Why Do Term Structures in Different Currencies Comove? *

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Abstract

Yield curve fluctuations across different currencies are highly correlated. This paper investigates this phenomenon by exploring the channels through which macroeconomic shocks are transmitted across borders. Macroeconomic shocks affect current and expected future short-term rates as central banks react to changing economic environments. However, policy reactions are not the only channel through which the macroeconomy may affect bond yields. Investors may also respond to these shocks by altering their required compensation for risk. Macroeconomic shocks thus influence bond yields both through a "policy" channel as well as through a "risk compensation" channel. In a no-arbitrage vector autoregressive framework, we employ deviations from the expectations hypothesis to identify the two transmission channels, with particular attention to the degree to which each channel contributes to the co-variation among term structures across the U.S., the U.K., and Germany. We find that a world inflation factor explains a sizable fraction of the co-variance of medium- to long-maturity yields between the U.S. and the U.K., as well as between the U.S. and Germany. Further, we find that the world inflation effect operates almost exclusively through the risk compensation channel for long-term bonds.

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1 Introduction

Yield curve fluctuations across different currencies are highly correlated (see Figure 1). We address this phenomenon by exploring how macroeconomic shocks are transmitted across borders. In a rational expectation equilibrium, shocks can affect the term structure in two ways. One is through the joint dynamics of state variables, including those upon which monetary policy actions are based. The other is through the market prices of risk (i.e. term premia) that investors assign to these state variables. We refer to these two channels as the "policy" and the "risk compensation" channels, respectively.

The first channel arises as central banks target the short-term rate, as a policy tool, in order to manage domestic growth and inflationary pressures. Thus, co-movement at the short end of the yield curve is likely to reflect co-movement in the economic fundamentals to which monetary authorities are responding. Indeed, Kose, Otrok, and Whiteman (2003) find that a common world factor is an important source of volatility for macro aggregates in most countries, providing evidence for a global business cycle.[1] If central banks indeed set interest rates as a function of global economic conditions, then the world business cycle would suggest that short-term interest rates are also highly correlated across countries.

However, a reasonable question is whether policy exposures to common global factors, coupled with the expectations hypothesis, can explain other equally important features of international yield curves across the maturity spectrum. While the policy channel would imply that long-maturity yields will also be correlated, it is unlikely that this channel alone can explain the significant degree of variation and cross-market co-variation observed for long maturity bond yields. In a single-country setting, several papers (see Ang and Piazzesi (2003), Kuttner (2001), Evans and Marshall (2007), and Dai and Philippon (2005) for example) have documented that macroeconomic shocks, including monetary and fiscal policy

[1] However, the evidence suggests that the cross-country correlations are significantly higher for interest rates than for macroeconomic variables.

surprises, do significantly impact bond yields. However, these shocks work their way to the long end of the term structure largely through market prices of risk, which allow long-maturity yields to load more or less on economic and monetary factors than predicted by the expectations hypothesis (see Dai and Singleton (2002), Le, Singleton, and Dai (2010), and Duffee (2002)).[2] These findings imply that macroeconomic factors may indeed be important in explaining the observed co-movement among yield curves in the international setting, but that these factors may impact long-maturity yields through the alternative risk compensation channel. Employing standard Campbell and Shiller (1991) regressions, Sutton (2000) provides early evidence that bond yields in different major currencies indeed display excess co-movement relative to that implied by the expectations hypothesis; we formally explore his conjecture that these findings are the result of positively correlated term premia.[3]

To provide some context, there are numerous examples where long-maturity bond yields tend to co-move across borders even when the short-term policy rates diverge, in contrast to the pattern implied by the standard policy (expectations hypothesis) channel. For instance, short-term policy rates diverged in the U.S. and U.K. in 1997, reflecting different inflationary pressures across the two countries, whereas ten-year bond yields in both currencies moved in concert (see Figure 2). Similarly, during the broad economic recession in 1990-1991, the Federal Reserve and the Bank of England responded by sharply cutting policy rates. During the same period, however, the German economy grew rapidly and the Bundesbank raised its policy rate by over 250 basis points from late 1989 to end of 1991. Despite these quite different policy directions, ten-year bond yields fell sharply across all three markets.

[2]In related work, Gallmeyer, Hollifield, Palomino, and Zin (2009) and Wachter (2006), in a single country setting, argue that long-maturity bond yield volatility is difficult to produce using a usual rational expectations model with stationary state variables. Both papers thus propose models with habit formation to generate large market prices of risk. For equity markets, Bansal and Lundblad (2002) and Colacito and Croce (2010) argue that time-varying expected returns explain the sharp contrast in correlation between international equity returns (high) versus aggregate market dividend growth rates (low).

[3]A few other papers examine the expectations hypothesis and the predictive power of term spreads in countries other than the U.S. However, they do not study linkages between term structures of different countries. Examples include Hardouvelis (1994) and Jorion and Mishkin (1991).

Generating these sorts of (admittedly anecdotal) patterns requires a more general model that facilitates fluctuations in market prices of risk.

In this paper, we quantify the contributions of both the policy and risk compensation channels by estimating a discrete-time affine term structure model with the observed underlying factors following a vector autoregressive (VAR) process. The model is similar in spirit to Ang and Piazzesi (2003), but allows for un-spanned macroeconomic variables as in Joslin, Priebsch, and Singleton (2010). We estimate the model (jointly) by maximum likelihood using macroeconomic and term structure data from the U.S., the U.K., and Germany (from 1987:04 to 2008:03, at monthly frequency).[4]

We find that measured macroeconomic and yield-curve factors are stationary and therefore have a limited impact on long-maturity yields under the expectations hypothesis. Further, we find that yield volatilities exceed those predicted by the expectations hypothesis and the physical dynamics of macroeconomic variables and yield-curve factors. The model attributes the remainder to risk premia fluctuations; similar to the single country research mentioned above, this is particularly important for long-maturity bond yields. With regards to cross-border *co-movement*, we find that global macro factors operate through both the policy channel (about one-third) and the risk compensation channel (about two-thirds) for five-year bond yields. As we move to the ten-year maturity, these numbers shift considerably in favor of the risk compensation channel.[5] The most important state variable is the world inflation rate as it alone explains over half of the cross-country co-variances in bond yields. Interestingly, the global inflation factor generally operates through the policy channel at the short end, as central banks respond to global inflationary pressures, but almost exclu-

[4] As discussed by Dong (2006), the operating procedures of the Federal Reserve prior to 1982 focused on meeting non-borrowed aggregate reserve targets instead of managing short interest rates in the post-Volcker era. For now, we exclude data during the recent sub-prime crisis (a better understanding of the link between policy, risk premia, and the term structure of interest rates during the crisis is an important avenue for future research).

[5] This is true despite the fact that by construction, the co-movement of short-term rates is 100% fundamentally driven.

sively through the risk compensation channel at the long end, as global investors alter their required compensation for risk.

This paper contributes to the finance and economics literature in three important ways. First, it is the first to examine the economics behind the international co-variation of the term structure of interest rates, providing a step in understanding the co-movement of other risky assets and, possibly, economic activity. Shiller (1989) studies the co-movement of the U.S. and U.K. stock prices in a simple present-value context and shows that the correlation between dividends in the two countries is too small to account for the high degree of co-variation in stock price. The macroeconomics literature also singles out the interest-rate channel as one of the most important linkages between different countries' stock markets. For example, Ehrmann and Fratzscher (2006) find that the transmission of U.S. monetary policy to other countries' equity markets occurs largely through the reaction of U.S. short-term interest rates, foreign short-term interest rates, and exchange rates.[6] Examining the impact of U.S. monetary policy on other countries' economic activity, Canova (2005) and Kim (2001) arrive at the same conclusion: interest rate reactions are the most important transmission channel.

Second, from a modeling standpoint, this paper is the first to apply an affine term structure model to the exploration of the cross-country transmission of macroeconomic shocks. Building on recent advances in the term structure modeling literature, pioneered by Joslin, Singleton, and Zhu (2010) and Joslin, Priebsch, and Singleton (2010), our model design allows us to jointly estimate multiple term structures denominated in different currencies in a particularly robust and efficient manner. Possibly due to the curse of dimensionality, most models studying the transmission of macroeconomic shocks and the term structures are employed in a single-currency setting. A few manage to extend the models to a two-countries setting, aiming at explaining the exchange rate dynamics and the failure of the uncovered

[6] See also Wongswan (2006).

interest rate parity.[7] However, due to the size of the parameter space, these papers typically model and estimate separate pairs of countries. If one considers the U.S.-U.K. and U.S.-Germany pairs as we do in this paper, the traditional models will almost surely imply different estimated dynamics for the U.S. term structure. Since joint maximum likelihood estimation can be easily implemented within our framework, this issue is completely avoided.

Third, the results in this paper may help to further develop our understanding of the channels through which crises and contagion spread. Forbes and Rigobon (2002) show that stock markets co-move highly in all periods but, due to the large volatility of shocks, they may

2 The Model

We build a model of multiple term structures of interest rates across countries, each indexed by i. Underlying each economy is a vector of macro variables M_t^i and yield curve pricing factors P_t^i. The length of these vectors can be unequal across currencies. In addition, we characterize the global economy through a vector of global macro variables M_t^W. The movement of M_t^W generates comovement among fundamentals of individual economies. Following Joslin, Singleton, and Zhu (2010) and Joslin, Priebsch, and Singleton (2010), among others, under the $\mathbb{Q}$-measures, only the pricing factors P can contemporaneously affect bond yields. Under the $\mathbb{P}$-measures, both M and P can contemporaneously influence and predict bond yields, through their joint dynamics. This way, we avoid over-fitting bond yields with a large number of pricing factors while allowing for both spanned and un-spanned macro risks.[8]

2.1 Short Rates and Bond Yields

We assume that the central bank of country i sets the short-term (policy) rate as a linear function of the world macro variables and all state variables of that country (similar in spirit to the backward-looking Taylor rule).[9] Based on the low dimensional structure of yields, we make the standard assumption that there exist N^i *unobservable* pricing factors X^i that drive yields at all maturities, $y_{n,t}^i$, of the i^{th} economy. Starting with the short rate y_1^i, we write:

$$y_{1,t}^i = r_\infty^i + \iota' X_t^i \tag{1}$$

[8]In reality, there may be several other omitted latent and observed variables. However, Duffee (2006) shows that the estimation is unbiased and consistent so long as the observed (and estimated) factors (i) do not affect the market price of risk of the omitted factors, and (ii) do not contain information about the future values of the omitted factors.

[9]Ang, Dong, and Piazzesi (2005) examine the term structure implications of backward-looking vs. forward-looking Taylor rules. Gallmeyer, Hollifield, and Zin (2005) study the implications of the Taylor rule vs. the McCallum rule. The latter allows information in the yield curve to be used in setting the policy interest rate.

where ι is a vector of ones. Following Joslin, Priebsch, and Singleton (2010), we assume the following risk-neutral dynamics for X^i:

$$X^i_{t+1} = \lambda^i X^i_t + \Sigma_{X^i} \epsilon^{\mathbb{Q}}_{i,t+1} \tag{2}$$

where λ^i is a diagonal matrix. Note that every VAR(1) specification can be transformed to (2) through standard shifts and rotations of the latent factors. For simplicity, we assume that elements of λ^i are real and distinct.[10] From (1) and (2), the yield of bond of country i with maturity n is given by[11]

$$y^i_{n,t} = A^i_{n,X^i}(\lambda^i, r^i_\infty, \Sigma_{X^i}) + B^i_{n,X^i}(\lambda^i)' X^i_t. \tag{3}$$

It is clear from (3) that the yields on all bonds of country i will only depend on X^i and a few parameters associated with its risk-neutral dynamics. For notational simplicity, we will drop the superscript i in showing how A^i_n and B^i_n are computed:

$$\begin{aligned} B_{n,X} &= \frac{1}{n} diag((I-\lambda)^{-1}(I-(\lambda)^n)), && (4)\\ A_{n,X} &= r_\infty - \frac{1}{2n} tr\left(\Sigma_X \Sigma_X' \sum_{k=0}^{n-1} k^2 B_{k,X} B_{k,X}'\right) && (5)\\ &= r_\infty - \frac{1}{2n} tr\left(\Sigma_X \Sigma_X' C_{n,X}\right) && (6) \end{aligned}$$

where

$$C_{n,X} = \sum_{k=0}^{n-1} k^2 B_{k,X} B_{k,X}'. \tag{7}$$

[10] Joslin, Singleton, and Zhu (2010) consider the case of complex λ^i. Real-valued λ^i's are found to be empirically adequate.

