Co-Movement, Spillovers and Excess Returns in Global Bond Markets*

Joseph P. Byrne[†], Shuo Cao[‡] and Dimitris Korobilis[§]

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Abstract

This paper investigates global term structure dynamics using a Bayesian hierarchical factor model augmented with macroeconomic fundamentals. More than half of the variation in bond yields of seven advanced economies is due to global co-movement, which is mainly attributed to shocks to non-fundamentals. Global fundamentals, especially global inflation, affect yields through a 'policy channel' and a 'risk compensation channel', but the effects through two channels are offset. This evidence explains the unsatisfactory performance of fundamentals-driven term structure models. Our approach delineates asymmetric spillovers in global bond markets connected to diverging monetary policies. The proposed model is robust as identified factors has significant explanatory power of excess returns. The finding that global inflation uncertainty is useful in explaining realized excess returns does not rule out regime changing as a source of non-fundamental fluctuations.

Keywords: Global Bond Markets, Term Structure of Interest Rates, Shocks to Fundamentals and Non-Fundamentals, Co-Movement, Contagion, Excess Return.

JEL Classification Codes: C11; C32; E43; F3; G12; G15.

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[†]School of Management and Languages, Heriot-Watt University, Edinburgh, UK. Email: j.p.byrne@hw.ac.uk

[‡]Adam Smith Business School, University of Glasgow, Glasgow, UK. Email: s.cao.1@research.gla.ac.uk

[§]Adam Smith Business School, University of Glasgow, Glasgow, UK. Email: dimitris.korobilis@glasgow.ac.uk

1 Introduction

Reduced-form factor models are widely used in analyzing the term structure of interest rates. These factor models assume the yield curve is driven by a few pricing factors and can be divided into two groups. The first group directly uses economic fundamentals as pricing factors, consistent with preference-based structural models such as Piazzesi and Schneider (2007). This group of fundamentals-driven models, such as Kozicki and Tinsley (2001), Dewachter and Lyrio (2008) and Orphanides and Wei (2012), help us understand how economic fundamentals affect asset prices. In contrast, term structure models using latent pricing factors have more successful empirical fit and avoid the mispricing indicated by Anh and Joslin (2013). However, the second group lacks economic rationale, as Diebold, Rudebusch and Aruoba (2006) indicate that latent factors are not explicitly linked to macroeconomic variables. Joslin, Priebsch and Singleton (2014) reconcile the above seemingly contradictory evidence by proposing a hybrid model that incorporates the joint dynamics of both fundamentals and latent factors. This important work paves a way for us to understand the linkage between fundamentals and non-fundamentals.

Non-fundamentals are essential to asset prices. Lee (1998) finds only 10% of the variance of stock prices are driven by fundamentals. There have been various theories proposed to explore shocks to non-fundamentals. A popular explanation relates non-fundamentals to sentiment or sunspots.[1] For example, Bansal and Shaliastovich (2010) show that the variance of returns are largely driven by a sentiment or confidence measure than fundamentals in the economy. Novy-Marx (2014) reviews the earlier literature and suggests other explanations in addition to sentiment. While most of the current empirical research focuses on the driving forces of domestic asset prices, only a few studies try to approach this topic from a global perspective. Hou, Karolyi and Kho (2011) indicate liberalization of financial markets around the world has increased market co-movement, but whether the co-movement is driven by global macroeconomic factors remains unanswered.

In this paper we aim to study the underlying sources driving global term structures by explicitly considering shocks to fundamentals and non-fundamentals.[2] Diebold, Li and Yue (2008) provide empirical evidence of strong co-movement in yield curves across countries, whereas Den Haan and Sumner (2004) reveal global co-movements in real activity and prices. One question is naturally raised: Whether the co-movement in bond yields is determined by global fundamentals? We specifically tackle this question in a global context, as Barberis, Shleifer and Wurgler (2005) suggest that common movement of asset prices among international markets may not be easily explained by a fundamentals-based

[1] Cass and Shell (1983) use the term intrinsic uncertainty, to refer to anything that affects economic fundamentals, and they use the term extrinsic uncertainty, or sunspots, to refer to anything that does not. Benhabib and Wang (2014) regard extrinsic uncertainty as sentiment or sunspots. Farmer (2014) use the term fundamental shock, to refer to intrinsic uncertainty and non-fundamental shock to refer to extrinsic uncertainty or sunspots.

[2] Lee (1998), Binswanger (2004) and Lanne and Lütkepohl (2010) define that a shock is called fundamental if it affects the real economy such as output and non-fundamental if its effect is transitory or trivial, which is consistent with our definitions in this paper. Our definitions should not be confused with the definitions from an econometric viewpoint such as Lütkepohl (2012), although coincidentally, 'nonfundamentalness' in econometrics can give rise the failure of fundamentals-driven models.

view. Our particular interest is twofold: How much of the variance in global bond yield co-movement is driven by global fundamentals, and why would it be the case? To answer the former, we identify structural shocks of global fundamentals. To further understand the underlying mechanism, we decompose long yield movements into two transmission channels, i.e. a 'policy channel' and a 'risk compensation channel'. These two standard channels are associated with short rate expectations and risk premia, respectively. We then evaluate the effects of global fundamentals through each channel.

Our main finding is in support of the sentiment-based theory favored by Kumar and Lee (2006), Bansal and Shaliastovich (2010) and Benhabib and Wang (2014), as shocks to non-fundamentals generate persistent fluctuations in bond pricing factors but are not followed by moves in global fundamentals. Among all fundamentals, global inflation has demonstrable influence on the co-movement of global short rates. Regarding the co-movements of long rates, there are no significant effects of any shocks to fundamentals. Intuitively, these shocks are offset through different transmission channels as suggested by Duffee (2011), i.e. the shocks driving up expected future short yields drive down term premiums. This empirical evidence explains why a standard structural model with purely shocks to fundamentals cannot generate substantial variability to match bond yield data, see for instance, Piazzesi and Schneider (2007).

To model global term structures of seven advanced economies, we extend the methodology of Moench, Ng and Potter (2013) and propose a novel 'Fundamentals-Augmented Hierarchical Factor Model' (FAHFM). The three-hierarchy structure is in fact an intuitive specification. At the highest global level, we allow global macroeconomic fundamentals to interact with global bond factors. At a lower level, national bond factors are driven by global bond factors and country-specific components. At the lowest level, the term structure of each country is driven by national bond factors and innovations. We specify the model in a setup of Unspanned Macro Risk established by Joslin, Priebsch and Singleton (2014) with global macro fundamentals, which is considered a more realistic and parsimonious specification, and we leave these details in the methodology section.

With our model specification, we jointly identify global and national bond pricing factors in a one-step Bayesian approach. We show that the method is robust as the identified factors help explain bond excess returns. We find two global yield factors can explain on average more than 60% of bond yields variance across our seven countries, and country-specific components contribute to most of the remaining variance. By conducting an analysis on country-specific components we duly unfold asymmetric spillovers among seven countries. While the co-movement of bond yields captures mainly the shocks to non-fundamentals that are closely related to 'monsoonal effects' contagion discussed in Calvo and Reinhart (1996), Masson (1998) and Kaminsky and Reinhart (2000), the spillover effects are linked to diverging monetary policies suggested by Jotikasthira, Le and Lundblad (2015). Lastly, we assess the argument of Bikbov and Chernov (2013) that non-fundamental fluctuations may be related to uncertainty about monetary policy regimes. The ability of global inflation uncertainty in explaining realized excess returns does not rule out this possibility.

Our work is related to the literature of global term structures. Bauer and Diez de los Rios (2012) model the unspanned macroeconomic risks driving international term premia and foreign exchange risk premia. In a similar framework but without international finance restrictions, Abbritti et al. (2013) reveal contrasting forces driving long- and short-term dynamics in yield curves. The most recent work of Jotikasthira, Le and Lundblad (2015) investigates the bond yield co-moments of three countries before the financial crisis. Building upon previous work, our approach jointly identifies robust latent factors with global fundamental augmentation, which is new to the literature. Our model structure follows and extends Diebold, Li and Yue (2008), and we particularly focus on assessing the internal link between co-movement in bond yields and shocks to fundamentals and non-fundamentals. We pin down that more than 70% of bond yield co-movement is driven by shocks to non-fundamentals, which implies standard structure models with purely shocks to fundamentals may fail to match the data. In particular, we show that the offsetting effect of fundamentals through two transmission channels is the key to understand the failure. In addition, our evidence cannot rule out the potential non-fundamental changes are connected with regime shifts in monetary policy rules.

The structure for the paper is as follows. In Section 2 we introduce our model, estimation and identification. In Section 3 we describe the data and present a preliminary data analysis. In Section 4 we explore the dynamics of global yield factors and provide an economic rationale for bond yield co-movement. In Section 5 we identify the asymmetric 'spillovers' in global bond markets. In Section 6 we perform robustness check by testing to what extent our identified factors can explain the expected and realized excess returns, and underline that the explanatory power global inflation uncertainty could be related to potential regime shifts. In Section 7 we conclude and summarize the implications of our analysis.

2 Methodology

2.1 Summary of the Model Structure

We begin by introducing our innovative hierarchical factor methodology. This framework shall model bond yields across countries, using global macro and yield factors. In the spirit of multi-level factor models, Kose, Otrok and Whiteman (2003) propose a dynamic factor model to study international business cycle co-movements, whereas Moench, Ng and Potter (2013) propose a hierarchical factor model to study the US housing market.[3] The hierarchical factor model is relatively more parsimonious in terms of parameters to be identified and factor structure, making it attractive for bond yield modeling. The model for bond yields X_{ibt} can be written as:

$$X_{ibt} = \Lambda_{ib}^F F_{bt} + e_{ibt}^X, \tag{2.1}$$

$$F_{bt} = \Lambda_b^G G_t + e_{bt}^F, \tag{2.2}$$

$$\begin{bmatrix} G_t \\ M_t \end{bmatrix} = \psi^G \begin{bmatrix} G_{t-1} \\ M_{t-1} \end{bmatrix} + \epsilon_t^G, \tag{2.3}$$

[3] Kose, Otrok and Whiteman (2003) identify regional factors that are uncorrelated with the global factors, while Moench, Ng and Potter (2013) aim to find the global factors driving the regional factors. See also Eickmeier, Gambacorta and Hofmann (2014) for multi-factor models of global monetary policy.

in which subscript i indicates the maturities of bond yields, subscript b indicates the countries and subscript t indicates different periods of time. In the above, Λ_{ib}^F, Λ_b^G and ψ^G are model parameters, and e_{ibt}^X, e_{bt}^F and ϵ_t^G are error terms. In the country-level Equation (2.1), X_{ibt} represent the bond yield of country b at maturity i, and F_{bt} are the latent yield factors of country b. In Equation (2.2), G_t are the latent global yield factors that drive the national yield factors F_{bt}. Finally Equation (2.3) describes the interactions between the yield factors and the global macro factors/fundamentals M_t using a Vector Autoregression (VAR).[4] In our model, we include four global macro variables: monetary policy rate, inflation, real activity and financial conditions, such that M_t is a 4×1 vector. The former three are standard macro fundamentals in term structure modeling, see for example, Ang and Piazzesi (2003) and Bikbov and Chernov (2013). Additional, we include financial conditions as the liquidity and credit risk measure suggested by Dewachter and Iania (2012).

Therefore, a key contribution of this paper is to extend the 'Dynamic Hierarchical Factor Model' proposed by Moench, Ng and Potter (2013) by augmenting the VAR system of global yield factors with global macro factors M_t. Our approach is the logical extension of existing work that exploits macro factors when characterizing national yield curves, see Ang and Piazzesi (2003), Diebold, Rudebusch and Aruoba (2006) and Bianchi, Mumtaz and Surico (2009). Moreover, we incorporate global macro factors to provide an economic interpretation of yield movements. The dynamics of the global factors are characterized by an unrestricted Factor Augmented Vector Autoregressive (FAVAR) model. The FAVAR at the global level allows our model to incorporate Unspanned Macro Risks, as proposed by Joslin, Priebsch and Singleton (2014).[5] The extended version of the hierarchical model is therefore defined as 'Fundamentals-Augmented Hierarchical Factor Model' (FAHFM).[6]

The model proposed in this paper has a similar structure to Diebold, Li and Yue (2008) but contrasts in that their approach uses two steps and does not include macro determinants.[7] We adopt a one-step Bayesian technique which should provide more accurate estimates. Diebold, Rudebusch and Aruoba (2006) and Pooter (2007) provide evidence that an one-step approach produces more effective estimates: Two-step estimation introduces bias if it does not fully consider the dynamics of the factors at a higher level. As shown in the previous literature, directly introducing macro fundamentals can provide a meaningful narrative which delineates the macro shocks that drive global term

[4]When referring to global macro fundamentals, 'fundamental' and 'factor' are used interchangeably in this paper.

[5]The setting of Unspanned Macro Risks is a more realistic assumption as suggested by Joslin, Priebsch and Singleton (2014). By definition, if there exists Unspanned Macro Risks, macro factors do not directly or contemporaneously impact yields and they influence current yields only through their correlation with the yield factors. Joslin, Priebsch and Singleton (2014) point out that the majority of previous macro-finance term structure models in reduced-form implied fully spanned macro factors, e.g. Ang and Piazzesi (2003), Ang, Dong and Piazzesi (2007) and Bikbov and Chernov (2010). The fully spanned assumption, i.e. the macro factors can be inverted as linear combinations of yields, is often questioned and might be counterfactual.

[6]Further details of our FAHFM are described in Appendix A.

[7]The first step is to extract the national Nelson-Siegel factors by fixing the loadings. The second step treats the national yield factors as data and then applies a Bayesian method to estimate the parameters and the higher-level latent global factors.

structures. Our hierarchical one-step framework allows us to jointly estimate the global factors and country-specific components, and hence builds upon the contribution of Bauer and Diez de los Rios (2012), Abbritti et al. (2013) and Jotikasthira, Le and Lundblad (2015).[8]

In this paper we estimate our global term structure model with macro fundamentals by a Bayesian estimation technique, specifically, the Gibbs sampling. In the Gibbs sampling, we begin with $50,000$ burn in draws and then save every 50th of the remaining $50,000$ draws. These 1000 draws are used to compute posterior means and standard deviations of the factors, as well as the posterior coverage intervals in the following sections.

2.2 Identification of Factors

To identify the global factors, a standard approach is the Principal Component method, but this lacks economic structure. More economically motivated restrictions include Nelson-Siegel and no-arbitrage constraints, see, for example, Diebold, Li and Yue (2008) and Jotikasthira, Le and Lundblad (2015), respectively. Duffee (2013) suggests these two alternative restrictions are equivalent in characterizing the cross section of interest rate term structure.

In this paper therefore we use an alternative identification scheme to Moench, Ng and Potter (2013) and impose cross-sectional restrictions. While these authors use zero restrictions which are of a statistical nature, we impose restrictions implied by the dynamic Nelson-Siegel (NS) term-structure model.[9] In other words, the loadings of country-level factors are fixed following Diebold, Li and Yue (2008).[10] The NS identification scheme has gained great popularity in term structure modeling, and we choose this scheme to fix our ideas. In the following sections we provide an economic interpretation of the latent pricing factors, so we can further understand the sources that drive the global yield co-movements.

[8]More sophisticated identification schemes of structural shocks can be introduced efficiently in this one-step approach, without running additional regressions that can potentially introduce bias.

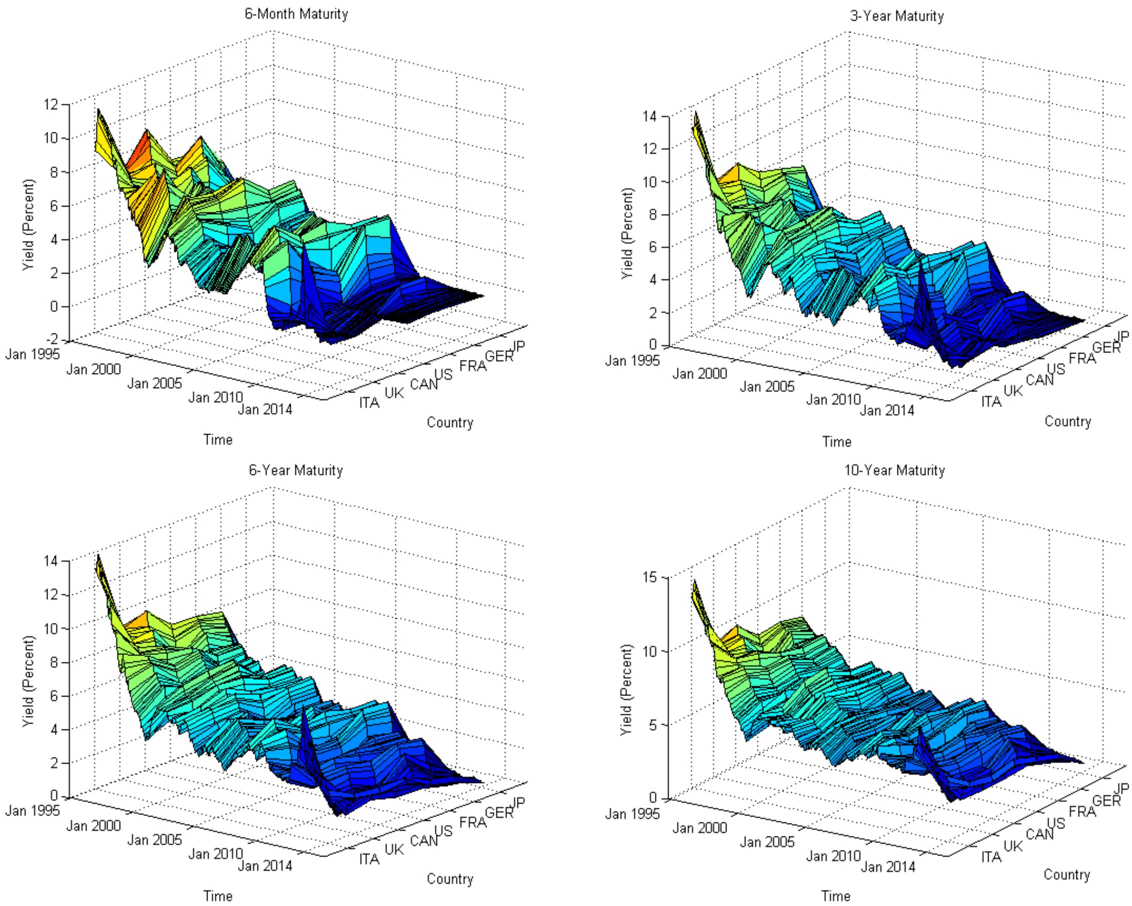
[9]We do not impose no-arbitrage constraints in our model. Joslin, Le and Singleton (2013) show that Gaussian no-arbitrage macro-finance models are close to factor-VAR models when risk premia dynamics are not constrained. Duffee (2014) also indicates that the no-arbitrage restrictions are unimportant if a model aims to pin down physical dynamics. Since our focus here is not on the structure of risk premia dynamics, we choose to impose no such restrictions to avoid potential misspecification. The potential drawback of no-arbitrage models is that it imposes very strong restrictions on the dynamics of risk prices, in order to 1) ensure no-arbitrage consumption and 2) identify the model with flat likelihood. Kim and Singleton (2012) and Jotikasthira, Le and Lundblad (2015) indicate the no-arbitrage framework may generate implausibly term premiums in the financial crisis. Instead, we impose Nelson-Siegel restrictions here, which provides a parsimonious structure and satisfactory performance in cross-sectional fittings of term structure.

[10]The details of the restrictions can be found in Appendix A.2. In fact, the two schemes, Diebold, Li and Yue (2008) and Moench, Ng and Potter (2013), share similar results, as shown in Appendix B. For more information regarding factor identification we refer the reader to Bai and Wang (2015).

3 Data Description and Preliminary Results

We obtain monthly bond yield data from Bloomberg for seven advanced countries: Italy, Canada, France, Germany, Japan, the UK and the US.[11] The empirical analysis focuses on yields of 11 maturities: 3, 6, 12, 24, 36, 48, 60, 72, 84, 96 and 120 months. Figure 1 shows the dynamics of bond yield at four maturities across all seven countries. All four maturities trend down from the beginning of the sample period, with the shorter rates displaying greater variance across time and countries.

Figure 1: Bond Yields of Seven Countries



Notes:
1. The above charts plot the bond yields for the seven countries in the sample. The sample includes Italy (ITA), Canada (CAN), France (FRA), Germany (GER), Japan (JP), the UK and the US, spanning from Dec. 1994 to Mar. 2014.
2. From top left clock-wise we have bond yields of maturities 6 months, 3 year, 10 years and 6 years. More information about the data is provided in Appendix C.

Our empirical model uses macroeconomic variables from Bloomberg, and indicators of financial condition from St. Louis Federal Reserve Economic Data (FRED). We construct

[11]The zero-coupon yields are calculated step-by-step using the discount factors that are derived from standard bootstrapping, given the set of coupon bonds, bills, swaps or a combination of these instruments. A minimum of four instruments at different tenors are required for each yield curve. The bootstrapping is similar to the Unsmoothed Fama-Bliss method, see Fama and Bliss (1987).

four global macro factors using a list of macro fundamentals among the seven countries, and the fundamentals include inflation (CPI), Industrial Production (IP) and monetary policy rates. We also use a large number of regional series of Financial Condition Index (FCI) to construct a global FCI. The full sample of monthly data is from December 1994 to March 2014.[12]

Before we implement our one-step estimation, the global macro factors M_t are extracted from regional macro series. There are four categories of regional macro series: the policy rate, indicator of real activity, inflation and Financial Condition Index (FCI). We employ a new approach proposed by Koop and Korobilis (2014) to extract the global macro indicators from regional series.[13] Figure 7 in Appendix D displays the estimated macro factors used to augment our proposed model.

In Table 1 we report summary statistics for bond yields at representative maturitiess. All yield curves are upward-sloping, suggesting positive term spreads. Apart from Japan, the yield volatility generally decreases with maturity, and all the yields are highly persistent for all countries, with first-order autocorrelation greater than 0.95. Japanese yields are typically the lowest, usually below two percent and are less persistent when compared to other yields.

[12]The details about the data are described in the Data Appendix C.

[13]Our main results are robust to the measure of global macro factors using Stock and Watson (2002).

Table 1: Descriptive Statistics of Bond Yields

Country	Maturity	Mean	Std. Dev.	Min.	Max.	$\hat{\rho}(1)$	$\hat{\rho}(12)$	$\hat{\rho}(30)$
US	3	2.82	2.27	0.01	6.39	0.99	0.74	0.26
	12	3.04	2.26	0.10	7.20	0.98	0.76	0.30
	60	3.92	1.88	0.59	8.03	0.97	0.76	0.42
	120	4.57	1.45	1.54	8.00	0.97	0.72	0.43
UK	3	3.91	2.32	0.28	7.50	0.99	0.77	0.47
	12	4.00	2.36	0.12	7.45	0.99	0.77	0.48
	60	4.51	2.00	0.58	8.94	0.98	0.75	0.43
	120	4.85	1.66	1.57	8.90	0.97	0.72	0.29
Germany	3	2.49	1.52	0.00	5.14	0.98	0.66	0.27
	12	2.63	1.53	0.01	5.82	0.98	0.64	0.25
	60	3.48	1.55	0.33	7.47	0.97	0.67	0.35
	120	4.17	1.46	1.22	7.69	0.97	0.70	0.35
France	3	2.63	1.68	0.01	7.93	0.98	0.56	0.21
	12	2.76	1.66	0.02	7.04	0.97	0.58	0.22
	60	3.67	1.49	0.69	7.87	0.96	0.61	0.29
	120	4.42	1.32	1.85	8.14	0.96	0.63	0.30
Italy	3	3.44	2.58	0.28	11.00	0.98	0.63	0.24
	12	3.73	2.50	0.60	11.74	0.98	0.57	0.17
	60	4.90	2.39	1.95	14.01	0.96	0.51	0.11
	120	5.61	2.22	3.42	14.14	0.97	0.54	0.09
Canada	3	3.10	1.91	0.21	8.88	0.96	0.59	0.28
	12	3.36	1.90	0.49	8.88	0.97	0.64	0.33
	60	4.23	1.80	1.19	9.40	0.97	0.74	0.45
	120	4.75	1.69	1.72	9.48	0.97	0.74	0.41
Japan	3	0.25	0.34	0.00	2.24	0.89	0.28	0.07
	12	0.31	0.37	0.01	2.48	0.89	0.39	0.07
	60	0.91	0.66	0.13	4.07	0.92	0.57	0.17
	120	1.66	0.77	0.55	4.79	0.94	0.60	0.18

Notes: This table presents descriptive statistics for monthly yields at different maturities across seven countries. The sample period is 1994:12–2014:03. We use the following abbreviations. **Std. Dev.**: Standard Deviation; **Min.**: Minimum; **Max.**: Maximum; $\hat{\rho}(\mathbf{k})$: Sample Autocorrelation for Lag k.

