains and briefly trading at 10.1 against the USD on December 23, it began to depreciate again and ended the year at 13.32.

¹⁰⁸ As illustrated in Figure 10, the 5-year term premium reached its maximum value in February 2022, slightly surpassing the level observed during the currency crisis in August 2018.

¹⁰⁹ Following the acquisition of a net amount of TRY 0.85 billion worth of fixed-coupon GDDS through the traditional auction method between April 1 and April 6, the CBT expanded its fixed-coupon GDDS holdings by purchasing a net amount of TRY 3.35 billion worth of bonds via the auctions conducted under the primary dealership outright purchase facility during the April 8-29 period, with a weighted average maturity of 4.82 years. This increase, aimed at supporting the primary dealership system and enhancing the effectiveness of the monetary transmission mechanism, marked a significant rise compared to its previous net purchases of TRY 1.01 billion, TRY 0.27 billion, and TRY 0.64 billion worth of fixed-coupon GDDS through the traditional auction method in January, February, and March 2022, respectively.

geopolitical tensions and expectations of a more rapid global monetary tightening. In May, even as Turkey's CDS spread widened further compared to April and the global risk appetite worsened, the term premiums for maturities exceeding 3-year continued to decline with the support of the CBT's increased bond purchases from the primary dealer banks through auctions, as well as the changes in collateral requirements for banks at the CBT under the "Liraization Strategy", which caused local banks to increase their demand for TRY-denominated fixed-coupon GDDS.¹¹⁰ In June 2022, as a complementary measure to enhance the weighting of the TRY-denominated fixed-coupon GDDS in the collateral pool within the scope of the liraization strategy, the CBT mandated banks to hold additional Turkish lira long-term fixed-coupon GDDS for foreign currency deposits. Accordingly, banks were required to maintain holdings between 3% and 10% of such deposits in TRY-denominated fixed-coupon GDDS with an original maturity of at least 5 years and a residual maturity of at least 4 years.¹¹¹ This regulation triggered increased demand for fixed-coupon GDDS, particularly at maturities beyond 4 years, leading to a fall in term premia across the yield curve.

¹¹⁰ In May 2022, the CBT purchased a net amount of TRY 5.72 billion worth of fixed-coupon GDDS via auctions under the primary dealership outright purchase facility, with a weighted average maturity of 4.99 years. On May 16, as an integral part of the liraization strategy, the CBT restricted the range of assets eligible for inclusion in the GDDS basket as collateral and increased its weighting in the collateral system for currency swap transactions with the CBT to a minimum of 30%, effective from May 27, 2022. The revised GDDS basket would include only TRY-denominated zero-coupon, fixed-coupon, floating-rate, and TLREF-indexed GDDS and lease certificates issued domestically by the Ministry of Treasury and Finance and the Asset Leasing Company of the Turkish Treasury. On the same date, the collateral discount rate applied to consumer price index (CPI)-linked, gold-backed, and foreign currency-denominated GDDS, as well as foreign currency-denominated and gold-backed assets, was increased from 5% to 15%. This discount rate was further increased to 30% on June 6, 2022.

¹¹¹ The regulation was publicly announced on June 10, detailed in a document distributed to banks on June 13, and officially became effective as of June 24, 2022. In accordance with this regulation, non-fixed coupon bonds, including CPI-indexed bonds, could not be held as collateral for foreign currency deposits. In the period from June 10 to June 24, 2022, Turkish banks' holdings of TRY-denominated securities as collateral increased from TRY 289.86 billion to TRY 360.45 billion, according to the BRSA weekly banking sector data.

1.5 Determinants of term premium and risk-neutral components

Equipped with the estimates of term premium and risk-neutral components of the Turkish government nominal bond yields, this section investigates their determinants through a series of regressions.

1.5.1 Sample sub-periods

I conduct the analysis for the full sample period as well as separately for two distinct sub-periods to assess whether there were changes in the significance or impact of explanatory variables over time. I split the sample in November 2017 as Turkey started to adopt an unorthodox economic and financial policy mix, marked by various episodes. These include periods of unconventional liquidity measures implemented by the CBT to tighten the monetary policy (such as market funding solely through late liquidity window between November 2017 and May 2018, and reduction of liquidity limits to primary dealers within the OMO framework to zero in August 2020), the liquidity squeeze in the offshore swap market for Turkish lira in March 2019, the unconventional bond purchases by the CBT between April and July 2020 and post-April 2022, the implementation of the AR regulation by the BRSA between April and December 2020, the monetary easing cycle between September and December 2021 despite a surge in domestic inflation when some other central banks in advanced and emerging market economies were in the monetary tightening phase in response to escalating inflation, the launch of a state-backed foreign currency-protected deposit scheme (KKM) in December 2021, and changes in collateral requirements for banks under the liraization strategy in May and June 2022.

1.5.2 Selection of explanatory variables

In this sub-section, I choose the explanatory variables to be utilized in the regression analyses of both term premium and risk-neutral components, based on theoretical and empirical considerations.

1.5.2.1 Explanatory variables for term premium

To account for the variation in the term premium component of bond yields, I consider a set of key factors discussed in the literature section. It is important to recall that, in the traditional two-part decomposition of bond yields, the term premium component implicitly incorporates credit risk premium, currency risk premium, liquidity risk premium, in addition to inflation risk premium. In light of this understanding, to represent the sovereign credit risk premium embedded in the term premium, I utilize the 5-year sovereign CDS spread of Turkey, recognized as the most liquid CDS contract across all maturities.¹¹² To proxy the currency risk premium inherent in the term premium, the implied volatility of the 6-month ATM USD/TRY option contract is employed.¹¹³ To measure the liquidity risk premium, I use a modified version of the noise (illiquidity) measure proposed by Hu et al. (2013).¹¹⁴ In contrast to their traditional approach, which assesses the noise in the

¹¹² As stated in footnote 20, the credit spread encompasses both the expected loss from credit events and the associated risk premium. This risk premium serves to compensate investors for the possibility that realized loss might surpass the expected loss. Given that risk premia often constitute a larger portion of emerging market credit spreads, exceeding the risk-neutral expected losses (Remolona, Scatigna, & Wu, 2007), I do not separately identify these components, and simply refer to their sum as credit risk premium.

¹¹³ For the utilization of implied volatility from ATM option contracts as a proxy for currency risk premium, see Gadanecz et al. (2018) and Özbek and Talaslı (2020).

¹¹⁴ As previously discussed in Section 1.3.6, Hu et al. (2013) argue that during market crises, the limited availability of arbitrage capital forces investors to scale back or entirely abandon their relative value trades. Consequently, yields fluctuate more independently in their respective habitats, leading to significant price deviations from the fundamental values. To assess the illiquidity of the overall market, they introduce a noise measure that aggregates cross-sectional pricing errors in the yield space. This measure is formulated by computing the root mean squared distance between actual market yields ($y_t^{i,actual}$) and model-implied yields ($y_t^{i,model}$):

yield space, I adapt the objective function value obtained during the derivation of the NSS model parameters in the construction of the zero-coupon yield curve.

Specifically, to capture illiquidity dynamics, I define the illiquidity measure by dividing the objective function by the number of bonds available for curve fitting at time t , denoted as N_t , and then taking the square root of this quantity:

$$\text{Illiquidity}_t = \sqrt{\frac{1}{N_t} \min_{(\beta_0, \beta_1, \beta_2, \beta_3, \tau_1, \tau_2)} \sum_{i=1}^{N_t} (P_t^{i,actual} - P_t^{i,model})^2 \frac{1}{\sqrt{D_t^i}}}$$

where $P_t^{i,actual}$ and $P_t^{i,model}$ represent the actual and model-implied prices of bond i at time t , and D_t^i is the duration of bond i .¹¹⁵ Compared to standard approach of measuring noise in the yield space, this method aggregates the deviations in the price space and thereby avoids the computational burden associated with numerically converting model-implied prices to model-implied yields. Turning to inflation risk premium embedded in the term premium, I use two different proxies to represent it. The first proxy is the inflation surprise measure, calculated as difference between the actual monthly CPI data released by the Turkish Statistical Institute and the (modified) mean expectation from the CBT's "Survey of Market Participants", scaled by the previous month's annual CPI rate.¹¹⁶ This scaling facilitates an assessment of the inflation surprise's magnitude relative to the previous month's

$$\text{Noise}_t = \sqrt{\frac{1}{N_t} \sum_{i=1}^{N_t} (y_t^{i,actual} - y_t^{i,model})^2}$$

where N_t is the number of bonds available for curve fitting at time t .

¹¹⁵ Hu et al. (2013) note that, apart from gauging noise in the yield space, they also explore it in the price space by employing squared pricing errors weighted by the inverse duration. They report that their main results are robust to both methods of noise measurement (see their footnote 13).

¹¹⁶ Gülşen and Kara (2019) demonstrate that most of the inflation uncertainty measures considered in their study are closely associated with the unscaled inflation surprise metric.

annual inflation rate, allowing for comparability across different time periods.¹¹⁷

Additionally, given that the CBT's surveys are typically conducted around the middle of a month, and inflation statistics for that specific month are generally released on the 3rd of the following month (see footnote 85), I employ a one-month lagged value of the inflation surprise measure in examining its impact on local term premium component.¹¹⁸ The second proxy to represent inflation risk premium is the exchange rate surprise measure. It is calculated as the difference between actual end-of-month USD/TRY exchange rate and the (modified mean) expectation among survey participants, scaled by previous month's exchange rate. Since the exchange rate is one of the key determinants of the inflation in Turkey, operating not only through the direct cost channel but also through the interaction between the inflation and inflation expectations (see Koç, Ögünç, & Özmen, 2021), an exchange rate surprise could potentially impact inflation uncertainty among economic agents during the subsequent month. This underscores the relevance of considering exchange rate surprise measure in the analysis of local term premia, particularly with a one-month lag structure, to better capture its influence on inflation uncertainty.

To investigate international monetary spillover effects from the major advanced economies, particularly the United States, on the term premia embedded in the Turkish Treasury bond yields, I include both the U.S. term premium and risk-neutral components as explanatory variables.¹¹⁹ Importantly, this enables a separate evaluation of the impact of each U.S. yield component on local term premia.

¹¹⁷ For instance, a 1% monthly inflation surprise in a period of 10% annual inflation holds more significance in the eyes of market participants, as opposed to the same 1% monthly inflation surprise in a period of 20% annual inflation.

¹¹⁸ From August 2001 to December 2012, the CBT's survey was conducted in the first and third week of each month. Since January 2013, the survey has been conducted monthly. For the period between February 2005 and Dec 2012, I use the results of the second survey conducted each month.

¹¹⁹ To decompose the U.S. nominal interest rates into term premium and risk-neutral components, I employ the ACM model. The NSS model parameters for the period from February 2005 to June 2022

In this study, the impacts of the CBT's unconventional bond purchases, the BRSA's AR regulation, and the CBT's liraization strategy on local term premia are also explored by including time dummy variables in the regression analysis. As previously outlined, over the course of my sample period, the CBT conducted unconventional bond purchases for its OMO portfolio on two distinct episodes, specifically between April and July 2020 and between April and June 2022. For the time dummy specification of unconventional bond purchases, I assign a value of 0 for July 2020 since the CBT's purchases in that month constituted less than 9% of the total unconventional bond purchases in 2020. Similarly, while the AR regulation was in effect from April to December 2020, nearly all local banks promptly complied with the rule by the end of June 2020.¹²⁰ Consequently, in specifying the time dummy for the AR regulation, I assign a value of 0 to the period from the July to December 2020. With regard to the time dummy for the liraization strategy, it is assigned a value of 1 for May and June 2022. Nevertheless, simultaneous inclusion of these three time dummies in a regression model gives rise to multicollinearity issues due to their temporal overlap, resulting in imprecise estimates of individual effects. To address this problem, I construct a single composite time dummy that captures the potential joint impact of the unconventional bond purchases, the AR regulation, and the liraization strategy, referring to it as the "Aggregate Policy" dummy.¹²¹ Alternatively, to strike a balance between preventing the multicollinearity

are sourced from the FED's website. In addition, in the regression analysis, the maturities of the U.S. term premia and risk-neutral rates are matched to those of the Turkish term premia.

¹²⁰ Only two banks, constituting approximately 2% of the total unconsolidated assets of banks under the AR regulation, faced administrative fines from the BRSA on July 29, 2020, for failing to meet the required AR ratio (see footnote 94 for the AR formula). This indicates that the majority of banks complied with the regulation by the end of June 2020.

¹²¹ The "Aggregate Policy" dummy, while also technically equivalent to time dummy for unconventional bond purchases, is specifically designed to emphasize the combined impact of the unconventional bond purchases, the AR regulation, and the liraization strategy whenever any of them is in effect. Therefore, the "Aggregate Policy" time dummy takes the value of 1 when any of the corresponding time dummies representing these policies is 1, and 0 otherwise.

issues and capturing the individual impact of each policy, I combine the unconventional bond purchases dummy and the AR regulation dummy into a single aggregate dummy variable. This combined time dummy variable is then included in the regression analysis along with the time dummy representing the liraization strategy.

Additionally, acknowledging Turkey's status as a net oil importer and its notable dependence on global energy markets, I incorporate the Brent crude oil prices into the analysis as an explanatory factor.¹²² The rationale behind this inclusion lies in the fact that they have a multifaceted impact on the Turkish economy, affecting economic growth, trade balance, and inflation. This, in turn, prompts an investigation of whether the Brent oil prices also contribute to the evolution of the term premium component in Turkey. Lastly, as a proxy for global risk sentiment, I include the VIX index in the explanatory variables to evaluate its influence on local term premia dynamics.

1.5.2.2 Explanatory variables for risk-neutral rate

For the regression of risk-neutral rate, I employ the same set of explanatory variables utilized in the regression of term premium. This decision is grounded not only in previous empirical evidence and standard economic theory that selected variables could also contribute to the variation of risk-neutral rate, but also in the recognition that maintaining consistency in the choice of explanatory variables ensures a thorough exploration of the shared drivers impacting both yield components.

For instance, Albagli et al. (2019) demonstrate that U.S. monetary policy shocks, identified as changes in the U.S. 2-year Treasury yields in their study,

¹²² In this study, I regard the Brent crude oil price as an indicator of oil prices, given the widespread use of this type of oil in Turkey.

significantly impact both term premium and risk-neutral components of 10-year bond yields in 12 emerging market economies in the post-GFC period. The effect is more pronounced on the term premium component, explaining around three-fourths of the overall spillover effect.¹²³ In a separate event study, Curcuru, Kamin, Li, and Rodriguez (2023) document that, during the post-GFC period, the U.S. 10-year yields' term premium and risk-neutral components account for a significant portion of the variation in 10-year government bond yields in Korea, Brazil, and Mexico. Furthermore, for all these three emerging market countries, the influence of U.S. risk-neutral rates significantly outweighs that of U.S. term premiums. Therefore, these findings collectively support the inclusion of both U.S. yield components as explanatory variables in the analysis of the risk-neutral rate in Turkey.

As discussed in the literature section, unconventional bond purchases by central banks can effectively reduce nominal bond yields not only through the term premium component (via the portfolio balance and liquidity channels) but also through the risk-neutral component (via the signaling channel).¹²⁴ Furthermore, the impact of the AR regulation on loan growth (see footnote 94) may potentially shape the expected path of future short-term interest rates. Thus, time dummies utilized in the regression of term premium are also deemed relevant for the regression of risk-neutral rate.

Finally, the dynamics of sovereign credit risk premium, exchange rate volatility, domestic macroeconomic surprises, domestic financial market stress,

¹²³ For the same period, they find that the spillover effects of U.S. monetary policy shocks on 12 developed economies are concentrated more on the risk-neutral component, accounting for more than half of the total impact.

¹²⁴ Christensen and Rudebusch (2012) argue that, within the framework of central bank bond purchase programs, the varying contributions of these channels in explaining changes in bond yields may be contingent on the institutional structure of financial markets and/or the communication policies of central banks.

international oil prices, and global risk perceptions can lead market participants to reassess and revise their expected path of future short-term interest rates. Hence, the inclusion of previously used proxy variables representing these factors is warranted for the regression of the risk-neutral rate.

1.5.3 Unit root tests for stationarity

The visual inspections of term premium and risk-neutral components (dependent variables) in Figure 10 and explanatory variables in Figure 11 underscore the necessity of conducting unit root tests. This precautionary step is essential to avoid the risk of spurious regressions, ensuring reliability and robustness of the subsequent empirical analyses. Consequently, I carry out an augmented Dickey-Fuller (ADF) test for each dependent and independent variable, excluding time dummies.

Table 13 presents the ADF unit root tests results with constant as well as constant and trend specifications. The findings reveal that, at the 5% level of statistical significance, the null hypothesis of a unit root cannot be rejected for any of the variables in their level form, except for the exchange rate and inflation surprise measures. On the other hand, when the ADF tests are conducted on the first differences, the null hypothesis of a unit root is strongly rejected for all of the variables, indicating that the variables which are non-stationary in their level forms become stationary in their first-difference forms.

1.5.4 Baseline model specification

Given the findings from the ADF unit root tests, to avoid the spurious regressions problem, my baseline time-series regression model employs the first differencing for all variables, with the exception of the exchange rate and inflation surprise measures,

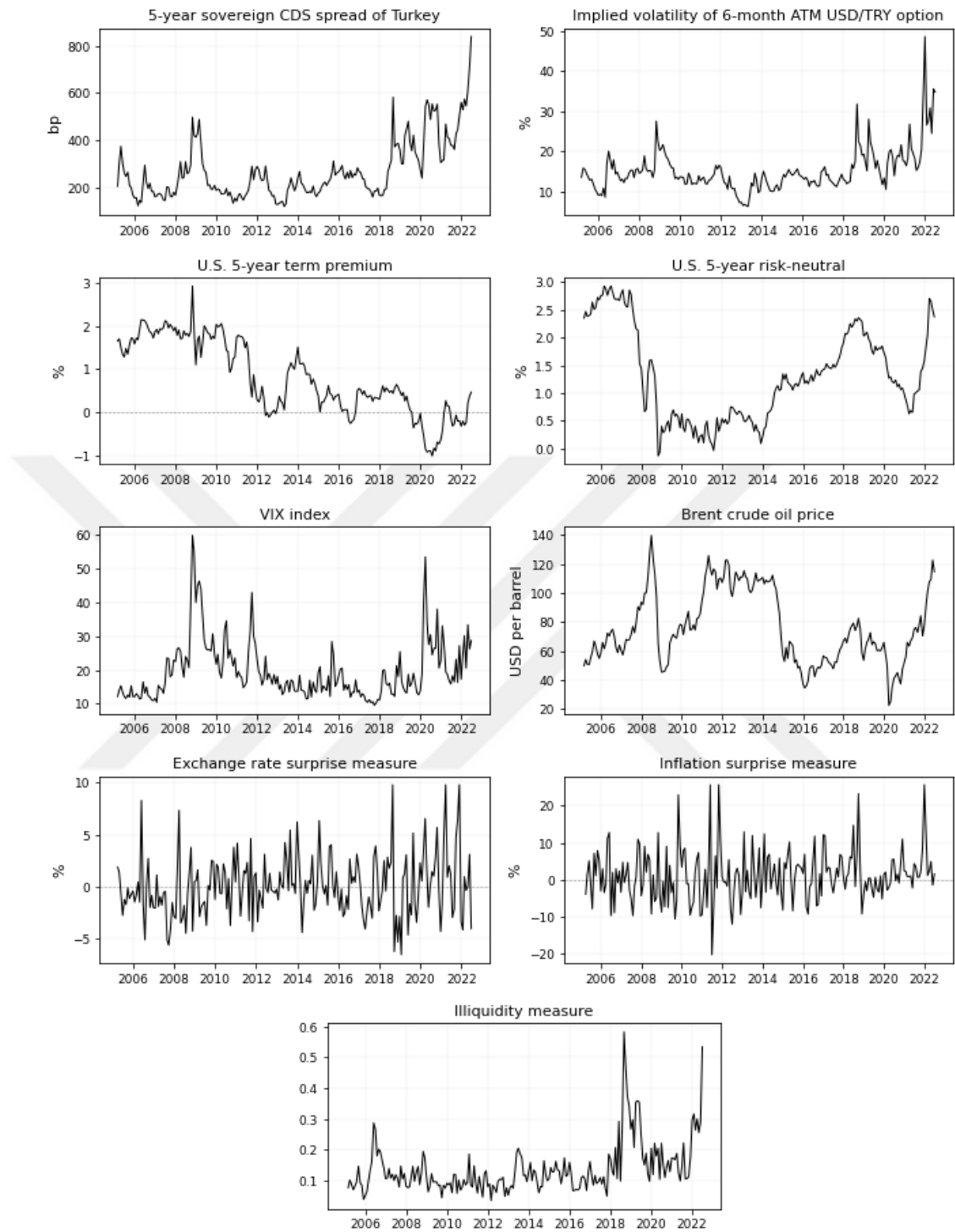


Figure 11. Explanatory variables for term premium and risk-neutral components

Note: To mitigate the potential impact of outliers on statistical analyses, I apply winsorization to both the exchange rate and inflation surprise measures, replacing the highest two values in their respective upper tails with the next highest values.

Table 13. ADF Unit Root Test Results

Variable	At Level		At First-Difference	
	Constant, No Trend	Constant & Trend	Constant, No Trend	Constant & Trend
Turkey 5-year Term Premium	-1.273 (0.642)	-1.228 (0.902)	-14.826 (0.000)	-14.800 (0.000)
Turkey 5-year Risk-Neutral	-1.515 (0.524)	-1.192 (0.909)	-4.819 (0.000)	-4.879 (0.001)
Turkey 5-year CDS Spread	-1.247 (0.654)	-2.331 (0.415)	-14.814 (0.000)	-14.921 (0.000)
Imp. Vol. 6m ATM USD/TRY	-0.401 (0.905)	-1.059 (0.932)	-17.493 (0.000)	-17.503 (0.000)
Illiquidity Measure	-1.704 (0.428)	-2.483 (0.336)	-17.637 (0.000)	-17.633 (0.000)
U.S. 5-year Term Premium	-1.710 (0.425)	-2.755 (0.216)	-14.318 (0.000)	-14.290 (0.000)
U.S. 5-year Risk-Neutral Rate	-1.855 (0.353)	-1.540 (0.813)	-7.416 (0.000)	-12.155 (0.000)
VIX Index	-2.663 (0.082)	-2.659 (0.255)	-16.257 (0.000)	-16.218 (0.000)
Brent Crude Oil	-2.619 (0.091)	-2.631 (0.267)	-10.707 (0.000)	-10.683 (0.000)
Exch. Rate Surprise Measure	-11.902 (0.000)	-12.155 (0.000)	-20.663 (0.000)	-20.611 (0.000)
Inflation Surprise Measure	-12.880 (0.000)	-12.962 (0.000)	-20.959 (0.000)	-20.907 (0.000)

The optimal lag length for the ADF tests is based on the modified Akaike information criterion (MAIC), as suggested by Ng and Perron (2001). The MacKinnon approximate p-values are reported in parentheses below the computed ADF statistics.

as well as the “Aggregate Policy” time dummy.¹²⁵ The baseline model is specified as follows:

$$\begin{aligned}
\Delta BYC_t^i = & \alpha_i + \beta_1^i \Delta CDS_t + \beta_2^i \Delta ATM6m_t + \beta_3^i \Delta ILLIQ_t + \beta_4^i \Delta US_{TP_t}^i + \beta_5^i \Delta US_{RN_t}^i \\
& + \beta_6^i \Delta VIX_t + \beta_7^i \Delta Brent_t + \beta_8^i USDTRY_Surprise_{t-1} \\
& + \beta_9^i CPI_Surprise_{t-1} + \beta_{10}^i Agg_Pol_t + \varepsilon_t^i
\end{aligned}$$

where Δ represents the first-difference operator, the dependent variable BYC_t^i is the bond yield component for maturity i , which is either the term premium or the risk-

¹²⁵ Choosing not to use first differences for the two surprise measures is grounded in their observed stationarity. This decision is theoretically motivated by the intention to avoid over-differencing, which could lead to the loss of information and inefficiencies in parameter estimates (Patra & Ray, 2010). Furthermore, employing them in their level forms aligns with my objective of capturing and examining their shock effects on changes in bond yield components. Notably, Camba-Méndez and Werner (2017) also make a similar choice, opting to keep the data surprise measure in levels rather than resorting to its first differences in their study.

neutral rate specific to that maturity in the Turkish local currency government bond market, CDS is the 5-year sovereign CDS spread of Turkey, $ATM6m$ is the implied volatility of the 6-month ATM USD/TRY option contract, $ILLIQ$ is the illiquidity measure for the local bond market, US_TP^i is the U.S. term premium for maturity i , US_RN^i is the U.S. risk-neutral rate for maturity i , VIX is the VIX index, $Brent$ is the Brent crude oil price, $USDTRY_Surprise$ is the exchange rate surprise measure, $CPI_Surprise$ is the inflation surprise measure, and Agg_Pol is the “Aggregate Policy” time dummy.

Acknowledging the potential heterogeneity in the impact of explanatory variables across different segments of the term structure, I conduct separate time series regressions for maturities spanning from 2-year to 5-year (denoted as i in the baseline model).¹²⁶ This approach allows for a more detailed examination of the relationship between changes in the yield components and their respective drivers at various maturity points along the yield curve.

In an effort to strike a balance between preventing multicollinearity issues arising from the temporal overlap of the unconventional bond purchases, the AR regulation and the liraization strategy, and capturing the individual impact of each of these policies, I also employ an alternative specification to the baseline regression model. Instead of utilizing the composite “Aggregated Policy” time dummy, I first aggregate the unconventional bond purchases dummy and the AR regulation dummy,

¹²⁶ In the Turkish local currency government bond market, 2-year and 5-year maturity bonds often serve as benchmark securities for short-term and medium-term interest rates, respectively. They are also commonly used as reference rates for pricing other fixed-income securities. Given the notable roles played by the 2-year and 5-year maturities, focusing on the 2-5 year segment is particularly important for a comprehensive understanding of the dynamics of yield components within this critical part of the term structure. In addition, similar to the 5-year maturity, the ADF unit root tests indicate that the 2-year, 3-year, and 4-year maturity term premium and risk-neutral components of both Turkish and U.S. nominal bond yields exhibit stationarity after first differencing.

and then incorporate this combined dummy along with the time dummy representing the liraization strategy into the model.

1.5.5 Baseline regression results

In this sub-section, I first report the estimation results of the baseline regression model for changes in the term premium and risk-neutral components across maturities ranging from 2-year to 5-year over the full sample period and the two sub-periods. Following this, I present the estimation results of the alternative specification for changes in both yield components across the same maturities and time periods.

1.5.5.1 Term premium regressions

Table 14 displays the baseline regression results for changes in the term premium component for the selected maturities within the specified time periods. The findings reveal a strong and statistically significant positive association between changes in the term premia and changes in the sovereign CDS spread. This relationship is significant not only for the entire sample period (February 2005 to June 2022) but also within the two sub-periods (pre- and post-November 2017), underscoring the key role of changes in the CDS spread in the dynamics of term premia changes (at the 1% level). A closer examination of the coefficients across different maturities also suggests that the impact of changes in the CDS spread becomes more pronounced for longer maturities. For instance, a 100 bp increase in the CDS spread corresponds to an 82 bp increase in the 2-year term premium and a 109 bp increase in the 5-year term premium over the full sample period. In the first and the second

Table 14. Baseline Regression Results for the Term Premium

Explanatory Variable	Full Sample Period (February 2005 – June 2022)				First Sub-Period (February 2005 – October 2017)				Second Sub-Period (November 2017 – June 2022)			
	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year
Δ CDS Spread	0.0082*** (5.430)	0.0109*** (6.285)	0.0116*** (6.798)	0.0109*** (6.621)	0.0066*** (2.844)	0.0087*** (3.225)	0.0098*** (4.019)	0.0099*** (4.568)	0.0092*** (5.702)	0.0122*** (6.866)	0.0131*** (6.479)	0.0125*** (5.530)
Δ Implied Vol.	0.0453* (1.680)	0.0519* (1.939)	0.0524** (2.042)	0.0483** (2.167)	0.0951** (2.487)	0.1014** (2.237)	0.0815* (1.921)	0.0292 (0.802)	0.0305 (1.576)	0.0349 (1.666)	0.0372 (1.439)	0.0399 (1.411)
Δ Illiquidity M.	0.7223 (0.931)	1.2121 (1.377)	1.5059 (1.552)	1.6786 (1.636)	-0.3810 (-0.429)	0.0265 (0.023)	1.2721 (1.027)	3.0528** (2.317)	0.9801 (1.218)	1.2133 (1.113)	1.0548 (0.803)	0.6455 (0.456)
Δ US Term Prem.	1.1421** (2.010)	1.1345*** (2.637)	0.8524*** (2.835)	0.5147** (2.529)	1.4176*** (3.335)	1.3692*** (4.004)	1.0425*** (4.258)	0.6745*** (3.761)	-0.7272 (-0.748)	-0.5759 (-0.685)	-0.4888 (-0.648)	-0.3431 (-0.465)
Δ US Risk-Neutral	0.3281* (1.767)	0.2041 (0.971)	0.0744 (0.343)	0.0430 (0.186)	0.3876* (1.941)	0.2588 (1.162)	0.1335 (0.640)	0.1315 (0.633)	-0.3006 (-0.767)	-0.3633 (-0.645)	-0.3407 (-0.466)	-0.2194 (-0.267)
Δ VIX	0.0005 (0.082)	-0.0049 (-0.696)	-0.0078 (-0.999)	-0.0080 (-0.840)	-0.0024 (-0.295)	-0.0037 (-0.397)	-0.0007 (-0.075)	0.0086 (0.809)	0.0047 (0.424)	-0.0036 (-0.249)	-0.0132 (-0.764)	-0.0226 (-1.108)
Δ Brent	0.0202*** (2.838)	0.0179** (2.436)	0.0150** (2.162)	0.0136** (2.027)	0.0110 (1.562)	0.0092 (1.173)	0.0077 (1.054)	0.0073 (1.038)	0.0554*** (3.759)	0.0543*** (2.798)	0.0444* (2.011)	0.0327 (1.457)
\$/TRY Surprise M.	0.0038 (0.309)	-0.0047 (-0.260)	-0.0097 (-0.461)	-0.0122 (-0.551)	0.0014 (0.089)	-0.0091 (-0.391)	-0.0121 (-0.461)	-0.0096 (-0.365)	0.0142 (0.607)	0.0063 (0.216)	-0.0049 (-0.148)	-0.0177 (-0.515)
CPI Surprise M.	0.0034 (0.682)	-0.0002 (-0.037)	-0.0015 (-0.240)	0.0004 (0.076)	-0.0012 (-0.240)	-0.0040 (-0.648)	-0.0034 (-0.574)	0.0011 (0.208)	0.0143 (0.941)	0.0090 (0.445)	0.0058 (0.260)	0.0076 (0.357)
Aggregate Policy	-0.8376** (-2.283)	-1.1922** (-2.149)	-1.4539* (-1.954)	-1.6050* (-1.807)					-0.8163*** (-3.018)	-1.2619** (-2.665)	-1.5463** (-2.246)	-1.6897* (-1.947)
Constant	-0.0027 (-0.099)	0.0054 (0.181)	0.0121 (0.431)	0.0157 (0.541)	-0.0009 (-0.036)	0.0009 (0.033)	0.0021 (0.077)	0.0025 (0.086)	-0.0607 (-0.815)	-0.0309 (-0.385)	0.0073 (0.090)	0.0369 (0.462)
Observations	207	207	207	207	151	151	151	151	56	56	56	56
Adjusted R ²	0.5839	0.6201	0.6382	0.5993	0.5324	0.5538	0.5899	0.5324	0.6767	0.6924	0.6668	0.6287

Note: The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period. Asterisks ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively. In the regression analysis, the exchange rate and inflation surprise measures lag one month.

sub-periods, these changes amount to 66 bp and 92 bp for the 2-year term premium, and 99 bp and 125 bp for the 5-year term premium, respectively.

