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International Yield Comovements

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Abstract

We decompose long-term nominal bond yields into real and inflation components in an international context using inflation-linked and nominal bonds. In contrast to extant results, real rate variation dominates the variation in inflation-linked and nominal yields. Cross-country nominal and inflation-linked yield correlations have declined since the Great Recession. Real rates are the main source of the correlation between nominal yields. Our results are robust to various alternative measurements of inflation expectations and the liquidity premium. They continue to hold when a no-arbitrage term structure model with real, nominal, and inflation factors is used to effect the yield decomposition.

I. Introduction

The nominal yield on a government bond can be decomposed into a real yield, expected inflation and an inflation risk premium. The decomposition is of critical economic interest because policymakers react very differently to expected inflation changes than to shifts in real yields or the inflation risk premium. However, if the market only trades nominal bonds, all three components are unobserved. The typical approach to identify these components involves estimating a term structure model, which imposes restrictions on the dynamics of state variables and risk compensation to achieve identification (see, e.g., Ang, Bekaert, and Wei (2008)). Instead, we follow recent literature which alleviates the identification problem by using survey data to (help) identify expected inflation, and inflation-linked bonds to help tie down real rates (see, e.g., D'Amico, Kim, and Wei (2018)). The older literature, which does not use inflation-linked debt, typically finds that inflation compensation (expected inflation and the inflation risk premium) accounts for most of the variation of nominal yields and nominal term spreads (see Bekaert and Wang (2010) for a survey). For example, Ang et al. (2008) find that variation in expected

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inflation and the inflation risk premium explains about 80% of the variation in nominal rates.

We reexamine this important decomposition with more recent data, ensuring that we can rely on inflation-linked debt yields. Yet, the use of inflation-linked bonds also creates various challenges. First, the time series sample is relatively short, starting only in 2004. Our main focus therefore is on comovements and variances, rather than level averages, which may be too sensitive to the time period at hand. Second, the secondary market for inflation-linked debt tends to be less liquid than that for nominal bonds, preventing the use of inflation-linked yields as direct proxies for real rates. Estimates by Gürkaynak, Sack, and Wright (2010) and D'Amico et al. (2018) show liquidity premiums on United States TIPS in the first 3–5 years after inception to vary substantially over time, dropping from over 1% annually to considerably smaller levels around 2004.[1] We estimate liquidity premiums using state-of-the-art methods. We examine the standard yield decomposition not only for the United States but also internationally and study the comovements of yields and their components across countries. We focus on France and the United Kingdom, because they have long time series on inflation-linked yields.

Our main findings are as follows. First, over the last 15 years, nominal and inflation-linked yields have decreased over time, and their standard deviations have mostly decreased. Internationally, these observed yields correlate highly but their correlation has decreased over time, in some cases turning negative. Second, expected inflation plays no role in these developments. Instead, real yields are the dominant component contributing to the variation in inflation-linked and nominal yields.[2] Liquidity premiums have only decreased in the U.S., but inflation risk premiums have decreased everywhere, and this decrease is statistically significant. Real yields are also the dominant component in explaining comovements across countries. These results are remarkably robust to various alternative measurements of expected inflation and the liquidity premium.

To preserve space, we focus all of our results on the 5-year yield, but the results hold for the 2- and 10-year maturities as well. In an effort to impose no arbitrage conditions across different yields, and to provide further robustness checks, we also set out a no-arbitrage term structure model. We largely follow the approach of Abrahams et al. (2016), who formulate a Gaussian affine term structure model with prices of risk that are linear in various pricing factors. Our factors include nominal and real factors extracted from nominal and inflation-linked yields, respectively; a liquidity factor and an inflation factor (the level of inflation). We purposedly do not use survey expectations in the model, so that inflation expectations and the inflation

[1] See also Sack and Elsasser (2004), Shen (2006), Abrahams, Adrian, Crump, Moench, and Yu (2016), and Pflueger and Viceira (2016) for similar results.

[2] This result is reminiscent but different from the result in Duffee (2018) who avoids the use of inflation-linked yields. He finds that expected inflation *news* contributes little to the variation in *shocks* to nominal bond yields. His computations (see his Table III) do suggest that these inflation variance ratios also decreased in 2008–2013 relative to earlier periods. Early signs of the increased importance of real yields can be observed from plots, for example, in Campbell, Shiller, and Viceira (2009), Abrahams et al. (2016), and D'Amico et al. (2018), but we establish it formally via variance decompositions. Compared to the earlier literature, we also have a substantially longer sample, control for the liquidity risk premium, and study an international cross-section.

risk premium are solely implied by the model. All our key results regarding the dominant role of the real rate in yield variance and comovement decompositions continue to hold.

Our contribution is twofold. First, while the literature on international asset return comovements is vast, surprisingly little research exists on yield correlations across countries. Jotikasthira, Le, and Lundblad (2015) examine correlations across nominal yields in the United States, United Kingdom, and Germany through the lens of a reduced-form term structure model with inflation and real activity as main factors. They mostly distinguish a "policy" channel (the short rate and its effect of long term yields through the expectations hypothesis) and "risk compensation" channel (term premiums). They find that nominal yields are highly correlated across countries, with both channels explaining roughly equal parts of the total variation for 5-year yields. In a contemporaneous paper, Berardi and Plazzi (2022) estimate a reduced-form term structure model to compute yield correlations across 4 major economies, focusing, similarly to Jotikasthira et al. (2015), on short rate expectations and term premiums. We extend these papers by decomposing the cross-country yield correlations into real yield, expected inflation and inflation risk premium components for nominal bonds and into real yield and liquidity premium components for inflation-linked bonds.

Our second contribution is to establish and economically interpret a set of stylized facts regarding yield decompositions. The extant literature performing such decompositions using inflation-linked yields either ignores the liquidity premium[3] or focuses on an individual time series, such as real yields in Campbell et al. (2009), arbitrage profits in Fleckenstein (2013) and Fleckenstein, Longstaff, and Lustig (2014), the inflation risk premium in Grishchenko and Huang (2013), expectation hypothesis violations in Pflueger and Viceira (2016), expected inflation in Kaminska, Zhuoshi, Relleen, and Vangelista (2018), the liquidity premium in D'Amico et al. (2018), or the issuance costs of inflation-linked versus nominal bonds in Ermolov (2021). Instead, we provide a comprehensive analysis on the relative importance of different yield components in multiple countries.

II. Decomposing Yields: On Liquidity and Inflation Risk Premiums

Throughout this article, we work with continuously compounded yields on zero-coupon government bonds. The main decomposition of interest is:

$$(1)\qquad \underbrace{y_t^n}_{\text{NOMINAL\_RATE}} = \underbrace{r_t^n}_{\text{REAL\_RATE}} + \underbrace{E_t\left[\pi_{t,t+n}^n\right]}_{\text{EXPECTED\_INFLATION}} + \underbrace{\varphi_t^n}_{\text{INFLATION\_RISK\_PREMIUM}},$$

[3]Among others, Christensen, Lopez, and Rudebusch (2010) and Chen, Engstrom, and Grishchenko (2016) for the United States, Evans (1998) (2003), Risa (2001), and Joyce, Lildhodlt, and Sorensen (2010), for the United Kingdom, and Garcia and Werner (2010), Hördahl and Tristani (2014), and Pericoli (2014), for the euro area. Haubrich, Pennacchi, and Richken (2012) use inflation swap rates instead of TIPS to estimate the various components assuming perfect liquidity in inflation swap markets.

where y_t^n is the yield on a nominal zero-coupon bond of maturity n, r_t^n is the yield on a perfectly indexed zero coupon bond of maturity n, and $\pi_{t,t+n}^n$ is (average) inflation from t to $t+n$. The difference between y_t^n and r_t^n is often called "inflation compensation" or sometimes "breakeven inflation rate." It consists of expected inflation and the inflation risk premium, the compensation investors demand to protect themselves against inflation risk.

The Fisher hypothesis holds that the inflation risk premium is zero, but is inconsistent with both modern asset pricing theory and recent empirical estimates of the inflation risk premium. In typical asset pricing models, the inflation risk premium depends on the covariance between the real pricing kernel and inflation. That is, the inflation risk premium is positive if inflation is high in "bad times," as the pricing kernel takes on high values in bad states of the world. Of course, this covariance between the wealth or consumption of agents and inflation may well vary through time inducing substantial variation in the conditional inflation risk premium. The premium may even be negative when inflation and stock returns (as an indicator of "wealth") are positively correlated (Campbell, Sunderam, and Viceira (2017)) or in an aggregate demand environment, where real activity is negatively correlated with inflation (Bekaert, Engstrom, and Ermolov (2021)).

Liquidity premiums in inflation-linked debt considerably complicate the identification problem embedded in equation (1), because inflation-linked yields do not deliver r_t^n. Let $r_t^{n,\mathrm{IL}}$ denote the zero-coupon yield derived from inflation linked debt. It consist of two components:

$$r_t^{n,\mathrm{IL}} = r_t^n + \mathrm{LIQ\_PR}_t^n, \tag{2}$$

where LIQ_PR represents a liquidity premium that may vary through time.

We partially resolve the identification problem by measuring inflation expectations from survey forecasts. Such forecasts, by either professionals or consumers, are now available for multiple countries. In fact, Ang, Bekaert, and Wei (2007) find that survey forecasts (in particular, the Survey of Professional Forecasters (SPF)) consistently beat other models in forecasting U.S. inflation out-of-sample. Assuming inflation expectations are observed, data on inflation-linked and nominal yields generate direct information on an interesting concept, which we dub the nominal debt premium. That is, $\mathrm{NDPR}_t^n = y_t^n - r_t^{n,\mathrm{IL}} - E_t\left[\pi_{t+n,n}^n\right]$. From equations (1) and (2), it follows that $\mathrm{NDPR}_t^n = \varphi_t^n - \mathrm{LIQ\_PR}_t^n$. The nominal debt premium is the difference between the inflation risk premium, priced in nominal bonds, and the liquidity premium, priced in inflation-linked debt. It represents the real cost advantage or disadvantage of the government issuing inflation-linked versus nominal debt. Full identification is then achieved by empirically estimating the liquidity premium.

III. Data and Initial Stylized Facts

Our yield data comprise end-of-month zero-coupon yields extracted from nominal and inflation-linked bonds from France, the United Kingdom, and the United States. The sample starts in 2004 because before this date an insufficient number of bonds are available to create yield curves, especially in France.

Zero-coupon yields for the U.S., both for nominal Treasuries and Treasury inflation-protected securities (TIPS), are from Gürkaynak, Sack, and Wright (2007) and (2010), respectively). The U.K. zero-coupon nominal and inflation-linked yields are from the Bank of England website.

For France, the nominal zero-coupon yields are from the Banque de France website. We use the Nelson and Siegel (1987) methodology to construct French zero-coupon yields from inflation-linked bond prices taken from Bloomberg. Under Nelson and Siegel (1987) parameterization, the time t n-period zero-coupon yield is

$$r_t^{n,\mathrm{IL}} = \beta_0 + \beta_1 \frac{1 - e^{-\frac{n}{\tau_1}}}{\frac{n}{\tau_1}} + \beta_2 \left(\frac{1 - e^{-\frac{n}{\tau_1}}}{\frac{n}{\tau_1}} - e^{-\frac{n}{\tau_1}} \right),$$

where $\beta_0, \beta_1, \beta_2$, and τ_1 are model parameters. We estimate end-of-month Nelson–Siegel parameters by minimizing the sum of squared deviations between observed and predicted bond prices weighted by the inverse of bond duration.[4] This procedure results in essentially the same yield curve as minimizing the sum of squared yield deviations, but is computationally much faster (Gürkaynak et al. (2007)). No bonds with residual maturity below 12 months are used, because their prices are strongly affected by indexation lags and seasonality effects. Our results are robust to excluding bonds with maturities less than 18 and 24 months. We only use the bonds linked to the euro zone harmonized index of consumer prices excluding tobacco as they are more common than bonds linked to domestic French inflation.

Because inflation-linked debt tends to be issued at relatively long maturities, the main security we focus on is the 5-year zero-coupon bond. Our results for longer maturities, discussed in Section IV.G (and also in Section V), are similar. As our analysis is based mainly on off-the-run long maturity bonds, we assume that the deflation protection and indexation lag premia embedded in inflation-linked bond prices are zero (see Risa (2001) and D'Amico et al. (2018)). We provide some institutional background regarding the three markets primarily analyzed in this article in Supplementary Appendix I.

In Table 1, we show the properties of the 5-year yields, with nominal yields on the left and inflation-linked yields on the right and three panels for the full sample, the first half of the sample (2004 till 2012) and second half of the sample (2012–2019).[5] We show standard errors from a GMM procedure, outlined in

[4]While the French nominal yield curves are constructed using the Svensson (1994) extension of the Nelson and Siegel (1987) methodology, we construct French inflation-linked yield curves using the original Nelson and Siegel (1987) methodology. The Svensson-methodology requires the availability of a sufficient number of long-term bonds to be reliably applied, which the French inflation-linked market lacks until the second half of our sample. For the last part of the sample, we confirmed that the differences between the Nelson–Siegel and Nelson–Siegel–Svensson methodologies are economically small at long maturities and essentially nonexistent at medium maturities, such as the 5-year maturity we study.

[5]We conduct the Bai, Lumsdaine, and Stock (1998) formal break tests on the 6 yields: three nominal and three inflation-linked yields. The test assumes a VAR with the lag order chosen by the BIC criterion (which is 1 in our case), formulates a test for one unknown break point, and provides a confidence interval on the break date. We find a significant break, with the null hypothesis of no break rejected at the 1% level. The break date is Feb. 2011, but the 95% confidence interval is ±29 months, which includes the mid-point of the sample. To preserve balanced samples, we chose the midpoint as the breakpoint.