In Table 2 we present descriptive statistics for the global and national factors estimated by our FAHFM approach. For each country, we extract a Level and Slope factor.[14] Diebold and Li (2006) interpret the Level as the long-term factor, while the Slope is

[14]The first factor is the 'Level' factor: a shock on the 'Level' changes the interest rates of all maturities by almost identical amounts, inducing a parallel shift that changes the level of the term structure. The other one is the 'Slope' factor: a shock on the 'Slope' factor increases short-term interest rates by larger amounts than the long-term interest rates, so that the yield curve becomes flatter.

a short-term factor. Posterior means and variances for the estimated Level factors are relatively similar because we use standardized data with zero mean and variance equal to one. Differences in the standard deviation of the Slope factors potentially reflect the impact of the national macroeconomic environment and monetary policy. The factor autocorrelations reveal that all factors display persistent dynamics, but the Level is more persistent than the Slope for all countries.

Table 2: Descriptive Statistics of Yield Factors

Country	Factor	Mean	Std. Dev.	Minimum	Maximum	$\hat{\rho}(1)$	$\hat{\rho}(12)$	$\hat{\rho}(30)$
Global	Level	0.00	1.05	-1.99	2.96	0.97	0.70	0.36
	Slope	0.00	0.58	-1.13	0.97	0.97	0.30	-0.25
US	Level	0.00	1.03	-2.11	2.42	0.97	0.72	0.43
	Slope	0.00	0.65	-1.53	1.22	0.94	0.26	-0.27
UK	Level	0.00	1.04	-2.14	2.69	0.97	0.70	0.35
	Slope	0.00	0.68	-1.84	1.58	0.95	0.31	-0.23
Germany	Level	0.00	1.01	-2.07	2.71	0.97	0.67	0.33
	Slope	0.00	0.44	-1.00	0.84	0.98	0.32	-0.24
France	Level	0.00	1.04	-2.22	3.08	0.96	0.60	0.28
	Slope	0.00	0.76	-1.55	1.42	0.96	0.30	-0.15
Italy	Level	0.00	1.05	-1.16	4.20	0.95	0.48	0.08
	Slope	-0.01	0.68	-3.04	1.07	0.93	0.42	0.07
Canada	Level	0.00	1.05	-1.90	2.79	0.97	0.74	0.40
	Slope	0.00	0.93	-2.06	1.45	0.98	0.36	-0.42
Japan	Level	0.00	1.08	-1.47	3.87	0.95	0.62	0.21
	Slope	0.00	1.03	-2.68	2.92	0.96	0.27	0.07

Notes: This table presents descriptive statistics for the global yield factors and national factors across seven countries. The Level and Slope factors are estimated from our proposed FAHFM. The sample period is 1994:12–2014:03. We use the following abbreviations. **Std. Dev.**: Standard Deviation; $\hat{\rho}(\mathbf{k})$: Sample Autocorrelation for Lag k.

3.1 Variance Decomposition of Model Hierarchies

As mentioned in the previous section, for each country we identify two latent pricing factors.[15] It is only the global Level factor in our model that drives the national Level factors. Similarly the global Slope drives national Slope factors. Table 3 displays the importance of the global-level ($Share_G$) and country-specific ($Share_F$) components in Eq. (2.3) and (2.2), as well as idiosyncratic noise ($Share_X$) in Eq. (2.1), relative to the total variation in the data of seven countries. It is clear that the global factors

[15] We only use two factors as two country-level principal components can account for more than 99% of the variance of bond yields across all countries.

explain the vast majority of country yields: $Share_G$ is greater than 0.6 for almost all countries.[16] Consequently, this characteristic leads us to believe the co-movement of international bond yields is generally very strong and dominates national or idiosyncratic movements.[17] The evidence is consistent with the importance of the global factors found in Diebold, Li and Yue (2008). As the global factors account for a large proportion of the information in national term structures, we are interested in the dynamics of the two global factors, Level and Slope, and seek to provide sensible economic interpretations for the factors in this study.

Although global factors clearly dominate yields, national factors remain important. The variance explained by country-specific components (i.e. $Share_F$) is non-trivial and more than two standard deviations from zero. This in turn implies, that the sum of $Share_G$ of global factors and $Share_F$ of country-specific components account for $96-99\%$ of bond variation across all countries.[18] The idiosyncratic noise is largely irrelevant and our model is doing a good job modeling yield (co)movement. It is consistent with the early evidence of Litterman and Scheinkman (1991) for bond markets. Having identified significant co-movement in yields using an empirical factor approach, we now seek to model international yields in more depth in the next section.

Table 3: Decomposition of Variance of Hierarchies

Country	Posterior Mean (Std. Dev.)		
	$Share_G$	$Share_F$	$Share_X$
Italy	0.36(0.10)	0.63(0.10)	0.01(0.00)
Canada	0.71(0.07)	0.27(0.07)	0.02(0.00)
France	0.76(0.07)	0.22(0.06)	0.02(0.00)
Germany	0.74(0.07)	0.22(0.06)	0.04(0.01)
Japan	0.68(0.08)	0.30(0.07)	0.03(0.01)
UK	0.85(0.05)	0.13(0.04)	0.02(0.01)
US	0.75(0.07)	0.24(0.07)	0.01(0.00)

Notes: This table summarizes the decomposition of variance for the three-level hierarchical model of bond yields. Std. Dev. denotes the posterior standard deviation of the posterior mean. For each country, $Share_G$, $Share_F$ and $Share_X$ denote the average variance shares across all maturities due to shocks of ϵ_G, ϵ_F and ϵ_X at different levels, respectively. Parentheses $(\cdot)$ contain the posterior standard deviation of shares in a specific block.

[16]The exception is Italy potentially as those yields bear higher sovereign and hence country-specific risks.

[17]We can refer to Figure 1 in last section, which gives a more intuitive impression.

[18]In other words, the sum equals to the share of variance of national yield factors. Note there is a clear distinction between national factors and country-specific components. Country-specific components are the movements in national factors that are not driven by global factors.

4 Co-Movement in Yields

4.1 Factor Dynamics

In this section, we depict the dynamics of the global yield factors estimated from our proposed 'Fundamentals-Augmented Hierarchical Factor Model'. As mentioned before, we extract two national yield factors that account for more than 96% of the variance of the term structure. We now focus on the global yield factors, as these factors typically drive the national Level and Slope factors. Firstly, we calculate the arithmetic sum of the global Level and Slope factors to evaluate the effect on the global short rate co-movement. This sum is denoted as the *global short rate factor*, and reflects the global co-movement in short rates across countries.[19] From the left panel in Figure 2, we can see the global short rate factor is strongly correlated with the first principal component of short rates across the seven advanced economies, also implying our model successfully captures the global co-movement of the short rates.[20] One feature of the movements of the global short rate factor is that it falls sharply after the Global Financial Crisis, consistent with a global expansion in monetary policy.

It is straightforward to decompose the global short rate factor into the global Level and Slope. The movements of these two factors are shown in the right panel in Figure 2, in which we also highlight some distinct historical events: January 1999 and the start of the euro area, US recessions in 2001 and 2008 as defined by NBER and the European sovereign debt crisis. As we have already discussed, Level and Slope factors control the shape of the term structure, which can be informative in revealing useful macroeconomic information. For example, before 1999 there is a downward trend for the Level factor and an upward trend for the Slope factor, which means the global term structure is moving down and flattening.[21] This phenomenon indicates a moderation in global term structure, possibly caused by greater integration.[22] We can observe two clear trends abstracting from temporary disturbances in the factors. Firstly, the downward-trending global Level seems to relate to the decreasing inflation level in the period of the Great Moderation, as suggested by Evans and Marshall (2007) and Koopman, Mallee and Van der Wel (2010). Secondly, the Slope factor is declining during US recessions, suggesting it is related to real economic activity, as indicated in Kurmann and Otrok (2013). Further evaluation of factor commonality can be found in Appendix E.

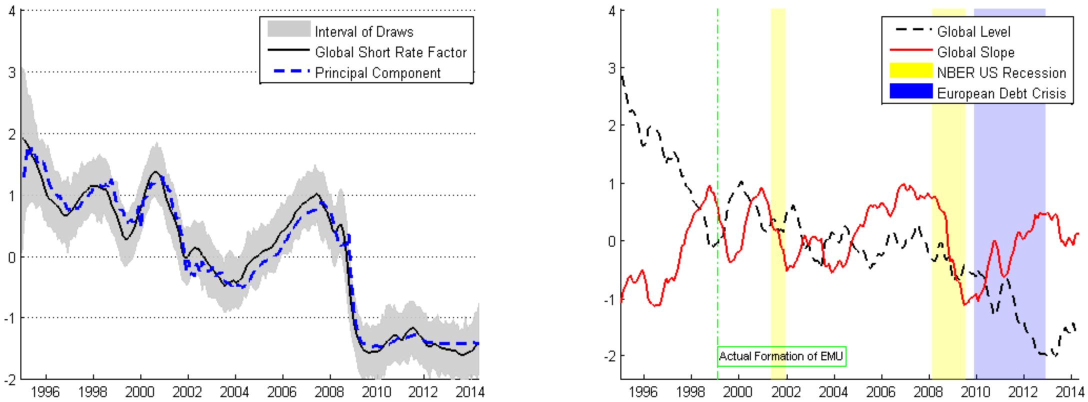
[19]By NS restrictions, for a bond at very short maturity, we have the equation that *short rate* $= \beta_1 L_t^{NS} + \beta_2 S_t^{NS}$, where the loadings equal to one, i.e. $\beta_1 = \beta_2 = 1$. Therefore, the short rate is directly driven by the sum of two factors in our model construction, see Appendix A.2 for details.

[20]Note that there is a smaller proportion of bond yield movements in country level that are not captured by the global yield factors. We find that these country-specific movements in national yield factors can be largely explained by the divergence of monetary policy in different countries. The results are consistent with the findings in Jotikasthira, Le and Lundblad (2015), but not shown here as we focus on the global co-movement.

[21]An increase in the level factor is consistent with higher yields on average. An increase in the slope factor is consistent with a flatter yield curve.

[22]The strong negative correlation between the Level and Slop disappears after 1999 and reappears after the financial crisis.

Figure 2: Global Short Rate Factor and the Decomposition



Notes: 1. The left panel shows the global short rate factor (i.e. an arithmetic sum of extracted global Level and Slope factors) and the first principal component of the national short-run policy rates (dashed line). The first principal component of national policy rates, which accounts for more than 84% of total variance of national policy rates. The gray areas cover all the draws of the global short rate factor (i.e. Level + Slope) from our model, and the solid black line is the median value of the draws. Data standardization implies yields can fall below zero.
2. The right panel shows the decomposition of the median of global short rate factor. We decompose the short rate factor into the global Level (dashed line) and the global Slope (solid red line). In general, the Level factor controls the level of the term structure whereas the Slope factor controls the slope of the term structure. The shaded areas cover some major recession periods in the US and Europe.

4.2 What Drives Term Structures across Countries?

In Section 3.1, we show that the global yield factors account for the majority of the variance of bond yields. There are important co-movements of yields, although the co-movements are primarily at the long end of the yield curve according to Byrne, Fazio and Fiess (2012) and Jotikasthira, Le and Lundblad (2015).[23] It may be due to the uncoupling of short-term policy rates in different countries.

To evaluate the relative importance of global fundamentals and non-fundamentals in driving the co-movement of bond yields, we further decompose the 40-period Forecast Error Variance (FEV) of yields driven by innovations of global factors. The global macro factors, i.e. IP, CPI and FCI, are extracted from the first stage of the FAHFM. Note that the shares are quantitatively similar for all countries based upon our model construction, except for Japan where inflation accounts for much less variance, i.e. around half of the shares of the other countries. Our results suggest that much of the FEV can be explained by the innovations of global factors as implied in Section 3.1. The remainder of FEV is explained by the country-specific components and the idiosyncratic innovations across yields at different maturities. The country-specific components in national yield factors are also meaningful as it may imply global 'spillovers', and hence we will discuss this in detail in Section 5. For now, this section focuses upon the global co-movement in yields.

In Table 4, we show the decomposition of the variance of all US yields explained by

[23]Our results also support that the variance accounted for by the global yield curve increases with yield maturity, see Table 20, 21 and 22 in Appendix F.

global factors. Our first finding is that shocks to non-fundamentals, i.e. shocks to the global Level and Slope factors, are the main sources driving interest rate movements. We observe the proportion of shocks to non-fundamentals increases with maturities. Moreover, we find that shocks to non-fundamentals are persistent and not followed by changes in fundamentals, which is consistent with Benhabib and Wang (2014).[24] Another observation is that the Level dominates, especially at longer maturities (around 79%), whereas the Slope is relatively more important for shorter maturities (up to 19%). Diebold and Rudebusch (2013) suggest the Slope factor is connected to investors' view about the stance of current monetary policy. If this argument is true, our empirical evidence does not show global Slope shocks, or the changing expectations indicated in Diebold and Rudebusch (2013), have detectable influence on global fundamentals. As suggested by Novy-Marx (2014), the underlying sources of shocks to non-fundamentals are not clear without further analysis.

Among all fundamentals, CPI accounts for a significant fraction of bond yield co-movement at shorter maturities, contributing to up to 21% of FEV of co-movement. The shares are considerably lower, however, for bonds at longer maturities. This finding is consistent with the results in the previous sections, because for a long-term bond, a negative inflation shock is likely to impose downward pressure on the risk premia, so the variance in short rate expectations is offset by the movements of risk premia. To have a deeper understanding of how global fundamentals affect the yields, in the next section we decompose the yield co-movements into two channels, in light of the results of Wright (2011) and Jotikasthira, Le and Lundblad (2015).

[24]More details about the identifications of different shocks in global yield factors can refer to Appendix G.

Table 4: Decomposition of US yield Variance Explained by Global Factors

Maturity (Month)	Posterior Mean (Standard Deviation)					
	IP	CPI	PR	Level	Slope	FCI
3	0.02(0.02)	0.21(0.12)	0.03(0.02)	0.51(0.11)	0.19(0.10)	0.05(0.04)
6	0.02(0.02)	0.19(0.12)	0.02(0.02)	0.54(0.12)	0.17(0.10)	0.05(0.04)
12	0.02(0.02)	0.17(0.12)	0.02(0.02)	0.60(0.13)	0.15(0.10)	0.05(0.04)
24	0.02(0.02)	0.13(0.11)	0.02(0.02)	0.67(0.14)	0.11(0.09)	0.04(0.04)
36	0.02(0.02)	0.11(0.11)	0.02(0.02)	0.72(0.14)	0.10(0.09)	0.04(0.04)
48	0.02(0.02)	0.10(0.11)	0.02(0.02)	0.74(0.14)	0.08(0.09)	0.03(0.04)
60	0.02(0.02)	0.09(0.10)	0.02(0.02)	0.76(0.14)	0.08(0.09)	0.03(0.04)
72	0.02(0.02)	0.08(0.10)	0.02(0.02)	0.77(0.14)	0.07(0.09)	0.03(0.04)
84	0.02(0.02)	0.08(0.10)	0.02(0.02)	0.78(0.13)	0.07(0.09)	0.03(0.04)
96	0.02(0.02)	0.08(0.09)	0.02(0.02)	0.78(0.13)	0.07(0.09)	0.03(0.04)
120	0.02(0.02)	0.07(0.09)	0.02(0.02)	0.79(0.13)	0.07(0.08)	0.03(0.04)

Notes: 1. This table summarizes the posterior mean of the decomposition of 40-period Forecast Error Variance of US bond yields driven by innovations of global yield and macro factors. In each parentheses $(\cdot)$ the posterior standard deviation of shares in a specific block is calculated from our draws, see Section 2. Larger Standard Deviation means higher uncertainty in the estimates, but we do not have an exact credible interval interpretation as the statistics do not necessarily follow (truncated) normal distributions.
2. IP, CPI, PR, Level, Slope and FCI denote the variance shares at different maturities in the country-level block due to global shocks of the Industrial Production growth rate (YoY), inflation, change of policy rate (YoY), global Level, global Slope and FCI, respectively. The shares in each row sum up to 1.
3. We employ Cholesky decomposition to identify the shocks using the following ordering: IP, CPI, PR, Level, Slope and FCI. The details can be found in Appendix A.3.

4.3 Policy and Risk Compensation Channels

From the previous section, inflation does not explain co-movement in long yields, although it explains a large proportion of short yields. We know short rates are generally anchored by the policy rate targets decided by national monetary authorities. On the other hand, the long rates reflect the effectiveness of central bank communications that anchor the future expectations of short rates and inflation. For instance, an 'inflation shock' can be transmitted to bond yields through two channels. The first channel is the influence on the current short rate and expected future short rates. The current short rate and future short rate expectations are closely connected to monetary policy, so we regard this channel as the 'policy channel'. The movements in this policy channel are in line with the 'Expectation Hypothesis'. The other channel is the 'risk compensation channel', through which the movements account for the bond market risk compensation for a bond at longer maturity. The compensation is also called 'term premia', which is the difference between the real long yield and the 'Expectation Hypothesis' consistent long yield. Following the approach of Wright (2011) and Jotikasthira, Le and Lundblad (2015), we aim to decompose the long yield movements into these two distinct channels and assess their relative importance.[25]

[25] Our definitions of these two channels are similar to Jotikasthira, Le and Lundblad (2015), although our model structure is different. See Appendix A.3 for technical details.

In summary, the policy channel determines expected short rates while the risk compensation channel accounts for movements of the term premia. Hence, Table 5 displays the proportion of variance of the 10-year bond driven by global factors accounted for by each channel. The tables also show the shares of the influence of each global yield factor or global macro factor though these two alternative channels.[26] Firstly, we find that the co-movements of the 10-year bond are largely driven by the risk compensation channel. For all seven countries, this risk channel accounts for more than 57% of the total variance of long rate co-movement. For Japan, the risk compensation channel even accounts for 97%. The relative importance of the risk compensation channel is in line with the results in Jotikasthira, Le and Lundblad (2015).[27] Secondly, we find that inflation is very important in driving the global co-movement of long yields through both channels. Take the US for example, recall Table 4, the joint contribution of CPI inflation to the 10-year bond co-movement is only 7%. But surprisingly, when we decompose the influence into the policy channel and risk compensation channel, the contribution through each channel is significantly increased, especially for the policy channel through which the share triples.

[26]The results for other long yields (maturities 5 to 9 years) do not vary much and therefore are not displayed here.

[27]Jotikasthira, Le and Lundblad (2015) indicate the risk compensation channel accounts for around 80% and 42% for the US and Germany, respectively. We include the financial crisis period in our sample so we have a decreased share for the US and an increased share for Germany.

Table 5: Decomposition of Variance through Two Channels (10-Year Bonds)

Country	Channel	Posterior Mean (Std. Dev.)					
		IP	CPI	PR	Level	Slope	FCI
US	Policy	0.02	0.30	0.04	0.49	0.09	0.06
	43%	(0.02)	(0.19)	(0.04)	(0.24)	(0.1)	(0.07)
	Risk Compensation	0.02	0.13	0.06	0.68	0.06	0.05
	57%	(0.02)	(0.11)	(0.03)	(0.17)	(0.06)	(0.05)
UK	Policy	0.02	0.32	0.04	0.46	0.09	0.07
	35%	(0.02)	(0.19)	(0.04)	(0.24)	(0.1)	(0.07)
	Risk Compensation	0.02	0.11	0.05	0.71	0.06	0.05
	65%	(0.02)	(0.09)	(0.03)	(0.15)	(0.06)	(0.05)
Germany	Policy	0.02	0.33	0.04	0.45	0.09	0.07
	21%	(0.02)	(0.19)	(0.04)	(0.24)	(0.1)	(0.07)
	Risk Compensation	0.02	0.08	0.04	0.77	0.06	0.04
	79%	(0.02)	(0.06)	(0.02)	(0.11)	(0.06)	(0.04)
France	Policy	0.02	0.32	0.04	0.46	0.09	0.07
	30%	(0.02)	(0.19)	(0.04)	(0.24)	(0.1)	(0.07)
	Risk Compensation	0.02	0.10	0.05	0.74	0.06	0.05
	70%	(0.02)	(0.08)	(0.02)	(0.13)	(0.06)	(0.04)
Italy	Policy	0.02	0.30	0.04	0.48	0.09	0.07
	26%	(0.02)	(0.19)	(0.04)	(0.24)	(0.1)	(0.07)
	Risk Compensation	0.02	0.08	0.04	0.76	0.06	0.04
	74%	(0.02)	(0.07)	(0.02)	(0.12)	(0.06)	(0.04)
Canada	Policy	0.02	0.30	0.04	0.48	0.09	0.07
	25%	(0.02)	(0.19)	(0.04)	(0.24)	(0.1)	(0.07)
	Risk Compensation	0.02	0.08	0.04	0.76	0.06	0.04
	75%	(0.02)	(0.07)	(0.02)	(0.12)	(0.06)	(0.04)
Japan	Policy	0.02	0.22	0.03	0.58	0.09	0.06
	3%	(0.02)	(0.17)	(0.04)	(0.23)	(0.1)	(0.07)
	Risk Compensation	0.02	0.06	0.03	0.81	0.06	0.03
	97%	(0.02)	(0.06)	(0.01)	(0.09)	(0.06)	(0.03)

Notes: 1. This table summarizes the decomposition of 40-period Forecast Error Variance of the 10-year bond yields driven by innovations of factors through two channels: the policy and risk premia channels. In each parentheses $(\cdot)$ the posterior standard deviation of shares in a specific block is calculated from our draws, see Section 2. Larger standard deviation means higher uncertainty in the estimates, but we do not have an exact credible interval interpretation as the statistics do not necessarily follow (truncated) normal distributions.
2. The global 10-year bond is obtained from our model. IP, CPI, PR, Level, Slope and FCI denote the variance shares at different maturities in the country-level block due to global shocks of the Industrial Production growth rate (YoY), inflation, change of policy rate (YoY), global Level, global Slope and FCI, respectively. The shares in each row sum up to 1.
3. We employ Cholesky decomposition to identify the shocks using the following ordering: IP, CPI, PR, Level, Slope and FCI. The details can be found in Appendix A.3.