Shifting the focus to the implied volatility of the 6-month ATM option, the results demonstrate a positive relationship between changes in the term premia and changes in the implied volatility.¹²⁷ For the full sample period and the first sub-period, almost all coefficients are statistically significant at the 5% or 10% levels, highlighting the importance of changes in the perceived exchange rate risk in influencing changes in the term premia. Specifically, a 1% increase in the implied volatility is associated with around 5 bp and 9 bp increases in the term premia for the full sample and pre-November 2017 periods, respectively. However, in the second sub-period, the statistical significance diminishes, suggesting potential variations in the relationship between currency risk premia and bond term premia over different time periods.

The examination of the coefficients associated with the illiquidity measure suggests that while the strength of the positive relationship between changes in the term premia and changes in the illiquidity measure increases with maturity for the full sample period, statistical significance is not achieved for any maturity at conventional levels.¹²⁸ In comparison of the two sub-periods, the impact of changes in the illiquidity measure attains statistical significance at the 5% level only for the 5-year maturity in the first sub-period, but fails to demonstrate any significance in the second sub-period.

With regard to international spillover impacts, the results indicate that changes in the U.S. term premia have a statistically significant impact on changes in

¹²⁷ This is consistent with Gadanecz et al. (2018), who find that higher exchange rate volatility leads investors to demand higher risk premia for local currency debt in emerging markets.

¹²⁸ It is noteworthy that the coefficient in the regression of changes in the 5-year local term premia exhibits marginal significance at the 10.4% level during the full sample period.

the Turkish term premia for the entire sample period (at the 1% and 5% levels). Importantly, in all regressions where the maturities of the U.S. and the Turkish term premia are matched (see footnote 119), the magnitude of the impact decreases with the maturity of the term premium. For example, a 100 bp shock in the U.S. 2-year term premium results in a 114 bp increase in the Turkish 2-year term premium, while the same shock in the U.S. 5-year term premium leads to a 51 bp increase in the Turkish 5-year term premium. Additionally, the key role of the U.S. term premia as a source of spillover effects to the Turkish term premia is only relevant in the pre-November 2017 period, with related coefficients greater and more significant than those in the full sample period. In the post-November 2017 period, characterized by Turkey's adoption of unorthodox economic and financial policy mix, this spillover channel ceases to persist, indicating a notable shift in the dynamics of the relationship between the U.S. and the Turkish term premia. Turning to the spillover effects from the U.S. risk-neutral rates to local term premia, the regression results show that these effects have significant impact only for the 2-year maturity during the full sample period and the pre-November 2017 period, albeit at the 10% significance level.

The baseline regression results in Table 14 also suggest that changes in the VIX index have no statistically significant effect on changes in local term premia throughout the entire sample period and the two sub-periods.¹²⁹ As for the impact of changes in the Brent crude oil prices, the corresponding coefficients are positive and statistically significant for all maturities in the full sample period. Furthermore, these coefficients indicate that the size and significance of the impact decrease with

¹²⁹ The CDS spread is likely to be influenced by changes in the VIX index (see Hibbert & Pavlova, 2017; Kocsis & Monostori, 2016; Longstaff, Pan, Pedersen, & Singleton, 2011; Pan & Singleton, 2008). By controlling for changes in the CDS spread in the baseline regression model, I aim to isolate and examine the direct impact of changes in the VIX index on changes in local term premia.

increasing maturity. When comparing the two sub-periods, it becomes apparent that the heightened responsiveness of local term premia to oil price fluctuations is particularly evident in the post-November 2017 period, a time marked by Turkey's experience with high and volatile inflation during the implementation of unconventional economic and financial policies. Another notable finding is that the exchange rate and inflation surprise measures, incorporated in the regression analysis as proxies of inflation uncertainty, do not produce statistically significant effects on changes in local term premia over the full sample period and the two sub-periods. While the study by Koç et al. (2021) document that exchange rate fluctuations and inflation surprises play very important roles in the formation of inflation expectations in Turkey, my analysis reveals that the employed inflation uncertainty indicators have no impact on changes in the term premium component of nominal bond yields for the considered maturities and time periods.

Focusing on the coefficients of the “Aggregate Policy” time dummy, which reflect the joint impact of the CBT's unconventional bond purchases, the BRSA's AR regulation, and the CBT's liraization strategy on changes in local term premia, the regression results show a negative and statistically significant effect. The decline in the term premia, resulting from the unconventional bond purchases by the CBT for its OMO portfolio, along with the bond purchases by local banks under the scope of the AR regulation and within the context of the liraization strategy, offers compelling evidence for the operation of the portfolio balance channel, which in principle works through scarcity and duration risk effects, as previously discussed in the literature section.¹³⁰ Importantly, the response of changes in the term premia to the “Aggregate

¹³⁰ Although the loan growth aspect of the AR regulation could add to upward pressure on term premia due to inflation uncertainties, the overall net impact, considering the bond purchases by the CBT and local banks, is negative, as evidenced by the regression results. Moreover, such government bond purchases could potentially reduce term premia through the liquidity channel. However,

Policy” becomes more pronounced for longer maturities, providing empirical support to the predictions implied by the duration risk effects of these specific bond purchases.¹³¹ Specifically, when the “Aggregate Policy” is in effect, the 2-year term premium declines by 82 bp (significant at 1% level), while the 5-year term premium experiences a more substantial drop of 169 bp (borderline significant at 5% level) in the second sub-period.

Overall, the adjusted R^2 values of the regressions for different maturities range from 58% to 64% in the full sample period, 53% to 59% in the pre-November 2017 period, and 63% to 69% in the post-November 2017 period, indicating that the baseline model has a strong explanatory power on changes in local term premia.

1.5.5.2 Risk-neutral rate regressions

Table 15 reports the baseline regression results for changes in the risk-neutral rate component. The coefficients on changes in the sovereign CDS spread are positive and statistically significant at the 5% level in the full-sample period. This implies that positive changes in the CDS spread prompt the market participants to revise their average expected level of future short-term interest rates upward, with the impact diminishing as the tenor of the risk-neutral rate increases. Specifically, a 100 bp increase in the CDS spread results in 36 bp and 26 bp increases in the 2-year and 5-year risk-neutral rates, respectively, over the entire sample period. Moreover, a

regressing changes in the illiquidity measure on the “Aggregate Policy” time dummy (with or without control variables) does not result in a statistically significant or negative coefficient. This suggests that the liquidity channel as a separate transmission mechanism of these specific bond purchases may not be operative in this context.

¹³¹ Such bond purchases reduce the aggregate duration risk available to bond investors, leading to an increase in the risk bearing-capacity of arbitrageurs and, therefore, a decline in the market price of duration risk demanded by arbitrageurs (see footnote 29). As the duration of the targeted bond increases, for a given reduction in the market price of duration risk, the corresponding term premium component falls more. For more details, see Vayanos and Vila (2009) and Altavilla, Carboni, and Motto (2015).

Table 15. Baseline Regression Results for the Risk-Neutral Rate

Explanatory Variable	Full Sample Period (February 2005 – June 2022)				First Sub-Period (February 2005 – October 2017)				Second Sub-Period (November 2017 – June 2022)			
	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year
Δ CDS Spread	0.0036** (2.095)	0.0033** (2.232)	0.0029** (2.289)	0.0026** (2.317)	0.0068*** (3.605)	0.0060*** (3.645)	0.0053*** (3.671)	0.0047*** (3.698)	0.0026 (1.227)	0.0025 (1.363)	0.0023 (1.431)	0.0021 (1.477)
Δ Implied Vol.	0.1078*** (4.844)	0.0920*** (4.916)	0.0799*** (4.921)	0.0701*** (4.911)	0.0730*** (2.722)	0.0619** (2.610)	0.0533** (2.555)	0.0461** (2.500)	0.1218*** (5.281)	0.1034*** (5.241)	0.0896*** (5.302)	0.0785*** (5.386)
Δ Illiquidity M.	1.1561* (1.845)	1.0917** (2.050)	0.9782** (2.110)	0.8680** (2.128)	2.2242** (2.092)	2.1115** (2.282)	1.9246** (2.372)	1.7353** (2.420)	0.8088 (0.930)	0.7328 (1.000)	0.6196 (0.982)	0.5353 (0.968)
Δ US Term Prem.	0.8879** (2.476)	0.4368* (1.879)	0.2615 (1.604)	0.1841 (1.488)	0.8119** (2.075)	0.4350 (1.646)	0.2830 (1.515)	0.2138 (1.502)	0.8363 (0.844)	0.1414 (0.202)	-0.0390 (-0.082)	-0.0783 (-0.229)
Δ US Risk-Neutral	-0.0232 (-0.143)	-0.1106 (-0.765)	-0.1303 (-0.938)	-0.1265 (-0.919)	0.1443 (0.816)	0.0637 (0.420)	0.0463 (0.327)	0.0529 (0.381)	-0.7009 (-1.231)	-0.7827 (-1.439)	-0.8155 (-1.482)	-0.8417 (-1.503)
Δ VIX	0.0115 (1.086)	0.0101 (1.100)	0.0089 (1.113)	0.0079 (1.124)	0.0008 (0.075)	0.0018 (0.179)	0.0020 (0.224)	0.0020 (0.254)	0.0240 (1.242)	0.0198 (1.214)	0.0168 (1.200)	0.0145 (1.186)
Δ Brent	0.0295*** (4.316)	0.0253*** (4.249)	0.0221*** (4.230)	0.0194*** (4.229)	0.0155*** (2.763)	0.0133*** (2.617)	0.0116** (2.563)	0.0101** (2.537)	0.0698*** (4.117)	0.0604*** (4.256)	0.0529*** (4.340)	0.0465*** (4.367)
$\$/TRY$ Surprise M.	0.0584*** (2.979)	0.0492*** (2.989)	0.0422*** (2.982)	0.0366*** (2.971)	0.0698*** (3.873)	0.0594*** (3.919)	0.0513*** (3.951)	0.0448*** (3.966)	0.0537 (1.420)	0.0440 (1.379)	0.0372 (1.361)	0.0321 (1.358)
CPI Surprise M.	0.0241*** (4.283)	0.0205*** (4.289)	0.0177*** (4.287)	0.0154*** (4.276)	0.0193*** (3.369)	0.0166*** (3.355)	0.0144*** (3.351)	0.0125*** (3.347)	0.0342 (1.638)	0.0291 (1.641)	0.0250 (1.631)	0.0216 (1.613)
Aggregate Policy	-0.6939* (-1.869)	-0.6081* (-1.894)	-0.5376* (-1.900)	-0.4781* (-1.901)					-0.7660* (-1.713)	-0.6647* (-1.847)	-0.5916* (-1.932)	-0.5277* (-1.965)
Constant	-0.0243 (-0.612)	-0.0208 (-0.605)	-0.0180 (-0.601)	-0.0156 (-0.595)	-0.0072 (-0.190)	-0.0062 (-0.189)	-0.0053 (-0.187)	-0.0045 (-0.180)	-0.0752 (-0.594)	-0.0634 (-0.587)	-0.0538 (-0.576)	-0.0462 (-0.563)
Observations	207	207	207	207	151	151	151	151	56	56	56	56
Adjusted R ²	0.5736	0.5784	0.5803	0.5812	0.55	0.5515	0.5522	0.5531	0.6256	0.635	0.6403	0.6431

Note: The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period. Asterisks ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively. In the regression analysis, the exchange rate and inflation surprise measures lag one month.

comparison of the two sub-periods reveals that the positive relationship between changes in the sovereign credit risk premium and changes in risk-neutral rates is statistically significant at the 1% level in the first sub-period, but its significance weakens during the implementation of unorthodox policies in the second sub-period, with p-values above conventional levels.

Turning to the impact of changes in the implied volatility of the 6-month ATM option, the corresponding coefficients imply that there is a statistically significant positive association between changes in risk-neutral rates and changes in the implied volatility for the full sample period and both sub-periods. This finding is consistent with the strong exchange rate pass-through effect on domestic inflation in Turkey in the sense that changes in uncertainty about the future path of exchange rate induce revisions in inflation expectations, which in turn lead to revisions in the expected path of future short-term interest rates. The response of changes in the risk-neutral rates to changes in the implied volatility also decreases with maturity across all sample periods. For instance, a 1% increase in the implied volatility increases the 2-year and 5-year risk-neutral rates by 11 bp and 7 bp in the full-sample period, 7 bp and 5 bp in the first sub-sample period, and 12 bp and 8 bp in the second sub-period. Notably, the impact magnitudes are more substantial in the post-November 2017 period.

Interestingly, changes in the illiquidity measure in the baseline regression exhibit a positive and statistically significant relationship with changes in the risk-neutral rates for the full sample period and the first sub-period. Hu et al. (2013) note that their illiquidity measure tends to spike during market crises, leading them to assert that this indicator not only captures the liquidity conditions specific to the bond market but also reflects the overall market conditions. As a policy response to

market crises in emerging markets, the central banks in those economies typically resort to monetary tightening measures due to concerns about the impact of capital outflows on financial stability. In light of this, it is intuitive that when the illiquidity measure surges during periods of market turmoil, investors may preemptively anticipate a tightening of monetary conditions in the short to medium term, which consequently results in an increase in the average expected level of future short-term interest rates. Additionally, the regression results highlight that the corresponding impact is more pronounced for shorter maturities. While this pattern is still observed in the post-November 2017 period, the statistical significance is not supported at conventional levels, suggesting that changes in the illiquidity indicator do not contribute to the temporal dynamics of changes in the risk-neutral rates during the second sub-period.

In the context of spillover effects from the U.S. bond yields to local currency government bond yields in Turkey, the regression results indicate that changes in the U.S. term premia also influence local bond yields through their impact on local risk-neutral rates, particularly at the short-end of the yield curve (at the 5% significance level), over the entire sample period and the pre-November 2017 period. For instance, a 100 bp increase in the U.S. 2-year term premium leads to an 89 bp and 81 bp increase in the Turkish 2-year risk-neutral rate in the full sample and first sub-sample period, respectively. However, in the second sub-period, changes in the U.S. term premia do not exhibit any significant impact on changes in the Turkish risk-neutral rates. This mirrors their effects on changes in local term premia during the same period. As for the spillover effects from the U.S. risk-neutral rates to local risk-neutral rates, there is no statistical evidence supporting the relevance of this channel in any of the sample periods examined.

Table 15 also shows that there is no statistically significant relationship between changes in the VIX index and changes in the Turkish risk-neutral rates across all maturities and sample periods. On the other hand, all coefficients on changes in the Brent crude oil prices are positive and statistically significant (at or near the 1% level) in the analysis of local risk-neutral rates. This aligns with previous empirical findings that fluctuations in oil prices have a well-established impact on agents' inflation expectations in Turkey. Given the intrinsic relationship between the inflation expectations and the expected path of future short-term interest rates, it is evident that the Brent oil prices influence local risk-neutral rates potentially through the inflation expectations channel. Consistent with economic intuition, the regression results also suggest that the impact of oil price shocks on changes in the risk-neutral rates diminishes with increasing maturity.

Moving to the effects of surprise measures in the baseline model, the results show that inflation surprises have a positive and statistically significant impact on changes in local risk-neutral rates at the 1% level in the full sample period and the first sub-period, and near 10% level in the second sub-period. This suggests that an unexpected upward or downward change in domestic inflation in the previous month leads investors to adjust their average expected level of future short-term interest rates in the same direction for the consequent month. Similar to inflation surprises, exchange rate surprises also have a positive and statistically significant relationship with changes in local risk-neutral rates at the 1% level, but only in the full-sample period and the first sub-period.¹³²

Regarding the impact of the “Aggregate Policy” time dummy on changes in local risk-neutral rates, the associated coefficients are negative and significant at the

¹³² Although the coefficients of the exchange rate surprise measure in the second sub-period are also positive, the corresponding p-values range between 0.16 and 0.18.

10% level. Although the impact of the AR regulation on loan growth may potentially exert an upward pressure on local risk-neutral rates, the negative coefficients suggest that the signaling channel of the CBT's unconventional bond purchases outweighs this effect. Moreover, the more pronounced downward effect on shorter maturities compared to longer maturities aligns with the operation of the signaling channel.¹³³ Specifically, when the “Aggregate Policy” is in effect, the 2-year and 5-year risk-neutral rates decline by 77 bp and 53 bp, respectively, in the second sub-period. Comparing the responses of local term premia in Table 14 with those of local risk-neutral rates in Table 15 reveals that the impact of the “Aggregate Policy” on nominal yield changes is more evident through local term premia than through local risk-neutral rates. Furthermore, the impact difference increases with maturity.

In terms of goodness of fit, the adjusted R^2 values of the baseline model are between 57% and 58% in the full sample period, around 55% in the pre-November 2017 period, and between 63% to 64% in the post-November 2017 period. These values indicate that the selected set of variables explains a considerable portion of the changes in local risk-neutral rates.

1.5.5.3 Alternative specification results

In the alternative specification, the composite “Aggregate Policy” dummy is replaced with two other time dummies: the first one combining the unconventional bond purchases and the AR regulation time dummies, and the second one exclusively representing the CBT's liraization strategy. While the task of disentangling the individual effects of these policies is complicated by multicollinearity issues

¹³³ The rationale behind this is that the central bank's commitment to keep the policy rate lower is likely to be more limited in the distant future (Altavilla et al., 2015).

stemming from their temporal overlap, this refined alternative approach facilitates a more concentrated analysis specifically in distinguishing and evaluating the impact of the liraization strategy on yield components. Additionally, since the VIX variable is not significant in the baseline regressions of both term premium and risk-neutral rate components, it is excluded from the alternative specification regressions.

Table 16 represents the estimation results for changes in the term premium component. The quantitative and qualitative findings are broadly similar to those observed in the baseline regressions in Table 14, albeit with some differences. Notably, in the second sub-period, the coefficients on changes in the implied volatility of the 6-month ATM option (except for the 2-year maturity) become statistically significant at either 5% or 10% level, primarily attributable to the omission of the VIX index. The other essential difference is that changes in the illiquidity measure appear to contribute significantly to the temporal dynamics of changes in local term premia at the 10% level over the full sample period, specifically for medium-term maturities.

Regarding the impacts of specified policies on term premia, the regression results demonstrate that the combined influence of the CBT's unconventional bond purchases and local banks' bond purchases under the AR regulation is statistically significant in lowering local term premia for the 2-year and 3-year maturities at the 5% level and for the 4-year maturity at the 10% level over the second sub-period. This is consistent with the fact that although there is a lack of detailed information on the maturity profile of banks' government bond purchases under the AR regulation, the CBT's unconventional bond purchases in the second sub-period are mostly concentrated on bonds with 2- to 4-year maturities. As for the liraization strategy, it has a statistically significant impact on reducing local term premia for the 3-year

Table 16. Alternative Specification Results for the Term Premium

Explanatory Variable	Full Sample Period (February 2005 – June 2022)				First Sub-Period (February 2005 – October 2017)				Second Sub-Period (November 2017 – June 2022)			
	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year
Δ CDS Spread	0.0083*** (5.434)	0.0107*** (6.401)	0.0113*** (7.031)	0.0106*** (7.022)	0.0065*** (2.736)	0.0086*** (3.243)	0.0097*** (4.185)	0.0102*** (4.993)	0.0094*** (5.719)	0.0122*** (6.683)	0.0129*** (6.245)	0.0121*** (5.167)
Δ Implied Vol.	0.0466 (1.629)	0.0563** (1.993)	0.0590* (2.218)	0.0556** (2.395)	0.0937** (2.537)	0.0992** (2.191)	0.0811* (1.892)	0.0346 (0.958)	0.0300 (1.468)	0.0392* (1.771)	0.0468* (1.869)	0.0539** (2.040)
Δ Illiquidity M.	0.8247 (1.143)	1.3866 (1.635)	1.7570* (1.814)	1.9778* (1.920)	-0.3769 (-0.425)	0.0322 (0.028)	1.2732 (1.034)	3.0402** (2.339)	1.0538 (1.376)	1.4168 (1.328)	1.3375 (0.987)	0.9489 (0.641)
Δ US Term Prem.	1.1887** (2.112)	1.1230*** (2.680)	0.8243*** (2.780)	0.4886** (2.334)	1.4192*** (3.358)	1.3749*** (4.013)	1.0436*** (4.269)	0.6606*** (3.706)	-0.5050 (-0.460)	-0.4617 (-0.519)	-0.4948 (-0.602)	-0.3894 (-0.462)
Δ US Risk-Neutral	0.3357* (1.788)	0.2128 (0.986)	0.0886 (0.402)	0.0595 (0.259)	0.3926* (1.950)	0.2676 (1.189)	0.1352 (0.653)	0.1069 (0.499)	-0.3001 (-0.767)	-0.3780 (-0.668)	-0.3486 (-0.477)	-0.1736 (-0.216)
Δ Brent	0.0200*** (2.936)	0.0191*** (2.734)	0.0171** (2.556)	0.0158** (2.488)	0.0113* (1.763)	0.0096 (1.303)	0.0077 (1.094)	0.0062 (0.917)	0.0532*** (3.681)	0.0554*** (3.347)	0.0498*** (2.709)	0.0418** (2.266)
$\$/TRY$ Surprise M.	0.0032 (0.257)	-0.0043 (-0.233)	-0.0089 (-0.418)	-0.0116 (-0.521)	0.0016 (0.103)	-0.0088 (-0.373)	-0.0120 (-0.454)	-0.0104 (-0.387)	0.0126 (0.571)	0.0070 (0.256)	-0.0018 (-0.054)	-0.0116 (-0.323)
CPI Surprise M.	0.0036 (0.711)	0.0003 (0.042)	-0.0008 (-0.120)	0.0013 (0.225)	-0.0010 (-0.202)	-0.0036 (-0.602)	-0.0033 (-0.563)	0.0002 (0.044)	0.0153 (1.048)	0.0095 (0.493)	0.0058 (0.267)	0.0065 (0.315)
UBP & AR Reg.	-0.6504* (-1.667)	-0.8648 (-1.527)	-0.9807 (-1.366)	-1.0469 (-1.231)					-0.7498** (-2.556)	-0.9877** (-2.192)	-1.0834* (-1.761)	-1.1098 (-1.399)
Liraization Strategy	-0.5979 (-1.010)	-0.9547 (-1.604)	-1.3695** (-1.985)	-1.6304** (-2.010)					-0.3291 (-0.668)	-0.8621* (-1.796)	-1.3252** (-2.177)	-1.5807** (-2.028)
Constant	-0.0028 (-0.103)	0.0033 (0.111)	0.0089 (0.308)	0.0122 (0.414)	-0.0010 (-0.042)	0.0007 (0.027)	0.0020 (0.076)	0.0030 (0.104)	-0.0563 (-0.774)	-0.0368 (-0.461)	-0.0103 (-0.127)	0.0098 (0.126)
Observations	207	207	207	207	151	151	151	151	56	56	56	56
Adjusted R ²	0.5884	0.6267	0.6503	0.6179	0.5354	0.5565	0.5927	0.5334	0.6786	0.7045	0.6884	0.6549

Note: The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period. Asterisks ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively. In the regression analysis, the exchange rate and inflation surprise measures lag one month. UBP stands for the CBT's "unconventional bond purchases".

maturity at the 10% level, and for the 4-year and 5-year maturities at the 5% level in the second sub-period. This aligns with the CBT's June 2022 mandate, requiring banks to maintain holdings between 3% and 10% of foreign currency deposits in TRY-denominated fixed-coupon GDDS with an original maturity of at least 5 years and a residual maturity of at least 4 years. This, in turn, prompted banks to increase their demand for such securities. Along with these results, examining the pattern of regression coefficients for both employed time dummies reveals that the response of changes in local term premia to corresponding policies becomes more pronounced for longer maturities, providing further support for the duration risk channel.

Turning to the alternative specification results for local risk-neutral rates, the coefficients and their significances in Table 17 are fairly similar to those presented in the baseline regressions in Table 15, though with some exceptions. A noteworthy difference is that, in the second sub-period, the coefficients on changes in the sovereign CDS spread (except for the 2-year maturity) attain statistical significance at the 10% level, driven mainly by the exclusion of the VIX index. Concurrently, the inflation surprise measure also achieves statistical significance at the 10% level during the same sub-period for the same reason.

The alternative specification results in Table 17 also highlight that, in the post-November 2017 period, the coefficients representing the joint impact of the CBT's unconventional bond purchases and local banks' bond purchases under the AR regulation on changes in the risk-neutral rates are negative and exhibit a decline in magnitude with maturity. Nevertheless, the statistical significance of these coefficients is weaker, with p-values ranging from 0.12 to 0.14, thus suggesting a marginal role of the signaling channel in lowering local bond yields in the second sub-period. Lastly, the regression results reveal that there is no statistically

Table 17. Alternative Specification Results for the Risk-Neutral

Explanatory Variable	Full Sample Period (February 2005 – June 2022)				First Sub-Period (February 2005 – October 2017)				Second Sub-Period (November 2017 – June 2022)			
	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year	2-year	3-year	4-year	5-year
Δ CDS Spread	0.0041*** (2.851)	0.0038*** (3.017)	0.0034*** (3.082)	0.0030*** (3.113)	0.0068*** (3.834)	0.0061*** (3.945)	0.0054*** (4.004)	0.0048*** (4.050)	0.0033 (1.599)	0.0031* (1.760)	0.0028* (1.832)	0.0025* (1.871)
Δ Implied Vol.	0.1032*** (5.047)	0.0879*** (5.133)	0.0763*** (5.117)	0.0670*** (5.083)	0.0734*** (2.732)	0.0629** (2.594)	0.0545** (2.526)	0.0473** (2.469)	0.1123*** (5.416)	0.0953*** (5.359)	0.0829*** (5.362)	0.0728*** (5.401)
Δ Illiquidity M.	1.1393* (1.734)	1.0814* (1.937)	0.9733** (2.004)	0.8666** (2.027)	2.2228** (2.109)	2.1087** (2.299)	1.9216** (2.390)	1.7324** (2.438)	0.8604 (0.956)	0.7915 (1.043)	0.6835 (1.041)	0.5960 (1.037)
Δ US Term Prem.	0.9523** (2.518)	0.4707* (1.960)	0.2809* (1.677)	0.1964 (1.548)	0.8114** (2.081)	0.4322 (1.647)	0.2798 (1.509)	0.2106 (1.490)	1.0626 (0.978)	0.2973 (0.418)	0.0635 (0.134)	-0.0090 (-0.027)
Δ US Risk-Neutral	-0.0420 (-0.257)	-0.1363 (-0.886)	-0.1590 (-1.048)	-0.1565 (-1.034)	0.1426 (0.818)	0.0595 (0.392)	0.0411 (0.287)	0.0472 (0.334)	-0.7899 (-1.308)	-0.8813 (-1.480)	-0.9270 (-1.525)	-0.9612 (-1.550)
Δ Brent	0.0269*** (4.732)	0.0230*** (4.687)	0.0200*** (4.661)	0.0176*** (4.651)	0.0154*** (3.094)	0.0131*** (2.922)	0.0114*** (2.859)	0.0099*** (2.829)	0.0604*** (4.196)	0.0525*** (4.417)	0.0462*** (4.549)	0.0407*** (4.635)
\$/TRY Surprise M.	0.0560*** (2.791)	0.0471*** (2.791)	0.0403*** (2.774)	0.0350*** (2.754)	0.0697*** (3.801)	0.0592*** (3.863)	0.0511*** (3.902)	0.0446*** (3.921)	0.0453 (1.152)	0.0374 (1.124)	0.0315 (1.104)	0.0270 (1.095)
CPI Surprise M.	0.0236*** (4.132)	0.0201*** (4.136)	0.0173*** (4.135)	0.0150*** (4.127)	0.0193*** (3.409)	0.0164*** (3.427)	0.0142*** (3.442)	0.0123*** (3.448)	0.0380* (1.864)	0.0320* (1.834)	0.0275* (1.814)	0.0238* (1.792)
UBP & AR Reg.	-0.7673 (-1.419)	-0.6690 (-1.422)	-0.5849 (-1.421)	-0.5149 (-1.418)					-0.8966 (-1.507)	-0.7755 (-1.536)	-0.6776 (-1.564)	-0.5973 (-1.580)
Liraization Strategy	0.0779 (0.127)	0.0643 (0.124)	0.0389 (0.086)	0.0194 (0.049)					0.0932 (0.153)	0.0933 (0.186)	0.0586 (0.135)	0.0376 (0.098)
Constant	-0.0215 (-0.546)	-0.0183 (-0.540)	-0.0158 (-0.537)	-0.0137 (-0.531)	-0.0072 (-0.191)	-0.0061 (-0.189)	-0.0052 (-0.185)	-0.0044 (-0.178)	-0.0537 (-0.416)	-0.0452 (-0.412)	-0.0386 (-0.406)	-0.0331 (-0.399)
Observations	207	207	207	207	151	151	151	151	56	56	56	56
Adjusted R ²	0.5692	0.5739	0.5756	0.5764	0.5531	0.5545	0.5551	0.5559	0.6103	0.6210	0.6268	0.6299

Note: The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period. Asterisks ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively. In the regression analysis, the exchange rate and inflation surprise measures lag one month. UBP stands for “unconventional bond purchases”.

significant relationship between the liraization strategy and changes in the risk-neutral rates, consistent with theoretical predictions.

