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ABSTRACT

Utilizing critical recent data advances, we analyze empirical evidence on long-run samples of short-maturity real interest rates as well as term spreads based on multi-century data. In contrast to an extensive literature on short-maturity real interest rates over the past few decades, we find strong and consistent evidence of trend stationarity in long horizon series, relatively fast adjustment speeds, and a paucity of structural breaks – results that we show to survive out of sample tests. The use of very long-run data offers a fresh perspective for ongoing monetary policy debates surrounding r*, and also provides a crucial missing link to reconstructing the long-run properties of term spreads. On balance and against limited post-COVID data, our evidence suggests caution on the idea of a break in short-term real interest rate behavior and instead points to elements of continuity over very long time periods. Relatedly, we show that term spreads are secularly rising while inflation volatility trends in the exact opposite direction – a finding questioning the emphasis of influential term structure models.

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1 Introduction

In this paper, we revisit the econometric properties of advanced economies' short-maturity real interest rates as well as term spreads, over the very long run, building on comprehensive advances by financial historians over recent years.[1]

Short-maturity real interest rates are, of course, very closely related to the concept of the underlying "neutral interest rate", "r*", the north star of modern monetary policy. At the same time, term spreads are important across a host of financial issues, not least for the maturity structure of public debt. As we shall demonstrate, a perspective based on centuries instead of decades of data, indeed gives fresh perspectives, for example challenging the consensus view that there is a significant permanent component to short-run real rate movements, and perhaps signalling caution in overstating the widespread belief that the sharp fall in r* after the financial crisis was mostly permanent. Relatedly, we show that centuries long trends in term spreads have been upward rather than downward, again raising questions about how much of the downward shift of recent decades can be expected to last.

Until now, broader types of investigations could simply not be undertaken, given the absence of a consistent long span of short-maturity rates. Seminal work on secular interest rate dynamics, such as Barro (1987), categorically declared: "A usable series on short-term interest rates is unavailable for the full sample". When more recent attempts were made to construct long-run series, they were inconsistent with regards to issuers and maturities – even for key economies such as the United States. The situation has fundamentally changed and this study exploits substantial recent data advances for both nominal rate and inflation series, as well as introducing new data, to leverage the radically improved statistical power that accompanies the use of multi-century, higher frequency, and conceptually consistent time series for the two key "safe economies" over the past centuries – the United Kingdom and the United States – and apply the insights to the major ongoing policy questions. Our data set is not only significantly longer than any previously studied in the literature, it is quarterly instead of annual, and based on consistent data construction.

Besides its basic properties, we are interested in the empirical evidence on structural breaks and half-lives – the time it takes for time series with such trend components to revert halfway back to the trendline: only if such half-lives (and higher order measures of mean reversion) are sufficiently short, and the frequency of structural breaks sufficiently low, are the results on stationarity also economically significant. We show that these properties indeed all apply, and not least we demonstrate that, on the basis of our new long-sample interest rate series, our main points survive an out of sample robustness check.

In an interpretative section following the econometric part, we apply our novel long sample insights to a key topical theoretical and empirical debate of recent years (section 6): arguably, short-maturity

[1] Recent application of long-dated historical data to *long-maturity* real rates (Rogoff et al., 2024), while perhaps suggestive, does not really address the question of how short-run real rates behave, given large and ongoing changes in term spreads, investor preferences and public debt issuance strategies.

real rates are the decisive variable for matters of monetary policy and also from the perspective of many financial intermediaries: we therefore discuss in particular what our re-assessment of real short-maturity interest rate properties implies for the characterization of r*, also known as the "neutral" rate of interest. Assessing the evidence from both long- and short-maturity sections of the term structure, we argue that r* should be modeled as a stationary process – whatever other specific modelling assumptions one incorporates for r*. Equally, the absence of clear evidence for a structural break in or around 1980 carries relevant policy implications, though it is very well conceivable that tests beyond our standard ones may nuance the break narratives.

There is a natural fundamental extension from our new short-maturity data with recent new data on long-maturity rates to derive term spreads over centuries. Though short-maturity debt is the decidedly more recent part of public debt, such a reconstruction using private commercial paper and also central bank discount rates is now possible for selected leading economies for more than two hundred years: in the process, the empirical backdrops afflicting the literature on term spreads can also be addressed, a literature that has equally operated with post-1945 data even when purportedly analyzing long-term structural trends and associations.

Crucially, we find that short-maturity rates have secularly fallen faster than long-maturity rates – resulting in clear evidence of secularly *rising* term spreads. This is surprising against a consensus literature that focused on falling term spreads over the recent decades in the context of the "great moderation" and falling inflation volatility. Indeed, we document that falling inflation volatility across advanced economies is unfolding against a backdrop of secularly *rising* term spreads over time. While we abstain from sweeping generalizations, we propose that, in order to jointly match long-run patterns in inflation volatility and term spreads, important qualifications in the existing literature are required: either term premia remain a close approximation of sovereign default risk – and default risk has in fact risen (in contrast to "safe asset" propositions) – say because financial repression risk has increased; or else, factors other than default risk are long-run determinants of term premia time variation – plausibly liquidity factors.

Overall, our evidence suggests that crucial stylized interest rate facts – including those most important to financial markets and monetary policy – may need to be reconsidered when much longer run data are applied.

2 Literature

Short-maturity real interest rates have been investigated extensively, especially for the U.S. and the U.K., though overwhelmingly they focus on post-1945 or even post-1980 data. A few examples of long sample econometric studies on short-maturity rates include Rose (1988), and Hamilton et al. (2016), with these two being distinct for their attempt to incorporate pre-1945 annual data. Both were unable to reject a unit root in their short-maturity real rate series, echoing results from short sample literature. However, while covering longer samples than related literature, it is not clear that either approach in fact exploited a sufficiently long sample to obtain significant statistical power, and in any case these two and all related "long-term" studies use spliced issuer series that mix together private and public short-maturity rates.[2] (The fact our new series are quarterly also offers the potential to uncover richer short-term dynamics than would be possible with only annual data.)

More recently, a prominent strand of literature is motivated by the construction of "neutral" interest rate series, often dubbed "r*". Prominent contributions, including Holston et al. (2017) and Holston et al. (2023), use evidence on short-maturity real interest rates combined with output growth and output gap estimates, to arrive at estimates of the "natural" rate. As is true for the general literature on short-maturity rates, existing analyses of r* broadly posit unit root behavior for this variable: Holston et al. (2023) model r* as a stochastic process allowing for highly persistent shocks; similarly, Jorda et al. (2022) are representative in their rates analysis, claiming that "the natural rate is a state variable that follows a random walk" (ibid., 169). Equally, Schmitt-Grohé and Uribe (2022) model the U.S. neutral rate as non-stationary over 1900-2021, in line with numerous other contributions. Importantly, while r* itself is unobservable in market data, much of the current literature relates r* to the time series behavior of actual real short-maturity rates (Bauer and Rudebusch, 2020).

Another well-established literature has focused on term structure dynamics as a separate time varying variable, which for our purposes is quite related to the question of whether the short-term real interest rate has very different time series properties than the long rate.

Some key relevant literature uses the concepts "term spreads" and "term premia" interchangeably.[3] However, in particular asset pricing literature often prefers to distinguish between observable term *spreads* on the one hand, which can directly be derived from long- and short-maturity yield differences – and term *premia* on the other hand, a measure that attempts to isolate the component embedded in term spreads that is not explained by changes in expected short-maturity rates over the lifetime of the long-maturity bond. This latter measure requires a variety of model assumptions of the term structure (such as assumptions on affine, spanned, or curvature model features), however, and disagreements over the appropriate model design exist even for contemporary estimates of term premia.[4] Not least because term premia strictly represent a component of term spreads over time, however, any reasonable model of term premia will accord with the econometric and secular empirical features that are established for directly observable term spreads: this is especially so beyond the horizon of very short horizons of several years, which is where different term premium estimates also disagree most among each other.[5] Against this backdrop, we propose

[2] Hamilton et al. (2016), for one, for the U.S. splice together annual commercial paper rates prior to 1913, then switch to Federal Reserve discount rates, in addition to switches in the inflation expectation basis to construct inflation expectations, from CPI to a GDP deflator. All relevant existing studies, including those on long-maturity real rates, in any case use annual level observations, while we also use consistent quarterly data.

[3] For instance Nakamura et al. (2013), who refer to the "term premium" as the difference in real rates of return of 10-year government bonds and those of bills.

[4] E.g. see the comment on Wright (2011) by Bauer et al. (2014), who vary the VAR model estimates and arrive at substantially different risk-neutral and term premium decompositions.

[5] The same discussion by Bauer et al. (2014) stresses that the secular association between inflation uncertainty and

that the conceptually cleanest approach is to focus on directly observable term spreads in all subsequent discussions, without taking a particular stance on term premium models – over the secular horizon that is of interest to us, results for term structure features are applicable to a broad array of derived term spread and term premium variations.

Overwhelmingly, in this literature that associates term structure variation with inflation volatility, the empirical focus has been on post-1945, or even post-1980s data. In representative fashion, Wright (2011) estimated international term premia trends since the year 1990, combining Gaussian model assumptions and survey evidence, and posited a "secular international decline... consistent with a decline in inflation uncertainty". The latter rationale – the association of term structure trends with inflation volatility – builds on an established literature that include, for example, Piazzesi and Schneider (2007) and Rudebusch and Swanson (2008). In the related "rare disaster" literature, equally, time-varying disaster frequency (inflation shocks) are posited to explain variation in the nominal term premium (Gabaix, 2012; Nakamura et al., 2013).[6]

More generally, of course, term spread and term premia have long been an active area of debate in the finance literature – with extensive efforts over the years to provide general frameworks for term structure shapes and curvatures: with much evidence tilted against the traditional "expectations hypothesis", for example, recent work has revolved around "preferred-habitat" (Vayanos and Vila, 2021), or investor and household disagreements over inflation expectations that could impact yield curve patterns (Ehling et al., 2018). But these recent analyses as well as virtually all seminal contributions in this strand of literature use at best post-1945 term structure data, including Campbell and Shiller (1987, 1991), who investigate U.S. term spreads over 1952-1987 – updated in Campbell (2018) for the period of 1961-2013.

Importantly, we do not take any particular stand on any of these debates, but it is important to highlight that what unites these seemingly disparate perspectives is the near-exclusive use of a small sample of term structure data. Whatever the exact preferred model for term structure dynamics, and their precise decomposition, we emphasize that they need to accord with the economically and econometrically relevant features that are, across key dimensions, only unearthed in sufficiently long samples that are constructed in a consistent fashion. The fact that the term structure literature near-exclusively relies on post-1945 data analysis could potentially lead to crucial mis-characterizations of the nature of term spreads. This concern extends by definition to measures that are unobservable directly, but derived from term structure data, importantly including estimates of the term premium.

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the exact decomposition of risk neutral rates and term premia is not affected by the model design. We note that we do not explore co-integration dynamics any further since that would only be of further interest when combining two non-stationary series. Any reasonable model of expected short-maturity interest rates over the long run would have to depart from stationary assumptions, however.

[6]Of course, a separate strand of literature has investigated the predictive power of term spreads for a variety of macroeconomic and financial dynamics, including bond returns and growth measures – literature that we do not directly aim to speak to. Perhaps most famously, Stock and Watson (1989) found that the difference between 10-year U.S. Treasury yields and 1-year U.S. Treasury yields over the period of 1960-1988 outperformed almost every other financial variable in predicting business cycle dynamics.

3 Long-run empirics in short-maturity rates and term spreads, 1704 – 2024

3.1 A short history of short maturity debt

To understand existing stylized facts arising from the predominant use of short-sample data in the literature on the properties of short-maturity rates and term spreads, and the limitations of such approaches, it is useful to briefly revisit the historical background behind short-maturity debt. The issuance and trade in merchant "bills" was known in Europe from the early modern period, when quarterly fairs recorded high turnovers in primary and secondary markets for such instruments, though prior to the early 1700s, the transactions lacked standardization and only occasional data is available. In the sovereign space, rulers do experiment with occasional short-maturity instruments – including *asientos* in Spain, and "debentures" in 17th century England (Gentles, 1969) – but none of these assets assume any consistent form, with the development also being hindered by a much more stringent interpretation of usury laws for short-maturity debt.[7] In private assets, a standardized market in short-maturity debt is gradually developing over the course of the 1600s and by the early 1700s it is standardized and liquid enough to assume a status as a "safe" asset in the key financial markets, including the London bill market (Flandreau et al., 2009). In the U.S., private short-maturity financial activities are first concentrated in and around the major economic centers in Boston, Philadelphia, New York, and Chicago, but by the early 1800s U.S. financial markets on the national level were "innovative, large, and perhaps the equal of any in the world"; the commercial paper market represented one of the regular channels to refinance, and was certainly larger than the market for bankers acceptances over the observation period (Foulke, 1931). These developmental contours allow us to trace data in these two safe financial markets in a consistent way through standardized instruments.

The sovereign space lags the private sector, and U.K. and U.S. Treasury bills are only being issued at regular and meaningful volumes from the 20th century, during and after World War I (Garbade, 2008). This evolution is mirrored in all other advanced economies over time: for instance in Italy, where high-frequency data exists thanks to Francese and Pace (2008), short-maturity debt averages less than 2% of the total consolidated debt stock between 1861-1914. In contrast, regular consolidated long-maturity sovereign debt issuance can be traced back centuries, with Italian city Republics issuing standardized and marketable perpetual debt instruments as early as the 13th century (Pezzolo, 2003; Eichengreen et al., 2021).

[7]While the Church widely tolerates long-maturity debt issuance in a variety of forms – and in fact widely engages in the issuance and trading itself – short-maturity debt is derided as a form of speculative "consumption loans" and therefore viewed much more critically. Nevertheless, ad-hoc "over the counter" short-maturity debt transactions are observed on a regular basis in the early modern period between sovereigns and private creditors. However, these transactions do not rise to the level of marketable instruments, and at no point in time represent a dominant form of sovereign credit. For a sample of hundreds of such "over the counter" short-maturity transactions over the period 1311-1850, and associated interest rate data points, see Schmelzing (2025), where the new long-maturity data for a broader set of geographies is also analyzed.

This lack of observable consolidated sovereign short-maturity yields has often led existing literature to employ central bank policy rates as a direct measure for "short-maturity" public market rates, for instance Hamilton et al. (2016), and others to splice public and private rates of different issuers and maturities (Jorda et al., 2017). However, given newly available data reconstructions, there is no longer any conceptual reason to let such series begin in 1870 or 1890, or to splice together public with private short-maturity rates and thus taper over fundamentally different risk, maturity, and liquidity features. And given both the sufficient liquidity in private short-maturity markets plus the availability of well-regarded higher frequency price indices – at least for wholesale prices – there is equally no methodological reason to restrict observations to annual level series for the two key "safe" financial markets over the past two centuries. At the same time, to the extent the U.K. is concerned, central bank discount rates during the very early years of the Bank's inception (ca. 1694-1820) show very little fluctuation: though it is not clear that the Bank regarded 5% as a legal ceiling for its discount rate, the actual stickiness around this particular level might raise questions: we have regarded the year 1822 (following the resumption of convertibility after the Napoleonic Wars) as the year when the U.K. Bank rate becomes a truly flexible, market-determined rate, and begin series involving the Bank rate at this point in time.[8]

3.2 Leading economy short-maturity real rate and term spread data

We turn to the new data reconstructions that now allow a significant and conceptually clean expansion of interest rate analyses, and potentially new insights into econometric short-maturity and term spread properties.

We discuss the construction and data sourcing from the country level up, focusing on the two key "safe asset" economies that are representative for broader advanced economy secular trends.[9] First, we construct a series for U.S. short-maturity real interest rates that follows a consistent definition of the concept – our series is therefore distinct not just due to its greater sample length and frequency, but also given its conceptual quality as opposed to mixed-issuer series with increasingly obscure origins – such as those used in Hamilton et al. (2016) and Jorda et al. (2017).

In particular, for U.S. short-rates, the widely used "JST database" (Jorda et al., 2017) relies on a website series of Officer (2021): Officer (2021), in turn, records an annual "surplus funds" series that splices call loan rates between 1870-1954 and then switches to Federal Funds rate over 1955-2023 – a dataset that therefore uses neither consistent issuer sectors nor consistent maturities.[10]

[8]For these early years of the Bank, and contemporary monetary policy discussions, see Richards (1964) and Goodhart (2010).

[9]U.S. and U.K. data combined account for close to 50% of advanced economy GDP between 1792-2024, and 26.8% of Global GDP using latest definitions and GDP in the Maddison Project Database ("MPD"), via Bolt and van Zanden (2023).

[10]The U.S. call loan rates refer to secured broker loans without a predetermined maturity date, which Officer himself traces back to a variety of sources including Mitchell (1916). We discuss the JST "STIR" data further, including to other advanced economies, and show econometric results in the Appendix (section 1.2).