[11] See proofs in Duffie and Kan (1996) and Dai and Singleton (2000) for continuous-time versions and Ang and Piazzesi (2003) for discrete-time versions.

It is important to note that $B_{n,X}$ and $C_{n,X}$ are dependent on λ only. This feature significantly reduces the dimensionality of the model and its estimation, which is particularly helpful in simultaneously dealing with multiple term structures. One shortcoming of this affine structure is that it leads to constant factor loadings $B_{n,X}$. This implies that our model cannot correctly capture the time-varying nature of the relative importance of the factors. In addition, the Gaussian structure of the state variables implies that the conditional covariances of bond yields will also be constant.

2.2 From Latent to Observed Pricing Factors

An affine model with latent factors can be difficult to estimate since the factor values have to be inferred from bond prices in each and every iteration of the estimation process. Therefore, maximum likelihood estimation requires that the parameters of both the physical and risk neutral dynamics be jointly estimated. Joslin, Singleton, and Zhu (2010) show that by assuming that a few portfolios of bonds are priced without error and using their yields as observed pricing factors, one can bypass this challenge. The reason is that the likelihood function can now be split into two smaller parts corresponding to the physical and risk-neutral dynamics, and each of these parts can (more or less) be maximized separately.[12]

We follow this approach and assume that fixed weight (W) portfolios of yields corresponding to the first N principal components of yields, P_t, are priced without error. Note again that all calculations in this subsection are specific to each currency i and so we drop the superscript i for simplicity. We first show that P_t is nothing more than a rotation of X_t.

$$P_t = W\, y_{\tilde{n},t} = W\, A_{\tilde{n},X} + W\, B'_{\tilde{n},X} X_t = U + V X_t \tag{8}$$

[12]The only parameter that appears in both is Σ_X. However, the covariances of residuals from the estimated physical dynamics provide good starting values for the estimation.

where $\tilde{n}$ is a vector of maturities that constitute the perfectly priced portfolios, and let

$$U = W\,A_{\tilde{n},X} \;\text{ and }\; V = W\,B'_{\tilde{n},X}. \tag{9}$$

From this, we can re-write the term structure model now with P_t, as opposed to X_t, as the pricing factors:

$$y_{n,t} = A_{n,X} + B'_{n,X}X_t = A_{n,X} + B'_{n,X}V^{-1}(P_t - U) = A_{n,P} + B'_{n,P}P_t \tag{10}$$

where

$$A_{n,P} = A_{n,X} - B'_{n,X}V^{-1}U \;\text{ and }\; B_{n,P} = (V^{-1})'B_{n,X}. \tag{11}$$

2.3 Physical Dynamics and Restrictions

Now that we can price all bonds as a function of *observed* pricing factors, P, we will present the physical dynamics of P as opposed to those of X (which simply reflects a rotation). To disentangle the part of $M_{i,t}$ that is correlated with the global economy, we first project M_t^i on M_t^W:

$$M_t^i = a^i M_t^W + M_t^{i,e} \tag{12}$$

to obtain $M_t^{i,e}$ as "residual" macro factors. Likewise, projecting P_t^i on M_t^W and $M_t^{i,e}$:

$$P_{i,t} = b^i M_t^{i,e} + c^i M_t^W + P_t^{i,e}, \tag{13}$$

we obtain $P_t^{i,e}$ as residual yield curve factors for the i^{th} economy.

We assume that the world macro variables follow an autonomous $VAR(1)$ process:

$$M_{t+1}^W = \kappa_{w,0} + \kappa_{w,1} M_t^W + \Sigma_w \epsilon_{w,t+1}^{\mathbb{P}}, \tag{14}$$

and model the joint physical dynamics of $Y_t^{i,e\prime} = (M_t^{W\prime}, M_t^{i,e\prime}, P_t^{i,e\prime})$ by a gaussian VAR(1):

$$Y_{t+1}^{i,e} = \kappa_0^{i,e} + \kappa_1^{i,e} Y_t^{i,e} + \Sigma_{Y^{i,e}} \epsilon_{i,t+1}^{\mathbb{P}}. \tag{15}$$

Similar to Ang and Piazzesi (2003) and Evans and Marshall (2007), we impose a block structure on the dynamics of $Y_t^{i,e}$. In particular, the residual yield curve factors of each economy are not allowed to feed back into the macro variables of any economy or the world. The residual macro variables of each economy do not feed back into the global factors or to the residual macro factors of any other economy, except its own. In addition, except for the contemporaneous correlation between macro variables of the i^{th} economy and the world economy, all other cross-economy contemporaneous correlations are shut down. With this structure, interactions between any two economies occur only through their respective correlations with the global macro variables. To be concrete, our restrictions imply that $\kappa_1^{i,e}$ and $\Sigma_{Y^{i,e}}$ take the following forms:

$$\kappa_1^{i,e} = \begin{bmatrix} xx & 0 & 0 \\ xx & xx & 0 \\ xx & xx & xx \end{bmatrix}, \; and \; \Sigma_{Y^{i,e}} = \begin{bmatrix} xx & 0 & 0 \\ 0 & xx & 0 \\ 0 & 0 & xx \end{bmatrix} \tag{16}$$

where xx denotes a non-zero block of a matrix.

The assumed dynamics of $Y_t^{i,e}$ imply a similar VAR(1) structure for $Y_t^i = (M_t^{W\prime}, M_t^{i\prime}, P_t^{i\prime})'$

because equations (12) and (13) together imply that $Y_t^i = \Pi^i Y_t^{i,e}$ where:

$$\Pi^i = \begin{bmatrix} I & 0 & 0 \\ a^i & I & 0 \\ c^i & b^i & I \end{bmatrix}. \tag{17}$$

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It follows that:

$$Y_{t+1}^i = \kappa_0^i + \kappa_1^i Y_t^i + \Sigma_{Y^i} \epsilon_{i,t+1}^{\mathbb{P}} \tag{18}$$

where

$$\kappa_1^i = \Pi^i \kappa^{i,e} \Pi^{i^{-1}} \textit{ and } \Sigma_{Y^i} = \Pi^i \Sigma_{Y^{i,e}} \textit{ and } \kappa_0^i = \Pi^i \kappa_0^{i,e}. \tag{19}$$

Given this structure and our bond pricing solution, bond yields are related across currencies by the exposures of pricing factors (P^i) to the common macro variables (M^W), both directly and indirectly through the influence of the world macro variables on the country-specific macro variables (M^i).

Our model differs from those of Ang and Piazzesi (2003) and Dong (2006) along several important dimensions. First, although we allow the macro variables to affect the pricing factors, they remain un-spanned by bond yields. This is consistent with the evidence provided by Joslin, Priebsch, and Singleton (2010) and others. Under a rigid application of the models in Ang and Piazzesi (2003) or Dong (2006), the macro variables can be expressed as linear functions of bond yields. Second, we adopt the risk-neutral setup of Joslin, Priebsch, and Singleton (2010), and therefore can restrict the number of pricing factors to a small number as required by the data. For example, the first two principal components of bond yields in the U.S. explain over 99% of their observed variation. Therefore, a term structure model that implies five or six factors for bond yields will likely overfit the data. Third, since

our yield curve factors are observable, the relationship between macro variables and bond yields will be more easily identified from the data. On this last point, although the contemporaneous relationship should, in theory, be apparent from the estimate of Σ_{Y^i}, identifying this relationship from the data can be challenging given measurement error, the exact timing of the macro variables, and the high degree of persistence of both the macro variables and bond yields.

2.4 Implied Monetary Policy Rule

Unlike Ang and Piazzesi (2003) and others, we do not start with a monetary policy rule (e.g. Taylor rule). However, under the assumed structure, our model implies a certain monetary rule given that the short rate loads on the pricing factors (as in (1)) and the pricing factors load on the macro variables (as in (12) and (13)). If we define a general form of monetary policy rule as

$$y_{1,t} = \delta M_t + y^*_{1,t} \tag{20}$$

where the coefficient δ represents the monetary policy responses to macroeconomic shocks, then our model implies that for country i, $\delta^{i\prime} = {B^i_{1,P}}' \, [c^i \; b^i]$ where both M^W and M^i affect country i's monetary policy. If M includes only the inflation rate and the output gap, then our model is similar to a Taylor rule.[13]

3 Empirical Implementation

3.1 Data

We study the U.S., the U.K., and German term structures, using LIBOR rates from the British Bankers Association and swap rates from Reuters. LIBOR-swap rates are more

[13]See also Joslin, Priebsch, and Singleton (2010) for a general discussion on this point.

readily available than government bond yields for a broader set of countries (which we plan to study in future research). Moreover, LIBOR rates are of much better quality than the government rates at short maturities, due to the lack of liquidity in the U.K. and German bill markets. In addition, several researchers have argued that government bonds are not an ideal proxy for the unobservable risk-free rates due to favorable taxation treatment, repo specials, scarcity premia, and benchmark status. Swaps may arguably provide a good alternative given that the credit risk priced into swap contracts is very small.[14] We extract the zero-coupon yields from the swap rates, using the un-smoothed Fama-Bliss procedure (or bootstrapping assuming constant forward rates over absent maturities).[15] We use the following maturities: 6 months, 1 year, 2 years, 5 years, 7 years, and 10 years. In general, the swap-implied zero-coupon rates look quite similar to (upward-sloping, etc.) and are highly correlated with government bond yields. To save space, general summary statistics are not reported, but certain important moments can be found in Table 5.

Following Ang and Piazzesi (2003) among others, we use two macroeconomic variables in the vector M for each country – the inflation rate and Industrial Production (IP) growth (denoted by π and g respectively), calculated as the overlapping 12-month (year-on-year) logged changes in the seasonally-adjusted CPI and IP indices, respectively. For the global counterparts, we obtain the two series for the IMF's "advanced economies" group, for which consumer price and industrial production aggregate indices are obtained as GDP or industry value-added, respectively, weighted averages of country-level indices.[16] The data are at the monthly frequency for the periods from 1987:04 to 2008:03. The starting date is restricted by the starting date of LIBOR.[17]

[14]See Blanco, Brennan, and Marsh (2005), Duffie and Huang (1996), and Hull, Predescu, and White (2004) for further elaboration.

[15]See Fama and Bliss (1987) for details.

[16]This group includes Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Australia, Canada, Denmark, Hong Kong, Israel, Japan, Korea, New Zealand, Norway, Singapore, Sweden, Switzerland, United Kingdom, and United States.

[17]Our starting date comes after most developed central banks have managed to better control inflation, so we do not face the problem of high and non-stationary inflation rates. Clarida, Gali, and Gertler (1998)

Table 1 reports the cross-country correlations of these variables. As discussed, the long-term yields are more highly correlated than both the short-term rates and the macro variables. This is not due to a simple explanation that the long-term yields are less volatile than short-term rates. First, the standard deviations of long-term yields are only slightly lower than those of the short-term rates. As pointed out by Gallmeyer, Hollifield, Palomino, and Zin (2009) and Wachter (2006), this pattern is difficult to produce in a usual rational expectation model with stationary state variables. Second, the covariances, not just the correlations, are higher at the long end (once the predictability has been accounted for) (see Figure 1). This means that long-term yields are likely to have greater loadings on common state variables than do short-term rates. Together, the data and the theoretical literature hints at the significance of volatile and correlated market prices of risk in explaining the cross-country covariation of yields. This paper examine this channel in detail under a no-arbitrage VAR framework.

Another interesting observation from Table 1 is that the correlations of short-term rates are only slightly higher than those of inflation rates, suggesting that inflation may drive the correlations at the short end of the term structure. Whether or not inflation is as important in explaining the co-movement across the long-maturity bond yields remains a question. For the U.S., Ang, Bekaert, and Wei (2008) find that the term structure of real interest rates is flat and that the inflation risk premium largely explains the variation in long-term yields.[18] These findings thus suggest that inflation is likely to also be an important driver of cross-country covariances of bond yields. Finally, both the macro variables and short-term rates exhibit higher correlations for the U.S. and U.K. pair than for the U.S. and Germany pair, implying possibly that the U.S. is more economically integrated with the U.K. than

provide some guidance on when controlling inflation became a major focus. For the Bundesbank, they identify March 1979, the time Germany entered the European Monetary System, at the starting point. They identify October 1979 for the Fed, when Volcker clearly signalled his intention to rein in inflation. Clarida et al. experiment with the post-1982 sample period, as the operating procedures of the Fed prior to 1982 focused on meeting non-borrowed aggregate reserve targets instead of managing short interest rates.