1.5.6 Robustness checks

In this sub-section, I perform robustness checks on the empirical results by introducing new explanatory variables, employing alternative measures for some of the existing explanatory variables, and conducting the Instrumental Variables Generalized Method of Moments (IV-GMM) regressions to control for simultaneity and reverse causality issues.

As a first robustness check, considering the positive correlations between the U.S. and the Euro area bond yield components, along with the potential spillover effects from the Euro area bond yields to local bond yields, I add changes in the Euro area term premia and risk-neutral components as explanatory variables to baseline regressions to mitigate the omitted variable bias concerns.¹³⁴ The results suggest that while the impacts of changes in the U.S. yield components on changes in local yield components remain broadly consistent with the baseline specification results, there is no evidence of a statistically significant positive association between changes in the Euro area yield components and changes in local yield components.

In the second robustness check, I incorporate the MOVE index, which measures the implied volatility of the U.S. Treasury market over the next 30 days and serves as a proxy for uncertainty in the global bond markets, into the baseline model. The earlier findings still hold, along with no statistically significant impact of changes in the MOVE index on changes in local yield components.

¹³⁴ To decompose the Euro area nominal interest rates into term premium and risk-neutral components, I employ the ACM model. The NSS model parameters for the period from February 2005 to June 2022 are obtained from the ECB's website.

In the third robustness check, I replace the exchange rate and inflation surprise measures with the inflation forecast dispersion measure from the CBT's "Survey of Market Participants". This measure can be used as a proxy for inflation uncertainty and is calculated as the cross-sectional standard deviation of 12-month-ahead inflation expectations, scaled by the (modified) mean of these expectations.¹³⁵ The regression results across all maturities and sample periods reveal that there is no statistically significant contribution of changes in the inflation forecast dispersion measure to changes in local term premia or risk-neutral rates.¹³⁶ Nonetheless, the conclusions derived from other explanatory variables in the baseline regressions remain largely unchanged.

As an additional check for the robustness of the empirical results, the sovereign CDS spread of Turkey is replaced with the J.P. Morgan Emerging Markets Bond Index Plus (EMBI+) spread for Turkey, providing an alternative measure of sovereign credit risk.¹³⁷ The results exhibit a high degree of similarity with the earlier findings.

I also investigate the potential endogeneity of the CDS spread due to concerns about simultaneity and reverse causality with local term premia and risk-neutral rates. To address these issues, I use the two-step IV-GMM estimation, employing changes in both the Markit CDX North America High Yield index spread (CDX.NA.HY index spread) and the VIX index as instruments for changes in the

¹³⁵ Carvalho and Minella (2012) employ this measure as a proxy for inflation uncertainty in Brazil.

¹³⁶ The ADF unit root tests results show that the inflation forecast dispersion measure becomes stationary after first differencing. Therefore, its first-difference form is employed in the time-series regressions.

¹³⁷ The J.P. Morgan EMBI+ tracks the aggregate performance of U.S. dollar-denominated sovereign bonds issued by a diverse range of emerging market countries and serves as a key benchmark for this debt market. The EMBI+ spread for a specific country measures the additional yield that investors require for holding that country's external debt compared to the yields of equivalent U.S. Treasury securities. This spread is used as an indicator of the credit risk associated with investing in that country's sovereign bonds.

Turkish CDS spread.¹³⁸ The Kleibergen-Paap rk LM statistic indicates the relevance of the instruments, while the Cragg-Donald and Kleibergen-Paap rk Wald F statistics demonstrate their strength. The Hansen J statistic confirms that the instruments are valid. However, the C-test fails to reject the null hypothesis that the CDS spread is exogenous.¹³⁹ These results remain largely consistent across the maturities and sample periods. Given the lack of evidence for endogeneity in the CDS spread, the OLS results can be considered consistent and unbiased.

Similarly, I examine the potential endogeneity of the 6-month ATM USD/TRY option volatility, considering the possibility of simultaneity and reverse causality with components of local bond yields. I repeat the two-step IV-GMM estimation, using changes in both the J.P. Morgan emerging markets currency volatility index (EM-VXY index) and the VIX index as instruments for changes in the 6-month ATM option volatility.¹⁴⁰ Although the test results confirm that the

¹³⁸ The CDX.NA.HY index is a credit default swap index composed of 100 North American high-yield issuers that are actively traded in the CDS market. Its spread serves as a gauge of market sentiment for entities with broadly similar credit quality, such as emerging market economies. In their panel analysis of emerging market sovereign CDS spreads, including those of Turkey, Fender, Hayo, and Neuenkirch (2012) identify this index as one of the most influential drivers of those spreads. In this study, I use the most liquid 5-year CDX.NA.HY index contract. Additionally, the VIX index, a proxy for global risk aversion, is recognized as a significant driver of emerging market sovereign CDS spreads (see Hibbert & Pavlova, 2017; Kocsis & Monostori, 2016; Pan & Singleton, 2008). Consequently, changes in both the CDX.NA.HY index spread and the VIX index are employed as instruments for changes in the Turkish CDS spread and are therefore treated as excluded exogenous variables in the IV-GMM model. Alternatively, I also consider changes in the aggregate J.P. Morgan EMBI+ spread in combination with changes in the VIX index as instruments for changes in the Turkish CDS spread.

¹³⁹ For instance, for the regression of changes in the 5-year local term premium over the full sample period, using changes in both the CDX.NA.HY index spread and the VIX index as instruments, the Kleibergen-Paap rk LM statistic has a p-value of 0.001, showing that the instruments are not underidentified. The Cragg-Donald F statistic is 22.813, exceeding the Stock and Yogo (2005) 10% critical value of 19.93, indicating that the instruments are sufficiently strong. The Kleibergen-Paap rk Wald F statistic is 12.715, surpassing the Staiger and Stock (1997) threshold of 10, further confirming the strength of the instruments. The Hansen J statistic has a p-value of 0.802, affirming that the instruments are not correlated with the residuals and, therefore, are valid. The C-test yields a p-value of 0.343, suggesting that the CDS spread is exogenous in the regression. Furthermore, using the set of changes in both the aggregate J.P. Morgan EMBI+ spread and the VIX index as instruments in the alternative IV-GMM specification produces qualitatively similar results, reinforcing the robustness of the findings.

¹⁴⁰ The J.P. Morgan EM-VXY index tracks the implied volatility of the 3-month ATM options on the exchange rate of emerging market currencies against the U.S. dollar. At the end of the sample period, the index includes 12 emerging market currencies, with TRY weighted at 5.7%.

instruments are relevant, sufficiently strong, and valid, the C-test does not reject the null hypothesis of exogeneity of the 6-month ATM option volatility. These findings hold broadly across the maturities and sample periods. Therefore, the OLS results are validated as consistent and unbiased.

1.6 Summary and conclusion

This chapter starts with the construction of end-of-month zero-coupon yield curves for Turkish government nominal bonds over the period from February 2005 to June 2022. Following that, I employ the ACM model to decompose zero-coupon yields into their risk-neutral rate and term premium components. Equipped with these components, I then proceed to investigate the key determinants of each component separately across the full sample period and two sub-periods (pre- and post-November 2017).

This study enriches the existing literature in several dimensions. In the calibration of the NSS model, I employ the objective function that incorporates Gimeno and Nave's (2006) weighting scheme, which strikes a much better balance between the short-maturity and the long-maturity bonds in the optimization process. Additionally, drawing inspiration from Manousopoulos and Michalopoulos (2009), who use various non-linear optimization algorithms to fit the NSS model, and from Wahlstrøm et al. (2022), who explore diverse parameter initialization approaches for the same purpose, I integrate both of these methods into the curve fitting procedure. In this context, I utilize four alternative non-linear optimization algorithms and three different parameter initialization approaches in the estimation of Turkish zero-coupon yield curves. Taken together, this comprehensive and integrated curve fitting methodology mitigates premature convergence to local optima and significantly

enhances in-sample accuracy in the estimation process, thus offering a valuable tool for policymakers, academics, and financial market practitioners.

Another contribution to the literature is that the model specification test results support the use of a five-factor specification within the ACM model for Turkey. The Wald test statistics suggest that the five-factor model is likely to provide a better fit for excess bond returns. Moreover, the in-sample RMSE statistics reveal that the five-factor model results in the smallest yield fitting errors. Importantly, additional results shed light on the significant role of the fourth and fifth PCs of yields in explaining the dynamics of bond risk premia in Turkey. Another noteworthy outcome is that the five-factor specification produces very reasonable maximal Sharpe ratios, with a peak at 3.27 and an average at 0.88 (below 1). This effectively mitigates overfitting concerns associated with the five-factor ATSMs.

To the best of my knowledge, this study is the first report to comprehensively investigate the determinants of both nominal yield components in the context of Turkey. The analysis involves separate regressions for term premia and risk-neutral rates across different maturities, employing an identical set of explanatory variables over the full sample period and two sub-periods.

The time-series regressions unveil several novel and key findings. Importantly, the decomposition of U.S. yields deepens the understanding of the spillover impacts on Turkish yield components. The magnitude and significance of the reported effects suggest that spillovers from the U.S. government bond market to the Turkish government local currency bond market primarily work through the channel from the U.S. term premia to the Turkish term premia, and solely during the pre-November 2017 period. During the implementation of unorthodox economic and financial policies in the post-November 2017 period, there is no statistically

significant evidence of pass-through from changes in the U.S. yield components to local yield components. It implies that any shift in Turkey's economic policy towards more orthodox approaches in the future may restore the transmission mechanism of the U.S. interest rate shocks, thereby increasing the local bond market's sensitivity to changes in the U.S. bond yields.

Another novel and key finding is that both baseline and alternative models reveal the operations of the signaling and duration risk channels in lowering local risk-neutral rates and term premia, respectively. From this perspective, to the best of my knowledge, this study is the first to analyze the impacts of the CBT's unconventional bond purchases, the BRSA's AR regulation, and the CBT's liraization strategy on local yield components during the post-November 2017 period. Additionally, by particularly distinguishing and analyzing the impact of the liraization strategy on yield components in the alternative specification, this study makes a useful contribution to the limited body of literature addressing the effects of financial sector regulations on bond yield components.

Notably, the regression results also indicate a significant positive relationship between the market illiquidity indicator and the local risk-neutral rates over the pre-November 2017 period. I argue that when there is a significant mispricing in the secondary bond market, stemming from the shortage of arbitrage capital in the marketplace during a financial market crisis, investors expect that the central bank will tighten the monetary conditions to uphold financial stability amid strong capital outflows. This subsequently increases the expectations component of local bond yields, with more pronounced effects on shorter maturities.

Furthermore, the analysis underscores the pivotal role of changes in sovereign credit risk and exchange rate volatility in shaping the dynamics of local term premia

and risk-neutral rates. I also identify the influence of variations in the Brent crude oil prices on local yield components, especially on the risk-neutral rates through the inflation expectations channel across all maturities and sample periods.

Overall, the empirical results contribute to a better understanding of the underlying drivers of nominal yield components in Turkey. Acquiring a deep understanding of these drivers is not only essential for accurately interpreting past changes in nominal yield components but also indispensable for forecasting their future movements, enabling policymakers, fixed-income analysts and investors to make well-informed and timely decisions in the dynamic landscape of financial markets.

CHAPTER 2

DECOMPOSITION OF BREAKEVEN INFLATION RATES AND DETERMINANTS OF INFLATION RISK PREMIA

2.1 Introduction and main results

As inflation erodes the real value of nominal bond payments, unanticipated changes in inflation pose a significant investment risk to bondholders. Therefore, bond investors require an inflation risk premium to compensate them for bearing the uncertainty associated with future inflation. While this premium provides a protective buffer to nominal bond investors against unanticipated changes in inflation beyond their expectations, those seeking to mitigate inflation risk may alternatively consider investing in inflation-indexed bonds.¹⁴¹ These bonds offer an explicit and more effective hedge against inflation by linking the coupon and principal payments to the value of an official price index of the issuing country, albeit with a time lag typically ranging from 1 to 3 months for most countries.¹⁴² Moreover, numerous sovereign issuers incorporate a deflation floor into their inflation-indexed bonds, ensuring that the principal payment remains at or above par (D’Amico, Kim, & Wei, 2018). This embedded put option provides an additional safeguard for investors, especially during prolonged periods of deflation over the tenure of inflation-indexed

¹⁴¹ The terms “inflation-indexed bonds”, “inflation-linked bonds”, “inflation-protected bonds”, “real bonds” and “linkers” are all used interchangeably in the literature and marketplace.

¹⁴² For instance, Turkish inflation-linked bonds are indexed to CPI, as announced by the Turkish Statistical Institute. The coupon and principal payments are multiplied by the ratio of reference CPI on the settlement date to that on the issue date. When the settlement date falls on day d_t of a month with d_n number of days, then the reference CPI on that day is calculated as:

$$CPI_{n-3} + \frac{d_t - 1}{d_n} (CPI_{n-2} - CPI_{n-3})$$

where CPI_{n-2} and CPI_{n-3} represent the CPI levels for the second and third months before the month in which settlement date falls, respectively. Due to the presence of the indexation lag in the inflation-linked bonds, full protection against inflation is not guaranteed. In general, this effect is greater when the indexation lag is longer, and the bond’s maturity is shorter.

bonds.¹⁴³ Despite these protections, the market for inflation-indexed bonds is generally smaller and less liquid than the market for nominal bonds. Therefore, bond investors demand an extra liquidity premium for bearing the liquidity risk associated with holding these bonds (Christensen & Gillan, 2012; Christensen & Rudebusch, 2019; Pflueger & Viceira, 2011).

Policymakers and investors closely monitor the spread between yields on comparable-maturity nominal and inflation-indexed bonds as a real-time measure of investors' future inflation expectations (D'Amico et al., 2018). More specifically, this yield spread, commonly known as the inflation compensation or breakeven inflation (BEI) rate in financial markets, represents the inflation rate that, if realized, would make an investor indifferent between holding a nominal bond and an inflation-indexed bond (Gürkaynak et al., 2010). However, as widely documented in the literature, the BEI rate includes not only expected inflation but also the inflation risk premium and the differential liquidity premium between these two types of bonds.

Building on this, the proper decomposition of BEI rates across various maturities is highly critical. First, the decomposition analysis allows for assessing the relative contribution of each component to the variation in BEI rates over time. Second, extracting the inflation risk premium and investigating its underlying determinants provide valuable insights for understanding its dynamics. Third, quantifying and comparing the inflation risk premium and the differential liquidity premium components enable governments to conduct an ex ante cost-benefit analysis when deciding whether to issue nominal or inflation-indexed bonds. While governments can refrain from paying inflation risk premia by issuing inflation-

¹⁴³ In some countries, the coupon payments are also protected against deflation, such as Turkey and Australia.

indexed bonds, their liquidity disadvantage relative to nominal bonds may negate this theoretical benefit and make their issuance more costly (Cardozo & Christensen, 2023; Christensen & Gillan, 2012; Dudley, Roush, & Ezer, 2009). Therefore, the BEI decomposition for different maturities holds significant importance for policymakers aiming to minimize the cost of local debt within the framework of sovereign debt management.

In this chapter, the main objective is to decompose the Turkish BEI rates across different maturities and to explore the local and global drivers of inflation risk premia over the full sample period from June 2009 to June 2022 and two sub-periods (pre- and post-November 2017). The other objective is to analyze, on an ex ante basis, the relative costs and benefits of issuing nominal versus inflation-indexed bonds from the perspective of the Turkish Treasury. Through rigorous investigations, this study aims to contribute to the limited existing literature on these topics in the context of emerging markets, with a specific focus on Turkey.

To derive the BEI rates over the sample period, the empirical analysis starts with the construction of end-of-month zero-coupon real yield curves. This involves the application of the same curve fitting methodology used for the estimation of zero-coupon nominal yield curves in Chapter 1. Accordingly, I employ the NSS model and the optimization objective function that incorporates the weighting scheme proposed by Gimeno and Nave (2006). In the calibration of six parameters of the NSS model, four optimization algorithms and three parameter initialization techniques are utilized. Separately, during the estimation phase, when trading activity in inflation-indexed bonds is limited on the last business day of certain months within the sample period, I implement a novel methodology to supplement the real

yield curve data for those months with prices derived from the internal rate of return (IRR) technique and, if necessary, selective use of Refinitiv evaluated prices.

Armed with the estimates of end-of-month zero-coupon nominal and real yields, I compute the BEI rates and then proceed to decompose them into expected inflation, inflation risk premium and differential liquidity premium components. For this purpose, similar to Shen (2006), Pflueger and Viceira (2011), Söderlind (2011), Grishchenko and Huang (2013), and Bekaert and Ermolov (2023), I employ an easy-to-implement “model-free” method that does not impose any no-arbitrage restrictions on nominal and real term structures. To capture the expected inflation component, I rely on inflation forecasts from the CBT’s “Survey of Market Participants”. As noted by Gülşen and Kara (2021), these survey forecasts predominantly reflect the institutional projections of the financial and real sector companies in Turkey, potentially undergoing several assessments and reviews in a structured manner.

Next, I explore the impact of liquidity differences between the Turkish nominal and inflation-indexed bonds on the BEI rates using two different regression-based approaches. As part of this investigation, a new liquidity proxy is introduced to gauge the differential liquidity between these two types of bonds in Turkey. This proxy accounts for both trading volumes and outstanding notional amounts, providing a more comprehensive measure of relative liquidity. The regression results indicate that changes in this proxy do not have a statistically significant impact on variations in the BEI or the spread between the BEI and inflation expectations. This suggests that although the inflation-indexed bonds in Turkey appear to have lower liquidity than their nominal counterparts, the local bond market does not factor in the differential liquidity premium between nominal and inflation-indexed bonds. A plausible explanation for this is that the buy-and-hold investors, such as local banks

and pension funds, who predominantly hold the inflation-indexed bonds in Turkey, prioritize the inflation-hedging benefits of these bonds over the short-term liquidity considerations. Thus, these investors are indifferent to the perceived liquidity differences between nominal and inflation-indexed bonds.

With the differential liquidity premium component deemed irrelevant in the BEI rate, I then calculate the inflation risk premium component by subtracting the survey-based average expected inflation from the BEI rate. Given that the average inflation expectations for the 3-, 4-, and 5-year horizons cannot be estimated before November 2017 due to limited forecast horizons in the CBT's survey, the estimates of the corresponding inflation risk premia are limited to the second sub-period in this study. The summary statistics show that, in the first sub-period, the 2-year inflation risk premium averages -6 bp, with a standard deviation of 74 bp. In the second sub-period, the term structure of inflation risk premia is upward sloping, with means ranging from 332 bp for the 2-year maturity to 377 bp for the 5-year maturity. Meanwhile, their volatilities decline with increasing maturity. The in-sample variance decomposition analysis of the BEI rates also reveals that, on average, the inflation risk premia are the key driver, contributing over 50% of the variation in levels and the vast majority of the variation in monthly changes. Additionally, comparing the model-based estimates of nominal term premia (obtained from the ACM model) in Chapter 1 with the model-free estimates of inflation risk premia in this chapter highlights their close alignment, reinforcing the validity and reliability of the model-free approach.

Given the estimates of inflation risk premia for different maturities in Turkey, I subsequently explore their determinants through a series of regressions. The baseline regression results indicate that, during the first sub-period, there is no

statistically significant relationship between changes in the 2-year inflation risk premium and changes in the sovereign credit risk, as proxied by changes in the Turkish 5-year CDS spread. However, in the second sub-period, coinciding with the implementation of unorthodox policies in Turkey, I observe that changes in the CDS spread have a statistically significant positive impact on changes in the country's inflation risk premia, albeit only for the 4- and 5-year maturities. Notably, a 100 bp increase in the CDS spread leads to around a 95 bp increase in the inflation risk premia for these maturities. Analyzing the impact of changes in exchange rate risk, where exchange rate risk is represented by the implied volatility of the 6-month ATM USD/TRY option contract, on changes in the inflation risk premia, the regression coefficients suggest a positive and statistically significant association for the full sample period and the second sub-period. Specifically, a 1% increase in the implied volatility corresponds to between 16 and 21 bp increases in the inflation risk premia over the second sub-period.

In the investigation of the potential monetary spillovers from the U.S. to Turkey, the baseline estimates show that changes in the U.S. term premia or risk neutral rates do not play a role in the variations of local inflation risk premia across the entire sample period and two sub-periods. Similarly, shifts in global risk sentiment, as captured by changes in the VIX index, do not contribute to changes in the inflation risk premia in Turkey over the examined time periods. On the other hand, changes in the Brent crude oil prices, reflecting fluctuations in global energy prices, have significant impact on the dynamics of the country's inflation risk premia across all maturities and specified time periods. Importantly, in line with theoretical predictions, the impact of oil price shocks on changes in local inflation risk premia decreases as maturity increases. A USD 10 per barrel increase in the Brent crude oil

prices results in 161 bp and 78 bp increases in the 2-year and 5-year local inflation risk premia, respectively, in the second-sub period.

Turning to the effects of the inflation uncertainty measures, one-month lagged inflation surprises positively and significantly influence changes in the inflation risk premia over the full sample period and the first sub-period. As for one-month lagged exchange rate surprises, they have a positive and statistically significant impact solely during the first sub-period. With the adoption of unorthodox policies in the second sub-period, both measures lose their significance.

The baseline regression results also show that while the CBT's unconventional bond purchases and the BRSA's AR regulation jointly do not affect changes in local inflation risk premia, the CBT's liraization strategy significantly reduces local inflation risk premia for the 3-, 4-, and 5-year maturities by around 360 bp. In fact, a regulation introduced in June 2022 as part of the liraization strategy prompted investors to buy nominal bonds and sell inflation-indexed bonds, which in turn led to a notable fall in the Turkish inflation risk premia. This finding is consistent with Gürkaynak, Sack, and Wright's (2008) argument that inflation risk premia can be influenced by factors beyond inflation uncertainty, such as exogenous shifts in demand for government bonds.

In the alternative specification model, I exclude variables that are statistically insignificant across all baseline regressions. The results closely resemble those of the baseline regressions, with two primary differences. First, the statistical significance of CDS coefficients for the 4- and 5-year maturities is weaker, suggesting a marginal impact of changes in the sovereign credit risk on changes in the corresponding inflation risk premia. Second, the coefficients associated with the liraization strategy not only attain statistical significance for all maturities but also imply a more

pronounced influence on local inflation risk premia. In a separate robustness check, I replace the exchange and inflation surprise measures with the inflation forecast dispersion measure from the CBT's "Survey of Market Participants". The findings indicate that changes in this inflation uncertainty measure do not impact variations in local inflation risk premia.

Lastly, I perform the ex ante cost-benefit analysis regarding the issuance of nominal versus inflation-indexed bonds in Turkey. Since the differential liquidity premium between these two types of bonds is not priced in the local bond market, the advantage or disadvantage of issuing inflation-indexed bonds over nominal bonds for the Turkish Treasury depends on the sign and magnitude of inflation risk premium. In the first sub-period, the mean value of the 2-year inflation risk premium was -6 bp, suggesting that, on average, the costs of issuing nominal and inflation-indexed bonds for this maturity were ex ante roughly equal. In the second period, the average inflation risk premia for the 2-, 3-, 4- and 5-year maturities were 332 bp, 363 bp, 376 bp, and 377 bp, respectively. This indicates that, on average, issuing inflation-indexed bonds became notably cheaper to issue compared to nominal bonds from the ex ante perspective.

The remainder of this chapter is structured as follows. Section 2 presents the literature review. Section 3 outlines the theoretical framework of BEI decomposition. Section 4 provides the construction of zero-coupon real yield curves in Turkey. Section 5 computes the corresponding BEI rates and compares them with inflation expectations from the CBT's survey. Section 6 investigates the impact of differential liquidity between the Turkish nominal and inflation-indexed bonds on the BEI rates. Section 7 and 8 respectively calculate the inflation risk premia and explore their determinants. Section 9 reports the results of ex ante cost-benefit analysis of issuing

nominal versus inflation-indexed debt in Turkey. Finally, Section 10 summarizes and concludes.

2.2 Literature review

Different methodologies are employed to decompose BEI rates. Some scholars adopt a regression-based approach to estimate liquidity premium component. For instance, Gürkaynak et al. (2010) regress 5-year and 10-year BEI rates on two liquidity measures, assuming that the differential liquidity premium component in U.S. BEI rates is solely attributable to illiquidity of inflation-indexed bonds.¹⁴⁴ The fitted values from the regression capture the time-varying liquidity impact of inflation-linked bonds on BEI rates. To quantify the level of liquidity premium, they normalize the negative of the fitted values to be zero in April 2005, a period when the liquidity premium embedded in inflation-linked bonds was at its minimum within their sample. Following this, they employ a state-space model, which integrates liquidity-adjusted BEI rates with two survey forecasts of inflation over 5-year and 10-year periods. This approach enables them to estimate the expected inflation component of BEI rates. The authors also acknowledge that inflation risk premium can be extracted in a similar way.

In another study, Söderlind (2011) computes the difference between the U.S. BEI rates and survey-based inflation expectations, referring to it as “breakeven deviation”, across 3-year and 10-year horizons.¹⁴⁵ Considering that this observed

¹⁴⁴ These liquidity measures are the ratio of trading volume of inflation-linked bonds to total Treasury trading volume, and the spread between Resolution Funding Corporation ((Refcorp) strips and Treasury strips. The second measure, known as the Refcorp spread, is assumed to be highly correlated with the liquidity premium in real bonds.

¹⁴⁵ The inflation expectations are from the Survey of Professional Forecasters (SPF), which is a quarterly survey of macroeconomic forecasts in the U.S. In the sensitivity analysis section, the author also uses 5-year inflation forecasts from the monthly Michigan Survey.

deviation, or spread, is related to liquidity premium on real bonds and inflation risk premium, it is regressed on proxies for real liquidity premium and inflation uncertainty.¹⁴⁶ While the author provides a time-series plot of the risk-adjusted BEI rates, the specific levels of the real liquidity premium and inflation risk premium are not explicitly stated.

Similarly, Pflueger and Viceira (2011) regress the difference between the U.S. 10-year BEI rate and the 10-year SPF inflation forecast on three different liquidity proxies, while controlling for the Chicago Fed National Activity Index as a proxy for short-term inflation expectations.¹⁴⁷ The liquidity premium in inflation-indexed bonds over the nominal Treasuries is then computed as the negative of the variation explained by these liquidity proxies. By adding this differential liquidity premium to BEI rate, they obtain the estimates of liquidity-adjusted BEI rate.

Focusing on the estimation of the U.S. inflation risk premia, Grishchenko and Huang (2013) initially make a correction on the inflation-linked bond yields for a 3-month indexation lag. The indexation-lag-adjusted real yields are then subtracted from the nominal Treasury yields to estimate the BEI rates. The authors subsequently calculate the inflation risk premium component by computing the difference between the BEI rates and expected inflations. As a proxy for expected inflation, they utilize

¹⁴⁶ In the main specification, the proxy for real liquidity premium is the yield difference between “off-the-run” and “on-the-run” 10-year nominal Treasury bonds, commonly referred to as “off-the-run” liquidity premium. According to the Federal Reserve Bank of Cleveland, this measure is likely to be strongly correlated with real liquidity premium. As for the proxy for inflation uncertainty, it is based on the estimations from the probability distributions (histograms) of inflation in the SPF. The additional analysis also explores alternative proxies for real liquidity premium (i.e., trading volume of inflation-linked bonds as a fraction of outstanding inflation-linked bonds volume and VIX index) and inflation uncertainty (i.e., implied volatility from bond options and disagreement among survey respondents).

¹⁴⁷ These three liquidity proxies in their benchmark specification are the nominal off-the-run spread, transaction volume of inflation-indexed bonds to the transaction volume of nominal Treasuries, and the difference between the asset-swap spread for inflation-indexed bonds and the asset-swap spread for nominal Treasuries. All liquidity proxies are normalized in such a way that they become zero under perfect liquidity conditions.

three alternative measures.¹⁴⁸ However, this inflation risk premium is not adjusted for liquidity. To derive the liquidity-adjusted inflation risk premium, they proceed to estimate the liquidity premium embedded in the real yields. For this purpose, they regress the spread between the survey-based inflation forecast and the 10-year BEI rate on three independent liquidity proxies.¹⁴⁹ The liquidity premium is then calculated as the fitted value from this regression.¹⁵⁰ Finally, by adding this liquidity premium to the previously calculated inflation risk premium, they obtain the liquidity-adjusted inflation risk premium.

Applying a similar procedure in previous studies, Vicente and Graminho (2015) decompose the BEI rates in Brazil by regressing them on three different liquidity proxies, while controlling for the expected inflation from the Brazilian survey.¹⁵¹ They find that the liquidity proxies, representing the relative liquidity of inflation-indexed bonds vis-à-vis nominal bonds, are not statistically significant at conventional levels. This implies that investors do not price any liquidity differences between these two types of bonds.¹⁵² Since the differential liquidity premia between

¹⁴⁸ These three proxies for expected inflation are average historical inflation, a VAR-based inflation forecast, and survey forecasts from the SPF and the Blue Chip Financial Forecasts (BCFF).

¹⁴⁹ The liquidity proxies are the noise measure of Hu et al. (2013) for inflation-linked bonds, the turnover ratio for inflation linked bonds (computed as the sum of all inflation-linked bond transactions divided by the amount of outstanding inflation-linked bonds), and the yield difference between off-the-run and on-the-run nominal Treasuries.

¹⁵⁰ Given that the fitted values may still incorporate some information about the inflation risk premium through the constant term of the regression, an adjustment is made to the derived liquidity premium before it is added to the inflation risk premium to estimate the liquidity-adjusted inflation risk premium.

¹⁵¹ The liquidity proxies are the ratio of difference between trading volumes of inflation-linked and nominal bonds to their total trading volumes, the difference between the bid and ask spreads of nominal and inflation-linked bonds in the secondary market, and the difference between noise measure of Hu et al. (2013) for nominal and that for inflation-linked bonds. Similar to Pflueger and Viceira (2011), they normalize all liquidity proxies to zero in the month when the liquidity of inflation-linked bonds reaches its maximum (see footnote 147). As a measure of expected inflation, they rely on the Focus survey conducted by the Central Bank of Brazil. The survey compiles macroeconomic forecasts from approximately 160 market participants (around 100 in 2015), including banks, asset managers, and various other institutions. Since the Focus survey provides inflation expectations for the next 18 months and 5 years, the monthly inflation expectations for horizons of 2, 3, and 4 years are derived by employing the flat forward interpolation technique.