TABLE 1

Annualized 5-Year Observed Zero-Coupon Bond Yields

Table 1 shows results obtained using monthly data. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	Nominal Yields			Inflation-Linked Yields		
	France	U.K.	U.S.	France	U.K.	U.S.
Full sample: Nov. 2004–Dec. 2019						
Average	1.61% (0.59%)	2.23% (0.56%)	2.32% (0.38%)	0.11% (0.43%)	−0.43% (0.60%)	0.48% (0.38%)
Std. Dev.	1.56% (0.12%)	1.58% (0.16%)	1.19% (0.14%)	1.12% (0.12%)	1.68% (0.14%)	1.10% (0.12%)
β wrt U.S.	0.89 (0.18)	1.13 (0.10)	1.00	0.68 (0.17)	1.19 (0.19)	1.00
Correlation with U.S.	0.67 (0.14)	0.85 (0.07)	1.00	0.66 (0.17)	0.78 (0.12)	1.00
Correlation with U.K.	0.93 (0.08)	1.00	0.85 (0.07)	0.94 (0.08)	1.00	0.78 (0.12)
Subsample 1: 2004M11–2012M5						
Average	2.99% (0.47%)	3.47% (0.72%)	2.95% (0.66%)	1.08% (0.36%)	0.93% (0.67%)	1.08% (0.57%)
Std. Dev.	0.86% (0.11%)	1.32% (0.18%)	1.30% (0.20%)	0.67% (0.09%)	1.24% (0.15%)	1.15% (0.15%)
β wrt U.S.	0.51 (0.09)	0.95 (0.09)	1.00	0.46 (0.05)	1.01 (0.05)	1.00
Correlation with U.S.	0.76 (0.14)	0.93 (0.09)	1.00	0.79 (0.09)	0.93 (0.05)	1.00
Correlation with U.K.	0.89 (0.12)	1.00	0.93 (0.09)	0.83 (0.14)	1.00	0.93 (0.05)
Subsample 2: 2012M6–2019M12						
Average	0.23%*** (0.18%)	0.99%*** (0.19%)	1.69%** (0.24%)	−0.86%*** (0.19%)	−1.80%*** (0.29%)	−0.12%*** (0.22%)
Std. Dev.	0.53% (0.11%)	0.46%** (0.12%)	0.58%** (0.17%)	0.42% (0.08%)	0.60%** (0.10%)	0.63%** (0.12%)
β wrt U.S.	−0.35*** (0.19)	0.06*** (0.07)	1.00	−0.33*** (0.08)	−0.24*** (0.13)	1.00
Correlation with U.S.	−0.38*** (0.21)	0.08*** (0.08)	1.00	−0.50*** (0.13)	−0.25*** (0.13)	1.00
Correlation with U.K.	0.61* (0.11)	1.00	0.08*** (0.08)	0.76 (0.11)	1.00	−0.25*** (0.13)

Supplementary Appendix II, incorporating 12 Newey–West (1987) lags. Asterisks in the second subsample indicate values statistically different from the first subsample. It should not be any surprise that yields have significantly decreased with inflation-linked yields becoming negative in the second subsample. The standard deviations of yields have decreased as well.

For the full sample, we confirm the result in Jotikasthira et al. (2015) that nominal yields are highly correlated across countries, with the correlation varying between 0.67 for the U.S. and France and 0.93 for France and the U.K. The inflation-linked yield correlations are of the same order of magnitude. When looking at the subsamples, however, we see that these correlations have decreased substantially in a statistically significant fashion, except for the correlation of French with U.K. yields. This is not due to a volatility effect, which we can infer from the statistics for U.S. betas. These betas represent the exposure of French and U.K. yields to U.S. yields (as implied by a linear regression with a constant).

The U.S. beta for French nominal (inflation-linked) yields has also decreased from 0.51 (0.46) to −0.35 (−0.33). For the U.K., betas of around 1 turn negative (inflation-linked yields) or become virtually zero (nominal yields) from the first to the second subsample.

Inflation-linked bonds may result in debt cost savings for the government, when the inflation risk premium is larger than the liquidity premium priced in inflation index-linked bonds. We now provide direct estimates of the relative interest rate cost of issuing nominal versus inflation-linked debt, by measuring inflation expectations. We take 5-year expected inflation from the SPF for the U.S., from the European Central Bank (ECB) SPF for France, and from Consensus Economics for the United Kingdom. For the U.S., SPF data is only available from 2005:Q3 onward, so we use Aruoba (2020) estimates, who aggregates data from multiple surveys, for 2004:Q4–2005:Q2. While the forecasts in Aruoba (2020) are "spot" forecasts and available every month, SPF and ECB SPF forecasts are available quarterly, and the Consensus Economics forecasts are only available twice a year. We assume that the forecasts do not change in between data releases. A robustness check, where we linearly interpolate between the forecasts, does not change any of our key findings. The ECB SPF provides 1-year-forward inflation forecasts, between year 0 and 1; between year 1 and 2 and between year 4 and 5. To distill an estimate for expected inflation over the next 5 years, we set expected inflation 2 and 3 years from now, equal to expected inflation 4 years from now. An alternative assumption where we linearly interpolate between the values for expected inflation 1 and 4 year(s) ahead, produces very similar empirical results. Figure 1 graphs these expectations, clearly showing inflation expectations to be higher in the U.K., followed by the U.S. and then France, where inflation expectations seem to vary the least. Section IV.F.2 considers further robustness checks to alternative data sources on inflation expectations.

In Table 2, we report the statistical results on expected 5-year inflation, as measured by the surveys (left-hand side). Average inflation expectations range from 1.80% in France, over 2.26% in the U.S. to 2.97% in the U.K. Therefore, the expectations are very near the inflation targets set by the European and

FIGURE 1

Annualized 5-Year Survey Expected Inflation

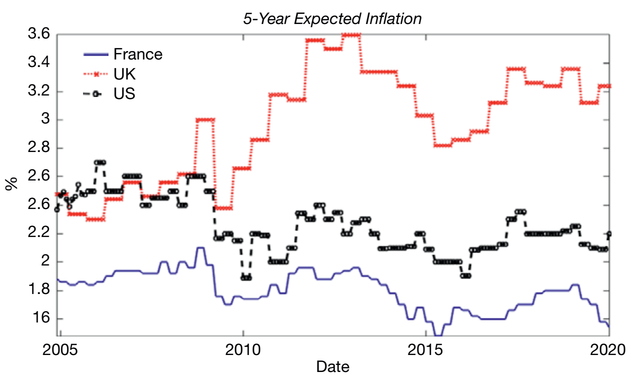


TABLE 2

Annualized 5-Year Survey Expected Inflation and Nominal Debt Premium

Table 2 shows results obtained using monthly data. The nominal debt premium is defined as the difference between nominal yields and the sum of expected inflation and inflation-linked yields. GMM standard errors computed using 12 Newey–West (1987) lags are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	Expected Inflation			Nominal Debt Premium		
	France	U.K.	U.S.	France	U.K.	U.S.
Full sample: 2004M11–2019M12						
Average	1.80%	2.97%	2.26%	−0.30%	−0.30%	−0.42%
	(0.05%)	(0.12%)	(0.06%)	(0.13%)	(0.14%)	(0.13%)
Std. Dev.	0.14%	0.39%	0.19%	0.47%	0.60%	0.61%
	(0.10%)	(0.12%)	(0.10%)	(0.09%)	(0.07%)	(0.08%)
β wrt U.S.	0.50	−0.81	1.00	0.36	0.63	1.00
	(0.10)	(0.43)		(0.10)	(0.09)	
Correlation with U.S.	0.70	−0.41	1.00	0.47	0.64	1.00
	(0.14)	(0.21)		(0.13)	(0.09)	
Correlation with U.K.	−0.11	1.00	−0.41	0.60	1.00	0.64
	(0.22)		(0.21)	(0.15)		(0.09)
Subsample 1: 2004M11–2012M5						
Average	1.88%	2.73%	2.36%	0.03%	−0.20%	−0.49%
	(0.04%)	(0.18%)	(0.10%)	(0.12%)	(0.26%)	(0.26%)
Std. Dev.	0.09%	0.38%	0.20%	0.41%	0.76%	0.82%
	(0.19%)	(0.22%)	(0.14%)	(0.11%)	(0.09%)	(0.12%)
β wrt U.S.	0.29	−0.70	1.00	0.41	0.74	1.00
	(0.10)	(0.43)		(0.02)	(0.06)	
Correlation with U.S.	0.63	−0.37	1.00	0.82	0.79	1.00
	(0.22)	(0.23)		(0.04)	(0.06)	
Correlation with U.K.	0.16	1.00	−0.37	0.91	1.00	0.79
	(0.15)		(0.23)	(0.04)		(0.06)
Subsample 2: 2012M6–2019M12						
Average	1.72%***	3.20%***	2.15%**	−0.64%***	−0.41%	−0.35%
	(0.04%)	(0.09%)	(0.04%)	(0.06%)	(0.09%)	(0.06%)
Std. Dev.	0.13%	0.21%*	0.11%*	0.24%	0.36%**	0.26%
	(0.12%)	(0.17%)	(0.12%)	(0.06%)	(0.09%)	(0.06%)
β wrt U.S.	0.72*	1.58***	1.00	0.70***	−0.09***	1.00
	(0.22)	(0.19)		(0.07)	(0.24)	
Correlation with U.S.	0.61	0.78***	1.00	0.75	−0.06***	1.00
	(0.18)	(0.09)		(0.07)	(0.17)	
Correlation with U.K.	0.78***	1.00	0.78***	−0.08***	1.00	−0.06***
	(0.14)		(0.09)	(0.21)		(0.17)

U.S. central banks (which are at 2%) but quite a bit above the target set by the Bank of England. There is a statistically significant decrease in the average expected inflation of about 20 basis points over the two sample periods for France and the U.S. In the U.K., inflation expectations have actually increased to 3.2% in the second sample half from 2.7% in the first, and the change is statistically significant. Inflation expectations are now very stable, as witnessed by the low standard deviations, especially in France and the U.S., where the standard deviation is less than 20 basis points. D'Amico et al. (2018) report standard deviations for expected inflation over various horizons invariably exceeding 1%. The low volatility of the inflation expectations may partially reflect the long horizons over which expectations are formed but may also derive from monetary policy's ability to anchor inflation expectations. In Supplementary Appendix III.A, we repeat the information in Table 2 but for 1-year-ahead inflation expectations. These expectations are more variable, exhibiting standard deviations in the 35–40 basis points range. While inflation is far from dead, as was sometimes claimed in media reports preceding the recent pandemic, its stochastic process has definitely changed.

Perhaps surprisingly, inflation expectations are not highly correlated across countries. The correlation between expected inflation in the U.K. and the U.S. is negative, which is also reflected in a very large but negative U.K. beta with respect to U.S. inflation. Expected inflation in the U.S. and France are more highly correlated at 0.7. The table does reveal that the low correlations are all due to the first part of the sample. More recently, expected inflation in France and the U.K. show a 0.78 correlation whereas U.S. expected inflation correlates 0.78 (0.61) with expected inflation in the U.K. (France). Generally, these correlation coefficients are of little economic importance, because, as can be seen from Figure 1 and Table 2, the standard deviations of long-term inflation forecasts are very low, a pattern we confirm for statistical long-term inflation expectations in Section IV.F.2. Supplementary Appendix III.A documents that 1-year-ahead inflation expectations are more highly but still modestly correlated. These results are consistent with alternative estimates in the literature. Monacelli and Sala (2009), for example, find that for inflation rates in the U.S., U.K., France, and Germany an international common factor explains less than 30% of the inflation variance. Förster and Tillman (2014) and Parker (2018) show that inflation correlations have decreased substantially in the 21st century. Kearns (2016), in contrast, suggests that inflation expectations are highly correlated but his results regard 1-year-ahead inflation expectations and a sample starting in 1992.

The right panel of Table 2 shows properties of the nominal debt premium. The nominal debt premium has been on average negative in all 3 countries, ranging from −30 basis points in France to −42 basis points in the U.S. In the two European countries nominal debt premiums decrease substantially in a statistically significant manner in the second subsample. It appears that governments have enjoyed a substantive yield advantage issuing nominal debt. While it is tempting to associate this finding with the unusual monetary policies in Europe and elsewhere, note that monetary policy should primarily affect real yields. However, low inflation risk premiums relative to liquidity premiums, may also arise from the ability of nominal bonds to hedge real risk in aggregate demand environments, and correlate negatively with equity returns in periods of market stress. From that perspective, it is surprising that the nominal debt premiums are lower in the second part of the sample, as the first half of the sample was dominated by the Great Recession, in which bond and stock returns were mostly negatively correlated and which is mostly characterized as an aggregate demand recession (see Mian and Sufi, 2010). Nominal debt premiums have become less correlated across countries, with the decrease significant for the pairs involving the U.K.

IV. Empirical Decomposition Results

To implement the decomposition in equation (1), we now estimate the liquidity premium and then show properties of the resulting liquidity and inflation risk premiums. With all the components in hand, we provide variance decompositions of the three yield components in the three countries and a decomposition of the correlation dynamics across countries of nominal and inflation-linked yields.

A. The Liquidity Premium

To estimate the liquidity premium, we follow Gürkaynak et al. (2010) and run the following regression:

$$\text{NDPR}^{n}_{t,i} = c_{1,i} + c_2' l_{t,i} + \varepsilon_{t,i}, \tag{3}$$

where $\text{NDPR}^{n}_{t,i}$ is the nominal debt premium in country i, $l_{t,i}$ is a vector of country-specific liquidity proxies and $\varepsilon_{t,i}$ is the error term. Recall that the debt premium equals the inflation risk premium minus the liquidity premium. Therefore, if the liquidity proxies indicate illiquidity we expect the coefficients to be negative. In addition, for the procedure to correctly identify the liquidity premium, the liquidity proxies should be uncorrelated with the inflation risk premium. Given the slope coefficients from equation (3), $\widehat{c}_2$, the liquidity premium in country i at time t can be computed as $-\widehat{c}_2' l_{t,i}$. Obviously, the mean of the liquidity premium is not identified through this procedure.