In contrast, we present a new conceptually consistent short maturity real interest rate series on a quarterly frequency for the U.S. spanning 1831:Q1 – 2024:Q2, based on the nominal side on 60-90 day maturity prime commercial bill rates. The nominal data is based on averaging monthly level observations in Bigelow (1862), Macaulay (1938), and Board (2024), and we are able to construct matching measures of inflation expectations employing year-on-year quarterly wholesale inflation via Warren and Pearson (1932), and BLS (2024) taking data all the way to the most recent 2024 quarters. We therefore ensure both a consistent issuer and maturity basis, for the highest frequency and covering a period with liquid national markets in these instruments. Figure 1 displays the raw quarterly data, together with a long-run trend component, estimated by the Müller and Watson (2018) approach. Inflation expectations are here constructed by a matching quarterly frequency year-on-year change in the wholesale/PPI price index, per the published sources in Appendix section 1.1. Visually, this series appears to show a downward trend, and often records volatile swings of up to 40% over several years in the first half of the 19th century. It should be mentioned that from the inception dates of our series, at the very least monthly level data points are reported in all underlying series – and at times even multiple observations each month. In all cases, for nominal rates and wholesale price changes, we average these monthly observations to arrive at quarterly data points (rather than taking year-end values, for instance).[11] While we concentrate on the quarterly frequency level here, which provides additional statistical power, we present the same data in annual form, together with results, in the Appendix (Figure A.1). Later, to construct nominal term premia for the United States, we use the new short-term interest rate data presented here together with the long-term interest rate data presented in Schmelzing (2025).[12]

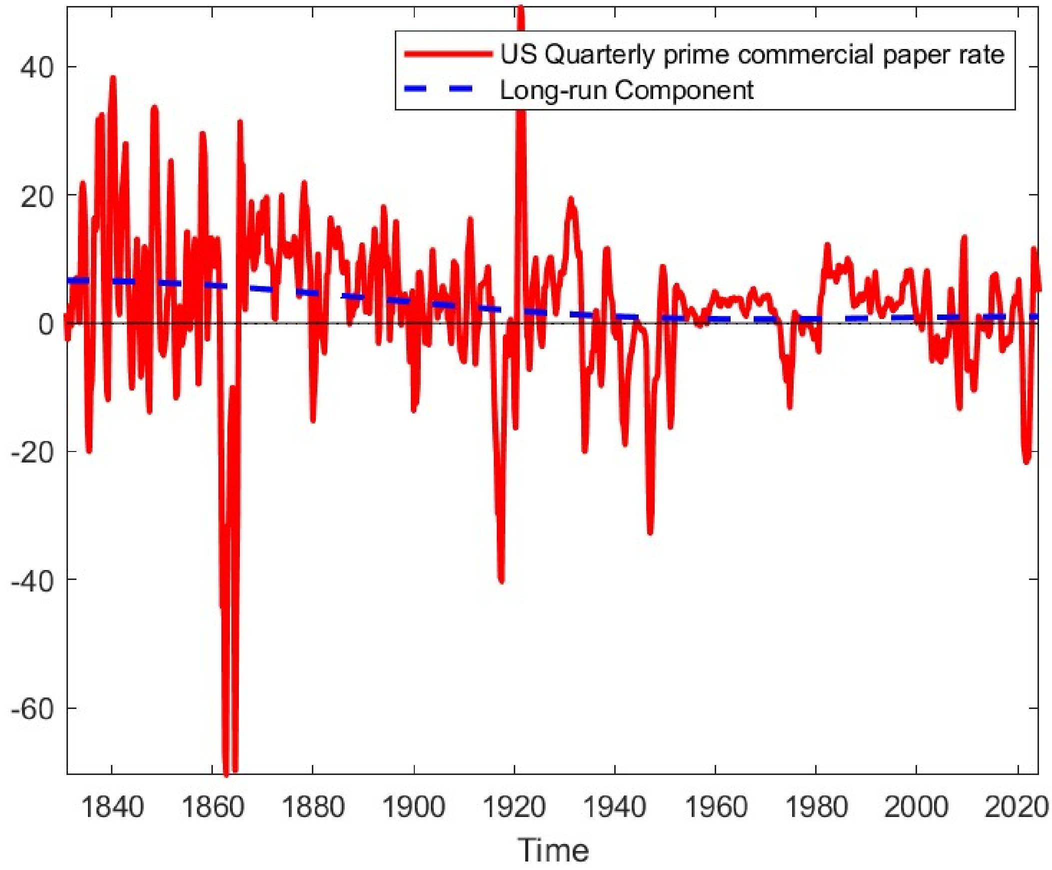
Second, we also analyze high-frequency short-maturity interest rate data for the U.K., constructing what perhaps constitutes the longest real private and real public rate series. On the basis of nominal interest rate and inflation data presented by Dimsdale and Thomas (2016), and straightforwardly extended to 2024, we construct two quarterly U.K. short-maturity real interest rate series spanning 1791:1Q – 2024:Q2, which adopt alternative inflation expectation approaches consistent with existing approaches in the literature. Figure 2 displays one of them, where inflation expectations used to construct the real interest rate are based the quarterly year-on-year change in the benchmark wholesale/PPI index, which is most closely in line with existing methodologies and in line also with the U.S. approach in Figure 1 (with the Appendix providing details on alternative constructions).

Third, again we use the U.K. data compiled by Dimsdale and Thomas (2016) to construct two initial measures of long-sample term spreads, extending their data through 2024 from official sources. Here we shift to nominal spreads to avoid dealing with how to consistently measure any difference

[11]See Appendix section 1.1 for additional details. Wherever multiple monthly level observations exist (e.g. in Bigelow (1862)), we follow the same approach to form a monthly average first. For the U.K., Dimsdale and Thomas (2016) form their quarterly observations which we use by averaging monthly observations, but the underlying monthly observations themselves are at times formed by averaging weekly data points (pre-1870), later by taking month-end values (post-1870).

[12]The long-term nominal interest rate data is used in Rogoff et al. (2024), who draw on Schmelzing (2025), where many related debates are addressed. They do not, however, deal with short-term interest rates, including the new U.S. data employed here.

Figure 1: U.S. real short-maturity interest rates, commercial paper basis, quarterly, 1831:Q1 – 2024:Q2.



Notes: Red line displays U.S. quarterly data based on nominal prime commercial paper rates, 60-90 day maturity. For nominal rates in quarter t=0 we form inflation expectations using quarterly year-on-year U.S. wholesale inflation rate in the same quarter (t=0), with sources including Bigelow (1862), Macaulay (1938), and Board (2024) for rates, and Smith and Cole (1935) for wholesale inflation. The blue line displays the underlying trend component using the methodology of Müller and Watson (2018), selecting fluctuations with periodicity of 100 years or more.

between long-term and short-term inflation expectations, and it also suffices for the main points we are making in this paper. We construct a "private" long-run version of term spreads, as well as a "public" version of U.K. term spreads, as well as a combination that best reflects liquid market pricing across the term structure. Specifically, it is possible to construct at least three respective combinations of the term structure over time for which multi-century consistent yield differentials can be obtained (for reasons of space, most of the subsequent econometric results focus on the second and third variation):

- First, it is possible to construct a quarterly term spread series, using consol rates over prime bill rates. Using the spread between consol rates and the prime bill rate, this version perhaps represents the most high frequency long-run term spread series available, and covers 1753:Q3 – 2024:Q2. It is displayed via Figure 3, which also includes its long-run component (blue dashed line) following the methodology of Müller and Watson (2018).

- Second, we define the annual spread between long-maturity consol yields and the short-maturity Bank of England rate as the measure of the **public term spread**. With consol yields constituting a form of perpetual debt, and the Bank rate constituting the official discounting rate, this term spread construction captures a relatively wide maturity interval on the yield curve.[13]

- Third, we define the spread between private mortgage rates, and short-maturity prime bill rates, as the measure of the **private term spread**. This series runs on an annual basis (as our mortgage rate series is only annual) between 1718 and 2024, and is displayed via Figure 4, which also includes its long-run component (blue dashed line) following the methodology of Müller and Watson (2018).

We can see visually already that for both series in Figures 3 and 4, it appears that the "recent" decline in term spreads since since 1980 (the focal point of key literature) is not necessarily consistent with longer-run trends, which if anything appear to be pointing towards an opposite (increasing) direction. Further, there appears to be an increase in volatility of the two series beginning around the early 19th century, and in absolute terms, the quarterly "public term spread" has also been lower, standing at an average of 0.4%, versus an average for the "private term spread" of 1.1% over the same period. Whether such general observations can be underpinned by econometrics is the focus in the next section.

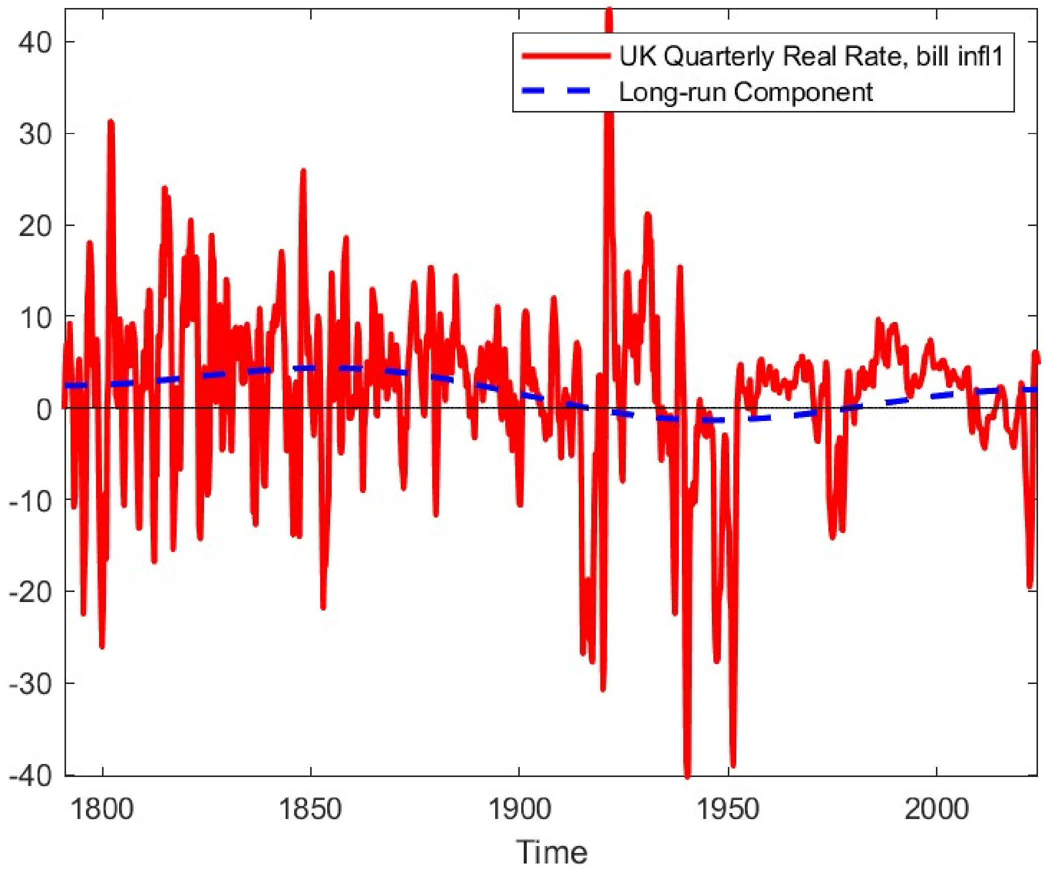
Although most of our focus will be on the nominal spread for reasons stated earlier, we do present a straightforward construction of a U.K. real term spread series spanning 1704-2024 on an annual frequency in the appendix and describe its basic properties, including its long-run component; the main messages are similar to what we find from the nominal series.[14]

Finally, in the Appendix (section 5) we also test different inflation expectation approaches for the construction of real short-maturity rates. Our simplest baseline approach typically involves an adjustment with either the current-year or a single-lagged realized inflation figure. Not least, we specifically present an autoregressive inflation expectation approach there, to replicate a recent methodology to estimate historical inflation expectations in Hamilton et al. (2016), who follow

[13]For this Bank rate series, the benchmark discounting rate prior to 1972 is used. The Bank itself describes the role of this rate as such: "Since Government receipts do not match Government expenditure, the Bank intervened in the markets through the Discount Houses, by issuing and buying Treasury Bills to ensure that the banking system was in balance at the end of the day. [the] Bank Rate...being the rate at which the Bank of England, acting as lender of last resort, would normally lend to members of the discount market against security (e.g. Treasury Bills and eligible bills). It was also a conventional reference point for the rates which the London Clearing Banks paid on deposits and charged on advances". After 1971, Dimsdale and Thomas (2016) use the Minimum Lending Rate 1972-1981, Minimum Band 1 Dealing Rate 1981-1997, Repo Rate 1997-2006, and the Bank Rate, 2006-2016. As with the U.S. series, monthly observations are averaged to create quarterly observations. This series is also available on an annual basis (1704-2024), and is featured in Figure 9 in section 7.

[14]See Appendix section 7 and Figure A.6.

Figure 2: U.K. real short-maturity interest rates, bill basis, quarterly 1791:Q1-2024:Q2.



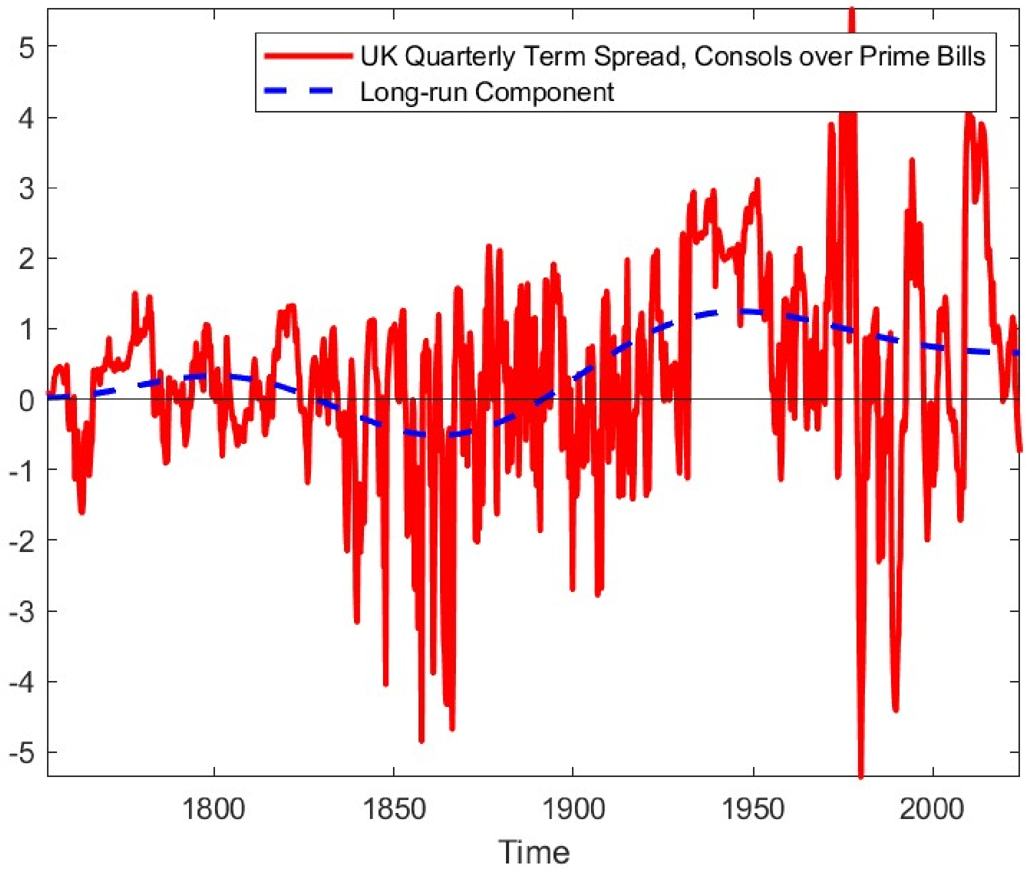
Notes: Red line displays U.K. quarterly real rates obtained via the underlying data in Dimsdale and Thomas (2016), extended to 2024 via Bank of England (2024) and ONS (2024). "Private basis" refers to prime bill rates with consistent maturities of 3-6 months. We adjust the quarterly nominal prime bill rates via a rolling year-on-year change in the wholesale quarterly price index, using the current year values (red line). The blue line displays the underlying trend component using the methodology of Müller and Watson (2018), selecting fluctuations with periodicity of 100 years or more.

Barro (2006): we demonstrate that our results hold for such methodological variations (Appendix tables A.8-10).

4 Statistical Properties

We undertake unit-root tests based on the Augmented Dickey-Fuller (ADF-GLS) approach by Elliott et al. (1996) and also test our series for employing standard structural break tests. Finally, we conclude this section with an investigation of long-run persistence (half-lives) of U.K. and U.S. short-term real interest rates and term spreads. A range of alternative specifications of the ADF-GLS (including a specification assuming no time trend, and using two alternative inflation expectations constructs from existing literature) and related robustness tests (including a Bai-Perron

Figure 3: U.K. nominal term spreads, consols over prime bills, quarterly, 1753:Q3-2024:Q2.



Notes: Nominal quarterly data based on Dimsdale and Thomas (2016) to 2016, and Bank of England (2024) after. "Public term spread" (red) displays U.K. Consols less Bank of England Bank rate. The blue line displays the underlying trend component using the methodology of Müller and Watson (2018), selecting fluctuations with periodicity of 100 years or more.

test) are to be found in the Appendix, broadly confirming our results.[15]

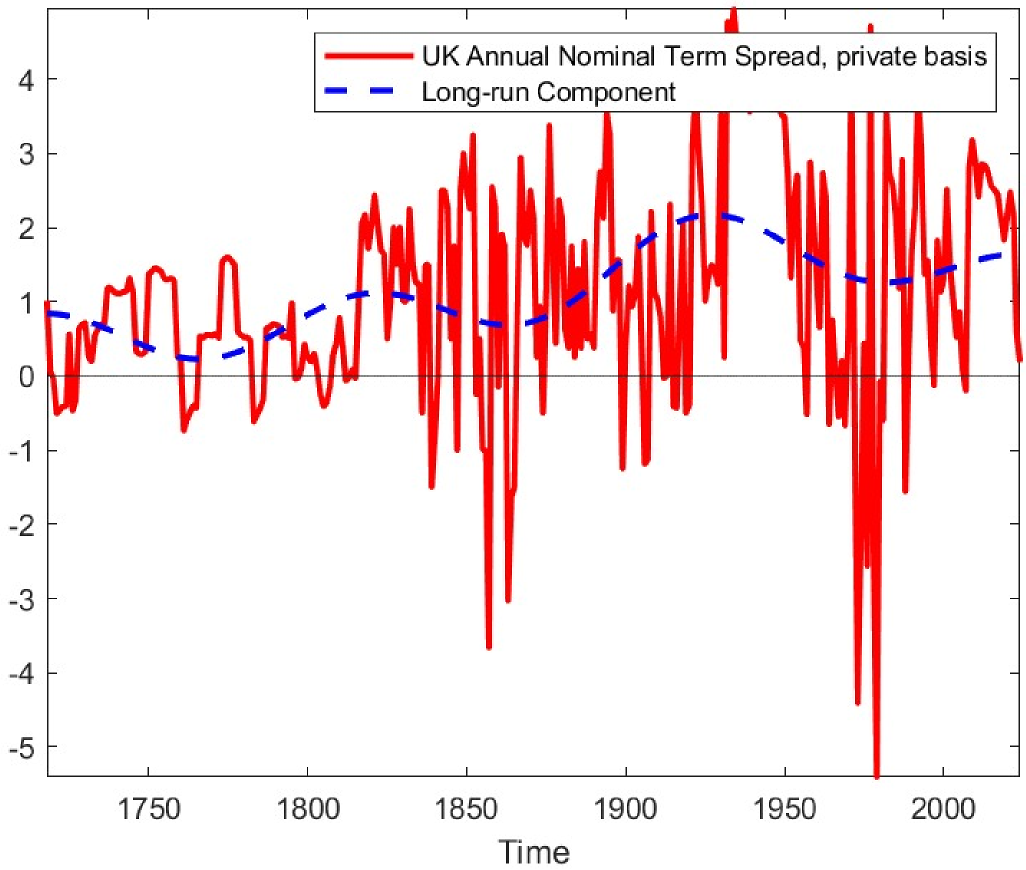
4.1 ADF-GLS and trend stationarity

We start our analysis by testing the econometric properties of short-maturity real interest rate series, as well as a variety of term spread variations. Our guiding question is whether short-term rates, a crucial measure for policy purposes, are stationary and wheter the full term structure exhibits trend stationarity – once multi-century time series are considered, and not just long-term instruments.