¹⁵² The authors argue that inflation-indexed bonds in Brazil are usually acquired by long-term investors who hold these securities until their maturities. Therefore, as these investors are not

inflation-indexed and nominal bonds have no influence on BEI rates, the authors compute the inflation risk premia across various maturities by simply subtracting the survey-based inflation expectations from the BEI rates.

To decompose BEI rates into the aforementioned components, several studies employ no-arbitrage ATSMs that jointly price both nominal and inflation-indexed bonds. Christensen, Lopez, and Rudebusch (2010) propose a four-latent-factor model in which two of them are level factors specific to nominal and real yield curves, while the other two are common slope and curvature factors. This model does not explicitly incorporate any liquidity factor to adjust for the relative illiquidity of U.S. inflation-linked bonds over nominal Treasuries. Consequently, it decomposes the BEI rates only into inflation expectations and inflation risk premia across different time horizons. Although the authors do not include any data on inflation or inflation expectations, the model-implied inflation expectations closely align with the survey-based measures of long-term inflation expectations, lending empirical support to the model's decomposition of the BEI rates.

Abrahams, Adrian, Crump, Moench, and Yu (2016) introduce a six-factor joint pricing model, consisting of the first three PCs from the cross-section of nominal yields, two PCs extracted from real yields, and one liquidity factor.¹⁵³ The liquidity factor is a composite liquidity indicator that addresses the relative illiquidity of inflation-linked bonds with respect to nominal Treasuries.¹⁵⁴ Despite the lack of

concerned about their illiquidity in the secondary market, they do not demand any additional compensation for holding them.

¹⁵³ The real pricing factors are the first two PCs extracted from the residuals of regressions of real yields on the first three nominal PCs and the liquidity factor. This orthogonalization step is necessary to mitigate the unconditional collinearity among the pricing factors utilized in the model.

¹⁵⁴ The liquidity factor is constructed by using two liquidity indicators. The first indicator is the noise measure of Hu et al. (2013) for inflation-linked bonds. The second one is the 13-week moving average of the ratio of primary dealers' transaction volume for nominal Treasuries to that for inflation-linked bonds. To construct the liquidity factor, these two liquidity indicators are first standardized and then averaged. Finally, the negative of the time series minimum is added to each observation to ensure the positivity of the liquidity premium.

any survey inflation data in the estimation, the model generates risk-adjusted inflation expectations that closely match the survey-based forecasts.

D'Amico et al. (2018) present a no-arbitrage four-factor asset pricing framework that jointly models nominal bond yields, inflation-indexed bond yields, and realized inflation.¹⁵⁵ As supplementary data inputs to the model, the authors incorporate survey forecasts for both short-term interest rates and inflation. These inclusions stabilize the estimation process and facilitate the generation of meaningful and reliable results.¹⁵⁶ Moreover, modeled as a function of the fourth factor, the pricing framework allows for a spread between the inflation-indexed bond yields and indexation-lag-adjusted real yields. When regressing this yield spread on a set of proxies for the relative illiquidity of inflation-indexed bonds versus nominal bonds, while controlling for CPI seasonality, deflation floor and other relevant factors, D'Amico et al. (2018) find that 86.3% of the variation in the spread is driven by changes in these liquidity proxies.¹⁵⁷ The result suggests that the specified spread predominantly reflects the liquidity premium in the inflation-indexed bonds. In assessing the economic significance of BEI components, the in-sample variance decomposition analysis reveals that expected inflation, inflation risk premium, liquidity premium, and indexation-lag effects account for 49.6%, 11.1%, 35.9%, and 3.5% of movement in the 10-year BEI rate, respectively.¹⁵⁸

¹⁵⁵ In their study, D'Amico et al. (2018) consider four model specifications. In the discussion above, I focus on the full model, denoted as "Model LII" in their research.

¹⁵⁶ See Kim and Orphanides (2012) for the usefulness of supplementing term structure data with survey forecasts in the estimation of dynamic term structure models.

¹⁵⁷ These liquidity proxies are the bid-ask spread in inflation-indexed bonds, the relative trading volumes of inflation-indexed bonds compared to nominal bonds, the noise measure proposed by Hu et al. (2013) for inflation-linked bonds, the difference between asset swap spreads of inflation-indexed and off-the-run nominal bonds, and the difference between inflation swap rates and BEI rates.

¹⁵⁸ Correspondingly, for the 5-year BEI rate, these components explain 38.5%, 7%, 48.6%, and 5.9% of the movement, respectively.

In a recent study, Beauregard, Christensen, Fischer, and Zhu (2021) employ an arbitrage-free ATSM to price both nominal and real bonds in Mexico. For the joint modeling of the two term structures, the model relies on seven factors. These include the first three PCs of nominal yields, the first two PCs of real yields, as well as nominal and real liquidity risk factors. By incorporating two separate liquidity risk factors in the state space, the study aims to identify liquidity risk premia present in both nominal and inflation-linked bonds. Similar to D'Amico et al. (2018), Beauregard et al. (2021) supplement the data inputs with long-term survey forecasts of inflation. Furthermore, to mitigate the potential overinfluence of survey inflation forecasts on the estimation process, they fix the standard deviation of the corresponding measurement errors at 75 bp. The decomposition results of 5-year forward BEI rate starting 5 years ahead indicate that while the model-implied expected inflation occasionally deviates from survey forecasts primarily due to the assumed standard deviation, their close alignment highlights the model's ability to accurately capture investors' inflation expectations in the Mexican bond markets. Cardozo and Christensen (2023) implement the same joint pricing model for nominal and real bonds in Colombia, allowing them to decompose the BEI rates into inflation expectations, inflation risk premia, and differential liquidity premia. They also evaluate the relative costs and benefits of issuing nominal and inflation-linked bonds in Colombia (and Mexico) by comparing the model-implied estimates of the latter two components across various maturities.

Another approach to decompose BEI rates is to use macro-finance models. Hördahl and Tristani (2014) employ an arbitrage free macro-finance model to estimate inflation expectations and inflation risk premia in both the U.S. and the Euro area. Their model utilizes data on nominal and real yields, inflation, output gap, and

survey forecasts of short-term interest rates and inflation. Given that the relative illiquidity of inflation-indexed bonds may distort the model estimates, the authors initially correct real yields for liquidity premia before incorporating them into the estimation process. To accomplish this task, they follow the methodology outlined by Gürkaynak et al. (2010), which involves regressing BEI rates on two liquidity proxies.¹⁵⁹ Subsequently, they normalize the negative of the fitted values from regression to be zero in April 2005 and utilize the resulting series as a measure of liquidity premia associated with inflation-indexed bonds. Using liquidity-adjusted real yield data, Hördahl and Tristani (2014) proceed to estimate the model and decompose BEI rates into inflation expectations and inflation risk premia components. The premium-adjusted BEI rates (i.e., the BEI rates adjusted for both inflation and liquidity risk premia) offer a more precise estimate of inflation expectations among investors.¹⁶⁰ Additionally, in both the U.S. and the Euro area, the adjusted BEI rates align more closely with the survey forecasts in comparison to the unadjusted rates.

The empirical estimation of inflation risk premium enables the exploration of the factors that drive its movements. Söderlind (2011) regresses the breakeven deviation on potential proxies of inflation risk premium and liquidity premium to identify variables that capture these premiums. The author considers three different proxies for inflation risk premia. The first proxy is the measure derived from the cross-sectional trimmed mean of individual standard deviations of inflation forecasts. These standard deviations are obtained from probability distributions provided by

¹⁵⁹ For the Euro area, they employ only one proxy for the liquidity premium due to the lack of volume data on inflation-indexed bonds.

¹⁶⁰ The premium-adjusted BEI rates do not precisely match the model-implied expected inflations due to yield measurement errors. Nevertheless, the difference between them is minimal, amounting to only a few basis points.

respondents in the SPF. The regression analyses reveal that this measure makes a statistically significant positive contribution to the variations in the breakeven deviations across medium and long maturities, suggesting that it effectively captures the inflation risk premium. However, regarding the other two alternative proxies for inflation uncertainty, while the implied volatility from bond options yields mixed results, the disagreement among the survey respondents in the Michigan survey, measured as the cross-sectional interquartile range of inflation forecasts, is not significant, meaning that the information from this variable does not provide a suitable proxy for inflation risk premium.

Hördahl and Tristani (2014) attempt to relate the dynamics of inflation risk premia to macroeconomic variables associated with the state variables employed in their model. They find that there is a positive correlation between inflation risk premia and actual inflation rates in both the U.S. and the Euro area. Moreover, inflation risk premia in both regions show somewhat procyclical behavior between 2008 and 2013, particularly during the last quarter of 2008 when the output contracts sharply. However, over a longer period, the U.S. inflation risk premia tend to display countercyclical dynamics, similar to corresponding term premia. In examining the response of inflation risk premia to structural macroeconomic shocks in their model, the authors document that a one standard deviation aggregate demand shock leads to increases of 10 bp and 2 bp in the inflation risk premia in the U.S. and Euro area, respectively, while equivalent cost-push shock prompts larger increases of 25 bp and 8 bp in the respective regions. This indicates that cost-push shocks exert more significant impact on the inflation risk premia compared to aggregate demand shocks.

Similarly, Abrahams et al. (2016) correlate their model-implied 5–10 year forward inflation risk premium with observable macroeconomic and financial variables. Accordingly, the study uncovers that the inflation risk premium exhibits significant positive correlations with both the SMOVE index, representing three-month swaption implied Treasury volatility from Merrill Lynch, and the disagreement about inflation outlook, measured as the cross-sectional standard deviation of individual one-year-ahead inflation forecasts from the BCFF survey. Additionally, it is positively correlated with the unemployment rate and negatively correlated with the consumer confidence from the Conference Board survey, aligning with the countercyclicality of risk premia.

Camba-Méndez and Werner (2017) also provide insights into the key drivers behind the inflation risk premium in both the Euro area and the U.S. According to the regression results, the measures of inflation uncertainty, such as bond volatility, realized inflation volatility, and disagreement on inflation forecasts, do not appear to significantly influence the inflation risk premium. On the other hand, deflation fears, proxied by the price of the one-year-ahead inflation floor with a strike at 0%, have a statistically negative impact on the inflation risk premium across most specifications. Similarly, inflation fears, proxied by the price of the one-year-ahead inflation cap with a strike at 4%, demonstrate a statistically positive impact on the inflation risk premium in some specifications. The opposite signs of coefficients on the measures of inflation and deflation fears suggest that the inflation risk premium is primarily determined by “the balance of inflation risk”. This result implies that observed low or even negative inflation risk premia are mainly linked to the pricing of deflation risk. They also find that while inflation surprise indicator, derived from the difference between actual and survey-based expected inflation, has a statistically

positive impact on the model-based estimates of inflation risk premium in the U.S., the standard deviation of corresponding inflation surprises is not significant. This suggests that the direction rather than the volatility of inflation surprises matters for inflation risk premia in the U.S.

Montes and Curi (2017) investigate the drivers of the Brazilian inflation risk premium. They find that the increases in both the average expected net public debt to GDP ratio and the disagreement in these expectations lead to increases in the inflation risk premium in Brazil.¹⁶¹ The authors also document that the Central Bank's ability to anchor inflation expectations to its official inflation target serves to mitigate the country's inflation risk premia.

Beauregard et al. (2021) shed light on the determinants of the Mexican long-term inflation risk premia. Their regression analysis for the 5-year forward inflation risk premium starting 5 years ahead reveals that while the foreign-held share of the Mexican nominal bond market has a significant negative impact on the inflation risk premium, oil prices proxied by the West Texas Intermediate Cushing crude oil price, the global risk aversion captured by the VIX index, and the interest rate level in the U.S. Treasury market represented by the U.S. 10-year yield have significant positive effects. Repeating a similar analysis for Colombia, Cardozo and Christensen (2023) find that the credit risk measured by the Colombian 5-year sovereign CDS spread, the global interest rate uncertainty level proxied by the MOVE index, and the global interest rate level represented by the U.S. 10-year Treasury yield have statistically significant and positive impacts on the Colombian 5–10 year forward inflation risk premia.

¹⁶¹ The disagreement in expectations about the net public debt to GDP ratio is derived from the range of the distribution of individual expectations obtained from the survey conducted by the Brazilian Central Bank. Specifically, it is computed as the first PC of the disagreement series across maturities ranging from 1 to 4 years.

2.3 Theoretical framework of BEI decomposition

In this section, I provide the theoretical framework of BEI decomposition using notation similar to presented in D'Amico et al. (2018) and Beauregard et al. (2021).

According to standard finance theory, in a world without any financial market frictions, the nominal yield $y_{t,\tau}^N$ can be decomposed into the real yield $y_{t,\tau}^R$, expected inflation $\pi_{t,\tau}^e$, and inflation risk premium $\phi_{t,\tau}$:

$$y_{t,\tau}^N = y_{t,\tau}^R + \pi_{t,\tau}^e + \phi_{t,\tau} \quad (37)$$

where τ is the number of years to maturity. Moreover, in a frictionless world, the real yield equals the sum of the risk-neutral real rate $RN_{t,\tau}^R$ and real term premium

$$TP_{t,\tau}^R.{}^{162}$$

$$y_{t,\tau}^R = RN_{t,\tau}^R + TP_{t,\tau}^R \quad (38)$$

Then, the frictionless BEI rate can be computed as difference between nominal and real yields of the same maturity:¹⁶³

¹⁶² The interpretation of risk-neutral rate and real term premium components is similar to that of their nominal counterparts.

¹⁶³ All yields discussed here are continuously compounded zero-coupon bond yields. Therefore, the BEI is expressed as continuously compounded zero-coupon rate. As per the standard Fisher equation, the expected inflation is calculated as:

$$\pi_{t,\tau}^{e,d} = \frac{(1 + y_{t,\tau}^n)}{(1 + y_{t,\tau}^r)} - 1$$

where yields are expressed as discretely compounded rates. The relationship between the continuously compounded rate and its discretely compounded equivalent is:

$$y_{t,\tau}^N = \ln(1 + y_{t,\tau}^n)$$

$$y_{t,\tau}^R = \ln(1 + y_{t,\tau}^r)$$

$$\pi_{t,\tau}^e = \ln(1 + \pi_{t,\tau}^{e,d})$$

or equivalently:

$$e^{y_{t,\tau}^N} = 1 + y_{t,\tau}^n$$

$$e^{y_{t,\tau}^R} = 1 + y_{t,\tau}^r$$

$$e^{\pi_{t,\tau}^e} = 1 + \pi_{t,\tau}^{e,d}$$

Consequently, it follows that:

$$e^{\pi_{t,\tau}^e} - 1 = \frac{e^{y_{t,\tau}^N}}{e^{y_{t,\tau}^R}} - 1$$

This indicates that the continuously compounded expected inflation rate is equal to:

$$\pi_{t,\tau}^e = y_{t,\tau}^N - y_{t,\tau}^R$$

The derivation presented here is adapted from Duran and Gülşen (2013), see their footnote 9.

$$BEI_{t,\tau} \equiv y_{t,\tau}^N - y_{t,\tau}^R = \pi_{t,\tau}^e + \phi_{t,\tau} \quad (39)$$

However, in a real-world setting, there exist trading frictions in financial markets.

Thus, the observed nominal yield $\tilde{y}_{t,\tau}^N$ may contain an unobserved liquidity premium

$\eta_{t,\tau}^N$:

$$\tilde{y}_{t,\tau}^N = y_{t,\tau}^N + \eta_{t,\tau}^N \quad (40)$$

where $\eta_{t,\tau}^N \geq 0$ for all t and relevant τ . Similarly, the observed real yield $\tilde{y}_{t,\tau}^R$ may

contain an unobserved liquidity premium $\eta_{t,\tau}^R$.¹⁶⁴

$$\tilde{y}_{t,\tau}^R = y_{t,\tau}^R + \eta_{t,\tau}^R \quad (41)$$

where $\eta_{t,\tau}^R \geq 0$ for all t and relevant τ . Substituting equation (37) and equation (38)

into equation (40) yields:

$$\tilde{y}_{t,\tau}^N = RN_{t,\tau}^R + TP_{t,\tau}^R + \pi_{t,\tau}^e + \phi_{t,\tau} + \eta_{t,\tau}^N \quad (42)$$

where $RN_{t,\tau}^R + \pi_{t,\tau}^e = RN_{t,\tau}^N$ and $TP_{t,\tau}^R + \phi_{t,\tau} = TP_{t,\tau}^N$.¹⁶⁵ Substituting equation (38)

into equation (41) leads to:

$$\tilde{y}_{t,\tau}^R = RN_{t,\tau}^R + TP_{t,\tau}^R + \eta_{t,\tau}^R \quad (43)$$

In turn, using equation (42) and equation (43), the observed BEI can be written as:

$$\widetilde{BEI}_{t,\tau} \equiv \tilde{y}_{t,\tau}^N - \tilde{y}_{t,\tau}^R = \pi_{t,\tau}^e + \phi_{t,\tau} + (\eta_{t,\tau}^N - \eta_{t,\tau}^R) \quad (44)$$

Equation (44) shows that the observed BEI comprises expected inflation, inflation risk premium, and differential liquidity premium between nominal and real bonds.

This equation highlights the challenge of regarding BEI as a proxy for expected inflation as it also incorporates both inflation risk premium and differential liquidity premium.

¹⁶⁴ The impacts of indexation lag, deflation protection, and CPI seasonality on real yields are ignored in this section and throughout the study.

¹⁶⁵ In contrast to the decomposition used in Chapter 1 with the ACM model, this approach separates the liquidity risk premium from the nominal term premium. In the ACM model, the nominal term premium includes not only the inflation risk premium but also the liquidity risk premium (and other risk premia).

2.4 Construction of zero-coupon real yield curves

The computation of BEI rates across different maturities requires zero-coupon nominal and real yield curve curves. Having estimated the nominal yield curves in the preceding chapter, I now proceed to construct the end-of-month Turkish real yield curves by using the inflation-indexed government bond data.

2.4.1 Estimation procedure

For the construction of real yield curves over the sample period, I employ the same comprehensive and integrated curve fitting procedure used for the construction of the nominal yield curves in Chapter 1. It relies on the NSS model as a curve fitting method. To the best of my knowledge, this is the first study to utilize the NSS model for estimating zero-coupon real yield curves in Turkey.¹⁶⁶ Furthermore, in the model calibration process, I again use the objective function that incorporates Gimeno and Nave's (2006) weighting scheme. This scheme strikes a much better balance between short-maturity and long-maturity bonds during the optimization. To solve this non-convex curve fitting problem, I continue to employ the same set of alternative optimization algorithms and parameter initialization approaches, as previously utilized in the estimation of nominal yield curves.¹⁶⁷ The optimal parameter set for each month in the sample is then determined by comparing the best solutions obtained from each optimization algorithm.

¹⁶⁶ Duran and Gülşen (2013) utilize the NS model to estimate the Turkish zero-coupon real yield curve at the daily frequency from October 2009 to July 2011. Güler, Keleş, and Polat (2017) use the smoothed cubic spline method to fit the real yield curve on a daily basis for the period between October 2012 and November 2015.

¹⁶⁷ In the estimation of real yield curves, the lower and upper bounds on the values of NSS parameters are dynamically adjusted during the optimization process. Moreover, in response to negative real yields observed as of November 2021, I relax the positivity constraints on β_0 and $\beta_0 + \beta_1$, allowing them to take on negative values.

2.4.2 Data

The Turkish Treasury launched its inaugural inflation-indexed bond with a 5-year maturity on February 20, 2007. In Turkey, the inflation-indexed bonds are linked to the national CPI.¹⁶⁸ Importantly, the coupon and principal payments of such bonds are protected not only against inflation but also in the event of deflation.¹⁶⁹

The universe of inflation-indexed bonds issued by Turkey is displayed in the top panel of Figure 12. Each inflation-indexed bond is indicated by a solid black line, with the horizontal axis representing time and the vertical axis denoting the bond's remaining maturity. When it is reissued, it is marked with a red star along the corresponding line. The chart demonstrates that, from February 2007 to June 2022, Turkey issued a total of 111 inflation-indexed bonds, consisting of 30 unique bonds with 81 reissues. The bottom panel of Figure 12 illustrates the number of outstanding inflation-indexed bonds at each point in time during the same period. It exhibits an upward trend over the years, albeit with some occasional plateaus or slight decreases.

For the construction of real yield curves through the NSS model, I choose to start my sample in October 2009, corresponding with the issuance of the fourth inflation-indexed bond by Turkey, as indicated by the vertical dashed green lines in both panels of Figure 12. While theoretical considerations may suggest starting the sample in either February or April 2010, during which Turkey issued its fifth and sixth inflation-indexed bonds, respectively, the flexibility of the NSS model still allows for accurate yield curve estimation when initiated in October 2009.

To estimate zero-coupon real yield curves, I collect the secondary bond market data for inflation-indexed securities from the BIST. The curve fitting process

¹⁶⁸ See footnote 142 for the calculation of reference CPI for a particular settlement date.

¹⁶⁹ If deflation occurs over the lifespan of an inflation-indexed bond, the principal payment at maturity is equal to the par value. Moreover, in case of deflation, the calculation of coupon payments is based on the par value and its semi-annual fixed real coupon rate.

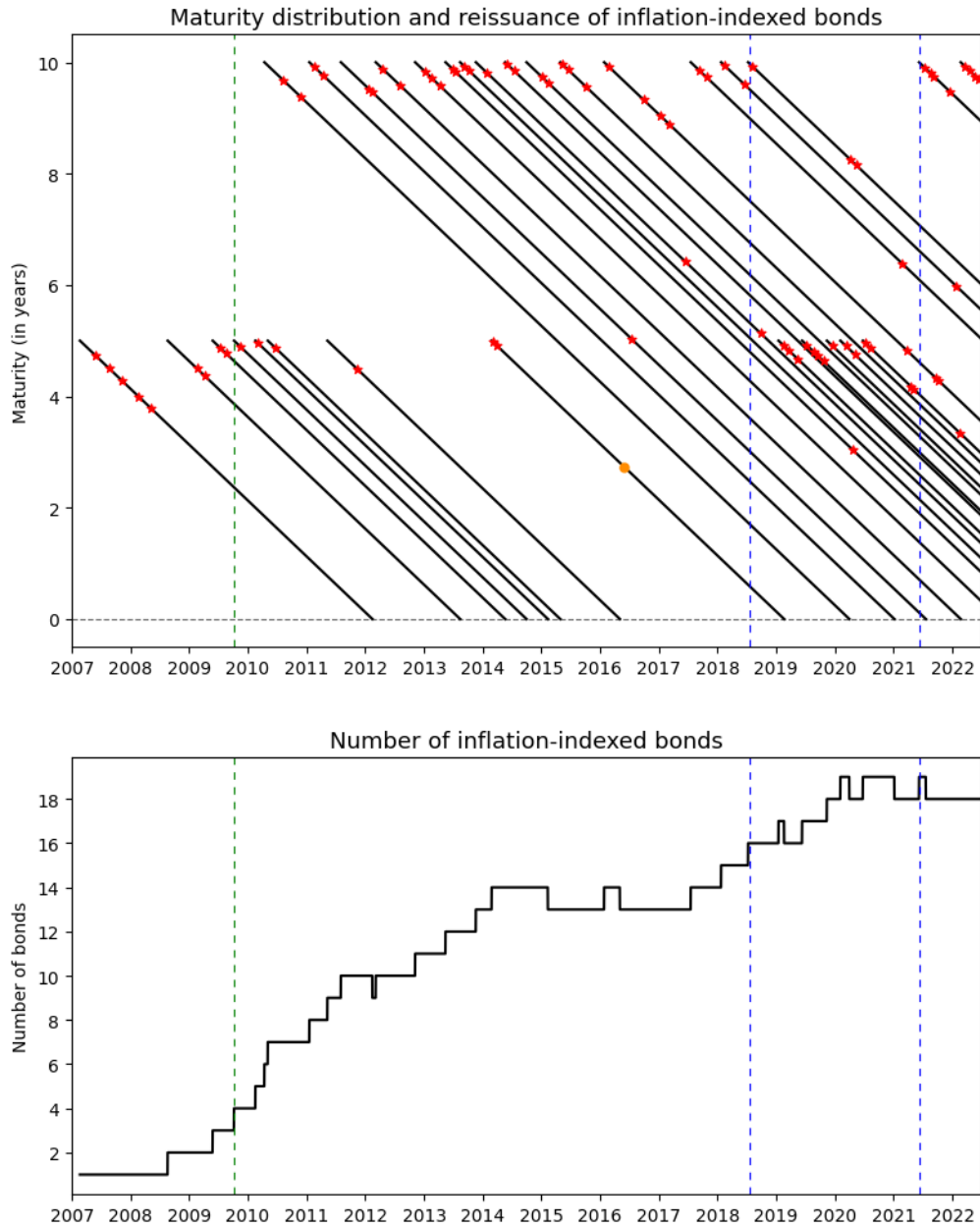


Figure 12. Overview of inflation-indexed bond market in Turkey

is conducted on the last full business day of each month from October 2009 to June 2022. After gathering the dataset, I refine it by filtering out real bonds that are clearly disconnected from the rest of the curve. This ensures that only relevant bonds are retained for real yield curve estimation.

During the estimation phase, when there are not enough distinct inflation-indexed bonds traded on the exchange on the last business day to reliably fit the real

yield curve, I examine the trading activity for inflation-indexed bonds over the preceding 4 business days to identify additional bonds that have traded during this period. For those bonds traded within that 4-day period but not on the last full business day of the month, I apply the IRR technique to estimate their end-of-month prices using their last daily weighted average prices as a reference.¹⁷⁰ Following the identification of those additional inflation-indexed bonds, I apply the same filtering process. Real bonds that are evidently disconnected from the rest of the curve are subsequently excluded from the yield curve estimation.¹⁷¹

Finally, if both actual month-end observations and those estimated using the IRR technique for inflation-indexed bonds are still insufficient due to prolonged extreme market illiquidity conditions, I then supplement the available data by selective use of the Refinitiv evaluated prices for yield curve estimation.

2.4.1 Estimation results

Table 18 reports the summary statistics for the end-of-month continuously compounded zero-coupon real yields. I do not provide the summary statistics for maturities below 2 years. The shortest maturity inflation-indexed bond outstanding in the real bond space at the end of May 2016 was set to mature on February 20, 2019,

¹⁷⁰ The IRR method provides a systematic approach to valuing bonds in the absence of trading activity on the valuation date. It enables financial professionals to determine the fair value of their bond holdings for reporting or regulatory purposes. For instance, as outlined in the “Communiqué on Principles Regarding Financial Reporting of Investment Funds” (II-14.2) published by the Capital Markets Board in Turkey, debt securities traded on the exchange but not subject to trading on the valuation date shall be valued at the IRR on the last transaction date.

¹⁷¹ The application of IRR technique assumes that market conditions and risk factors affecting the bond’s value remain relatively stable between the last trading date and the valuation date. However, the addition of bonds from the preceding 4 business days to those trading on the last full business day may introduce potential inconsistencies in bond prices or valuations if significant market events occur during this period. In such cases, the filtering operation helps to mitigate the discrepancies arising from these conditions, thereby contributing to the accuracy and reliability of the yield curve estimation.

with a remaining life of 2.73 years, illustrated as the orange dot in the top panel of Figure 12. Extrapolating backward from the available data points to estimate yields for maturities significantly lower than the shortest observed maturity may lead to inaccuracies and undermine the reliability of the estimations. Similarly, I also do not provide the summary statistics for 10-year real yield. The Turkish Treasury did not issue any 10-year maturity inflation-indexed bonds between July 11, 2018 and June 9, 2021, a period indicated by the dashed blue lines in Figure 12. Throughout this non-issuance period of 10-year bonds, the longest maturity real bond outstanding was set to mature on June 28, 2028. Therefore, particularly towards the end of this period, such as in May 2021, extrapolating forward from the available data points to estimate yields for maturities well beyond the longest observed maturity in the dataset may raise concerns about the accuracy of extrapolations and compromise the reliability of the estimations, especially when considering potential turning or inflection points within 7-10 year segment of the real yield curve.

Table 18. Summary Statistics for Zero-Coupon Real Yields

Maturity	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
2-year	0.0180	0.0271	-0.1373	0.0170	0.0229	0.0297	0.0511	153
3-year	0.0202	0.0227	-0.1191	0.0192	0.0248	0.0299	0.0468	153
4-year	0.0220	0.0192	-0.0907	0.0203	0.0258	0.0301	0.0470	153
5-year	0.0235	0.0161	-0.0623	0.0214	0.0267	0.0307	0.0451	153
7-year	0.0252	0.0120	-0.0280	0.0229	0.0267	0.0310	0.0439	147

The summary statistics cover the period from October 2009 to June 2022 for maturities ranging from 2 to 5 years, and from April 2010 to June 2022 for the 7-year maturity.