We use three types of liquidity proxies: the nominal off-the-run spread, the relative outstanding amount of inflation-linked bonds, and the inflation swap spread. The nominal off-the-run spread is the difference between yields of the most recent and older nominal bonds of the same maturity offering almost identical cash flows (see Krishnamurthy (2002) for the U.S. and Geyer, Kossmeier, and Pichler (2004) internationally). Following Pflueger and Viceira (2016), we construct the off-the-run spread by estimating a nominal yield for a particular maturity using the cross-section of all bonds (most of which are off-the-run) and subtracting the on-the-run yield from Bloomberg. Pflueger and Viceira (2016), among others, propose this spread as an indicator of the overall demand for liquidity (higher spreads indicating stronger demand and higher liquidity premiums). Although the off-the-run spread is not directly linked to the liquidity of inflation-linked bonds, a voluminous literature, starting with Chordia, Sarkar, and Subrahmanyam (2005), suggests that there is strong commonality in liquidity between different markets.

The relative outstanding amount of inflation-linked bonds is the nominal value of outstanding inflation-linked bonds divided by the nominal value of outstanding inflation-linked and nominal bonds. This variable reflects the general market development of the inflation-linked market, as more debt outstanding likely implies a more comprehensive term structure of inflation-linked debt, more regular issue calendars and so forth The outstanding amount may also be correlated with trading volumes, a variable we were unable to track down for the French market. We obtain the data on nominal outstanding amounts of nominal and inflation-linked debt from the Agence France Trésor, for France, from the United Kingdom Debt Management Office for the U.K., and from the Bank of International Settlements for the U.S.

The inflation-swap spread is defined as the rate on a zero-coupon inflation swap position paying fixed and receiving floating minus the difference between the zero-coupon nominal and inflation-linked yields. The inflation swap quote is the risk neutral expectation of future inflation and thus essentially represents inflation compensation (including expected inflation and the inflation risk premium). The nominal yield which is subtracted, also incorporates inflation compensation. In the absence of market frictions and liquidity differences, this spread should therefore be

0, as the inflation-linked yield would simply correct for the real yield. The inflation swap spread should therefore reflect the liquidity premium embedded in inflation-linked yields. Fleckenstein et al. (2014) show that the spread varies substantially over time. Because the spread is in principle arbitrageable, they also link it to the strength of arbitrage activity in debt markets.[6] Following Pflueger and Viceira (2016), we use the end-of-month 5-year inflation swap spread, with the inflation swap rate taken from Bloomberg.

Finally, we include a coarse measure of market development, the log of the number of months since inception. It is typically the case that early inflation-linked programs are associated with poor liquidity, uncertainty about the viability of the market, incomplete availability of bonds along the maturity spectrum and irregular issuance calendars (e.g., Campbell et al. (2009)). From that perspective the U.K. has a much longer experience with inflation-linked debt than the other countries. However, because the variable only depends on time, it may also reflect trending behavior in the inflation risk premium. For example, Bekaert and Wang (2010) survey a large number of empirical studies on the inflation risk premium and show that studies with shorter, more recent samples tend to find smaller average inflation risk premiums.

We run a panel regression with monthly data including all three countries, imposing c_2 to be the same across countries. The results are very similar using quarterly observations. Table 3 first shows univariate results for each independent variable, with and without country fixed effects, before showing the full specification with all 4 independent variables and country fixed effects.

In line with economic intuition, the off-the-run spread and inflation swap spreads have statistically significant negative coefficients, while the log share of total government debt accounted for by inflation-linked debt has a significantly positive coefficient. Months since inception has a negative coefficient which is only significant when country fixed effects are introduced. Thus, the nominal debt premium results suggest an upward trend in the liquidity premium that is stronger for the countries with more recent inflation-linked debt programs, but perhaps an economically more plausible explanation is that the negative coefficient captures the decrease in the inflation risk premium documented by several recent articles (among others, Chen et al. (2016), Song (2017), and D'Amico et al. (2018)). In specification 9, we use all 4 variables and country fixed effects. First, note that the adjusted R^2 is 56% suggesting the fit is good. The off-the-run spread is the only variable that is not significant. We therefore use specification 10 in our estimates of the liquidity premium, which only features the inflation swap spread, the log(share of inflation-linked debt) and log(months since exception) and actually has a higher adjusted R^2 than specification 9. Given that the log(months since inception of the inflation-linked bond program) variable is statistically significant, we include it as a regressor in the nominal debt premium regression in order to avoid an omitted variable bias on our estimated coefficients. However, we do not include it to compute the liquidity premium given that it likely captures inflation risk premium

[6]To and Tran (2019) argue that inflation swaps may exhibit overpricing, which would bias our liquidity premium estimates upward; Fleckenstein (2013), describing the G7 inflation swap markets, argues that the 5–10-year tenors are the most liquid.

TABLE 3

Inflation-Linked Bonds Liquidity Premia Estimation

Table 3 shows results obtained using the monthly data over the period of 2004M11 to 2019M12. The panel regression is $y_{t,i} - y^{IL}_{t,i} - \pi^{e}_{t,i} = c_{1,i} + c_2' l_{t,i} + \varepsilon_{t,i}$, where $y_{t,i}$ is zero-coupon yield in country i at time t, π^{e}_{t} is expected inflation, and l_t is the vector of liquidity proxies, which are assumed to be uncorrelated with the inflation risk premium, and $\varepsilon_{t,i}$ is the error term. Regressions are for 5-year zero-coupon yields. Driscoll and Kraay (1998) standard errors computed with 12 lags are in parentheses. *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	Spec 1	Spec 2	Spec 3	Spec 4	Spec 5	Spec 6	Spec 7	Spec 8	Spec 9	Spec 10
OFF-THE-RUN_PREMIUM	−2.91***	−2.90***							−0.33	
	(0.97)	(0.93)							(0.47)	
INFLATION_SWAP_SPREAD			−1.50***	−1.74***					−1.77***	−1.80***
			(0.34)	(0.27)					(0.31)	(0.33)
log(SHARE_OF_INFLATION-LINKED_DEBT)					0.24*	1.01**			0.51**	0.55**
					(0.14)	(0.44)			(0.23)	(0.23)
log(MONTHS_SINCE_INCEPTION)							−0.19	−0.66***	−0.63***	−0.63***
							(0.13)	(0.20)	(0.10)	(0.10)
Country-fixed effects	No	Yes	No	Yes	No	Yes	No	Yes	Yes	Yes
Adj.R^2	9.40%	9.88%	37.73%	43.85%	2.00%	5.65%	2.01%	10.52%	58.08%	58.14%

dynamics. That is, the liquidity premium is only computed using the inflation swap spread and log(share of inflation-linked debt)-variables from that regression. While we think this approach represents the most sensible choice economically, given that the interpretation of the log(months since inception of the inflation-linked bond program)-variable is somewhat ambiguous, we conduct two robustness checks. First, we do not include the log(months since inception of the inflation-linked bond program)-variable into the nominal debt premium regression and do not include it into the liquidity premium estimation. Second, we include the log(months since inception of the inflation-linked bond program)-variable into the nominal debt premium regression and also use it to compute the liquidity premium along the inflation swap spread and log(share of inflation-linked debt)-variables. Our main results hold in both cases and are presented in Supplementary Appendix IV.A. Graph A of Figure 2 shows the three estimated liquidity premiums over time. Further robustness checks are described in Section IV.F.1.

B. Liquidity and Inflation Risk Premiums

Given the presence of fixed effects and the use of the nominal debt premium on the left hand side, it is impossible to tie down the level of the liquidity premium. This is immaterial for most of our results, which focus on volatility, variance decompositions, and comovements. However, to graph the liquidity and inflation risk premiums, and report averages, we follow Gürkaynak et al. (2010), normalizing the level of the liquidity premium to zero at the point in time at which it was the lowest. Once the liquidity premium is identified, we identify real rates from inflation-linked yields using equation (2), and the inflation risk premium is simply

$$\varphi_t^n = y_t^n - E_t\left[\pi_{t+n,n}^n\right] - r_t^n. \tag{4}$$

In Table 4, we produce characteristics of liquidity (left panel) and inflation risk premiums (right panel). Liquidity premiums are around 50 basis points on average

FIGURE 2

Annualized 5-Year Zero-Coupon Liquidity and Inflation Risk Premia

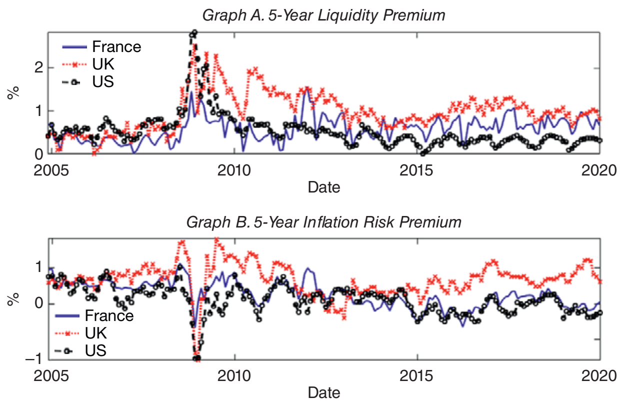


TABLE 4

Annualized 5-Year Zero-Coupon Liquidity and Inflation Risk Premia

Table 4 shows results obtained using the monthly data. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	Liquidity Premium			Inflation Risk Premium		
	France	U.K.	U.S.	France	U.K.	U.S.
Full sample: 2004M11–2019M12						
Average	0.58%	0.97%	0.52%	0.27%	0.67%	0.10%
	(0.05%)	(0.12%)	(0.12%)	(0.10%)	(0.10%)	(0.07%)
Std. Dev.	0.29%	0.45%	0.41%	0.36%	0.45%	0.39%
	(0.06%)	(0.11%)	(0.09%)	(0.07%)	(0.09%)	(0.07%)
β wrt U.S.	0.17	0.53	1.00	0.60	0.42	1.00
	(0.07)	(0.10)		(0.11)	(0.25)	
Correlation with U.S.	0.24	0.48	1.00	0.66	0.36	1.00
	(0.10)	(0.09)		(0.12)	(0.22)	
Correlation with U.K.	0.47	1.00	0.48	0.38	1.00	0.36
	(0.16)		(0.09)	(0.16)		(0.22)
Subsample 1: 2004M11–2012M5						
Average	0.50%	0.99%	0.74%	0.52%	0.79%	0.25%
	(0.10%)	(0.27%)	(0.18%)	(0.04%)	(0.10%)	(0.09%)
Std. Dev.	0.34%	0.61%	0.47%	0.28%	0.52%	0.44%
	(0.09%)	(0.10%)	(0.12%)	(0.09%)	(0.11%)	(0.11%)
β wrt U.S.	0.39	0.71	1.00	0.37	0.62	1.00
	(0.05)	(0.16)		(0.04)	(0.25)	
Correlation with U.S.	0.54	0.55	1.00	0.59	0.53	1.00
	(0.07)	(0.13)		(0.07)	(0.21)	
Correlation with U.K.	0.54	1.00	0.55	0.46	1.00	0.53
	(0.13)		(0.13)	(0.21)		(0.21)
Subsample 2: 2012M6–2019M12						
Average	0.65%	0.95%	0.31%***	0.02%***	0.55%*	−0.04%**
	(0.03%)	(0.07%)	(0.02%)	(0.08%)	(0.12%)	(0.07%)
Std. Dev.	0.20%***	0.19%***	0.13%	0.23%	0.34%	0.26%*
	(0.03%)	(0.08%)	(0.08%)	(0.08%)	(0.08%)	(0.09%)
β wrt U.S.	0.34	0.55	1.00	0.56	−0.44***	1.00
	(0.16)	(0.14)		(0.10)	(0.30)	
Correlation with U.S.	0.22**	0.36	1.00	0.62	−0.33***	1.00
	(0.11)	(0.09)		(0.11)	(0.23)	
Correlation with U.K.	0.36	1.00	0.36	−0.04*	1.00	−0.33***
	(0.06)		(0.09)	(0.27)		(0.23)

in France and the U.S., but around 1% in the U.K. In the U.S., the liquidity premium decreases significantly from 74 to 31 basis points, which is not surprising given that liquidity premiums were high during the Great Recession. However, in the U.K., the liquidity premium is stable and in France, it increases slightly. This may be related to the continuing poor economic conditions in Europe.[7] Liquidity premiums vary substantially through time, especially in the U.S. and the U.K. where the standard deviation is 40–45 basis points. In France, it is only 29 basis points. In the U.K. and the U.S., the standard deviation of the liquidity premiums has

[7]Pflueger and Viceira (2016) use a different normalization for the liquidity premium, but their estimates are 82% correlated with ours over the overlapping sample. Our U.K. liquidity premium estimates are not comparable to Pflueger and Viceira (2016). First, they only study bond liquidity for 20-year maturity bonds, while we focus on maturities of 10 years and below. Second, they do not use the inflation swap spread as a liquidity proxy for the U.K., while we find it to be the most statistically and economically significant liquidity proxy.

decreased to French levels (U.K.) or even lower (the U.S.) in the second part of the sample. Liquidity premiums comove positively across countries, with correlations in the 0.25–0.50 range. Such correlation is not surprising: Panyanukul (2010) shows a strong commonality of liquidity risk in international government bond markets, finding the U.S. market to be an important source of global liquidity risk. These correlations have decreased over time, particularly between France and the other countries. This decrease is statistically significant for the U.S.–France pair.