[18] See also Wright (2009) for international evidence.

to Germany. In general, the patterns in the data seem to indicate that macro variables are critical in the determination of bond yield co-movements.

3.2 Construction of Pricing Factors

We construct the vector of pricing factors P as the first two principal components of yields for each currency. Table 2 reports the variations explained by these principal components and their loadings on the zero-coupon yields. In all currencies, P^1 explains around 95% of yield variations, and P^1 and P^2 together explain over 99% of yield variations. Thus, without imposing no-arbitrage, it appears that using these factors will allow us to fit the term structures quite well. The loadings of P on yields suggest that in all currencies, P^1 is the "level" factor since it loads almost equally on all yields. P^2 appears to be the "slope" factor since it loads negatively on the short-term rates and positively (increasingly so) on the long-maturity rates. Our results are thus consistent with what others have documented about the behavior of the term structure, particularly that over 90% of the time, the term structure moves in a parallel fashion.[19] Note that these pricing factors are *observed*, resulting in an easier estimation environment relative to the latent factors found in most term structure models.

In order to understand how important the macro variables are in explaining these pricing factors, we regress each pricing factor on the macro variables and document the R-squared in the last two columns of Table 2. In all currencies, over 60% of the variation in P^1 can be accounted for by the world macro variables. This proportion declines for P^2, particularly for the U.K. Still, given that P^1 explains over 95% of variation bond yields, the high R^2's from regressing P^1 on the world macro variables suggest that they are quite important in explaining the movements of country-level term structures. Country-level macro variables appear to add additional explanatory contribution, particularly for the second (slope) factor.

[19]See Litterman and Scheinkman (1991) for example.

We will perform variance decompositions as part of our analysis to confirm this conjecture.

Together, dropping the superscript i as we stack together the state variables of all three countries, the vector Y in (18) is 14×1, including 2 world macro variables, 2 country-specific macro variables for each country, and 2 pricing factors for each country. Given the structure of our model, the challenge in estimation lies in searching for λ (this will be clear in the next subsection), and given that λ is diagonal, it only grows linearly with the number of pricing factors. These 14 state variables, including only 6 pricing factors, will not yield a specification too large to be estimated.

3.3 Estimation

Starting with the observed pricing factors (P), we obtain the residual yield curve factors (P^e) by the projections described in (12) and (13). We estimate the model parameters jointly by maximum likelihood (ML). We write down explicitly the log-likelihood function of the observed data for all three countries– state variables Y_t^e, and bond yields $y_{n,t}$ – as follows:

$$\begin{aligned}\mathcal{L} &= \sum_{t=1}^{T-1} ln(f(Y_{t+1}^e, ye_{t+1}|Y_t^e)) \\ &= \sum_{t=1}^{T-1} \left[ln(f(Y_{t+1}^e|Y_t^e)) + ln(f(ye_{t+1}|Y_{t+1}^e)) \right] \qquad (21)\end{aligned}$$

where $f(y|x)$ is the Gaussian density function and ye is a vector of all pricing errors. Y_t^e is the union of all the $Y_t^{i,e}$ vectors described earlier. In our estimation, we assume the remaining four higher order yield portfolios are priced with IID errors. Within each country, these pricing errors are independent but have the same variance. The pricing error variances are different across countries. This choice of IID errors and constant variance structure is standard. The rationale is that it forces the estimation to fit the data as much as possible through the structure of the model as opposed to pricing errors. The first term in equation

(21) corresponds to the physical dynamics (time series fit) and the second term to the risk-neutral pricing (cross-sectional fit). It is easy to see that unlike models with latent factors, the two terms in equation (21) are *almost* independent (since Y^e_{t+1} is observed). This feature will be helpful in our estimation, given the total number of parameters we need to estimate is quite large.

To save notation, denote the Gaussian density function by:

$$f^G(X, \Sigma) = -\frac{1}{2} ln((2\pi)^N |\Sigma\Sigma'|) - \frac{1}{2} tr \left(X'(\Sigma\Sigma')^{-1} X \right)$$

where N is the size of Σ. We can then write equation (21) as a function of parameters to be estimated below:

$$\mathcal{L} \ = \textstyle\sum_{t=1}^{T-1} \ \left[f^G(Y^e_{t+1} - \kappa^e_0 - \kappa^e_1 Y^e_t, \Sigma_{Y^e}) + f^G(y_{t+1} - (A_P + B'_P \Pi Y^e_{t+1}), \Sigma_{ye}) \right] \qquad (22)$$

where Σ_{ye} is a diagonal matrix whose diagonal elements correspond to the standard deviations of the pricing errors, y_{t+1} is the vector of yields priced with errors, and Π is obtained by appropriately combining together Π^i, described in the model section.

We implement our estimation procedure, relying on the fact that holding λ (the risk-neutral drift fixed), the ML estimates of all other parameters can be derived (quasi-)analytically. We omit the detailed proof but briefly sketch out our procedure here.[20] Given the functional forms of A_P and B_P, the second term only depends on λ, r_∞, σ_{ye}, and Σ_{Y^e}. Among these, only Σ_{Y^e} shows up in the first term. Therefore, if Σ_{Y^e} is given, then the two parts of equation (22) can be separately estimated. The physical dynamics parameters in the first part can be estimated analytically by typical least squares procedures. In doing this, we force the model-implied means to match the sample means of the pricing factors. Among the risk-neutral parameters that appear in the second part, the global estimates of r_∞ and Σ_{ye} can be solved

[20]The detailed proof is available from the authors upon request.

for analytically since the first order conditions with respect to these parameters have affine forms. These leave us with only Σ_{Y^e}, which is shared by both the physical dynamics part and the risk-neutral dynamics part of the likelihood function. The first order condition of Σ_{Y^e} is non-linear. We solve this non-linear equation by iteration, starting with the values implied by the VAR(1) of Y^e_{t+1}. We find in our implementation that convergence is very fast, typically obtained after four or five iterations.

Since the estimates of all other parameters follow (quasi-)analytically given the values of λ, maximizing $\mathcal{L}$ only requires numerically searching over λ which has only six unique elements in our case. Our maximization is fast and the convergence is robust. For this reason, we are able to compute the small-sample standard errors of our estimates by Monte Carlo simulations.

4 Empirical Results

4.1 Model Estimates and Fit

Table 3 reports the ML estimates of model parameters. The first panel provides estimates of the dynamic process for the global macro factors. The remaining panels show the country-level processes for the U.S., the U.K., and Germany, respectively. Overall, the model fits the data quite well as indicated by the yield measurement errors of 11-16 basis points, in line with others in the literature. The risk-neutral persistence parameters λ are around one for all unobserved pricing factors X, suggesting that for pricing purposes shocks to these factors are treated as if they are permanent. This means that long-term yields respond strongly to shocks. These persistence parameters, transformed to obtain the counterparts for P^i are higher than their physical values, implying that the market price of risk is important in explaining observed variation in long-maturity bond yields.

The next set of parameters (a, b, and c in equations (12) and (13)) show that both the

country-level macro variables and pricing factors load strongly on the world macro variables, particularly the *global* inflation rate. Recall from Table 2 that the observed pricing factors for each currency P^i (which are linear transformations of X^i) can be interpreted as the level and slope factors, and that global macro variables explain nearly two-thirds of the variation in the level factors. The loadings documented here thus demonstrate that, among the macro variables, inflation rates are most important in explaining bond yield variation. Indeed, Table 4 shows that the model-implied monetary policy rule looks similar to a global version of the Taylor rule (taking the one-month rate as the policy rate) – the coefficients on global inflation are large and statistically significant for all countries. However, the coefficients on global industrial production growth are mixed.. For the U.K. and Germany, the estimated responses to local inflation are also positive and significant, suggesting a role for country-specific policy considerations beyond the global dynamics.

We report estimates of the feedback matrix (κ_1) and the contemporaneous relationships (Σ_Y) by country. Every state variable is persistent, though less so than under the risk-neutral dynamics. The cross feedbacks are weak with only few exceptions. First, in all countries, the effects of lagged global factors are limited. Second, local IP growth in the U.S and U.K. responds negatively to lagged local inflation, though this is not true for Germany. Finally, the residual contemporaneous relationships are also relatively limited.

We next examine the general goodness of fit of the model in Table 5. To assess the pricing fit (or fit of the risk-neutral parameters), we compare the sample moments of fitted yields with those of the data. The table shows that, conditional on observing all state variables, the model provides a close match to the means, standard deviations, first-order autocorrelations, and cross-country correlations of the bond yields for all countries and maturities. We conduct the tests of difference in moments using GMM standard errors with twelve lags (Newey and West (1987)'s weights). The differences are statistically significant for a few moments, but even for these moments, the economic differences are very small. Most importantly, the

model appears to match the cross-currency correlations of bond yields well. The fitted values captures between 93-96% of cross-country yield covariance.

We also assess the fit of the overall model, including both the physical and risk-neutral dynamics, by comparing the model-implied unconditional moments at the ML estimates with the sample moments from the data. The model matches the averages of bond yields by construction, but overestimates the volatilities, particularly at the short end of the U.S. term structure and the long end of German term structure. The model slightly underestimates the persistence of yields. Finally, the model significantly underestimates the cross-currency correlations of yields and produces unconditional correlation structures that are too flat. Thus, our covariance analysis based on the estimated model parameters may only partially capture the observed co-movement of yields.

In Table 6, we examine how well the model explains the behavior of risk premiums observed in the data using Campbell and Shiller (1991)'s, henceforth C-S, regressions for a single term structure and Sutton (2000)'s modification for a pair of term structures. These coefficients speak to the correlations between risk premia and the short rate and among risk premia in different currencies. Consistent with the evidence in C-S and Dai and Singleton (2000) among others, the coefficients estimated from the data are generally negative (increasingly so at longer maturities).[21] The magnitudes are smaller than those reported in earlier studies, possibly due to risk premia becoming smaller and less volatile in recent years. For each country, our model implied C-S coefficients do demonstrate this increasingly negative patterns; however, the point estimates are significantly different from those obtained from the actual data in some cases. That said, These results indicate that our model may actually underestimate either the negative correlations between risk premia and short rates or risk premium volatilities or both. In fact, given that we attempt to match the data with the

[21] C-S and Dai and Singleton (2000) show that the expectations hypothesis would predict that the C-S regression coefficients be unity for all maturities. Negative coefficients imply positive correlations between bond excess returns and term spreads, which Dai and Singleton (2000) argue are driven by negative correlations between the short rates and risk premia.

least number of parameters (as we are dealing with several term structures), our model's performance in each currency is surprisingly good compared to many models considered by Dai and Singleton (2000).

We now turn to the model's performance in matching the co-movement of excess bond returns. The last two rows of Table 6 show that the data exhibit positive cross-currency correlations of risk premia as shown in Sutton (2000). Our model does capture this feature of the data, which is important for our study given that our goal is to measure the importance of risk premia in explaining cross-currency co-movements of bond yields. Still, in both currency pairs, our model produces risk premia correlations that are (i) too low and (ii) too flat in structure, compared to the empirical estimates from the data. This pattern resembles the mismatch in the overcall cross-country correlations of yield changes in Table 5, suggesting that underestimating the correlations of term premia contributes to underestimating the overall co-movements. This means that the extent to which we may overestimate the importance of risk premia is perhaps limited.

4.2 Impulse Responses

To understand the joint dynamics of bond yields, we first examine the joint dynamics of the state variables. In Figure 3, we plot the responses of inflation rates, IP growth rates, and level factors in all three countries to a one standard deviation shock in the world macro variables. A one standard deviation shock to the world inflation rate causes the U.S. inflation rate to initially increase by approximately 0.30%. This response dies down over 2-3 years, suggesting that under the expectations hypothesis, long-term yields should not respond much to this shock. Inflation rates in the U.K. and Germany also respond to the world inflation shock, with the former effect somewhat larger than that in the U.S. and the latter somewhat smaller. The world inflation rate also has a negative impact on U.S. IP growth, but no discernible impact on the U.K. and German IP growth rates. It is also important to point

out that the level factors of all three countries respond strongly (0.5-1%) and in the same direction to the world inflation shock. This suggests that the world inflation is important in explaining the cross-country co-movements of bond yields. Global IP growth, however, demonstrates a limited effect on country-level inflation, IP growth rate, and the level factor.

In Figure 4, we turn our attention to the responses of bond yields to various economic shocks. Since the country-level slope factors explain only a small portion of movement in yields, we focus only on the global and local inflation and IP growth rates and the country-level level factors. We also decompose the responses into those that operate through the policy channel and those that go through the risk compensation channel. The policy component is calculated based on the expectations hypothesis: the long-term yield is the expected average of the short-term rates from now to maturity (i.e. the market price of risk is zero). Thus, the policy component is a result of the monetary policy loadings to various state variables and how these state variables evolve through time. The risk compensation component is the difference between the total responses and the policy component. Note that the two components can be of opposite signs, in which case the sign of the total responses will depend on which component dominates.