The summary statistics in Table 18 reveal that the average term structure of real interest rates is upward sloping. However, their volatilities decrease with maturity. Figure 13 shows the estimates of the selected set of real yields over the sample period. During the early phase of the sample period, the real yields generally followed a downward trend and dipped below 1% for all maturities in April 2013,

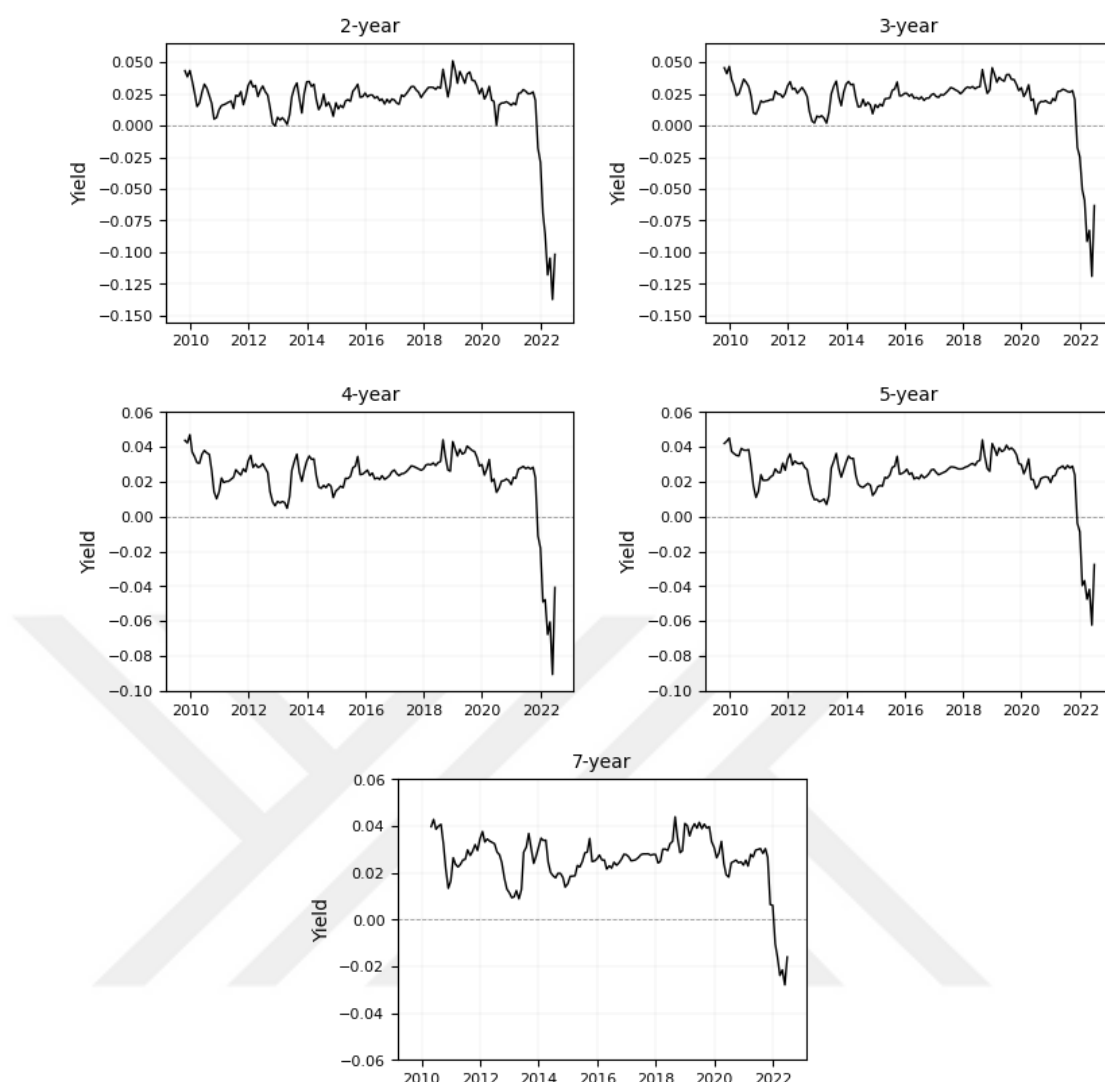


Figure 13. Zero-coupon real yields

shortly before the onset of the taper tantrum episode. In the subsequent period from late 2013 to the first half of 2018, they fluctuated within a relatively stable range. Nonetheless, there was a notable increase in these yields during the local currency crisis in August 2018. After a brief decline, the real yields surged again in December 2018, and remained elevated in the first half of 2019. Following this, as the medium-term inflation outlook improved due to favorable supply conditions and sluggish domestic demand, the CBT reduced its policy rate from 24% to 10.75% between July 2019 and February 2020, exceeding market expectations. This reduction contributed

to the decline in real interest rates. Thereafter, in response to economic challenges posed by the coronavirus pandemic, the CBT opted for additional rate cuts totaling 250 bp between March and May 2020. These policy rate adjustments, coupled with the unconventional bond purchases by the central bank and the bond purchases by local banks under the scope of AR regulation during the second quarter, extended the downward pressure on real yields. In the third quarter of 2020, prompted by the deterioration in inflation expectations, the CBT began tightening monetary policy. In this context, from September 2020 to March 2021, the policy rate was raised from 8.25% to 19%. This reversed real yields to pre-pandemic levels. However, despite the heightened inflation and inflation expectations, the CBT kept the policy rate on hold until August 2021, and then unexpectedly lowered it to 18% in September 2021. This was followed by additional cuts totaling 400 bp in the last quarter of 2021. In the wake of this, although the conflict between Russia and Ukraine in late February 2022 worsened the already high domestic inflationary pressures, the CBT maintained its policy rate at 14%. Consequently, after hitting negative levels in November 2021, the real yields plunged further into negative territory in 2022, reflecting expectations of no change in the policy rate and surging inflation in the near term.

Figure 14 depicts a three-dimensional plot of the zero-coupon real yield curve data, specifically focusing on the 2-7 year segment. It presents both the cross-sectional variation of the real yields across different maturities and the dynamic evolution of the real yield curve over the sample period.

Figure 15 displays the estimated zero-coupon real yield curves together with the continuously compounded actual and fitted yields to maturity for the specific dates. It is evident that the estimation framework can capture different yield curve shapes. Notably, the comparison of actual and fitted yields offers key insights into

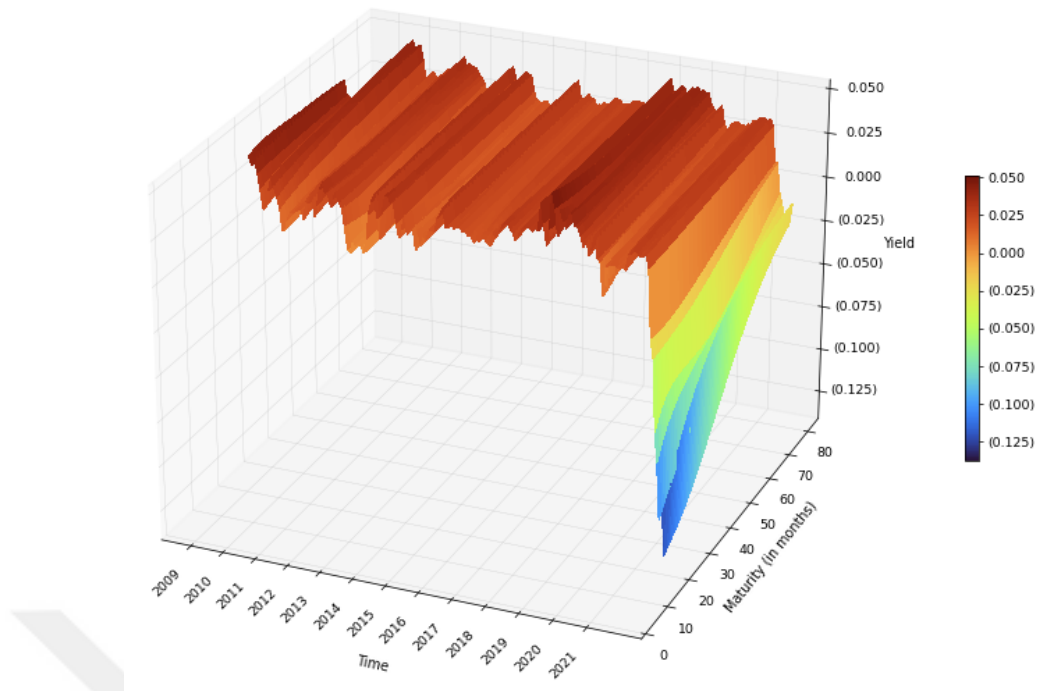


Figure 14. Zero-coupon real yield surface

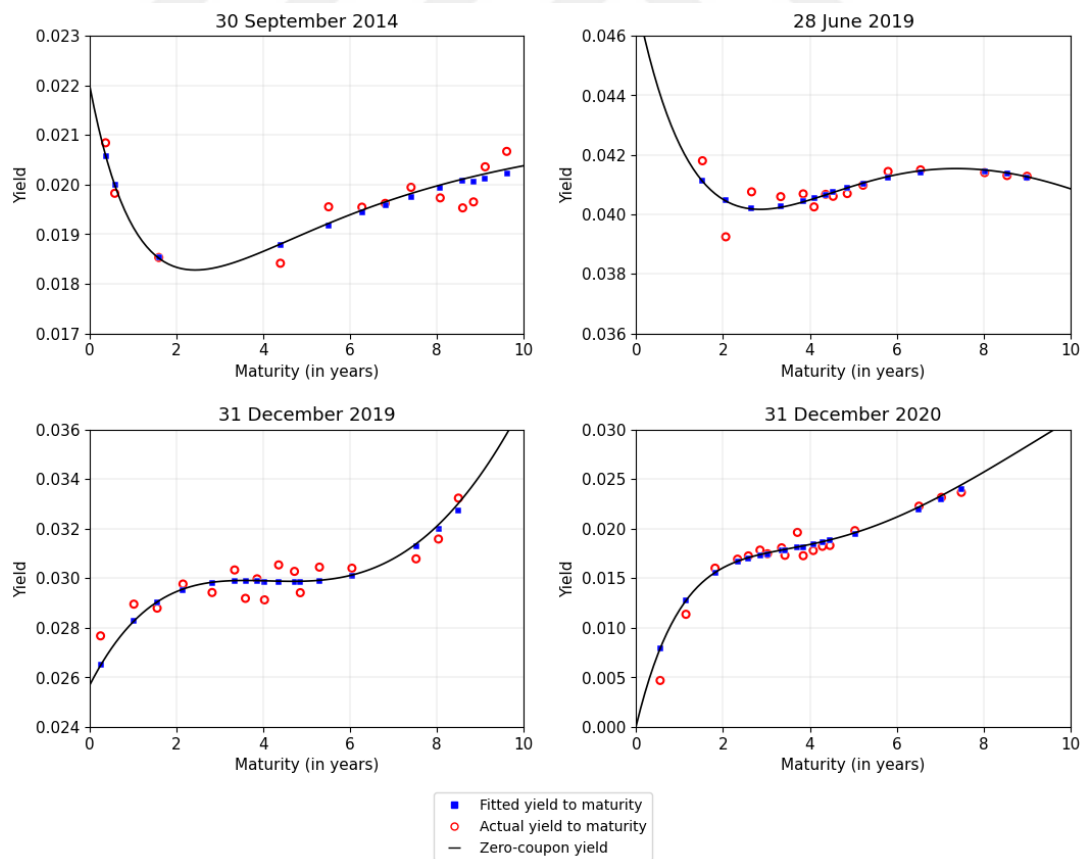


Figure 15. Real yield curve estimation for selected dates

the quality of fit. Overall, the real yield curve estimation demonstrates a high degree of accuracy.

2.5 BEI rates

In this section, I first calculate the BEI rates and then compare them with the survey-based inflation expectations from the CBT’s “Survey of Market Participants”.

2.5.1 Computation of BEI rates

Equipped with the estimates of zero-coupon nominal and real yields, $\tilde{y}_{t,\tau}^N$ and $\tilde{y}_{t,\tau}^R$, respectively, it is straightforward to compute the BEI rates ($\widetilde{BEI}_{t,\tau}$) using equation (44).

Table 19 presents the summary statistics for the end-of-month continuously compounded zero-coupon BEI rates across maturities ranging from 2 to 7 years. The average term structure of BEI rates is downward sloping. This is consistent with my previous findings that the average nominal and real yield curves beyond 2-year maturity exhibit a downward and upward slope, respectively (see Tables 2 and 18). Regarding the volatility of BEI rates, it diminishes with increasing maturity.

Table 19. Summary Statistics for BEI Rates

Maturity	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
2-year	0.0964	0.0590	0.0460	0.0635	0.0745	0.1074	0.3679	153
3-year	0.0941	0.0544	0.0491	0.0636	0.0723	0.1053	0.3348	153
4-year	0.0919	0.0501	0.0486	0.0633	0.0711	0.1028	0.2987	153
5-year	0.0897	0.0461	0.0474	0.0625	0.0705	0.1002	0.2792	153
7-year	0.0876	0.0408	0.0462	0.0626	0.0716	0.0999	0.2543	147

The summary statistics cover the period from October 2009 to June 2022 for maturities ranging from 2 to 5 years, and from April 2010 to June 2022 for the 7-year maturity.

Figure 16 shows the estimates of the selected set of Turkish BEI rates over the sample period. These rates demonstrated a generally stable trend until the end of April 2018, after which they experienced four notable periods of sharp increases. The initial spike occurred during the currency crisis in August 2018, and the second one coincided with the post-local election uncertainty in April 2019. The third spike, specifically for the medium-term maturities, emerged in March 2021 after the replacement of the central bank governor. Finally, the fourth spike took place in March 2022 for the medium-term maturities, and in May 2022 for the short-term maturities, following the renewed currency crisis in December 2021 and the Russia-Ukraine conflict in late February 2022.

2.5.2 Comparison of BEI rates with survey-based inflation expectations

As shown in equation (44), the BEI rate is driven by expected inflation, inflation risk premium and differential liquidity premium between nominal and real bonds. One commonly utilized proxy for expected inflation in the literature is survey-based inflation forecasts.¹⁷² The CBT has been conducting the “Survey of Expectations” since August 2001 to gather forecasts on inflation and other key macroeconomic variables from a panel of market participants (renamed “Survey of Market Participants” in June 2021). While initially conducted bi-monthly, during the first and third week of each month until December 2012, the survey transitioned to a monthly frequency starting in January 2013. Gülşen and Kara (2021) report that the survey participants are primarily from financial institutions, including commercial banks, investment banks, asset management companies, insurance and factoring

¹⁷² See Shen (1998), Shen (2006), Pflueger and Viceira (2011), Söderlind (2011), Grishchenko and Huang (2013), Vicente and Graminho (2015), Montes and Curi (2017), Gülşen and Kara (2019), and Bekaert and Ermolov (2023).

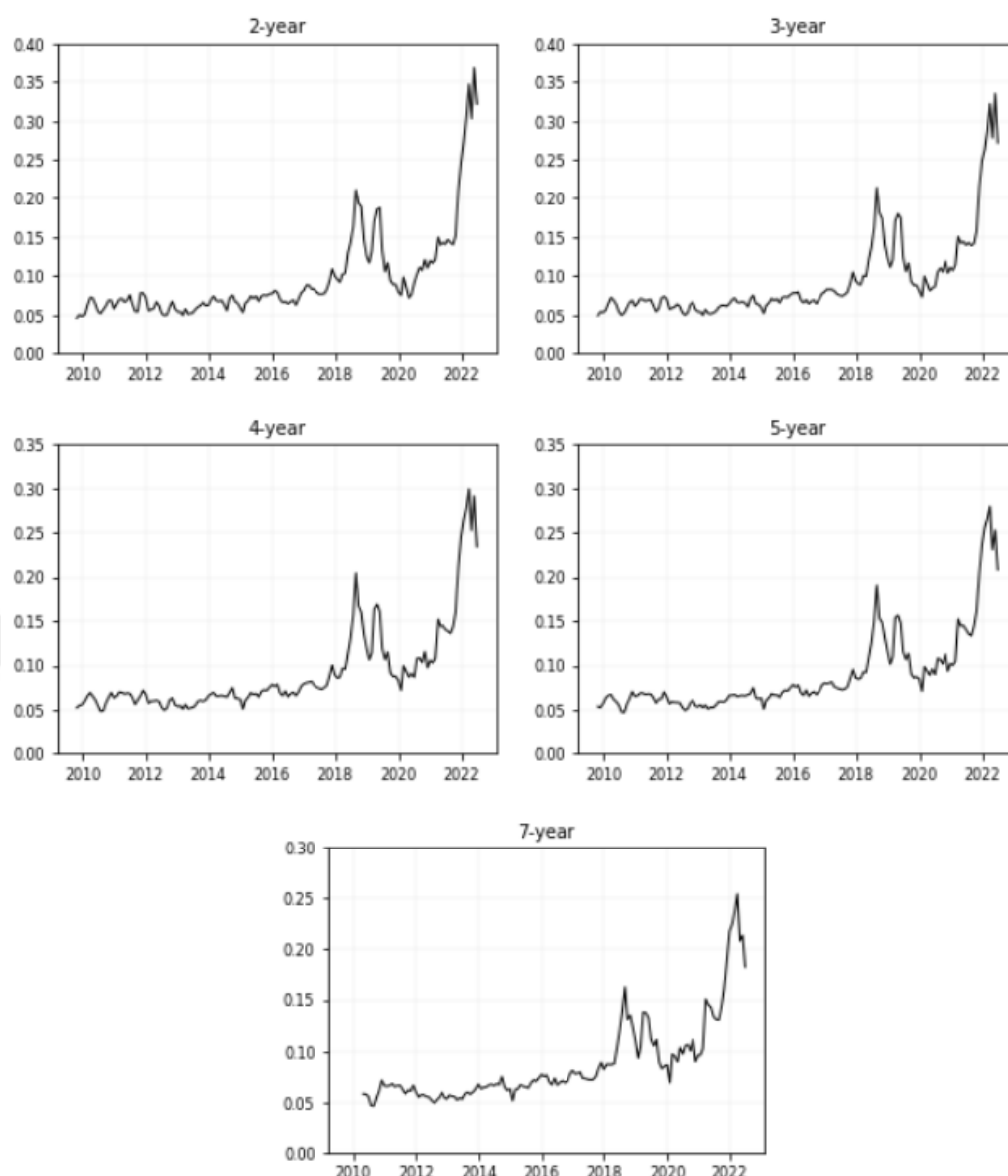


Figure 16. BEI rates

companies, pension funds, as well as real sector companies, including large firms and conglomerates.¹⁷³ Their study further highlights that the survey, distributed to around 100 participants, has maintained a response rate of 60%-70% since 2006.

Additionally, the same study provides insights from a feedback survey conducted among participants, revealing that approximately three-fourths of the respondents

¹⁷³ Financial institutions comprise approximately 80% of the survey participants, as indicated by Gülşen and Kara (2021).

indicate that their forecasts represent institutional projections, potentially involving rigorous checks and reviews. Considering the sophistication of the forecasting process, the authors argue that survey responses should be less vulnerable to issues such as cheap-talk problems, herd behavior, and strategic misreporting. Therefore, in the comparison of BEI rates and inflation expectations, I rely on inflation forecasts of the professional and representative agents in the CBT's "Survey of Market Participants".

For the period from August 2001 to March 2006, the maximum horizon for inflation forecasts available in the survey was 1-year. With the CBT's adoption of a full-fledged explicit inflation targeting regime at the beginning of 2006, the survey was expanded to incorporate the 2-year-ahead annual inflation expectations in April of the same year.¹⁷⁴ While the 5-year-ahead (between year 4 and year 5) and the 10-year-ahead (between year 9 and year 10) horizons were also added to the survey in November 2017, the latter one was excluded in August 2021.¹⁷⁵ Given the evolution of forecast horizons in the survey data, average annual inflation expectations can only be estimated for the 2-year horizon over the entire sample period from October 2009 to June 2022. Estimations for the 3-, 4- and 5-year horizons are only feasible starting from November 2017, corresponding to the availability of the 5-year-ahead inflation expectations in the survey.

To compute the average inflation expectations, I first collect the (modified mean) inflation expectations data for the 1-, 2-, and 5-year-ahead horizons.¹⁷⁶ Next, I

¹⁷⁴ The 2-year-ahead annual inflation expectations, denoted as $\pi_{t|t+1 \rightarrow t+2}^e$, represent the anticipated annual inflation rate at the end of a 2-year period, specifically covering time interval between year 1 and year 2.

¹⁷⁵ The addition of the 5- and 10-year-ahead inflation expectations to the survey in November 2017 coincided with the adoption of unorthodox policies in Turkey, as discussed in Chapter 1.

¹⁷⁶ Since the survey was conducted bi-monthly until December 2012, I utilize the inflation forecasts from the second survey conducted each month for the period between October 2009 and December 2012.

employ logarithmic interpolation to approximate the 3- and 4-year-ahead annual inflation forecasts from these data points. Afterwards, I estimate the average inflation expectations for any horizon by geometrically averaging the annual expected inflation rates within that horizon. Lastly, the resulting average inflation expectations are converted into continuously compounded rates to align them with the BEI rates, ensuring consistency for accurate comparison.

Figure 17 shows the estimates of the 2-, 3-, 4- and 5-year BEI rates (solid black lines) and corresponding average inflation expectations (solid blue lines). The plots reveal strong correlations between these two sets of estimates. During the period from October 2009 to October 2017 (the left side of vertical dashed grey lines), the correlation for the 2-year horizon stands at 0.738, while from November 2017 to June 2022, the correlations are notably high, with values of 0.915, 0.888, 0.859, and 0.831 for the 2-, 3-, 4-, and 5-year horizons, respectively.¹⁷⁷ Figure 17 also illustrates that the BEI rates exhibited greater volatility compared to the average inflation expectations.¹⁷⁸ This pattern may be attributed to variations in the inflation risk premium and/or the differential liquidity premium between nominal and inflation-indexed bonds over time.¹⁷⁹ Importantly, the BEI rates mostly exceeded the average expected inflation rates in the period following November 2017, aligning with Turkey's implementation of unorthodox policies. In view of equation (44), the positive spread between the BEI rate and the survey-based inflation expectation provides evidence of the positive and time-varying inflation risk premium

¹⁷⁷ The correlations between changes in BEI rates and changes in corresponding average inflation expectations across both periods range from 0.182 to 0.372. However, excluding June 2022 shifts the range to 0.313-0.459.

¹⁷⁸ The BEI rates, on average, are 1.8 times more volatile than the average inflation expectations over the observed periods.

¹⁷⁹ The impacts of indexation lag, deflation protection, and CPI seasonality on real yields are disregarded. See footnote 164.

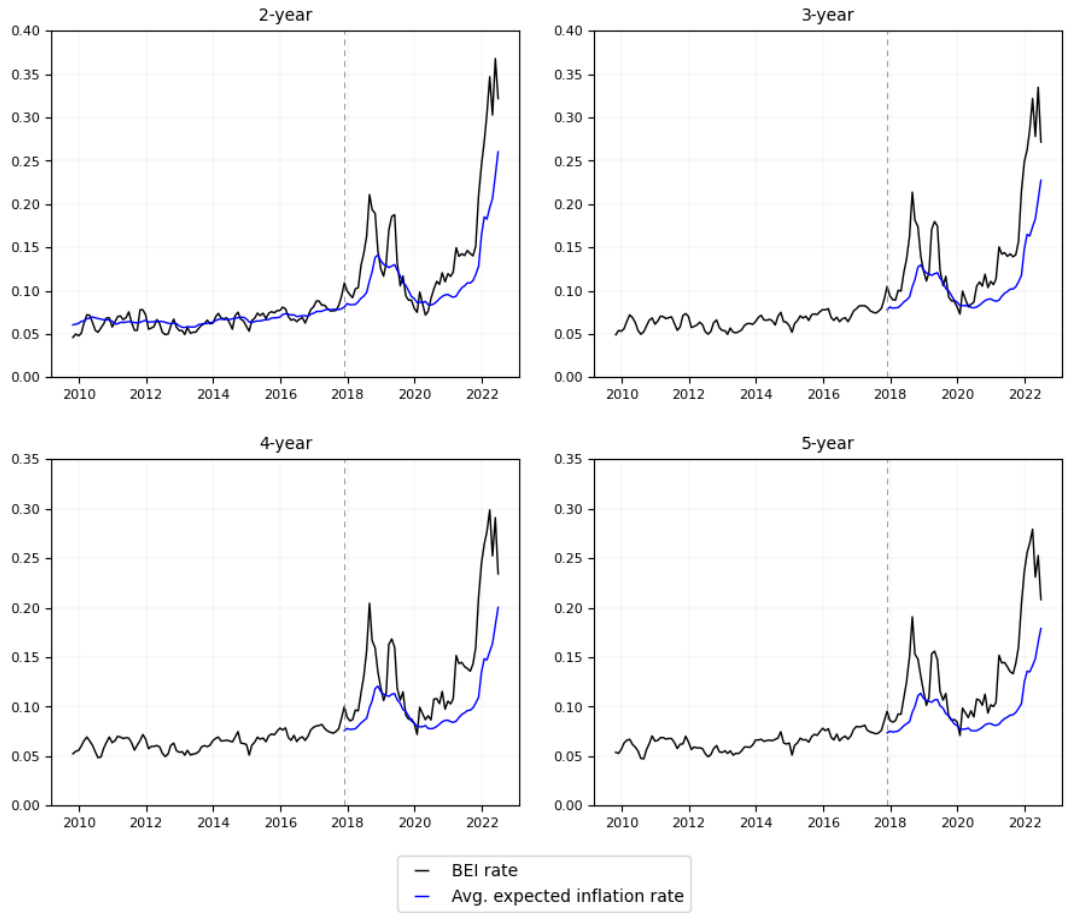


Figure 17. BEI rates versus average inflation expectations from the CBT's survey

component ($\phi_{t,\tau} > 0$), which offsets the non-positive differential liquidity premium between nominal and inflation-indexed bonds.¹⁸⁰

2.6 Impact of differential liquidity on BEI

In this section, I investigate the impact of differential liquidity between the Turkish nominal and inflation-indexed bonds on the BEI rates across different maturities using two different regression-based approaches described in the literature section.

¹⁸⁰ Given that inflation-indexed bonds tend to be less liquid than nominal bonds, investors may demand a higher liquidity premium for holding inflation-indexed bonds relative to nominal bonds (Pflueger & Viceira, 2011; Shen, 1998), resulting in $\eta_{t,\tau}^N - \eta_{t,\tau}^R \leq 0$. This, in turn, may push BEI rates down.

2.6.1 Proxy for differential liquidity

To identify the differential liquidity between nominal and inflation-indexed bond markets and estimate its impact on the BEI rates, various liquidity proxies are utilized in the literature (see, for example, footnotes 144, 146, 147, 149, and 151).

One widely employed proxy is the ratio of trading volume of inflation-indexed bonds to that of nominal bonds. However, this measure fails to consider the outstanding notional amounts of the bonds. Therefore, to provide a more comprehensive assessment of relative liquidity, I choose to compare the trading volumes of inflation-indexed and nominal bonds, each scaled by their respective outstanding notional amounts. In essence, this metric involves comparing the turnover ratio of inflation-indexed bonds to that of nominal bonds, where the turnover ratio represents the trading volume divided by the outstanding notional amount.¹⁸¹ However, for my monthly analysis, considering the observed market illiquidity on the last full business day of certain months over the sample period, particularly notable in the inflation-indexed bond market, I adjust the turnover ratios of both types of bonds by recalculating them using the total trading volumes over the last 5 business days of each month, with these volumes scaled by their respective end-of-month outstanding notional amounts.¹⁸² By implementing this adjustment methodology tailored to the Turkish market, I aim to provide a more accurate measure of relative liquidity and effectively mitigate the impact of intermittent market liquidity problems.

The top left and right panels of Figure 18 compare the trading volumes of the Turkish nominal and inflation-indexed bonds over the last 5 business days of each

¹⁸¹ When comparing turnover ratios, I consider their ratios, rather than their differences, in contrast to Güler et al.'s (2017) approach.

¹⁸² Only bonds with the same trade and settlement dates are included in the calculation of the total trading volumes over this 5-day period. For nominal bonds, I use both zero-coupon and fixed-coupon bonds.

month and their end-of-month outstanding notional amounts, respectively. Regarding the trading volumes, the nominal bonds generally exhibited a downtrend over the sample period. In contrast, the inflation-indexed bonds initially showed fluctuations without a clear trend until mid-2014, followed by a subsequent downtrend. In terms of the outstanding notional amounts, both types of bonds displayed an almost monotonic upward trend throughout the sample period.

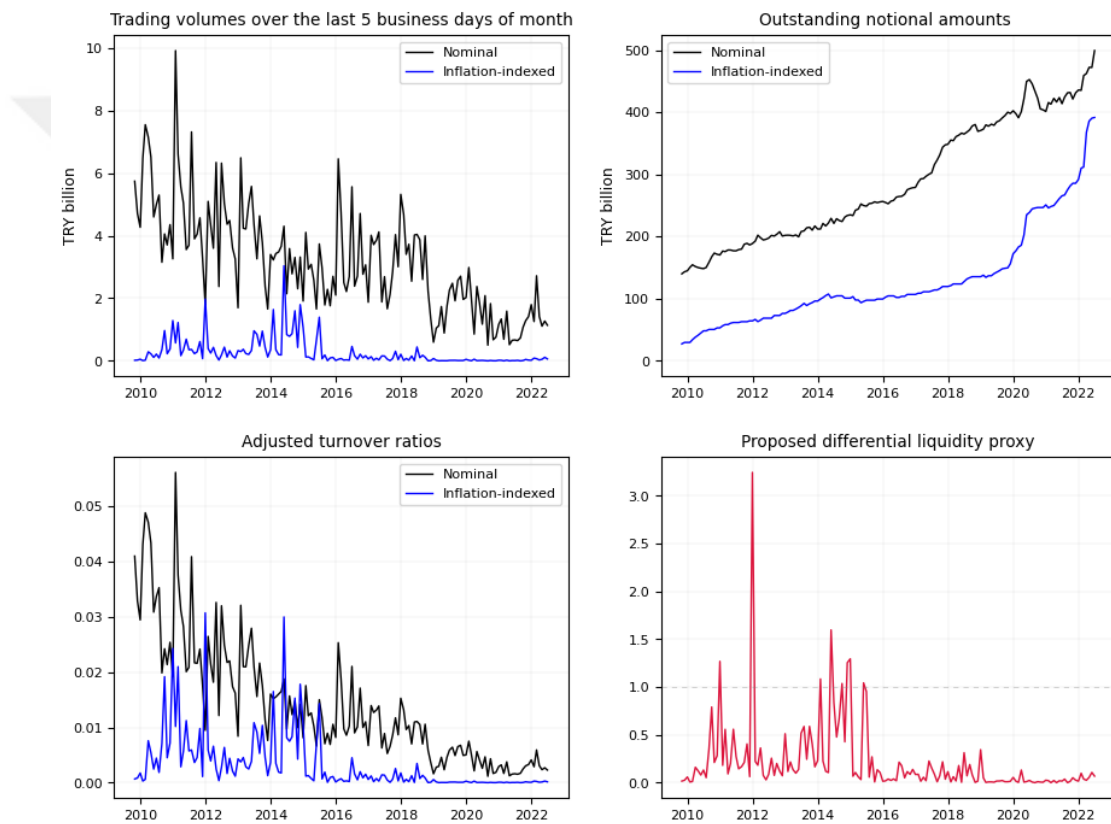


Figure 18. Comparison of nominal and inflation-indexed bonds, and proposed differential liquidity proxy

Moving to the bottom left panel of Figure 18, it depicts the respective turnover ratios. Consistent with the trends observed in trading volumes and outstanding notional amounts, the turnover ratio of the nominal bonds demonstrated a downward trajectory, while that of the inflation-indexed bonds illustrated occasional spikes until mid-2014, followed by a notable decline and stabilization at

particularly low levels. The bottom right panel of Figure 18 manifests the proposed differential liquidity proxy, calculated by dividing the adjusted turnover ratio of the inflation-indexed bonds by that of the nominal bonds. During the sample period, it predominantly remained below 1, suggesting relatively lower liquidity of the inflation-indexed bonds compared to the nominal bonds in Turkey.