To preview our results, the ADF-GLS test finds unambiguous evidence in favor of trend stationarity for both U.S. and U.K. long sample real short-maturity interest rates, and also for multiple term spread variations for these two key "safe" markets over time. In particular,

[15]For constructions without time trends for U.K. and U.S. data, see Tables A.5 and A.9 in the appendix.

Figure 4: U.K. nominal term spreads, private rate basis, annual, 1718-2024.



Notes: Nominal annual data based on Dimsdale and Thomas (2016) to 2016, and extended afterwards via Bank of England (2024). Private term spread basis (red) constructs spread between U.K. mortgage rates less U.K. prime bill rate. The blue line displays the underlying trend component using the methodology of Müller and Watson (2018), selecting fluctuations with periodicity of 100 years or more.

- Per Table 1, all U.K. quarterly data constructions, for both Bank and Bill bases, and for both inflation adjustment approaches, are robustly trend stationary at the 1% level, using our multi-century series (1791:Q1 – 2024:Q2). And as with U.S. series, this result holds for lower level frequencies, including annual.

- Per Table 2, the null hypothesis of a unit root can also be rejected at the 1 percent significance level for both the annual U.K. "Public term spread" (spanning 1704-2024), as well as the "Private term spread" (1718-2024) equivalent.

- Per Table 3, we equally find that quarterly U.S. short-maturity private interest rates over 1831:Q1 – 2024:Q2 are trend stationary: the ADF-GLS test soundly rejects the presence of a unit root at the 1% significance level in real rates (again these are results robust to higher or lower frequencies, including annual or monthly constructions).[16]

[16]Tables 1-3 include a time trend in the estimated model; however, the results are robust to excluding the time trend,

Table 4 confirms stationarity for U.K. and U.S. data when using the Zivot and Andrews (1992) stationarity test (again at a highly significant 1% level, across all series). The Zivot and Andrews (1992) test is robust to the possibility of a structural break even though it is not itself a test for a structural break. Section 5 of the Appendix will discuss more closely the possibility that our series may have undergone structural changes and we will draw from the historical literature for a set of likely break dates to gain deeper insights about the long-term patterns of term spreads over time. While all our structural break tests are standard ones and point towards consistent results – and also span tests where the break date is unknown – we are by no means exhaustive in the full range of possible additional tests which may well nuance our break evidence to a degree.

Finally, the Appendix material confirms that the stationarity results hold for a variety of other specifications, including when no allowance is made for a time trend.[17] In addition, the Appendix (section 6) features a Phillips-Perron test for our headline time series, once again confirming the stationarity results.

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as well as using other observation frequencies such as annual for the U.S.

[17]Per Table A.4, which includes a time trend, we can confirm trend stationarity at the 1% level for both U.S. real rates, as well as for nominal term spread bases – this result is confirmed in Table A.7, where we equally find stationarity at the 1% significance level for real U.S. short-maturity series, as well as for nominal term spreads, all over the period 1831-2023. In the Appendix, we equally obtain trend stationarity when testing ADF-GLS without a time trend, albeit at lower significance levels, see Appendix Table A.5.

Table 1: ADF-GLS Test with Time Trend for U.K. quarterly data			
U.K. quarterly short-maturity real rates, and the nominal term spread.			
	Number of lags	ADF-GLS test statistic	Optimal lag
Bill real, Infl2 (1792:Q1–2024:Q2)	3	-8.817	Seq, SIC
	2	-7.765	
	1	-7.566	MAIC
Bill real, Infl1 (1791:Q1–2024:Q2)	3	-11.347	
	2	-11.687	
	1	-11.901	Seq, SIC, MAIC
Bank real, Infl1 (1822:Q1–2024:Q2)	3	-8.363	MAIC
	2	-9.050	
	1	-9.745	SIC, Seq
Bank real, Infl2 (1822:Q1–2024:Q2)	3	-5.544	Seq
	2	-5.100	
	1	-5.185	SIC, MAIC
Public t-spread (1822:Q1–2024:Q2)	3	-6.564	MAIC
	2	-7.165	
	1	-7.763	SIC

Notes: The table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of 3 lags). The regression includes a constant and a deterministic time trend. The critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -3.480 (1%); -2.890 (5%); -2.570 (10%). "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the Bayesian Information Criterion (SIC), or the Modified Information Criterion (MAIC). The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. For definitions and details on data construction, see the accompanying text. Both the bill and the bank bases matched with two distinct ways of estimating inflation expectations: "Infl1" denotes an approach that estimates inflation expectations for nominal rates in the current year (t) with the matching current-year (t) year-on-year change in the consumer price index; "Infl2" denotes an approach that estimates inflation expectations for nominal rates in current year t through equal-weighted current and previous year year-on-year change in the consumer price index change (that is, the average of the t and t-1 year-on-year index change).

Table 2: ADF-GLS Test with Time Trend

U.K. nominal term spreads, annual, 1704-2024

	Number of lags	ADF-GLS test statistic	Optimal lag
Public t-spread	3	-5.62	Seq
	2	-5.26	MAIC
	1	-6.84	SIC
Private t-spread	3	-6.032	
	2	-6.176	Seq, SIC, MAIC
	1	-7.871	

Notes: The table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of 3 lags). The regression includes a constant and a deterministic time trend. The critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -3.48 (1%); -2.89 (5%); -2.57 (10%). "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the Bayesian Information Criterion (SIC), or the Modified Information Criterion (MAIC). The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. "Public term spread" is defined as the nominal difference between U.K. long-maturity consol yields, and U.K. Bank of England bank rate; "Private term spread" is defined as the nominal difference between U.K. long-maturity mortgage rates, and U.K. prime bill rates. All sources for these rates via Appendix section 1.1.

Table 3: ADF-GLS Test with Time Trend

U.S. real short-maturity rates and nominal term spreads, 1831:Q1 – 2024:Q2

	Number of lags	ADF-GLS test statistic	Optimal lag
Real rates (quarterly)	3	-10.636	
	2	-11.255	
	1	-11.582	Seq, SIC, MAIC
T-spread (annual)	3	-6.373	SIC, Seq.
	2	-5.431	MAIC
	1	-6.018	

Notes: The table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of 3 lags). The regression includes a constant and a deterministic time trend. The critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -3.480 (1%); -2.890 (5%); -2.570 (10%). "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the Bayesian Information Criterion (SIC), or the Modified Information Criterion (MAIC). The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. For the real rate series, nominal rates in quarter t are matched to inflation expectations that are constructed through the current-quarter (t) year-on-year change in the wholesale/PPI price index. "T-spread" denotes nominal spread between long-maturity U.S. sovereign bond yields, and short-maturity 60-90 day prime commercial paper rates.

Table 4: Zivot-Andrews Test Results, U.K. rates and term spreads, and U.S. rates	
	Test statistic
U.K. annual Public term spread nominal	-7.127
U.K. annual Private term spread nominal	-6.370
U.K. quarterly Real Rate Bank Infl 2	-7.609
U.K. quarterly Real Rate Bank Infl 1	-8.271
U.K. quarterly Public term spread nominal	-7.965
U.S. annual term spread nominal	-7.187
U.S. quarterly real short-maturity rates	-8.594

Notes: The table reports the Zivot and Andrews (1992) unit root test statistic, which allows for a break in both the mean and the trend under the alternative. The critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -5.57 (1%); -5.08 (5%); -4.82 (10%). The trimming parameter is 0.10. The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. "Public term spread", "Private term spread", "Infl 1" and "Infl 2" defined as per Table 1.

4.2 Half-lives and higher order convergence measures

In Laubach and Williams (2016), among others, the question of adjustment speeds for the neutral rate is considered. The authors argue that forecasts of mean-reversion following the Great Financial Crisis (GFC) consistently failed to predict realized movements in real short-maturity rates, at least when considering the 5-7 year window after 2009 – therefore supporting the notion that the GFC could represent one of the permanent shocks to which short-maturity rates are subject (the other being the "secular stagnation" shock). Over a longer data sample, however, a failure to detect a clear mean-reversion over a seven-year horizon following the GFC could in general remain entirely consistent with the existence of trend stationary real short-maturity rates. To gauge the situation more precisely, we proceed with a persistence exercise for both U.S. and U.K. real short-maturity rates for the long samples introduced, and for term spreads. Tables 7 and 8 show U.S. and U.K. real short-maturity half-life estimates. Of course, very long adjustment speeds would imply that the finding of trend stationarity remains a conceptually relevant result – but that the economic implications would be of minor economic relevance.

However, per Tables 5 and 6, in fact the adjustment speeds (half-lives) of short-maturity rates and term spreads are generally comparable to the speeds found in long-maturity rates – that is, they are on the order of several years, not decades. In fact, for the U.K. quarterly series adjustment speeds between 1-3 years are generally observable (4-12 quarters). That order of magnitude holds for all our individual constructions – including annual data and including nominal term spreads. We also emphasize that the test, the bootstrap method of Hansen (1999) as in Steinsson (2008), is robust to the presence of unit roots and/or high persistence.

Therefore, the finding of trend stationarity for short-maturity series and term spreads also assumes economic significance, certainly in the context of existing claims in the literature.

Together with the half-life measure, we also report the 3/4 half-life and the 7/8 half-life. The 3/4 and 7/8 half lives are the number of periods after which a shock has decreased its initial impact to 1/4 and to 1/8, respectively. Per Table 5, we observe that our results hold when using these alternative bands, with 3/4 lives generally being recorded at 2-5 year ranges for real rates and 14 years in the last century for nominal term spreads; and 7/8 lives equivalents standing at 2-6 years for real rates, and up to 20 years for nominal term spreads.

Table 5: Half-Lives of Short-Maturity Real Rates and Term Spreads

U.K. quarterly series (half-life in years)								
	α	CI	1/2 -life	CI	3/4 -life	CI	7/8 -life	CI
Real Rate Bill, Infl 1								
1791-2024	0.83	(0.80,0.85)	1.15	(1,1.3)	1.5	(1.3,1.7)	1.73	(1.5,2.1)
1791-1914	0.75	(0.70,0.79)	0.83	(0.7,0.9)	1.1	(0.9,1.3)	1.24	(1.1,1.6)
1915-2024	0.88	(0.85,0.90)	1.58	(1.3,1.9)	1.9	(1.6,2.4)	2.21	(1.8,3.2)
Real Rate Bill, Infl 2								
1792-2024	0.92	(0.90,0.94)	2.31	(1.9,2.9)	3.5	(2.8,4.6)	4.8	(3.7,6.3)
1792-1914	0.87	(0.83,0.90)	1.38	(1.1,1.8)	2.2	(1.7,3)	3	(2.2,4.2)
1915-2024	0.95	(0.93,0.97)	3.4	(2.4,5.2)	4.8	(3.3,7.8)	6.2	(4,10.2)
Real Rate Bank, Infl 1								
1822-2024	0.83	(0.81,0.85)	1.21	(1.1,1.4)	1.5	(1.3,1.7)	1.73	(1.5,2.1)
1822-1914	0.71	(0.66,0.76)	0.76	(0.7,0.9)	0.96	(0.8,1.1)	1.1	(1,2)
1915-2024	0.88	(0.85,0.90)	1.56	(1.3,1.9)	1.9	(1.6,2.4)	2.2	(1.8,3.2)
Real Rate Bank, Infl 2								
1822-2024	0.92	(0.90,0.94)	2.43	(1.9,3.1)	3.7	(2.8,4.8)	4.9	(3.7,6.5)
1822-1914	0.84	(0.8,0.88)	1.15	(0.9,1.5)	1.84	(1.4,2.5)	2.53	(1.8,3.6)
1915-2024	0.95	(0.92,0.96)	3.26	(2.3,4.9)	4.5	(3.1,7.1)	5.8	(3.7,9.4)
Nominal term spread, consols over bank rate (public)								
1822-2024	0.87	(0.83,0.90)	1.02	(0.7,1.4)	2.5	(1.9,3.4)	3.9	(2.9,5.4)
1822-1914	0.73	(0.64,0.81)	0.38	(0.3,0.5)	1.06	(0.6,1.6)	1.91	(1.3,3)
1915-2024	0.91	(0.88,0.94)	2.12	(1.5,3.2)	3.6	(2.3,5.7)	5	(3.1,8.3)

Notes: The table reports median unbiased estimates and 90% confidence intervals of α based on Hansen (1999)'s grid-bootstrap as well as median unbiased estimates and 90% confidence intervals of: the half-life (column labeled 1/2-life), the (3/4)-life (column labeled 3/4-life) and the (7/8)-life (column labeled 7/8-life) based on Steinsson (2008). The regression is: $y_t = \mu_0 + \mu_1 t + \alpha y_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta y_{t-j} + \varepsilon_t$ where α is the largest root. The table reports both the full sample as well as sub-sample estimates. The number of lags is the same as in Tables 1-3 (MAIC). The median unbiased estimates of γ_j in the full sample are: Real Rate Bill, Infl 1: 0.44; Real Rate Bill, Infl 2: 0.34; Real Rates, Bank Infl 1: 0.48; Real Rates, Bank Infl 2: 0.35; Nominal Term Spread, Public: -0.007,-0.051,-0.056. The half-life is estimated as the first horizon at which the impulse response equals one-half of the initial impact effect (similarly, the 3/4- and 7/8-lives are estimated as the first horizon at which the impulse response equals one-quarter and one-eight of the initial impact effect, respectively). The half-life can be a non-integer number (similarly for the 3/4- and 7/8-lives). The half-lives are separately estimated in each sub-sample. "nom" = nominal; "CI"=confidence interval. "2y-w" refers to nominal rate adjustment via quarterly year-on-year wholesale inflation rates, averaging eight observations over the current and most recent past year; "1y-w" refers to nominal rate adjustment via year-on-year change in the quarterly wholesale price index, using rolling figures. All data sources per section 3.2 and further details in Appendix section 1.1.

Table 6: Half-Lives of Short-Maturity Real Rates and Term Spreads								
U.S. annual series (half-life in years)								
	α	CI	1/2 -life	CI	3/4 -life	CI	7/8 -life	CI
Real rate, annual								
1831-2024	0.49	(0.39,0.61)	0.77	(0.66,0.94)	2.1	(1.1,2.9)	3.28	(2.5,4.9)
1831-1914	0.30	(0.15,0.53)	0.57	(0.51,0.74)	0.85	(0.76,2.5)	2.45	(0.9,4.3)
1915-2024	0.63	(0.52,0.74)	2.06	(1.61,2.80)	2.94	(2.3,4.3)	3.68	(2.8,6.1)
Nominal term spread, annual								
1831-2024	0.63	(0.52,0.75)	0.85	(0.72,1.34)	3.04	(2.02,4.9)	5.2	(3.5,8.6)
1831-1914	0.58	(0.40,0.81)	0.68	(0.56,1.28)	2.64	(0.9,6.4)	4.7	(2.7,12.7)
1915-2024	0.74	(0.62,0.87)	2.49	(1.78,4.74)	4.2	(2.7,9.7)	6.2	(3.6,15.1)

*Notes:*The table reports median unbiased estimates and 90% confidence intervals of α based on Hansen (1999)'s grid-bootstrap as well as median unbiased estimates and 90% confidence intervals of: the half-life (column labeled 1/2-life), the (3/4)-life (column labeled 3/4-life) and the (7/8)-life (column labeled 7/8-life) based on Steinsson (2008). The regression is: $y_t = \mu_0 + \mu_1 t + \alpha y_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta y_{t-j} + \varepsilon_t$ where α is the largest root. The table reports both the full sample as well as sub-sample estimates. The number of lags is the same as in Tables 2-4 (MAIC). The median unbiased estimates of γ_j in the full sample are: U.S. real: -0.14076; U.S term spread: -0.21222, -0.02931. The half-life is estimated as the first horizon at which the impulse response equals one-half of the initial impact effect (similarly, the 3/4- and 7/8-lives are estimated as the first horizon at which the impulse response equals one-quarter and one-eight of the initial impact effect, respectively). The half-life can be a non-integer number (similarly for the 3/4- and 7/8-lives). "CI"=confidence interval. For annual data 1-H values are used for the 2024 value.