Inflation risk premiums are on average quite small in France (27 basis points) and the U.S. (10 basis points), but larger in the U.K. (67 basis points). They have decreased substantially over time, in a statistically significant manner, in all 3 countries and are now negative in the U.S. and France. This confirms the claim in Chen et al. (2016) that the recent decrease in inflation compensation in the U.S. can be attributed to lower inflation risk premiums. Their volatility is about 35–45 basis points, which has decreased over time to 25–35 basis points. Inflation risk premiums are positively correlated across countries (e.g., the correlation between U.S. and French inflation risk premiums is around 60%). However, U.K. inflation risk premiums of late have decoupled from the ones in the U.S. and France, with correlations dropping to −0.04 for the U.K. and France, and −0.33 for the U.K. and the U.S.

We graph the inflation risk premiums in Graph B of Figure 2. At the height of the Great Recession (roughly from the end of 2008 to the first quarter of 2009), there is a downward spike in the inflation risk premium, which coincides with an upward spike in the liquidity premium (see Graph A of Figure 2). While highly positive liquidity and low inflation risk premiums are not surprising during such crisis times, the particular sharp decrease of the inflation risk premium may have been partially affected by dislocations in the TIPS market (e.g., Huebscher (2009)). The unwinding of Lehman's portfolio after its bankruptcy caused a sharp increase in TIPS yields, which, in turn, contributed to the sharp decrease in the inflation risk premium.

C. Real Rates

Given liquidity premium estimates, we can now identify real rates. In Table 5, we report the statistical properties of real yields. Real rates have been unusually low during this sample period, being negative on average for the full sample (note again that unconditional levels are not exactly identified, given that the level of the liquidity premium is not pinned down). Real rates dropped steeply over the sample period, being on average in the −6 to 59 basis points range for the first sample period, while being robustly negative in the second (varying between −42 basis points for the U.S. to −2.76% for the U.K.). The decreases are statistically significant for France and the U.K.

The volatility of real rates decreases sharply from the first to the second sample for all 3 countries with the decrease steepest for the U.K. (by almost 1%). Overall, real rates are highly correlated across countries (correlations between 0.54 for the U.S. and France and 0.91 for the U.K. and France). When viewed over the two sample periods, correlations decrease everywhere, mildly so for France and the U.K., but more steeply for the other correlations. These correlations are now quite

TABLE 5

Annualized 5-Year Zero-Coupon Real Yields

Table 5 shows results obtained using the monthly data. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	France	U.K.	U.S.
Full sample: 2004M11–2019M12			
Average	−0.46% (0.59%)	−1.41% (0.56%)	−0.04% (0.38%)
Std. Dev.	1.23% (0.12%)	1.83% (0.16%)	0.97% (0.14%)
β wrt U.S.	0.69 (0.18)	1.34 (0.10)	1.00
Correlation with U.S.	0.54 (0.14)	0.71 (0.07)	1.00
Correlation with U.K.	0.91 (0.08)	1.00	0.71 (0.07)
Subsample 1: 2004M11–2012M5			
Average	0.59% (0.47%)	−0.06% (0.72%)	0.34% (0.66%)
Std. Dev.	0.77% (0.11%)	1.61% (0.18%)	1.07% (0.19%)
β wrt U.S.	0.55 (0.09)	1.36 (0.09)	1.00
Correlation with U.S.	0.77 (0.14)	0.90 (0.09)	1.00
Correlation with U.K.	0.81 (0.12)	1.00	0.90 (0.09)
Subsample 2: 2012M6–2019M12			
Average	−1.51%*** (0.18%)	−2.76%*** (0.19%)	−0.42% (0.24%)
Std. Dev.	0.48% (0.11%)	0.68%** (0.12%)	0.66%* (0.17%)
β wrt U.S.	−0.38*** (0.19)	−0.21*** (0.07)	1.00
Correlation with U.S.	−0.51*** (0.21)	−0.20*** (0.08)	1.00
Correlation with U.K.	0.75 (0.11)	1.00	−0.20*** (0.08)

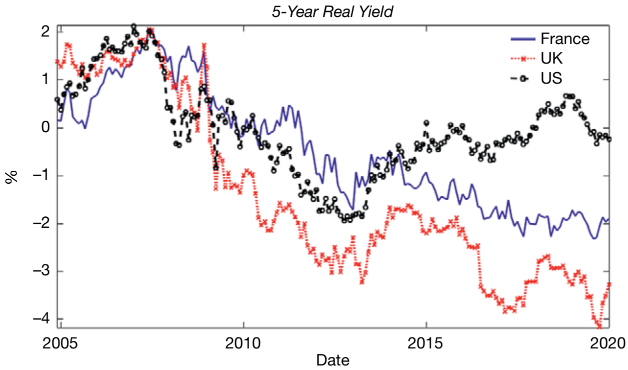
low at −0.51 for the France–U.S. and − 0.20 for the U.K.–U.S. correlation, and the betas with respect to U.S. rates are significantly negative in the second subsample for both France and the U.K. Figure 3 graphs the real yields illustrating these patterns. Overall, cleansing inflation-linked yields from the effects of liquidity premiums does not change the properties of real yields all that much (compare the right panel of Table 1 with Table 5).

D. Variance Decompositions

With all the components identified, we now determine their contribution to the observed yield variation. We begin with nominal yields. In Panel A of Table 6, we show variance decompositions of nominal yield variation into the variation of its 3 components. To keep the decompositions simple, we report $\frac{\mathrm{cov}(\mathrm{COMPONENT}_i,\mathrm{NOMINAL\_YIELD})}{\mathrm{var}(\mathrm{NOMINAL\_YIELD})}$, for i =real yield, expected inflation, and inflation risk premium, respectively. This assigns the covariance terms equally to the three components and the three components add to 100%. The main message is that real

FIGURE 3

Annualized 5-Year Real Yields



yield variation is the dominant component of the yield variation for all three countries and both sample periods. The exact magnitude, the variation over time, and what other components are important vary across countries.

For France, real yield variation accounts for about 77% of nominal yield variation in the first sample period, 85% in the second, and 75% overall. The remainder is mostly accounted for by the inflation risk premium, except in the second subsample where expected inflation accounts for almost 15% of the total variation. For the U.K., real yield variation accounts for about 115% of nominal yield variation in the first sample period, 135% in the second, and 110% overall. Ratios above 100% occur when there are other, negative covariance contributions, and for the U.K. real yields and expected inflation are negatively correlated. A negative correlation between real yields and expected inflation is inconsistent with activist monetary policy, but during our sample may be partially driven by longer term offsetting trends in expected inflation (up) and real yields (down). The contribution of the inflation risk premium is positive overall and in the first subsample, but it is strongly negative ($<-30\%$) in the second subsample), because the correlation between the inflation risk premium and real yields was negative then.

For the U.S., real yield variation accounts for about 75% of nominal yield variation in the first sample period, almost 100% in the second, and 75% overall. The contribution of expected inflation is mostly larger overall than in the other countries, but still smallish at about 10%, and it drops to close to zero in the second subsample. The inflation risk premium accounts for about 15% of the total nominal yield variation over the full sample, but the share becomes slightly negative in the second period. Even though real yields and inflation risk premiums decreased over time, they show a positive covariation in the first sample period but a negative covariation in the second subsample resulting in the negative contribution of the inflation risk premium to nominal yield variation.

We can do a similar decomposition for inflation-linked yields, splitting them in real yields and liquidity premiums. The results are in Panel B of Table 6. Again, real yields dominate, accounting for between 82% and 122% of the total variation of inflation-linked yields. The numbers are often above 1, because the covariance between liquidity premiums and real yields is mostly negative.

TABLE 6

5-Year Zero-Coupon Yield Variance Decompositions

Table 6 shows results obtained using the monthly data over the period of 2004M11 to 2019M12. Subsample 1 is 2004M11–2012M5. Subsample 2 is 2012M6–2019M12. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	France			U.K.			U.S.		
	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2
Panel A. Nominal Yield Variance Decomposition									
cov(REAL_YIELD, NOMINAL_YIELD) / var(NOMINAL_YIELD)	77.22%	84.80%	74.92%	112.67%	116.73%	134.58%	75.27%	77.68%	99.36%
	(3.62%)	(7.91%)	(8.12%)	(4.06%)	(8.41%)	(11.34%)	(7.57%)	(7.93%)	(14.10%)
cov(EXPECTED_INFLATION, NOMINAL_YIELD) / var(NOMINAL_YIELD)	6.29%	5.64%	14.44%	−19.65%	−23.21%	−2.92%*	10.34%	9.19%	0.80%
	(1.18%)	(2.87%)	(3.29%)	(3.00%)	(6.75%)	(2.44%)	(2.75%)	(2.99%)	(4.20%)
cov(INFLATION_RISK_PREMIUM, NOMINAL_YIELD) / var(NOMINAL_YIELD)	16.49%	9.56%	10.64%	6.98%	6.48%	−31.66%***	14.39%	13.13%	−0.17%
	(2.74%)	(5.69%)	(5.94%)	(4.61%)	(7.93%)	(12.85%)	(5.22%)	(6.91%)	(10.07%)
Panel B. Inflation-Linked Yield Variance Decomposition									
cov(REAL_YIELD, INFLATION-LINKED_YIELD) / var(INFLATION-LINKED_YIELD)	107.30%	104.08%	105.77%	105.59%	121.94%	109.63%	81.74%	85.45%	102.31%
	(5.00%)	(14.53%)	(2.95%)	(6.43%)	(11.75%)	(8.19%)	(12.07%)	(16.46%)	(1.51%)
cov(LIQUIDITY_PREMIUM, INFLATION-LINKED_YIELD) / var(INFLATION-LINKED_YIELD)	−7.30%	−4.08%	−5.77%	−5.59%	−21.94%	−9.63%	18.26%	14.55%	−2.31%
	(5.00%)	(14.53%)	(2.95%)	(6.43%)	(11.75%)	(8.19%)	(12.07%)	(16.46%)	(1.51%)

The real yield dominance is such that if we simply split up nominal interest rates in inflation-linked yields and the remainder (inflation compensation), it is still the case that inflation-linked yields dominate the variation of nominal interest rates. This decomposition does not require measurement of the liquidity premium, but, unlike our main decompositions, is difficult to interpret economically. These results are in Supplementary Appendix IV.B.

E. Comovement Decompositions

In Table 7, we report a decomposition of the international correlation between nominal yields. Recall that these correlations were generally high, but decreased in the second part of the sample, with the decrease particularly dramatic for the French–U.S. correlation. Because the nominal yield has three components, there would be nine components in a full correlation decomposition. We therefore report a simpler decomposition, analogous to the one reported for the variance. We compute the covariance with the three components of the nominal yield in one country with the nominal yield in the other country, but scale all of them by the product of the nominal yield standard deviations so that the three numbers sum to the total correlation. We can do this decomposition from the perspective of the two countries, but they provide a clear picture of what component dominates the correlation. Overwhelmingly, it is the covariance between real yields across countries that constitutes the most important component of the total correlation.

This is always true over the full sample with the real yield covariance accounting for 0.74 (French perspective) or 1.03 (U.K. perspective) of a 0.93 correlation for French–U.K. nominal yields, for 0.53/0.40 of a 0.67 French–U.S. correlation and for 0.94/0.58 of a 0.85 U.K.–U.S. correlation. The second most important component is invariably the inflation risk premium with expected inflation even providing a negative contribution to the France–U.K. correlation.

For the first subsample, the overall correlation remains positive and the overall result in terms of relative contribution, with real rates dominating followed by inflation risk premiums, mostly holds. For the second subsample, the results are slightly different. For France and the U.K., the nominal yield correlation is still robustly positive and dominated by real yield comovements: 0.63 of 0.61 from the French perspective and 0.89 of 0.61 from the U.K. perspective. Expected inflation is now more important than inflation risk premiums, with the latter's contribution now negative. For France and the U.S., the nominal yield correlation drops to -0.38. Clearly, the negative correlation is driven by the highly negative covariance between the French and U.S. real yields, which show up in the real yield component from both perspectives. Again, expected inflation is not important, but from the U.S. perspective there is a strong positive contribution of the inflation risk premium component. The U.K.–U.S. nominal yield correlation is close to zero: relatively large and positive contributions of inflation risk premiums are partially offset by negative contributions of real yields and expected inflation.