2.6.2 Unit root tests for stationarity

Before estimating the regressions, I perform the ADF unit root tests to assess the stationarity of the time series variables (BEI rates, breakeven deviation rates, and winsorized differential liquidity proxy) to be included in the models.¹⁸³ This precautionary step is essential to avoid the risk of spurious regressions, ensuring reliability and validity of the subsequent empirical analyses. Table 20 presents the tests results with constant only and constant and trend specifications. The findings suggest that, at the 5% level of statistical significance, the null hypothesis of a unit root cannot be rejected for any of the variables in their level form. Nevertheless, when the ADF tests are applied to the first differences, the null hypothesis of a unit root is strongly rejected for all variables, indicating that the variables which are non-stationary in their level forms become stationary upon first differencing.

2.6.3 Estimating the impact of differential liquidity on BEI

I estimate the impact of differential liquidity on the BEI rates across various maturities using two regression-based approaches. In the first approach, following

¹⁸³ Breakeven deviation rate refers to the difference between BEI rate and average inflation expectation from the CBT's survey. Moreover, to mitigate the potential impact of outliers on statistical analyses, I winsorize the proposed differential liquidity proxy. This entails replacing the highest two values with the third highest value. For the remainder of the study, I employ this winsorized form instead of the raw data.

Table 20. ADF Unit Root Test Results

Variable	At Level		At First-Difference	
	Constant, No Trend	Constant & Trend	Constant, No Trend	Constant & Trend
2-year BEI	0.511 (0.987)	-0.267 (0.991)	-6.742 (0.000)	-6.881 (0.000)
3-year BEI	0.695 (0.992)	-0.808 (0.962)	-6.915 (0.000)	-6.998 (0.000)
4-year BEI	0.264 (0.976)	-0.676 (0.972)	-6.852 (0.000)	-6.891 (0.000)
5-year BEI	0.006 (0.957)	-0.590 (0.978)	-6.877 (0.000)	-6.895 (0.000)
2-year breakeven deviation	-0.009 (0.956)	-1.180 (0.910)	-7.881 (0.000)	-7.843 (0.000)
3-year breakeven deviation	-0.363 (0.907)	-0.687 (0.968)	-4.269 (0.001)	-4.160 (0.001)
4-year breakeven deviation	-0.363 (0.907)	-0.823 (0.956)	-3.983 (0.003)	-3.888 (0.019)
5-year breakeven deviation	-0.398 (0.901)	-0.976 (0.938)	-3.876 (0.004)	-3.805 (0.024)
Differential liquidity proxy	-2.066 (0.259)	-3.007 (0.134)	-18.933 (0.000)	-18.873 (0.000)

The optimal lag length for the ADF tests is based on the modified Akaike information criterion (MAIC), as suggested by Ng and Perron (2001). The MacKinnon approximate p-values are reported in parentheses below the computed ADF statistics.

the methodology outlined by Gürkaynak et al. (2010), I regress the BEI rates on the differential liquidity proxy introduced in Section 2.6.1.¹⁸⁴ However, given the findings from the ADF unit root tests, to avoid the spurious results, I use the first differences of the variables instead of their levels in the regressions.

Furthermore, similar to Chapter 1, I split the sample in November 2017 to account for Turkey's adoption of the unorthodox economic and financial policy mix. The analysis is conducted for the full sample period (October 2009 to June 2022) and two sub-periods (pre- and post-November 2017), allowing for the identification of potential changes in both the statistical significance and the magnitude of the differential liquidity's impact over time.

¹⁸⁴ To accurately identify the impact of differential liquidity on the BEI rates, the chosen differential liquidity proxy must be uncorrelated with both the expected inflation and inflation risk premium. Otherwise, there is a risk of omitted variable bias in the regression analysis.

Tables 21a, 21b and 21c show the regression results for the 2-, 3-, 4-, and 5-year horizons over the full sample period, the first sub-period, and the second sub-period, respectively. The findings indicate that changes in the differential liquidity proxy exhibit the expected positive sign, with the exception of the 2-year horizon in the second sub-period. Nevertheless, across all horizons and sample periods, these changes do not demonstrate a statistically significant effect on changes in the BEI rates.¹⁸⁵ In other words, the results suggest that changes in the differential liquidity between nominal and inflation-indexed bonds have no impact on the BEI rates.

In the second approach, similar to Shen (2006), Pflueger and Viceira (2011), Söderlind (2011), Grishchenko and Huang (2013), and Bekaert and Ermolov (2023), I examine the impact of differential liquidity proxy on the observed breakeven deviation rates.¹⁸⁶ Since these rates, which represent the spread between BEI rates and average inflation expectations from the CBT's survey for the corresponding horizons, are non-stationary and become stationary after first differencing (see Table 20), I employ a regression model in which I regress changes in the breakeven deviation rates on changes in the differential liquidity proxy.¹⁸⁷ Moreover, considering that the average inflation expectations for the 3-, 4-, and 5-year horizons can only be estimated starting from November 2017, the regressions of the corresponding breakeven deviation rates are conducted solely for the post-November 2017 period. Tables 22a, 22b, and 22c report the regression results for the specified time periods. It is evident that there is no statistically significant

¹⁸⁵ While not presented in Table 21, the results also extend to the 7-year horizon.

¹⁸⁶ As also highlighted by Bekaert and Ermolov (2023), to accurately identify the impact of differential liquidity on breakeven deviation rates, the chosen differential liquidity proxy must be uncorrelated with the inflation risk premium. Otherwise, the regression analysis may be susceptible to omitted variable bias.

¹⁸⁷ Shen (2006) also employs the regression of changes to avoid spurious results. While Pflueger and Viceira (2011) estimate regressions in levels, they also conduct the regressions with all variables expressed in quarterly changes due to concerns about spuriousness, particularly because some of the liquidity proxies are persistent.

Table 21. Regressions of Δ BEI Rates on Δ Differential Liquidity Proxy

(a) Full Sample Period (October 2009 – June 2022)

	Δ BEI (%)			
	2-year	3-year	4-year	5-year
Δ Differential Liquidity Proxy	0.1294 (0.720)	0.1942 (1.278)	0.1776 (1.249)	0.1263 (0.888)
Constant	0.1813 (1.132)	0.1463 (1.130)	0.1195 (0.992)	0.1014 (0.915)
Observations	152	152	152	152
R ²	0.0007	0.0017	0.0017	0.0011

(b) First Sub-Period (October 2009 – October 2017)

	Δ BEI (%)			
	2-year	3-year	4-year	5-year
Δ Differential Liquidity Proxy	0.1582 (0.921)	0.1947 (1.388)	0.1602 (1.222)	0.1046 (0.777)
Constant	0.0493 (0.892)	0.0431 (0.884)	0.0372 (0.849)	0.0337 (0.816)
Observations	96	96	96	96
R ²	0.0099	0.0231	0.0204	0.0096

(c) Second Sub-Period (November 2017 – June 2022)

	Δ BEI (%)			
	2-year	3-year	4-year	5-year
Δ Differential Liquidity Proxy	-0.5585 (-0.282)	0.2663 (0.129)	0.7267 (0.373)	0.7858 (0.450)
Constant	0.4057 (1.154)	0.3235 (0.960)	0.2620 (0.832)	0.2195 (0.760)
Observations	56	56	56	56
R ²	0.0005	0.0001	0.0011	0.0016

Note: The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period.

Table 22. Regressions of Δ Breakeven Deviations on Δ Differential Liquidity Proxy

(a) Full Sample Period (October 2009 – June 2022)

	Δ Breakeven Deviation (%)
	2-year
Δ Differential Liquidity Proxy	0.0836 (0.489)
Constant	0.0499 (0.442)
Observations	152
R ²	0.0003

(b) First Sub-Period (October 2009 – October 2017)

	Δ Breakeven Deviation (%)
	2-year
Δ Differential Liquidity Proxy	0.1100 (0.655)
Constant	0.0299 (0.586)
Observations	96
R ²	0.0053

(c) Second Sub-Period (November 2017 – June 2022)

	Δ Breakeven Deviation (%)			
	2-year	3-year	4-year	5-year
Δ Differential Liquidity Proxy	-0.6325 (-0.354)	0.6326 (0.308)	1.0642 (0.552)	1.0454 (0.613)
Constant	0.0822 (0.273)	0.0310 (0.098)	0.0165 (0.054)	0.0127 (0.044)
Observations	56	55	55	55
R ²	0.0008	0.0007	0.0023	0.0027

Note: The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period.

relationship between changes in the differential liquidity proxy and changes in the breakeven deviation rates.

The findings from both approaches reveal that changes in the liquidity differences between these two types of bonds in Turkey do not contribute significantly to the variations in the BEI or breakeven deviation rates during the analyzed time periods. This result is in line with Duran and Gülşen (2013), who document that daily changes in liquidity conditions in the Turkish nominal and real bond markets have no significant effect on inflation compensation between January 2010 and May 2011. Furthermore, the lack of significant impact is also broadly consistent with Güler et al. (2017), who report that the average liquidity premium difference between 10-year nominal and inflation-indexed bonds in Turkey is -3 bp over the period from October 2012 to November 2015.¹⁸⁸

Overall, my results imply that although the Turkish inflation-indexed bonds seem to be less liquid than their nominal counterparts, the differential liquidity between them is not priced within the local bond market. This finding resonates with the situation observed in Brazil, as documented by Vicente and Graminho (2015). Similar to their explanation, one plausible reason for this phenomenon is that the inflation-indexed bonds are mainly held by buy-and-hold investors. In Turkey, where inflation dynamics exhibit high volatility, investors such as local banks and pension funds prioritize the inflation-hedging benefits of inflation-indexed bonds over short-term liquidity concerns. This preference for inflation protection, in turn, mitigates the perceived importance of liquidity risk differences between nominal and inflation-

¹⁸⁸ Despite this, Güler et al. (2017) also note this difference ranges from -31 bp to 43 bp.

indexed bonds, leading to the absence of a statistically significant differential liquidity premium.

2.6.4 Robustness analysis

To check the robustness of my results, I re-estimate the regressions using two alternative differential liquidity proxies. The first proxy measures the ratio of total trading volume of inflation-indexed bonds to that of nominal bonds over the last 5 business days of each month. The second proxy considers the difference between the adjusted turnover ratios of inflation-indexed and nominal bonds. The results of these alternative regressions consistently support the conclusion that changes in the differential liquidity between these two types of bonds do not have a statistically significant impact on the dynamics of the BEI or breakeven deviation rates in Turkey.

2.7 Inflation risk premium

Given the lack of a statistically significant relationship between the differential liquidity of inflation-indexed versus nominal bonds and the BEI rates across various maturities, suggesting the irrelevance of the differential liquidity premium component (i.e., $\eta_{t,\tau}^N - \eta_{t,\tau}^R \approx 0$) in the BEI rates, then the inflation risk premium component can be computed by subtracting the average expected inflation from the BEI rate (see equation 44).¹⁸⁹

¹⁸⁹ This implies that the breakeven deviation rate in Turkey effectively represents the inflation risk premium. Gülşen and Kara (2019) consider the difference between BEI and average expected inflation across a 2-year horizon as a market-based proxy for inflation uncertainty in Turkey, while acknowledging that this measure may contain the liquidity premium and other factors. However, my regression results indicate the absence of statistically significant differential liquidity premia in the BEI or breakeven deviation rates.

Tables 23a and 23b provide the summary statistics for the computed inflation risk premium for the periods from October 2009 to October 2017 and from November 2017 to June 2022, respectively. As the average annual inflation expectations for the 3-, 4-, and 5-year horizons cannot be estimated before November 2017 due to the limited forecast horizons in the CBT's survey (see Section 2.5.2), the computations of the corresponding inflation risk premia are confined to the second time interval.¹⁹⁰

According to Table 23a, in the first sub-period, the 2-year inflation risk premium has a mean of -6 bp with a standard deviation of 74 bp. During the second sub-period, as indicated by Table 23b, the average term structure of inflation risk premia is upward sloping, with means ranging from 332 bp for the 2-year maturity to 377 bp for the 5-year maturity. The volatility of inflation risk premia also declines with increasing maturity. Notably, the standard deviation of the 2-year inflation risk premium in the second sub-period is 385 bp, representing a 5.2-fold increase compared to the first sub-period.

Table 23. Summary Statistics for Inflation Risk Premium

(a) Between October 2009 and October 2017								
Maturity	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
2-year	-0.0006	0.0074	-0.0159	-0.0059	-0.0009	0.0054	0.0145	97

(b) Between November 2017 and June 2022								
Maturity	Mean	Std. Dev.	Min.	Lower Quartile	Median	Upper Quartile	Max.	Obs.
2-year	0.0332	0.0385	-0.0157	0.0078	0.0256	0.0533	0.1510	56
3-year	0.0363	0.0381	-0.0097	0.0092	0.0261	0.0509	0.1480	56
4-year	0.0376	0.0370	-0.0079	0.0114	0.0263	0.0535	0.1430	56
5-year	0.0377	0.0351	-0.0065	0.0125	0.0257	0.0518	0.1377	56

Note: The computations of 3-, 4-, and 5-year inflation risk premiums are limited to the period from November 2017 to June 2022.

¹⁹⁰ The periods from October 2009 to October 2017 and from November 2017 to June 2022 also coincide with the first and second sub-periods, as previously discussed when splitting the sample in November 2017 in response to Turkey's adoption of unorthodox policies.

Figure 19 displays the temporal evolution of the inflation risk premia. In the first sub-period, the 2-year inflation risk premium remained within a relatively stable range. In the second sub-period, similar to the movements observed in the BEI rates, the inflation risk premia underwent substantial fluctuations, characterized by four notable periods of sharp increases. The first spike took place during the currency crisis in August 2018, followed by the second spike amid the post-election uncertainty in April 2019. The third spike, particularly pronounced for the medium-term horizons, coincided with the replacement of the central bank governor in March 2021. Lastly, the fourth spike emerged in March 2022, in the aftermath of the renewed currency crisis in December 2021 and the conflict between Russia and Ukraine in late February 2022.

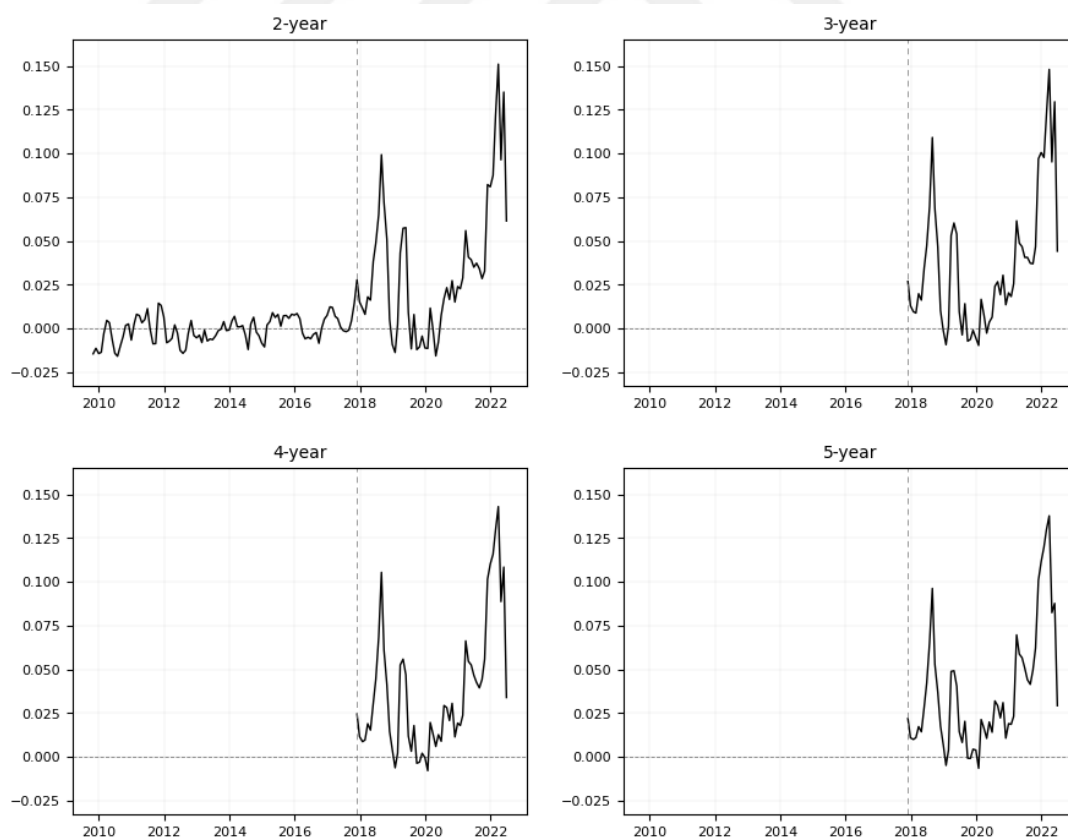


Figure 19. Inflation risk premium

With the identification of the inflation risk premium, I can conduct an in-sample variance decomposition analysis of the BEI rate. This allows for assessing the relative contribution of each component of the BEI rate to its variation over time.

Given that $\eta_{t,\tau}^N - \eta_{t,\tau}^R \approx 0$, utilizing equation (44), the variance of the BEI rate can be written as follows:

$$Var[\widetilde{BEI}_{t,\tau}] = Cov[\widetilde{BEI}_{t,\tau}, \widetilde{BEI}_{t,\tau}] = Cov[\widetilde{BEI}_{t,\tau}, \pi_{t,\tau}^e] + Cov[\widetilde{BEI}_{t,\tau}, \phi_{t,\tau}] \quad (45)$$

Dividing each side of equation (45) by $Var[\widetilde{BEI}_{t,\tau}]$ gives the following expression:

$$1 = \frac{Cov[\widetilde{BEI}_{t,\tau}, \pi_{t,\tau}^e]}{Var[\widetilde{BEI}_{t,\tau}]} + \frac{Cov[\widetilde{BEI}_{t,\tau}, \phi_{t,\tau}]}{Var[\widetilde{BEI}_{t,\tau}]} \quad (46)$$

Tables 24a and 24b present the variance decompositions of the BEI rates, both in their level and first-difference forms, as per equation (46). The results in Table 24a indicate that, in the first sub-period, the inflation risk premium accounts for approximately two-thirds of the variation observed in the 2-year BEI rate. In the second sub-period, while the expected inflation explains half of movements at the 2-year horizon, its contribution decreases with maturity, dropping to 37.3% at the 5-year horizon. Since the BEI rates are highly persistent, I also estimate the variance decompositions of their monthly changes. Table 24b reveals that when the BEI rates and their components are expressed in the first-difference forms, the inflation risk premium emerges as the dominant factor, explaining between 89.2% and 95.3% of the variation in the changes in the BEI rates.

In concluding this section, I undertake a comparative analysis between the nominal term premia estimated by the ACM model in Chapter 1 and the inflation risk premia derived from the model-free approach in this chapter, which does not rely on any standard no-arbitrage ATSM.¹⁹¹

¹⁹¹ Recall from equation (42) that $TP_{t,\tau}^R + \phi_{t,\tau} = TP_{t,\tau}^N$. However, the nominal term premia estimates from the ACM model implicitly accommodate the liquidity risk premia.

Table 24. Variance Decomposition of BEI Rates

(a) Levels			
	Period	Expected Inflation	Inflation Risk Premium
2-year	October 2009 – October 2017	0.374	0.626
2-year	November 2017 – June 2022	0.504	0.496
3-year	November 2017 – June 2022	0.453	0.547
4-year	November 2017 – June 2022	0.407	0.593
5-year	November 2017 – June 2022	0.373	0.627
(b) First-Differences			
	Period	Expected Inflation	Inflation Risk Premium
2-year	October 2009 – October 2017	0.065	0.935
2-year	November 2017 – June 2022	0.108	0.892
3-year	November 2017 – June 2022	0.063	0.937
4-year	November 2017 – June 2022	0.053	0.947
5-year	November 2017 – June 2022	0.047	0.953

Given that $\eta_{t,\tau}^N - \eta_{t,\tau}^R \approx 0$, the variance decomposition analysis relies on equation (46).

Figure 20 illustrates the positive co-movement between the nominal term premia and the inflation risk premia across the specified maturities. In the first sub-period, the correlation between the model-based estimates of the nominal term premium and the model-free estimates of the inflation risk premium for the 2-year maturity stands at 0.535. Moving to the second sub-period, the correlations are notably strong, with values of 0.90, 0.913, 0.931, and 0.944 for the 2-, 3-, 4-, and 5-year maturities, respectively. When considering the relationships in the first differences, the 2-year maturity exhibits a correlation of 0.448 in the first sub-period, while in the second sub-period, the correlation coefficients for the 2-, 3-, 4-, and 5-year maturities reach 0.761, 0.753, 0.792, and 0.848, respectively.¹⁹² Overall, the close alignment of the estimates of inflation risk premia with the estimates of term premia reinforces the validity and reliability of the model-free approach.

¹⁹² The relatively lower correlation between the level or first difference estimates of inflation risk premia and nominal term premia in the first sub-period may primarily stem from the impact of real term premia, although the presence of liquidity risk premia embedded in the nominal term premia could also contribute.

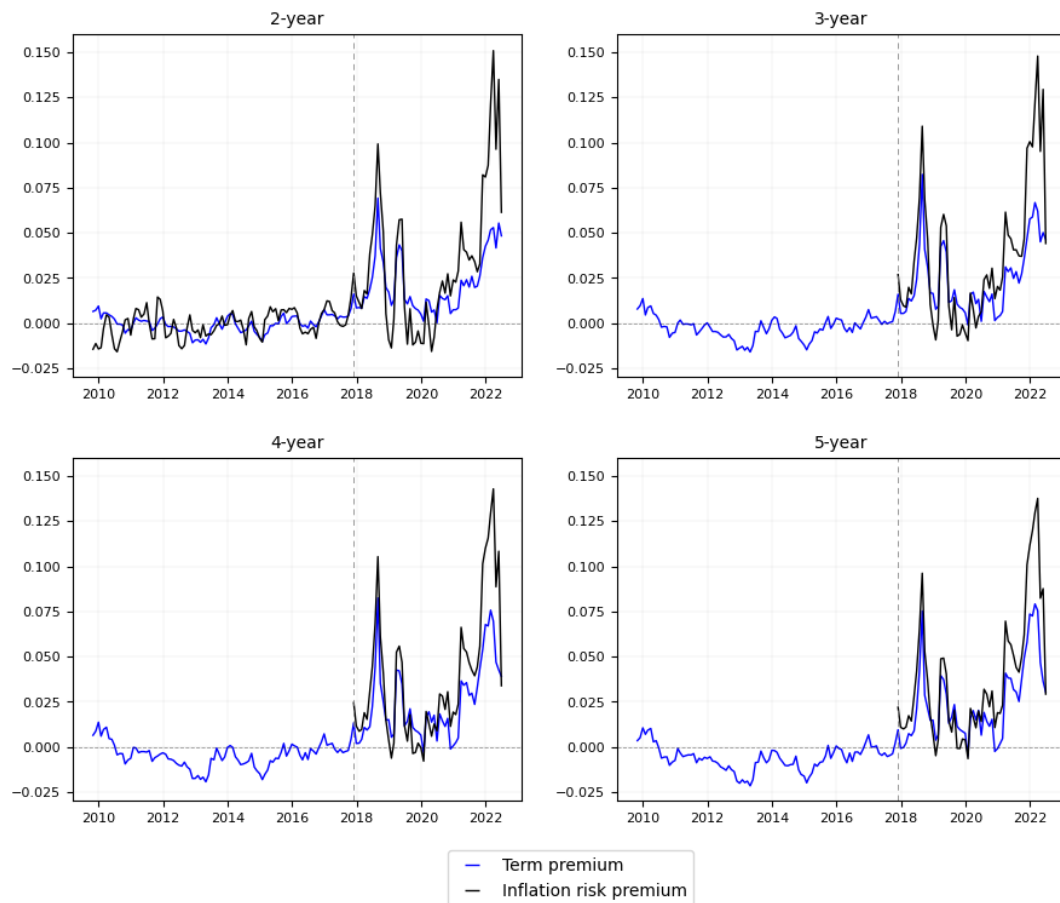


Figure 20. Model-based term premia versus model-free inflation risk premia

2.8 Determinants of inflation risk premium

Armed with the estimates of the inflation risk premium component, this section investigates its determinants through a series of regressions.

2.8.1 Unit root tests for stationarity

Before running the time-series regressions, I conduct the ADF unit root tests to evaluate the stationarity of the inflation risk premia. The results in Table 25 indicate that, at the 5% level of statistical significance, the null hypothesis of a unit root cannot be rejected for the level forms of inflation risk premia. However, when the ADF tests are applied to the first differences, the null hypothesis of a unit root is

strongly rejected for all, indicating that the nonstationary time series of inflation risk premia become stationary upon first differencing.

Table 25. ADF Unit Root Test Results

Variable	At Level		At First-Difference	
	Constant, No Trend	Constant & Trend	Constant, No Trend	Constant & Trend
2-year inflation risk premium	-0.009 (0.956)	-1.180 (0.910)	-7.881 (0.000)	-7.843 (0.000)
3-year inflation risk premium	-0.363 (0.907)	-0.688 (0.968)	-4.269 (0.001)	-4.160 (0.001)
4-year inflation risk premium	-0.363 (0.907)	-0.822 (0.956)	-3.983 (0.003)	-3.888 (0.019)
5-year inflation risk premium	-0.398 (0.901)	-0.976 (0.938)	-3.876 (0.004)	-3.805 (0.024)

The optimal lag length for the ADF tests is based on the modified Akaike information criterion (MAIC), as suggested by Ng and Perron (2001). The MacKinnon approximate p-values are reported in parentheses below the computed ADF statistics.

2.8.2 Explanatory variables and baseline model specification

For the regression of inflation risk premium, I employ the same set of explanatory variables utilized in the baseline regression model of term premium in Chapter 1, albeit with two modifications.¹⁹³ First, I replace the illiquidity measure for the nominal bond market with the proposed differential liquidity proxy that measures the liquidity differences between the Turkish nominal and inflation-indexed bond markets. Second, similar to approach taken in the alternative model specification in Chapter 1, I replace the composite “Aggregate Policy” dummy with two other time dummies. The first time dummy captures the joint impact of the CBT’s unconventional bond purchases and the BRSA’s AR regulation, while the second time dummy exclusively represents the CBT’s liraization strategy.

¹⁹³ Refer to Section 1.5.2.1 for the detailed list of the explanatory variables used in the regression of term premium in Turkey.

Based on the results of the ADF unit root tests (see Tables 13 and 25), to avoid the spurious regressions problem, my baseline regression model of inflation risk premium employs the first differences of all the variables, except for the exchange rate and inflation surprise measures, as well as the two time dummies.¹⁹⁴ Therefore, to explain the monthly changes in the Turkish inflation risk premium, I use the following regression model:

$$\begin{aligned}\Delta IRP_t^i = & \alpha_i + \beta_1^i \Delta CDS_t + \beta_2^i \Delta ATM6m_t + \beta_3^i \Delta DIFF_LIQ_t + \beta_4^i \Delta US_TP_t^i \\ & + \beta_5^i \Delta US_RN_t^i + \beta_6^i \Delta VIX_t + \beta_7^i \Delta Brent_t + \beta_8^i \Delta USDTRY_Surprise_{t-1} \\ & + \beta_9^i \Delta CPI_Surprise_{t-1} + \beta_{10}^i \Delta UBP_AR_t + \beta_{11}^i \Delta Liraization_t + \varepsilon_t^i\end{aligned}$$

where Δ represents the first-difference operator, the dependent variable IRP^i denotes the Turkish inflation risk premium for maturity i , CDS is the 5-year sovereign CDS spread of Turkey, $ATM6m$ is the implied volatility of the 6-month ATM USD/TRY option contract, $DIFF_LIQ$ is the differential liquidity proxy, US_TP^i is the U.S. term premium for maturity i , US_RN^i is the U.S. risk-neutral rate for maturity i , VIX is the VIX index, $Brent$ is the Brent crude oil price, $USDTRY_Surprise$ is the exchange rate surprise measure, $CPI_Surprise$ is the inflation surprise measure, UBP_AR is the time dummy variable reflecting joint impact of the CBT's unconventional bond purchases and the BRSA's AR regulation, and $Liraization$ is the time dummy variable representing the CBT's liraization strategy.

In line with the methodology utilized in the previous analyses, I perform the regressions over the full sample period and two sub-periods. This allows for the assessment of potential changes in both the statistical significance and the magnitude of the impacts of the explanatory variables over time.

¹⁹⁴ See footnote 125 for the explanation of why I choose not to use first differences for the two surprise measures.

Additionally, since the inflation risk premia for the 3-, 4-, and 5-year maturities can only be estimated starting from November 2017, the corresponding time-series regressions are conducted solely for the second sub-period.

2.8.3 Baseline regression results

Table 26 presents the baseline regression results for changes in the inflation risk premium component in Turkey across the selected maturities within the specified time periods. The findings suggest that changes in the sovereign credit risk, proxied by changes in the Turkish 5-year CDS spread, do not exhibit a statistically significant effect on changes in the 2-year inflation risk premium over the full sample period and the first-sub period. During the second sub-period, characterized by Turkey's adoption of unorthodox policies, the strength of the positive relationship between changes in the inflation risk premia and changes in the CDS spread increases with maturity. Nonetheless, the statistical significance is achieved only for the 4- and 5-year maturities. Specifically, a 100 bp increase in the CDS spread is associated with approximately a 95 bp increase in the inflation risk premia for these respective maturities.