Panel A-1 shows that a one-standard deviation increase in the world inflation rate (π^W) positively affects the U.S., U.K., and German 6-month rates by about 20-40 basis points. Since these responses are relatively large and in the same direction (positive), the contributions of π^W to the co-movement of short-term rates must be significant and positive. Importantly, at this maturity, the world inflation affects yields almost exclusively through the policy channel.

The responses of the U.S., U.K., and German 10-year yields to π^W are in the same direction and slightly smaller than those of the short-term yields. Due to the stationarity of π^W, the policy components of the responses are indistinguishable from zero. With the exception of Germany, almost all of the responses come from the risk compensation channel.

At long maturities, the world inflation rate is an important factor in explaining variation in bond yields. In contrast to what we observe at short maturities, it works largely through the risk compensation channel.

Panel A-2 of Figure 4 plots the responses of each country's yields to their own *local* inflation rate. For the U.S., the effects are quite small. However, the response to a shock to local inflation in the U.K. and Germany are much larger, but largely limited to the short-end policy channel. The local inflation rate has limited impact on the 10-year yields. Panels B-1 and B-2 of Figure 4 illustrate the responses of yields to shocks in the world IP growth and the local IP growth rates, respectively. Three interesting patterns emerge. First, the yield responses to g^W are generally somewhat smaller than those to π^W. Second, for the U.K. and Germany, the impacts of g^W on yields manifests slowly, becoming larger at longer horizons. This implies that, for these countries, g^W may lead other factors that matter in pricing by several months. Third, the local IP growth rates have limited impact on teh U.S. and German yields but large impact on the U.K. yields. Finally, Panel C of Figure 4 illustrate the responses of yields to shocks in the local level factor. The responses to these factors appear to be relatively large (from 35-60 basis points).

Taking all these impulse-responses together, the world inflation rate seems to be the dominant factor that affects, in the same direction, yields for all maturities and countries. We thus expect the world inflation to account for the largest fraction of the observed yield covariances. Finally, the world inflation rate operates primarily through the policy channel at the short end but through the risk compensation at the long end. This is consistent with the general findings of Ang, Bekaert, and Wei (2008) for the U.S. and Wright (2009) for several developed countries.

4.3 Variance and Covariance Decomposition

Following Ang and Piazzesi (2003) and others, we construct variance and covariance decompositions to obtain the contribution of each state variable to the variances and covariances of forecast errors of bond yields. These decompositions measure the proportion of the forecast variance or covariance (typically positive, but can be negative for the covariance) attributable to each factor, and are closely related to the impulse responses at different horizons. In addition, we decompose each coefficient matrix in the MA(∞) representation of a yield into the matrix resulting from the physical dynamics alone and that resulting from the time-varying market price of risk. The relative contributions of the policy channel (again assuming the expectations hypothesis holds) and the risk compensation channel for each shock can then be obtained. Finally, we also calculate the proportion of unconditional variation and co-variation of yields explained by macro variables. We follow the methodology of Bikbov and Chernov (2010), henceforth BC R^2, but implement it through Monte Carlo simulation. As Bikbov and Chernov (2010) point out, BC R^2 is theoretically the same as the infinite horizon variance decomposition if the model is a correct description of the data.

Table 7 reports the sum of proportions of yield forecast variance attributable to all macro variables (left panel), exclusively to global macro variables (middle panel), and to global macros variables through the risk compensation channel (right panel). The contributions of macro variables to yield variances are of similar magnitudes to those reported by Ang and Piazzesi (2003), and the importance of macro variables becomes greater at longer forecast horizons. Second, consistent with the evidence presented above, much of the total contribution comes from the global macro variables. Third, the contributions to yield variances from the global macro variables arising through the risk compensation channel show a clear pattern that is similar for all countries. These contributions are very small for six-month yields. However, as we move to five-year yields, the risk compensation channel becomes more important. The most striking finding is that as we move to ten-year yields, the contributions

through the risk compensation channel coming from the global macro variables are sizeable. This can be explained by the pattern of impulse-responses discussed in the previous section, consistent with the fact that the state variables are stationary and therefore converge to the means in the long run. This evidence indicates that the longer-term yields move with the state variables more than predicted by the expectations hypothesis. In line with this finding, Kuttner (2001) finds that although unanticipated monetary policy shocks affect bond yields of all maturities, these shocks contain little information about future monetary policy actions. The importance of time-varying market prices of risk has been discussed in detail by Campbell and Shiller (1991), Dai and Singleton (2002), and Duffee (2002).

Finally, we turn to the *covariance* decompositions reported in Table 8. The results for the U.S.-U.K. pair (in Panel A) show that, consistent with the variance decompositions presented above, the total contribution of the global macro factors that operates through the market price of risk is small for short-term rates (even negative in some cases) and quite large for long-term yields. The risk-compensation channel contributes 28-70% to the co-movements of the U.S. and U.K. five-year yields, depending on the forecast horizon. These contributions increase to roughly 80% at the ten-year maturity. The results for the U.S.-Germany pair in Panel B are broadly similar. The contributions of the risk compensation channel to the covariances between the U.S. and German yields appear to be even larger. Collectively, the evidence suggests that international macro-models need to facilitate a role for not just for fundamental macroeconomic variations and monetary policy but also for time-varying risk compensation (implicitly the role of international investors).

To provide some guidance on this point, the state variable that unambiguously contributes the most to the co-movements (and hence risk-premia variation) at long maturities is the world inflation rate. At the long end, the world inflation clearly dominates, alone accounting for 50-95% of yield covariances across five and ten-year maturities. Taken together, we conclude that the evidence suggests that the world inflation rate is the single most important

variable for explaining the cross-country co-movements of bond yields (at least for the portion of co-movements captured by our model). The world inflation works mostly through the policy channel at the short end but almost exclusively through the risk compensation channel at the long end.[22]

5 Additional Consideration

5.1 Policy Divergence

We document the primacy of global macro factors in determining correlations among bond markets. It is, however, important to acknowledge that additional global factors, beyond inflation and industrial production, might be important. In the spirit of our principle components exercise detailed above, we conduct an additional component extraction for the entire set of bonds across all markets at once (these are available upon request). While the components extracted in this manner again suggest a level and slope effect, an additional important component is related to a "policy divergence" factor that buys U.S. bonds and shorts German bonds. To explore the robustness of our findings to the inclusion of a policy divergence factor, we incorporate an additional global variable, r^{US-DE}, constructed as the difference between the 3-month U.S. and German bill yields.

Table 9 provides a covariance decomposition of the U.S.-U.K. (Panel A) and the U.S.-German (Panel B) bond pairs under an augmented model that adds the policy divergence factor to M_t^W. Consistent with the evidence presented above, the contribution of the world inflation and IP factors in explaining correlations across bond markets remains elevated. For the correlations among short-term bonds, the policy channel dominates, whereas the

[22]Consistent with the reported impulse-responses, the world IP growth rate is unimportant at short forecast horizons but somewhat important at long forecast horizons. For example, at 5-year forecast horizons, the world IP growth rate explains roughly one-third of 10-year yield covariances both for the U.S.-U.K. pair and for the U.S.-Germany pair.

risk compensation channel is first-order for correlations among long-term bonds. The policy divergence variable is relevant for the correlations among short-term bonds; however, the effect is largely limited to the policy channel. The inclusion of the policy divergence channel is not relevant at all for long-term bonds and our main results regarding the importance of the global inflation factor are unchanged.

5.2 Exchange Rates

Given our focus on cross-country interest rates, one natural question is what our model implies for the dynamics of exchange rates. If we were to assume that bond markets are complete and fully integrated, then the difference in nominal SDFs of two countries (used in our bond pricing) determines the change in their exchange rate (see Backus, Foresi, and Telmer (2001) and Inci and Lu (2004) among others).[23] Let S_t^i denote the time-t exchange rate, expressed as the number of dollars per one unit of non U.S. currency i, and $s_t^i = \log(S_t^i)$. Since we have fully specified the $\mathbb{P}$ and $\mathbb{Q}$ dynamics of the pricing factors as well as the short rate for each economy, the stochastic discount factor for the i^{th} economy can be written as:

$$SDF_{t+1}^i = e^{-y_{1,t}^i} \frac{f_t^{\mathbb{Q}}(P_{t+1}^i)}{f_t^{\mathbb{P}}(P_{t+1}^i)} \tag{23}$$

where $f_t^{\mathbb{Q}}(.)$ and $f_t^{\mathbb{P}}(.)$ denote the density under the risk-neutral and physical measures, respectively. Then, no arbitrage implies that

$$\begin{aligned} \Delta s_{t+1}^i &\equiv s_{t+1}^i - s_t^i = \log(SDF_{t+1}^i / SDF_{t+1}^{US}) \\ &= (y_{1,t}^{US} - y_{1,t}^i) + \log\left(\frac{f_t^{\mathbb{Q}}(P_{t+1}^i)}{f_t^{\mathbb{P}}(P_{t+1}^i)} \frac{f_t^{\mathbb{P}}(P^US_{t+1})}{f_t^{\mathbb{Q}}(P^US_{t+1})}\right) \end{aligned} \tag{24}$$

Notice that the expected change in exchange rate does not follow the uncovered interest rate parity (UIRP). For an investor with dollar as the numeraire currency, the second term in

[23]In fact, the assumption that the modeled SDFs are minimum-variance SDFs are sufficient.

expression (24) is the excess return on currency i, the foreign exchange risk premium, which is generally state-dependent. Dong (2006) shows that his model, which in spirit is similar to ours, produces the path of exchange rates that matches the data very well. Since fitting the exchange rate path is not of direct interest in this paper and doing so may possibly worsen the fitting of bond yields (plus additional assumptions are needed), we do not use equation (24) in our estimation.[24] For future research, we plan to further explore the implications of our model for exchange-rate movements.

6 Conclusions

This paper provides a possible explanation for the observed degree of co-movement in yield curves across different currencies and maturities. This issue is important as a first step in better understanding the linkages among international asset markets. We jointly model the term structures of interest rates in several countries using an affine model in which the state variables follow a vector autoregressive (VAR) process. We use only a few pricing factors based on linear combinations of yields, but also allow macro variables (including the inflation rate and industrial production growth) to influence the pricing factors in the physical dynamics. The model is estimated to match the data for the U.S., the U.K., and Germany. The policy channel refers to the propagation of shocks through the physical dynamics of short-term rates, as predicted by the expectations hypothesis. The risk compensation channel refers to the responses of bond yields to shocks that are in excess of those transmitted through the policy channel. We find that the role for global macro variables in the policy and the risk compensation channels account for a sizable portion of the cross-country covariances of 5-year bond yields. As we move to longer maturities, these numbers shift considerably in favor of the risk compensation channel. The most important state variable is the world inflation rate, which alone explains over half of the co-variances among bond yields at almost

[24] See Brandt, Cochrane, and Santa-Clara (2006).

all maturities and forecast horizons. An acknowledgement of the risk compensation channel, often absent or secondary in the macroeconomic literature, is critical for understanding the behavior of the yield curve, in general, and long maturity (benchmark) bonds, in particular.

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	Level	1-Month Change	1-Year Change	Conditional (Macros Only)	Conditional (Macros and Yields)
Panel A: U.S. - U.K.					
Inflation Rate	0.742	0.174	0.386	-	-
IP Growth	0.635	0.024	0.488	-	-
6-Month Yield	0.767	0.233	0.528	0.414	0.039
1-Year Yield	0.783	0.268	0.536	0.470	0.106
2-Year Yield	0.843	0.387	0.615	0.563	0.210
5-Year Yield	0.908	0.456	0.691	0.562	0.272
7-Year Yield	0.919	0.499	0.705	0.527	0.304
10-Year Yield	0.924	0.495	0.694	0.496	0.328
Panel B: U.S. - Germany					
Inflation Rate	0.319	0.270	0.075	-	-
IP Growth	0.095	0.203	0.314	-	-
6-Month Yield	0.407	0.399	0.230	0.509	0.265
1-Year Yield	0.454	0.432	0.305	0.548	0.269
2-Year Yield	0.560	0.497	0.441	0.644	0.328
5-Year Yield	0.745	0.554	0.607	0.720	0.398
7-Year Yield	0.797	0.573	0.660	0.721	0.423
10-Year Yield	0.832	0.594	0.691	0.711	0.450

Table 1: Cross-Country Correlations. This table reports the cross-country correlations of (1-year moving) inflation rate, (1-year moving) IP growth rate, and the LIBOR-swap zero-coupon yields for maturities 6 months, 1 year, 2 years, 5 years, 7 years, and 10 years. The sample period is from 1987:04 to 2008:03. The frequency is monthly. The correlations are calculated based on: (i) the level, (ii) the monthly change, (iii) the annual change (overlap), (iv) the forecast errors based on a linear model with inflation and output gap up to 12 lags, and (v) the forecast errors based on a linear model with inflation, output gap, and bond yields up to 12 lags. Panel A reports the correlations between the U.S. and the U.K. variables. Panel B reports the correlations between the U.S. and German variables.