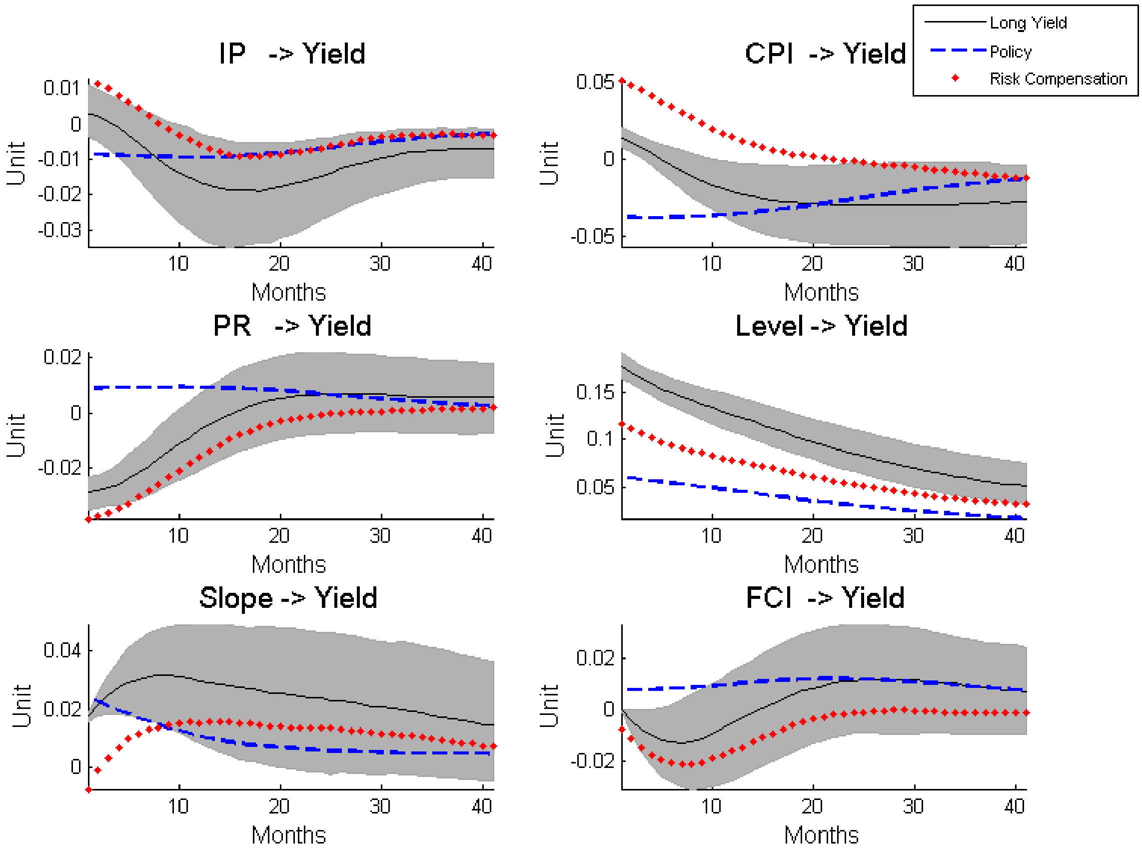
Why might this be the case? Impulse responses help us understand why. There is a sizable reduction in the overall influence of inflation because risk and policy channels counteract one another, which also applies to other global fundamentals. We plot the impulse response of the co-movement of the US 10-year bond to global shocks in Figure 3 and the two offsetting effects are revealed. A global shock that has positive effects through the policy channel usually has negative effects through the risk compensation channel. These opposite effects of macroeconomic shocks (i.e. IP, CPI, PR and FCI) can potentially explain why the current shocks on yield factors are not driven by the macroeconomic shocks indicated in Appendix G. This observation highlights the importance of shocks to non-fundamentals, and we need to go beyond a fundamentals-driven model to capture bond yield movements.

Wright (2011) suggested that inflation uncertainly is closely connected to global bond yield movements, through both the policy and the risk compensation channels. Specifically, Wright (2011) points out that the term premia is positively correlated with inflation expectation, and our findings confirm this mechanism. The top-right panel in Figure 3 indicates that a positive shock on current inflation immediately drives up the term premia, as the increase in inflation may raise inflation uncertainty and hence the risk premia. Inflation is the most important driver of the term premia among all macro variables, both in terms of the quantity and persistency.

Why might the term premium be sensitive to inflation? It is worth mentioning a positive financial shock drives down the term premia, and the effects are quite persistent; see the bottom-right panel in Figure 3. The underlying mechanism is that a positive or malignant financial shock will drive down future long-run inflation, and hence this expectation lowers the term premia. See Appendix G.2 for more extensive discussion.

Figure 3 also shows how the global policy channel reacts to changes in inflation. It is interesting that a positive shock to global inflation is accompanied by a decrease in the policy rate, which seems to contrast a standard Taylor rule. This observation is because for the G7 countries, our identified global inflation shock is not orthogonal to changes in global bond yield factors, and a positive inflation shock is in fact a negative shock to global Level factor. The story dates back to Blanchard and Quah (1989), in which a supply shock to industrialized economies would result in cost-push inflation and recession. As shown in Appendix G.1.2, the positive inflation shock is accompanied by undesired movements in Industrial Production (IP) growth rate or FCI, so the global policy rate level decreases to offset to these expected movements.

Figure 3: Impulse Responses of US 10-Year Bond Co-Movement to Global Shocks



Notes: 1. This table presents impulse responses of the US 10-year bond to shocks on global factors. The solid blue lines in the above panels show the impulse response of the 10-year long yield co-movements driven by six orthogonal global shocks (positive), respectively. Cholesky decomposition is employed to identify the shocks, see Appendix A.3. The 16 to 84 percent posterior coverage intervals of the long yield are are shaded in gray.
2. The long yield co-movements can be decomposed into two channels: the policy channel (blue dashed line) and the risk compensation channel (red dotted line).

5 Contagion

Apart from the global shocks, how would the country-specific components of one country affect other countries? To shed light on this question, we employ the concepts of 'contagion' proposed by Calvo and Reinhart (1996), Masson (1998) and Kaminsky and Reinhart (2000). They distinguish between 'monsoon effects' contagion, which arises because of common shocks, and 'spillover effects' contagion which arises after controlling for common shocks. The 'monsoon effects' contagion in global bond markets can be viewed as the bond yield movements driven by the common factors, which could be caused by the changes in sentiment suggested by these earlier studies. These movements are well captured by our identified global Level and Slope factors augmented with global macro information. Given the common shocks have been controlled for, we are able to analyze the 'spillovers', which are induced by country-specific components and may be asymmetric among our sample of countries. Jotikasthira, Le and Lundblad (2015) suggest country-specific components are caused by the uncoupling of policy rates, so this analysis can help us understand the spillovers of diverging monetary policies, which may be closely related to country-specific fundamentals.[28]

Our constructed model allows us to separate the country-specific components driving the national yield factors from the global yield factors. Although the country-specific components are modeled in individual processes which do not explicitly specify the interdependence among countries, it does not preclude potential correlation among these components. In fact, if these identified components are truly correlated, the interdependent relations imply 'spillovers' among countries apart from the common shocks.

It is evident that there exists strong cross country correlations.[29] For example, US Level and Slope factors are related to Canadian, German, Italian, and Japanese factors. German factors are also related France and Italy. The strong correlations encourage us to explore the inner mechanism of potential 'spillovers'. How would the country-specific components in yield factors of one country affect the movements of the components of another country and to what extent? Are these effects symmetric or asymmetric? The answers for these questions are desirable and we will conduct the following evaluation process and try to provide sensible evidence.

To further investigate the nature of cross country contagion, we employ the 'Granger Causality' test as a robustness check, in order to see whether movements of one series are followed by movements of another series. If one country-specific component Granger-cause another country-specific component, then it is very likely that there exist 'spillover effects' apart from the common shocks, as 'Granger causality' is consistent with the concept of contagion. In Appendix H, we find significant 'Granger causality' among country-specific components and the directions of spillovers among countries are also revealed.

We would like to further analyze the global connectedness by quantifying the 'spillover effects' among countries, so we employ the approach proposed by Diebold and Yilmaz (2009, 2014); see Appendix A.5 for details. In other words, we construct a VAR(1) system using the country-specific components, and then conduct generalized variance decomposition of the form proposed by Koop, Pesaran and Potter (1996). The decomposition helps us delineate connectedness, because this arises through the covariance matrix that can reveal contemporaneous correlation.[30] To quantify connectedness we follow Diebold and Yilmaz (2009, 2014) and calculate Spillover Indexes based on the variance decomposition. The results are reported in Table 6.

[28]Our empirical results affirm that the country-specific components influencing the national yield factors are largely accounted by the divergence of policy rates in different countries. We can construct the indicators of the divergence of policy rates by subtracting the principal component of all policy rate series from each national policy rate series, and the residuals indicate the monetary policy divergence. For each country, adding the indicator of divergence as an additional explainable variable in the regressions of global yield factors can greatly improve the explanatory power of the regression model, and the usefulness of this local divergence variable is distinguished by the high significance. The finding is robust as it holds for all countries across yield maturities, especially for the short yields. The results are not shown here due to the limited space, but are available under request.

[29]Table 23 in Appendix H displays the correlation matrix of the country-specific components in national Level and Slope factors.

[30]Alternative schemes, for example, network connectedness measures based on Granger-causal patterns of Billio et al. (2012), can be employed as robustness checks. See Appendix H for details.

Table 6: Spillover Table of the Country-Specific Components

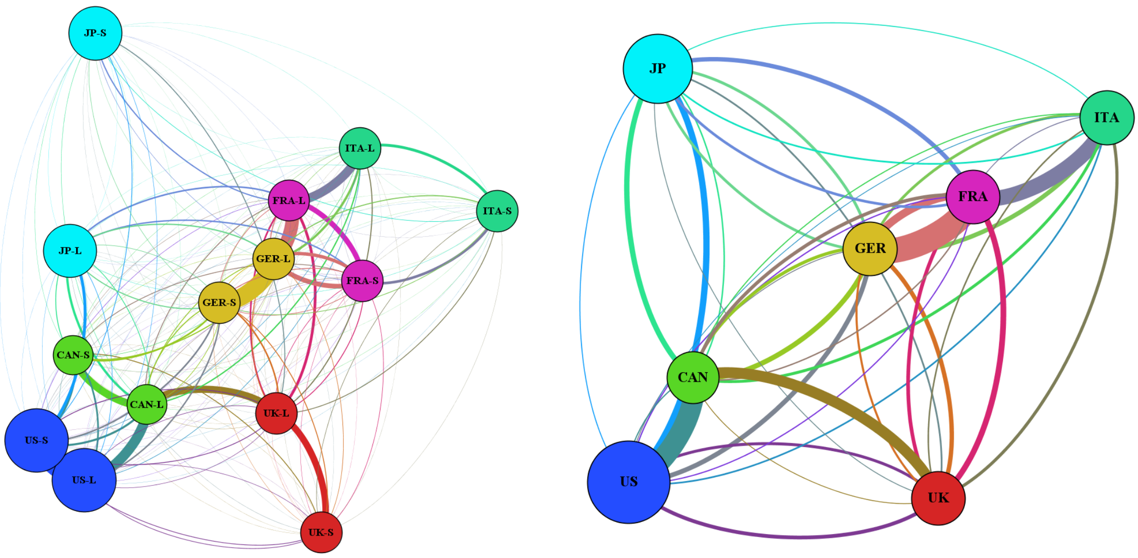
	From														Contribution	
To	ITA_L	ITA_S	CAN_L	CAN_S	FRA_L	FRA_S	GER_L	GER_S	JP_L	JP_S	UK_L	UK_S	US_L	US_S	From Others	
ITA_L	96.31	0.00	0.26	0.54	0.02	0.44	0.89	0.82	0.06	0.11	0.24	0.18	0.06	0.08	3.69	9.52
ITA_S	7.08	87.09	0.00	0.32	0.21	0.14	2.46	0.20	0.58	0.09	0.36	1.17	0.00	0.30	12.91	
CAN_L	3.14	0.48	90.17	0.28	0.98	0.17	0.47	1.40	1.79	0.05	0.53	0.02	0.27	0.24	9.83	18.76
CAN_S	0.95	0.10	19.22	71.57	0.08	0.45	2.70	4.17	0.00	0.01	0.00	0.02	0.71	0.03	28.43	
FRA_L	22.37	0.00	2.12	0.78	61.73	0.25	2.17	1.57	3.89	0.00	4.04	0.09	0.02	0.98	38.27	64.04
FRA_S	0.01	6.61	0.93	1.91	12.98	60.99	8.42	2.06	3.52	0.08	1.29	0.75	0.43	0.01	39.01	
GER_L	5.74	0.20	2.04	0.64	34.03	12.14	37.55	1.66	3.21	0.01	2.09	0.00	0.02	0.67	62.45	69.27
GER_S	0.08	1.63	2.42	0.22	0.16	1.57	43.48	48.04	1.28	0.06	0.30	0.67	0.04	0.05	51.96	
JP_L	0.31	0.65	5.76	2.17	0.12	0.01	0.02	2.63	87.05	0.13	0.03	0.53	0.58	0.01	12.95	21.86
JP_S	0.14	1.21	0.87	0.62	3.08	0.97	0.58	1.11	0.39	90.57	0.07	0.03	0.03	0.33	9.43	
UK_L	2.67	2.01	18.17	0.00	6.90	2.07	0.23	3.79	0.01	2.25	58.47	0.02	1.28	2.14	41.53	50.12
UK_S	0.02	0.38	0.04	2.09	0.00	1.22	1.26	2.20	0.12	0.02	16.47	74.92	1.10	0.17	25.08	
US_L	1.24	0.00	25.34	4.21	0.24	0.52	0.00	3.61	1.25	0.97	1.12	1.54	59.96	0.02	40.04	68.92
US_S	0.96	0.04	5.17	4.95	0.31	0.27	0.54	4.00	7.44	1.60	1.92	1.70	39.74	31.36	68.64	
Contribution	44.70	13.32	82.37	18.72	59.12	20.22	63.22	29.22	23.54	5.36	28.46	6.70	44.28	5.02	**444.22**	
To others	50.93		81.58		66.10		47.29		28.39		18.66		9.54			**302.48**
Contribution including own	141.01	100.41	172.53	90.29	120.85	81.21	100.77	77.25	110.59	95.94	86.93	81.62	104.23	36.39	Spillover Index	
															= 31.7%	= 24%

Notes: 1. This table summarizes the spillover table of the country-specific components among the Level and Slope factors of all countries: Italy (ITA), Canada (CAN), France (FRA), Germany (GER), Japan (JP), the UK and the US. Subscripts L and S are for Level and Slope factors respectively.
2. The underlying variance decomposition (reported in percentage) is based upon a monthly VAR of order 1, identified using a generalized variance decomposition. The (i, j)-th value is the estimated contribution to the variance of the 12-month-ahead forecast error of country-specific component i coming from innovations to the component j.
3. The spillover index is the cross variance share, i.e. the variance due to the shocks on $j, j \neq i$ relative to total forecast error variation of i. Two indexes that measure spillovers among components and countries are calculated, respectively. See Appendix A.5 for details.
4. The final two columns set out the fraction of movements in a component due to shocks on other components or in a country due to shocks on other countries, respectively. The final three rows respectively set out the contribution of a component to other components, a country to other countries or the total contribution to components including to its own.

According to the spillover table, we find that the movements in bond markets of France, Germany, UK and US are susceptible to changes in other countries. Moreover, the bond markets of Italy, Canada, France and Germany contribute to large proportions of changes in other countries. We find that around one fourth of the variance across home countries is due to the shocks on the country-specific components of foreign countries.

To have a more intuitive understanding of the asymmetric spillovers among the country-specific components, we use the evidence in Table 6 to plot the network graphs in Figure 4. This graph displays idiosyncratic connections based upon the distance and the thickness of connections.[31] We find that there are two main clusters: Europe and North America. The European markets are united, but the UK is relatively unconnected to Europe. In general the UK market has a similar link to European markets and the markets of North America. The Japanese market is also very independent, as the edges connected to the node of Japan are relatively thin and hence the node is further away from the other clusters. We can see there is a large distance between the US and Italy, and it seems there is no significant direct connection between these two countries. This means the country-specific components of Italy are not likely to directly affect the components of the US, implying there might not be strong spillover effects of the sovereign crisis from Italy to the US through a direct channel. However, if the sovereign crisis affects all European markets, the US market will also be influenced, but the spillovers may boil down to global co-movements that we have discussed in the last section.

Figure 4: Network of Global Spillovers



Notes: The left panel shows the spillovers of country-specific components among seven countries, which is constructed according to the results in Table 6. In the right panel, the spillover effects of the same country are grouped. The mnemonics are defined as in Table 6.

[31]The node location is determined by the *ForceAtlas2* algorithm of Jacomy et al. (2014): We assume that nodes repel each other, but edges attract the nodes according to average pairwise directional connectedness 'to' and 'from' in Table 6. The algorithm finds a steady state in which repelling and attracting forces reach a balance.

By construction, the directions and quantity of the 'spillover effects' contagion are identified, which can be used for further analysis of the network model. The measured asymmetries in contagion effects controlling for co-movement are by far new to related research of financial contagion.

6 Robustness

6.1 Excess Returns Explained by Yield Factors

Dahlquist and Hasseltoft (2013) indicate that the traditional Level, Slope and Curvature factors have relatively weak power in explaining the realized excess returns, which is also implied in Cochrane and Piazzesi (2005) and Cochrane and Piazzesi (2009). The results are puzzling as in an ordinary setup of affine term structure models, both yields and the price of risk are spanned by the yield factors, so a large share of term premia should be explained by the factors. With mild restrictions, the term premia can be spanned perfectly by the factors. If the model-implied term premia is a good predictor of the future holding period excess returns, then a proportion of the excess returns can be spanned by the factors. Particularly, if the factors are estimated using the full sample data, the explanatory power of factors should be sufficiently significant as the information of realized excess returns has been contained.

Based on the above inference, we conduct robustness checks in this section. Our proposed model, augmented with global macro information, successfully identifies global and national yield factors. The movements of these factors have meaningful economic implications. We consider excess bond returns and how much variability can be accounted for by the identified global or national factors. If we can explain the variance of excess returns using these factors, then our method is considered robust.

We begin by considering expected excess returns across countries. The expected excess return, i.e. the model-implied term premia, is the difference between the long-term yield and the average of expected future short rates.[32] To construct the expected excess return for each country, we use the model proposed in Bauer, Rudebusch and Wu (2012, 2014), which corrects the small sample bias in an arbitrage-free model. In total seven models are constructed, and in each model we only use the bond yields of a single country. We augment the factor dynamics in each country with country-specific macro information (i.e. monetary policy rate, CPI inflation and IP growth rate) and global FCI. Based upon our model construction, if our factors are correctly identified, the explanatory power for the term premia should be strong.[33]

We estimate 2-, 5- and 8-year term premia, and then we proceed by regressing the term premia on our identified factors for each country. The regression results are shown in Table 7 and 8. We find that the global factors have significant explanatory power for the expected excess returns across all countries, on average accounting for more than half of the total variance. When adding the country-specific components in yield factors, the vast majority of variance in term premia can be explained in almost all countries. The

[32] Technical details about the term premia can be found in Appendix A.3.1.

[33] In the arbitrage-free model we employ here, both the short and long yields are mainly driven by latent pricing factors, so approximately should the term premia.

results highlight the usefulness of our identified factors in driving the model-implied term premia.

We then proceed to assess the explanatory power for realized excess returns across countries. The one-period holding period return for a τ-period bond is given by

$$rx_{t+1}^{(\tau)} = p_{t+1}^{(\tau-1)} - p_t^{(\tau)} - y_t(1),$$

where $p_t^{(\tau)}$ is the log price of a zero bond, $p_t^{(\tau)} = -\tau y_t(\tau)$, and $y_t(1)$ is the one-period continuously compounded rate. Following Cochrane and Piazzesi (2005), we firstly construct the forward rate factor (henceforth denoted CP factor) and regress the realized one-year holding period excess return on the CP factor. The results are shown in Table 9. As Cochrane and Piazzesi (2005) and Dahlquist and Hasseltoft (2013) suggest, the explanatory power of the CP factor is unrelated to the traditional bond factors that may have relatively weak performance in explaining the realized excess returns. Table 9 confirms their findings: Adding the traditional Level and Slope factors extracted by the principal component method does not significantly improve the explanatory power in most cases, and most of the coefficients are not significant (not reported for the sake of brevity). We then add the identified global factors and country-specific components in national factors to the regressions to see whether the explanatory power is significantly increased. There are indeed significant improvements, especially in France.

Table 10 and 11 show that although adding the global factors can significantly increase the Adjusted R^2 in some countries, such as the US, it is not the case for all. It appears the information of the global factors is dominated by the CP factor, except in Italy where the coefficient of global Level is significantly negative, implying a 'flight-to-quality' demand. However, when we augment the regressions with the country-specific components in national yield factors, the explanatory power demonstrably increases: The Adjusted R^2 rises by up to 50%. We are therefore persuaded that the national factors capture useful information orthogonal to the CP factor in predicting the holding period excess returns. These results differentiate our identified factors from the traditional Level and Slope factors, as Cochrane and Piazzesi (2005) and Dahlquist and Hasseltoft (2013) suggest these traditional yield factors have relatively weak explanatory power in explaining excess returns. There are two major reasons for the enhanced explanatory ability: 1) The factors are augmented with macro information that can potentially increase the predictability of excess returns, as suggested by Ludvigson and Ng (2007); 2) Our unified modeling approach may better characterize the dynamics of the latent factors, as Duffee (2011) indicates that the dynamic factors estimated from more robust techniques can improve explanatory power.

From the above regressions, we observe that the fraction of the realized excess returns explained by the identified factors are generally lower than the fraction of expected excess returns. In fact, it is true even when the expected excess return is an unbiased predictor of the realized excess returns. The following endeavors to explain this empirical observation.

The relation between the τ-period model-implied term premia (expected excess re-

Table 7: Expected Excess Return Regressions (US, UK, Japan and Canada)

		2-Year		5-Year		8-Year	
US	Constant	1.53(0.17)	1.53(0.06)	1.52(0.17)	1.52(0.07)	1.52(0.17)	1.52(0.08)
	G_L	-0.40(0.14)	-0.44(0.04)	-0.34(0.14)	-0.38(0.05)	-0.31(0.14)	-0.35(0.05)
	G_S	-1.25(0.23)	-0.84(0.13)	-1.17(0.24)	-0.76(0.18)	-1.14(0.24)	-0.73(0.19)
	F_L^I		-1.04(0.31)		-1.06(0.38)		-1.05(0.40)
	F_S^I		-1.81(0.17)		-1.84(0.21)		-1.83(0.22)
	R^2	0.45	0.88	0.37	0.81	0.36	0.79
UK	Constant	0.85(0.19)	0.85(0.14)	0.86(0.21)	0.85(0.16)	0.86(0.21)	0.86(0.17)
	G_L	-0.24(0.16)	-0.29(0.13)	-0.11(0.20)	-0.16(0.17)	-0.09(0.21)	-0.14(0.17)
	G_S	-1.38(0.25)	-1.10(0.21)	-1.22(0.33)	-0.92(0.28)	-1.20(0.34)	-0.90(0.29)
	F_L^I		-0.39(0.94)		-0.10(1.08)		-0.06(1.11)
	F_S^I		-1.76(0.33)		-1.86(0.34)		-1.87(0.35)
	R^2	0.43	0.59	0.34	0.52	0.33	0.51
Japan	Constant	1.17(0.06)	1.17(0.01)	1.17(0.07)	1.17(0.01)	1.17(0.07)	1.17(0.01)
	G_L	0.54(0.06)	0.55(0.01)	0.58(0.07)	0.61(0.01)	0.60(0.07)	0.63(0.01)
	G_S	-0.12(0.07)	-0.07(0.02)	-0.09(0.08)	-0.03(0.01)	-0.08(0.08)	-0.02(0.01)
	F_L^I		0.57(0.02)		0.65(0.02)		0.68(0.02)
	F_S^I		0.03(0.01)		0.1(0.00)		0.11(0.00)
	R^2	0.83	0.99	0.80	1.00	0.79	1.00
Canada	Constant	1.53(0.14)	1.53(0.10)	1.54(0.15)	1.54(0.10)	1.54(0.15)	1.54(0.10)
	G_L	0.16(0.14)	0.20(0.09)	0.31(0.15)	0.35(0.12)	0.39(0.14)	0.42(0.12)
	G_S	-0.99(0.17)	-0.55(0.14)	-0.84(0.26)	-0.43(0.20)	-0.81(0.27)	-0.41(0.20)
	F_L^I		0.29(0.48)		0.69(0.45)		0.77(0.44)
	F_S^I		-0.76(0.13)		-0.62(0.13)		-0.6(0.12)
	R^2	0.46	0.69	0.44	0.64	0.47	0.66

Notes: 1. This table summarizes regression results for the following regression: $TP_t(\tau) = \beta_0^{(\tau)} + \beta_1^{(\tau)} G_{Lt} + \beta_2^{(\tau)} G_{St} + \beta_3^{(\tau)} F_{Lt}^I + \beta_4^{(\tau)} F_{St}^I + \epsilon_t^{(\tau)}$, where $TP_t(\tau)$ are expected excess returns (model-implied term premia) and $\tau = 2, 5, 8$ years. G_L and G_S are medium values of global Level and Slope; F_L^I and F_S^I are country-specific components in national Level and Slope. The sample is from 1984:12 to 2014:03 at monthly frequency.
2. The Newey and West (1987) standard errors are given in parentheses $(\cdot)$ and the Adjusted R_2 are reported.