5 Long sample short-maturity real rates and implications for r*

Conceptually, the real short-maturity rate as analyzed thus far is closely related to the concept of the "neutral interest rate", or r*. Alternatively referred to as the "equilibrium real rate", it describes the real short-maturity interest rate consistent with full employment and real output growth at potential (Holston et al., 2017; Kiley, 2020). r* – unlike the real interest rate – cannot be straightforwardly constructed – which is often constructed so that it coincides with the observed real interest rate in the long run (e.g. Bauer and Rudebusch (2020)). Of course, r*, in almost any model, is by construction a smoother version of r, so it must automatically inherit the trend and stationarity properties of r. True, the half life of shocks to r* can be potentially be longer than shocks to r, given that the latter is more volatile. But it cannot be too different from the one to three years we find for r, given that r – r* essentially fluctuates at business cycle frequencies, although the shocks to r – r* could be longer lasting in the case of a pandemic of financial crisis. But even then the difference would be a matter of years, not decades or generations.

Yet until now, the literature on r* has typically modeled it as a nonstationary series, which makes sense if one accepts the existing consensus in the literature that r itself is nonstationary (see Laubach and Williams (2016), DelNegro et al. (2017), Jorda et al. (2022), or Schmitt-Grohé and

Uribe (2022)).[18] Needless to say, our evidence clearly supports the notion that shocks to r* are transitory, rather than permanent – and that half-lives (and higher-order measures of convergence) are relatively short, making this observation also economically significant. In the very prominent applied monetary and macroeconomic debates on whether the GFC and more recently COVID have accounted for a new short-rate interest rate regime, on balance and against limited post-COVID data our evidence would warrant caution on the idea of a break and clearly leans towards emphasizing the elements of continuity over very long time periods.

It goes without saying that the stationarity of r is also very important for models of asset pricing, but we will not delve into that here.

One might ask whether our results on the stationarity of r (and by implication r*) survive more challenging specification tests such as out-of-sample robustness tests (see Rossi (2013) for a review of the literature in the context of empirical exchange rate models). We undertake such an analysis next.

6 Forecasting r and r*

Having focused on the in-sample properties, we proceed to demonstrate that the stationarity features identified in short-maturity rates also survive out-of-sample robustness tests, a relatively high bar to pass in the context of standard exercises. In the short-run, of course r and r* are different, but long-run forecasts should be very similar for any reasonable model of r*.

We consider two classes of models. A first class is "stationary" models, that is models that do not have a unit root component; we compare the stationary models with a second class of models that are "non-stationary", that is, models that impose a unit root. At each point in time, the models are re-estimated in a real-time fashion using a rolling window procedure, mimicking what a forecaster would have done in real time, using only past information to forecast future values.[19] Below, we describe in detail each of the models that we consider.

Stationary Models. The class of stationary models includes autoregressive models as well as a deterministic trend model. The *Autoregressive model* is:

$$E_{t-1}r_t = \alpha_0 + \alpha_1(t/T) + \beta_1 r_{t-1} + \beta_2 r_{t-2} + ... + \beta_p r_{t-p},$$

where E_{t-1} denotes expectation based on information available at time $(t-1)$.[20] We consider two versions of the autoregressive model: one version (labeled 'ARp-trend') includes the deterministic

[18] Laubach and Williams (2016) consider arguments in favor of a stationary modeling of the neutral rate in the U.S. post-war context, and show that it could still be consistent with a trend decline in neutral rates, but on balance find the evidence in favor of stationarity unconvincing. In DelNegro et al. (2017), the authors posit that a DSGE model with a stationary neutral rate assumption captures low frequency movements in the short-term – but that a VAR model with non-stationarity provides the appropriate fit for long horizons.

[19] The window size used for estimation is half of the total sample size.

[20] The number of lags of the Autoregressive model, p, is determined recursively by the Bayesian Information Criterion. The h-step-ahead forecast is obtained by iterating h-steps-ahead.

trend (as in the above equation) while the second version (labeled 'ARp-no trend') does not include the deterministic trend (that is, $\alpha_1 = 0$.) The *deterministic trend model* (labeled 'trend') only includes a deterministic trend (that is, it sets $\beta_1 = ... = \beta_p = 0$.)[21]

Non-stationary Models: The competing class non-stationary models includes the *Random Walk* (labeled 'RW'), which is:

$$E_{t-1}r_t = r_{t-1},$$

as well the *Unit root model*, which is an autoregressive model that imposes a unit root:

$$E_{t-1}r_t = \alpha_0 + \alpha_1(t/T) + r_{t-1} + \beta_2\Delta r_{t-1} + ... + \beta_{p-1}\Delta r_{t-p+1}$$

We consider two versions of the unit root model: one version (labeled 'Unit root w/trend') includes the deterministic trend (as in the above equation) while the second version (labeled 'Unit root') does not include the deterministic trend (that is, $\alpha_1 = 0$.)[22] We compare the models using the Diebold and Mariano (1995) test for equal predictive ability in the Mean Square Forecast Error sense, comparing each of the stationary models to each of the non-stationary ones.[23]

Figure 5 begins with the U.K. Bill Rate horse race, comparing forecasting accuracy over a forecasting horizon of 1 to 40 quarters. At all forecasting horizons from 8 quarters onwards, the best stationary model beats all alternative models (random walk, unit root and unit root with trend). The Appendix, section 2, reports results for the remainder of series, which all confirm our basic forecasting results here for alternative series constructions.

It is also crucial to note that our quest is not to find the best forecasting model for r. Rather, it is to see if the out-of-sample results are consistent with the stationarity properties (or lack thereof) we find for in-sample r. Several papers have focused on out-of-sample forecasts of r* using complex models that involve non-stationary specifications, e.g. Holston et al. (2017, 2023). However, the latter analyses are based on small samples, while we demonstrate that r* is almost certainly stationary once we rely on longer samples and hence have more powerful in-sample tests. Therefore, the fact that stationary models forecast r* better than unit root models in an out-of-sample race provides additional support to our main result.

Our results have several important implications for policymakers. First, they directly provide a benchmark estimate of medium to long-run r* that does not involve estimating complex and potentially time-varying models. Second, they show that modeling r* as a stationary variable provides not only better in-sample fit (as per our ADF-GLS results) but also better out-of-sample forecasts.

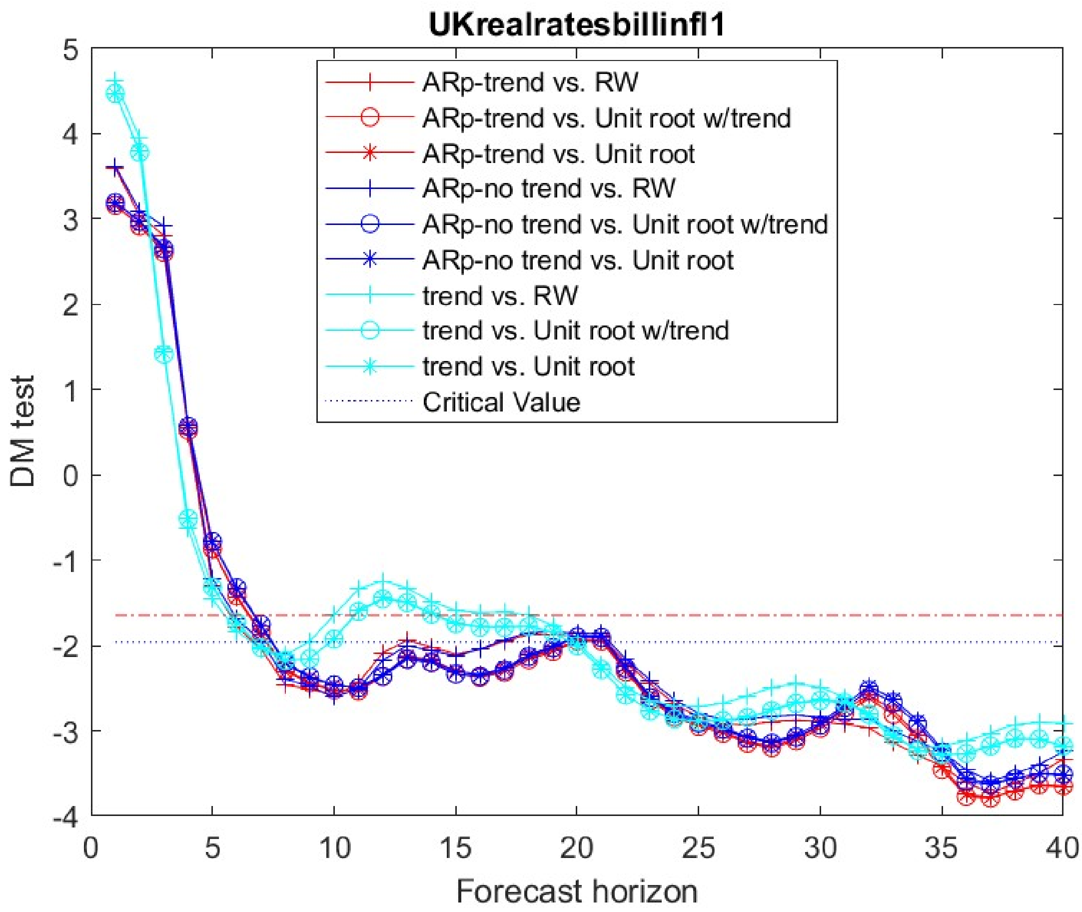
Figures 6 and 7 show that our results are similar for alternative constructions, such as the 2-year equal-weighted inflation adjustment for quarterly U.K. Bill rates (Figure 6), as well as the quarterly nominal term spread series since 1753 (Figure 7).

[21]Thus, the deterministic trend model is: $E_{t-1}r_t = \alpha_0 + \alpha_1(t/T)$.

[22]For comparability, the number of lags of the unit root model is the same as that of the Autoregressive model.

[23]The critical values are based on Giacomini and White (2006).

Figure 5: Out-of-Sample Forecast Comparisons: U.K. Quarterly Real Rate, Bill Infl 1, 1972-2024.

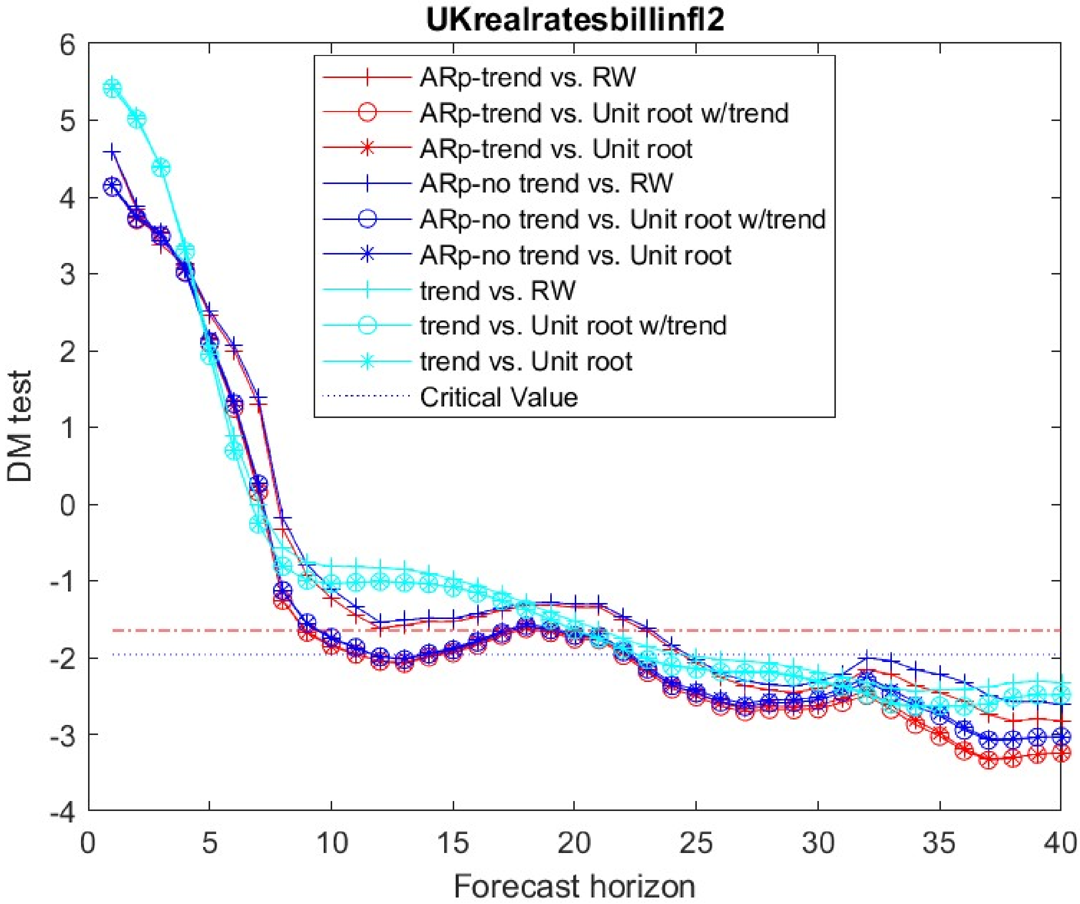


Notes: The picture reports the Diebold and Mariano (1995) test for equal predictive ability for stationary vs. non-stationary models, described in the legend. Negative values indicate that the stationary model has a lower mean squared forecast error than the non-stationary model. The picture also reports the critical value based on the Giacomini and White (2006) approximation (one-sided is -1.645 (dotted line) and two-sided is -1.96 (dashed-dotted line)). Forecasts are obtained in rolling windows with size equal to half of the total sample size.

7 Term spreads, inflation uncertainty, and rare disaster dynamics over time

The time-varying nature of term premia is a salient observation in the literature. Indeed, this represents one of the key reasons that the econometric properties found for long-maturity rates cannot be assumed to hold for short-maturity rates. However, there is remarkably little evidence on the secular stylized facts of this time variation going beyond a few most recent decades. Our reconstructions of longer horizon term spreads already raises fundamental questions on whether the existing consensus alleging a "secular decline" in term spreads over the long run is actually representative (Wright, 2011). We want to explore the long-run dynamics in term spreads more systematically in this section, and test a key stylized fact that is prominent in finance and macroeconomic literature, namely that nominal term spreads are closely linked to inflation

Figure 6: Out-of-Sample Forecast Comparisons: U.K. Quarterly Real Rate, Bill Infl. 2, 1792-2024.



Notes: The picture reports the Diebold and Mariano (1995) test for equal predictive ability for stationary vs. non-stationary models, described in the legend. Negative values indicate that the stationary model has a lower mean squared forecast error than the non-stationary model. The picture also reports the critical value based on the Giacomini and White (2006) approximation (one-sided is -1.645 (dotted line) and two-sided is -1.96 (dashed-dotted line)). Forecasts are obtained in rolling windows with size equal to half of the total sample size.

uncertainty.

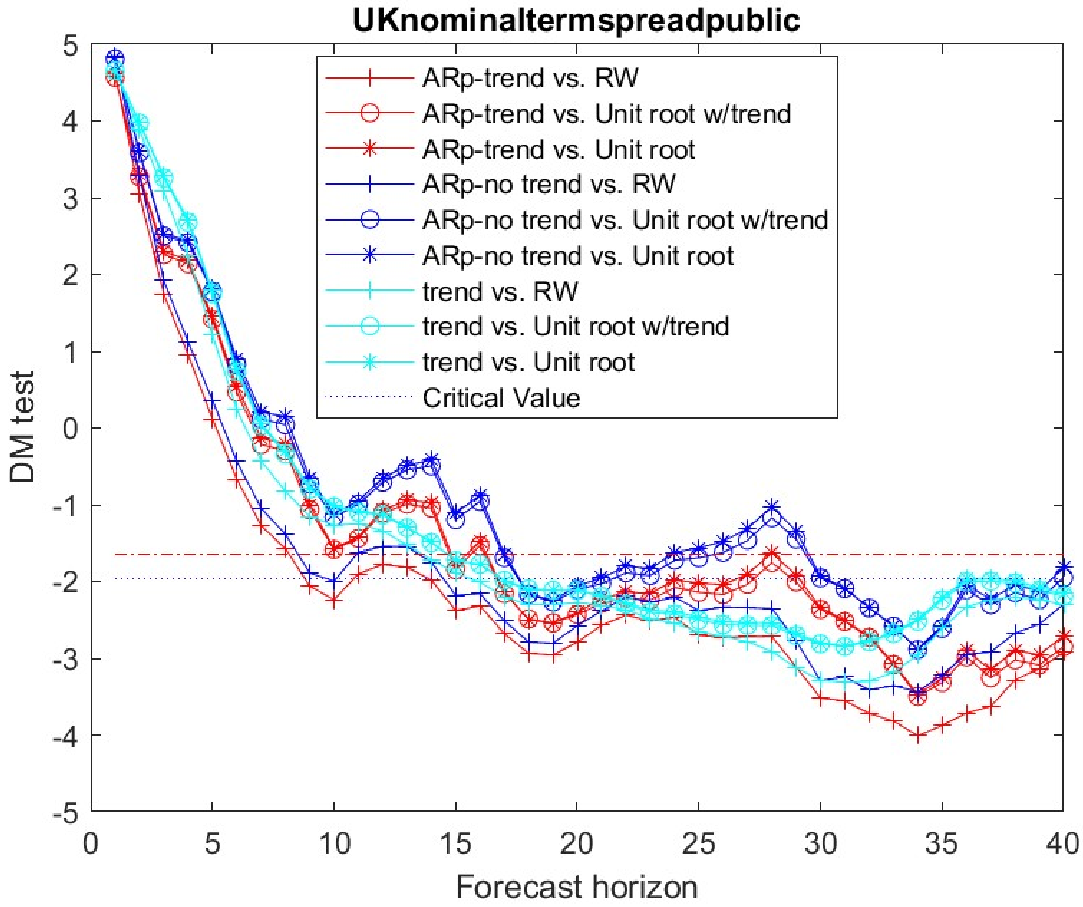
7.1 Are nominal term spreads upwards sloping?

Characteristically, on the basis of this post-1980 U.S. and international data, Wright (2011) has highlighted a "secular" decline in nominal term premia over 1990-2009.

Before we proceed with a more nuanced long-run discussion, we explore the long-run context of this apparent post-1980s decline, empirically and through an econometric/regression exercise. First, we construct annual level nominal term spread series for these two "safe" economies over time empirically via Figure 8, data that we then analyze econometrically.