	% Variance Explained	Loading on Yield						% Explained by Global Macros	% Explained by Global and Country-Specific Macros
		6-Month	1-Year	2-Year	5-Year	7-Year	10-Year		
U.S.									
1st PC	94.35	0.45	0.44	0.44	0.39	0.37	0.35	61.41	64.43
2nd PC	5.45	0.51	0.40	0.13	-0.30	-0.43	-0.53	19.35	28.46
U.K.									
1st PC	96.49	0.48	0.45	0.42	0.38	0.37	0.36	77.36	82.25
2nd PC	3.28	0.52	0.39	0.09	-0.29	-0.43	-0.55	0.18	48.43
Germany									
1st PC	94.87	0.48	0.46	0.45	0.38	0.34	0.31	66.13	75.74
2nd PC	4.80	0.46	0.37	0.14	-0.30	-0.47	-0.57	15.24	44.17

Table 2: Principal Components of Yields. This table reports (i) the variance of yields explained by the first few principal components in each currency, (ii) the loadings of these principal components on yields, and (iii) the relationships between these principal components with the world and country-specific macro variables. The macro variables (M) include inflation (π) and IP growth (g). The country-specific macro variables are constructed as the residuals from projecting the actual macro variables for each country on the global macro variables.

	κ_0	κ_1 π^W	κ_1 g^W	Σ π^W	Σ g^W
π^W	0.000	0.981	0.017	0.002	0
	(0.000)	(0.016)	(0.005)	(0.000)	
g^W	0.004	-0.059	0.893	0.001	0.012
	(0.003)	(0.096)	(0.033)	(0.001)	(0.001)

Table 3: Maximum Likelihood Estimates. This table reports maximum likelihood estimates of the model parameters (notations as described in the model section). The priced factors (P) are the principal components of yields. The macro variables (M) which partially explain the dynamics of the priced factors include inflation (π) and IP growth (g). The superscripts W, US, UK, and DE denote the world, the U.S., the U.K., and Germany, respectively. Small-sample standard errors are calculated by Monte Carlo simulation at the maximum likelihood parameter estimates, and are in parentheses.

Parameters Specific to U.S. Term Structure

	$X^{US,1}$	$X^{US,2}$
λ	0.995	0.969
	(0.001)	(0.001)

r_∞	Measurement Error (bps)
0.111	11.200
(0.011)	(0.252)

Factor Loadings

	π^{US}	g^{US}	$P^{US,1}$	$P^{US,2}$
π^W	0.812	-0.645	2.623	0.261
	(0.111)	(0.437)	(0.775)	(0.361)
g^W	0.051	0.765	0.628	-0.160
	(0.033)	(0.149)	(0.247)	(0.094)
$\pi^{US,e}$	1	0	-0.793	-0.722
			(1.333)	(0.489)
$g^{US,e}$	0	1	0.325	0.037
			(0.239)	(0.090)

Physical Dynamics

		κ_1					
	κ_0	π^W	g^W	$\pi^{US,e}$	$g^{US,e}$	$P^{US,1,e}$	$P^{US,2,e}$
$\pi^{US,e}$	0.001	-0.016	0.010	0.881	-0.009	0	0
	(0.001)	(0.009)	(0.006)	(0.038)	(0.007)		
$g^{US,e}$	0.007	-0.013	0.048	-0.502	0.837	0	0
	(0.003)	(0.036)	(0.028)	(0.170)	(0.034)		
$P^{US,1,e}$	0.004	0.009	0.043	0.017	-0.002	0.925	-0.042
	(0.007)	(0.050)	(0.040)	(0.260)	(0.053)	(0.037)	(0.098)
$P^{US,2,e}$	0.002	0.000	-0.033	-0.024	0.004	0.005	0.942
	(0.003)	(0.023)	(0.014)	(0.077)	(0.020)	(0.013)	(0.034)

	Σ					
	π^W	g^W	$\pi^{US,e}$	$g^{US,e}$	$P^{US,1,e}$	$P^{US,2,e}$
$\pi^{US,e}$	0	0	0.0019	0	0	0
			(0.0001)			
$g^{US,e}$	0	0	0.0026	0.0092	0	0
			(0.0007)	(0.0004)		
$P^{US,1,e}$	0	0	0	0	0.0105	0
					(0.0010)	
$P^{US,2,e}$	0	0	0	0	-0.0003	0.0030
					(0.0005)	(0.0003)

Table 3, cont'd: Maximum Likelihood Estimates.

Parameters Specific to U.K. Term Structure

	$X^{UK,1}$	$X^{UK,2}$
λ	1.000	0.946
	(0.000)	(0.002)

r_∞	Measurement Error (bps)
2.977	15.710
(1.437)	(0.327)

Factor Loadings

	π^{UK}	g^{UK}	$P^{UK,1}$	$P^{UK,2}$
π^W	1.434	-0.092	4.625	-0.041
	(0.337)	(0.366)	(0.981)	(0.379)
g^W	0.086	0.549	0.204	-0.002
	(0.083)	(0.114)	(0.212)	(0.112)
$\pi^{UK,e}$	1	0	1.009	-0.644
			(0.561)	(0.167)
$g^{UK,e}$	0	1	0.655	0.153
			(0.216)	(0.062)

Physical Dynamics

		κ_1					
	κ_0	π^W	g^W	$\pi^{UK,e}$	$g^{UK,e}$	$P^{UK,1,e}$	$P^{UK,2,e}$
$\pi^{UK,e}$	-0.001	-0.001	0.021	0.938	0.000	0	0
	(0.001)	(0.023)	(0.014)	(0.031)	(0.015)		
$g^{UK,e}$	-0.001	-0.021	0.012	-0.206	0.749	0	0
	(0.004)	(0.038)	(0.024)	(0.092)	(0.046)		
$P^{UK,1,e}$	0.003	0.072	-0.033	0.012	0.126	0.882	0.081
	(0.005)	(0.056)	(0.041)	(0.092)	(0.066)	(0.043)	(0.112)
$P^{UK,2,e}$	0.002	0.017	-0.005	0.007	0.040	-0.007	0.890
	(0.002)	(0.018)	(0.013)	(0.028)	(0.020)	(0.015)	(0.054)

	Σ					
	π^W	g^W	$\pi^{UK,e}$	$g^{UK,e}$	$P^{UK,1,e}$	$P^{UK,2,e}$
$\pi^{UK,e}$	0	0	0.0034	0	0	0
			(0.0003)			
$g^{UK,e}$	0	0	0.0009	0.0115	0	0
			(0.0007)	(0.0006)		
$P^{UK,1,e}$	0	0	0	0	0.0121	0
					(0.0007)	
$P^{UK,2,e}$	0	0	0	0	0.0008	0.0037
					(0.0003)	(0.0002)

Table 3, cont'd: Maximum Likelihood Estimates.

Parameters Specific to German Term Structure

	$X^{DE,1}$	$X^{DE,2}$	r_∞	Measurement Error (bps)
λ	0.995	0.975	0.120	13.883
	(0.001)	(0.002)	(0.018)	(0.290)

Factor Loadings

	π^{DE}	g^{DE}	$P^{DE,1}$	$P^{DE,2}$
π^W	0.567	0.452	3.153	-0.024
	(0.273)	(0.710)	(0.922)	(0.294)
g^W	-0.202	0.874	-0.353	0.153
	(0.070)	(0.202)	(0.227)	(0.072)
$\pi^{DE,e}$	1	0	1.273	-0.416
			(0.652)	(0.189)
$g^{DE,e}$	0	1	-0.126	-0.141
			(0.114)	(0.034)

Physical Dynamics

		κ_1					
	κ_0	π^W	g^W	$\pi^{DE,e}$	$g^{DE,e}$	$P^{DE,1,e}$	$P^{DE,2,e}$
$\pi^{DE,e}$	0.001	0.006	-0.023	0.920	-0.003	0	0
	(0.001)	(0.018)	(0.013)	(0.033)	(0.009)		
$g^{DE,e}$	-0.004	-0.026	0.105	0.016	0.805	0	0
	(0.006)	(0.087)	(0.043)	(0.163)	(0.045)		
$P^{DE,1,e}$	-0.003	0.060	-0.006	0.024	-0.001	0.923	0.114
	(0.005)	(0.047)	(0.027)	(0.094)	(0.031)	(0.038)	(0.091)
$P^{DE,2,e}$	0.004	0.000	0.012	-0.016	-0.031	0.012	0.851
	(0.002)	(0.017)	(0.010)	(0.032)	(0.011)	(0.015)	(0.043)

	Σ					
	π^W	g^W	$\pi^{DE,e}$	$g^{DE,e}$	$P^{DE,1,e}$	$P^{DE,2,e}$
$\pi^{DE,e}$	0	0	0.0034	0	0	0
			(0.0003)			
$g^{DE,e}$	0	0	-0.0042	0.0180	0	0
			(0.0014)	(0.0009)		
$P^{DE,1,e}$	0	0	0	0	0.0081	0
					(0.0006)	
$P^{DE,2,e}$	0	0	0	0	0.0004	0.0037
					(0.0004)	(0.0002)

Table 3, cont'd: Maximum Likelihood Estimates.

	U.S.	U.K.	Germany
World Inflation	1.023	2.298	1.577
	(0.439)	(0.467)	(0.409)
World IP Growth	0.394	0.102	-0.267
	(0.138)	(0.100)	(0.109)
Country-Specific Inflation	0.121	0.969	0.879
	(0.744)	(0.249)	(0.274)
Country-Specific IP Growth	0.124	0.209	0.022
	(0.134)	(0.096)	(0.050)
R-Squared	0.560	0.829	0.799

Table 4: Implied Monetary Policy Rules. This table reports the implied monetary policy rules at the maximum likelihood parameter estimates. The loadings are calculated taking the 1-month interest rates as the policy rates. Small-sample standard errors, calculated by Monte Carlo simulation at the maximum likelihood parameter estimates, are in parentheses.

	U.S.			U.K.			Germany		
	6-Month	5-Year	10-Year	6-Month	5-Year	10-Year	6-Month	5-Year	10-Year
Mean									
Data	0.051	0.063	0.067	0.072	0.073	0.073	0.047	0.055	0.060
Fitted	0.051	0.063	0.067	0.071	0.073	0.073	0.047	0.054	0.060
Different? (p-value)	(0.007)	(0.283)	(0.470)	(0.053)	(0.017)	(0.114)	(0.037)	(0.037)	(0.108)
Model Implied	0.051	0.063	0.067	0.071	0.073	0.073	0.047	0.054	0.060
Standard Deviation									
Data	0.021	0.018	0.017	0.030	0.023	0.023	0.022	0.017	0.015
Fitted	0.021	0.018	0.017	0.029	0.023	0.022	0.022	0.017	0.015
Different? (p-value)	(0.189)	(0.000)	(0.000)	(0.048)	(0.004)	(0.000)	(0.789)	(0.128)	(0.005)
Model Implied	0.027	0.021	0.018	0.031	0.025	0.024	0.025	0.020	0.018
AR(1) Coefficient									
Data	0.992	0.987	0.986	0.991	0.992	0.994	0.995	0.992	0.993
Fitted	0.991	0.988	0.984	0.990	0.994	0.993	0.994	0.993	0.992
Different? (p-value)	(0.266)	(0.000)	(0.000)	(0.038)	(0.001)	(0.000)	(0.740)	(0.033)	(0.001)
Model Implied	0.964	0.963	0.963	0.964	0.960	0.957	0.966	0.972	0.966
Yield Correlation									
Data	-	-	-	0.767	0.908	0.924	0.407	0.745	0.832
Fitted	-	-	-	0.762	0.905	0.928	0.405	0.741	0.832
Different? (p-value)	-	-	-	(0.181)	(0.005)	(0.001)	(0.984)	(0.037)	(0.007)
Model Implied	-	-	-	0.600	0.624	0.594	0.224	0.503	0.534
			-						
Portion of Covariance Explained by Fitted Values	-	-	-	0.934	0.954	0.937	0.951	0.958	0.930

Table 5: Goodness of Fit. This table compares moments of actual yields, model-fitted yields, and unconditional moments implied by the model at the maximum likelihood parameter estimates. The test of difference in value of the sample moment is based on GMM with one degree of freedom. Corresponding p-values are in parentheses. The last row reports the proportion of yield covariance explained by the model-fitted values for each pair of country and each maturity.