Panel B of Table 7 investigates the cross-country correlation decompositions of inflation-linked yields. It shows that, as for nominal yields, real yield correlations are the main driver of correlation levels and their time variation. This is also true for

TABLE 7

5-Year Zero-Coupon Yield Correlation Decompositions

Table 7 shows results obtained using the monthly data over the period of 2004M11 to 2019M12. Subsample 1 is 2004M11–2012M5. Subsample 2 is 2012M6–2019M12. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	France–U.K.			France–U.S.			U.K.–U.S.		
	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2
Panel A. Nominal Yield Correlation Decomposition									
Panel A1. Country 1 Perspective									
$\frac{\text{cov}(\text{REAL\_YIELD}_1,\text{NOMINAL\_YIELD}_2)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.74	0.78	0.63	0.53	0.67	−0.43***	0.94	1.12	−0.27***
	(0.04)	(0.10)	(0.05)	(0.11)	(0.14)	(0.15)	(0.09)	(0.12)	(0.11)
$\frac{\text{cov}(\text{EXPECTED\_INFLATION}_1,\text{NOMINAL\_YIELD}_2)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.05	0.03	0.01	0.03	0.02	−0.03	−0.19	−0.24	−0.03
	(0.01)	(0.03)	(0.05)	(0.02)	(0.02)	(0.07)	(0.04)	(0.06)	(0.12)
$\frac{\text{cov}(\text{INFLATION\_RISK\_PREMIUM}_1,\text{NOMINAL\_YIELD}_2)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.14	0.08	−0.03**	0.11	0.07	0.07	0.09	0.06	0.38*
	(0.02)	(0.06)	(0.05)	(0.03)	(0.06)	(0.12)	(0.05)	(0.07)	(0.10)
Panel A2. Country 2 Perspective									
$\frac{\text{cov}(\text{REAL\_YIELD}_2,\text{NOMINAL\_YIELD}_1)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	1.03	0.98	0.89	0.40	0.58	−0.73***	0.58	0.73	−0.05***
	(0.07)	(0.11)	(0.25)	(0.14)	(0.15)	(0.17)	(0.11)	(0.11)	(0.10)
$\frac{\text{cov}(\text{EXPECTED\_INFLATION}_2,\text{NOMINAL\_YIELD}_1)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	−0.16	−0.17	0.23**	0.11	0.09	0.05	0.11	0.09	−0.05***
	(0.02)	(0.07)	(0.08)	(0.02)	(0.04)	(0.03)	(0.02)	(0.03)	(0.03)
$\frac{\text{cov}(\text{INFLATION\_RISK\_PREMIUM}_2,\text{NOMINAL\_YIELD}_1)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.06	0.07	−0.52**	0.16	0.09	0.30*	0.16	0.11	0.18
	(0.04)	(0.09)	(0.16)	(0.05)	(0.03)	(0.11)	(0.04)	(0.05)	(0.06)
Panel A3. Total Correlation									
	0.93	0.89	0.61*	0.67	0.76	−0.38***	0.85	0.93	0.08***
	(0.08)	(0.12)	(0.11)	(0.14)	(0.14)	(0.21)	(0.07)	(0.09)	(0.08)

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TABLE 7 (continued)

5-Year Zero-Coupon Yield Correlation Decompositions

Panel B. Inflation-Linked Yield Correlation Decomposition

	France–U.K.			France–U.S.			U.K.–U.S.		
	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2
Panel B1. Country 1 Perspective									
$\frac{\text{cov(REAL\_YIELD}_1\text{, INFLATION-LINKED\_YIELD}_2)}{SD\text{(INFLATION-LINKED\_YIELD}_1)SD\text{(INFLATION-LINKED\_YIELD}_2)}$	1.03	0.97	0.86	0.72	0.88	−0.59***	0.83	1.10	−0.24***
	(0.05)	(0.16)	(0.06)	(0.15)	(0.23)	(0.15)	(0.12)	(0.17)	(0.14)
C $\frac{\text{cov(LIQUIDITY\_PREMIUM}_1\text{, INFLATION-LINKED\_YIELD}_2)}{SD\text{(INFLATION-LINKED\_YIELD}_1)SD\text{(INFLATION-LINKED\_YIELD}_2)}$	−0.10	−0.14	−0.10	−0.06	−0.10	0.09	−0.06	−0.17	−0.01
	(0.05)	(0.15)	(0.04)	(0.06)	(0.19)	(0.02)	(0.07)	(0.15)	(0.03)
Panel B2. Country 2 Perspective									
$\frac{\text{cov(REAL\_YIELD}_2\text{, INFLATION-LINKED\_YIELD}_1)}{SD\text{(INFLATION-LINKED\_YIELD}_1)SD\text{(INFLATION-LINKED\_YIELD}_2)}$	0.96	0.94	0.83	0.44	0.62	−0.55***	0.57	0.82	−0.27***
	(0.08)	(0.15)	(0.06)	(0.17)	(0.19)	(0.15)	(0.15)	(0.16)	(0.16)
$\frac{\text{cov(LIQUIDITY\_PREMIUM}_2\text{, INFLATION-LINKED\_YIELD}_1)}{SD\text{(INFLATION-LINKED\_YIELD}_1)SD\text{(INFLATION-LINKED\_YIELD}_2)}$	−0.02	−0.12	−0.07	0.23	0.17	0.05	0.21	0.11	0.02
	(0.06)	(0.12)	(0.07)	(0.10)	(0.15)	(0.02)	(0.10)	(0.14)	(0.02)
Panel B3. Total Correlation									
	0.94	0.83	0.76	0.66	0.79	−0.50***	0.78	0.93	−0.25***
	(0.08)	(0.14)	(0.11)	(0.17)	(0.09)	(0.13)	(0.12)	(0.05)	(0.13)

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the negative correlations observed between French and U.S. and between U.K. and U.S. real yields in the second sample.

F. Robustness Checks

In this section, we summarize the results of a number of robustness exercises regarding the measurement of various inputs to the computations.

1. Measurement of the Liquidity Premium

We conduct four robustness checks. First, we consider a model where we only use the inflation swap spread as an independent variable. The temporal evolution of the other variables we use may correlate with that of inflation risk premiums, which would bias our estimates. For example, the inflation-linked debt issuance volumes may depend on the magnitude of the inflation risk premium. Our results, reported in Supplementary Appendix IV.C, are robust, which is not surprising given that the inflation swap spread has by far the highest explanatory power for variation in the nominal debt premium.[8] Our results also remain robust when we explicitly impose the coefficient of -1 on the inflation swap spread in the liquidity premium regression (3). This would correspond to the assumption that inflation swap spreads are perfectly liquid, as in Haubrich et al. (2012), an assumption nonetheless criticized by Christensen and Gillan (2012) and To and Tran (2019).

Second, we verify the usefulness of an alternative general liquidity measure, namely, Hu, Pan, and Wang's (2013) "noise measure," which measures the pricing errors of fitted yields. We describe the calculations in Supplementary Appendix IV. D and then repeat the liquidity premium estimation exercise in Table 3. Because the Hu et al. (2013) measure is insignificant when we add our other liquidity proxies, we do not include it in the specification we report. However, including it does not affect any of our conclusions.

Third, in Supplementary Appendix IV.E we reestimate the models country by country rather than in a panel. All our results remain robust.

Finally, the liquidity premium is estimated with error as it relies on the fitted value of a panel regression. To assess the effect of this sampling error, we draw 10,000 possible coefficient vectors from their asymptotic distribution, recompute the liquidity premiums, and then recompute our variance and correlation decompositions. Supplementary Appendix IV.F shows 95% confidence intervals for our estimates. The confidence intervals are rather tight, and our conclusions regarding the primary importance of real rate variation remain intact. For the variance decomposition, both for nominal yields and inflation-linked yields, the confidence intervals for the other components do not overlap at all with that for the real yield contribution. As a concrete example, the U.S. real yield accounts for 75% of the variation of the U.S. nominal yield, with a confidence interval of [67%, 83%], whereas the upper bound of the confidence interval for the inflation risk premium contribution is 23%. Similarly, with one exception, we get nonoverlapping

[8]Note that the inflation swap *spread* is the rate on a zero-coupon inflation swap position paying fixed and receiving floating minus the difference between the zero-coupon nominal and inflation-linked yields. Thus, unlike the inflation swap *rate*, it does not have a strong relationship with the spread between nominal and index-linked debt by (near) arbitrage.

confidence intervals for the different contributions to the yield correlations across countries. For example, for the France–U.K. nominal yield correlation (from the French perspective), the real yield contribution is 0.74 with a confidence interval of [0.66, 0.82], whereas the upper bound to the inflation risk premium contribution is 0.22. In many cases, the confidence intervals are quite a bit tighter than the ones reported here.

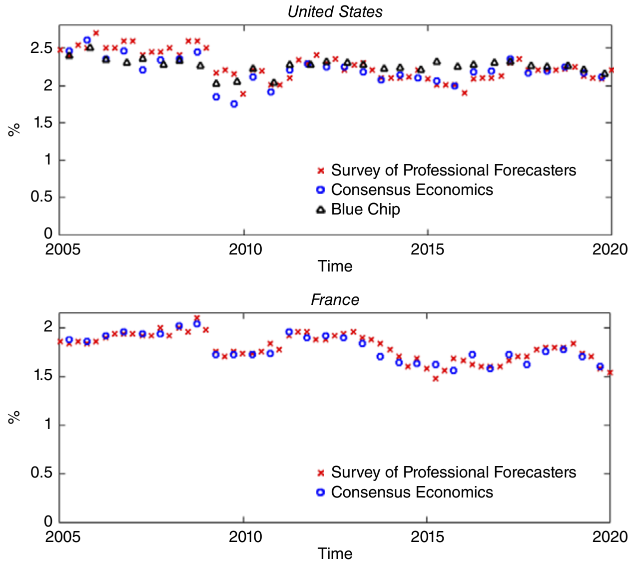
2. Measurement of Inflation Expectations

As a first robustness check, we use alternative inflation surveys to the ones used in our main specification. There also exist Consensus Economics survey forecasts for France and the U.S., and we recompute our variance and correlation decompositions using survey long-term expected inflation from the semi-annual Consensus Economics survey for all three countries. The results are in Supplementary Appendix III.B for the case where we do not change the forecasts in between data releases (but they are almost identical under linear interpolation in between dates). For the U.S., the Blue Chip survey also produces long-term inflation forecasts, but they are also only available semi-annually. In Supplementary Appendix III.B, we also generate our results using that survey, coupled with the Consensus Economics surveys for the U.K. and France. Again, our results prove robust. This is not surprising, as the forecasts produced by the surveys are very similar. To illustrate this fact, Figure 4 shows the forecast from the three different surveys for the U.S. This similarity is likely driven by the fact that many survey respondents in the various surveys overlap.

As an alternative to survey expectations, we compute a statistical estimate of 5-year expected inflation, based on a country-specific vector autoregression (VAR) with quarterly data. The VAR variables include the 3-month nominal short rate, realized quarterly inflation, and either one quarter (for the U.S.) or 1-year inflation survey forecasts (for France and the U.K.). We use a lag length of 2 (the BIC optimal length for the realized quarterly inflation is 1 or 2 depending on the country). We also correct the parameter estimates for small sample bias. This is done by resampling 10,000 time series of historical length under the estimated parameters (i.e., the VAR residuals are bootstrapped and the actual series recreated using the VAR parameters), reestimating the VAR parameters for each time series, and computing the bias as the difference between the average of the parameters estimated from the 10,000 sampled series and the parameters used to simulate. From the VAR, we can compute the forecast for 5-year (20 quarter) future (annualized) inflation. Because our yield data are monthly, we simply keep the forecast the same within a quarter. However, linearly interpolating the forecasts to monthly data has no material impact on our results. Statistical expected inflation is slightly more (less) variable in the U.S. and France (the U.K.), compared to the survey forecasts, but its overall variability remains very low. It is therefore not surprising that our key result, that real rates dominate the variation in nominal yields and are the main driver of nominal yield comovements across countries, remains robust with the statistical measure of inflation. These results are reported in Supplementary Appendix III.C.

FIGURE 4

Annualized 5-Year Expected Inflation From Different Surveys



G. Results for Other Maturities

We repeat our reduced-form analysis for 2- and 10-year bonds. The 2-year maturity is the shortest for which we can reliably use inflation-linked bonds (Gürkaynak et al.(2010)). This is because at shorter maturities indexation lag effects and specific trading associated with them become a significant concern. This is especially true in the U.K., where the indexation lag for bonds issued pre-2005 is 8 months (it is 3 months for post-2005 bonds; in France and the U.S. the indexation lag is 2.5 months). D'Amico et al. (2018) argue that the indexation lag effects can still be significant at maturities between 2 and 4 years. Thus, we provide 2-year results only as suggestive additional evidence. The second new maturity we introduce is 10 years. This choice is driven by the lack of longer-term survey inflation expectations.

The results in Supplementary Appendix V confirm the dominant role of real yields in variance and correlation decompositions of 2- and 10-year yields. The 2- and 10-year yields decompositions are very similar to 5-year yields decompositions. For example, for French 10-year nominal yields, $\frac{\mathrm{cov(REAL\_YIELD,NOMINAL\_YIELD)}}{\mathrm{var(NOMINAL\_YIELD)}} = 74.14\%$, $\frac{\mathrm{cov(EXPECTED\_INFLATION,NOMINAL\_YIELD)}}{\mathrm{var(NOMINAL\_YIELD)}} = 3.80\%$, and $\frac{\mathrm{cov(INFLATION\_RISK\_PREMIUM,NOMINAL\_YIELD)}}{\mathrm{var(NOMINAL\_YIELD)}} = 22.07\%$, while these numbers for 5-year bonds in Table 6 are 77.22%, 6.29%, and 16.49%, respectively.

V. A No-Arbitrage Term Structure Model

Our previous results rely on empirical proxies for inflation expectations and liquidity premiums to estimate the components of nominal yields in a model-free fashion. In this section, we use a no-arbitrage model with real and nominal factors to provide alternative estimates of our key decompositions. While the results so far appear very robust to different choices made, the use of a term structure model has the benefit of more efficiently using information in actual yields and imposing no-arbitrage restrictions across yields of different maturities. Section V.A outlines the model and discusses its estimation; the decomposition results are discussed in Section V.B. Finally, in Section V.C we explore implications for exchange rate dynamics.

A. A No-Arbitrage Term Structure Model With Real and Nominal Factors

Our model follows closely the Gaussian affine term structure literature (see, e.g., Joslin, Singleton, and Zhu (2011)) where principal components of yields are used as factors, and prices of risk are linear in the factors. Our particular application closely follows the set up in Abrahams et al. (2016).