Indeed, Figure 8 confirms the visual upwards drift of nominal term spreads for the two key safe economies over the past several centuries (here constructed on annual frequencies). In combination

Figure 7: Out-of-Sample Forecast Comparisons: U.K. Quarterly Nominal Term Spread, Public, 1753:Q3-2024:Q2.



Notes: The picture reports the Diebold and Mariano (1995) test for equal predictive ability for stationary vs. non-stationary models, described in the legend. Negative values indicate that the stationary model has a lower mean squared forecast error than the non-stationary model. The picture also reports the critical value based on the Giacomini and White (2006) approximation (one-sided is -1.645 (dotted line) and two-sided is -1.96 (dashed-dotted line)). Forecasts are obtained in rolling windows with size equal to half of the total sample size.

with our Figures 1 and 2, it therefore appears that rising secular term spreads imply meaningful differences in the downward slopes of short- and long-maturity nominal rates, with the short-maturity rate slope being much sharper. Specifically, while we observe that all maturity points trend in the same direction (downwards) over time – it becomes apparent that the decline in short-maturity rates is meaningfully steeper over the past centuries compared to long-maturity rates: in the case of the U.S., the long-maturity real interest rate trend fall since 1831 stands at 1.75 basis points per annum, while the short-maturity real interest rate fall stands at a much steeper 4.06 basis points per annum since 1831; for the U.K., the equivalent drop in long-maturity real rates since 1704 stands at 1.01 basis points per annum, while for short-maturity real rates it stands at 1.81 basis points per annum. Potentially, this faster short-maturity decline could be explained by a greater increase in liquidity in short-maturity debt (public and private), though we observe that this steeper short-maturity drop also holds for variations that take central bank overnight rates on

Table 7: Trend Estimates, U.K. and U.S. annual nominal term spreads.		
	μ_1	Confidence Interval
U.S. spread, annual		
1831-2024	0.024	(0.020, 0.029)
U.K. spread, private		
1718-2024	0.005	(0.003, 0.007)
U.K. spread, public		
1822-2024	0.013	(0.009, 0.016)

Notes: The table reports regression coefficients together with their 95% confidence intervals. The regression is: $spread_t = \mu_0 + \mu_1 t + \varepsilon_t$. For U.K., then "public" term spread is used as defined in section 3.2; for data sources see Appendix section 1 and notes to Figures 3 and 8.

the short-maturity section of the curve.

Whatever the precise driver, it is this faster decline over time of short rates ("bull steepening") – and not *rising* long-maturity rates ("bear steepening") – that account for the observed secular increase of term spreads.

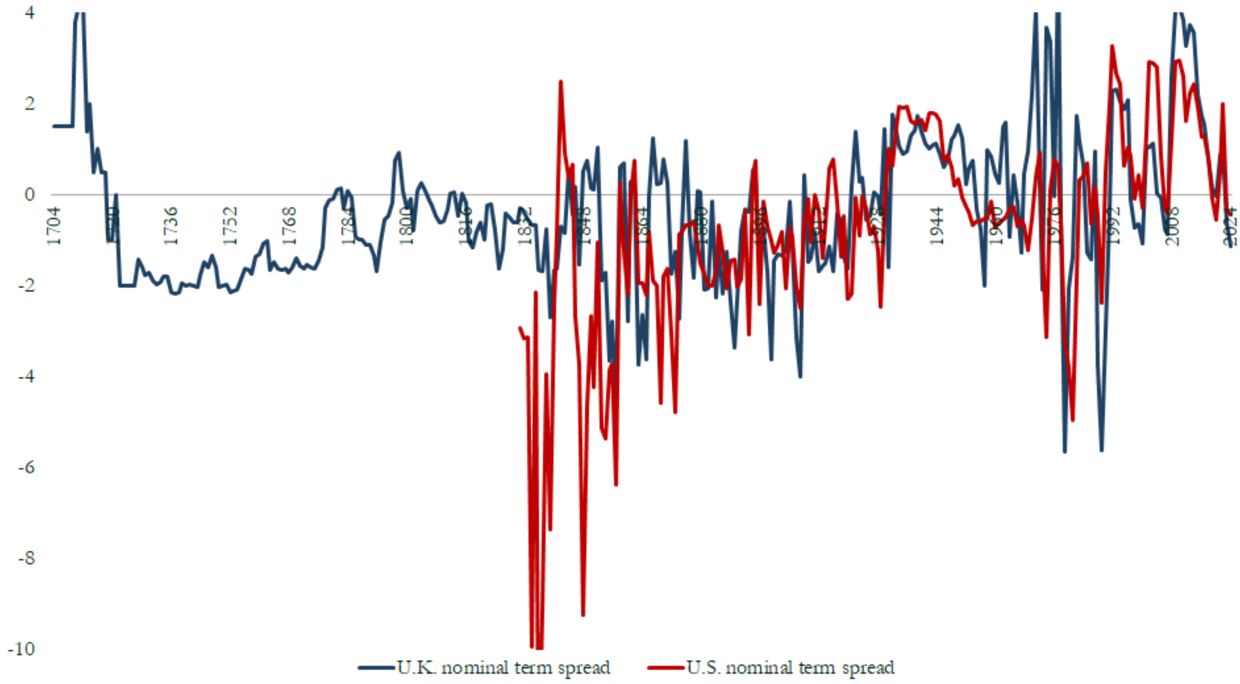
Before we proceed, can we confirm these visual impressions statistically? Indeed, table 7 conducts trend estimates for our U.K. and U.S. term spread data over the respective long horizons. Indeed, we can confirm through this exercise that the visual impressions in Figure 8 are holding: both U.K. and U.S. term spreads are secularly upwards trending – the thesis of an alleged "secular" international decline in term spreads, in other words, appears to be fully reversed once long horizons are properly integrated.

7.2 Term spreads and inflation uncertainty

We now link term spread evidence to broader macro and finance theses. In existing discussions, it is widely accepted that the decline in term spreads is driven by declining inflation uncertainty – in fact even where strong disagreements over the precise decomposition of term spreads into term premia and risk-neutral rates exist, the association with inflation uncertainty is agreed upon (Wright, 2011; Bauer et al., 2014): "bond risk premia mainly reflect uncertainty about future inflation; inflation erodes the value of a nominal bond in precisely those states of the world in which investors' marginal utility is high. In any of these models [Piazzesi and Schneider (2007); Rudebusch and Swanson (2008); Campbell et al. (2017)], reducing inflation uncertainty, then, ought to lower term premia".

In Piazzesi and Schneider (2007) and Miller et al. (2022)'s terminology, among others, it is "rare disaster" inflation shocks that explain the equity premium, but are also consistent with an upward sloping nominal yield curve in both advanced and emerging economies over the long run. Prominently, Gabaix (2012) makes the link explicit. A rare disaster raises default

Figure 8: U.K. and U.S. nominal term spreads, 1704/1831-2024.



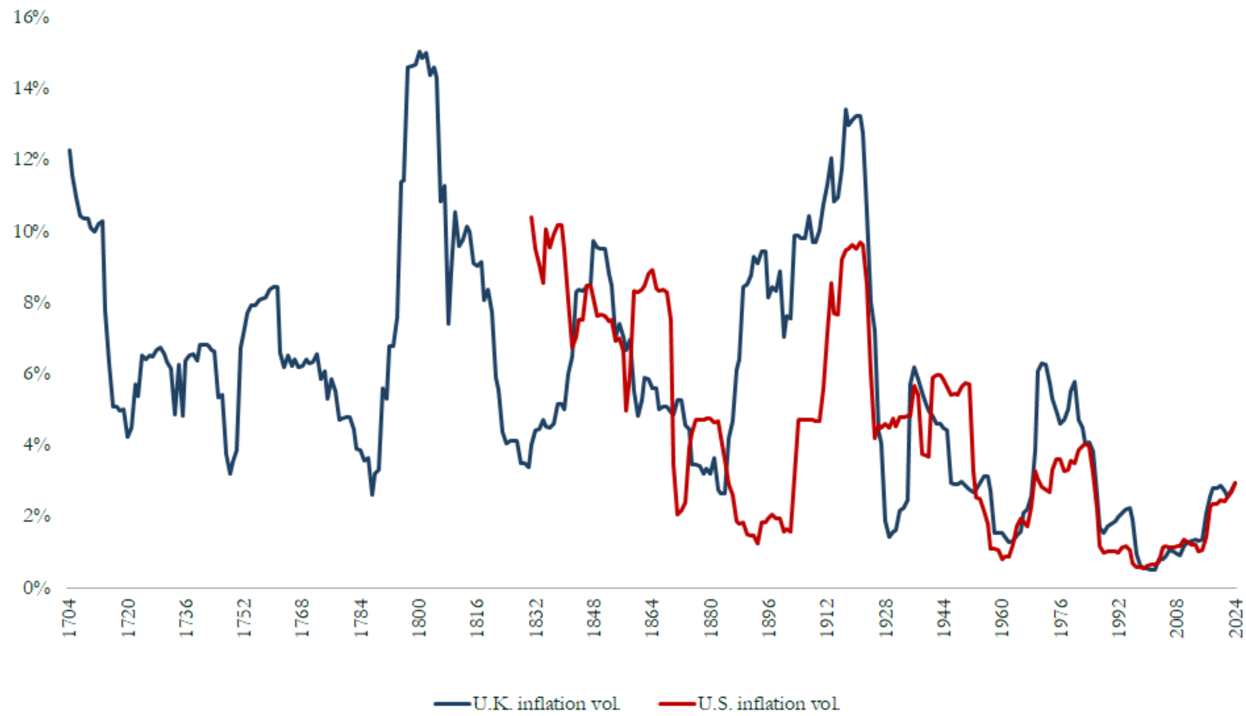
Notes: The figure displays the nominal term spreads, non-smoothed, annual basis. For the U.K., the series uses the "Public term spread" as defined in table 1, and for the U.S. the term spread as defined in table 3. The U.K. series starts in 1704, the U.S. series starts in 1831, equivalent to the inception dates for the inflation volatility series in Figure 8.

risk – expressed as inflation risk – of sovereigns: in the presence of such default (inflation) risk, long-maturity bonds are seen as a riskier asset compared to short-maturity bonds, since the latter allow creditors to regularly reprice the bond in the context of a slow-moving inflation acceleration. Long-maturity bonds would therefore carry a premium compensating investors for the additional risk. In these standard models, secularly declining rare disaster risk should be associated with a secular decline in the nominal term spreads and premia. Via Gabaix (2012): "Since very short-term bills are essentially immune to inflation risk and long-term bonds lose value when inflation is higher, long-term bonds are riskier, so they yield a higher risk premium. Thus, the yield curve slopes up. Moreover, the magnitude of the surge in inflation is time-varying, which generates a time-varying bond premium". Such theoretical evidence appears consistent, for instance, with data from previous U.S. war events, which were disproportionately refinanced via the inflation channel (Hall and Sargent, 2023).

Thus following this literature, Figure 9 displays annual level U.S. and U.K. inflation volatility over the long-run, letting all series begin at the start date of our associated term spread series (1704 for the U.K., 1831 for the U.S.). We observe that inflation volatility on this measure – where we use a simple rolling window of a 11-year centered standard deviation of annual year-on-year changes in CPI – has trended within a wide range in the early part of the sample, and reached peak values of above 10% in both countries, even under gold standard conditions. But trend-wise, the series both appear to point *downwards*, and both reach all-time lows by the early 21st century, amid "Great

Moderation" dynamics. A partial reversal takes place amid the COVID inflation acceleration in both countries, but it remains moderate in the long-run context. Consistent with such longer-term evidence, Miller et al. (2022) observe that option prices have not reflected any secular increase in rare disaster risk, but that data from index-linked Gilts indicate an economically large decline in inflation risk from at least the 1980s. The decline in inflation volatility over the 20th century is echoed by a decline in other financial volatility measures such as exchange rate volatility, though not in equity volatility (Ilzetzki et al., 2020). We do incorporate the most recent inflation surge in the wake of the pandemic: while it does appear to qualify as an above-median inflation shock in the Barro and Ursua (2008) disaster range, and inflation volatility visibly rises, there is no obvious indication that it upsets the more general secular trends described.[24]

Figure 9: U.K. and U.S. inflation volatility, 11-year centered average, 1704-2024.



Notes: The figure displays the volatility of inflation, as defined by the standard deviation of the realized annual year-on-year change in the consumer price index, as rolling 11-year centered average. The U.K. series starts in 1704, the U.S. series starts in 1831, equivalent to the inception dates for the term spread series.

On the contrary, as we saw, Figure 8 displayed nominal term spread trends for the same two economies: in both cases, the series display a clear *upward-sloping* trend over the long-run, a visual finding that can be underpinned by statistical testing. In other words, a falling inflation and default risk trend is in fact co-existing with a secularly rising premium for long-maturity bond yields in the two key "safe" economies over recent centuries. While term spreads in advanced economies did indeed fall when just isolating the period of 1990-2010 (Wright, 2011), and continued to fall

[24]In Barro and Ursua (2008), the typical disaster raises inflation by a median of 2.4% above the normal long sample rate of 4.2%. That median is topped in both U.S. and U.K. annual CPI data over 2020-2.

thereafter, they remain far higher than long-run averages.[25] However, our structural break analyses (in the appendix) indicate that there is little intuitive reason to begin time series in the 1980s or 1990s: the most recent break detected in any nominal U.S. or U.K. term spread series is 1963.

Visually speaking, following Figure 8, the upward slope is particularly sharp between ca. 1850-1914, for both the U.S. and the U.K., with a subsequent era of lower and more volatile term spreads, lasting into the 1980s. Especially since that latter date, nominal term spreads appear to have "picked up" their pre-1914 tendencies again, and once more resumed an upwards drift. The contradictions between term spreads and inflation volatility are confirmed for various key historical episodes. For instance, it appears that the conclusion of the U.S. Civil War – the first of three major inflation volatility events over the past two centuries in the U.S., next to World War I and the 1980s – did not lead to a repricing of inflation risk: on the contrary, falling inflation volatility in the U.S. following 1865 was mirrored by a secular increase in term spreads from around -3%, to around 0% at the eve of World War I.[26]

How can this key empirical inconsistency possibly be reconciled? We cannot provide a definitive explanation, but for one, of course, the structure of U.S. public debt has evolved substantially over the past 200 years, including since World War II. For one, the average maturity of government debt held by the public has declined meaningfully since 1941, with the average maturity declining from close to 10 years initially, to just over 3 years by 2009. Second, more and more long-maturity debt is held by the non-public, the most important entities of which are the Social Security Funds and the Federal Reserve. The average maturity of government debt assets held by the non-public is several years higher than the average maturity of government debt assets held by the public.

Combined, with the rise in inflation-protected debt,[27] this secular development implies two features for financial risks for public government debt holders: first, they are more and more short-maturity oriented, and thus less inflation sensitive; second, even investors on the long-maturity spectrum are less inflation sensitive, with more and more of them holding inflation-protected instruments. With regular consolidated short-maturity debt issuance only beginning in earnest during the late 19th and early 20th centuries, short-maturity assets by construction acquired a shorter "track record" to establish credibility, and build a structural investor base. While short-term debt transactions occurred regularly between sovereign debtors and private creditors, save for exceptions such business fully resembled "over-the-counter" dealings.

[25]Per FRED data, using monthly frequency, in June 2023 the spread between the U.S. 10-year Treasury bond yield and the 2-year equivalent reached the lowest level since August 1981. In January 2011, the spread reached the highest level over at least five decades, at 2.84%.

[26]In line with such observations, despite the fall in inflation risk, nominal total return volatility of U.S. long-maturity bonds has been secularly increasing over the nominal total return volatility of short-maturity assets. An absolute difference between long-maturity and short-maturity total returns should be expected in the presence of general inflation risk (a fact confirmed empirically for U.S. post-1945 data via Hall and Sargent (2011)) – however, secular trends in realized inflation volatility would lead one to expect a *narrowing* of the relative volatility.

[27]Alongside the rise of debt held by the non-public the rise in inflation-protected instruments can be observed: in the U.S, TIPS instruments (exclusively issued in maturities of 5, 10, and 30 years) constituted less than 2% of aggregate public debt stocks in the mid-1990s, but have risen to account for close to 7% as of 2024, see Hilscher et al. (2022).

Liquidity factors are against this backdrop an important plausible driver of the steepening, though we note that steepening trends appear to be already present before the formal introduction of consolidated debt instruments (Schmelzing, 2025), and are in any case also present when one uses Bank rates, for instance, as a short-maturity component.

Indeed, it is known that the share of short-maturity debt as a percentage of total general government debt outstanding has risen sharply over the long run: for instance, for Italy, Francese and Pace (2008) constructed perhaps the most granular breakdown of debt decomposition by maturity over more than 150 years with consistent definitions. The data confirms that the end of the traditional gold standard in 1914 set off an unprecedented surge in the (relative and absolute) issuance of short-maturity debt – with another wave following in the late 1930s, in conjunction with World War II. During the post-war decades, the share of short-maturity debt remains highly elevated compared to the pre-1914 peacetime periods; however, from the all-time peaks reached in the early 1980s (July 1982: 64.4%), the relative importance once more declines. In that sense, relative and absolute short-maturity liquidity increased sharply over a 150-year horizon – and more dramatically than long-maturity equivalents. While the partial reversal since the 1980s (an observation not unique to Italy[28]) requires further rationalization in the context of a continued increase in term spreads, this data suggests that liquidity factors may be able to explain at least part of the apparent contradiction between term spreads and inflation risk – given that these spreads are driven by sharply falling short-maturity rates, in the presence of less sharply declining long-maturity rates.

Overall, we do not attempt to offer any robust "solution" to these empirical contradictions in long-run data. However, we posit that our observation that a secular term spread trend – driven by aggressively falling short-maturity rates – has been underway over centuries, is new and significant. We highlight that the long-run evidence is only consistent with one of two hypothesis, both of which appear to require qualifications in existing debates.

- The first possible conclusion is that markets may have internalized the notion that it is increasingly hard to "inflate away the debt" (Hilscher et al., 2022) and thus decoupled debt pricing from inflation risk – but that such default risk is in fact rising secularly, driven by sources other than inflation; the fall in long-maturity rates is thus stickier than that in short-maturity rates.
- Alternatively, it means that rising term premia over time are primarily reflecting factors other than default risk considerations – such as relative liquidity across the term structure or shifting investor preferences –, and that sovereign default risks are indeed secularly falling.