Table 8: Expected Excess Return Regressions (Germany, France and Italy)

		2-Year		5-Year		8-Year	
Germany	Constant	1.53(0.07)	1.53(0.07)	1.54(0.07)	1.54(0.07)	1.54(0.07)	1.54(0.07)
	G_L	0.19(0.06)	0.18(0.06)	0.29(0.07)	0.29(0.07)	0.32(0.07)	0.32(0.07)
	G_S	-1.02(0.10)	-1.28(0.16)	-0.71(0.14)	-0.67(0.19)	-0.66(0.14)	-0.61(0.19)
	F_L^I		0.50(0.39)		-0.03(0.33)		-0.04(0.32)
	F_S^I		-0.75(0.36)		0.14(0.36)		0.21(0.36)
	R^2	0.73	0.74	0.68	0.67	0.68	0.68
France	Constant	1.56(0.09)	1.56(0.07)	1.56(0.09)	1.56(0.07)	1.57(0.10)	1.56(0.07)
	G_L	-0.23(0.10)	-0.20(0.07)	-0.12(0.11)	-0.09(0.07)	-0.11(0.11)	-0.08(0.07)
	G_S	-1.05(0.14)	-0.76(0.10)	-0.81(0.21)	-0.54(0.17)	-0.82(0.22)	-0.54(0.18)
	F_L^I		0.95(0.29)		0.92(0.3)		0.93(0.31)
	F_S^I		-0.57(0.32)		-0.53(0.3)		-0.55(0.30)
	R^2	0.54	0.71	0.37	0.54	0.36	0.54
Italy	Constant	1.98(0.16)	1.97(0.13)	1.99(0.14)	1.98(0.10)	1.99(0.15)	1.99(0.08)
	G_L	-0.51(0.19)	-0.30(0.14)	0.07(0.17)	0.11(0.10)	0.36(0.19)	0.37(0.08)
	G_S	-1.13(0.22)	-0.78(0.16)	-0.51(0.19)	-0.19(0.12)	-0.49(0.18)	-0.11(0.10)
	F_L^I		0.34(0.11)		0.66(0.10)		0.87(0.07)
	F_S^I		-0.56(0.28)		0.05(0.23)		0.21(0.16)
	R^2	0.45	0.67	0.18	0.55	0.39	0.82

Notes: 1. This table summarizes regression results for the following regression: $TP_t(\tau) = \beta_0^{(\tau)} + \beta_1^{(\tau)} G_{Lt} + \beta_2^{(\tau)} G_{St} + \beta_3^{(\tau)} F_{Lt}^I + \beta_4^{(\tau)} F_{St}^I + \epsilon_t^{(\tau)}$, where $TP_t(\tau)$ are expected excess returns (model-implied term premia) and $\tau = 2, 5, 8$ years. G_L and G_S are medium values of global Level and Slope; F_L^I and F_S^I are country-specific components in national Level and Slope. The sample is from 1984:12 to 2014:03 at monthly frequency.
2. The Newey and West (1987) standard errors are given in parentheses $(\cdot)$ and the Adjusted R_2 are reported.

Table 9: R^2 of Cochrane-Piazzesi Excess Return Regressions

	2-Year			5-Year			8-Year		
	CP	CP+PCs	CP+G,F	CP	CP+PCs	CP+G,F	CP	CP+PCs	CP+G,F
US	0.13	0.17	0.24	0.28	0.27	0.31	0.31	0.31	0.33
UK	0.04	0.07	0.18	0.24	0.24	0.30	0.31	0.31	0.34
Japan	0.75	0.85	0.85	0.70	0.70	0.71	0.52	0.54	0.55
Canada	0.27	0.35	0.36	0.41	0.40	0.40	0.40	0.42	0.42
Germany	0.20	0.25	0.48	0.32	0.32	0.47	0.34	0.36	0.44
France	0.26	0.32	0.72	0.39	0.39	0.73	0.44	0.46	0.71
Italy	0.49	0.51	0.64	0.65	0.65	0.69	0.64	0.64	0.67

Notes: 1. This table summarizes the Adjusted R^2 statistics for the following regression: $rx_{t+h}(\tau) = \beta_0^{(\tau)} + \beta_1^{(\tau)} CP_t + \epsilon_t^{(\tau)}$, $rx_{t+h}(\tau) = \beta_0^{(\tau)} + \beta_1^{(\tau)} CP_t + \beta_2^{(\tau)} PC_{Lt} + \beta_3^{(\tau)} PC_{St} + \epsilon_t^{(\tau)}$ or $rx_{t+h}(\tau) = \beta_0^{(\tau)} + \beta_1^{(\tau)} CP_t + \beta_2^{(\tau)} G_{Lt} + \beta_3^{(\tau)} G_{St} + \beta_4^{(\tau)} F^I_{Lt} + \beta_5^{(\tau)} F^I_{St} + \epsilon_t^{(\tau)}$, where $rx_{t+h}(\tau)$ are realized excess returns, $h = 12$ months (1-year holding period) and $\tau = 2, 5, 8$ years. For each country, PC_L and PC_S are traditional Level and Slope factors, respectively extracted by principal component; for each country, CP is the Cochrane and Piazzesi (2005) factor, which is constructed using one-year yield and 2- to 8-year forward rates. G_L and G_S are medium values of global Level and Slope; F^I_L and F^I_S are country-specific components in national Level and Slope. The sample is from 1984:12 to 2014:03 at monthly frequency.

turn) $TP_t(\tau)$ and the realized excess return rx is given by

$$TP_t(\tau) = y_t(\tau) - \frac{1}{\tau}\sum_{i=0}^{\tau-1} \mathrm{E}_t y_{t+i}(1) = \frac{1}{\tau}\mathrm{E}_t\Big(\sum_{i=0}^{\tau-1} rx_{t+i+1}^{(\tau-i)}\Big), \tag{6.1}$$

where $rx_{t+i+1}^{(\tau-i)}$ are one-period holding period excess returns at different time $t+i+1$, and $y_t(\tau)$ is the τ-period continuously compounded rate.

Alternatively, the realized one-period holding period excess return of an τ-maturity bond can be written as

$$rx_{t+1}^{(\tau)} = -(\tau - 1)\mathrm{E}_t\ TP_{t+1}^{\tau-1} + \tau TP_t^{(\tau)} + error. \tag{6.2}$$

Our results demonstrate that, even the first two terms of the right hand side of Equation 6.2 are fully spanned by some factors, if the variance of the error term is sufficiently large and not spanned by the same factors, then the explanatory power of the regressions on these factors will be largely decreased.

Table 10: Realized Excess Return Regressions (US, UK, Japan and Canada)

		2-Year		5-Year		8-Year	
US	Constant	0.12(0.35)	0.20(0.45)	-0.35(0.94)	0.06(1.23)	-1.05(1.11)	-0.23(1.45)
	CP_{US}	0.28(0.09)	0.25(0.11)	1.25(0.22)	1.08(0.33)	1.90(0.25)	1.58(0.46)
	G_L	0.24(0.13)	0.24(0.13)	-0.04(0.46)	-0.02(0.47)	-0.07(0.87)	-0.07(0.87)
	G_S	0.78(0.40)	0.75(0.39)	1.29(1.18)	1.35(1.38)	1.11(1.69)	0.85(2.05)
	F_L^I		0.00(0.64)		-0.79(2.31)		-0.26(3.44)
	F_S^I		-0.24(0.48)		-1.73(1.38)		-2.63(1.79)
	R^2	0.24	0.24	0.29	0.31	0.31	0.33
UK	Constant	0.20(0.54)	1.37(0.61)	-0.92(1.45)	0.88(1.74)	-1.58(1.65)	-1.23(2.43)
	CP_{UK}	0.20(0.19)	-0.32(0.22)	1.43(0.55)	0.65(0.71)	2.32(0.69)	2.19(1.06)
	G_L	0.10(0.13)	0.15(0.11)	-0.28(0.42)	-0.24(0.45)	-0.32(0.90)	-0.40(0.89)
	G_S	0.30(0.64)	-0.72(0.58)	1.26(1.79)	-0.41(1.84)	1.47(2.19)	0.90(2.52)
	F_L^I		1.14(1.11)		4.86(3.21)		8.18(5.33)
	F_S^I		-2.22(0.66)		-3.05(2.28)		0.23(3.70)
	R^2	0.05	0.18	0.25	0.30	0.31	0.34
Japan	Constant	-0.18(0.07)	0.02(0.07)	-0.16(0.25)	0.52(0.54)	0.33(0.77)	0.79(1.17)
	CP_{JP}	0.28(0.04)	0.16(0.04)	1.10(0.08)	0.70(0.33)	1.71(0.35)	1.44(0.71)
	G_L	-0.08(0.05)	0.12(0.06)	-0.18(0.20)	0.39(0.48)	-0.33(0.49)	-0.13(1.06)
	G_S	0.07(0.06)	0.04(0.05)	-0.19(0.29)	-0.26(0.34)	-0.90(0.58)	-0.91(0.68)
	F_L^I		0.12(0.07)		0.61(0.57)		0.78(1.34)
	F_S^I		0.15(0.03)		0.28(0.25)		-0.22(0.52)
	R^2	0.77	0.85	0.70	0.71	0.54	0.55
Canada	Constant	-0.21(0.36)	-0.42(0.40)	-0.35(1.17)	-0.36(1.11)	0.48(1.94)	0.96(1.55)
	CP_{CA}	0.38(0.11)	0.46(0.13)	1.15(0.35)	1.16(0.34)	1.37(0.56)	1.18(0.45)
	G_L	0.09(0.19)	-0.01(0.16)	-0.06(0.47)	-0.06(0.45)	0.03(0.76)	0.28(0.74)
	G_S	0.58(0.43)	0.49(0.47)	0.42(1.20)	0.43(1.25)	-1.09(1.93)	-0.8(1.87)
	F_L^I		0.42(0.81)		-0.18(1.89)		-2.43(2.62)
	F_S^I		0.53(0.28)		-0.04(0.67)		-1.66(0.89)
	R^2	0.31	0.36	0.40	0.40	0.40	0.42

Notes: 1. This table summarizes regression results for the following regression: $rx_{t+h}(\tau) = \beta_0^{(\tau)} + \beta_1^{(\tau)} CP_t + \beta_2^{(\tau)} G_{Lt} + \beta_3^{(\tau)} G_{St} + \beta_4^{(\tau)} F_{Lt}^I + \beta_5^{(\tau)} F_{St}^I + \epsilon_t^{(\tau)}$, where $rx_{t+h}(\tau)$ are realized excess returns, $h = 12$ months (1-year holding period) and $\tau = 2, 5, 8$ years. For each country, CP is the Cochrane and Piazzesi (2005) factor. G_L and G_S are medium values of global Level and Slope; F_L^I and F_S^I are country-specific components in national Level and Slope. The sample is from 1984:12 to 2014:03 at monthly frequency.
2. The Newey and West (1987) standard errors are given in parentheses $(\cdot)$ and the Adjusted R_2 are reported.

Table 11: Realized Excess Return Regressions (Germany, France and Italy)

		2-Year		5-Year		8-Year	
Germany	Constant	-0.04(0.38)	0.47(0.44)	-0.13(1.30)	0.93(1.46)	0.84(1.92)	1.71(2.06)
	CP_{GE}	0.25(0.12)	0.06(0.13)	1.07(0.41)	0.67(0.48)	1.35(0.61)	1.03(0.72)
	G_L	0.01(0.23)	0.22(0.26)	-0.49(0.65)	-0.08(0.76)	-0.60(0.92)	-0.29(1.02)
	G_S	0.08(0.40)	-0.59(0.47)	-0.62(1.20)	-1.21(1.61)	-2.18(1.62)	-1.32(2.36)
	F_L^I		2.35(0.64)		4.15(2.03)		2.22(2.97)
	F_S^I		0.94(0.96)		5.23(3.53)		9.91(5.01)
	R^2	0.20	0.48	0.34	0.47	0.37	0.44
France	Constant	-0.16(0.42)	0.18(0.27)	-0.97(1.18)	0.69(0.87)	-1.25(1.64)	1.70(1.37)
	CP_{FR}	0.35(0.11)	0.22(0.09)	1.40(0.33)	0.76(0.27)	2.10(0.50)	0.95(0.43)
	G_L	-0.09(0.15)	0.13(0.11)	-0.97(0.55)	0.08(0.33)	-1.42(0.91)	0.39(0.56)
	G_S	0.35(0.34)	0.11(0.30)	0.45(0.86)	-0.04(0.73)	0.37(1.30)	0.07(0.92)
	F_L^I		2.90(0.37)		8.96(1.03)		11.52(1.60)
	F_S^I		0.98(0.44)		0.46(1.10)		-2.69(1.61)
	R^2	0.27	0.72	0.43	0.73	0.48	0.71
Italy	Constant	0.22(0.34)	1.27(0.63)	-0.50(1.05)	0.10(2.05)	-0.25(1.60)	-0.71(2.85)
	CP_{IT}	0.32(0.04)	0.04(0.13)	1.19(0.13)	1.04(0.42)	1.62(0.23)	1.75(0.59)
	G_L	-0.44(0.13)	0.01(0.17)	-0.95(0.48)	-1.27(0.59)	-0.45(1.00)	-2.04(0.95)
	G_S	0.14(0.21)	0.09(0.27)	-0.07(0.84)	-0.25(0.95)	-0.85(1.41)	-1.19(1.5)
	F_L^I		1.81(0.67)		2.00(2.13)		1.57(3.00)
	F_S^I		0.90(0.42)		2.66(1.13)		4.81(1.72)
	R^2	0.54	0.64	0.66	0.69	0.64	0.67

Notes: 1. This table summarizes regression results for the following regression: $rx_{t+h}(\tau) = \beta_0^{(\tau)} + \beta_1^{(\tau)} CP_t + \beta_2^{(\tau)} G_{Lt} + \beta_3^{(\tau)} G_{St} + \beta_4^{(\tau)} F^I_{Lt} + \beta_5^{(\tau)} F^I_{St} + \epsilon_t^{(\tau)}$, where $rx_{t+h}(\tau)$ are realized excess returns, $h = 12$ months (1-year holding period) and $\tau = 2, 5, 8$ years. For each country, CP is the Cochrane and Piazzesi (2005) factor. G_L and G_S are medium values of global Level and Slope; F_L^I and F_S^I are country-specific components in national Level and Slope. The sample is from 1984:12 to 2014:03 at monthly frequency.
2. The Newey and West (1987) standard errors are given in parentheses $(\cdot)$ and the Adjusted R_2 are reported.

6.2 Non-Fundamental Change: Inflation Uncertainty and Regime Shifts

According to the results of the previous sections, we find that global inflation plays an important role in driving bond yield co-movement among all global fundamentals. However, we also observe the standard deviation of global inflation in variance decomposition is relatively high. Suppose we consider a Taylor-type policy rules, this evidence implies the response to inflation may be uncertain. As suggested in Ang et al. (2011) and Bikbov and Chernov (2013), potential regime shifts may lead the economy to be prone to non-fundamental sunspot fluctuations. Dai, Singleton and Yang (2007) argue that investors should require a premium for possible changes of future regimes even if there is no uncertainty about the state variables. This conjecture is further confirmed by the finding of Wright (2011) that inflation uncertainty is capable of explaining the excess return across countries.

We follow the above lead to conduct further robustness analysis. In a single-regime model structure, the information of the first moment has been captured by the factor dynamics. If investors truly require compensation for the uncertainty in the response to global inflation that we do not allow for, then second moment information of global inflation should be useful in explaining realized excess returns (but not the model-implied expected excess returns).

To capture the information of the second moment of the global information, we employ a simple GARCH$(1, 1)$ model to calculate the conditional volatility of all draws from our model.[34] The median value of the conditional volatilities is added as an additional explanatory variable to the full model regressions of expected or realized excess returns in the last section.

Table 13 displays the significant increases in explanatory power for realized excess returns across countries, especially for bonds at shorter maturities; Table 12 shows the increases in the expected excess returns are not as obvious as the realized, because the single-regime model used to construct the country-specific term premia does not explicitly allow for inflation uncertainty. In other words, inflation uncertainty seems to largely span the error term in Eq. (6.2).

In addition, we can see that the coefficients of the conditional volatility of inflation are all positive, which means the global inflation uncertainty increases the excess return across all countries as Wright (2011) suggests.[35] Earlier findings of Ang et al. (2011) and Bikbov and Chernov (2013) have related non-fundamental sunspot fluctuations to regime shifts, and our empirical evidence shown in this section cannot rule out this possibility from a global perspective. The conjecture still needs to be firmed by further analysis, but it implies shocks to global fundamentals accountable for a higher proportion in a regime switching model.

[34]Our results are not sensitive to the method used to capture conditional volatility. Our conclusion does not vary whether we use EGARCH$(1, 1)$ or the method similar to Bansal and Shaliastovich (2013) to conduct robustness checks.

[35]This empirical evidence is also consistent with the findings in Bansal and Yaron (2004) and Bansal and Shaliastovich (2013), although they argue that investors demand higher risk premia because of the Long-Run Risks in inflation.

Table 12: Expected Excess Return Regressions with Conditional Volatility of Inflation

		Expected Excess Return					
		2-Year		5-Year		8-Year	
		Factors	$+I_{CV}$	Factors	$+I_{CV}$	Factors	$+I_{CV}$
US	R^2	0.88	0.89	0.81	0.83	0.79	0.81
	$\beta_{I_{CV}}$		1.04(0.64)		1.25(0.79)		1.31(0.82)
UK	R^2	0.59	0.61	0.52	0.52	0.51	0.51
	$\beta_{I_{CV}}$		2.14(1.04)		1.16(0.88)		1.09(0.87)
Japan	R^2	0.99	0.99	1.00	0.99	1.00	1.00
	$\beta_{I_{CV}}$		-0.01(0.05)		-0.01(0.04)		-0.03(0.04)
Canada	R^2	0.69	0.70	0.64	0.64	0.66	0.67
	$\beta_{I_{CV}}$		1.26(0.8)		1(0.73)		0.94(0.7)
Germany	R^2	0.74	0.74	0.67	0.67	0.68	0.68
	$\beta_{I_{CV}}$		0.86(0.58)		0.7(0.51)		0.66(0.49)
France	R^2	0.71	0.72	0.54	0.54	0.54	0.53
	$\beta_{I_{CV}}$		0.91(0.57)		0.42(0.58)		0.35(0.58)
Italy	R^2	0.67	0.67	0.55	0.55	0.82	0.82
	$\beta_{I_{CV}}$		0.87(0.98)		0.87(0.72)		0.81(0.52)

Notes: 1. This table summarizes the regression results for the following regression: $TP_t(\tau) = \beta_0^{(\tau)} + \mathbb{B}^{(\tau)}\mathbb{F}_t + \beta_{I_{CV}}^{(\tau)} I_{CV\,t} + \epsilon_t^{(\tau)}$, where $TP_t(\tau)$ are expected excess returns (model-implied term premia) and $\tau = 2, 5, 8$ years. For each country, $\mathbb{F}$ include global and national Level and Slope factors; I_{CV} is the conditional volatility of global inflation calculated using a GARCH$(1, 1)$ model. The sample is from 1984:12 to 2014:03 at monthly frequency.
2. The Newey and West (1987) standard errors are given in parentheses $(\cdot)$ and the Adjusted R_2 are reported.

Table 13: Realized Excess Return Regressions with Conditional Volatility of Inflation

		One-Year Holding Period Excess Return					
		2-Year		5-Year		8-Year	
		Factors	$+I_{CV}$	Factors	$+I_{CV}$	Factors	$+I_{CV}$
US	R^2	0.24	0.28	0.31	0.35	0.33	0.35
	$\beta_{I_{CV}}$		2.47(1.12)		7.36(3.45)		6.55(6.01)
UK	R^2	0.18	0.27	0.30	0.34	0.34	0.37
	$\beta_{I_{CV}}$		3.97(1.49)		9.9(4.02)		12.96(6.15)
Japan	R^2	0.85	0.82	0.71	0.68	0.55	0.53
	$\beta_{I_{CV}}$		0.08(0.18)		1.52(1.13)		2.93(2.40)
Canada	R^2	0.36	0.40	0.40	0.41	0.42	0.42
	$\beta_{I_{CV}}$		2.59(0.98)		4.9(2.21)		3.87(3.32)
Germany	R^2	0.48	0.49	0.47	0.46	0.44	0.44
	$\beta_{I_{CV}}$		1.67(1.08)		4.45(3.84)		4.71(6.26)
France	R^2	0.72	0.74	0.73	0.74	0.71	0.71
	$\beta_{I_{CV}}$		1.79(0.74)		3.81(2.25)		3.74(3.61)
Italy	R^2	0.64	0.67	0.69	0.69	0.67	0.67
	$\beta_{I_{CV}}$		2.81(1.10)		5.77(3.51)		5.13(5.31)

Notes: 1. This table summarizes the regression results for the following regression: $rx_{t+h}(\tau) = \beta_0^{(\tau)} + \mathbb{B}^{(\tau)}\mathbb{F}_t + \beta_{CP}^{(\tau)}CP + \beta_{I_{CV}}^{(\tau)} I_{CV\,t} + \epsilon_t^{(\tau)}$, where $rx_{t+h}(\tau)$ are realized excess returns, $h = 12$ months (1-year holding period) and $\tau = 2, 5, 8$ years. $\mathbb{F}$ include global and national Level and Slope factors for each country; I_{CV} is the conditional volatility of global inflation calculated using a GARCH$(1,1)$ model; CP is the Cochrane and Piazzesi (2005) factor, which is constructed using one-year yield and 2- to 8-year forward rates. The sample is from 1984:12 to 2014:03 at monthly frequency.
2. The Newey and West (1987) standard errors are given in parentheses $(\cdot)$ and the Adjusted R_2 are reported.

7 Conclusion

We propose a new 'Fundamentals-Augmented Hierarchical Factor Model' to jointly identify global and national Level and Slope factors augmented with global fundamentals: inflation, real activity, changes in policy rate and financial conditions. Our method is considered robust: The identified global yield factors can significantly explain the variance in expected excess returns across countries and therefore, should be distinguished from the traditional Level and Slope factors that with relatively weak explanatory power.

Our approach identifies structural shocks of global fundamentals and non-fundamentals to global bond yields. Shocks to non-fundamentals are persistent and account for the majority of global term structure movement. The global inflation, among all macro fundamentals, contributes the most to the long yield co-movements through both the policy channel and risk compensation channel, but the effects are offset though these two channels. Hence, the transmission mechanism of global fundamentals explains the failure of term structure models with merely shocks to fundamentals. Moreover, we find that the global inflation uncertainty can significantly explain realized excess returns across countries, which means it is possible that regime shifts serve as a source of non-fundamental fluctuations.

There are many possible avenues for future work. We do not affirm that shocks to non-fundamentals are exclusively due to sentiment changes, as Novy-Marx (2014) investigates several additional sources. We have applied a convenient procedure to identify the asymmetric 'spillover effects' mainly caused by divergence in policy rates, but it may also be interesting to specifically evaluate whether the contagion across different countries is driven by additional macroeconomic information. Lastly, this paper does not explicitly model potential time-varying nonlinear dynamics of yield factors such as regime shifts. Allowing for nonlinearity can be promising in unfolding various sources of non-fundamental fluctuations.

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Appendix A Econometric Methods

In this paper we propose a novel approach which extends the hierarchical factor model of Moench, Ng and Potter (2013) by augmenting the model with macro factors. We apply the NS restrictions similar to Diebold, Li and Yue (2008) for the yield factor identification. The estimation of our model is in one step, which should provide more accurate estimates when compared to other multi-step estimations. We call the new model 'Fundamentals-Augmented Hierarchical Factor Model' (FAHFM).

Our proposed hierarchical model has three levels of factor dynamics, but we only focus on the global level that are augmented with global macro factors. At the global level, the dynamics of the global yield factors can be regarded as an unrestricted Factor-Augmented Vector Autoregressive (FAVAR) system. To provide economic interpretations for the estimated global yield factors, we employ the identification schemes in Kurmann and Otrok (2013) to identify the desired shocks in the FAVAR.