Turning to the impact of changes in exchange rate risk, where exchange rate risk is represented by the implied volatility of the 6-month ATM USD/TRY option contract, the corresponding regression coefficients imply that there is a statistically significant positive relationship at the 5% level between changes in the inflation risk premia and changes in the implied volatility for the full sample period and the second sub-period. Importantly, the impacts are more pronounced for short-term maturities compared to medium-term maturities in the second sub-period. For instance, a 1% increase in the implied volatility corresponds to 21 bp and 16 bp increases in the 2-

Table 26. Baseline Regression Results for the Inflation Risk Premium

Explanatory Var.	Full Sample	First Sub-Period	Second Sub-Period			
	2-year	2-year	2-year	3-year	4-year	5-year
Δ CDS Spread	0.0038 (0.768)	-0.0005 (-0.131)	0.0049 (0.781)	0.0069 (1.191)	0.0092* (1.817)	0.0097** (2.187)
Δ Implied Vol.	0.1855** (2.293)	0.0868 (1.129)	0.2104** (2.403)	0.2409** (2.556)	0.1982** (2.333)	0.1604** (2.292)
Δ Diff. Liq. Proxy	0.0627 (0.382)	0.0465 (0.251)	-0.5278 (-0.402)	0.4436 (0.293)	0.8529 (0.647)	0.8337 (0.762)
Δ US Term Prem.	-1.4143 (-0.623)	0.6865 (0.628)	-5.5808 (-1.391)	-2.3863 (-0.960)	-1.2172 (-0.701)	-0.8423 (-0.573)
Δ US Risk-Neutral	-0.2532 (-0.347)	0.3397 (0.746)	-1.1233 (-0.822)	-0.4748 (-0.379)	-0.5400 (-0.500)	-0.5094 (-0.446)
Δ VIX	0.0090 (0.349)	-0.0112 (-0.581)	0.0614 (1.368)	0.0278 (0.703)	0.0013 (0.033)	-0.0121 (-0.293)
Δ Brent	0.0839*** (3.024)	0.0220** (2.533)	0.1610*** (3.000)	0.1336** (2.554)	0.1041** (2.359)	0.0781** (2.122)
$\$/TRY$ Surprise M.	-0.0016 (-0.037)	0.0480** (2.134)	-0.0164 (-0.196)	-0.0345 (-0.464)	-0.0303 (-0.534)	-0.0321 (-0.704)
CPI Surprise M.	0.0394*** (2.636)	0.0200*** (2.866)	0.0724 (1.332)	0.0623 (1.131)	0.0627 (1.270)	0.0584 (1.386)
UBP & AR Reg.	-0.9276 (-0.696)		-0.6643 (-0.509)	-0.9146 (-0.779)	-1.2072 (-0.957)	-1.2687 (-0.932)
Liraization Strgy.	-2.1586 (-0.991)		-1.7004 (-0.763)	-3.5453 (-1.615)	-3.7573* (-1.958)	-3.5002** (-2.115)
Constant	-0.0132 (-0.159)	-0.0055 (-0.126)	-0.2368 (-1.074)	-0.1607 (-0.859)	-0.1140 (-0.776)	-0.0736 (-0.585)
Observations	152	96	56	55	55	55
Adjusted R ²	0.3518	0.1150	0.4272	0.5105	0.5489	0.5557

Note: The full sample period, spanning from October 2009 to June 2022, is divided into two sub-periods: pre- and post-November 2017. The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period. Asterisks ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively. In the regression analysis, the exchange rate and inflation surprise measures lag one month. UBP stands for “unconventional bond purchases”.

year and 5-year local inflation risk premia, respectively, in the post-November 2017 period.

Moreover, the regression results indicate that changes in the differential liquidity proxy, which reflect changes in the liquidity differences between nominal and inflation-indexed bonds in Turkey, have no statistically significant effect on changes in the inflation risk premia across the entire sample period and the two sub-periods. This finding is consistent with the assumptions made earlier in the

regression analyses of the BEI and the breakeven deviation rates in Section 2.6.3 (see footnotes 184 and 186).¹⁹⁵

The estimation results in Table 26 also offer valuable insights into the sensitivity of local inflation risk premia to the global risk factors. Regarding the spillover effects from the U.S. bond yields to local currency government bond yields, the findings show that changes in the U.S. term premia or risk-neutral rates do not have a statistically significant impact on changes in the Turkish inflation risk premia throughout the full sample period and two sub-periods. Similarly, there is no statistically significant relationship between changes in the VIX index, capturing shifts in the global risk sentiment, and changes in the inflation risk premia in Turkey over the examined time periods. However, changes in the Brent crude oil prices, serving as a proxy for fluctuations in global energy prices, have a strong positive effect on changes in the country's inflation risk premia across all maturities and specified time periods. The comparison of regression coefficients of the 2-year inflation risk premium in the first and second sub-periods suggests a notable increase in sensitivity to changes in the Brent crude oil prices in the latter period, which coincides with Turkey's implementation of unorthodox economic and monetary policies. Furthermore, in line with theoretical expectations, the regression results reveal a diminishing influence of oil price shocks on changes in local inflation risk premia as maturity increases. Specifically, during the post-November 2017 period, a USD 10 per barrel increase in the Brent crude oil prices results in 161 bp and 78 bp increases in the 2-year and 5-year local inflation risk premia, respectively.

¹⁹⁵ The lack of significance in the impact of changes in the differential liquidity proxy on changes in the inflation risk premia supports the validity of the regression analyses conducted in Section 2.6.3.

Examining the roles of surprise measures, it is evident from the regressions that inflation surprises exert a positive and statistically significant impact on changes in the Turkish inflation risk premia at the 1% level throughout the full sample period and the first sub-period. These results imply that, during these respective time periods, unexpected upward or downward changes in domestic inflation in the previous month drive movements in local inflation risk premia in the same direction for the subsequent month. Regarding exchange rate surprises, they demonstrate a positive and statistically significant relationship with changes in local inflation risk premia at the 5% level, albeit only during the pre-November 2017 period.

Moving to the joint impact of the CBT's unconventional bond purchases and the BRSA's AR regulation on changes in local inflation risk premia, the corresponding regression coefficients are not statistically significant. On the other hand, the CBT's liraization strategy has a statistically significant negative impact on changes in local inflation risk premia for the 3-year maturity at around 10% level, and for the 4- and 5-year maturities at around 5% level. When this strategy is in effect, the 3-, 4-, and 5-year inflation risk premia decline by 355 bp, 376 bp, and 350 bp, respectively. These declines are primarily attributed to the CBT's new regulation introduced in June 2022 as part of the liraization strategy. This regulation (publicly announced on June 10, detailed in a document sent to banks on June 13, and officially came into effect as of June 24, 2022) mandated banks to maintain additional Turkish lira long-term fixed-coupon GDDS for foreign currency deposits, on top of required reserves. Accordingly, banks were required to hold between 3% and 10% of such deposits in fixed-coupon GDDS with an original maturity of at least 5 years and a residual maturity of at least 4 years. Moreover, the new rule also prohibited the use of non-fixed coupon bonds, including CPI-indexed bonds, as

collateral for foreign currency deposits. Consequently, the significant buying activity for nominal bonds, coupled with the selling activity for inflation-indexed bonds, contributed to a notable compression in local inflation risk premia in June 2022. This finding aligns with the assertion of Gürkaynak et al. (2008), who note that the inflation risk premium may not solely reflect inflation uncertainty. They highlight that exogenous shifts in demand for long-term nominal bonds (by foreign central banks) may affect nominal bond yields and subsequently impact the inflation risk premium, regardless of any change in inflation uncertainty. Similarly, Shen (2006) indicates that regulatory changes may influence the demand for nominal and inflation-indexed bonds differently, prompting changes in BEI rates. In the context of Turkey, the decline in the BEI rates, attributed to the divergent impact of the CBT's new regulation on demands for nominal and inflation-indexed bonds, led to a reduction in local inflation risk premia for the cited maturities, as the new regulation had no discernible impact on expected inflation rates.

Overall, the adjusted R^2 values of the baseline model for the 2-year inflation risk premium are 35% for the full sample period and 12% for the pre-November 2017 period. For the post-November 2017 period, the goodness of fit values range from 43% to 56% for maturities from 2 to 5 years. These values indicate that while the baseline model exhibits limited explanatory power in the first sub-period, its performance improves significantly in the second sub-period.

2.8.4 Alternative specification results

In the alternative specification, any variable found to be statistically insignificant across all regressions in Table 26 is removed. Consequently, the differential liquidity proxy, the U.S. term premium, the U.S. risk-neutral rate, the VIX, and the time

dummy variable representing the joint impact of the CBT's unconventional bond purchases and the BRSA's AR regulation are excluded from the baseline model.

Table 27 reports the estimation results for the alternative specification. The quantitative and qualitative findings are very similar to those observed in the baseline regressions in Table 26, albeit with two main differences. First, in the second sub-period, the statistical significance of the coefficients on changes in the sovereign CDS spread for the 4- and 5-year maturities is weaker, with p-values of 0.15 and 0.12, respectively, suggesting a reduced impact of changes in the sovereign credit risk on changes in the corresponding local inflation risk premia as compared to the baseline regression results. Second, there is a notable increase in the statistical significance of the coefficients related to the liraization strategy.¹⁹⁶ Specifically, the coefficient for the 2-year maturity achieves significance at the 5% level, while the coefficients for the 3-, 4-, and 5-year maturities reach 1% level. This indicates a marked improvement in the relationship between the liraization strategy and changes in local inflation risk premia relative to the baseline regression. Additionally, the coefficients imply a larger impact of the liraization strategy on the Turkish inflation risk premia across all maturities. When it is in effect, the 2-, 3-, 4-, and 5-year inflation risk premia decrease by 365 bp, 474 bp, 489 bp, and 462 bp, respectively.

2.8.5 Robustness checks

I conclude this section by exploring the robustness of the baseline regression results. In the first robustness check, I introduce the MOVE index into the baseline model as a measure of the uncertainty in the global bond markets.¹⁹⁷ The estimation results

¹⁹⁶ This is primarily attributable to the exclusion of the time dummy variable reflecting the joint impact of the CBT's unconventional bond purchases and the BRSA's AR regulation in the alternative specification.

¹⁹⁷ Recall that the MOVE index measures the implied volatility of the U.S. Treasury market.

Table 27. Alternative Specification Results for the Inflation Risk Premium

Explanatory Var.	Full Sample	First Sub-Period	Second Sub-Period			
	2-year	2-year	2-year	3-year	4-year	5-year
Δ CDS Spread	0.0036 (0.662)	-0.0017 (-0.354)	0.0055 (0.847)	0.0070 (1.141)	0.0084 (1.457)	0.0085 (1.571)
Δ Implied Vol.	0.1947** (2.024)	0.0966 (1.236)	0.2155** (2.019)	0.2487** (2.314)	0.2162** (2.140)	0.1838** (2.069)
Δ Brent	0.0729*** (3.403)	0.0263*** (3.789)	0.1094** (2.177)	0.1014** (2.344)	0.0836** (2.457)	0.0648** (2.419)
\$/TRY Surprise M.	-0.0088 (-0.195)	0.0540** (2.117)	-0.0406 (-0.507)	-0.0487 (-0.691)	-0.0402 (-0.737)	-0.0379 (-0.857)
CPI Surprise M.	0.0392*** (2.627)	0.0224*** (2.863)	0.0805 (1.551)	0.0658 (1.304)	0.0632 (1.328)	0.0585 (1.389)
Liraization Strgy.	-3.3780** (-2.476)		-3.6495** (-2.544)	-4.7438*** (-3.045)	-4.8920*** (-4.028)	-4.6192*** (-5.876)
Constant	-0.0281 (-0.349)	-0.0059 (-0.133)	-0.2064 (-0.952)	-0.1892 (-1.027)	-0.1853 (-1.141)	-0.1606 (-1.084)
Observations	152	96	56	55	55	55
Adjusted R ²	0.3553	0.1390	0.4154	0.5304	0.5638	0.5636

Note: The full sample period, spanning from October 2009 to June 2022, is divided into two sub-periods: pre- and post-November 2017. The t-statistics are shown in parentheses below each coefficient estimate and based on Newey-West heteroscedasticity and autocorrelation consistent (HAC) standard errors with three lags for the full sample period and the first sub-period, and two lags for the second sub-period. Asterisks ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively. In the regression analysis, the exchange rate and inflation surprise measures lag one month.

suggest that the main findings hold firmly, while changes in the MOVE index have no significant impact on changes in local inflation risk premia.

In the second robustness check, I employ the actual U.S. nominal yields instead of the U.S. term premia and risk-neutral rates in the regressions. The baseline regression results remain intact, along with no statistically significant impact of changes in the U.S. yields on changes in the Turkish inflation risk premia.

In the third robustness check, the exchange rate and inflation surprise measures are replaced by the 12-month-ahead inflation forecast dispersion measure from the CBT's "Survey of Market Participants".¹⁹⁸ The regression results reveal that changes in this proxy for inflation uncertainty do not contribute significantly to

¹⁹⁸ Refer to Section 1.5.6 for the computation of the 12-month-ahead forecast dispersion measure.

variations in local inflation risk premia.¹⁹⁹ Nevertheless, the findings related to other explanatory variables in the baseline regressions are largely unaltered.

As an additional robustness check, I substitute the sovereign CDS spread with the J.P. Morgan's country-specific EMBI+ spread for Turkey. The main results remain practically the same across all maturities and sample periods.

I also investigate the potential endogeneity of the CDS spread due to concerns about simultaneity and reverse causality with local inflation risk premia. To address these issues, I use the two-step IV-GMM estimation, employing changes in both the aggregate J.P. Morgan EMBI+ spread (see footnote 137) and the VIX index (see footnote 138) as instruments for changes in the Turkish CDS spread.²⁰⁰ The Kleibergen-Paap rk LM statistic confirms the relevance of the instruments, while the Cragg-Donald and Kleibergen-Paap rk Wald F statistics indicate their strength. The Hansen J statistic verifies that the instruments are valid. However, the C-test fails to reject the null hypothesis that the CDS spread is exogenous.²⁰¹ These results remain largely consistent across the maturities and sample periods. Given the lack of evidence for endogeneity in the CDS spread, the OLS results can be considered consistent and unbiased.

Similarly, I examine the potential endogeneity of the 6-month ATM USD/TRY option volatility, considering the possibility of simultaneity and reverse

¹⁹⁹ I also utilize the 24-month-ahead inflation forecast dispersion measure and reach the same conclusion.

²⁰⁰ Using daily data from June 2009 to July 2016 for a panel of 34 countries, including Turkey, Hibbert and Pavlova (2017) find that positive changes in the J.P. Morgan's EMBI+ index (or negative changes in the EMBI+ spread) lead to a decrease in the sovereign CDS spreads.

²⁰¹ For instance, for the regression of changes in the 2-year inflation risk premium over the full sample period, using changes in both the J.P. Morgan EMBI+ spread and the VIX index as instruments, the Kleibergen-Paap rk LM statistic has a p-value of 0.000, showing that the instruments are not underidentified. The Cragg-Donald F statistic is 38.512, exceeding the Stock and Yogo (2005) 10% critical value of 19.93, indicating that the instruments are sufficiently strong. The Kleibergen-Paap rk Wald F statistic is 16.035, surpassing the Staiger and Stock (1997) threshold of 10, further confirming the strength of the instruments. The Hansen J statistic has a p-value of 0.296, affirming that the instruments are not correlated with the residuals and are therefore valid. The C-test yields a p-value of 0.121, suggesting that the CDS spread is exogenous in the regression.

causality with local inflation risk premia. I repeat the two-step IV-GMM estimation, using changes in both the J.P. Morgan emerging markets currency volatility index (EM-VXY index, see footnote 140) and the VIX index as instruments for changes in the 6-month ATM option volatility. Although the test results confirm that the instruments are relevant, sufficiently strong, and valid, the C-test does not reject the null hypothesis of exogeneity of the 6-month ATM option volatility. These findings hold broadly across the maturities and sample periods, affirming that the OLS results are consistent and unbiased.

2.9 Ex ante cost-benefit analysis of issuing nominal versus inflation-indexed debt

In this section, I compare the direct issuance costs of nominal versus inflation-indexed government bonds in Turkey from an ex ante perspective.²⁰² If the terms of equation (44) are rearranged:

$$(\tilde{y}_{t,\tau}^N - \pi_{t,\tau}^e) - \tilde{y}_{t,\tau}^R = \phi_{t,\tau} + (\eta_{t,\tau}^N - \eta_{t,\tau}^R) \quad (45)$$

The first term in parentheses on the left-hand side of equation (45) represents the interest rate paid on a nominal bond minus expected inflation, while the second term represents the interest rate paid on an inflation-indexed bond. The spread between these two terms indicates an ex ante or expected issuance cost difference between nominal and inflation-indexed bonds (Ermolov, 2021).²⁰³ Put differently, it reflects the real cost advantage or disadvantage of issuing a nominal bond versus an

²⁰² I focus on an ex ante cost-benefit analysis of issuing nominal versus inflation-indexed bonds. It is important to acknowledge that actual inflation outcomes, to which inflation-indexed debt payments are linked, may vary considerably from inflation expectations made at the time of inflation-indexed bond issuances. This variation can result in significant discrepancies between ex ante and ex post analyses. Moreover, my analysis does not take into consideration certain difficult-to-quantify benefits of issuing inflation-indexed bonds. As outlined by Dudley et al. (2009), these extra benefits include optimizing risk allocation, enhancing monetary policy effectiveness, incentivizing responsible fiscal policies, and broadening the investor base. For an in-depth discussion of differences between ex ante and ex post analyses and these other benefits of inflation-indexed issuances, refer to Dudley et al. (2009).

²⁰³ This spread is actually the breakeven deviation rate, as defined previously.

inflation-indexed bond for the government on an ex ante basis (Bekaert & Ermolov, 2023). A positive difference implies that the government achieves cost reduction by issuing the inflation-indexed bond.

The decomposition of BEI rates allows for a clearer understanding of the components that determine the net expected cost or benefit of issuing inflation-indexed bonds over nominal bonds. In fact, the right-hand side of equation (45) shows that an ex ante issuance cost difference between nominal and inflation-indexed bonds is equal to the sum of the inflation risk premium and the differential liquidity premium components. This breakdown reveals that the government faces a tradeoff when deciding whether to issue nominal or inflation-indexed bonds. If the sum of the inflation risk premium and the liquidity premium embedded in nominal bond yields is larger than the liquidity premium embedded in inflation-indexed bond yields (i.e., $\phi_{t,\tau} + \eta_{t,\tau}^N > \eta_{t,\tau}^R$), issuing inflation-indexed bonds becomes more advantageous for the government (Cardozo & Christensen, 2023).²⁰⁴

In the context of Turkey, given the irrelevance of the differential liquidity premium component (i.e., $\eta_{t,\tau}^N - \eta_{t,\tau}^R \approx 0$) in the BEI rates (see Section 2.6.3), it is the sign and magnitude of the inflation risk premium component that determine whether inflation-indexed bond issuances are a more cost-effective choice compared to nominal bond issuances. As reported in Table 23a, the 2-year inflation risk premium averaged -6 bp during the first sub-period. This suggests that, on average, the direct issuance costs of nominal and inflation-indexed bonds for this specific maturity were ex ante approximately equal between October 2009 and October

²⁰⁴ Rephrasing the same condition, if the inflation risk premium embedded in nominal bond yields exceeds the difference in the liquidity premiums between the inflation-indexed and nominal bonds (i.e., $\phi_{t,\tau} > \eta_{t,\tau}^R - \eta_{t,\tau}^N$), then the cost of issuing inflation-indexed bonds will be lower than that of nominal bonds on an ex ante basis.

2017.²⁰⁵ Turning to the second sub-period, Table 23b demonstrates that the average term structure of inflation risk premia exhibits an upward slope, with mean values ranging from 332 bp for the 2-year maturity to 377 bp for the 5-year maturity. This indicates that, on average and from the ex ante perspective, inflation-indexed bonds became substantially cheaper to issue vis-à-vis nominal bonds in the post-November 2017 period, with the relative cost advantage growing with increasing maturity. This phenomenon is consistent with the economic theory in the sense that, during periods of high and volatile inflation, such as that experienced in Turkey over this time frame, investors prioritize preserving the purchasing power of their investments. Consequently, the growing demand for inflation protection prompts investors to accept lower yields on inflation-indexed bonds, ultimately leading to cheaper issuance costs relative to nominal bonds. During the second sub-period, in an effort to optimize the sovereign debt financing strategy, the Turkish Treasury effectively capitalized on this ex ante cost advantage by issuing more inflation-indexed bonds compared to nominal bonds. Remarkably, throughout this period, the outstanding notional amount of the inflation-indexed bonds surged from TRY 119.6 billion to TRY 391.5 billion, while that of the nominal (fixed-coupon) bonds increased from TRY 343.7 billion to TRY 499.1 billion (see the top right panel of Figure 18).²⁰⁶

2.10 Summary and conclusion

In this study, I employ the model-free approach to decompose the Turkish BEI rates into expected inflation, inflation risk premium, and differential liquidity premium

²⁰⁵ However, over the course of this first sub-period, the direct ex ante issuance cost difference between these two types of bonds also demonstrated notable time series variation, reflecting fluctuations in the 2-year inflation risk premium, which ranged from -159 bp and 145 bp.

²⁰⁶ Over the same period, the share of inflation-indexed bonds in the Treasury's TRY-denominated debt stock rose from 22.5% to 34.6%, while that of nominal (fixed-coupon) bonds declined from 64.8% to 44.1%.

components across various maturities. To capture expected inflation, I rely on the inflation forecasts provided by the professional and representative agents in the CBT's "Survey of Market Participants", considering the sophistication of their methodology in forming expectations. To estimate the impact of liquidity differences between the Turkish nominal and inflation-indexed bonds on the BEI and breakeven deviation rates, I run a series of regressions. Although the inflation-indexed bonds in Turkey seem to be less liquid than their nominal counterparts, the regression results indicate that the differential liquidity premium between these two types of bonds is not priced in the local bond market. This phenomenon can be explained by the fact that the inflation-indexed bonds are primarily held by buy-and-hold investors, such as local banks and pension funds, who place greater importance on the inflation-hedging benefits of these bonds over short-term liquidity concerns. Consequently, these investors do not require additional compensation for the lower market liquidity of inflation-indexed bonds compared to nominal bonds. Given the irrelevance of differential liquidity premium component in the BEI rate, I then compute the inflation risk premium component as the difference between the BEI rate and the corresponding survey-based average expected inflation.

This study pioneers a comprehensive exploration of the determinants of inflation risk premium within the context of Turkey. The time-series regressions provide several novel and key findings. Notably, in assessing the influences of global risk factors on the Turkish inflation risk premia, I find that there are no spillover impacts from the U.S. bond yields and their components on local inflation risk premia over the entire sample period and two sub-periods. Additionally, changes in the VIX index also show no significant relationship with changes in the country's inflation risk premia. On the other hand, the fluctuations in the Brent crude oil prices

strongly impact the variations in local inflation risk premia across all maturities and specified time periods.

Another important finding is the negative impact of the CBT's liraization strategy on changes in local inflation risk premia. The new regulation, introduced in June 2022 as part of this strategy, increased the demand for nominal bonds and decreased the demand for inflation-indexed bonds, resulting in a significant reduction on local inflation risk premia. This finding adds to the limited body of literature addressing the effects of financial sector regulations on bond yield components, particularly on inflation risk premia.

Regarding the impacts of inflation uncertainty measures, while one-month lagged exchange rate and inflation surprises exhibit statistically significant positive effects on changes on local inflation risk premia during the first sub-period, they lose their statistical significance over the second sub-period. Importantly, as discussed in the robustness checks section, there is no observed relationship between changes in the survey-based inflation forecast dispersion measures and changes in the Turkish inflation risk premia over time.

The regression results also suggest that fluctuations in the exchange rate volatility play a significant role in shaping the dynamics of local inflation risk premia in the post-November 2017 period. However, changes in the sovereign credit risk exert a relatively subdued influence on these dynamics during the same period.

Lastly, I conduct the ex ante cost-benefit analysis of issuing nominal versus inflation-indexed bonds from the standpoint of the Turkish Treasury. On average, during the first sub-period, the direct ex ante issuance costs of both types of bonds were roughly equal. Nonetheless, in the second sub-period, characterized by persistent and erratic inflation dynamics, the inflation-indexed bonds proved to be a

more favorable issuance option compared to nominal bonds, both on average and from the ex ante perspective.

Overall, this study presents a straightforward method for estimating the inflation risk premium in Turkey. The findings on its determinants, together with the results of ex ante cost-benefit analysis, offer policymakers, fixed-income analysts, and investors valuable insights to guide their decision-making processes.



REFERENCES

- Adrian, T., Crump, R. K., & Moench, E. (2013). Pricing the term structure with linear regressions. *Journal of Financial Economics*, 110(1), 110–138. <https://doi.org/10.1016/j.jfineco.2013.04.009>
- Aguilar-Argaez, A., Diego-Fernández, M., Elizondo, R., & Roldán-Peña, J. (2022). Term premium dynamics and its determinants: the Mexican case. *BIS Working Papers*, No. 993.
- Albagli, E., Ceballos, L., Claro, S., & Romero, D. (2019). Channels of US monetary policy spillovers to international bond markets. *Journal of Financial Economics*, 134(2), 447–473. <https://doi.org/10.1016/j.jfineco.2019.04.007>
- Altavilla, C., Carboni, G., & Motto, R. (2015). Asset purchase programmes and financial markets: Lessons from the euro area. *ECB Working Paper*, No. 1864.
- Amihud, Y., & Mendelson, H. (2006). Stock and bond liquidity and its effect on prices and financial policies. *Financial Markets and Portfolio Management*, 20(1), 19–32. <https://doi.org/10.1007/s11408-006-0001-y>
- Amihud, Y., Mendelson, H., & Pedersen, L. H. (2006). Liquidity and asset prices. *Foundations and Trends® in Finance*, 1(4), 269–364. <https://doi.org/10.1561/05000000003>
- Andrade, P., Breckenfelder, J., De Fiore, F., Karadi, P., & Tristani, O. (2016). The ECB's asset purchase programme: An early assessment. *ECB Working Paper*, No. 1956.
- Ang, A., & Piazzesi, M. (2003). A no-arbitrage vector autoregression of term structure dynamics with macroeconomic and latent variables. *Journal of Monetary Economics*, 50(4), 745–787. [https://doi.org/10.1016/s0304-3932\(03\)00032-1](https://doi.org/10.1016/s0304-3932(03)00032-1)
- Bank for International Settlements (2005). Zero-coupon yield curves: technical documentation. *BIS Papers*, No. 25.
- Başçı, E., & Kara, H. (2011). Financial stability and monetary policy. *Central Bank of the Republic of Turkey Working Papers*, No. 11/08.
- Bauer, M. D., & Hamilton, J. D. (2018). Robust bond risk premia. *The Review of Financial Studies*, 31(2), 399–448. <https://doi.org/10.1093/rfs/hhx096>
- Bauer, M., & Rudebusch, G. D. (2014). The signaling channel for Federal Reserve bond purchases. *International Journal of Central Banking*, 10(3), 233–289.
- Bauer, M. D., & Rudebusch, G. D. (2017). Resolving the spanning puzzle in macro-finance term structure models. *Review of Finance*, 21(2), 511–553. <https://doi.org/10.1093/rof/rfw044>

- Bauer, M. D., Rudebusch, G. D., & Wu, J. C. (2012). Correcting estimation bias in dynamic term structure models. *Journal of Business & Economic Statistics*, 30(3), 454–467. <https://doi.org/10.1080/07350015.2012.693855>
- Bauer, M. D., Rudebusch, G. D., & Wu, J. C. (2014). Term premia and inflation uncertainty: Empirical evidence from an international panel dataset: Comment. *American Economic Review*, 104(1), 323–337. <https://doi.org/10.1257/aer.104.1.323>
- Beauregard, R., Christensen, J. H. E., Fischer, E., & Zhu, S. (2021). Inflation expectations and risk premia in emerging bond markets: Evidence from Mexico. *FRB of San Francisco Working Papers*, 2021-08. <https://doi.org/10.24148/wp2021-08>
- Bekaert, G., Engstrom, E., & Ermolov, A. (2021). Macro risks and the term structure of interest rates. *Journal of Financial Economics*, 141(2), 479–504. <https://doi.org/10.1016/j.jfineco.2021.03.011>
- Bekaert, G., & Ermolov, A. (2023). International yield comovements. *Journal of Financial and Quantitative Analysis*, 58(1), 250–288. <https://doi.org/10.1017/S0022109022000515>
- Bekaert, G., & Hodrick, R. J. (2001). Expectations hypotheses tests. *The Journal of Finance*, 56(4), 1357–1394. <https://doi.org/10.1111/0022-1082.00371>
- Bekaert, G., Hodrick, R. J., & Marshall, D. A. (1997). On biases in tests of the expectations hypothesis of the term structure of interest rates. *Journal of Financial Economics*, 44(3), 309–348. [https://doi.org/10.1016/s0304-405x\(97\)00007-x](https://doi.org/10.1016/s0304-405x(97)00007-x)
- Bekaert, G., Hodrick, R. J., & Marshall, D. A. (2001). Peso problem explanations for term structure anomalies. *Journal of Monetary Economics*, 48(2), 241–270. [https://doi.org/10.1016/s0304-3932\(01\)00075-7](https://doi.org/10.1016/s0304-3932(01)00075-7)
- Bekaert, G., Hoerova, M., & Lo Duca, M. (2013). Risk, uncertainty and monetary policy. *Journal of Monetary Economics*, 60(7), 771–788. <https://doi.org/10.1016/j.jmoneco.2013.06.003>
- Bekker, P. A., & Bouwman, K. E. (2009). Arbitrage smoothing in fitting a sequence of yield curves. *International Journal of Theoretical and Applied Finance*, 12(05), 577–588. <https://doi.org/10.1142/S0219024909005373>
- Beltran, D. O., Kretchmer, M., Marquez, J., & Thomas, C. P. (2013). Foreign holdings of US Treasuries and US Treasury yields. *Journal of International Money and Finance*, 32, 1120–1143. <https://doi.org/10.1016/j.jimonfin.2012.09.005>
- Bernanke, B.S. (2015, April 13). Why are interest rates so low, part 4: Term premiums. *Brookings*. <https://www.brookings.edu/articles/why-are-interest-rates-so-low-part-4-term-premiums/>