	U.S.			U.K.			Germany		
	2-Year	5-Year	10-Year	2-Year	5-Year	10-Year	2-Year	5-Year	10-Year
C-S Regression Coefficient									
Data	-0.448	-0.757	-1.066	0.656	0.313	-0.233	-0.122	-0.707	-1.087
	(0.330)	(0.378)	(0.475)	(0.235)	(0.353)	(0.467)	(0.289)	(0.369)	(0.416)
Model Implied	0.171	-0.283	-0.780	0.546	0.224	-0.338	1.827	0.924	-0.021
Excess Bond Return Correlation									
Data	-	-	-	0.543	0.697	0.693	0.437	0.577	0.664
	-	-	-	(0.013)	(0.006)	(0.006)	(0.021)	(0.012)	(0.007)
Model Implied	-	-	-	0.418	0.372	0.316	0.045	0.188	0.232

Table 6: Deviations from Expectations Hypothesis. This table compares (i) coefficients of Campbell and Shiller (1991) regression and (ii) correlations of excess bond returns across currencies, estimated from the data, with those implied by the model at the maximum likelihood parameter estimates. Excess returns are calculated over one-year holding period. Standard errors are in parentheses underneath the corresponding estimates.

	All Macros Total Contribution			World Macros Total Contribution			World Macros Through Market Price of Risk		
Horizon	6-Month	5-Year	10-Year	6-Month	5-Year	10-Year	6-Month	5-Year	10-Year
U.S.									
3 Months	0.54	0.45	0.40	0.52	0.39	0.31	-0.10	0.16	0.23
1 Year	0.67	0.56	0.49	0.66	0.51	0.40	-0.05	0.28	0.33
5 Years	0.75	0.70	0.66	0.74	0.67	0.60	0.02	0.48	0.54
BC R-Squared	0.75	0.72	0.69	0.74	0.68	0.62	0.02	0.50	0.57
U.K.									
3 Months	0.57	0.54	0.52	0.34	0.28	0.24	-0.07	0.03	0.15
1 Year	0.71	0.67	0.65	0.55	0.45	0.38	-0.07	0.10	0.26
5 Years	0.84	0.81	0.79	0.75	0.66	0.59	-0.03	0.28	0.46
BC R-Squared	0.86	0.82	0.80	0.77	0.68	0.61	-0.03	0.30	0.48
Germany									
3 Months	0.60	0.50	0.45	0.41	0.27	0.17	0.05	0.00	0.05
1 Year	0.67	0.56	0.51	0.48	0.37	0.26	0.02	0.00	0.10
5 Years	0.81	0.74	0.68	0.71	0.63	0.52	0.01	0.17	0.31
BC R-Squared	0.79	0.72	0.67	0.69	0.61	0.52	0.01	0.16	0.30

Table 7: Variance Decomposition. This table reports the contributions to yield variance of (i) all macro variables (global and country-specific), (ii) all global macro variables only, and (iii) all global macro variables only through the market price of risk channel. The macro variables (M) include inflation (π) and IP growth (g). The variance contributions through the market price of risk channel are calculated by subtracting the forecast error variance predicted by the expectations hypothesis from the total forecast error variance implied by the model. BC R-Squared is the unconditional yield variance explained by the variables of interest, calculated by Bikbov and Chernov (2010)'s methodology. The bond maturities are in the columns and the forecast horizons are in the rows.

Panel A: U.S.-U.K.

		Contribution Through Physical Dynamics			Contribution Through Market Price of Risk		
Maturity	Horizon	π^W	g^W	Total	π^W	g^W	Total
6-Month	3 months	0.54	0.75	1.28	0.04	-0.32	-0.28
	1 year	0.36	0.77	1.13	0.05	-0.17	-0.13
	5 years	0.24	0.76	1.00	0.06	-0.06	0.00
	BC R-Squared	-	-	1.03	-	-	0.01
5-Year	3 months	0.15	0.57	0.72	0.59	-0.31	0.28
	1 year	0.10	0.45	0.56	0.54	-0.10	0.44
	5 years	0.05	0.29	0.34	0.50	0.17	0.66
	BC R-Squared	-	-	0.33	-	-	0.70
10-Year	3 months	0.05	0.22	0.27	0.77	-0.04	0.73
	1 year	0.04	0.18	0.21	0.72	0.06	0.79
	5 years	0.01	0.10	0.12	0.63	0.25	0.88
	BC R-Squared	-	-	0.11	-	-	0.91

Panel B: U.S.-Germany

		Contribution Through Physical Dynamics			Contribution Through Market Price of Risk		
Maturity	Horizon	π^W	g^W	Total	π^W	g^W	Total
6-Month	3 months	-1.10	1.89	0.79	-0.03	0.24	0.21
	1 year	-11.34	9.04	-2.30	-1.02	4.32	3.30
	5 years	0.65	0.35	1.00	0.15	-0.15	0.00
	BC R-Squared	-	-	1.08	-	-	0.02
5-Year	3 months	0.51	0.74	1.24	1.25	-1.50	-0.24
	1 year	0.25	0.54	0.79	0.90	-0.69	0.21
	5 years	0.07	0.27	0.34	0.61	0.05	0.66
	BC R-Squared	-	-	0.33	-	-	0.68
10-Year	3 months	0.11	0.25	0.36	0.95	-0.31	0.64
	1 year	0.07	0.21	0.28	0.86	-0.14	0.72
	5 years	0.02	0.11	0.13	0.66	0.22	0.87
	BC R-Squared	-	-	0.12	-	-	0.87

Table 8: Covariance Decomposition. This table reports detailed decompositions of covariance of yields in two countries for the same maturity. Factors include world inflation (π^W) and IP growth (g^W). The contribution of each factor to yield covariance is further decomposed into the part that works through the factor physical dynamics and the part that works through the market price of risk. The latter is calculated by subtracting the former from the total yield covariance implied by the model. BC R-Squared is the unconditional yield covariance explained by the variables of interest, calculated by a modified Bikbov and Chernov (2010)'s methodology. Panels A and B report the yield covariance decompositions for the U.S.-U.K. pair and for the U.S.-Germany pair, respectively.

Panel A: U.S.-U.K.

		Contribution Through Physical Dynamics				Contribution Through Market Price of Risk			
Maturity	Horizon	π^W	g^W	r^{US-DE}	Total	π^W	g^W	r^{US-DE}	Total
6-Month	3 months	0.86	0.06	0.11	1.03	0.06	-0.06	-0.04	-0.03
	1 year	0.67	0.21	0.14	1.02	0.09	-0.08	-0.04	-0.02
	5 years	0.37	0.30	0.33	1.00	0.07	-0.02	-0.05	0.00
	BC R-Squared	-	-	-	1.02	-	-	-	-0.04
5-Year	3 months	0.02	0.17	0.22	0.41	0.92	-0.12	-0.21	0.59
	1 year	-0.02	0.15	0.23	0.36	0.89	-0.04	-0.21	0.64
	5 years	0.19	0.09	0.28	0.56	0.46	0.14	-0.16	0.44
	BC R-Squared	-	-	-	0.71	-	-	-	0.27
10-Year	3 months	-0.01	0.01	0.11	0.11	0.96	0.04	-0.10	0.89
	1 year	0.01	0.01	0.10	0.11	0.91	0.07	-0.09	0.89
	5 years	0.16	0.02	0.08	0.27	0.57	0.18	-0.01	0.73
	BC R-Squared	-	-	-	0.31	-	-	-	0.67

Panel B: U.S.-Germany

		Contribution Through Physical Dynamics				Contribution Through Market Price of Risk			
Maturity	Horizon	π^W	g^W	r^{US-DE}	Total	π^W	g^W	r^{US-DE}	Total
6-Month	3 months	1.22	0.16	-0.29	1.09	0.03	-0.12	0.00	-0.09
	1 year	0.89	0.34	-0.20	1.03	0.08	-0.11	0.00	-0.03
	5 years	0.32	0.51	0.03	0.86	0.18	-0.02	-0.02	0.14
	BC R-Squared	-	-	-	0.88	-	-	-	0.12
5-Year	3 months	0.04	0.26	0.02	0.32	0.96	-0.18	-0.10	0.68
	1 year	-0.06	0.22	0.05	0.21	0.99	-0.09	-0.12	0.79
	5 years	-0.18	0.09	0.15	0.06	0.86	0.17	-0.09	0.94
	BC R-Squared	-	-	-	0.18	-	-	-	0.84
10-Year	3 months	-0.13	0.04	0.06	-0.03	1.07	0.06	-0.10	1.03
	1 year	-0.13	0.02	0.07	-0.04	1.06	0.08	-0.10	1.04
	5 years	-0.04	-0.01	0.07	0.02	0.77	0.22	0.00	0.98
	BC R-Squared	-	-	-	0.07	-	-	-	0.95

Table 9: Covariance Decomposition with Three World Factors. This table reports detailed decompositions of covariance of yields in two countries for the same maturity. Factors include world inflation (π^W), world IP growth (g^W), and the difference between the U.S. and German 3-month interest rates (r^{US-DE}). The contribution of each factor to yield covariance is further decomposed into the part that works through the factor physical dynamics and the part that works through the market price of risk. The latter is calculated by subtracting the former from the total yield covariance implied by the model. BC R-Squared is the unconditional yield covariance explained by the variables of interest, calculated by a modified Bikbov and Chernov (2010)'s methodology. Panels A and B report the yield covariance decompositions for the U.S.-U.K. pair and for the U.S.-Germany pair, respectively.

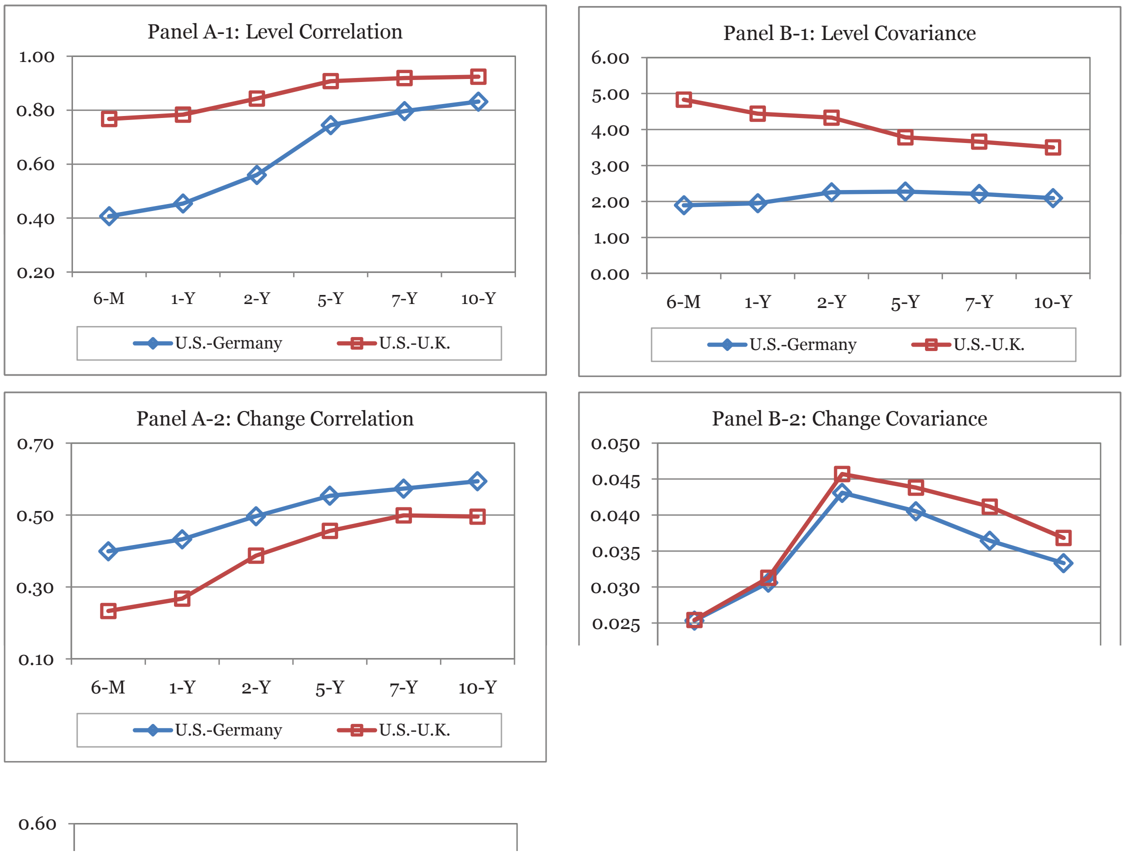

Figure 1: Cross-Country Correlations and Covariances. This figure plots cross-country correlations and covariances for the interest rates and bond yields in two pairs of countries: (a) the U.S. and Germany, and (b) the U.S. and the U.K. The frequency is monthly. The sample period is 1987:04-2008:03. The correlations and covariances are calculated based on: (i) the level, (ii) the monthly change, and (iii) the forecast errors based on a linear model with state variables and bond yields up to 12 lags. The diamond markers and blue lines are for the U.S.-Germany pair, and the square markers and red lines are for the U.S.-U.K. pair. Panel A reports the correlations. Panel B reports the covariances.