We conduct the analysis in two steps. The first step is to extract the latent global yield factors, using the proposed 'Fundamentals-Augmented Hierarchical Dynamic Factor Model'. The second step is to directly use the estimation results of FAVAR at the global level to identify the shocks of interest. We believe our proposed approach can shed light on the transmission mechanism between global bond yield factors and macro variables.

A.1 Fundamentals-Augmented Hierarchical Factor Model

To extract the latent factors, a principal component method is commonly utilized. Bai and Ng (2006) have shown that the estimated factors from the principal components method can be treated as though they are observed, if $\sqrt{T}/N \to \infty$ as $T, N \to \infty$. However, the method of principal components is not well suited for the present analysis, because the number of series available[36] is much smaller than the large dimensions that the principal component method typically requires. Accordingly, the FAHFM is proposed to extract the latent global factors.

A.1.1 A Three-Level Hierarchical Factor Model

Following the framework developed by Moench, Ng and Potter (2013), a three-level model is considered here. Level one is the national level, which describes how national yield factors drive the yields at different maturities. Level two is the global-national level, illustrating how the global yield factors govern the national yield factors. Level three displays the autoregressive dynamics of the global factors.

Firstly, we treat a block (identified as b) as one of the seven countries, so $b = 1, 2, ..., B$ where $B = 7$. At the national level, the bond yield data for a specific country are stacked in the vector X_{bt}, and the dynamic representation is given by

$$X_{b,t} = \Lambda_b^F F_{b,t} + e_{b,t}^X, \qquad \text{(A.1)}$$

where $X_{b,t}$ is an $N_b \times 1$ vector of yields of country b at different maturities, $F_{b,t}$ is a $k_b \times 1$ vector of latent common yield factors at national level, Λ_b^F is an $N_b \times k_b$ coefficient matrix

[36]There are only seven countries so $N = 7$.

and $e^X_{b,t}$ is the vector of idiosyncratic components. Note that in our model $N_b = 11$ and $k_b = 2$ for $b = 1, 2, ..., B$; in other words, for each country, we use yield data of 11 different maturities and assume that 2 factors can explain most of the yield variance.

Stacking up $F_{b,t}$ across seven countries produces a $K^F \times 1$ vector F_t.[37] At the global-national level, it is assumed that

$$F_t = \Lambda^G G_t + e^F_t, \tag{A.2}$$

where K^G global common factors are collected into the vector G_t, Λ^G is a $K^F \times K^G$ coefficient matrix and e^F_t are country-specific components at the global-national level.

The dynamics of the global factors G_t are described at level three:

$$G_t = \Psi^G G_{t-1} + \epsilon^G_t, \tag{A.3}$$

where Ψ^G is the coefficient matrix and the innovations $\epsilon^G_t \sim N\Big(\mathbf{0}, \Sigma^G\Big)$[38].

The model is completed by specifying the dynamics of idiosyncratic and country-specific components $e^X_{b,t}$ and e^F_t.

$$e^X_{b,t} = \Psi^X_b e^X_{b,t-1} + \epsilon^X_{b,t}, \tag{A.4}$$
$$e^F_t = \Psi^F e^F_{t-1} + \epsilon^F_t, \tag{A.5}$$

where Ψ^X_b is an $N_b \times N_b$ diagonal coefficient matrix, Ψ^F is a $K^F \times K^F$ diagonal coefficient matrix, the innovations $\epsilon^X_{b,t} \sim N\Big(\mathbf{0}, \Sigma^X_b\Big)$ and $\epsilon^F_t \sim N\Big(\mathbf{0}, \Sigma^F\Big)$.[39]

A.1.2 An Extension with Macro Factor Augmentation

Assuming at level three, i.e. the level that describes the global factor dynamics, the factor dynamics are augmented with Macro information. So the Equation (A.3) can be rewritten as

$$\begin{bmatrix} G_t \\ M_t \end{bmatrix} = \psi^G \begin{bmatrix} G_{t-1} \\ M_{t-1} \end{bmatrix} + u_t, \tag{A.6}$$

$$u_t \sim N\Big(\mathbf{0}, \Sigma^G\Big),$$

where Σ^G is the variance-covariance matrix of u_t. The evolution of the global factors displayed here uses only one lag here for simplicity; in practice, more lags can be used to estimate the factor dynamics. The Equation (A.6) is indeed a factor-augmented vector autoregressive (FAVAR) system. The estimates from this system will be used for the identification of shocks for the structural analysis.

[37] $K^F = \sum_{b=1}^{B} k_b$ and $F_t = \Big(F_{1,t} \quad F_{2,t} \quad ... \quad F_{B,t}\Big)'$

[38] $\Sigma^G = \text{diag}\big((\sigma^G_1)^2, ..., (\sigma^G_{K^G})^2\big)$.

[39] $\Sigma^X_b = \text{diag}\big((\sigma^X_{b,1})^2, ..., (\sigma^X_{b,N_b})^2\big)$ and $\Sigma^F = \text{diag}\big((\sigma^F_1)^2, ..., (\sigma^F_{K^F})^2\big)$.

A.1.3 Estimation via Gibbs Sampling

Before we proceed with the estimation scheme, the parameters needed to be estimated are summarised for better illustration. Collect $\{\Lambda_1^F, ..., \Lambda_B^F\}$ and Λ^G into $\mathbf{\Lambda}$, $\{\Psi_1^X, ..., \Psi_B^X\}$, Ψ^F and Ψ^G into $\mathbf{\Psi}$, and $\{\Sigma_1^X, ..., \Sigma_B^X\}$, Σ^F, Σ^G into $\mathbf{\Sigma}$. To sum up, the parameters we need to estimate are $\mathbf{\Lambda}$, $\mathbf{\Psi}$ and $\mathbf{\Sigma}$.

A Bayesian method, i.e., Markov Chain Monte Carlo (MCMC), is used to estimate the model. A simple extension of the algorithm in Carter and Kohn (1994) is proposed here. Based on the observed values of M_t, and the initial values of $\{F_{b,t}\}$ and G_t from the method of principal components, for each iteration we construct the Gibbs sampler in the following steps:

1. Draw G_t, conditional on F_t, $\mathbf{\Lambda}$, $\mathbf{\Psi}$ and $\mathbf{\Sigma}$.
2. Draw Ψ^G, conditional on Σ^G, G_t and M_t.
3. Draw Σ^G, conditional on Ψ^G, G_t and M_t.
4. Draw Λ^G, conditional on G_t and F_t.
5. For each b, draw $F_{b,t}$, conditional on $\mathbf{\Lambda}$, $\mathbf{\Psi}$, $\mathbf{\Sigma}$ and G_t.
6. For each b, draw b_{th} elements of Ψ^F and Σ^F, conditional on G_t and F_t.
7. For each b, draw the Λ_b^F, Ψ_b^X and Σ_b^X, conditional on F_t and $X_{b,t}$.

Similar to Diebold, Li and Yue (2008) and Moench, Ng and Potter (2013), the elements of $\mathbf{\Lambda}$ and $\mathbf{\Psi}$ are set to have normal priors, and $\mathbf{\Sigma}$ follow inverse gamma priors. Given the conjugacy, the posterior distributions are not difficult to compute. Regarding the factors G_t and F_t, we follow Carter and Kohn (1994) and Kim and Nelson (1999) to run the Kalman filter forward to obtain the estimates in period T and then proceed backward to generate draws for $t = T-1, ..., 1$. It is worth noting that, if we impose hard restrictions on Λ^G and Λ_b^F, then there is no need to draw these parameters in the above Gibbs sampling.

A.2 Nelson-Siegel Restrictions

Following Diebold, Li and Yue (2008) we use two factors summarize most of the information in the term structure of interest rates. In the following, we describe a three-factor Nelson-Siegel model for generalization, but we only use the first two in our application, i.e. we do not include the third factor 'Curvature' in Equation (A.7). As we show in the Section 3.1, two factors have accounted for around 99% of the bond yield variance across all countries.

The below Equation (A.7) describes how restrictions are imposed; the restrictions used in our hierarchical factor model are in fact fixing the loading of the factors. Let $y_t(\tau)$ denote yields at maturity τ, then the factor model for a single country we use is of the form

$$y_t(\tau) = L_t^{NS} + \frac{1-e^{-\tau\lambda}}{\tau\lambda} S_t^{NS} + \left(\frac{1-e^{-\tau\lambda}}{\tau\lambda} - e^{-\tau\lambda}\right) C_t^{NS} + \varepsilon_t(\tau), \tag{A.7}$$

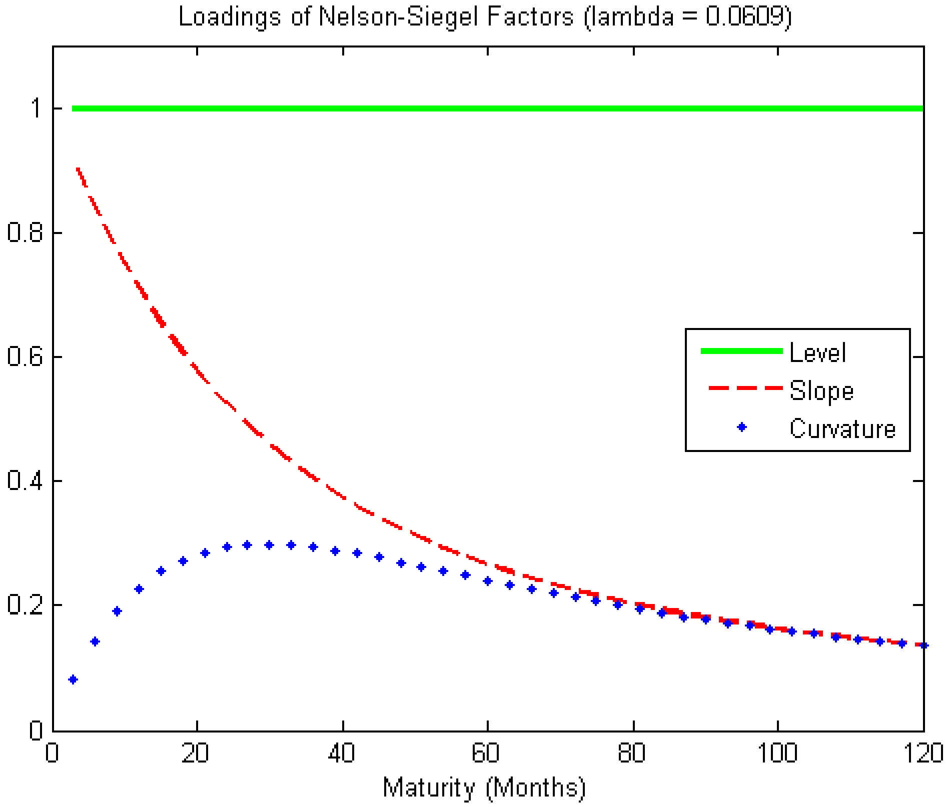
where L_t^{NS} is the "Level" factor, S_t^{NS} is the "Slope" factor, C_t^{NS} is the "Curvature" factor and ε_t is the error term. Additionally, λ in the exponential functions controls the shapes of loadings for the NS factors; following Diebold and Li (2006) and Bianchi, Mumtaz and Surico (2009), we set $\lambda = 0.0609$.[40]

The interpretations of Nelson-Siegel factors are of empirical significance. The Nelson-Siegel Level factor L_t^{NS} is identified as the factor that is loaded evenly by the yields of all maturities. The Slope factor S_t^{NS} denotes the spread between the yields of a short- and a long-term bond, and its movements are captured by putting more weights on the yields with shorter maturities. The Curvature factor C_t^{NS} focuses on changes that have their largest impact on medium term maturity yields, and therefore, is loaded more heavily by bonds of medium-term maturities; particularly, if $\lambda = 0.0609$, the C_t^{NS} has the largest impact on the bond at 30-month maturity.

The following Figure 5 depicts the shapes of the loadings of the NS factors. In our model estimation, we fixed the Λ_b^F in Equation (A.1) by the NS loadings. We further set the Λ^G in Equation (A.2) to a diagonal matrix to identify the global factors, and the intuition behind is that the country-level Level (Slope, Curvature) factor is only driven by the global Level (Slope, Curvature) factor.

[40] Alternatively, we can select the value of λ from a grid of reasonable values by comparing the goodness of fit. However, if we do not specify the factor dynamics and fit the Nelson-Siegel model in a static way, the selection may not be optimal. Also we choose a single value of λ for all the countries, as Nelson and Siegel (1987) indicate that there is little gain in practice by fitting λ individually. Therefore, we set $\lambda = 0.0609$ to fix the ideas because 1) this value is the mostly used in the related literature so revealing the dynamics the associate latent factors is more desirable, and 2) using this value we have a relatively better fit of the 'global short rate factor'. To ensure the robustness, we also try a grid of reasonable values; we find the results are qualitatively similar and hence our findings are robust to the selection of λ.

Figure 5: Loadings of Nelson-Siegel Factors



Notes: The solid green line, red dashed line and blue dotted line are the loadings for Level, Slope and Curvature factors, respectively ($\lambda = 0.0609$). The horizontal axis shows the maturities of bonds, and the unit is month.

A.3 Decomposition of Variance Driven by Global Factors

Recall Equation (A.3) that describes the dynamics of the global factors G_t at level three in Section A.1:

$$G_t = \Psi^G G_{t-1} + \epsilon_t^G,$$

We can rewrite this as an implied Wold MA(∞) representation:

$$G_t = \sum_{i=0}^{\infty} \psi_i \mu_{t-i}, \tag{A.8}$$

where μ_t are the orthogonal innovations and Cholesky decomposition is needed to take into account the contemporaneous correlation of the shocks.[41]

[41]The ordering of our global VAR system is the following: Industrial Production growth rate, inflation rate, change of policy rate, Level, Slope and FCI. The sequence of the first three variable is standard in the related literature, for example Christiano, Eichenbaum and Evans (2005). These three are followed by the financial variables Level, Slope and FCI, so the financial variables can react to the contemporaneous macro shocks in the first three variables. The Level, Slope and FCI are placed lower in the ordering

With simple algebra, we can write the bond yield co-movements driven by the global factors X_t^G as the following equation:

$$X_t^G = B \sum_{i=0}^{\infty} \psi_i \mu_{t-i}, \tag{A.9}$$

where B is the product of the loadings Λ^F (in Equation A.1) and Λ^G (in Equation A.2). The impulse response at time $t+h$ is therefore:

$$X_{t+h}^G = B \sum_{i=0}^{\infty} \psi_i \mu_{t+h-i}. \tag{A.10}$$

It is easy to have the error of the optimal h-step ahead forecast at time t:

$$X_{t+h}^G - \hat{X}_{t+h|t}^G = B \sum_{i=0}^{h-1} \psi_i \mu_{t+h-i}. \tag{A.11}$$

The mean squared error of X_{t+h}^G is given by

$$\text{MSE}(X_{t+h}^G) = \text{diag}\Big(B(\sum_{i=0}^{h-1} \psi_i \psi_i')B'\Big). \tag{A.12}$$

Therefore, the contribution of the kth factor to the MSE of the h-step ahead forecast of the yield at the jth maturity is

$$\Omega_{jk,h} = \sum_{i=0}^{h-1} R_{jk,i}^2 / \text{MSE}(X_{t+h}^G), \tag{A.13}$$

where $R_{jk,i}$ is the element in row j, column k of $R_i = B\psi_i$.

A.3.1 Decomposition of Policy Channel and Risk Compensation Channel

The policy channel is consistent with the 'Expectation Hypothesis' (EH). The EH consistent long yield is given by

$$y_t(\tau)^{EH} = \frac{1}{\tau} \sum_{i=0}^{\tau-1} \text{E}_t y_{t+i}(1), \tag{A.14}$$

where $y_t(\tau)$ is the element of yield data X_t at maturity τ. That is to say, the EH consistent long yield is equal to the average of expected short yields $\text{E}_t y_{t+i}(1)$. If we only focus on the part driven by global factors, then after some iterations, the above equation can be written as

$$y_t(\tau)^{EH} = \frac{1}{\tau} B(I + \Psi^G + \Psi^{G^2} + \ldots + \Psi^{G^{\tau-1}}) \sum_{i=0}^{\infty} \psi_i \mu_{t-i}. \tag{A.15}$$

because Hubrich, D'Agostino et al. (2013) argue that the monetary policy only react to asset price movements if there are prolonged, while the bond yields react immediately to policy change.

The term premia (risk compensation channel) is given by

$$TP_t(\tau) = y_t(\tau) - y_t(\tau)^{EH}. \tag{A.16}$$

In other words, the term premia is the difference between the long yield and the EH consistent long yield. We can use similar transformations as in Equations (A.10) and (A.13) to compute the impulse response and variance decomposition of the above two channels.

A.4 Identification of Current Shocks and Long-Run Shocks in a FAVAR

In the second step, we proceed the structural analysis with the shock identification schemes similar to Kurmann and Otrok (2013). With the estimates of the FAVAR (Equation (A.6)), we can immediately identify the shocks we are interested in.

The identification schemes in Kurmann and Otrok (2013) are extensions of the method proposed by Uhlig (2004). The idea of Uhlig (2004) is to extract the largest 1 or 2 shocks that explain the maximal amount of the prediction variance of the global yield factors. This identification can evaluate how the shocks influence other macro indicators, and therefore it is possible to reveal the economic meaning of these global yield factors.

The schemes used in Kurmann and Otrok (2013) are also closely related to Barsky and Sims (2011). If we are interested in the long- and short-run effects of macro shocks, we can introduce the restrictions used by Barsky and Sims (2011) to separately identify the 'News Shocks' (i.e. the 'Long-Run Shock') and the 'Current Shocks' of global yield factors.[42]

To describe the above identification schemes, we firstly rewrite Equation (A.6) into a vector moving average representation of a reduced-form VAR:

$$Y_t = C(L)u_t, \tag{A.17}$$

where $Y_t = \begin{bmatrix} G_t \\ M_t \end{bmatrix}$ is a vector of m variables in the FAVAR at time t, and u_t is a vector of prediction errors with covariance matrix Σ^F. $C(L) = I + \psi_1^F L + \psi_2^F L^2 + ...$ is a lag polynomial, where L is the lag operator.

Generally, identification of the structural shocks is to find a mapping A between the prediction errors u_t and a vector of mutually orthogonal shocks ε_t, i.e., $u_t = A\varepsilon_t$. It needs to satisfy the key restriction that $\Sigma^F = E[A\varepsilon_t\varepsilon_t'A']$. More restrictions are needed to identify A as there exists some alternative matrix $\tilde{A}$ that also satisfies $\Sigma^F = \tilde{A}\tilde{A}'$. In other words, we can find some matrix $\tilde{A}$ such that $\tilde{A}Q = A$, where Q is an orthonormal matrix, to identify another vector of mutually orthogonal shocks $\tilde{\varepsilon}$ such that $u_t = \tilde{A}\tilde{\varepsilon}_t$.

[42] The 'Long-Run Shocks' for yield factors are conceptually similar to the 'News Shocks' described in Barsky and Sims (2011) and Kurmann and Otrok (2013), though 'News Shocks' are usually used to describe Total Factor Productivity. In the remaining of this section we only use the term 'News Shock' to avoid confusion.

Therefore, we set $\tilde{A}$ as the Cholesky decomposition of Σ^F,[43] and then identification reduces to choosing an orthonormal matrix Q.

Uhlig (2004) proposes an approach aiming to find the matrix Q in which the first $n < m$ columns (recap that m is the number of variables) that meet the identification restrictions. That is to say, the first n columns in Q define the n mutually orthogonal shocks that can explain most of the Fraction of Forecast Error Variance (FEV) of some variable in Y_t over forecast horizon $\underline{k}$ to $\overline{k}$. Formally, the k-step ahead forecast error of the ith variable $y_{i,t}$ in Y_t is given by

$$y_{i,t+k} - \mathbf{E}_t y_{i,t+k} = \mathbf{e}_i' \Big[\sum_{l=0}^{k-1} C_l \tilde{A} Q \varepsilon_{t+k-l} \Big] \tag{A.18}$$

where $\mathbf{e}_i$ is a selection vector with 1 in the ith position and zeros elsewhere, and $\mathbf{E}$ is the expectation operator. We then need to solve

$$Q_n^* = \arg\max_{Q_n} \mathbf{e}_i' \Big[\sum_{k=\underline{k}}^{\overline{k}} \sum_{l=0}^{k-1} C_l \tilde{A} Q_n Q_n' \tilde{A}' C_l' \Big] \mathbf{e}_i \tag{A.19}$$

subject to $Q_n' Q_n = I$, where Q_n contains the columns of Q defining the n most important shocks. It can be shown that the above problem be formulated as a principal components problem. The desired n columns of Q_n correspond to the eigenvectors associated with the n largest eigenvalues of the transformed objective in the objective function. More details can be found in Uhlig (2004) or the online Appendix of Kurmann and Otrok (2013).

Extending the Uhlig (2004)'s approach, we can identify the current shocks and news shock of our desired variables, which are the global yield factors in this paper. The following identification schemes are the same as Barsky and Sims (2011) and Kurmann and Otrok (2013).

Assume that the movements of target variable follow the following exogenous moving average process:

$$y_{i,t} = v(L)\varepsilon_t^{current} + d(L)\varepsilon_t^{news}, \tag{A.20}$$

where $\varepsilon_t^{current}$ and ε_t^{news} are orthogonal innovations, and $v(L)$ and $d(L)$ are lag polynomials.

The above process implies that in a VAR with the desired variable ordered first, the $\varepsilon_t^{current}$ is identified as the shock associated with the first column of the matrix $\tilde{A}$ obtained from a Cholesky decomposition. The current shock identification is not sensitive to the order of variables entering the VAR system. With arbitrage ordering, the one-period maximization problem will equate the identified current shock to the first column of the matrix $\tilde{A}$ obtained from the Cholesky decomposition with the target variable ordered first (the order of the rest of variables can also be arbitrage).

The only restriction on the news shock ε_t^{news} is $d(0) = 0$. We are trying to identify the innovation that explains all remaining variation in the desired variable that is orthogonal

[43] Uhlig (2004) indicates that arbitrary decomposition would work equally well, and choosing Cholesky decomposition is only for computational convenience.

to the contemporaneous shock. As mentioned before, the method of Uhlig (2004) is extended to achieve the goal. Based on a VAR with the desired variable y_i ordered first and Cholesky decomposition $\tilde{A}$, we can select the column q of Q that maximizes the FEV of y_i over forecast horizon $\underline{k}$ to $\overline{k}$, subject to $q'q = 1$ and $q(1) = 0$.

Intuitively, the current shocks are identified as the contemporary shocks that only affect the desired variable at the beginning (time t) and do not account for the lagged expected movements that stem from the changes of other forward-looking variables (time $t+1$ or later). In contrast, the news shocks materialize the anticipated changes in the desired variable only at time $t+1$ or later and have no contemporaneous effects at time t.

A.5 Spillover Table and Generalized Variance Decomposition

The generalized variance decomposition (GVD) framework of Koop, Pesaran and Potter (1996) produces variance decompositions invariant to ordering. The GVD approach accounts for correlated shocks using the historically-observed error distribution, under a normality assumption. The GVD matrix has entries

$$\delta_{ij} = \frac{\delta_{jj}^{-1} \sum_{h=0}^{H-1} (e_i' A_h \Sigma e_j)^2}{\sum_{h=0}^{H-1} (e_i' A_h \Sigma A_h' e_i)} \tag{A.21}$$

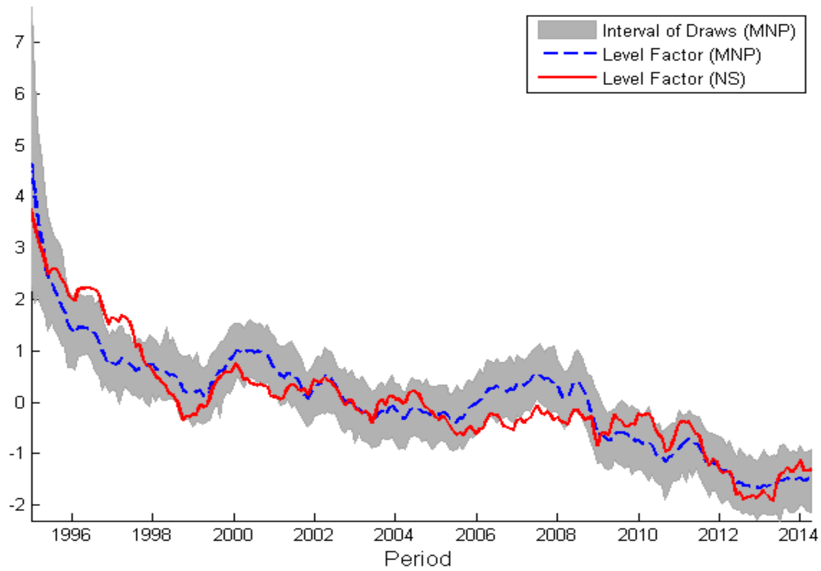
where δ_{jj} is the standard deviation of VAR shock ε_j, Σ is the covariance matrix of VAR shocks, A_h are MA(∞) coefficient matrices and e_j is a selection vector with jth element unity and zeros elsewhere. It means that shocks to variable j are responsible for $100 \times \delta_{ij}$ percent of the H-step-ahead forecast error variance in variable i.