- Bernanke, B., Reinhart, V., & Sack, B. (2004). Monetary policy alternatives at the zero bound: An empirical assessment. *Brookings Papers on Economic Activity*, 2004(2), 1–100. <https://doi.org/10.1353/eca.2005.0002>
- Bolder, D., & Stréliski, D. (1999). Yield curve modelling at the Bank of Canada. *Bank of Canada Technical Report*, No. 84.
- Broyden, C. G. (1970). The convergence of a class of double-rank minimization algorithms 1. general considerations. *IMA Journal of Applied Mathematics*, 6(1), 76–90. <https://doi.org/10.1093/imamat/6.1.76>
- Camba-Méndez, G., & Werner, T. (2017). The inflation risk premium in the post-Lehman period. *ECB Working Paper*, No. 2033.
- Campbell, J. Y. (1986). A defense of traditional hypotheses about the term structure of interest rates. *The Journal of Finance*, 41(1), 183–193. <https://doi.org/10.1111/j.1540-6261.1986.tb04498.x>
- Campbell, J. Y., & Shiller, R. J. (1991). Yield spreads and interest rate movements: A bird's eye view. *The Review of Economic Studies*, 58(3), 495–514. <https://doi.org/10.2307/2298008>
- Cardozo, C. H. R., & Christensen, J. H. E. (2023). The Benefit of inflation-indexed debt: Evidence from an emerging bond market. *FRB of San Francisco Working Papers*, 2023-04. <https://doi.org/10.24148/wp2023-04>
- Carvalho, F. A., & Minella, A. (2012). Survey forecasts in Brazil: A prismatic assessment of epidemiology, performance, and determinants. *Journal of International Money and Finance*, 31(6), 1371–1391. <https://doi.org/10.1016/j.jimonfin.2012.02.006>
- Ceballos, L., & Romero, D. (2016). Decomposing long-term interest rates: An international comparison. *The Journal of Fixed Income*, 26(1), 61–73. <https://doi.org/10.3905/jfi.2016.26.1.061>
- Chari, A., Stedman, K. D., & Lundblad, C. (2017). Taper Tantrums: QE, its aftermath and emerging market capital flows. <https://doi.org/10.3386/w23474>
- Chen, R. R., & Scott, L. (1992). Pricing interest rate options in a two-factor Cox–Ingersoll–Ross model of the term structure. *Review of Financial Studies*, 5(4), 613–636. <https://doi.org/10.1093/rfs/5.4.613>
- Chordia, T., Sarkar, A., & Subrahmanyam, A. (2004). An empirical analysis of stock and bond market liquidity. *Review of Financial Studies*, 18(1), 85–129. <https://doi.org/10.1093/rfs/hhi010>
- Christensen, J. H. E., & Gillan, J. M. (2012). Could the US Treasury benefit from issuing more TIPS? *FRB of San Francisco Working Papers*, 2011-16.
- Christensen, J. H.E., & Rudebusch, G. D. (2012). The response of interest rates to US and UK quantitative easing. *The Economic Journal*, 122(564), F385–414. <https://doi.org/10.1111/j.1468-0297.2012.02554.x>

- Christensen, J. H.E., & Rudebusch, G. D. (2019). A new normal for interest rates? Evidence from inflation-indexed debt. *The Review of Economics and Statistics*, 101(5), 933–949. https://doi.org/10.1162/rest_a_00821
- Cieslak, A., & Povala, P. (2015). Expected returns in Treasury bonds. *Review of Financial Studies*, 28(10), 2859–2901. <https://doi.org/10.1093/rfs/hhv032>
- Cochrane, J. H., & Piazzesi, M. (2005). Bond risk premia. *American Economic Review*, 95(1), 138–160. <https://doi.org/10.1257/0002828053828581>
- Cochrane, J. H., & Piazzesi, M. (2009). Decomposing the yield curve. In *AFA 2010 Atlanta Meetings Paper*.
- Cohen, B. H., Hördahl, P., & Xia, F. D. (2018). Term premia: models and some stylised facts. *BIS Quarterly Review*, September, 79–91.
- Cooper, I., & Priestley, R. (2009). Time-varying risk premiums and the output gap. *Review of Financial Studies*, 22(7), 2801–2833. <https://doi.org/10.1093/rfs/hhn087>
- Coroneo, L., Nyholm, K., & Vidova-Koleva, R. (2008). How arbitrage-free is the Nelson–Siegel model? *ECB Working Paper*, No. 874.
- Corradin, S., Grimm, N., & Schwaab, B. (2021). Euro area sovereign bond risk premia during the COVID-19 pandemic. *ECB Working Paper*, No. 2561.
- Cox, J. C., Ingersoll, J. E., & Ross, S. A. (1985). A theory of the term structure of interest rates. *Econometrica*, 53(2), 385–407. <https://doi.org/10.2307/1911242>
- Crump, R. K., Eusepi, S., & Moench, E. (2016). The term structure of expectations and bond yields. *Federal Reserve Bank of New York Staff Report*, No. 775.
- Csonto, B., & Tovar, C.E. (2017). Uphill capital flows and the international monetary system. *IMF Working Papers*, 17(174). <https://doi.org/10.5089/9781484311424.001>
- Culbertson, J. M. (1957). The term structure of interest rates. *The Quarterly Journal of Economics*, 71(4), 485–517.
- Curcucu, S. E., Kamin, S. B., Li, C., & Rodriguez, M. (2023). International spillovers of monetary policy: Conventional policy vs. quantitative easing. *International Journal of Central Banking*, 19(1), 111–158.
- Cuthbertson, K., & Nitzsche, D. (2004). *Quantitative financial economics: Stocks, bonds and foreign exchange* (2nd ed.). Wiley.
- Çepni, O., & Küçüksaraç, D. (2017). Optimal mix of the extended Nelson Siegel model for Turkish sovereign yield curve. *Economics Bulletin*, 37(2), 1133–1142.

- Dai, Q., & Singleton, K. J. (2000). Specification analysis of affine term structure models. *The Journal of Finance*, 55(5), 1943–1978.
<https://doi.org/10.1111/0022-1082.00278>
- D’Amico, S., English, W. B., López-Salido, J. D., & Nelson, E. (2012). The Federal Reserve’s large-scale asset purchase programs: Rationale and effects. *Finance and Economics Discussion Series*, 2012(85), 1–56.
<https://doi.org/10.17016/feds.2012.85>
- D’Amico, S., Kim, D. H., & Wei, M. (2018). Tips from TIPS: The informational content of Treasury Inflation-Protected Security prices. *Journal of Financial and Quantitative Analysis*, 53(1), 395–436.
<https://doi.org/10.1017/s0022109017000916>
- De Lange, P. E., Rissstad, M., & Westgaard, S. (2022). Term premia in Norwegian government bond yields. *Beta*, 36(1), 1–21.
- De Santis, R. A. (2019). Redenomination risk. *Journal of Money, Credit and Banking*, 51(8), 2173–2206. <https://doi.org/10.1111/jmcb.12582>
- Domowitz, I., Glen, J., & Madhavan, A. (1998). Country and currency risk premia in an emerging market. *The Journal of Financial and Quantitative Analysis*, 33(2), 189–216. <https://doi.org/10.2307/2331307>
- Du, W., & Schreger, J. (2016). Local Currency Sovereign Risk. *The Journal of Finance*, 71(3), 1027–1070. <https://doi.org/10.1111/jofi.12389>
- Dudley, W., Roush, J. , & Ezer, M.S. (2009). The case for TIPS: An examination of the costs and benefits. *Federal Reserve Bank of New York Economic Policy Review*, 15(1), 1–17.
- Duffee, G. R. (2011). Sharpe ratios in term structure models. *Johns Hopkins University, Department of Economics Working Papers*, No. 575.
- Duffee, G. R. (2013). Bond pricing and the macroeconomy. *Handbook of the Economics of Finance*, 2, 907–967. <https://doi.org/10.1016/b978-0-44-459406-8.00013-5>
- Duffie, D., & Kan, R. (1996). A yield-factor model of interest rates. *Mathematical Finance*, 6(4), 379–406. <https://doi.org/10.1111/j.1467-9965.1996.tb00123.x>
- Duran, M., & Gülşen, E. (2013). Estimating inflation compensation for Turkey using yield curves. *Economic Modelling*, 32, 592–601.
<https://doi.org/10.1016/j.econmod.2013.02.036>
- Ermolov, A. (2021). When and where is it cheaper to issue inflation-linked debt? The Review of Asset Pricing Studies, 11(3), 610–653.
<https://doi.org/10.1093/rapstu/raab016>
- Fama, E. F., & Bliss, R. R. (1987). The information in long-maturity forward rates. *American Economic Review*, 77(4), 680–692.

- Fender, I., Hayo, B., & Neuenkirch, M. (2012). Daily pricing of emerging market sovereign CDS before and during the global financial crisis. *Journal of Banking & Finance*, 36(10), 2786–2794.
- Fletcher, R. (1970). A new approach to variable metric algorithms. *The Computer Journal*, 13(3), 317–322. <https://doi.org/10.1093/comjnl/13.3.317>
- Fontaine, J.-S., & Garcia, R. (2012). Bond Liquidity Premia. *Review of Financial Studies*, 25(4), 1207–1254. <https://doi.org/10.1093/rfs/hhr132>
- Fratzscher, M., Lo Duca, M., & Straub, R. (2018). On the international spillovers of US quantitative easing. *The Economic Journal*, 128(608), 330–377. <https://doi.org/10.1111/eoj.12435>
- Froot, K. A. (1989). New hope for the expectations hypothesis of the term structure of interest rates. *The Journal of Finance*, 44(2), 283–305. <https://doi.org/10.1111/j.1540-6261.1989.tb05058.x>
- Gadanecz, B., Miyajima, K., & Shu, C. (2018). Emerging market local currency sovereign bond yields: The role of exchange rate risk. *International Review of Economics & Finance*, 57, 371–401. <https://doi.org/10.1016/j.iref.2018.02.004>
- Gagnon, J., Raskin, M., Remache, J., & Sack, B. P. (2010). Large-scale asset purchases by the Federal Reserve: Did they work? *Federal Reserve Bank of New York Staff Report*, No. 441.
- Gertler, M., & Karadi, P. (2015). Monetary policy surprises, credit costs, and economic activity. *American Economic Journal: Macroeconomics*, 7(1), 44–76. <https://doi.org/10.1257/mac.20130329>
- Gilli, M., Große, S., & Schumann, E. (2010). Calibrating the Nelson-Siegel-Svensson model. *COMISEF Working Papers*, No. 31.
- Gilli, M., & Schumann, E. (2010). A note on ‘good starting values’ in numerical optimisation. *COMISEF Working Papers*, No. 44.
- Gimeno, R., & Nave, J. M. (2006). Genetic algorithm estimation of interest rate term structure. *Banco de España Research Paper*, No. 0634.
- Goldfarb, D. (1970). A family of variable-metric methods derived by variational means. *Mathematics of Computation*, 24(109), 23–26. <https://doi.org/10.1090/s0025-5718-1970-0258249-6>
- Goodfriend, M. (1998). Using the term structure of interest rates for monetary policy. *FRB Richmond Economic Quarterly*, 84(3), 13–30.
- Greenwood, R., & Vayanos, D. (2010). Price pressure in the government bond market. *American Economic Review*, 100(2), 585–590. <https://doi.org/10.1257/aer.100.2.585>

- Gregory Mankiw, N., & Shapiro, M. D. (1986). Do we reject too often? *Economics Letters*, 20(2), 139–145. [https://doi.org/10.1016/0165-1765\(86\)90161-8](https://doi.org/10.1016/0165-1765(86)90161-8)
- Grishchenko, O. V., & Huang, J.-Z. (2013). The inflation risk premium: Evidence from the TIPS market. *The Journal of Fixed Income*, 22(4), 5–30. <https://doi.org/10.3905/jfi.2013.22.4.005>
- Güler, M. H., Keleş, G., & Polat, T. (2017). An empirical decomposition of the liquidity premium in breakeven inflation rates. *The Quarterly Review of Economics and Finance*, 63, 185–192. <https://doi.org/10.1016/j.qref.2016.04.002>
- Gülşen, E., & Kara, H. (2019). Measuring inflation uncertainty in Turkey. *Central Bank Review*, 19(2), 33–43. <https://doi.org/10.1016/j.cbrev.2019.06.003>
- Gülşen, E., & Kara, H. (2021). Policy performance and the behavior of inflation expectations. *International Journal of Central Banking*, 17(70), 179–224.
- Gürkaynak, R. S., Sack, B., & Wright, J. H. (2007). The U.S. Treasury yield curve: 1961 to the present. *Journal of Monetary Economics*, 54(8), 2291–2304. <https://doi.org/10.1016/j.jmoneco.2007.06.029>
- Gürkaynak, R. S., Sack, B., & Wright, J. H. (2008). The TIPS yield curve and inflation compensation. *Finance and Economics Discussion Series*, 2008(05), 1–40. <https://doi.org/10.17016/feds.2008.05>
- Gürkaynak, R. S., Sack, B., & Wright, J. H. (2010). The TIPS yield curve and inflation compensation. *American Economic Journal: Macroeconomics*, 2(1), 70–92. <https://doi.org/10.1257/mac.2.1.70>
- Gürkaynak, R. S., & Wright, J. H. (2012). Macroeconomics and the term structure. *Journal of Economic Literature*, 50(2), 331–367. <https://doi.org/10.1257/jel.50.2.331>
- Hamilton, J. D., & Wu, J. C. (2012). Identification and estimation of Gaussian affine term structure models. *Journal of Econometrics*, 168(2), 315–331. <https://doi.org/10.1016/j.jeconom.2012.01.035>
- Hardouvelis, G. A. (1994). The term structure spread and future changes in long and short rates in the G7 countries. *Journal of Monetary Economics*, 33(2), 255–283. [https://doi.org/10.1016/0304-3932\(94\)90003-5](https://doi.org/10.1016/0304-3932(94)90003-5)
- Hibbert, A. M., & Pavlova, I. (2017). The drivers of sovereign CDS spread changes: Local versus global factors. *Financial Review*, 52(3), 435–457. <https://doi.org/10.1111/fire.12140>
- Ho, T. S. Y. (1992). Key rate durations. *The Journal of Fixed Income*, 2(2), 29–44. <https://doi.org/10.3905/jfi.1992.408049>
- Hofmann, B., & Takáts, E. (2015). International monetary spillovers. *BIS Quarterly Review*, September, 105–118.

- Hördahl, P. (2008). The inflation risk premium in the term structure of interest rates. *BIS Quarterly Review*, 3, 23–38.
- Hördahl, P., Sobrun, J., & Turner, P. (2016). Low long-term interest rates as a global phenomenon. *BIS Working Papers*, No. 574.
- Hördahl, P., & Tristani, O. (2014). Inflation risk premia in the Euro Area and the United States. *International Journal of Central Banking*, 10(3), 1–47.
- Hu, G. X., Pan, J., & Wang, J. (2013). Noise as information for illiquidity. *The Journal of Finance*, 68(6), 2341–2382. <https://doi.org/10.1111/jofi.12083>
- Joslin, S., Priebsch, M., & Singleton, K. J. (2014). Risk premiums in dynamic term structure models with unspanned macro risks. *The Journal of Finance*, 69(3), 1197–1233. <https://doi.org/10.1111/jofi.12131>
- Joslin, S., Singleton, K. J., & Zhu, H. (2011). A new perspective on Gaussian dynamic term structure models. *Review of Financial Studies*, 24(3), 926–970. <https://doi.org/10.1093/rfs/hhq128>
- Joyce, M. A., Lasasosa, A., Stevens, I., & Tong, M. (2011). The financial market impact of quantitative easing in the United Kingdom. *International Journal of Central Banking*, 7(3), 113–161.
- Joyce, M.A., Miles, D., Scott, A., & Vayanos, D. (2012). Quantitative easing and unconventional monetary policy – an introduction. *The Economic Journal*, 122(564), 271–288. <https://doi.org/10.1111/j.1468-0297.2012.02551.x>
- Joyce, M. A., & Tong, M. (2012). QE and the gilt market: a disaggregated analysis. *The Economic Journal*, 122(564), 348–384. <https://doi.org/10.1111/j.1468-0297.2012.02552.x>
- Kanlı, İ. B., Küçüksaraç, D., & Özel, Ö. (2013). Yield curve estimation for corporate bonds in Turkey. *Central Bank of the Republic of Turkey Working Papers*, No. 13/26.
- Kara, H. (2013). Monetary policy after the global crisis. *Atlantic Economic Journal*, 41(1), 51–74. <https://doi.org/10.1007/s11293-012-9358-7>
- Kara, H. (2021). Monetary policy in the Covid era: The Turkish experience. In B. English, K. Forbes, & A. Ubide (Eds.), *Monetary Policy and Central Banking in the Covid Era* (pp. 293–310). London, UK: Center for Economic Policy Research.
- Kim, D. H., & Orphanides, A. (2007). The bond market term premium: What is it, and how can we measure it? *BIS Quarterly Review*, June, 27–40.
- Kim, D. H., & Orphanides, A. (2012). Term structure estimation with survey data on interest rate forecasts. *Journal of Financial and Quantitative Analysis*, 47(1), 241–272. <https://doi.org/10.1017/S0022109011000627>

- Kim, D. H., & Wright, J. H. (2005). An arbitrage-free three-factor term structure model and the recent behavior of long-term yields and distant-horizon forward rates. *Finance and Economics Discussion Series*, 2005(33), 1–21. <https://doi.org/10.17016/feds.2005.33>
- Klemm, A., Meier, A., & Sosa, S. (2014). Spillovers to Latin America from the normalization of U.S. monetary policy. In D.M. Iakova, L. M. Cubeddu, G. Adler, & S. Sosa (Eds.), *Latin America New Challenges to Growth and Stability* (pp. 213–229). Washington, DC: International Monetary Fund. <https://doi.org/10.5089/9781498328166.071>
- Kocsis, Z., & Monostori, Z. (2016). The role of country-specific fundamentals in sovereign CDS spreads: Eastern European experiences. *Emerging Markets Review*, 27, 140–168. <https://doi.org/10.1016/j.ememar.2016.05.003>
- Koç, Ü., Ögünç, F., & Özmen, M. U. (2021). The Role of expectations in the inflation process in Turkey: Have the dynamics changed recently? *Central Bank of the Republic of Turkey Working Papers*, No. 21/02.
- Kolasa, M., & Wesolowski, G. (2020). International spillovers of quantitative easing. *Journal of International Economics*, 126, 1–32. <https://doi.org/10.1016/j.jinteco.2020.103330>
- Krishnamurthy, A., Nagel, S., & Vissing-Jorgensen, A. (2018). ECB policies involving government bond purchases: Impact and channels. *Review of Finance*, 22(1), 1–44. <https://doi.org/10.1093/rof/rfx053>
- Krishnamurthy, A., & Vissing-Jorgensen, A. (2011). The effects of quantitative easing on interest rates: Channels and implications for policy. *Brookings Papers on Economic Activity*, 2011(2), 215–265.
- Krishnamurthy, A., & Vissing-Jorgensen, A. (2013). The ins and outs of LSAPs. *Kansas City Federal Reserve Symposium on Global Dimensions of Unconventional Monetary Policy*, 57–111.
- Langetieg, T. C. (1980). A multivariate model of the term structure. *The Journal of Finance*, 35(1), 71–97. <https://doi.org/10.2307/2327182>
- Lasdon, L. S., Waren, A. D., Jain, A., & Ratner, M. (1978). Design and testing of a generalized reduced gradient code for nonlinear programming. *ACM Transactions on Mathematical Software (TOMS)*, 4(1), 34–50. <https://doi.org/10.1145/355769.355773>
- Lemke, W., & Werner, T. (2020). Dissecting long-term Bund yields in the run-up to the ECB's public sector purchase programme. *Journal of Banking & Finance*, 111, 105682. <https://doi.org/10.1016/j.jbankfin.2019.105682>
- Li, C., & Wei, M. (2012). Term structure modelling with supply factors and the Federal Reserve's large scale asset purchase programs. *Finance and Economics Discussion Series*, 2012(37), 1–33. <https://doi.org/10.17016/feds.2012.37>

- Litterman, R. B., & Scheinkman, J. (1991). Common factors affecting bond returns. *The Journal of Fixed Income*, 1(1), 54–61. <https://doi.org/10.3905/jfi.1991.692347>
- Liu, W. (2006). A liquidity-augmented capital asset pricing model. *Journal of Financial Economics*, 82(3), 631–671. <https://doi.org/10.1016/j.jfineco.2005.10.001>
- Liu, D. C., & Nocedal, J. (1989). On the limited memory BFGS method for large scale optimization. *Mathematical programming*, 45(1-3), 503–528. <https://doi.org/10.1007/bf01589116>
- Longstaff, F. A., Pan, J., Pedersen, L. H., & Singleton, K. J. (2011). How sovereign is sovereign credit risk? *American Economic Journal: Macroeconomics*, 3(2), 75–103. <https://doi.org/10.1257/mac.3.2.75>
- Longstaff, F. A., & Schwartz, E. S. (1992). Interest rate volatility and the term structure: A two-factor general equilibrium model. *The Journal of Finance*, 47(4), 1259–1282. <https://doi.org/10.1111/j.1540-6261.1992.tb04657.x>
- Ludvigson, S. C., & Ng, S. (2009). Macro factors in bond risk premia. *Review of Financial Studies*, 22(12), 5027–5067. <https://doi.org/10.1093/rfs/hhp081>
- Malik, S., & Meldrum, A. (2016). Evaluating the robustness of UK term structure decompositions using linear regression methods. *Journal of Banking & Finance*, 67, 85–102. <https://doi.org/10.1016/j.jbankfin.2016.02.006>
- Mankiw, N. G., & Shapiro, M. D. (1986). Do we reject too often?: Small sample properties of tests of rational expectations models. *Economics Letters*, 20(2), 139–145. [https://doi.org/10.1016/0165-1765\(86\)90161-8](https://doi.org/10.1016/0165-1765(86)90161-8)
- Manousopoulos, P., & Michalopoulos, M. (2009). Comparison of non-linear optimization algorithms for yield curve estimation. *European Journal of Operational Research*, 192(2), 594–602. <https://doi.org/10.1016/j.ejor.2007.09.017>
- McLaren, N., Banerjee, R. N., & Latto, D. (2014). Using changes in auction maturity sectors to help identify the impact of QE on gilt yields. *The Economic Journal*, 124(576), 453–479. <https://doi.org/10.1111/econj.12141>
- Mishra, P., Moriyama, K., N'Diaye, P.M., & Nguyen, L. (2014). Impact of Fed tapering announcements on emerging markets. *IMF Working Papers*, 14(109). <https://doi.org/10.5089/9781498361484.001>
- Miyajima, K., Mohanty, M. S., & Yetman, J. (2014). Spillovers of US unconventional monetary policy to Asia: the role of long-term interest rates. *BIS Working Papers*, No. 478.
- Modigliani, F., & Sutch, R. (1966). Innovations in interest rate policy. *The American Economic Review*, 56(1/2), 178–197.

- Monfort, A., & Renne, J.-P. (2013). Decomposing Euro-area sovereign spreads: Credit and liquidity risks. *Review of Finance*, 18(6), 2103–2151. <https://doi.org/10.1093/rof/rft049>
- Montes, G. C., & Curi, A. (2017). Disagreement in expectations about public debt, monetary policy credibility and inflation risk premium. *Journal of Economics and Business*, 93, 46–61. <https://doi.org/10.1016/j.jeconbus.2017.06.003>
- Naudon, A., & Yany, A. (2016). The impact of the US term premium on emerging markets. In L.I. Jácome, Y. Carriere-Swallow, H. Faruquee, & K. Srinivasan (Eds.), *Challenges for Central Banking: Perspectives for Latin America* (pp. 109–130). Washington, DC: International Monetary Fund.
- Nawalkha, S. K., Soto, G. M., & Beliaeva, N. A. (2005). *Interest rate risk modeling*. John Wiley & Sons.
- Nelson, C. R., & Siegel, A. F. (1987). Parsimonious modeling of yield curves. *Journal of Business*, 60(4), 473–489.
- Ng, S., & Perron, P. (2001). Lag length selection and the construction of unit root tests with good size and power. *Econometrica*, 69(6), 1519–1554.
- Nymand-Andersen, P. (2018). *Yield curve modelling and a conceptual framework for estimating yield curves: Evidence from the European Central Bank's yield curves*. ECB Statistics Paper, No. 27.
- Obstfeld, M. (2015). Trilemmas and trade-offs: living with financial globalisation. *BIS Working Papers*, No. 480.
- Özbek, İ., & Talaslı, İ. (2020). Term premium in emerging market sovereign yields: Role of common and country specific factors. *Central Bank Review*, 20(4), 169–182. <https://doi.org/10.1016/j.cbrev.2020.09.003>
- Pan, J., & Singleton, K. J. (2008). Default and recovery implicit in the term structure of sovereign CDS spreads. *The Journal of Finance*, 63(5), 2345–2384. <https://doi.org/10.1111/j.1540-6261.2008.01399.x>
- Patra, M. D., & Ray, P. (2010). Inflation Expectations and Monetary Policy in India, *IMF Working Papers*, 2010(84). <https://doi.org/10.5089/9781451982640.001.A001>
- Pericoli, M., & Taboga, M. (2008). Canonical term-structure models with observable factors and the dynamics of bond risk premia. *Journal of Money, Credit and Banking*, 40(7), 1471–1488. <https://doi.org/10.1111/j.1538-4616.2008.00167.x>
- Pflueger, C. E., & Viceira, L. M. (2011). An empirical decomposition of risk and liquidity in nominal and inflation-indexed government bonds. *NBER Working Papers*, 16892.

- Price, K., Storn, R. M., & Lampinen, J. A. (2005). *Differential evolution: A practical approach to global optimization*. Springer-Verlag. <https://doi.org/10.1007/3-540-31306-0>
- Remolona, E. M., Scatigna, M., & Wu, E. (2007). Interpreting sovereign spreads. *BIS Quarterly Review*, March, 27–39.
- Rudebusch, G. D. (1998). Monetary policy and the term structure of interest rates: An overview of some recent research. In R. Rovelli & I. Angeloni (Eds.), *Monetary Policy and the Term Structure of Interest Rates* (pp. 263–271). New York: St. Martin's Press.
- Sarno, L., Thornton, D. L., & Valente, G. (2007). The empirical failure of the expectations hypothesis of the term structure of bond yields. *Journal of Financial and Quantitative Analysis*, 42(1), 81–100. <https://doi.org/10.1017/s0022109000002192>
- Sarr, A., & Lybek, T. (2002). Measuring liquidity in financial markets. *IMF Working Papers*, 2002(232), 1–61. <https://doi.org/10.5089/9781451875577.001>
- Shanno, D. F. (1970). Conditioning of quasi-Newton methods for function minimization. *Mathematics of Computation*, 24(111), 647–656. <https://doi.org/10.1090/s0025-5718-1970-0274029-x>
- Shen, P. (1998). How important is the inflation risk premium? *Economic Review*, 83(4), 35–47.
- Shen, P. (2006). Liquidity risk premia and breakeven inflation rates. *Economic Review*, 91(2), 29–54.
- Shiller, R. J., Campbell, J. Y., Schoenholtz, K. L., & Weiss, L. (1983). Forward rates and future policy: Interpreting the term structure of interest rates. *Brookings Papers on Economic Activity*, 1983(1), 173–223. <https://doi.org/10.2307/2534355>
- Sierra, J. (2014). International capital flows and bond risk premia. *Quarterly Journal of Finance*, 4(1), 1–36. <https://doi.org/10.1142/s2010139214500013>
- Smith, J. M., & Taylor, J. B. (2009). The term structure of policy rules. *Journal of Monetary Economics*, 56(7), 907–917. <https://doi.org/10.1016/j.jmoneco.2009.09.004>
- Söderlind, P. (2011). Inflation risk premia and survey evidence on macroeconomic uncertainty. *International Journal of Central Banking*, 7(2), 113–133.
- Söderlind, P., & Svensson, L. (1997). New techniques to extract market expectations from financial instruments. *Journal of Monetary Economics*, 40(2), 383–429. [https://doi.org/10.1016/s0304-3932\(97\)00047-0](https://doi.org/10.1016/s0304-3932(97)00047-0)
- Staiger, D., & Stock, J.H. (1997). Instrumental variables regression with weak instruments. *Econometrica*, 65(3), 557–586. <https://doi.org/10.2307/2171753>

- Stock, J.H., & Yogo, M. (2005). Testing for weak instruments in linear IV regression. In D.W.K. Andrews & J.H. Stock (Eds.), *Identification and inference for econometric models: Essays in honor of Thomas Rothenberg* (pp. 80–108). New York: Cambridge University Press.
- Storn, R., & Price, K. (1997). Differential evolution - a simple and efficient heuristic for global optimization over continuous spaces. *Journal of Global Optimization*, 11(4), 341–359. <https://doi.org/10.1023/A:1008202821328>
- Svensson, L. E. O. (1994). Estimating and interpreting forward interest rates: Sweden 1992-1994. *IMF Working Papers*, 94(114), 1–29. <https://doi.org/10.5089/9781451853759.001>
- Vasicek, O. (1977). An equilibrium characterization of the term structure. *Journal of Financial Economics*, 5(2), 177–188. [https://doi.org/10.1016/0304-405x\(77\)90016-2](https://doi.org/10.1016/0304-405x(77)90016-2)
- Vayanos, D., & Vila, J. L. (2009). A preferred-habitat model of the term structure of interest rates. *NBER Working Paper*, No. 15487.
- Vicente, J. V. M., & Graminho, F. M. (2015). Decompondo a inflação implícita. *Revista Brasileira de Economia*, 69(2), 263–284. <https://doi.org/10.5935/0034-7140.20150012>
- Wahlström, R. R., Paraschiv, F., & Schürle, M. (2022). A comparative analysis of parsimonious yield curve models with focus on the Nelson-Siegel, Svensson and Bliss versions. *Computational Economics*, 59(3), 967–1004. <https://doi.org/10.1007/s10614-021-10113-w>
- Weiss, C. R. (2022). Foreign demand for U.S. Treasury securities during the Pandemic. *FEDS Notes*, 3046. <https://doi.org/10.17016/2380-7172.3046>
- Wright, J. H. (2011). Term premia and inflation uncertainty: Empirical evidence from an international panel dataset. *American Economic Review*, 101(4), 1514–1534. <https://doi.org/10.1257/aer.101.4.1514>
- Wu, T. (2014). Unconventional monetary policy and long-term interest rates. *IMF Working Papers*, 14(189), 1–50. <https://doi.org/10.5089/9781498317245.001>
- Zhu, C., Byrd, R. H., Lu, P., & Nocedal, J. (1997). Algorithm 778: L-BFGS-B: Fortran subroutines for large-scale bound-constrained optimization. *ACM Transactions on Mathematical Software (TOMS)*, 23(4), 550–560.