Figure 2: Time Series of 3-Month Rates and 10-Year Yields. This figure plots the time series of 3-month LIBOR rates and 10-year swap-implied zero-coupon yields for the U.S., Germany, and the U.K. for the period from 1987:04 to 2008:03. The blue solid line represents the U.S. The red solid line with square markers represent Germany. The green solid line with triangle markers represent the U.K. The dashed rectangular boxes identify two periods during which the monetary policies among these countries diverged. Panel A presents the 3-month LIBOR rates. Panel B presents the 10-year swap-implied zero-coupon yields.

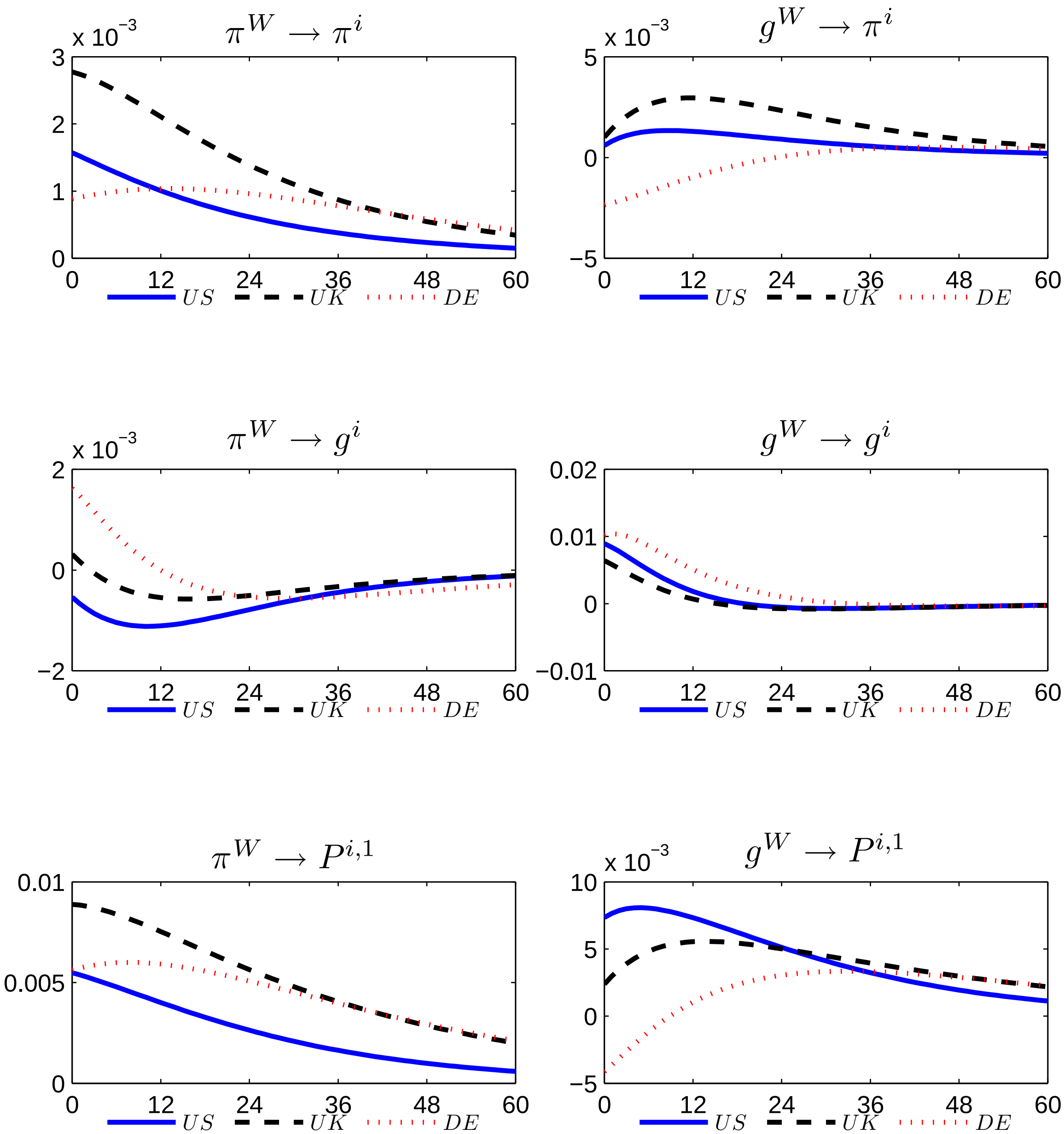

Figure 3: Impulse Responses of Macroeconomic Variables and the Level Factor. This figure plots the impulse responses of inflation (π^{US}, π^{UK}, and π^{DE}), IP growth (g^{US}, g^{UK}, and g^{DE}) and the first principal component of yields, or the level factor, ($P^{US,1}$, $P^{UK,1}$, and $P^{DE,1}$) to a Choleski one standard deviation innovation in the world inflation and world IP growth (π^W and g^W).

Panel A-1: World Inflation on Yields

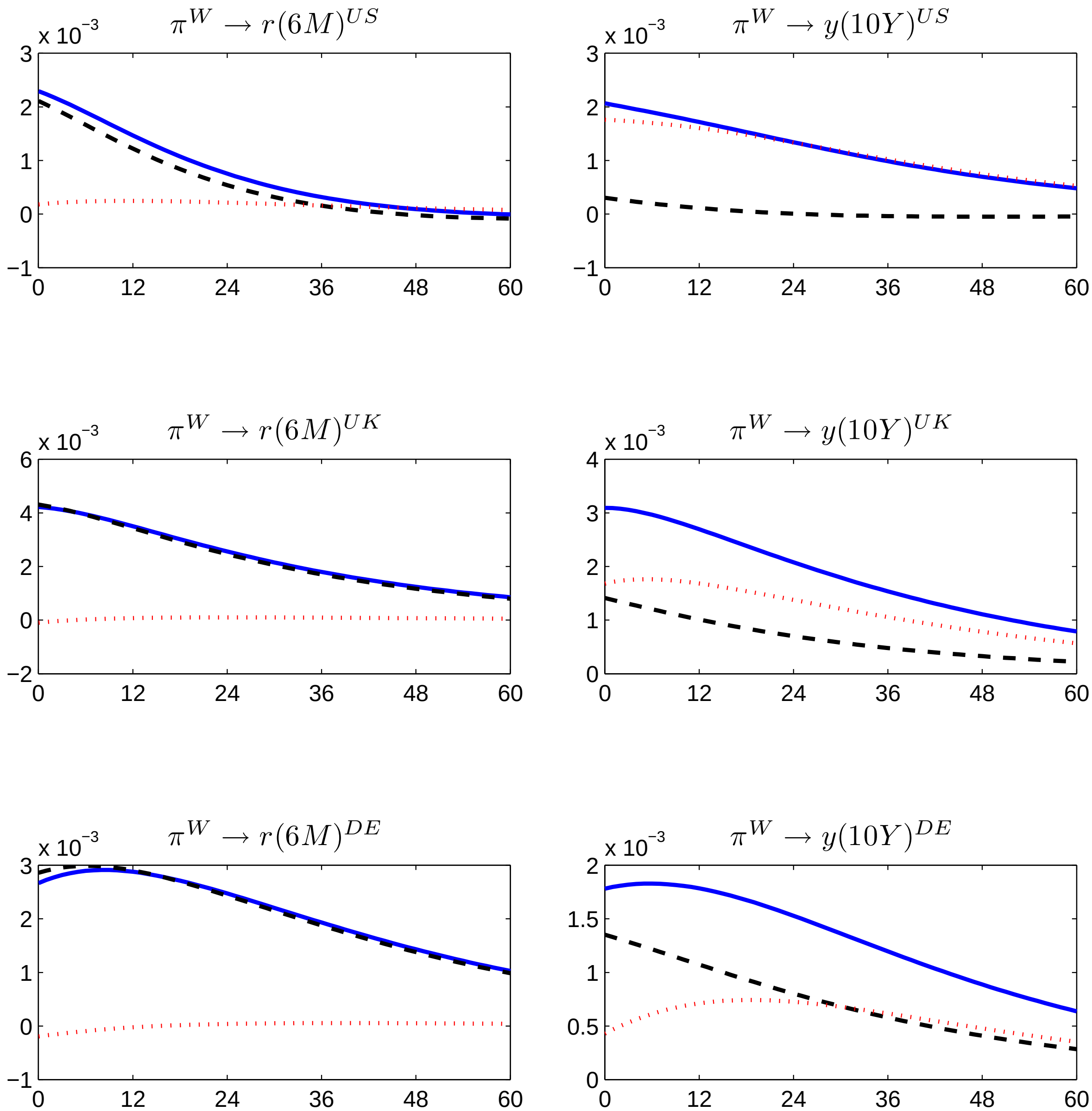

Figure 4: Impulse Responses of Yields. This figure plots the impulse responses of interest rates and swap-implied zero-coupon yields to a Choleski one standard deviation innovation in each of the following state variables: world inflation (π^W) and country-specific inflations ($\pi^{i,e}$) in Panels A-1 and A-2, world IP growth (g^W) and country-specific IP growths ($g^{i,e}$) in Panels B-1 and B-2, and the country-specific level factor ($P^{i,1,e}$) in Panel C. The blue solid line represents the total responses. The black dashed line represents the physical responses. The dotted red line represents the market price of risk responses.

Panel A-2: Country-Specific Inflations on Yields

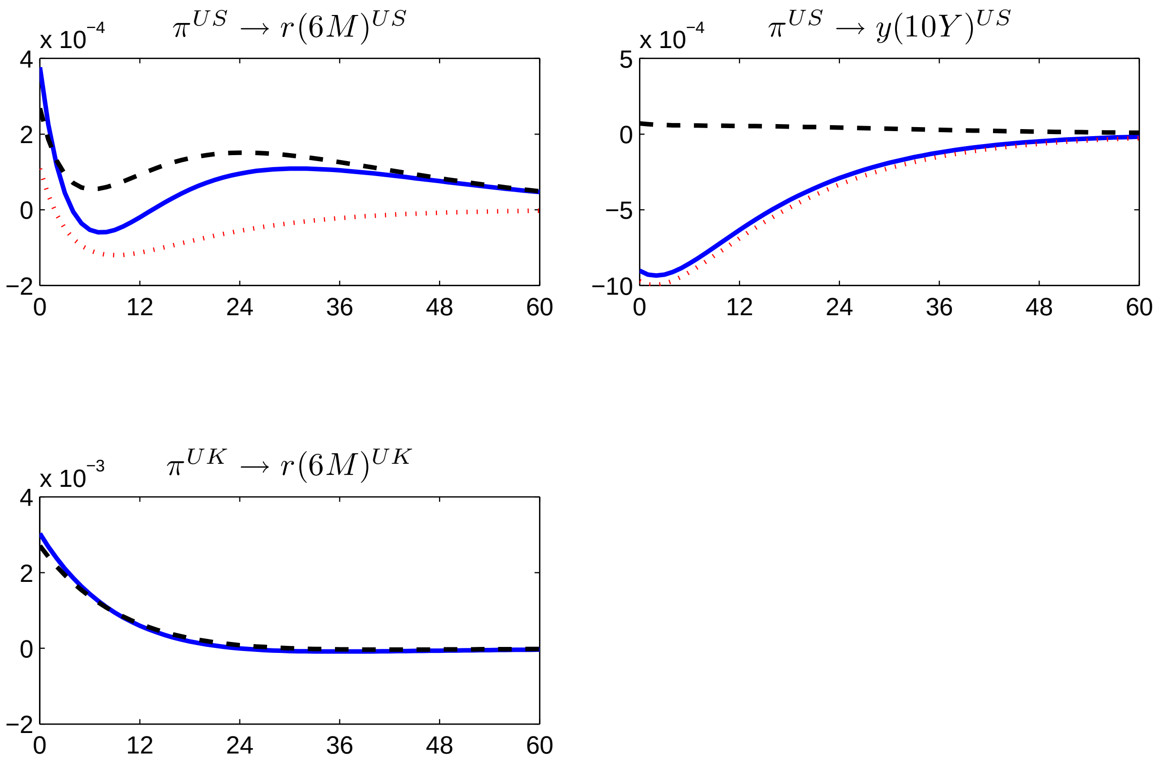

Figure 4, cont'd: Impulse Responses of Yields.