Indeed, there appears to be at least some evidence in support of greater uncertainty over advanced economy debt sustainability over longer horizons (Rogoff and Schmelzing, 2024), compared to a

[28]According to Treasury direct data, the share of short-maturity bills as a percentage of total marketable U.S. debt stood at just under 18% by June 2023, but stood at just under 23% at the beginning of the century, having risen to levels above 30% during the 2008-9 financial crisis.

more benign recent consensus. On the contrary, it is relevant to point out that related measures of political risk at least also indicate secularly falling risk (e.g. Brecke (2020), who documents a steady decline among advanced economies in conflict intensity over 1400-2000, including the period ca. 1700-2000 under consideration here). In any case, we note that the opposing linkages between both term spreads and inflation risk precede the most recent rise of inflation-protected bonds, as well as rises in the share of government debt held by the non-public.

Two lessons appear to emerge from the new long-run term premia evidence: first, the need to far better rationalize the time variance in inflation and rare disaster risk, as the empirics appear to contradict stylized predictions. Second, the observation that it is a secular "bear steepening" – driven by the faster decline in short-maturity yields – that is unfolding in financial markets over time, and accounts for rising term premia. This stands in contrast to post-1980s analyses, and possibly the role of (relative) liquidity is a promising variable to link to this particular shape of the term structure over the past centuries.

8 Conclusion

This paper has provided a first conceptually consistent long-sample investigation of advanced economy short-maturity real interest rates and term spreads – utilizing significant data improvements in recent years. It departs from the observation that the econometric properties of key financial and macroeconomic time series – including interest rate series – may change dramatically once long samples are utilized. The new data is designed to overcome conceptual drawbacks in existing spliced series (they include significant issuer and maturity switches), and allows one to explore a much longer time period than was possible previously.

On a practical level, our results on the short-term real interest rate are especially relevant for the debate on the properties and structural evolution of the neutral rate, "r*", with virtually all key literature assuming that r* is subject to very persistent or permanent shocks (such as 1981 or 2008). Evidence from long-run data puts doubts on such assumptions, and the qualifications suggested by our results are not just conceptually but also economically significant, with short half-lives and very few (if any) structural breaks. Of course, r* (like the term premium) is an unobservable variable, and does not necessarily overlap identically with r in the short-run – but whatever particular construction approach is chosen, it is clear that such r* properties need to adhere to the trend and stationarity features uncovered over our new long samples for actual short maturity rates.

For financial market research, as well as public debt management, our results on the term structure are also important. Indeed, our evidence on inflation volatility and term spreads represents an excellent exhibit for employing newly available historical time series to give fresh perspectives on fundamental long-standing questions. Namely, we have used our new long-sample interest rate data to flag a relevant conceptual contradiction in the existing literature: we documented that in long-run data, a declining long-run inflation volatility (sovereign default risk) in fact coincides with a secularly *rising* term spread. Canonical existing literature in macroeconomics and asset pricing predict the exact opposite – including both "rare disaster" debates, and those focused on the "great moderation" dynamics in the international economy. In contrast, our long run view implies that either falling inflation volatility is not the primary approximation of sovereign default risk – and that such default risk is in fact rising secularly – for instance, because of rising financial repression risk; or alternatively, it means that rising term spreads and term premia over time are primarily reflecting factors other than default risk considerations, and that default risks are indeed secularly falling. While we do not definitively resolve this dichotomy in the present paper, and are agnostic about the precise decomposition estimates of term spreads into risk-neutral rates and term premia (a decomposition that cannot be observed), the serious possibility that the first alternative may hold should perhaps lead one to pause, in the context of assumptions of almost mechanically structurally improving safe asset debt sustainability.

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APPENDIX

1. Details on data sources, robustness tests, and alternative econometric specifications

.1 Details on long-run short-maturity real rates and term spreads

Here we elaborate on our data construction details.

- For the U.S., on the **rates side**, we begin nominal short-maturity observations based on the data of Bigelow (1862), who reports between 1-3 monthly observations for prime commercial paper rates in Boston starting in 1831. We arithmetically average all monthly observations to arrive at a quarterly nominal short-maturity rate series. There is one single year in Bigelow (1862) where data is missing for several months (February-November 1834) – here we linearly interpolate between January and December.
- After 1856, we switch to Macaulay (1938, A.142-160), who also reports monthly frequency commercial paper rates, 60-90 day maturity for New York City, all the way to January 1937. As with the other inputs, we average monthly observations to arrive at quarterly (or annual, for the respective exercises) data points.
- After December 1936, we switch first to NBER Macrohistory data (to December 1971, series m13002, NY prime commercial bills 60-90 day maturity), then to the consistent Board (2024) series for AA non-financial prime commercial paper rates, 60-90 day maturity. Again, the underlying observation frequency is at the monthly level, and we average monthly observations to arrive at quarterly (annual) data points. Some of the reported NBER and Fed data from this period, which these authors report at the monthly level, is in itself based on averaging daily offering rates of dealers.
- On the **inflation side**, we use consistent U.S. wholesale index changes, on a monthly reporting frequency, to obtain year-on-year change. These monthly changes are then averaged for quarterly (annual) data points, consistent with our nominal rates approach. Source-wise, we use the classic Warren and Pearson (1932) wholesale price index, which is constructed on the monthly level, up to December 1913. From January 1914, we switch to the harmonized Bureau of Labor Statistics (BLS) Producer Price Index for all commodities, which is equally constructed on the monthly level, and for which we equally average monthly year-on-year changes in the index to arrive at quarterly (annual) inflation rates.

For the U.K., the following equivalent sources are used.

- On the **rates side**, Dimsdale and Thomas (2016), sheet Q.1 "Discount rate on prime short-term paper" is used, and for this consistent construction the authors in turn rely on splicing data in Weiller and Mirowski (1990), Nishimura (1971), and Capie and Webber (1985) to 1970, and afterwards use Bank of England data on eligible bills, to 4Q-2016. For our extensions all the way to 2Q-2024, we equally use the subsequent Bank of England releases for short-maturity (<1 year) non-financial weighted rates and Bank rates (series CFMBI79 and IUQABEDR), monthly underlying observation frequency that we average to form quarterly data points.
- As mentioned previously, for the U.K., Dimsdale and Thomas (2016) form their quarterly observations that we use by averaging monthly observations, but the underlying monthly observations themselves

are at times formed by averaging weekly data points (pre-1870), later by taking month-end values (post-1870).

- On the **inflation side**, for quarterly wholesale inflation rates, we equally use Dimsdale and Thomas (2016), sheet Q.1, who in turn construct their consistent series on the basis of Gayer et al. (1953), Klovland (1993), and ONS: their series runs until 4Q-2016, and is based on averaged monthly observations to yield quarterly (annual) data points, consistent with the approach for the U.S. From 1Q-2017, we add the continued wholesale price change observations from ONS (2024), all the way up to Q2-2024.

.2 Alternative inflation, test specification, frequency, and data approaches

We present a range of robustness tests for our benchmark series – testing alternative observation frequencies, inflation expectation approaches, and stationarity tests that do not assume a time trend, among other robustness tests.

First, Figure A.1 displays our main U.S. short-maturity real rate series over 1831-2024 on the annual basis, rather than quarterly, using the same exact sources. We show in the associated tables below that all our main results hold when the frequency is changed to this level, and can be visually confirmed, the general trends are identical. These identical features can also be confirmed econometrically, however, not least across ADF-GLS and Zivot-Andrews tests.

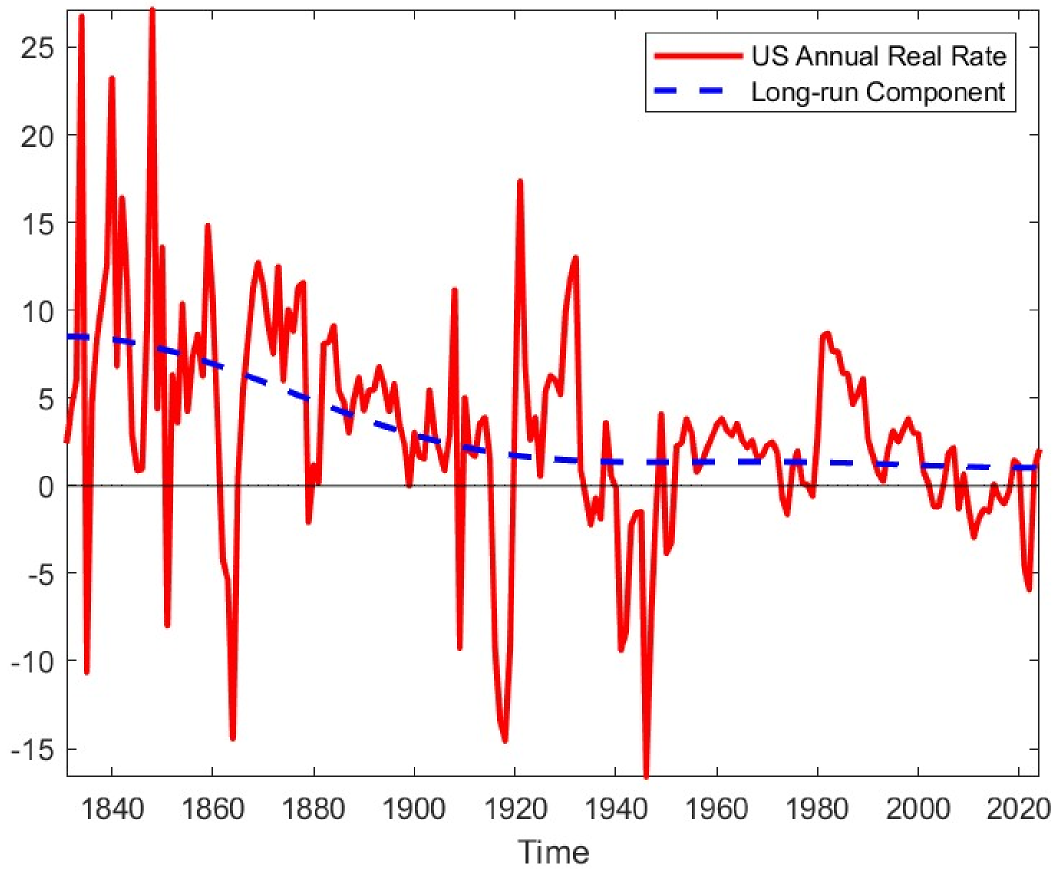
Next, while we highlighted a number of conceptual concerns with the data, we here conduct ADF-GLS tests (with time trend) using "JST"'s short-maturity interest rate series over the period 1870-2020 (Jorda et al., 2017). Tables A.1 and A.2 display the results. We note that across all country-level series, as well as the two "global" series, we confirm stationarity at the 1% significance level.

The only two exceptions are the two series for which multiple annual observations are missing in JST: Spain (missing: 1937-9), and France (missing: 1944-7). When we split these two countries two-way – before and after the gaps – we can no longer reject unit roots in the more recent halves of the samples: France 1945-2020, and Spain 1940-2020. While we do not include these specific countries in our new sample, this evidence is supportive of the more general notion that using shorter sample lengths for term spread data can lead one to over-confirm unit root properties – when in fact short-rates and term spreads are trend stationary.

The JST database offers a "STIR" (Short-term interest rates) and a "LTIR" (Long-term interest rates) annual series for a variety of advanced economies. For "STIR", the database uses an amalgamation of definitions and assets to splice together nominal headline series, so it is critical to elaborate on some representative examples.

- For Germany, JST splice open market rate quotations in Berlin over 1875-1879 with open market discount rates in Homer and Sylla (2005) over the period 1880-1913. For 1924-1940 money market rates in Morawietz (1994) are then used.
- For Italy, Weidenmier and Neal (2003) is used between 1870-1 and 1885-1914, recording private short-term rates, followed by T-bill rates between 1966-8 in Homer and Sylla (2005), subsequently money market rates over 1969-77, and afterwards a switch back to T-bill rates over 1978-2020 occurs, based on IMF IFS data.
- For the United States, surplus funds yields are used via the website of Officer (2021): the latter, in

Figure A.1: U.S. real short-maturity interest rates, consistent concept basis, annual, 1831-2024.



Notes: Red line displays U.S. annual data based on sources including Bigelow (1862), Macaulay (1938), and Board (2024), measuring consistent 60-90 day prime commercial paper rates, adjusted with current-year year-on-year U.S. inflation rate via David and Solar (1977) and BLS (2024). The blue line displays the underlying trend component using the methodology of Müller and Watson (2018), selecting fluctuations with periodicity of 50 years or more.

turn, traces call loan rates in this series over 1870-1954, and then switches to Federal Funds rates.[29] In this sense there is a conceptual switch involved, from a private sector short-maturity rate, onto a public sector rate. In addition, U.S. call loan rates have no predetermined maturity date, complicating efforts to consistently capture a constant short-maturity basis.

For most series, therefore, the JST database not just uses inconsistent maturity definitions, but also splices series from different issuer sectors, with potentially very different risk profiles during certain periods.

As with the Dimsdale data, we also undertake a Zivot-Andrews test for the JST data series, via table A.3. Seven out of the ten individual issuer series are once again confirmed as trend stationary at the 1% level. The one full exception is Japan – which is confirmed only at the 10% level. Once more, France and Spain cannot be confirmed as trend stationary once split two-way given the World War-related data gaps.

[29] Officer (2021) refers to multiple sources as his own underlying source, including for the call loan basis Mitchell (1916).

Table A.1: JST term spreads: ADF-GLS Test with Time Trend, 1/2

	N. lags	ADF-GLS test statistic	Critical Values 1%	5%	10%	Optimal lag
Global Series						
Global AW	3	-3.710	-3.519	-2.979	-2.689	
	2	-3.858	-3.519	-2.979	-2.689	
	1	-4.423	-3.519	-2.979	-2.689	SIC,MAIC
Global GW	3	-3.067	-3.519	-2.979	-2.689	
	2	-3.266	-3.519	-2.979	-2.689	MAIC
	1	-3.652	-3.519	-2.979	-2.689	SIC
Individual Countries						
Italy	3	-4.128	-3.537	-2.994	-2.704	
	2	-4.277	-3.537	-2.994	-2.704	
	1	-4.327	-3.537	-2.994	-2.704	SIC,MAIC
UK	3	-4.712	-3.519	-2.979	-2.689	Seq
	2	-4.215	-3.519	-2.979	-2.689	MAIC
	1	-5.475	-3.519	-2.979	-2.689	SIC
Germany	3	-3.533	-3.519	-2.979	-2.689	MAIC
	2	-4.105	-3.519	-2.979	-2.689	
	1	-4.756	-3.519	-2.979	-2.689	SIC
US	3	-7.903	-3.519	-2.979	-2.689	Seq
	2	-7.738	-3.519	-2.979	-2.689	MAIC
	1	-9.522	-3.519	-2.979	-2.689	SIC
Japan	3	-2.306	-3.530	-2.988	-2.698	MAIC
	2	-2.756	-3.530	-2.988	-2.698	
	1	-3.070	-3.530	-2.988	-2.698	Seq,SIC
Holland	3	-4.897	-3.519	-2.979	-2.689	
	2	-5.036	-3.519	-2.979	-2.689	MAIC
	1	-6.154	-3.519	-2.979	-2.689	SIC*

Notes: Ng and Perron's Sequential criterion selects zero lags for the following series (test statistic in parenthesis): Global AW (4.876), Global GW (-4.270), Germany (-6.099), Italy (-5.263).

Table A.2: ADF-GLS Test with Time Trend 2/2

	Sample	N. lags	ADF-GLS test statistic	Critical values 1%	5%	10%	Optimal lag
France	1870-1943	3	-3.227	-3.679	-3.113	-2.818	
		2	-3.048	-3.679	-3.113	-2.818	
		1	-3.216	-3.679	-3.113	-2.818	SIC,MAIC
	1948-2020	3	-2.110	-3.683	-3.116	-2.821	
		2	-2.372	-3.683	-3.116	-2.821	Seq,SIC, MAIC
		1	-3.375	-3.683	-3.116	-2.821	
Spain	1870-1936	3	-4.223	-3.705	-3.136	-2.839	Seq,SIC
		2	-2.315	-3.705	-3.136	-2.839	MAIC
		1	-2.777	-3.705	-3.136	-2.839	
	1940-2020	3	-2.168	-3.652	-3.091	-2.797	Seq,MAIC
		2	-3.048	-3.652	-3.091	-2.797	
		1	-3.559	-3.652	-3.091	-2.797	SIC

Notes: For France, 1870-1943, the lag length chosen by the Ng and Perron's sequential criterion is zero and the corresponding test statistic is -3.099. Term spreads are constructed by subtracting the "STIR" series from the "LTIR" series in JST (Jorda et al., 2017), and definitions therein.

Table A.3: Zivot-Andrews Test Results: JST Term Spread

Series	Sample	Test Statistic	
Global AW	1870-2020	-5.799	1914
Global GW	1870-2020	-5.326	1893
Italy	1870-2020	-6.344	1922
UK	1870-2020	-6.655	1979
Germany	1870-2020	-8.859	1932
US	1870-2020	-8.135	1931
Japan	1870-2020	-4.939	1957
Holland	1870-2020	-7.753	1914
France	1870-1943	-5.042	1914
	1948-2020	-4.923	1994
Spain	1870-1936	-5.780	1880
	1940-2020	-3.898	1982

Notes: The table reports the Zivot-Andrews unit root test allowing for break in both intercept and trend. Critical values at the following significance levels are: 1%: -5.57; 5%: -5.08; 10%: -4.82. (For Holland, the missing value in 1945 was imputed using the average of adjacent observations both before and after the missing value, using a centered window of six observations; for France and Spain, given the sequence of missing values around World War Two, we separately tested in the two sub-samples before and after the war).