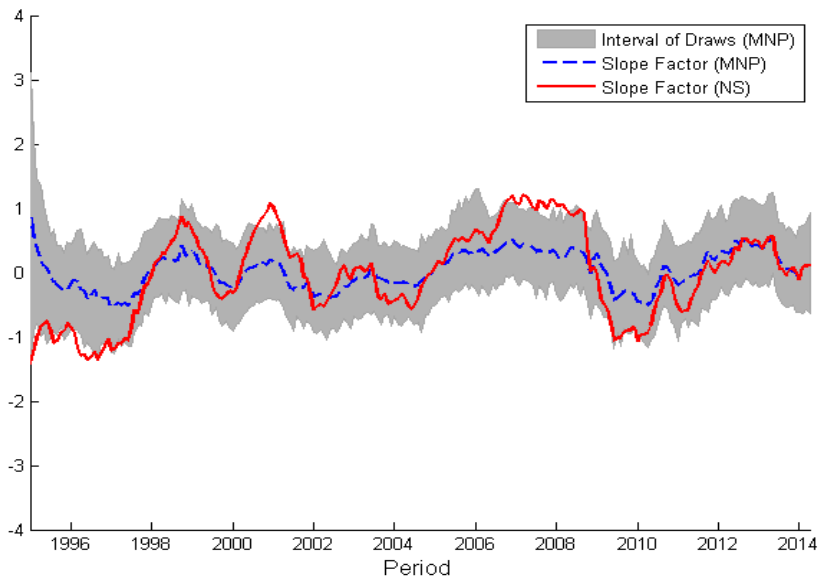
Because shocks are not necessarily orthogonal in the GVD environment, sums of forecast error variance contributions, i.e. row sums in GVD matrices, are not necessarily unity. Therefore, the (i,j)-th entry in the spillover table is given by $100 \times \widetilde{\delta}_{ij} = 100 \times \frac{\delta_{ij}}{\sum_{j=1}^{N} \delta_{ij}}$, where N is the number of shocks. The Spillover Index is calculated from

$$SOI = \frac{\sum\limits_{\substack{i,j=1 \\ i \neq j}}^{N} \widetilde{\delta}_{ij}}{\sum_{i,j=1}^{N} \widetilde{\delta}_{ij}}. \tag{A.22}$$

Appendix B Comparison of Factor Indentifiaction Schemes

Figure 6: Identified Factors from Different Schemes (MNP vs. NS)





Notes:
1. In the above two charts, the factors identified by the scheme of Moench, Ng and Potter (2013) are plotted against the factors identified by the NS scheme of Diebold, Li and Yue (2008). To better serve the comparison purpose, the factors are extracted from a less complicated system without macro factor augmentation.
2. The upper chart shows the Level factors, while the lower chart displays the Slope factor. The dashed blue lines are the median values of MNP identified factors and the gray areas cover all the draws from the MCMC estimation. The solid red lines are the median values of NS identified factors.

Appendix C Data Appendix

Table 14: List of Yields

Series ID	Description
ITA	Italy Sovereign (IYC 40) Zero Coupon Yields [1]
CAN	Canada Sovereign (IYC 7) Zero Coupon Yields [1]
FRA	France Sovereign (IYC 14) Zero Coupon Yields [1]
GER	German Sovereign (IYC 16) Zero Coupon Yields [1]
JP	Japan Sovereign (IYC 18) Zero Coupon Yields [1]
UK	United Kingdom (IYC 22) Zero Coupon Yields [1]
US	Treasury Actives (IYC 25) Zero Coupon Yields [1]

Notes:
1. In square brackets [·] we have a code for data transformations used in this data set: [1] means original series is used. The series are not seasonally adjusted.
2. Data are attained from Bloomberg, spanning from Dec. 1994 to Mar. 2014. The yields are of the following 11 maturities: 3 months, 6 months, 1 year, 2 years, 3 years, 4 years, 5 years, 6 years, 7 years, 8 years and 10 years.

Table 15: List of Financial Condition Indexes

Series ID	Description
STLFSI	St. Louis Fed Financial Stress Index [1]
KCFSI	Kansas City Financial Stress Index [1]
ANFCI	Chicago Fed Adjusted National Financial Conditions Index [1]
CFSI	Cleveland Financial Stress Index [1]
VIX	CBOE S&P Volatility Index [1]
BFCIUS	Bloomberg United States Financial Conditions Index [1]
BFCIEU	Bloomberg Euro-Zone Financial Conditions Index [1]
GFSI	BofA Merrill Lynch Global Financial Stress Index [1]
EASSF	Euro Area Systemic Stress Indicator Financial Intermediary [1]
WJF	Westpac Japan Financial Stress Index [1]
GSF	Goldman Sachs Financial Index [1]
BCF	Bank of Canada Financial Conditions Index [1]

Notes:
1. In square brackets [$\cdot$] we have a code for data transformations used in this data set: [1] means original series is used. The series are not seasonally adjusted.
2. Data are attained from Bloomberg, spanning from Jan. 1990 to Mar. 2014. The data may be unbalanced. The first five series can also be attained from St. Louis Federal Reserve Economic Data (http://research.stlouisfed.org/).

Table 16: List of Real Activity Indicators

Series ID	**Description**
IMFIPUS	IMF US Industrial Production SA [5]
IMFIPUK	IMF UK Industrial Production SA [5]
IMFIPJP	IMF Japan Industrial Production SA [5]
IMFIPGER	IMF Germany Industrial Production SA [5]
IMFIPFR	IMF France Industrial Production SA [5]
IMFIPITA	IMF Italy Industrial Production SA [5]
IMFIPCAN	IMF Canada Industrial Production SA [5]

Notes:
1. In square brackets [$\cdot$] we have a code for data transformations used in this data set: [5] means log first-order difference (annually growth rate) is used.
2. Data are attained from Bloomberg, spanning from Jan. 1990 to Mar. 2014. The data may be unbalanced.

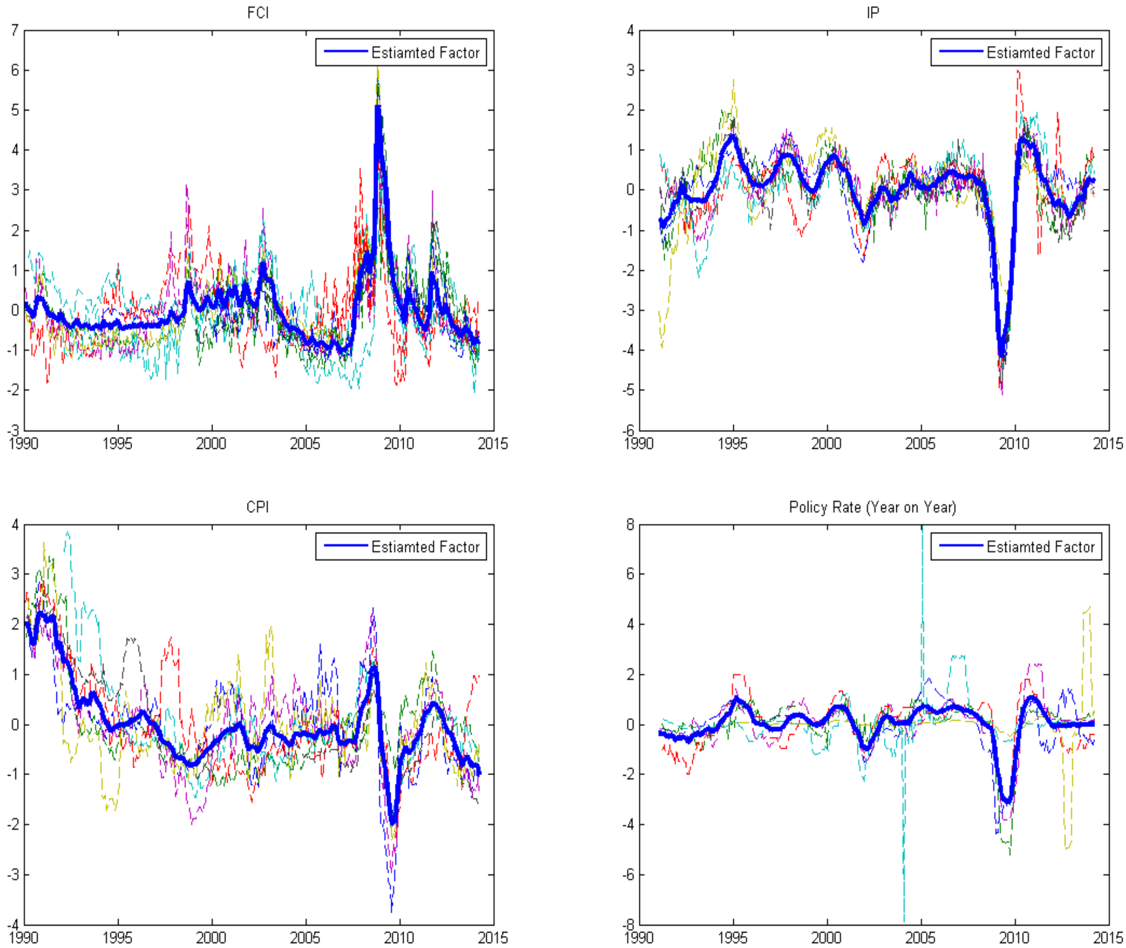
Table 17: List of CPI and Policy Rates

Series ID	**Description**
IMFCPIUS	IMF US CPI % Change in Percent per Annu [1]
IMFCPIUK	IMF UK CPI % Change in Percent per Annu [1]
IMFCPIJP	IMF Japan CPI % Change in Percent per Annu [1]
IMFCPIGER	IMF Germany CPI % Change in Percent per Annu [1]
IMFCPIFR	IMF France CPI % Change in Percent per Annu [1]
IMFCPIITA	IMF Italy CPI % Change in Percent per Annu [1]
IMFCPICAN	IMF Canada CPI % Change in Percent per Annu [1]
IMFFUNDUS	IMF US Federal Funds Rate in Percent per Annu [5]
IMFFUNDUK	IMF UK Bank of England Official Bank Rate [5]
IMFFUNDJP	IMF Japan Official Rate in Percent per Annu [5]
IMFFUNDCAN	IMF Canada Official Rate in Percent per Annu [5]
IMFFUNDEU	IMF Euro Area Official Rate in Percent per Annu [5]

Notes:
1. In square brackets [·] we have a code for data transformations used in this data set: [1] means original series is used. The series are all seasonally adjusted; [5] means log first-order difference (annually) is used.
2. Data are attained from Bloomberg, spanning from Jan. 1990 to Mar. 2014. The data may be unbalanced.

Appendix D Global Macro Factors

Figure 7: Estimated Global Macro Factors



Notes:
1. In the above charts, the thick blue lines are the global macro factors, which are estimated using the method proposed by Koop and Korobilis (2014). The Matlab code can be obtained in website https://sites.google.com/site/dimitriskorobilis/matlab/. The other thin lines with different colors are the standardized series for the estimation.
2. From top left clock-wise we have global factors of financial condition indexes, real activity, policy rates and inflation. The data used for the factor estimation are described in Appendix C, spanning from Jan. 1990 to Mar. 2014.

Table 18: Correlations between the National Series and Global Factors

FCI	STLFSI	KCFSI	ANFCI	CFSI	VIX	BFCIUS	BFCIEU	GFSI	EASSF	WJF	GSF	BCF
Correlation	0.945	0.952	0.568	0.695	0.845	0.935	0.848	0.866	0.701	0.528	0.671	0.814

IP	IMFIPUS	IMFIPUK	IMFIPJP	IMFIPGER	IMFIPFR	IMFIPITA	IMFIPCAN
Correlation	0.899	0.889	0.767	0.831	0.940	0.946	0.731

CPI	IMFCPIUS	IMFCPIUK	IMFCPIJP	IMFCPIGER	IMFCPIFR	IMFCPIITA	IMFCPICAN
Correlation	0.805	0.810	0.761	0.525	0.887	0.891	0.739

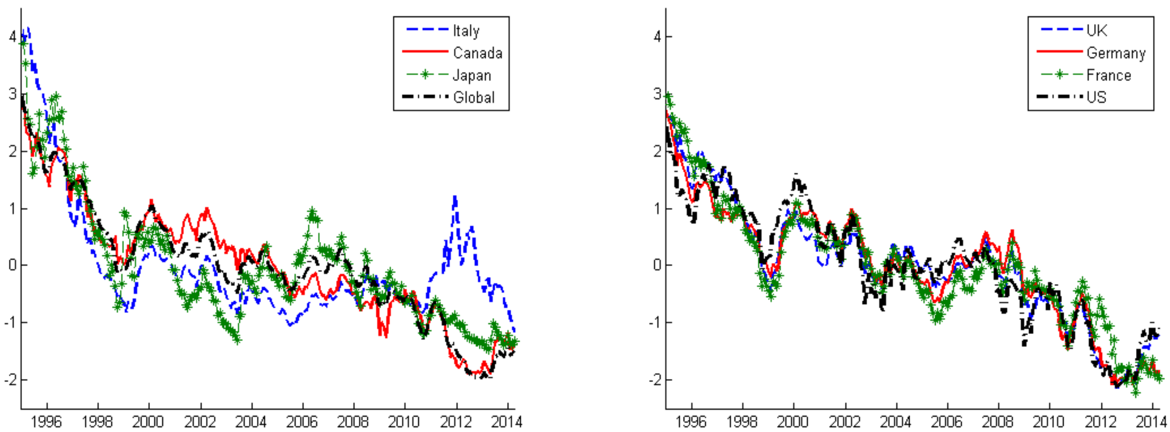
PR	IMFFUNDUS	IMFFUNDUK	IMFFUNDJP	IMFFUNDCAN	IMFFUNDEU
Correlation	0.844	0.911	0.330	0.914	0.073

Notes: This table summarizes the correlations between the national macro series in Data Appendix and the global macro factors shown in Figure 7, for four categories: Financial Condition Index, Industrial Production growth rate, CPI and the change (YoY) of policy rate.

Appendix E Commonality of Level and Slope

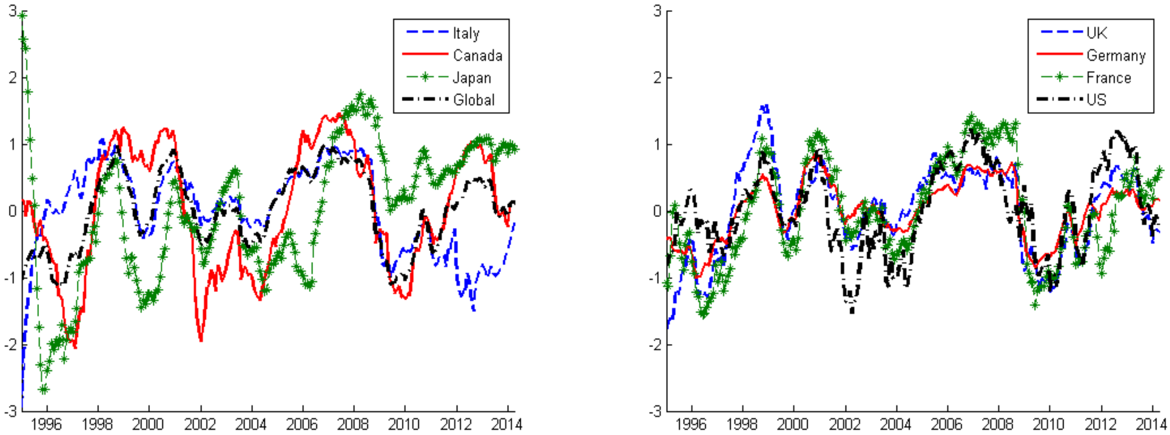
We firstly plot our identified Level and Slope factors in Figure 8 and 9, respectively, in order to evaluate the commonalities in country-level yield factors. The Slope factors are relatively less persistent than the Level factors. From the figures it is evident that a strong co-movement in Level factor dynamics exists, but some also exists for the Slope. We also calculate the communality statistics for all countries in Table 19 to better quantify matters. That is we calculate the proportion of national level or slope factor explained by the global equivalent. This indicates that the commonality in Level factor dynamics is stronger but co-movement remains in the Slope. Generally, we find significant co-movement among Germany, France, Canada, UK and US. In contrast, the Level and Slope factors of Italy are relatively more divorced from the global factors, consistent with Table 3 above; the Japanese Slope factor is much less common among all Slope factors as the communality statistic is nearly zero. The above findings are reassuringly in line with the results in Diebold, Li and Yue (2008).

Figure 8: Estimated Global and National Level factors



Notes: The left chart shows the median values of global Level factor and the national Level factors of Italy, Canada and Japan. The right chart shows the median values of the national Level factors of the UK, Germany, France and the US.

Figure 9: Estimated Global and National Slope factors



Notes: The left chart shows the median values of global Slope factor and the national Slope factors of Italy, Canada and Japan. The right chart shows the median values of the national Slope factors of the UK, Germany, France and the US.

Table 19: Communality Table of Level and Slope

Level		Slope	
Country	Communality	Country	Communality
Italy	0.45	Italy	0.24
Canada	0.94	Canada	0.35
France	0.94	France	0.67
Germany	0.94	Germany	0.91
Japan	0.80	Japan	0.04
UK	0.98	UK	0.77
US	0.90	US	0.51
Average	0.85	**Average**	0.50

Notes: This table summarizes for all countries the communality statistics of global Level and Slope factors for national Level and Slope factors. For example, the communality for a given country is interpreted as the proportion of the variation in the national Level factor explained by the global Level factor. Likewise for the Slope communality.

Appendix F Variance Decomposition across Maturities

Table 20: Decomposition of Variance (US)

Maturity	Posterior Mean (Standard Deviation)		
(Month)	$Share_G$	$Share_F$	$Share_X$
3	0.65(0.08)	0.32(0.08)	0.02(0.01)
6	0.68(0.08)	0.32(0.08)	0.01(0.00)
12	0.71(0.08)	0.29(0.08)	0.00(0.00)
24	0.74(0.07)	0.26(0.07)	0.01(0.00)
36	0.76(0.07)	0.24(0.07)	0.01(0.00)
48	0.77(0.07)	0.22(0.07)	0.01(0.00)
60	0.78(0.07)	0.22(0.06)	0.00(0.00)
72	0.79(0.06)	0.21(0.06)	0.00(0.00)
84	0.79(0.06)	0.21(0.06)	0.00(0.00)
96	0.79(0.06)	0.21(0.06)	0.01(0.00)
120	0.78(0.07)	0.20(0.06)	0.03(0.01)

Notes: This table summarizes the decomposition of variance for the three-level hierarchical model of US bond yields. $share_G$, $share_F$ and $share_Z$ denote the variance shares at different maturities in the country-level block due to shocks of ϵ_G, ϵ_F and ϵ_X, respectively. In each parentheses $(\cdot)$ the posterior standard deviation of shares in a specific block is calculated from our draws, see Section 2. Larger standard deviation means higher uncertainty in the estimates, but we do not have an exact credible interval interpretation as the statistics do not necessarily follow (truncated) normal distributions.

Table 21: Decomposition of Variance

Maturity	UK			Germany			France		
(Month)	$Share_G$	$Share_F$	$Share_X$	$Share_G$	$Share_F$	$Share_X$	$Share_G$	$Share_F$	$Share_X$
3	0.80(0.06)	0.20(0.06)	0.01(0.00)	0.70(0.08)	0.23(0.06)	0.07(0.02)	0.66(0.08)	0.27(0.07)	0.07(0.02)
6	0.81(0.06)	0.19(0.06)	0.00(0.00)	0.71(0.08)	0.23(0.06)	0.06(0.02)	0.70(0.08)	0.28(0.07)	0.03(0.01)
12	0.83(0.05)	0.17(0.05)	0.01(0.00)	0.72(0.08)	0.23(0.07)	0.05(0.02)	0.73(0.07)	0.27(0.07)	0.00(0.00)
24	0.84(0.05)	0.14(0.05)	0.01(0.00)	0.75(0.08)	0.23(0.07)	0.02(0.01)	0.75(0.07)	0.24(0.07)	0.01(0.00)
36	0.86(0.05)	0.13(0.04)	0.01(0.00)	0.76(0.07)	0.23(0.07)	0.01(0.00)	0.77(0.07)	0.22(0.06)	0.02(0.00)
48	0.88(0.04)	0.12(0.04)	0.00(0.00)	0.77(0.07)	0.23(0.07)	0.00(0.00)	0.78(0.06)	0.21(0.06)	0.01(0.00)
60	0.89(0.04)	0.11(0.04)	0.00(0.00)	0.77(0.07)	0.22(0.07)	0.01(0.00)	0.79(0.06)	0.20(0.06)	0.00(0.00)
72	0.89(0.04)	0.11(0.04)	0.00(0.00)	0.76(0.07)	0.21(0.07)	0.02(0.01)	0.80(0.06)	0.20(0.06)	0.00(0.00)
84	0.88(0.04)	0.10(0.04)	0.01(0.00)	0.75(0.08)	0.21(0.06)	0.04(0.01)	0.80(0.06)	0.20(0.06)	0.00(0.00)
96	0.86(0.04)	0.10(0.03)	0.04(0.01)	0.73(0.08)	0.20(0.06)	0.06(0.02)	0.80(0.06)	0.19(0.06)	0.01(0.00)
120	0.80(0.06)	0.09(0.03)	0.11(0.03)	0.71(0.08)	0.19(0.06)	0.10(0.03)	0.78(0.06)	0.18(0.05)	0.04(0.01)

Notes: This table summarizes the decomposition of variance for the three-level hierarchical model of bond yields. For each country, $share_G$, $share_F$ and $share_Z$ denote the variance shares at different maturities in the country-level block due to shocks of ϵ_G, ϵ_F and ϵ_X, respectively. In each parentheses $(\cdot)$ the posterior standard deviation of shares in a specific block is calculated.