1. Model

We assume that a $K \times 1$ vector of pricing factors, X_t, evolves under the physical measure as

$$X_{t+1} = \mu_X + \Phi(X_t - \mu_X) + v_{t+1}, \tag{5}$$

where μ_X is the $K \times 1$ unconditional mean vector, Φ is a $K \times K$ matrix, and $v_{t+1} \sim \mathcal{N}(0, \Sigma)$ with 0 a $K \times 1$ zero vector and Σ a $K \times K$ covariance matrix. We assume that the log-nominal stochastic discount factor is

$$m_{t+1} = -r_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\Sigma^{-\frac{1}{2}}v_{t+1}, \tag{6}$$

where r_t is the nominal short rate and $\lambda_t = \Sigma^{-\frac{1}{2}}(\lambda_0 + \lambda_1 X_t)$ with λ_0 a $K \times 1$ vector and λ_1 a $K \times K$ matrix. Thus, the price of risk, λ_t, is linear in the state variables but shocks are homoscedastic (Duffee (2002)). This implies that the state variable dynamics under the risk-neutral measure are governed by the following parameters:

$$\begin{aligned} \tilde{\mu} &= (I_{K\times K} - \Phi)\mu_X - \lambda_0, \\ \tilde{\Phi} &= \Phi - \lambda_1, \end{aligned} \tag{7}$$

where $I_{K \times K}$ is a $K \times K$ identity matrix.

It follows that the time t price of an n-period risk-free nominal zero-coupon bond (P_t^n) satisfies a set of recursive equations (with $A_0 = 0, B_0 = 0$):

$$\begin{aligned} \log P_t^n &= A_n + B_n' X_t, \\ A_n &= A_{n-1} + B_{n-1}'\tilde{\mu} + \frac{1}{2}B_{n-1}{}'\Sigma B_{n-1} - \delta_0, \\ B_n' &= B_{n-1}'\tilde{\Phi} - \delta_{1'}, \end{aligned} \tag{8}$$

where δ_0 is a constant and δ_1 is a $K \times 1$ vector.

So far, the model is a standard nominal term structure model, pricing nominal yields. To price zero coupon inflation-linked bonds, we use the fact that such bonds pay out gross inflation at maturity. This requires a model for inflation. Following

Abrahams et al. (2016), we assume that inflation is an affine function of the K pricing factors, X_t: $\pi_0 + \pi_1' X_t$. It follows that the time t price of an n-period risk-free inflation-linked zero-coupon bond, $P^n_{t,\mathrm{il}}$, satisfies a set of recursive equations, with $A_{0,\mathrm{il}} = 0, B_{0,\mathrm{il}} = 0$:

$$(9)\quad \log P^n_{t,\mathrm{il}} = A_{n,\mathrm{il}} + B'_{n,\mathrm{il}} X_t,$$

$$A_{n,\mathrm{il}} = A_{n-1,\mathrm{il}} + (B_{n-1,\mathrm{il}} + \pi_1)'\tilde{\mu} + \frac{1}{2}(B_{n-1,\mathrm{il}} + \pi_1)'\Sigma(B_{n-1,\mathrm{il}} + \pi_1) - \delta_0 + \pi_0,$$

$$B'_{n,\mathrm{il}} = (B_{n-1,\mathrm{il}} + \pi_1)'\tilde{\Phi} - \delta_1'.$$

2. Factors

We estimate the model country-by-country. To bring the model to the data, the factors X_t must be determined. All factors are assumed to be observed. First, we specify nominal factors as the smallest number of principal components of nominal yields, which together explain over 99.5% of the yield variation. We use 1 month, and 1–10-year nominal zero-coupon yields for each country (11 yields in total) to extract principal components.[9] This procedure results in 3 factors for each country. We scale the factors to have zero mean and unit variance.

Second, because we cannot assume that TIPS produce the real interest rate factors, we must specify a liquidity factor. The different sensitivity of nominal and real yields to this factor then implicitly determines the liquidity premium; so that, liquidity-adjusted yields can be constructed. We assume that the liquidity factor is an affine function of the inflation swap spread. Compared to alternative specifications such as using a weighted average of several standardized liquidity proxies (the inflation swap spread, the log of the share of inflation-linked debt, and the nominal off-the run premium) or the first principal component of these proxies, this results in a slightly better model fit, but our conclusions are robust across these alternative specifications. We use the 5-year inflation swap spread, but the maturity has only a minor effect on the results. We also do not find that adding other liquidity proxies as separate pricing factors improves the model fit or affects the decomposition outcomes.

Third, given these two sets of factors, we determine a set of real factors, by regressing zero-coupon inflation-linked yields on our nominal and liquidity factors and pick the smallest number of principal components which explain over 99.5% of the residual variation as our real factors. We use 2–10-year inflation-linked zero-coupon yields (9 in total) for the U.S. and France and 4–10-year inflation-linked zero-coupon yields (7 in total) for the U.K. We do not use shorter-maturity inflation-linked yields due to the indexation lag factors discussed above. For the U.K., we

[9]For France only bonds of maturity 1 month, 3, 6, 9 months, 1 year, 2, 5, 10, and 30 years are available. We construct the missing maturities (3, 4, 6, 7, 8, and 9 years) by fitting the Svensson (1994) zero-coupon model to French nominal bonds. This estimation procedure, excluding bonds with maturities below (above) 6 months (15 years) and all strips, achieves a near perfect fit for maturities reported on the Banque de France website.

also do not use 2- and 3-year zero-coupon yields.[10] The procedure results in 2 real factors for the U.S. and 3 real factors for France and the U.K. We again scale the factors to have zero mean and unit variance.

These three sets of factors suffice to estimate the model. In this model, inflation, expected inflation, and the inflation risk premium are completely driven by financial market information and no inflation data is used. Alternatively, we can incorporate survey inflation forecasts by adding expected and realized inflation as pricing factors and impose the corresponding restrictions on π_0 and π_1 (see, e.g., Kim and Orphanides (2012)). However, we prefer to obtain purely financial markets-implied inflation forecasts and contrast them with our reduced-form results which are obtained using only survey inflation expectations. While we estimate the inflation-less model with similar results to our reported model, our main model also includes 1month of realized inflation as a factor, because it improves the fit of the model. Measures of real growth, such as industrial production or consumption growth, do not improve the fit, after controlling for the inflation factor. Importantly and on purpose, we do not impose any restrictions that inflation in the model should be related to our inflation factor: the model-implied inflation is just a linear combination of factors, which best fits the inflation-linked yields.

3. Estimation Results

We discuss the estimation procedure, which largely follows Adrian et al. (2013), and report the parameter estimates in Supplementary Appendix VI. Figure 5 shows the actual zero-coupon yields together with the model-implied yields for the 5-year maturity, showing them to be nearly indistinguishable. Table 8 shows the mean absolute and root-mean-square errors for 2-,[11] 5- and 10-year yields. The errors are always lower than 0.12%, and mostly 0.05% or lower; they tend to be slightly higher for the lowest maturity.

Given the estimated parameters, expected inflation can be obtained as the difference between nominal and inflation-linked zero-coupon yields computed under the physical measure. This can be done by replacing μ and Φ with $(I_K - \Phi)\mu_X$ and Φ, respectively, in the nominal and inflation-linked bond pricing equations (8) and (9). The inflation risk premium can then be computed as the difference between the nominal and inflation-linked zero-coupon yields under the risk-neutral measure, computed exactly as in equations (8) and (9), minus expected inflation. To adjust for liquidity effects, we remove the component due to the liquidity factor from the nominal-inflation-linked yield spread under both the risk-neutral and physical measures while computing expected inflation and the inflation risk premium. Formally, in our model both nominal and inflation-linked bond prices load on the liquidity factor. Therefore, we refer to the difference between liquidity components of inflation-linked and nominal yields as the liquidity premium. Empirically, we find the contribution of the liquidity factor to nominal yields negligible.

[10]The Bank of England data have several missing values on yields and returns for those maturities, preventing their use.

[11]Four-year yields for the U.K. inflation-linked bonds due to the data issues discussed earlier.

FIGURE 5

Model Yield Fit

The sample is monthly 2004M11–2019M12. Yields are annualized zero-coupon yields.

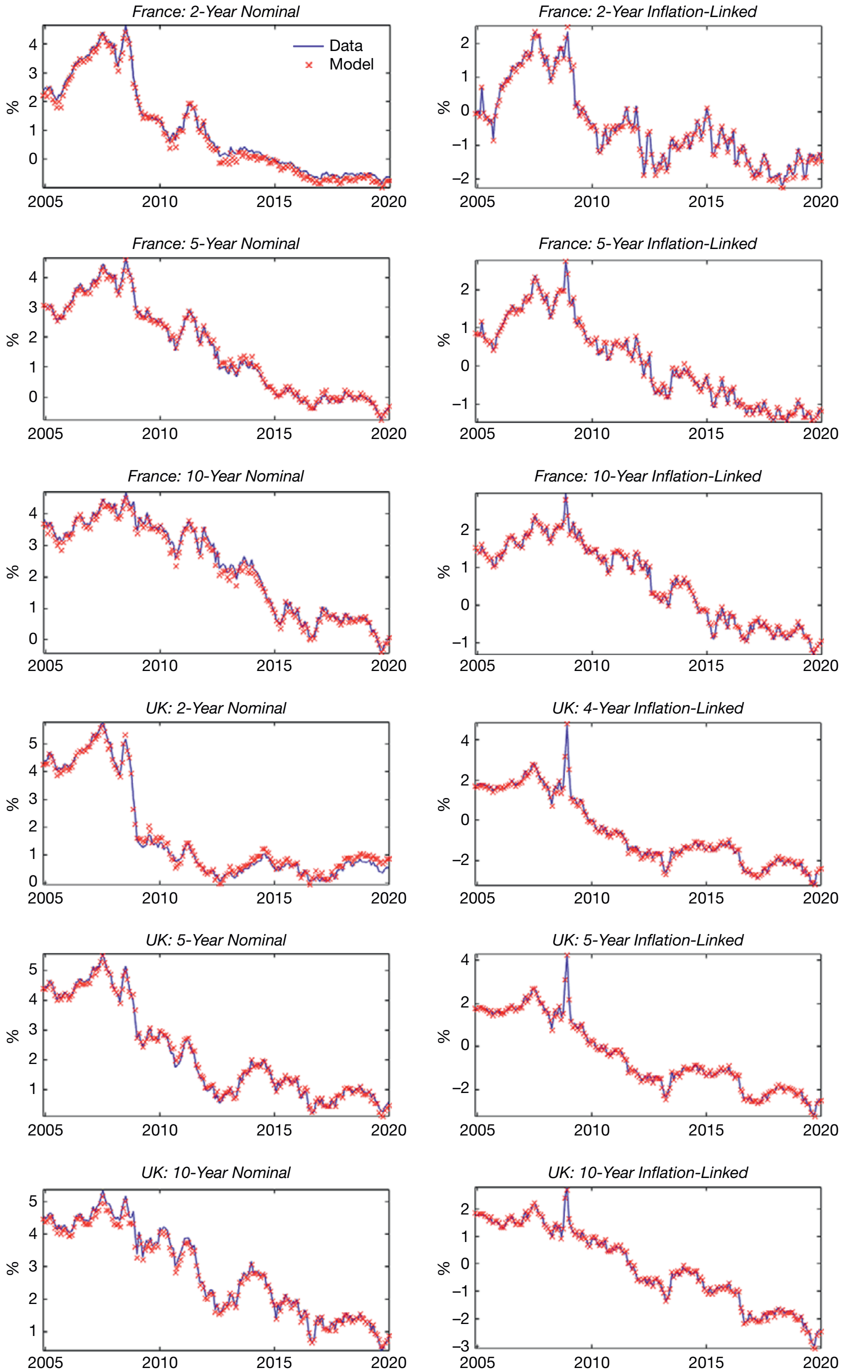


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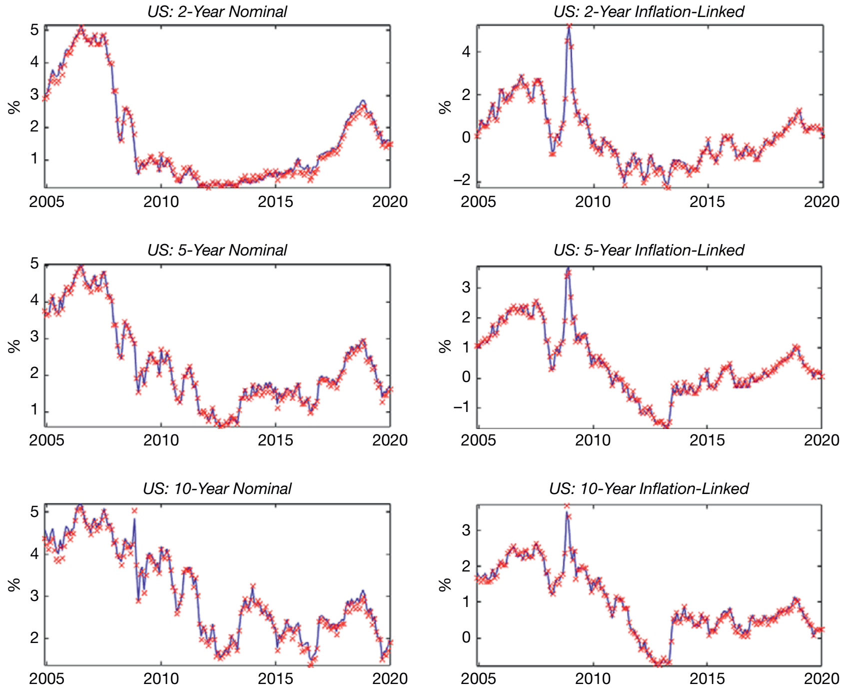
FIGURE 5 (continued)

Model Yield Fit



B. Variance and Correlation Decompositions

The model generates results for any maturity. To enhance comparability with our previous results, we focus on 5-year yields, and also report results for 10-year yields. Table 9 reports our variance decompositions for the 5- and 10-year nominal and inflation-linked yields. Panels A and B focus on 5-year nominal and inflation-linked yields, which can be compared with the results in Table 6. The dominance of the real yield contributions is again very apparent. In Panel A, its smallest contribution to the nominal yield variance is 60.58% for France in the second subsample. This is also the entry with the largest difference relative to Table 6, as the real yield contribution is 74.92% in our main specification in Section IV. The other numbers are remarkably close to the numbers in Table 6. In Panel B, the contribution of the real yield to inflation-linked yields is never lower than 73.80% (full sample for the U.S.); the largest difference is the real yield contribution for the U.K. in the first subsample being 111.90% whereas it is 121.94% in the main specification. Overall, given the sampling error in these estimates, it is obvious that our main conclusions continue to hold. These results also hold for 10-year yields in Panels C and D, with the relative real contributions being quite similar to those for 5-year yields.