Panel B-1: World IP Growth on Yields

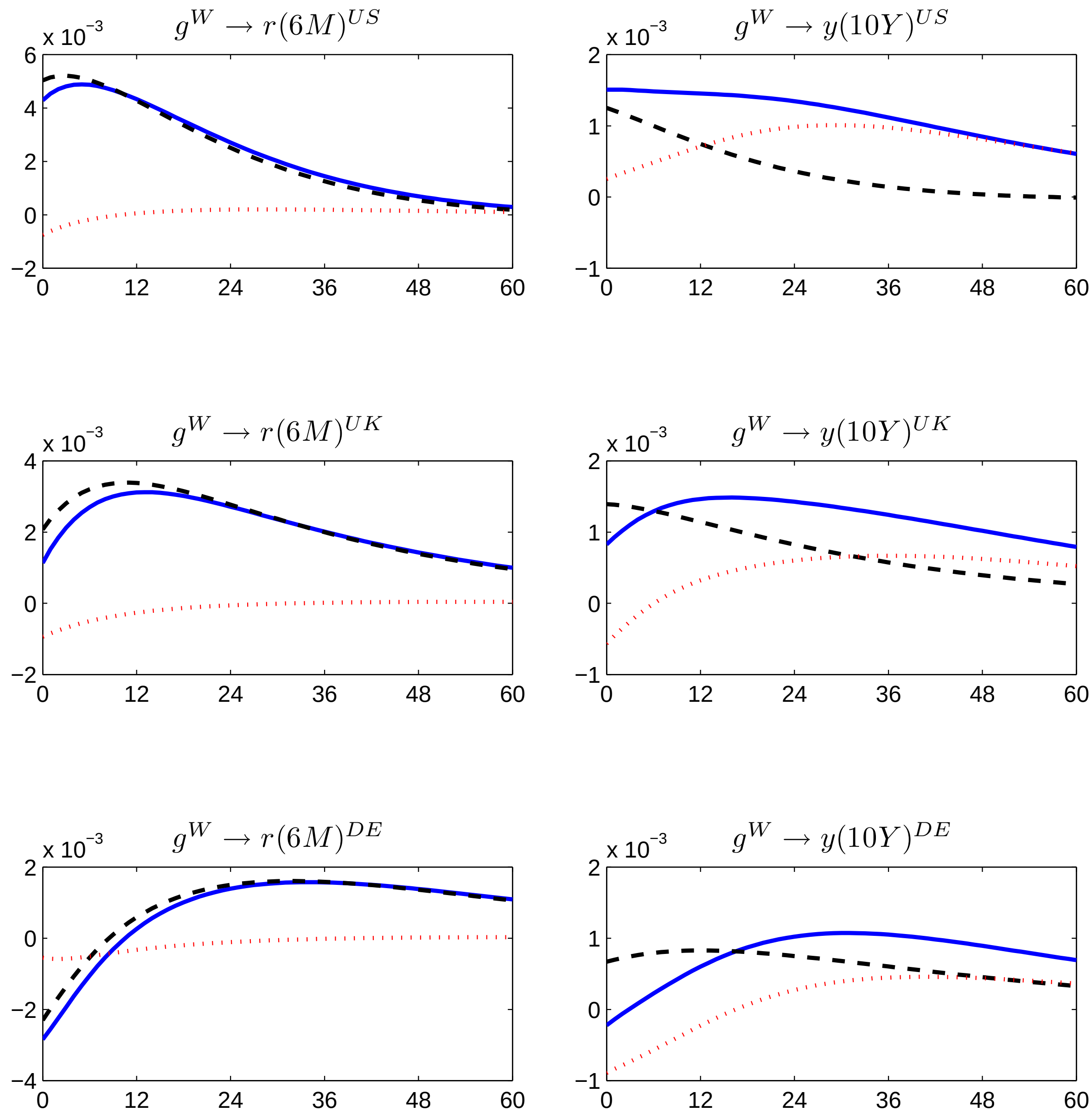

Figure 4, cont'd: Impulse Responses of Yields.

Panel B-2: Country-Specific IP Growths on Yields

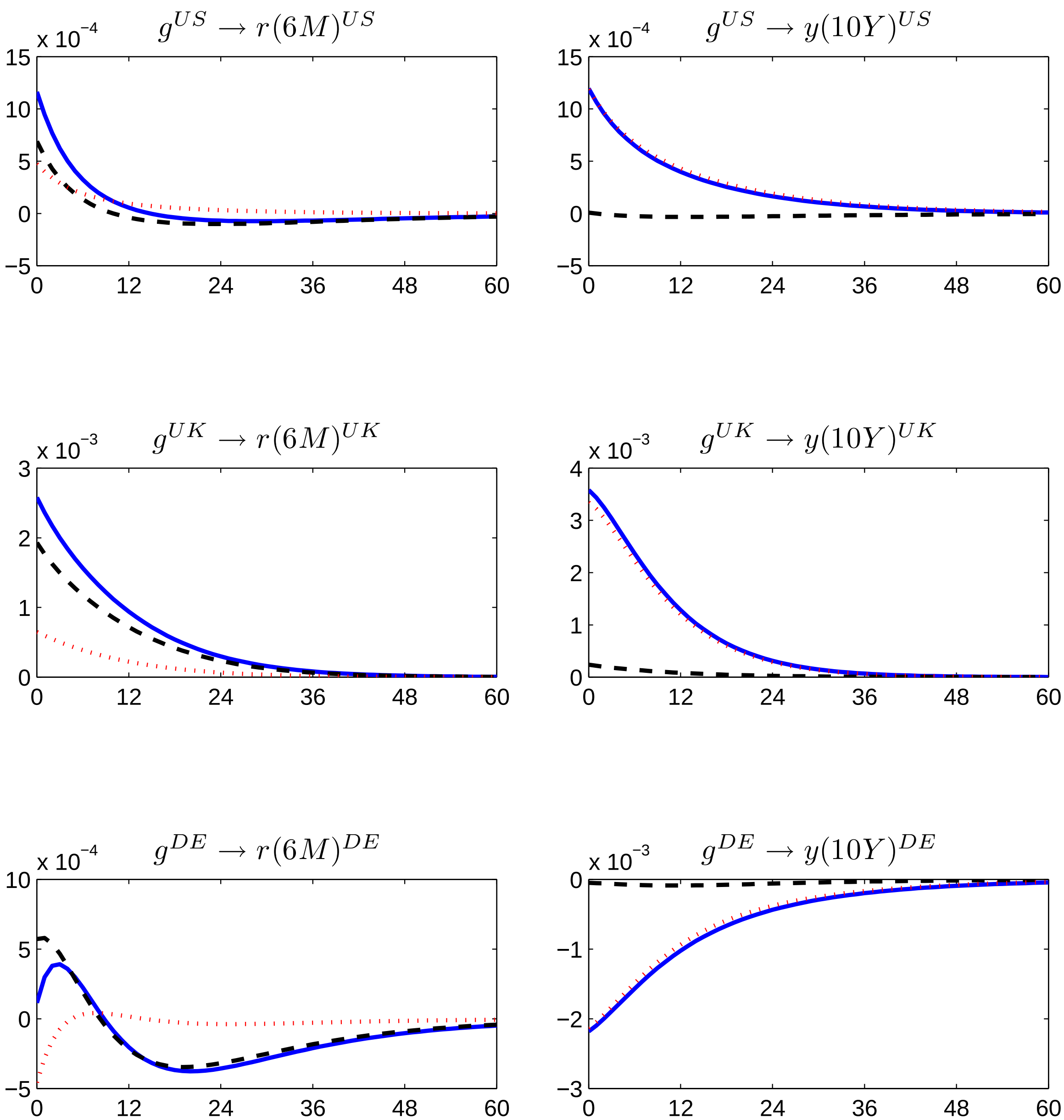

Figure 4, cont'd: Impulse Responses of Yields.

Panel C: Country-Specific Level Factor on Yields

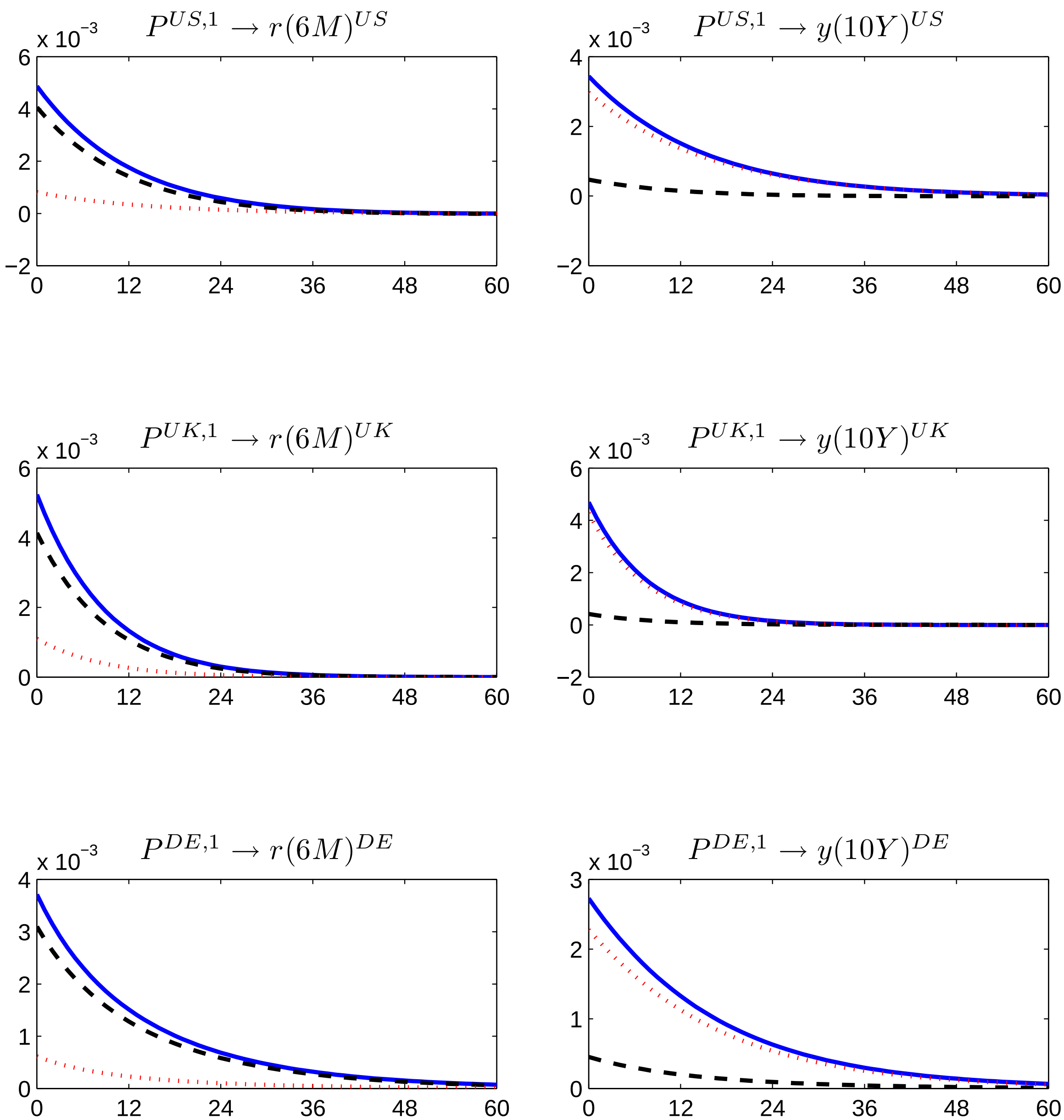

Figure 4, cont'd: Impulse Responses of Yields.