Table 22: Decomposition of Variance

Maturity	Italy			Canada			Japan		
(Month)	$Share_G$	$Share_F$	$Share_X$	$Share_G$	$Share_F$	$Share_X$	$Share_G$	$Share_F$	$Share_X$
3	0.31(0.09)	0.66(0.09)	0.03(0.01)	0.52(0.10)	0.36(0.08)	0.12(0.03)	0.5(0.10)	0.44(0.09)	0.06(0.01)
6	0.32(0.10)	0.67(0.09)	0.01(0.00)	0.57(0.09)	0.36(0.08)	0.07(0.02)	0.54(0.10)	0.43(0.09)	0.03(0.01)
12	0.34(0.10)	0.66(0.10)	0.00(0.00)	0.63(0.09)	0.35(0.08)	0.02(0.01)	0.60(0.09)	0.39(0.09)	0.02(0.00)
24	0.35(0.10)	0.64(0.10)	0.00(0.00)	0.70(0.08)	0.30(0.08)	0.00(0.00)	0.65(0.08)	0.31(0.08)	0.04(0.01)
36	0.36(0.10)	0.63(0.10)	0.00(0.00)	0.74(0.07)	0.26(0.07)	0.00(0.00)	0.69(0.08)	0.28(0.07)	0.03(0.01)
48	0.37(0.10)	0.62(0.10)	0.00(0.00)	0.76(0.07)	0.24(0.07)	0.00(0.00)	0.72(0.07)	0.26(0.07)	0.02(0.00)
60	0.38(0.10)	0.62(0.10)	0.00(0.00)	0.77(0.07)	0.23(0.07)	0.00(0.00)	0.75(0.07)	0.25(0.07)	0.01(0.00)
72	0.38(0.10)	0.61(0.10)	0.00(0.00)	0.78(0.06)	0.22(0.06)	0.00(0.00)	0.76(0.07)	0.24(0.07)	0.00(0.00)
84	0.39(0.10)	0.61(0.10)	0.01(0.00)	0.79(0.06)	0.21(0.06)	0.00(0.00)	0.76(0.07)	0.23(0.06)	0.01(0.00)
96	0.39(0.10)	0.60(0.10)	0.01(0.00)	0.79(0.06)	0.21(0.06)	0.00(0.00)	0.75(0.07)	0.22(0.06)	0.02(0.01)
120	0.39(0.10)	0.59(0.10)	0.02(0.00)	0.79(0.06)	0.20(0.06)	0.01(0.00)	0.73(0.07)	0.21(0.06)	0.06(0.02)

Notes: This table summarizes the decomposition of variance for the three-level hierarchical model of bond yields. For each country, $share_G$, $share_F$ and $share_Z$ denote the variance shares at different maturities in the country-level block due to shocks of ϵ_G, ϵ_F and ϵ_X, respectively. In each parentheses $(\cdot)$ the posterior standard deviation of shares in a specific block is calculated.

Appendix G Shock Identifications of Global Yield Factors

G.1 Current and Long-Run Shocks of Yield Factors

As we mentioned in the previous sections, there are no significant conclusions on the underlying macro drivers. To our knowledge, none of current literature constructs a comprehensive analysis in identifying the structural shocks of global yield factors to explore the intrinsic mechanism. In the following sections, we seek to fill the gaps in previous research. Similar to method in Barsky and Sims (2011) and Kurmann and Otrok (2013), we can identify the 'current/contemporaneous shocks' and 'long-run shocks' of global yield factors, which explain the short-run and long-run movements, respectively. We would like to be clear about the nomenclature employed here. The contemporaneous shock, i.e. the current shock, is the shock that directly affects a variable only at current time, propagates forward and gradually dies out, but itself does not affect the future variance of the variable though the reactions of other forward-looking variables. The long-run shock, on the contrary, does not have current influence on a variable, but the impact will realize in the future through the immediate reactions of other forward-looking variables to the current 'news'. In our global-level VAR, a contemporaneous shock and a long-run shock as identified above can explain the vast majority of global yield factor variance.[44]

G.1.1 Current Shocks

The first important finding is that the contemporaneous shocks on the yield factors are not likely to be driven or transmitted to the global macroeconomic fundamentals. According to our identification, we find that the contemporaneous shocks on the bond factors are very persistent, which implies Long Memory in bond yields. The contemporaneous shocks of one yield factor may also persistently affect the other yield factor (see Figure 10 and 11), but these shocks can hardly explain any variance in macro factors, and hence, have no effects on the real economy. However, the current shock on the Slope factor accounts slightly more variance of macro factors than the current shock on Level factor, which implies that the Slope factor is relatively more sensitive to macro changes.

It is worth considering the reason why the current movements of global macro factors can hardly be accounted by the current shocks that span the current movements of global Level or Slope factor. Several explanations can be offered for this observation. From a financial perspective, these instantaneous shocks on the bond market are purely disturbances within the bond market, and hence, have no spillovers to the real economy. Conversely, a macroeconomic shock on the bond yields might be offset through different channels, so it will not account for much of the variance of the bond yield factors. From an economic viewpoint, the current shocks on the yield factors may be caused by a temporary monetary policy without inertia, or an expected exogenous process, so it would not have permanent real effects on the real economy.[45] Therefore, the current shocks on global yield

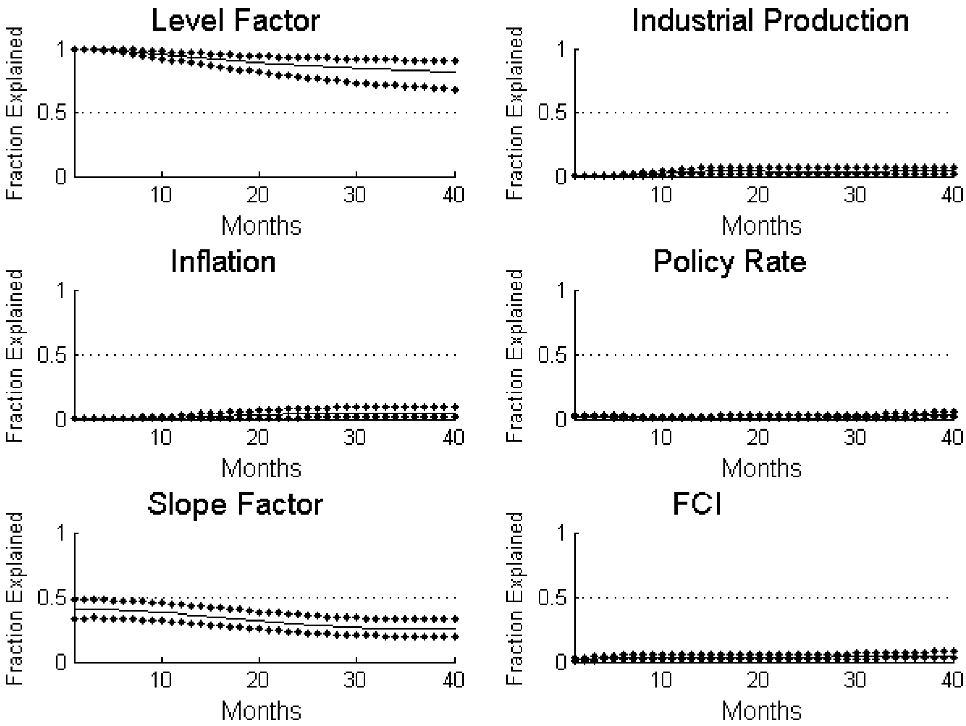
[44] As indicated in Kurmann and Otrok (2013), at a given point in time, a variable in the VAR system can move for three possible reasons. First, a contemporaneous shock hits. Second, past changes in the variable innovations propagate forward to affect the current value of the variable. Third, the shocks realize in the long run through the presence of forward-looking variables that react immediately to 'news'. In the context of global yield factor movements, we can separately identify the 'current shock' that collects the first two kinds of shocks above, and the last kind of shocks defined as the 'long-run shock'. The long-run shocks, which have no effect on a variable at current time t, would affect the variable at time $t+1$ or later. The technical details are discussed in Appendix A.4.

[45] For example, if the Level factor truly contains the information of inflation level, the current shock can be interpreted as an exogenous process expected by the agents, for instance, a downward trend of inflation expectation in the period of 'Great Moderation'; this shock will anchor the Level movements

factors are not related to the movements of other global macroeconomic fundamentals.

However, the above does not mean bond yield movements are not at all related to the macro factors. The other forward-looking macro factors can propagate the shocks on yield factors in the long run, which we successfully identify in the next section.

Figure 10: Fraction of Forecast Error Variance (FEV) Explained by Level Current Shock



Notes: The solid lines in the above panels are the posterior median values. The dotted lines indicate 16 to 84 percent posterior coverage intervals.

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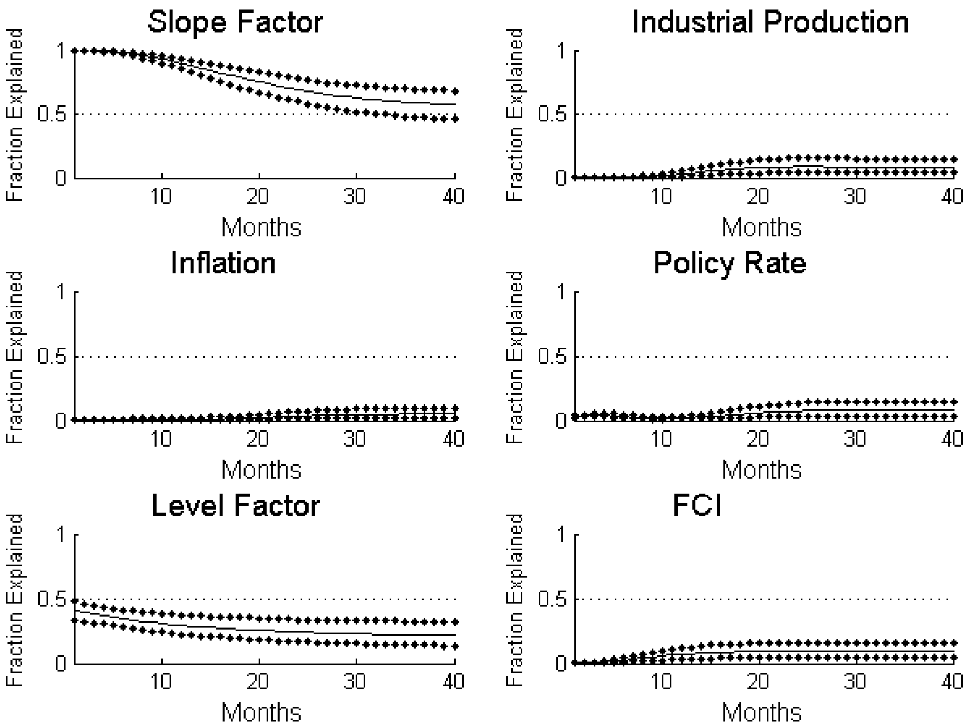
but is not likely to affect real global economy. For the Slope factor, as suggested by Wu (2001), it has a strong correlation with monetary policy surprises, so the Slope current shock can be regarded as a temporary monetary policy, for instance, a temporary liquidity release to meet the seasonal demand of banks. This shock would not affect the real economy, which is in line with the theory from the new Keynesian perspective.

Figure 11: Fraction of Forecast Error Variance (FEV) Explained by Slope Current Shock



Notes: The solid lines in the above panels are the posterior median values. The dotted lines indicate 16 to 84 percent posterior coverage intervals.

G.1.2 Long-Run Level and Slope Shocks are Inflation Shocks

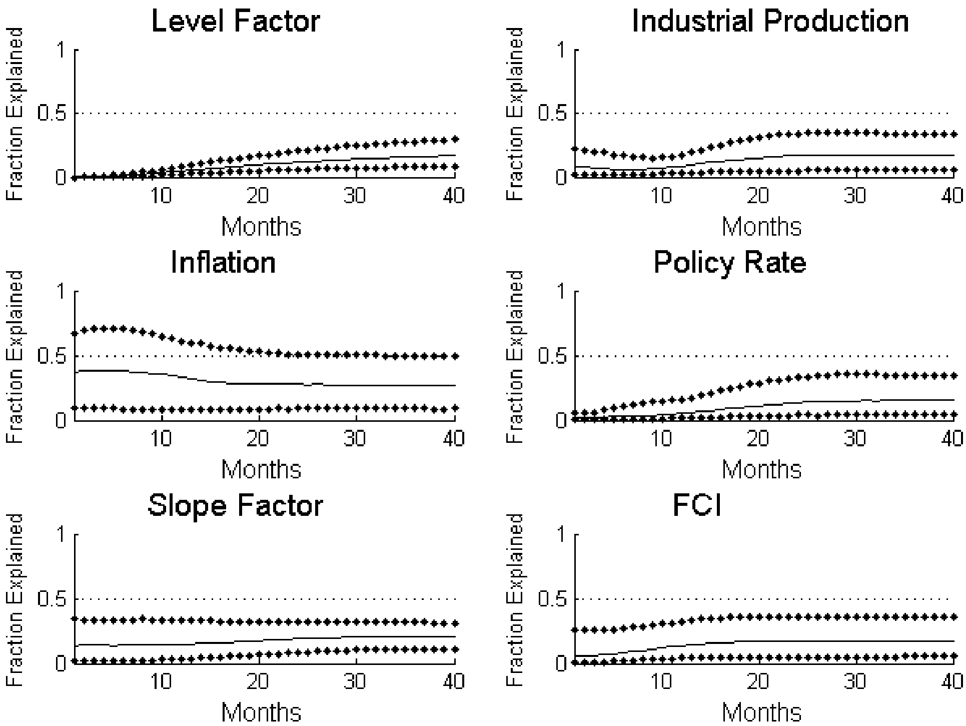
Next, we start analyzing the 'long-run shocks' of global yield factors. Long-run shocks are the innovations on a variable that are generally anticipated by market participants, but with lags between the recognition of the innovations and the eventual impact on the variable.[46] In a traditional view, asset prices are forward-looking variables, so they will react to news immediately. However, the presence of other forward-looking variables, such as inflation and FCI, may cause future fluctuations in asset prices, especially for bond yields that are more associated with the real economy. That is to say, the current reactions of inflation or FCI to some news can cause long-term movements in global yield factors.

Our empirical results show that the long-run shocks of global factors exist and should not be overlooked. As the Figure 12 shows, the effect of long-run Level shock on the economy is not negligible. The top left panel indicates the fraction of the long-run shock is continuously increasing and finally reaches 17% of the FEV at the end of 40-month

[46]Beaudry and Portier (2006) and Schmitt-Grohé and Uribe (2012) indicate that forward-looking agents will, in general, react to news about future changes in different fundamentals by adjusting some forward-looking variables. The anticipated changes in exogenous fundamentals may need sufficient time to materialize. In a VAR context, it is possible to identify the shocks as described in the Appendix A.4.

forecast horizon. In particular, this shock also accounts for almost 40% FEV of inflation at the short end of forecast horizon and 30% at the long end. This shock does not account for much variance of other global Macro variables. Therefore, this shock can be identified as an 'inflation shock'.

Figure 12: Fraction of Forecast Error Variance (FEV) Explained by Long-Run Level Shock



Notes: The solid lines in the above panels are the posterior median values. The dotted lines indicate 16 to 84 percent posterior coverage intervals.

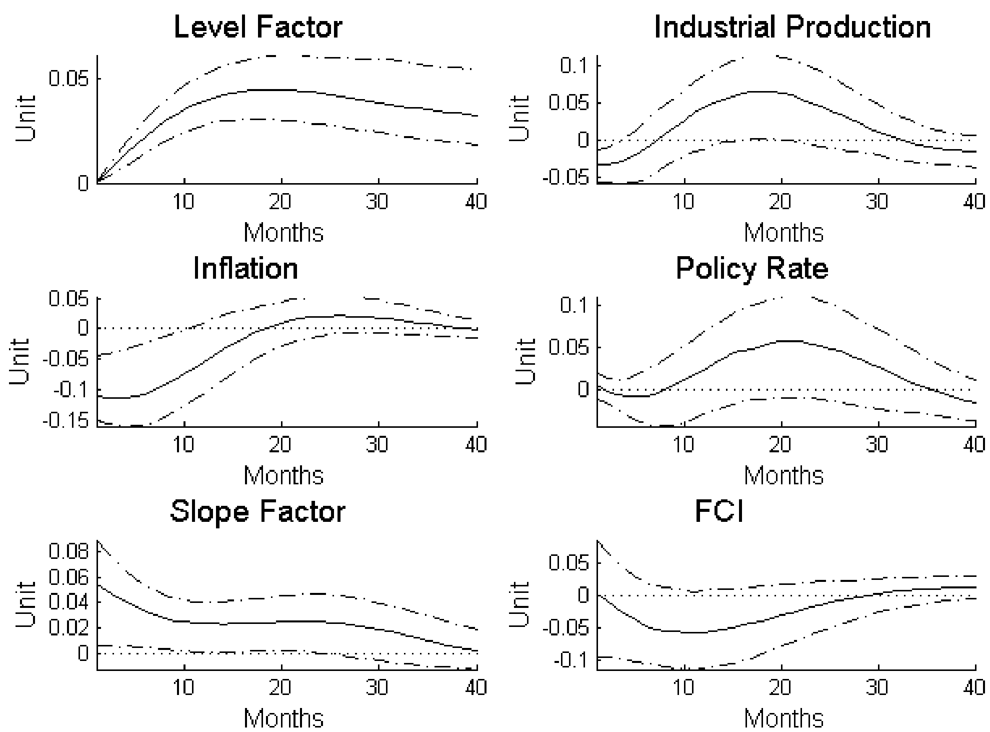
Generally, this 'inflation shock' can be transmitted to bond yields through two channels. The first channel is the influence on the current short rate and expected future short rates. The current short rate and future short rate expectations are closely connected to the conduction of monetary policy, so we regard this channel as 'policy channel'. The movements in this policy channel are in line with the 'Expectation Hypothesis'. The other channel is the 'risk compensation channel', through which the movements account for the bond market risk compensation for a bond at longer maturity. The compensation is also call 'term premia', which is the difference between the real long yield and the 'Expectation Hypothesis' consistent long yield.[47]

Figure 13 sets out the results of IRFs to the long-run Level shock. From the middle-left panel we see that the long-run Level factor is a negative shock on inflation rate.

[47]Our definitions of these two channels are similar to Jotikasthira, Le and Lundblad (2015), although our model structure is different.

The top-left panel shows the long-run Level shock is a positive shock on Level factor which reaches the maximum about a year and a half, so this shock will gradually increase the level of bond yield term structure. The bottom-left panels shows this shock also positively affect the Slope factor, meaning that the short rate increases more than the long rates. Therefore, this shock changes not only the level of term structure, but also the slope. The increase in Slope implies that this negative inflation shock affects the bond yields through both policy channel and risk compensation channel, as it tends to increase the 'Expectation Hypothesis' consistent long yield and decrease the 'term premia'. This finding is consistent with the argument in Wright (2011) that the term premia is positively correlated to the inflation expectation. This argument is also consistent with Duffee (2011) that the shock immediately driving up the expected future short yields drives down the term premiums. However, with the realized increase of future short rates and the recover of inflation (expectation), the term premiums may decrease after a short period, and we will further discuss about it in the following sections.

Figure 13: Impulse Responses to Long-Run Level Shock

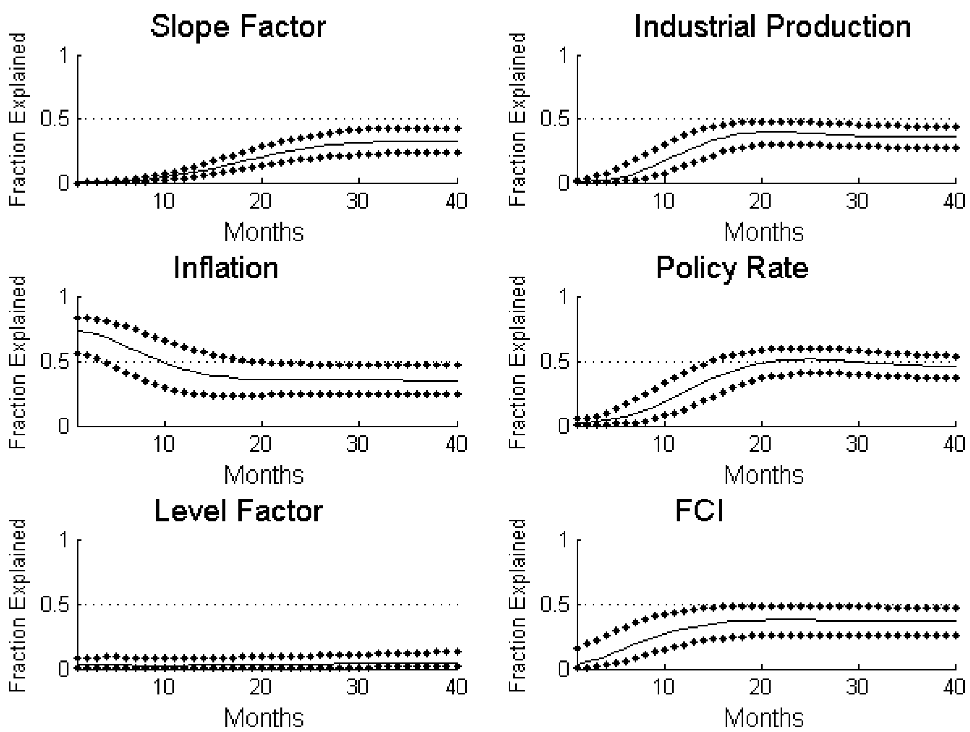


Notes: The solid lines in the above panels are the posterior median values. The dashed lines indicate 16 to 84 percent posterior coverage intervals.

As we mentioned above, the long-run Level factor is a negative shock on inflation rate. Why would this shock drive up future short rates? Because in a global context, our VAR implies that the changes of policy rate is more sensitive to the objectives of Industrial Production (IP) growth rate and FCI, when compared with global inflation rate. This negative inflation shock is accompanied by future favorable movements in Industrial Production (IP) growth rate or FCI, and therefore the global policy rate level will increase to offset to these movements. Moreover, the forward-looking global Slope immediately reacts to this expectation under this circumstance.[48]

We then proceed with the examinations of the long-run shocks of the global Slope factor. Figure 14 shows that the long-run Slope shock is more significant. It increases over the forecast horizon and accounts for 36% of the FEV of the Slope factor after two years. The long-run Slope shock is also most closely related to the inflation, as it accounts for 73% of the FEV of inflation at the beginning and more than 35% at the end. It is also worth noting that this shock is basically orthogonal to the movements of Level factor, so it can barely affect the level of the term structure.

Figure 14: Fraction of Forecast Error Variance (FEV) Explained by Long-Run Slope Shock



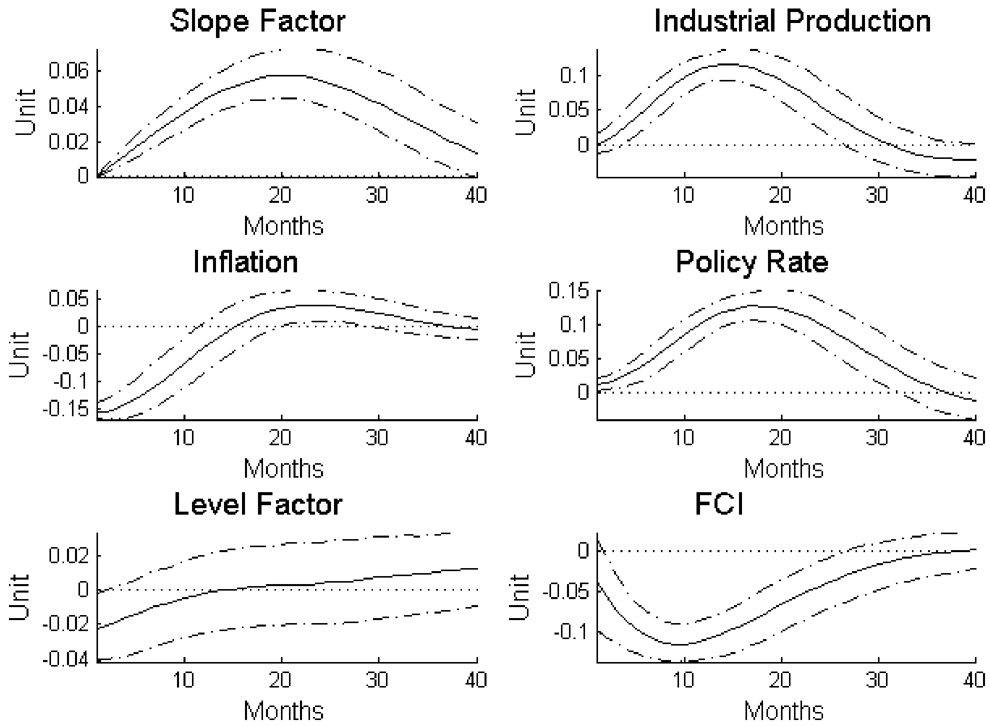
Notes: The solid lines in the above panels are the posterior median values. The dotted lines indicate 16 to 84 percent posterior coverage intervals.

From Figure 15, we see that the long-run Slope shock starts with a negative inflation

[48]The long-run Level shock can be viewed as a markup shock in a Dynamic Stochastic General Equilibrium (DSGE) model. Liu, Waggoner and Zha (2011) show that when a negative shock hits the markup, the output gap is slowly increasing and hence the investment; if the liquidity condition in financial market is improved under this circumstance, then the investment even increases more.

shock, and this shock gradually drives up the IP growth rate and slope of policy rate. This shock is similar to the long-run Level shock, but differs in a way that it does not affect the long end of term structure too much and, hence, has no significant effects on the slope of term structure. Though in a global context, this result is still in line with Kurmann and Otrok (2013), as the shock influencing long-term movements of Slope factor can be viewed as a news shock on IP growth rate.