In terms of the relative contributions of the other components to nominal yield variation, we do obtain sometimes somewhat different results. For the full sample, it

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TABLE 8

Model Yield Fitting Errors

Table 8 shows results obtained using the monthly data over the period of 2004M11 to 2019M12. Yields are annualized zero-coupon yields.

Panel A. France

	2-Year Nominal	5-Year Nominal	10-Year Nominal
MEAN_ABSOLUTE_ERROR	0.10%	0.06%	0.07%
ROOT-MEAN-SQUARE_ERROR	0.12%	0.07%	0.09%
	2-Year Inflation-Linked	5-Year Inflation-Linked	10-Year Inflation-Linked
MEAN_ABSOLUTE_ERROR	0.04%	0.02%	0.02%
ROOT-MEAN-SQUARE_ERROR	0.06%	0.03%	0.03%

Panel B. U.K.

	2-Year Nominal	5-Year Nominal	10-Year Nominal
MEAN_ABSOLUTE_ERROR	0.06%	0.05%	0.06%
ROOT-MEAN-SQUARE_ERROR	0.09%	0.07%	0.09%
	4-Year Inflation-Linked	5-Year Inflation-Linked	10-Year Inflation-Linked
MEAN_ABSOLUTE_ERROR	0.05%	0.03%	0.03%
ROOT-MEAN-SQUARE_ERROR	0.06%	0.04%	0.04%

Panel C. U.S.

	2-Year Nominal	5-Year Nominal	10-Year Nominal
MEAN_ABSOLUTE_ERROR	0.08%	0.03%	0.05%
ROOT-MEAN-SQUARE_ERROR	0.10%	0.05%	0.07%
	2-Year Inflation-Linked	5-Year Inflation-Linked	10-Year Inflation-Linked
MEAN_ABSOLUTE_ERROR	0.05%	0.03%	0.03%
ROOT-MEAN-SQUARE_ERROR	0.07%	0.04%	0.04%

remains the case that the inflation risk premium has a higher variance contribution than expected inflation for all three countries, with the absolute magnitude even rather close (except for the U.S.). The results differ the most for the second subsample. For example, for France, we now find a very large contribution of the inflation risk premium, but a negative one for expected inflation, whereas we obtain relatively similar positive 10%–15% contributions for both components in the main specification. For the U.K., the contribution of expected inflation is essentially 0 in the main specification with the inflation risk premium contributing a negative 32%. The estimates of the term structure model make the inflation risk premium contribution even more negative and the expected inflation contribution solidly positive (at 29%). Similar results hold for 10-year yields. We conclude that the variance contributions generated by the term structure model are consistent with our main results, except for the relative contributions of expected inflation and the inflation risk premium in the second subsample.

Table 10 reports the correlation decomposition for 5-year yields, and should be compared with Table 7 for our main specification. Focusing first on Panel A, for nominal yields, the contribution of the real yield to the total correlation is remarkably similar to our numbers in Table 7. For the full sample, the real rate scaled covariance with the nominal yield (first country perspective) is respectively 0.69 for the France–U.K. pair; 0.50 for the France–U.S. pair and 0.90 for the U.K.–U.S. pair. The corresponding numbers in Table 7 are 0.74, 0.53, and 0.94. Overall, the real rate contributions are insignificantly different from what we observe in Table 7. Again, there are somewhat large differences for the contributions of expected inflation and the inflation risk premium, especially in the second subsample. This is particularly the case for the pairs involving France, where the relative importance rank among expected inflation and the inflation risk premium changes between Tables 7 and 10.

TABLE 9

Model-Implied Variance Decompositions

Table 9 shows results obtained using the monthly data over the period of 2004M11 to 2019M12. Subsample 1 is 2004M11–2012M5. Subsample 2 is 2012M6–2019M12. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	France			U.K.			U.S.		
	Whole Sample	Subsample 1	Subsample 2	Whole Sample	Subsample 1	Subsample 2	Whole Sample	Subsample 1	Subsample 2
Panel A. 5-Year Nominal Yield									
cov(REAL_YIELD, NOMINAL_YIELD) / var(NOMINAL_YIELD)	71.67%	85.60%	60.58%	110.55%	109.80%	128.39%	69.48%	75.30%	96.39%
	(3.18%)	(6.68%)	(7.21%)	(3.23%)	(5.86%)	(11.27%)	(6.59%)	(5.17%)	(12.52%)
cov(EXPECTED_INFLATION, NOMINAL_YIELD) / var(NOMINAL_YIELD)	6.45%	15.40%	−15.53%***	−13.83%	−37.12%	28.78%***	1.59%	9.35%	16.58%
	(1.95%)	(2.68%)	(2.32%)	(4.93%)	(8.23%)	(7.29%)	(3.01%)	(2.61%)	(5.99%)
cov(INFLATION_RISK_PREMIUM, NOMINAL_YIELD) / var(NOMINAL_YIELD)	21.88%	−1.00%	54.95%***	3.28%	27.32%	−57.17%***	28.93%	15.36%	−12.98%
	(3.74%)	(6.78%)	(5.15%)	(7.43%)	(10.51%)	(14.03%)	(9.31%)	(5.31%)	(18.05%)
Panel B. 5-Year Inflation-Linked Yield									
cov(REAL_YIELD, INFLATION-LINKED_YIELD) / var(INFLATION-LINKED_YIELD)	100.25%	100.38%	99.02%	103.29%	111.90%	106.71%	73.80%	76.62%	97.44%
	(3.03%)	(8.61%)	(3.52%)	(3.46%)	(6.81%)	(4.12%)	(12.63%)	(16.42%)	(3.55%)
cov(LIQUIDITY_PREMIUM, INFLATION-LINKED_YIELD) / var(INFLATION-LINKED_YIELD)	−0.25%	−0.38%	0.98%	−3.29%	−11.90%	−6.71%	26.20%	23.38%	2.56%
	(3.03%)	(8.61%)	(3.52%)	(3.46%)	(6.81%)	(4.12%)	(12.63%)	(16.42%)	(3.55%)
Panel C. 10-Year Nominal Yield									
cov(REAL_YIELD, NOMINAL_YIELD) / var(NOMINAL_YIELD)	72.75%	77.06%	64.40%	108.64%	98.27%	117.71%	67.50%	82.98%	69.62%
	(1.85%)	(5.69%)	(3.42%)	(4.90%)	(7.74%)	(11.53%)	(7.10%)	(4.61%)	(11.81%)
cov(EXPECTED_INFLATION, NOMINAL_YIELD) / var(NOMINAL_YIELD)	2.72%	10.10%	−6.45%***	−4.14%	−24.12%	13.88%***	−4.28%	2.62%	−6.60%
	(1.21%)	(2.58%)	(0.81%)	(3.77%)	(7.37%)	(2.48%)	(2.53%)	(2.35%)	(5.22%)
cov(INFLATION_RISK_PREMIUM, NOMINAL_YIELD) / var(NOMINAL_YIELD)	24.53%	12.84%	42.05%**	−4.49%	25.85%	−31.60%***	36.78%	14.41%	36.98%
	(2.34%)	(5.31%)	(2.71%)	(7.38%)	(11.93%)	(13.53%)	(9.43%)	(6.06%)	(16.73%)
Panel D. 10-Year Inflation-Linked Yield									
cov(REAL_YIELD, INFLATION-LINKED_YIELD) / var(INFLATION-LINKED_YIELD)	99.95%	97.93%	99.96%	101.63%	111.68%	102.50%	79.51%	80.10%	98.03%
	(1.16%)	(5.72%)	(1.26%)	(2.05%)	(5.46%)	(1.84%)	(8.88%)	(13.34%)	(1.75%)
cov(LIQUIDITY_PREMIUM, INFLATION-LINKED_YIELD) / var(INFLATION-LINKED_YIELD)	0.05%	2.07%	0.04%	−1.63%	−11.68%	−2.50%	20.49%	19.90%	1.97%
	(1.16%)	(5.72%)	(1.26%)	(2.05%)	(5.46%)	(1.84%)	(8.88%)	(13.34%)	(1.75%)

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TABLE 10

Model-Implied 5-Year Zero-Coupon Yield Correlation Decompositions

Table 10 shows results obtained using the monthly data over the period of 2004M11 to 2019M12. Subsample 1 is 2004M11–2012M5. Subsample 2 is 2012M6–2019M12. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	France–U.K.			France–U.S.			U.K.–U.S.		
	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2	Full Sample	Subsample 1	Subsample 2
Panel A. Nominal Yield Correlation Decomposition									
Panel A1. Country 1 Perspective									
$\frac{\text{cov}(\text{REAL\_YIELD}_1,\text{NOMINAL\_YIELD}_2)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.69	0.79	0.56	0.50	0.66	−0.36***	0.90	1.02	−0.26***
	(0.03)	(0.08)	(0.07)	(0.08)	(0.14)	(0.12)	(0.07)	(0.10)	(0.11)
$\frac{\text{cov}(\text{EXPECTED\_INFLATION}_1,\text{NOMINAL\_YIELD}_2)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.07	0.15	−0.14***	0.09	0.17	0.11	−0.18	−0.32	−0.19
	(0.02)	(0.03)	(0.03)	(0.02)	(0.04)	(0.04)	(0.03)	(0.06)	(0.06)
$\frac{\text{cov}(\text{INFLATION\_RISK\_PREMIUM}_1,\text{NOMINAL\_YIELD}_2)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.16	−0.05	0.18*	0.08	−0.07	−0.16	0.11	0.23	0.47
	(0.04)	(0.06)	(0.09)	(0.05)	(0.06)	(0.10)	(0.04)	(0.08)	(0.13)
Panel A2. Country 2 Perspective									
$\frac{\text{cov}(\text{REAL\_YIELD}_2,\text{NOMINAL\_YIELD}_1)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	1.02	0.94	0.80	0.31	0.50	−0.76***	0.50	0.68	−0.07***
	(0.05)	(0.08)	(0.18)	(0.12)	(0.15)	(0.14)	(0.10)	(0.09)	(0.10)
$\frac{\text{cov}(\text{EXPECTED\_INFLATION}_2,\text{NOMINAL\_YIELD}_1)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	−0.09	−0.37	0.33***	−0.11	0.04	−0.35**	−0.07	0.07	−0.21**
	(0.06)	(0.08)	(0.08)	(0.05)	(0.03)	(0.11)	(0.05)	(0.03)	(0.08)
$\frac{\text{cov}(\text{INFLATION\_RISK\_PREMIUM}_2,\text{NOMINAL\_YIELD}_1)}{SD(\text{NOMINAL\_YIELD}_1)SD(\text{NOMINAL\_YIELD}_2)}$	0.00	0.32	−0.53***	0.47	0.22	0.69	0.41	0.18	0.30
	(0.08)	(0.12)	(0.19)	(0.11)	(0.09)	(0.23)	(0.11)	(0.07)	(0.16)
Panel A3. Total Correlation									
	0.93	0.89	0.60*	0.67	0.76	−0.42***	0.84	0.93	0.02***
	(0.09)	(0.09)	(0.09)	(0.13)	(0.15)	(0.22)	(0.07)	(0.09)	(0.12)
Panel B. Inflation-Linked Yield Correlation Decomposition									
Panel B1. Country 1 Perspective									
$\frac{\text{cov}(\text{REAL\_YIELD}_1,\text{INFLATION-LINKED\_YIELD}_2)}{SD(\text{INFLATION-LINKED\_YIELD}_1)SD(\text{INFLATION-LINKED\_YIELD}_2)}$	0.96	0.92	0.79	0.68	0.84	−0.49***	0.81	1.03	−0.21***
	(0.03)	(0.10)	(0.07)	(0.11)	(0.14)	(0.11)	(0.09)	(0.09)	(0.12)
$\frac{\text{cov}(\text{LIQUIDITY\_PREMIUM}_1,\text{INFLATION-LINKED\_YIELD}_2)}{SD(\text{INFLATION-LINKED\_YIELD}_1)SD(\text{INFLATION-LINKED\_YIELD}_2)}$	−0.02	−0.08	−0.04	−0.02	−0.06	−0.01	−0.03	−0.10	0.00
	(0.03)	(0.09)	(0.02)	(0.04)	(0.10)	(0.02)	(0.04)	(0.08)	(0.02)
Panel B2. Country 2 Perspective									
$\frac{\text{cov}(\text{REAL\_YIELD}_2,\text{INFLATION-LINKED\_YIELD}_1)}{SD(\text{INFLATION-LINKED\_YIELD}_1)SD(\text{INFLATION-LINKED\_YIELD}_2)}$	0.95	0.90	0.80	0.36	0.49	−0.50***	0.49	0.71	−0.19***
	(0.04)	(0.09)	(0.09)	(0.15)	(0.20)	(0.10)	(0.14)	(0.16)	(0.11)
$\frac{\text{cov}(\text{LIQUIDITY\_PREMIUM}_2,\text{INFLATION-LINKED\_YIELD}_1)}{SD(\text{INFLATION-LINKED\_YIELD}_1)SD(\text{INFLATION-LINKED\_YIELD}_2)}$	−0.01	−0.06	−0.05	0.30	0.29	0.00	0.28	0.22	−0.02
	(0.03)	(0.07)	(0.04)	(0.11)	(0.17)	(0.03)	(0.11)	(0.15)	(0.03)
Panel B3. Total Correlation									
	0.94	0.83	0.75	0.66	0.78	−0.49***	0.77	0.93	−0.21***
	(0.08)	(0.13)	(0.12)	(0.17)	(0.09)	(0.12)	(0.12)	(0.06)	(0.14)

Yet, these differences are economically and statistically not that large, given the dominance of the real yield contributions.