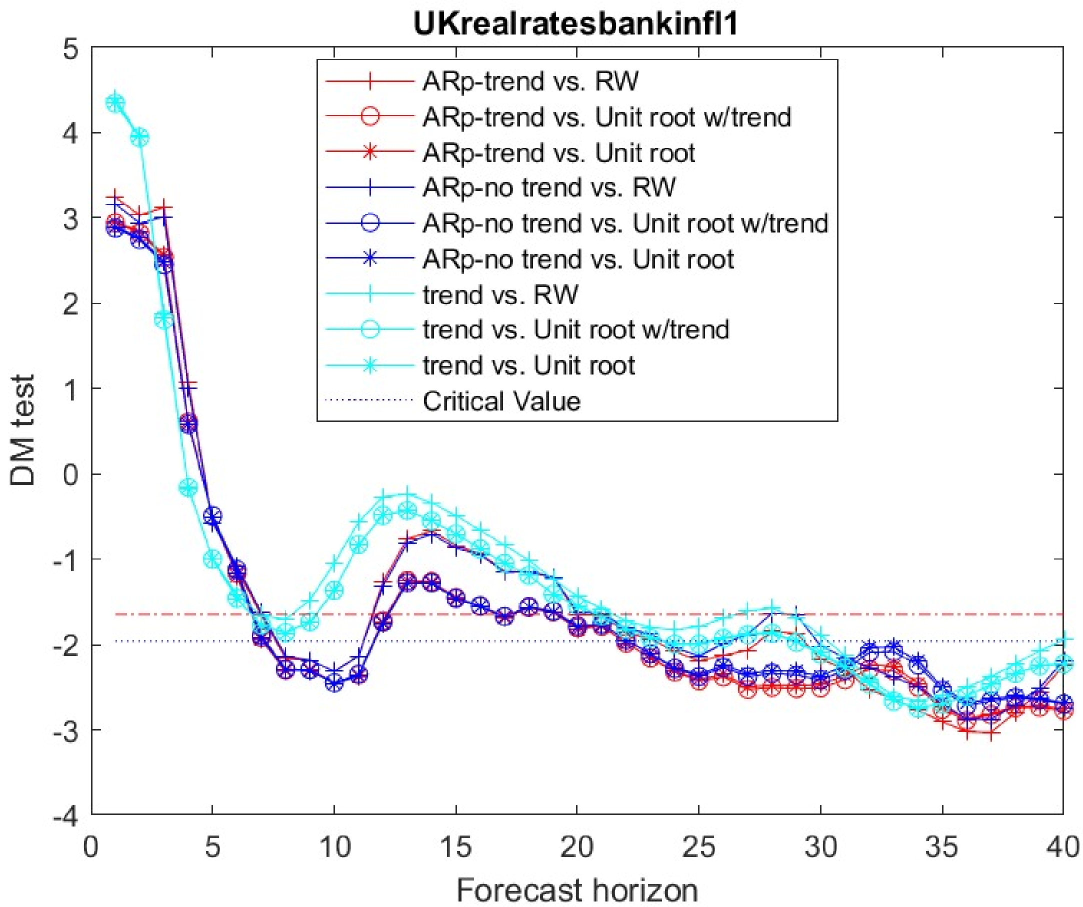
2. Additional Out-of-Sample Results

We report additional results for the out-of-sample exercise introduced in section 7 of the main paper. First, Figure A.2 displays results for the U.K. quarterly private real rate series, on the Bank rate 1-year basis.

Second, Figure A.3 below displays the out-of-sample exercise for the U.K. quarterly term spread on the 1-year bank rate ("Spread Public") basis. We observe that our trend stationary model continues to beat stationary models – autoregressive models with trend, labelled ARp-trend –, as well as models that impose a unit root – labelled ARp-no trend.

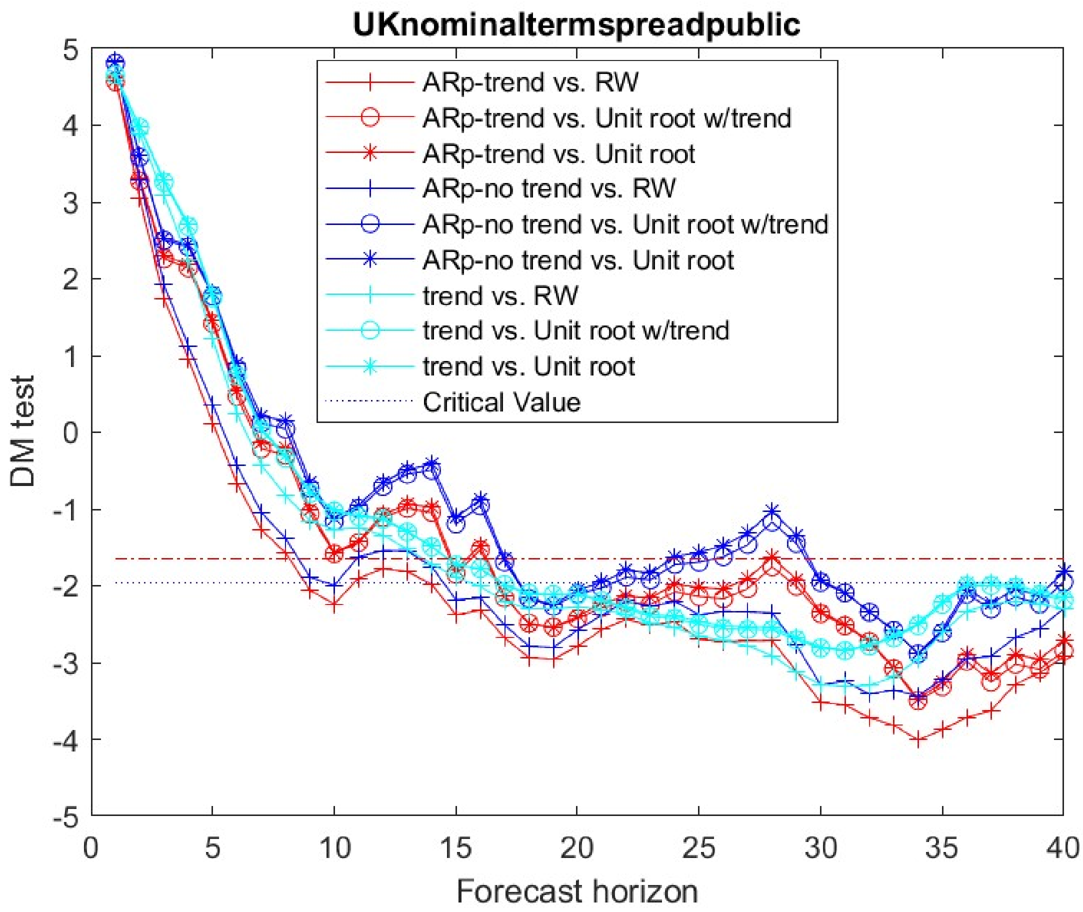
Third, Figure A.4 displays results for the U.K. quarterly real term spread series, with the spread calculated as the difference between consol rates on the long-end, and bank rates on the short end – with these rates inflation-adjusted through the 2Y basis (taking the average of t-1 and t=0 year-on-year price change at each quarter).

Figure A.2: Out-of-Sample Forecast Comparisons: U.K. Quarterly Real Rate Bank Infl1.



Notes: The picture reports the Diebold and Mariano (1995) test for equal predictive ability for stationary vs. non-stationary models, described in the legend. Negative values indicate that the stationary model has a lower mean squared forecast error than the non-stationary model. The picture also reports the critical value based on the Giacomini and White (2006) approximation (one-sided is -1.645 (dotted line) and two-sided is -1.96 (dashed-dotted line)). Forecasts are obtained in rolling windows with size equal to half of the total sample size.

Figure A.3: Out-of-Sample Forecast Comparisons: U.K. Nominal Term Spread Public.



Notes: The picture reports the Diebold and Mariano (1995) test for equal predictive ability for stationary vs. non-stationary models, described in the legend. Negative values indicate that the stationary model has a lower mean squared forecast error than the non-stationary model. The picture also reports the critical value based on the Giacomini and White (2006) approximation (one-sided is -1.645 (dotted line) and two-sided is -1.96 (dashed-dotted line)). Forecasts are obtained in rolling windows with size equal to half of the total sample size.

All our key results hold when testing the Officer (2021) series via ADF-GLS, and subsequently Zivot and Andrews (1992): per table A.4, which includes a time trend, we can confirm trend stationarity at the 1% level for both U.S. nominal and real rates, as well as for nominal and real term spread bases – this result is confirmed in Table A.7, where we equally find stationarity at the 1% significance level for real U.S. short-maturity series, as well as for nominal term spreads, all over the period 1831-2021.

Table A.5 now reports stationarity test results for headline U.S. series based on the ADF-GLS test without including a time trend. The results confirm our main results in the body of the paper, with stationarity holding at the 1% significance level throughout.

3. Further econometric results – ADF-GLS, Bai-Perron, Zivot-Andrews, and Chow test

Bai-Perron results follow in Table A.6, via the standard methodology of Bai and Perron (1998). A variety of structural break tests are in principle available to conduct such tests, some of which could well nuance the

Table A.4: ADF-GLS Test with Time Trend, Officer (2021).

U.S. short-maturity (ST) rates and term spreads, 1831-2021

	Number of lags	ADF-GLS test statistic	Optimal lag
U.S. annual short rate real	3	-6.144	
	2	-7.243	Seq, SIC
	1	-6.721	MAIC
U.S. an. nominal t-spread	3	-5.530	
	2	-5.364	
	1	-5.655	SIC, MAIC

Notes: The table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of 3 lags). The regression includes a constant and a deterministic time trend. The critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -3.471 (1%); -2.939 (5%); -2.649 (10%). "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the Bayesian Information Criterion (SIC), or the Modified Information Criterion (MAIC). The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. All observations at annual frequency, and in nominal terms. "OF" refers to annual "ordinary funds" series in Officer (2021), long-maturity nominal data for term spread via Schmelzing (2025). Short-maturity real interest rates are constructed by adjusting current-year inflation to nominal rates, while long-maturity real rates use seven-year progressively lagged inflation expectation approach, following Homer and Sylla (2005).

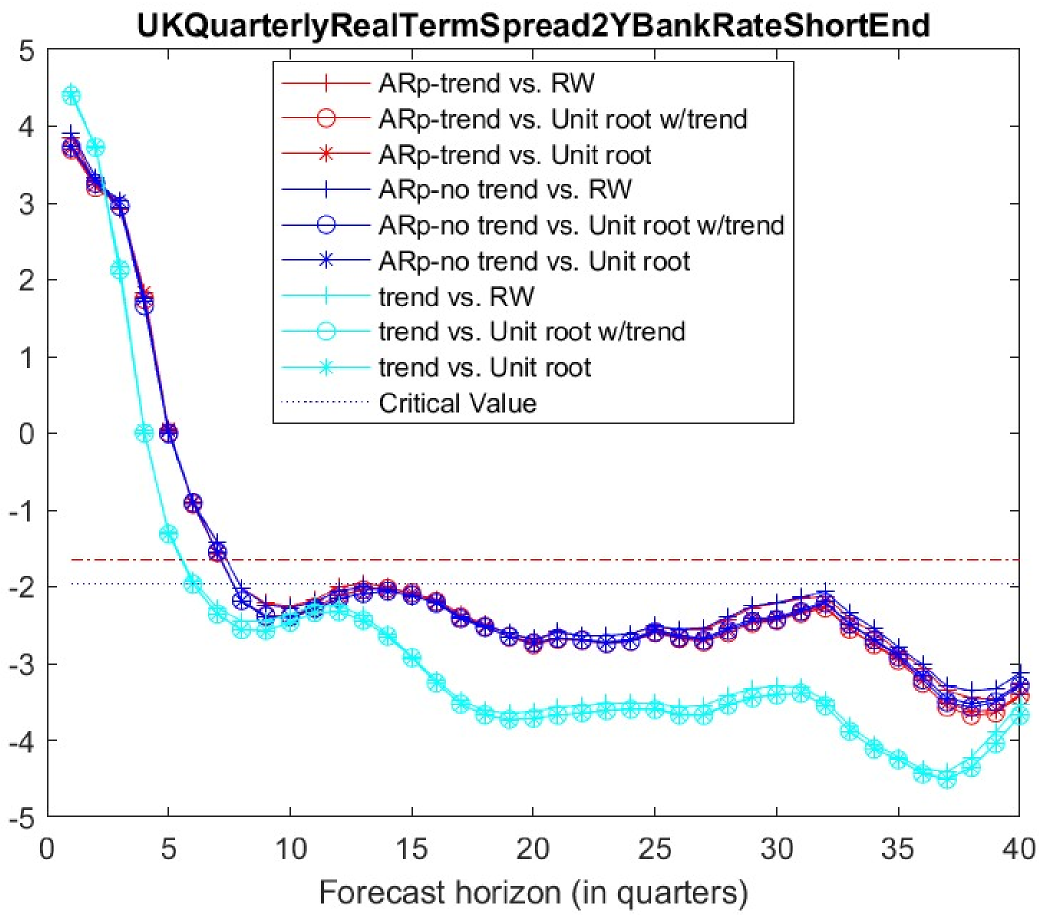
Table A.5: ADF-GLS Test without Time Trend, selected U.S. series.

U.S. short-maturity (ST) rates and term spreads, 1831-2024

	Number of lags	ADF-GLS test statistic	Optimal lag
U.S. annual real, OF	3	-5.919	
	2	-7.014	Seq, SIC
	1	-6.557	MAIC
U.S. annual nominal t-spread	3	-5.465	
	2	-5.313	
	1	-5.608	SIC, MAIC
U.S. quarterly real, RRS	3	-10.623	
	2	-11.244	
	1	-11.572	Seq, SIC, MAIC

Notes: The table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of 3 lags). The regression includes a constant but no deterministic time trend. The critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -2.588 (1%); -1.950 (5%); -1.616 (10%). "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the Bayesian Information Criterion (SIC), or the Modified Information Criterion (MAIC). The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. The series "OF" refers to Officer (2021) "ordinary funds" series and is annual over 1831-2021; the "RRS" quarterly real series constructed as described in main text (spanning 1831:Q1 – 2024:Q2); and nominal term spread uses annual consistent concept commercial paper series on short-end, and long-maturity nominal rates used via Schmelzing (2025). Short-maturity real interest rates are constructed by adjusting current-year inflation to nominal rates for annual basis, and year-on-year change in wholesale price index for quarterly real basis, as per main text.

Figure A.4: Out-of-Sample Forecast Comparisons: U.K. Quarterly Real Term Spread 2Y Bank Rate Short End.



Notes: The picture reports the Diebold and Mariano (1995) test for equal predictive ability for stationary vs. non-stationary models, described in the legend. Negative values indicate that the stationary model has a lower mean squared forecast error than the non-stationary model. The picture also reports the critical value based on the Giacomini and White (2006) approximation (one-sided is -1.645 (dotted line) and two-sided is -1.96 (dashed-dotted line)). Forecasts are obtained in rolling windows with size equal to half of the total sample size.

standard tests we choose to follow in the following. However, much recommends the use of a Bai-Perron test, via Bai and Perron (1998), a prominent mainstream test that results in "best fit" observations up to a defined maximum number of breaks, without relying on imputed dates. Especially over the course of our observation period in question (the past 200-300 years), tests that are relying on such imputed dates display high levels of sensitivity on the specific imputed dates, which we want to avoid.[30]

We observe that in annual level U.K. short-maturity real rates, the years 1977 and 2007 are confirmed as significant breaks at the 1% level, but only with U.K. Bank rate, but not the much longer commercial paper rate. Remarkably, even for the longest and most high-frequency series, the quarterly U.K. bases, with one exception (public nominal term spread) we consistently fail to identify a single break, across all specific iterations of short-maturity rates, as well as term spreads – and across all plausible variations of inflation adjustments.

[30] We present the results for a Chow test with such sensitive properties in a subsequent Appendix section, section 5.

Table A.6: Bai-Perron Test results, short-maturity rates and term spreads	
U.K. annual series	Estimated break dates
U.K. public term spread nom.	2007
U.K. private term spread nom.	none
U.K. real public	none
U.K. real private	none
Quarterly series	
U.K. private real 1	none
U.K. private real 2	none
U.K. public real 1	none
U.K. public real 2	none
U.K. public nominal term spread	none
U.S. annual series	
U.S. short-maturity real	none
U.S. term spread nominal	none

Notes: The table reports the results of the sequential Bai and Perron's test (Bai and Perron, 1998) for the U.S. real short-maturity rate data in Officer (2021), which is combined with long-maturity data for the U.S. as based on Schmelzing (2025) for the term spread series. U.K. data per text. The test is implemented in Matlab. The model includes a constant and a deterministic trend (as a fraction of the total sample size) and we test for breaks in both the constant as well as the slope of the deterministic trend. The test is implemented using a HAC variance estimator, the trimming parameter is 15 percent and the maximum number of break points is 5 for all series. All dates refer to significance at the 1% level. Real rates, unless otherwise stated, uses current year inflation adjustment.

Crucially, the same pattern repeats itself when we test the U.S. data. Importantly, Bai-Perron do not indicate a single relevant structural break for either the annual nominal term spread series or the real short-maturity rate series (1831-2024).

In sum, across countries, the Bai-Perron results at least point towards a far greater degree of continuity in short-maturity interest rate and term spread trends on the real basis than thus far recognized. For a majority of series – across different inflation expectation approaches and observation frequencies (annual and quarterly data) – we in fact find zero breaks over time. Importantly, across all series, it is difficult to clearly confirm a real interest rate inflection point during the 1970s or 1980s for any of these series – such breaks generally appear more sensitive to the frequency and issuer levels than one would expect.

Table A.7: Zivot-Andrews Test Results, annual U.S. data	
	Test statistic
U.S. annual real, Officer	-7.942
U.S. nominal t-spread, Officer	-6.861
U.S. annual real, RRS	-7.953

Notes: The table reports the Zivot and Andrews (1992) unit root test statistic, which allows for a break in both the mean and the trend under the alternative. The critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -5.57 (1%); -5.08 (5%); -4.82 (10%). The trimming parameter is 0.10. The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. Underlying "OF" data sourced from Officer (2021) "ordinary funds" series for short-maturity. "RRS" refers to construction approach described for quarterly rates in main text, transposed to annual basis. Short-maturity annual rates are adjusted with current-year year-on-year wholesale price index changes, per Appendix section 1.1.