Figure 15: Impulse Responses to Long-Run Slope Shock



Notes: The solid lines in the above panels are the posterior median values. The dashed lines indicate 16 to 84 percent posterior coverage intervals.

To sum up, our empirical results conclude that the long-run movements of global Level and Slope factors are both driven by inflation shocks. The long-run shocks are not negligible, but the influence is more significant for the global Slope. On the contrary, the current movements of global yield factors seem to be normal market disturbances and do not have clear economic implications, as they are isolated in the bond market and not likely to be transmitted to the real economy.

G.2 Financial Shock and Inflation News Shock

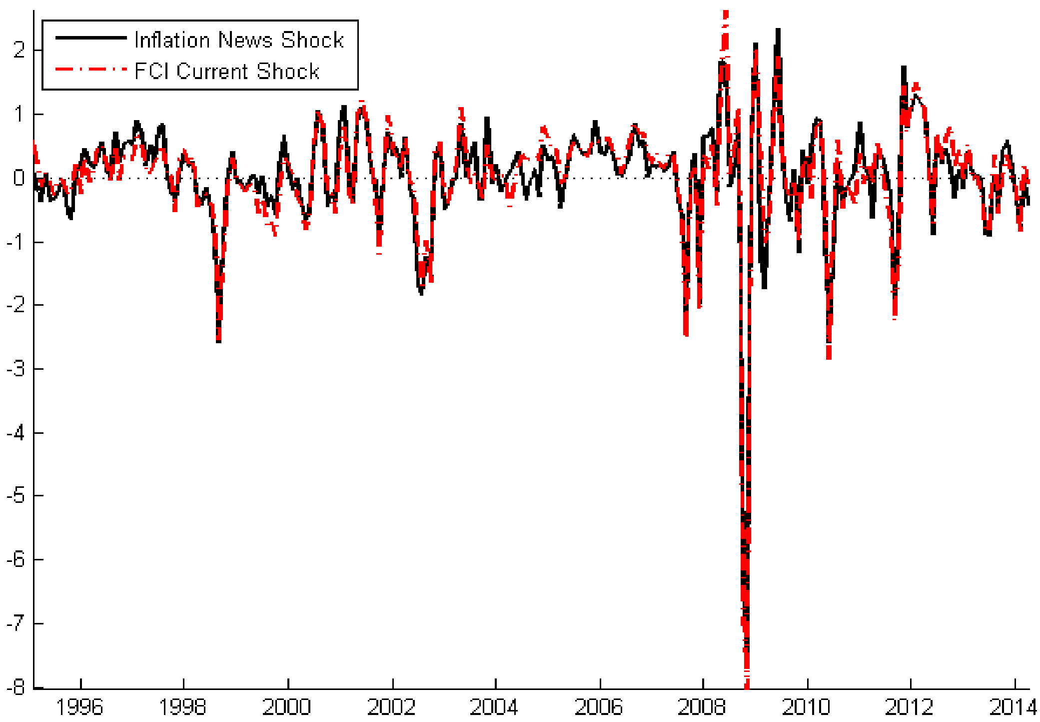
Generally, our discussion is about the nominal rates of bonds. Inflation is considered to be important in influencing the nominal rates if we think real rates are relatively stable.

Hence, we would like to see whether the short-term or long-term shock on the global inflation has significant effects on bond yield factors.

Using the similar identification methods mentioned in last two sections, we can identify the current shock and news shock of global inflation. However, we find that these two kinds of shocks on inflation account for very small proportion of global yield factor movements.[49] This might seen odd, but the reason might be that the inflation shocks that drive up the short rate may also drive down the risk premia, so the effects of these shocks are offset.

Additionally, we have a surprising finding that the inflation news shock is basically the financial current shock, see Figure 16. Although the identification criteria behind the two shocks are completely different from each other, the correspondence is amazingly close (see the following graphs in this section for more comparison). It means the forward-looking variables FCI immediately reacts to the news about the future movements of inflation, and hence the changes of inflation can gradually materialize in the future.

Figure 16: Comparison of Inflation News Shock and FCI Current Shock



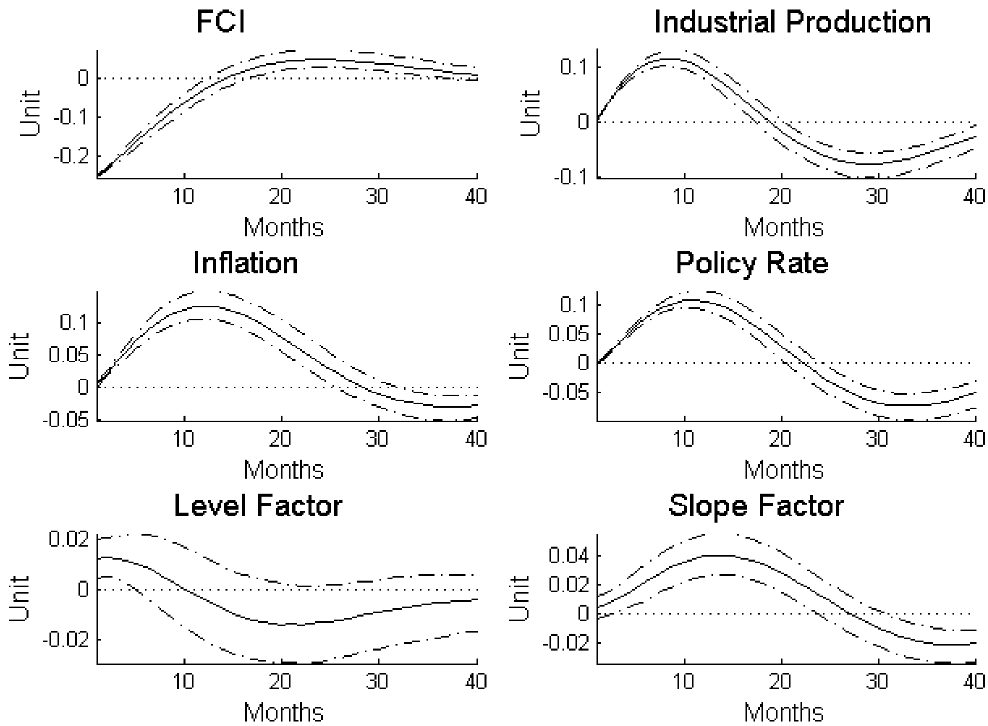
Notes: The solid black line is the posterior median value of all draws of inflation news shocks, whereas the red dashed line is the median of FCI current shocks.

From our empirical evidence, we can conclude that when excluding the financial disturbances in bond markets, the Level factor is more related to the real changes in inflation,

[49]For example, as can be seen in Figure 20, the inflation news shock only accounts for a small proportion of the variance of Slope factor . The other results are not shown here for sake of brevity.

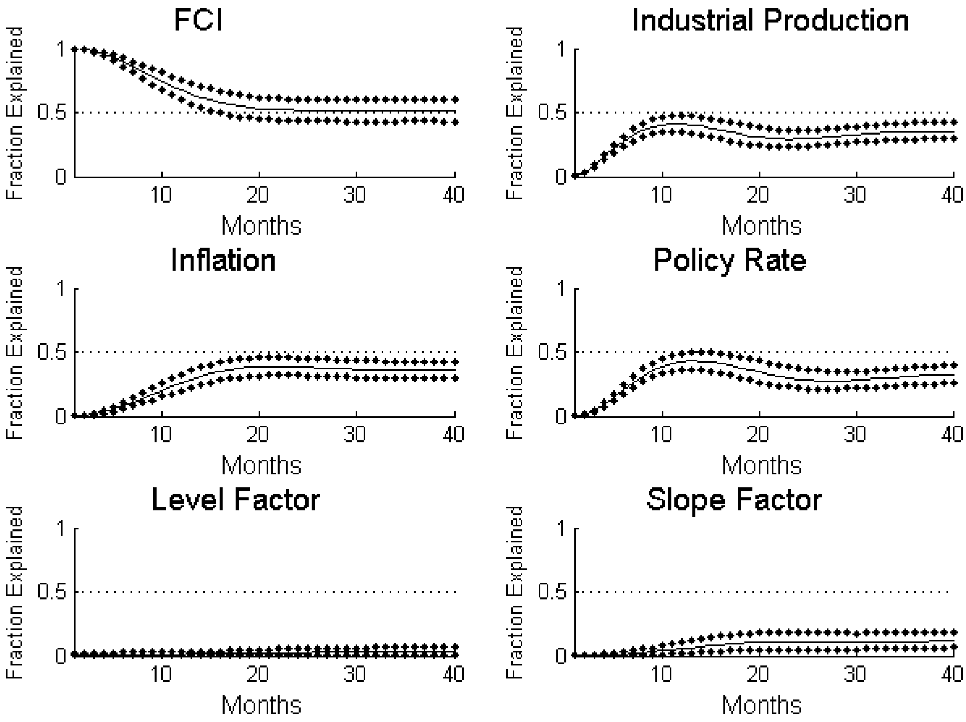
whereas the Slope factor is not only sensitive to the changes of inflation but also the news of inflation, as well as the changes of other macroeconomic information as implied by Figure 14.

Figure 17: Impulse Responses to Financial Current Shock



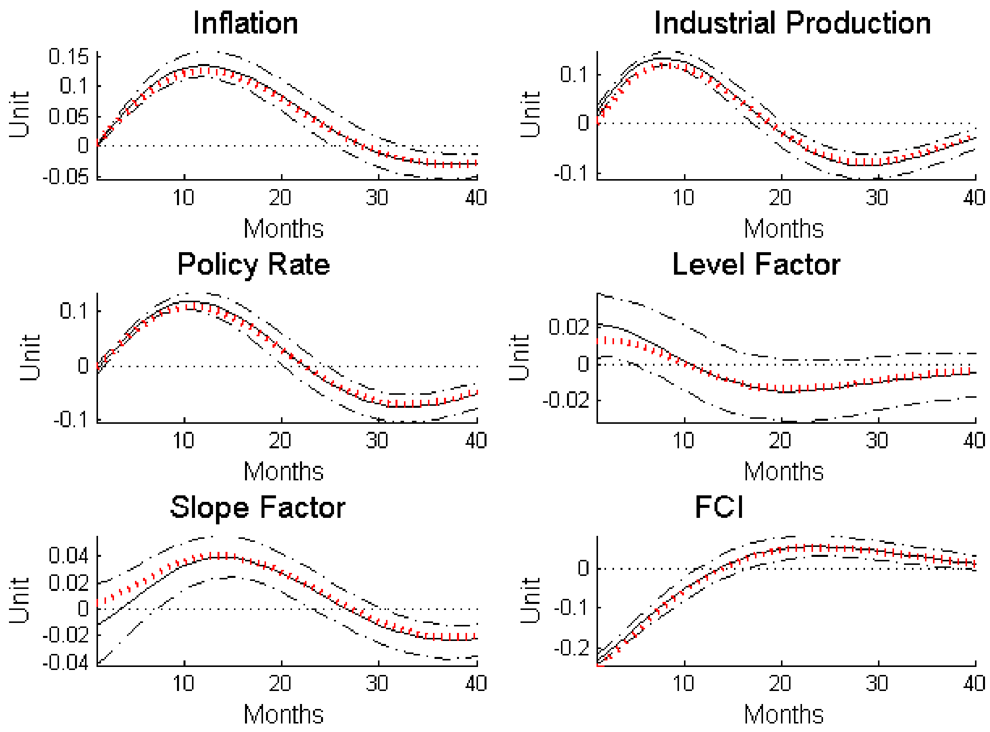
Notes: The solid lines in the above panels are the posterior median values. The dashed lines indicate 16 to 84 percent posterior coverage intervals.

Figure 18: Fraction of Forecast Error Variance (FEV) Explained by FCI Current Shock



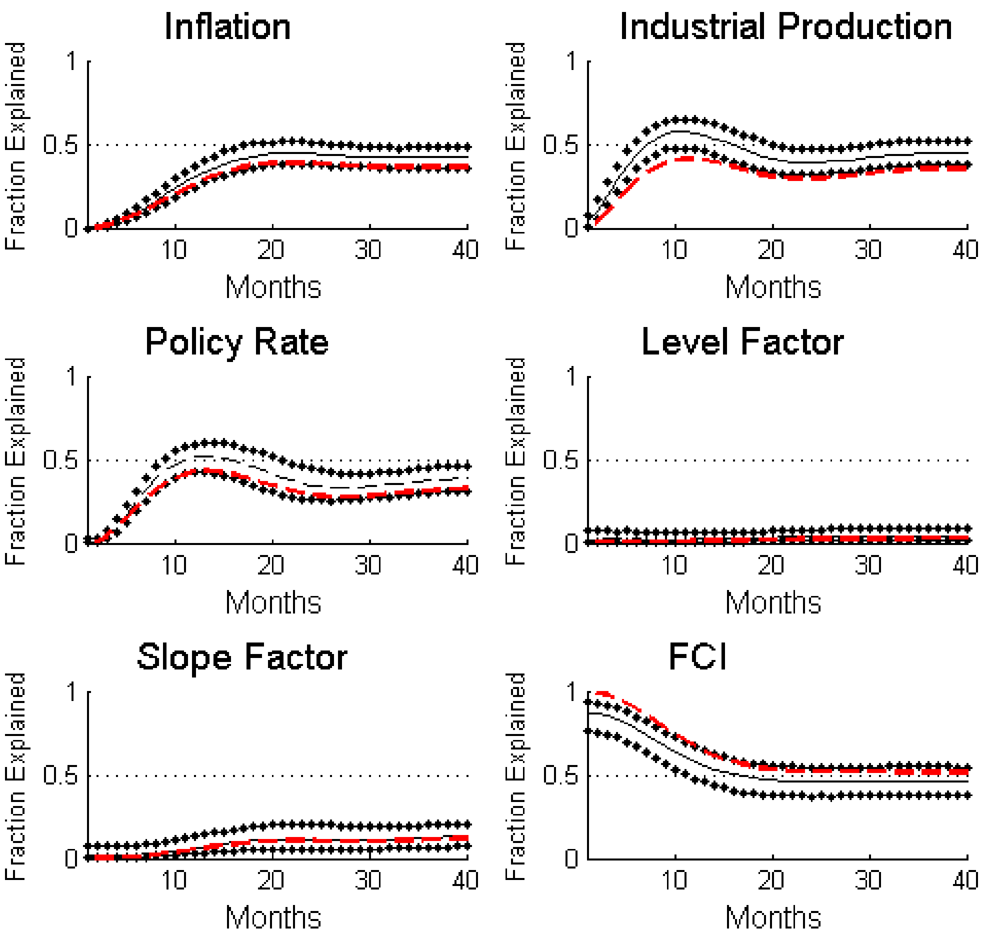
Notes: The solid lines in the above panels are the posterior median values. The dotted lines indicate 16 to 84 percent posterior coverage intervals.

Figure 19: Impulse Responses to Inflation News Shock



Notes: The solid lines in the above panels are the posterior median values. The black dashed lines indicate 16 to 84 percent posterior coverage intervals. The red dotted lines indicate the impulse response to the FCI Current Shock.

Figure 20: Fraction of Forecast Error Variance (FEV) Explained by Inflation News Shock



Notes: The solid lines in the above panels are the posterior median values. The black dotted lines indicate 16 to 84 percent posterior coverage intervals. The red dashed lines indicate the FEV Explained by the FCI Current Shock.

Appendix H Country-Specific Components: Correlation and Granger Causality

Table 23 displays the correlation matrix of the country-specific components in national Level and Slope factors, which implies potential 'Granger causality' among country-specific components.

We set strict criteria for the 'Granger causality' to reveal 'spillovers'. The critical value of the test is set to be 0.01, and the maximum lag is set to be one as the transmission in financial market is considered very rapid. We conduct the causality test for all the draws obtained from our model. We then construct two directed graphs according to the results of Granger causality test in Figure 21.[50]

The upper graph in Figure 21 displays the asymmetric 'spillovers' among Level factors. One obvious observation is that the country-specific components of the UK Granger-cause the country-specific movements in Level factors of all other countries, which implies that the country-specific movements of the UK bond factors release some signals to other markets and cause different degrees of shifts in term structures. But the interpretations of the signals are heterogeneous in different markets, so the 'spillovers' are not captured by the global co-movement.

The lower graph in Figure 21, in contrast, displays the asymmetric 'spillovers' among Slope factors. It is evident that the country-specific movements of Italy in Slope factor are susceptible to all the country-specific components of other countries, which suggests the vulnerability of the Italy bond market.[51]

More interesting observations are shown in Figure 21. Regarding the sovereign risks of Italy, it seems that the risks can influence the levels of bond yields of the US and Germany, but the contagion to Germany market is more evident, as it also affects the movements in Slope. It is possible that the spillovers from Italy to the US arise through the Germany market. Regarding the bond market of Japan, it is clear that the market is closely connected to the US market, as the movements in the US Granger-cause the movements of Japan in both Level and Slope.[52] Nevertheless, the response from Japan is not significant as the US market is solely affected by European markets in Level.

In terms of the reflective mechanisms among the markets, we should pay attention to three pairs: UK-France, Japan-Canada and Italy-Germany. However, to confirm whether there are amplifications or counteractions, more work needs to be done. Moreover, a more complicated mechanism lies in the relation between the movements in the UK, Italy and Canada and the reactions in Germany; the first three countries are likely to affect the Level of Germany, while the feedback from Germany is through the effects on the Slope of the three.

[50] The test results can be found in Table 24 and 25 in Appendix H.

[51] Generally, the changes in Slope factor, i.e. the changes in the shape of the term structure, potentially reflect more severe changes in investor sentiment than the case of parallel shifts.

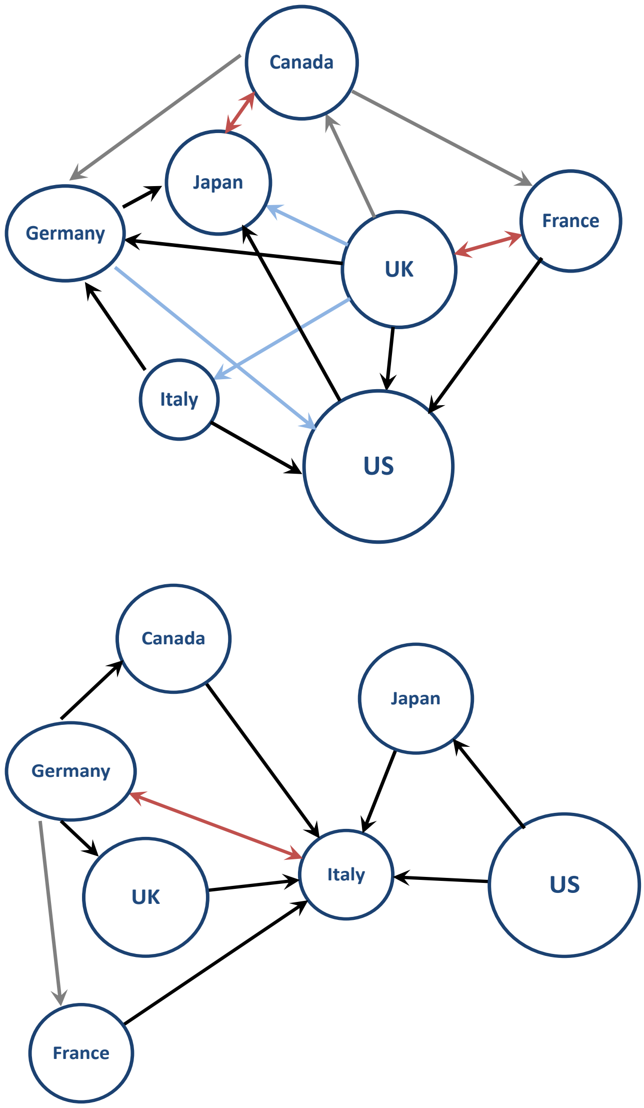
[52] The reason why the US country-specific movements in yield factors do not affect other markets is that the 'fundamental' effects from the US have been captured in the global factor movements.

Table 23: Correlation Matrix of the Country-Specific Components

	ITA_L	ITA_S	CAN_L	CAN_S	FRA_L	FRA_S	GER_L	GER_S	JP_L	JP_S	UK_L	UK_S	US_L
ITA_L													
ITA_S	-0.65(0.03)												
CAN_L													
CAN_S			-0.47(0.04)										
FRA_L	0.49(0.07)												
FRA_S													
GER_L	-0.34(0.07)				0.46(0.05)	0.40(0.07)							
GER_S													
JP_L	0.47(0.04)		-0.41(0.06)	0.37(0.03)			-0.42(0.04)						
JP_S		-0.47(0.03)					0.31(0.05)						
UK_L													
UK_S	-0.34(0.07)												
US_L	-0.57(0.04)				-0.31(0.07)								
US_S	0.47(0.04)		-0.44(0.05)	0.55(0.03)			-0.36(0.06)		0.65(0.03)				-0.34(0.05)

Notes: 1. This table summarizes the correlation matrix of the country-specific components among the Level and Slope factors of all countries: Italy (ITA), Canada (CAN), France (FRA), Germany (GER), Japan (JP), the UK and the US. In each parentheses $(\cdot)$ the posterior standard deviation of the correlation element is calculated from our draws, see Section 2. Larger standard deviation means higher uncertainty in the estimates, but we do not have an exact credible interval interpretation as the statistics do not necessarily follow (truncated) normal distributions.
2. The diagonal elements are dismissed and we only show the elements in the lower triangular part of the correlation matrix with absolute values larger than 0.30. Subscripts L and S are for Level and Slope factors respectively.

Figure 21: Directed Graphs of 'Spillovers' in Country-Specific Components



Notes: The upper figure shows how each country-specific component in Level are affected by components of other countries, whereas the lower figure displays the influence in Slope. The graphs are constructed according to the results in Table 24 and 25.

Table 24: Granger Causality of the Country-Specific Components (Level)

	ITA_L	CAN_L	FRA_L	GER_L	JP_L	UK_L	US_L
ITA_L	NA			***			***
CAN_L		NA	**	**	***		
FRA_L			NA			**	***
GER_L				NA	***		*
JP_L		**			NA		
UK_L	*	**	***	***	*	NA	***
US_L					**		NA
ITA_S							
CAN_S							
FRA_S							
GER_S							
JP_S							
UK_S							***
US_S					***		

Notes: 1. This table summarizes the Granger causality statistics of the Country-Specific Components of all countries: Italy (ITA), Canada (CAN), France (FRA), Germany (GER), Japan (JP), the UK and the US. The subscript of each country indicates the factor Level (L) or Slope (S).
2. The diagonal elements in the upper half table are not applicable. Each column indicates whether the component of one country is Granger-caused by other components.
3. The significance level of the Granger causality test is set to be 0.01, and the lag is set to be 1. *, ** and *** indicate 70%, 80% and 90% of the posterior draws reject the test, respectively.

Table 25: Granger Causality of the Country-Specific Components (Slope)

	ITA_S	CAN_S	FRA_S	GER_S	JP_S	UK_S	US_S
ITA_L	***				***		***
CAN_L	***						
FRA_L	***						
GER_L	***		**				
JP_L	***						
UK_L	***					**	
US_L	***				***		
ITA_S	NA						
CAN_S	***	NA					
FRA_S	***		NA				
GER_S	***	***		NA		***	
JP_S	***				NA		
UK_S	***					NA	
US_S	***						NA

Notes: 1. This table summarizes the Granger causality statistics of the Country-Specific Components of all countries: Italy (ITA), Canada (CAN), France (FRA), Germany (GER), Japan (JP), the UK and the US. The subscript of each country indicates the factor Level (L) or Slope (S).
2. The diagonal elements in the lower half table are not applicable. Each column indicates whether the component of one country is Granger-caused by other components.
3. The significance level of the Granger causality test is set to be 0.01, and the lag is set to be 1. *, ** and *** indicate 70%, 80% and 90% of the posterior draws reject the test, respectively.