In Panel B, the scaled covariances of real yields with inflation-linked yields are again very close between Tables 7 and 10. For the full sample, the real rate scaled covariance with the inflation-linked yield (first country perspective) is respectively 0.98 for the France-U.K. pair; 0.68 for the France–U.S. pair and 0.81 for the U.K.–U.S. pair. The corresponding numbers in Table 7 are 1.03, 0.72, and 0.83. Again, these differences are trivial, especially given the sampling error involved in the estimates. The largest difference we observe is that for the France–U.S. pair; the real rate covariance contributes only 0.49 in the first subsample (second country perspective), whereas that number is 0.62 in Table 7.

We report the correlation decompositions for the 10-year yields in Supplementary Appendix V. We again confirm that the real yield contributions dominate, and that the results are not that different from our results reported earlier.

C. Exchange Rate Dynamics

Let M_t^i be a time t nominal stochastic discount factor in country i and S_t the time t spot exchange rate between countries A and B. If the markets are complete, the well-known no-arbitrage condition (e.g., Lucas (1982)) implies

$$\frac{M_{t+1}^A}{M_{t+1}^B} = \frac{S_{t+1}}{S_t}, \tag{10}$$

or $m_{t+1}^A - m_{t+1}^B = s_{t+1} - s_t$ in logs.

Equation (10) implies that state variables and shocks which span stochastic discount factors should also span exchange rate changes. However, Jotikasthira et al. (2015) find that stochastic discount factor state variables and shocks which price the term structure of international interest rates in Germany, the U.K., and the U.S. only explain 6.7%–10.5% of the variation in the corresponding exchange rate changes.

We repeat the exercise for our model by regressing monthly log nominal exchange rate changes on stochastic discount factor realizations computed following equation (6). We estimate the R^2s at 16.52%, 18.06%, and 13.74% for the USD-GBP, EUR-GBP, and EUR-USD pairs, respectively. These are somewhat higher values than in Jotikasthira et al. (2015). Of course, our values are far away from 100%, indicating that additional work is needed to understand the joint dynamics of interest and exchange rates.

This raises the question whether a model which fits exchange rates perfectly can fit yield dynamics. This question was addressed in Chernov and Creal (2022), in a model similar to ours. We therefore apply their "WFX model" to our specification in Section V.A, adding monthly USD-EUR and USD-GBP exchange rate changes as factors. Furthermore, we impose international no-arbitrage restrictions as in Chernov and Creal (2022). The model fits exchange rates by construction. In unreported results, we find that our variance and correlation decomposition results are almost indistinguishable from the results in Section V.B. This is very much in line with the conclusions of Chernov and Creal (2022) that the models with and without exchange rates imply almost identical bond yield dynamics. We choose not to include the WFX model into this article instead of/in addition to the model we estimate above, as it is less parsimonious, but has the same implications for the issues we are studying.

VI. Extensions

In this section, we consider 3 extensions of our analysis. First, because the real yield in this article is a long-term yield, we decompose the results in a component due to expected future short rates (the expectations hypothesis or EH component), and a component due to the real term premium. Second, we redo all the analysis using slopes rather than levels. Third, we expand our sample of countries to include Germany, Australia, and Sweden (over a shorter sample period).

A. The Expectations Hypothesis Decomposition

Real interest rates vary because of variation in expected future rates or variation in the real term premium. Unfortunately, neither component is observed. We identify expected future short rates using two approaches: a pure statistical model, and survey expectations. The first approach uses a VAR(2) model with short-term nominal interest rates, inflation and inflation expectations. The second approach uses Blue Chip survey forecasts for nominal rates in the U.S. and then extrapolates a linear model on observables (expected real GDP growth and expected inflation) to France and the U.K. as in Wright (2011). Both methodologies forecast nominal short rates and use survey expected inflation to correct for expected inflation, but we check robustness with a statistical estimate of expected inflation. Full details on the methodology and results are provided in Supplementary Appendix VII; here we provide a summary.

For the full sample, the EH component roughly contributes between 54% and 75% of the total variance of the real yield depending on the country and the methodology. Its dominance is very prominent in the first subsample (where its contribution is at least 70%), but in the second subsample its contribution in the U.K. and France decreases in a statistically significant fashion to 6%–12% in the U.K. and about 36% in France. In the U.S., its contribution decreases from 67%–75% to 46%–52%. We also compute the contribution of the two components to the correlations of real yields across countries. For the overall sample and the first subsample, the EH component is again the dominant component in driving the high and positive correlation between real yields across countries. The term premium is again relatively more important in the second subsample, with the results generally more mixed and somewhat dependent on the country's perspective.

B. Evidence for Slopes

Slopes over a very short time period may be more informative about the economic forces driving interest rates, especially when interest rates show (near) nonstationary behavior. For example, if the short-term interest rate follows a driftless random walk, the term spread equals the term premium. More generally, the large changes observed in the means of nominal and inflation-linked interest rates across subsamples may not affect spreads as strongly.

We replicate all our earlier work on levels for slopes. Our measure of the nominal term spread is simply the 5-year yield minus the one-quarter Treasury bill rate. Furthermore, we assume that the one-quarter nominal yield equals the real yield plus one quarter expected inflation, where the latter is measured from survey data. That is, we assume the Fisher hypothesis holds at the one-quarter horizon, rendering the inflation and liquidity risk premiums 0 at the one-quarter horizon.

Table 11 provides a summary of the results for nominal slopes, computed from the data, and real slopes, which follow from correcting the inflation-linked slopes for the liquidity premium. First, we show the usual statistical properties in terms of averages, standard deviations, and correlations. It is indeed the case that neither nominal nor real slopes show statistically significant changes in means over the two subsamples, although real slopes increase by about 40 basis points in the U.S., and nominal and real slopes decrease substantially in France in the second subsample. We do observe that the standard deviations of both nominal and real slopes decreased over time, and in a statistically significant fashion. As with interest rate levels, nominal slopes are highly correlated across countries, but real interest rate slopes are somewhat less correlated (with the correlations varying between 26% and 70%). Almost invariably, slope correlations decrease over time, mostly in a statistically significant fashion. The real slope correlation between France and the U.S. is negative in the second subsample.

We also show how much our derived real slopes contribute to the variance of the nominal and inflation-linked slopes. We confirm our main result for levels: the real slope dominates. In the U.K., the real rate slope contribution is 93% over the full sample but increases from 85 to slightly over 130% from the first subsample to the second. In the U.S. and France, the real rate slope contribution varies between 73% and 86% over the three samples. The real rate slope also dominates inflation-linked yield slopes, with its contribution being never lower than 68%, which occurs for the U.K. in the first subsample. It exceeds 100% for the U.S. in the second subsample.

In Supplementary Appendix VIII, we further verify the international correlation decomposition, analogous to what we did for levels, finding that the real part of the nominal slope dominates overall correlations over the full sample. For the France–U.S. pair, the contribution of the expected inflation slope is close to that of the real slope, however. The dominance of the real yield slope continues to hold for the cross-country inflation-linked slopes correlations.

C. Additional International Evidence

Many inflation-linked debt markets have relative short histories and/or an insufficient number of bonds to reliably compute zero coupon bonds in the early stages of their development (see Ermolov (2021) for more details). To expand our sample internationally to Germany, Sweden, and Australia, we are forced to start the sample later, in 2011. Thus, this sample roughly coincides with our second subsample during which, for example, correlations between yields decrease. Nonetheless, we can still verify whether our main result, namely, the dominance of real rate variation, holds up for these other countries.

TABLE 11

Annualized 5-Year – 1-Quarter Zero-Coupon Yield Curve Slopes

Table 11 shows results obtained using the monthly data. GMM standard errors, computed using 12 Newey–West (1987) lags, are in parentheses. For subsample 2, *, **, and *** indicate if statistics are different from subsample 1 at the 10%, 5%, and 1% significance levels, respectively.

	Nominal Slope			Real Slope		
	France	U.K.	U.S.	France	U.K.	U.S.
Full sample: 2004M11–2019M12						
Average	0.83% (0.20%)	0.61% (0.27%)	1.00% (0.24%)	0.35% (0.15%)	−0.29% (0.25%)	0.52% (0.22%)
Std. Dev.	0.60% (0.11%)	0.85% (0.16%)	0.76% (0.13%)	0.52% (0.10%)	0.93% (0.10%)	0.77% (0.11%)
$\frac{\text{cov(REAL\_SLOPE,NOMINAL\_SLOPE)}}{\text{var(NOMINAL\_SLOPE)}}$	73.24% (6.80%)	92.99% (12.59%)	80.16% (13.96%)			
$\frac{\text{cov(REAL\_SLOPE,INFLATION-LINKED\_SLOPE)}}{\text{var(INFLATION-LINKED\_SLOPE)}}$				75.41% (5.11%)	72.82% (4.01%)	72.62% (12.40%)
Correlation with U.S.	0.55 (0.20)	0.76 (0.12)	1.00	0.26 (0.19)	0.58 (0.12)	1.00
Correlation with U.K.	0.76 (0.14)	1.00	0.76 (0.12)	0.70 (0.12)	1.00	0.58 (0.12)
Subsample 1: 2004M11–2012M5						
Average	1.10% (0.36%)	0.64% (0.57%)	1.04% (0.47%)	0.47% (0.29%)	−0.29% (0.47%)	0.31% (0.42%)
Std. Dev.	0.70% (0.12%)	1.09% (0.18%)	0.91% (0.15%)	0.63% (0.14%)	1.11% (0.13%)	0.92% (0.13%)
$\frac{\text{cov(REAL\_SLOPE,NOMINAL\_SLOPE)}}{\text{var(NOMINAL\_SLOPE)}}$	78.87% (7.79%)	84.75% (12.49%)	85.34% (11.81%)			
$\frac{\text{cov(REAL\_SLOPE,INFLATION-LINKED\_SLOPE)}}{\text{var(INFLATION-LINKED\_SLOPE)}}$				72.14% (14.93%)	67.55% (14.55%)	68.93% (16.86%)
Correlation with U.S.	0.69 (0.10)	0.76 (0.15)	1.00	0.52 (0.07)	0.65 (0.08)	1.00
Correlation with U.K.	0.91 (0.07)	1.00	0.76 (0.15)	0.86 (0.07)	1.00	0.65 (0.08)
Subsample 2: 2012M6–2019M12						
Average	0.56% (0.18%)	0.58% (0.21%)	0.97% (0.29%)	0.22% (0.06%)	−0.28% (0.29%)	0.73% (0.23%)
Std. Dev.	0.32%* (0.11%)	0.50%*** (0.14%)	0.57%** (0.19%)	0.34%** (0.09%)	0.70%* (0.15%)	0.51%** (0.14%)
$\frac{\text{cov(REAL\_SLOPE,NOMINAL\_SLOPE)}}{\text{var(NOMINAL\_SLOPE)}}$	86.32% (9.67%)	134.12% (12.48%)	72.57% (12.80%)			
$\frac{\text{cov(REAL\_SLOPE,INFLATION-LINKED\_SLOPE)}}{\text{var(INFLATION-LINKED\_SLOPE)}}$				88.08% (14.56%)	103.27% (17.90%)	97.97% (12.35%)
Correlation with U.S.	0.27* (0.22)	0.83 (0.13)	1.00	−0.21*** (0.19)	0.47 (0.23)	1.00
Correlation with U.K.	0.47*** (0.17)	1.00	0.83 (0.13)	0.35*** (0.12)	1.00	0.47 (0.23)

Full estimation details and the results are provided in Supplementary Appendix IX. Note that we identify liquidity premiums in the three new countries separately as we could not assemble a consistent set of liquidity proxies for all three countries. First, we confirm for these countries that the bulk of nominal yield variation derives from the real rate, which accounts for 64% of nominal yield variation in Germany and Australia and over 100% in Sweden. In Germany and Australia, the inflation risk premium is the second most important component. Second, nominal yields are generally highly correlated, with one exception. U.S. yields, during this more recent period, show negative correlation with the yields of these other countries. All other correlations vary between 0.70 (Sweden and the U.K.) and 0.97 (France and Germany). While the latter high correlation is not surprising, note that Australian and German/French yields are also more than 90% correlated. The correlation decomposition shows the real rate consistently to be the main variable behind these high correlations.

VII. Conclusion

We reconsider an important decomposition of nominal bond yields into its real and inflation components in an international context, focusing on the United States, the United Kingdom, and France. We start the sample in 2004, because we want to alleviate the identification problem in the decomposition by using inflation-linked debt. With this period dominated by unusual monetary policies, we primarily focus on the 5-year yield, rather than short term rates. The key finding relative to earlier work is that the roles of expected inflation and real rates have changed. Inflation expectations show little variation and thus variance and cross-country comovement decompositions show that expected inflation accounts for little of the variation in nominal yields. With stable inflation expectations, and moderately variable inflation risk premiums, real rate variation dominates the variation in nominal yields. Real rates correlate highly across countries for most of the sample period, but more recently the correlations have decreased substantially.

We establish these results in a model-free way, using survey inflation expectations and empirical estimates of the liquidity premium. We then confirm that they continue to hold from the perspective of a no-arbitrage Gaussian term structure model, where principal components of nominal and liquidity-adjusted real yields serve as factors (together with the level of inflation).

Supplementary Material

Supplementary Material for this article is available at https://doi.org/10.1017/S0022109022000515.

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