4. Variations of inflation expectations

In the main body of this paper, we adjusted nominal short-maturity interest rates with current-year realized year-on-year inflation based on David and Solar (1977) in the case of the U.S., and via Dimsdale and Thomas (2016) in the case of the U.K. However, Barro (2006) and Hamilton et al. (2016) use an autoregressive inflation expectations approach to obtain ex ante U.S. short-maturity real rates.

Here we therefore test such an alternative approach for both U.S. and U.K. short-maturity real interest rate series, applying an autoregressive AR(1) inflation adjustment approach in line with Hamilton et al. (2016) both to the nominal interest rate series on the U.S. conceptual-consistency basis, and that of Officer (2021) over 1831-2021, as well as for the two real short-maturity rate variations for the U.K.

Our ADF-GLS results hold for the AR(1) inflation adjustment basis, as per table A.8, which includes a time trend. We observe that for both U.S. and for both U.K. variations of short-maturity real rates, at least 10% trend stationarity significance is always obtained – and for multiple iterations, even 1% or 5% significance is confirmed. The same is true when ADF-GLS is tested without a time trend, as presented in table A.9: once more, we can confirm stationarity here, at the very least at a 10% significance level, but for the majority of series at a 1% level.

Overall, these results suggest that our trend stationarity results for short-maturity real rates are not driven by our particular – widely used – inflation adjustment approach in the main text, where current-year realized inflation is used to adjust nominal rates.

We equally note that using the Zivot-Andrews test for autoregressively-derived inflation expectations results in a confirmation of trend stationarity: per table A.10, trend stationarity is confirmed at the 1% level for all four real rate iterations.

5. Chow test results, and autoregression approach

In an earlier section of the appendix, we presented Bai-Perron test results to investigate structural breaks in short-maturity (real and nominal) interest rtae series. We argued that the use of an open-ended test such as Bai-Perron is better suited for our purposes because over the observation horizon in question (the

Table A.8: ADF-GLS Test, with time trend.			
Short-maturity real rates, autoregressive variations			
	Number of lags	ADF-GLS test statistic	Optimal lag
U.S. annual real, Officer	3	-3.921	MAIC, Seq, SIC
	2	-5.625	
	1	-6.701	
U.S. annual real, RRS	3	-3.387	Seq, SIC, MAIC
	2	-4.551	
	1	-5.794	
U.K., Public real rate	3	-5.017	MAIC
	2	-6.009	
	1	-7.230	Seq, SIC
U.K., Private real rate	3	-5.692	MAIC
	2	-7.013	Seq, SIC
	1	-9.058	

Notes: The table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of 3 lags). The regression includes a constant and a deterministic time trend. For the U.S. series, the critical values at the 1, 5 and 10 percent significance levels are the following: -3.504 (1%); -2.967 (5%); -2.677 (10%). For the U.K. series, the critical values at the 1, 5 and 10 percent significance levels are the following: -3.480 (1%); -2.890 (5%); -2.570 (10%). "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the Bayesian Information Criterion (SIC), or the Modified Information Criterion (MAIC). The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. All observations at annual frequency, and in nominal terms. U.S. short-maturity nominal rate data via ordinary funds series in Officer (2021), all U.K. data sourced via Dimsdale and Thomas (2016) to 2016, and extended to 2024 via Bank of England (2024). Short-maturity real interest rates are constructed by applying an autoregressive AR(1) approach in line with Hamilton et al. (2016) to the nominal interest rate data.

Table A.9: ADF-GLS Test, without time trend.			
Short-maturity real rates, autoregressive variations, annual data			
	Number of lags	ADF-GLS test statistic	Optimal lag
U.S., Officer (2021) basis	3	-2.891	SIC, MAIC, Seq
	2	-4.303	
	1	-5.327	
U.S., RRS	3	-1.773	Seq, SIC, MAIC
	2	-2.603	
	1	-3.541	
U.K., public basis	3	-3.963	Seq, MAIC
	2	-4.892	
	1	-6.072	SIC
U.K., private basis	3	-5.115	Seq, SIC, MAIC
	2	-6.396	
	1	-8.393	

Notes: The table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of 3 lags). The regression includes a constant and a deterministic time trend. For the two U.K. series, the critical values at the 1, 5 and 10 percent significance levels are the following for all observations: -2.580 (1%); -1.950 (5%); -1.620 (10%). For the two U.S. series, the critical values at the 1,5 and 10 percent significance levels are the following for all observations: -2.588 (1%); -1.950 (5%); -1.616 (10%). "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the Bayesian Information Criterion (SIC), or the Modified Information Criterion (MAIC). The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. All observations at annual frequency. U.S. short-maturity nominal rate data via ordinary funds series in Officer (2021), all U.K. data sourced via Dimsdale and Thomas (2016) to 2016, and extended to 2024 via Bank of England (2024). Short-maturity real interest rates are here constructed by applying an autoregressive AR(1) approach in line with Hamilton et al. (2016) to the nominal interest rate data.

Table A.10: Zivot-Andrews Test results, autoregression approaches, annual data	
	Test statistic
U.S. real rate, Officer (2021) basis	-5.096
U.S. real rate, RRS	-4.608
U.K. real rate, public	-8.543
U.K. real rate, private	-5.835

Notes: The table reports the Zivot and Andrews (1992) unit root test statistic, which allows for a break in both the mean and the trend under the alternative. The critical values at the 1, 5 and 10 percent significance levels are the following: -5.57 (1%); -5.08 (5%); -4.82 (10%). The trimming parameter is 0.10. The test rejects when the test statistic is negative and larger (in absolute value) than the critical value. Underlying data sourced from Officer (2021) for U.S. short-maturity real rates, and Dimsdale and Thomas (2016) for U.K. short-maturity real rates. "RRS" refers to construction approach described for quarterly rates in main text, transposed to annual basis. Short-maturity rates are adjusted with autoregressively-derived inflation expectation estimate in line with Hamilton et al. (2016), with data sources as described in Appendix section 1.1.

past 300 years), the variation of sample length appears to impact dramatically the confirmation of break dates in other test setups. To illustrate this phenomenon, we here present the results of the Chow structural break test – which relies on an "imputed" set of structural break dates, which we define ourselves based on existing literature. We test the following five key years, all of which are linked to strong existing claims in financial and historical analyses covering the period over 1800-2021:

- **1871** – the global switch to the classical gold standard, which occurred throughout the 1870s, but is often directly associated with the German decision to transition away from silver after its military victory over France (Bordo and Schwartz, 1999).
- **1915** – the founding of the Fed, and the end of the classical gold standard. The proposition of 1914-5 as a major inflection point existed for some time, expressed for instance by Mankiw and Miron (1986) and more recently Bernstein et al. (2010).
- **1933** – the Great Depression, and FX regime shifts. The year 1933 marks the bottom of the banking crisis and is the final year of the free exchange of U.S. Dollars for gold, ended by the Jan 1934 Gold Reserve Act.
- **1981** – the Volcker "war on inflation" and the start of the Great Moderation (Bernanke, 2004).
- **2008** – the GFC, and the arrival of negative policy rates.

Table **??** displays the results, testing both breaks in trends and in the mean. Once we focus on this panel, and allow for a time trend, the following dates appear as significant in U.S. real short-maturity rates: 1915, 1933, 1981, and 2008.[31]

However, we notice in table **??** that once longer samples are tested – the U.K. real rate series runs from 1704-2016, here on the Bank rate basis – our confirmation of structural breaks via Chow is dramatically reduced. Specifically, we test five dates that incorporate significant expected structural breaks over this time horizon in the U.K. context, adding 1756 (outbreak of the Seven-Years War) and 1866 (the Overend Gurney financial crisis) to the earlier sample. Over the 1704-2016 U.K. real short-maturity sample, we can now only confirm a single marginal break, that of 2008.

This radical reduction of structural breaks does not depend on the exact short-maturity measure used: as we demonstrate in table **??**, we equally confirm only a single marginal break (once more for 2008) when switching to a U.K. private short-maturity basis, testing real prime bill rates.

6. Phillips Perron

Table A.13 presents the results of a Phillips Perron Test, with a time trend.

[31]We also obtained Chow results for the panel without a time trend, and for the nominal sample. For the nominal sample, allowing for a trend, the years 1933, 1981, and 2008 are significant breaks in the Chow test setup.

Table A.11: Chow Test Results - U.S. real short-maturity rates, annual			
Regressor	Coefficient	Std error	t-statistic
trendbreak 1871	8.88	15.71	0.57
trendbreak 1915	-251.52	24.45	-10.29
trendbreak 1933	200.38	24.77	8.09
trendbreak 1981	78.92	9.46	8.34
trendbreak 2008	-65.90	12.63	-5.22
meanbreak 1871	-5.44	2.33	-2.34
meanbreak 1915	118.31	12.26	9.65
meanbreak 1933	-90.48	12.63	-7.17
meanbreak 1981	-66.04	7.55	-8.75
meanbreak 2008	60.84	11.26	5.40
trend	7.30	7.63	0.96
constant	-7.97	7.21	-1.11

Notes: The model includes a constant and a deterministic trend (as a fraction of the total sample size). The standard errors are based on Newey and West's (1987) HAC estimator is chosen according to Lazarus et al. (2018). The critical value of the (absolute value of the) t-statistic is 2.24 at the 5 percent significance level and 3.089 at the 1 percent. The coefficient associated with "trendbreak 1981" denotes the difference of the estimated trend coefficients before and after a break in 1981 (allowing for breaks in all the other potential break dates); hence, the associated t-statistic is the Chow test for the absence of a structural break in trend in 1981. Similarly, "meanbreak 1981" refers to a break in the mean in 1981. The results refer to the consistent-concept real interest rate data as described in the main text.

Table A.12: Chow Test Results - U.K. real short-maturity rates, public, annual			
Regressor	Coefficient	Std error	t-statistic
trendbreak 1866	-1.56	10.51	-0.15
trendbreak 1915	-1.58	12.55	-0.13
trendbreak 1981	22.43	14.78	1.52
trendbreak 2008	-12.53	11.95	-1.05
meanbreak 1866	2.54	2.88	0.88
meanbreak 1915	3.44	7.83	0.44
meanbreak 1981	-22.96	11.08	-2.07
meanbreak 2008	16.46	10.86	1.52
trend	-7.37	8.65	-0.85
constant	4.85	8.2	0.59

Notes: The model includes a constant and a deterministic trend (as a fraction of the total sample size). The standard errors are based on Newey and West's (1987) HAC estimator where the lag length is chosen according to Lazarus et al. (2018). The critical value of the (absolute value of the) t-statistic is 2.168 at the 5 percent significance level and 2.957 at the 1 percent. The coefficient associated with "trendbreak 1981" denotes the difference of the estimated trend coefficients before and after a break in 1981 (allowing for breaks in all the other potential break dates); hence, the associated t-statistic is the Chow test for the absence of a structural break in trend in 1981. Similarly, "meanbreak 1981" refers to a break in the mean in 1981. The results refer to the U.K. real short-maturity interest rate data on the Bank rate basis, with data sourced from Dimsdale and Thomas (2016) and with nominal rates adjusted with current-year realized U.K. wholsesale inflation.

Regressor	Coefficient	Std error	t-statistic
trendbreak 1756	-25.71	6.41	-4.01
trendbreak 1866	8.33	13.98	0.60
trendbreak 1915	-9.12	20.05	-0.45
trendbreak 1981	34.65	21.65	1.60
trendbreak 2008	1.41	16.41	0.09
meanbreak 1756	4.48	1.24	3.62
meanbreak 1866	-1.24	7.85	-0.16
meanbreak 1915	8.91	15.00	0.59
meanbreak 1981	-35.15	18.00	-1.95
meanbreak 2008	3.26	15.52	0.21
trend	-27.87	12.80	-2.18
constant	24.98	12.38	2.02

Table A.13: Chow Test Results - U.K. real short-maturity rates, private

Notes: The model includes a constant and a deterministic trend (as a fraction of the total sample size). The standard errors are based on Newey and West's (1987) HAC estimator where the lag length is chosen according to Lazarus et al. (2018). The critical value of the (absolute value of the) t-statistic is 2.168 at the 5 percent significance level and 2.957 at the 1 percent. The coefficient associated with "trendbreak 1981" denotes the difference of the estimated trend coefficients before and after a break in 1981 (allowing for breaks in all the other potential break dates); hence, the associated t-statistic is the Chow test for the absence of a structural break in trend in 1981. Similarly, "meanbreak 1981" refers to a break in the mean in 1981. The results refer to the U.K. real short-maturity interest rate data on the prime bill basis, with data sourced from Dimsdale and Thomas (2016) and with nominal rates adjusted with current-year realized U.K. inflation.

Table A.14: Phillips Perron Test with Time Trend					
Real Rate and nominal term spread series, U.K. annual bases					
		Test Statistic	Critical Value		
			1%	5%	10%
UK nominal term spread, public basis (1822-2024)	Z_ρ	-106.438	-28.08	-21.08	-17.84
	Z_τ	-8.377	-4.006	-3.436	-3.136
UK nominal term spread, private basis (1718-2024)	Z_ρ	-159.814	-28.512	-21.345	-18.022
	Z_τ	-10.275	-3.988	-3.428	-3.130
UK real rate, public basis (1822-2024)	Z_ρ	-119.678	-28.08	-21.08	-17.84
	Z_τ	-9.151	-4.006	-3.436	-3.136
UK real rate, private basis (1718-2024)	Z_ρ	-190.503	-28.512	-21.345	-18.022
	Z_τ	-12.258	-3.988	-3.428	-3.130

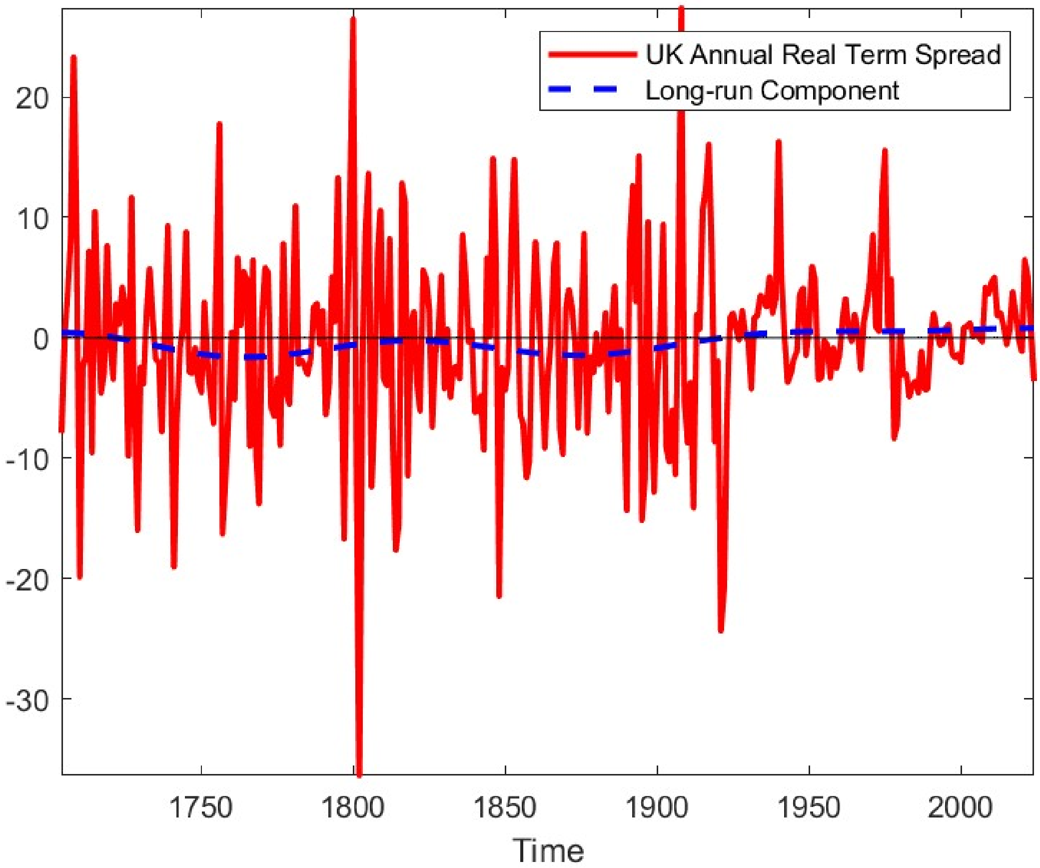
Notes: The table reports the Phillips-Perron Z_ρ and Z_τ test statistics and their critical values at the 1, 5 and 10 percent significance levels. The regression includes a constant and a deterministic time trend. The standard errors are obtained by Newey–West's HAC estimator. The test rejects when the test statistic is negative and larger (in absolute value) than the critical value.

7. Real Rate term spread, U.K.

We rationalized the focus on nominal term spreads in the paper, not least given bigger assumptions that need to be made regarding inflation expectations – but we mentioned that a construction of real term spreads is in principle possible in light of most recent data advances. In light of occasional interest in the literature on this real basis (e.g. Harvey (1988)), Figure A.5 displays a construction of an annual frequency real term spread series over 1704-2024. This is done on the basis of the U.K. real long-maturity data in Rogoff et al. (2024), which approximates a benchmark 10-year real interest rate over the very long-run, and the real short-maturity series (public basis) introduced in this paper.

While basic constructions on this real level are thus possible, we do not analyze this series further here, beyond noting that it also appears to show a moderate upwards slope over time, in line with the nominal bases shown in the main part of this paper, and stands at an all-time average of -0.37%.

Figure A.5: Real term spread construction, public basis, U.K., 1704-2024.



Notes: The picture displays a real term spread series for the U.K. as constructed on the basis of long-maturity real rates for the U.K. in Rogoff et al. (2024), and U.K. real public short-maturity rates as introduced in this paper. The blue dashed line represents the long-run component in the series, as based on the Müller and Watson (2018) methodology, selecting fluctuations with periodicity of 100 years or more.