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ABSTRACT

Taking advantage of key recent advances in long-run financial and economic data, this paper analyzes the statistical properties of global long-maturity real interest rates over the past seven centuries. In contrast to existing consensus, which has overwhelmingly concentrated on short samples for short-maturity rates, we find that long-maturity real interest rates across advanced economies are in fact trend stationary, and exhibit a persistent downward trend since the Renaissance. We investigate structural breaks in real interest rates over time using multiple statistical approaches, and find that only the Black Death and the "Trinity default" of 1557 appear as consistent inflection points in capital markets on both global and country levels. While a 1914 break is also suggested in multiple series (though less robust than existing literature would lead one to expect), the evidence for an inflection point in 1981 appears much weaker. We further examine trends in persistence, as well as commonly-invoked drivers of global real rates: exploiting significant data advances, we argue that historically, demographic and productivity factors appear to show no promising causal role, and in fact diverge from real interest rates over the long run.

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1 Introduction

More than a decade after the Global Financial Crisis (GFC), and following a sharp rebound in economic activity from a worldwide pandemic shock, global real interest rates remain at historically depressed levels. This ongoing fall in real interest rates preceded those major macroeconomic shocks as well as the era of "unconventional monetary policy" - and has led to intense academic discussions attempting to identify explanations for the phenomenon. Despite disagreements over specific drivers, many researchers have posited that current real interest rate dynamics can be traced to a key inflection point during the 1970s or 1980s, when the transition from fixed exchange rates coincided with a deregulation spurt in international finance and an acceleration of globalization, culminating in the "great moderation" phase (Rachel and Summers, 2019). Other contributions emphasize structural global factors, including financial flows Bernanke (2005), demographics (Goodhart and Pradhan, 2021), trend growth (Holston et al., 2017), or monetary policy (Garcia and Perron, 1996; Bernanke, 2004). Summaries are provided by Blanchard (2022) and Lunsford and West (2019), with the latter finding the strongest positive association between demographic factors and the safe rate between 1890-2016.

This paper will take advantage of recent advances in long-run financial and economic data to approach the debate from a fundamentally different angle - investigating real interest rate trends from a multi-century perspective. Beyond contextualizing trends and drivers for the recent decades, this paper is motivated by a literature that sought to characterize the fundamental statistical properties of (real) interest rates, and their relation to other macroeconomic variables. For both areas of interest, (very) long-run datasets can potentially offer far superior results compared to an existing literature that has typically focused on relatively short sample periods, typically 75 to 150 years at most. First, it is clear that a multi-century overview encompasses a much richer tail event, monetary, and fiscal variation over time than an exclusive post-War, or post-Bretton Woods assessment - and this longer perspective is the obvious desideratum to assess "secular" forces in the global economy, and the structural impact of generational "macro shocks" like the GFC or COVID. However, such investigations were until recently impossible given a lack of adequate data.

Second, there are key econometric reasons that recommend a very long-run examination of the statistical properties of real rates: most importantly, it is known that the power of unit root and other statistical tests improves significantly over longer sample periods; the advantages remain, even relative to cases where one can get more observations by using a shorter sample that has higher frequency observations (Perron, 1989). Frankel (1986) in his exposition on PPP demonstrated that the rejection of the random walk for U.S.-U.K. exchange rate time series crucially depended on the time horizon: the statistical power significantly improved once horizons of over 100 years were chosen, finally allowing for a rejection of the unit root.[1] Therefore, even though recent

[1]For a practical example, see Lothian and Taylor (1996). The authors there show that the existence of a unit root for real exchange rates cannot be rejected over the sub-periods of 1946-1990, and 1974-1990 - but this result disappears when the sample length is extended to 1803-1990, for which the series can clearly reject the unit root. The authors use a range of other random 45-year intervals, for which the rejection of unit root is also impossible: this strongly underscores the role of sample length, rather than regime-specific features involved in the post-1945 episodes. We test

contributions have expanded the time span of statistical real interest rate assessments into the 19th century (Hamilton et al., 2016), ideally longer datasets should be used.

Third, existing analyses of real interest rate dynamics overwhelmingly focus on short-maturity interest rates, with very few exceptions. In fact, all key reference contributions have reached generalizations on interest rate properties on the basis of central bank policy or bill rates. Much, however, suggests that using *long-maturity* interest rates, which reflect fundamental growth and inflation expectations, is more appropriate. In any event, different ends of the term structure need not necessarily exhibit identical structural statistical features. Not least, there are crucial data reasons that recommend the use of long-maturity data: long-maturity financial instruments - perpetual consols, for instance - have been in existence far longer than either short-term consolidated debt or rate-setting policy institutions, with the issuance of consolidated sovereign long-maturity debt dating back at least to the first Venetian issuance of negotiable *Monte Vecchio* bonds in 1262 (Pezzolo, 2003; Stasavage, 2011; Eichengreen et al., 2021). In addition, long-maturity data properly reflect structural trends and expectations by market participants: long-maturity rates incorporate structural growth, rare disaster, inflation, demographic, and capital flow expectations - as opposed to short-maturity rates, which are much more likely to be subject to noise or random business cycle shocks. For the assessment of "secular", "generational" forces affecting large parts of the advanced international system over long investment horizons, therefore, it is not clear that the existing literature has been operating with a plausible empirical basis.[2] Not least, historical studies have repeatedly demonstrated that as early as the 14th century, households and institutions treated sovereign long-maturity assets as a relatively "safe asset", to which they increased financial exposure in times of crisis (de Moor and Zuijderduijn, 2013).[3]

With recourse to a new multi-century real interest rate dataset constructed by Schmelzing (2022) and based on long-maturity market rates, this paper will seek to analyze more fundamental dynamics underpinning both the secular evolution, as well as the present features of the global interest rate environment. Following a literature review, section three of this paper will describe the new long-run interest rate data, with section four proceeding with an analysis of the general statistical properties of global real rates since the birth of sovereign debt in 1311.

There, we will present three key statistical results. First, we show that - in contrast to the existing consensus - we are able to reject non-stationarity at conventional significance levels once one allows for a deterministic trend and long-maturity real interest rate data are employed: for both headline global real rates, as well as for all advanced economy country-level series, this result is shown to

our basic conclusions for different sample lengths in the appendix in an analogous way, an exercise which confirms the importance of sample length for real interest rates.

[2] These dynamics are underpinned by asset pricing theory: as risk-averse investors move from single-periods to infinite horizons, the riskless asset is no longer a real Treasury bill, but a real perpetuity (Campbell, 2018, chapter 9).

[3] For instance, de Moor and Zuijderduijn (2013) recently showed that Dutch households systematically increased exposure to long-maturity annuities in times of heightened pandemic, natural disaster, or war frequencies - at the expense of exposure to livestock, houses (though not land), and cash. Mulho (1971) and Mueller (1997) showed separately that not insignificant shares of Venetian and Florentine debt in the 14th and 15th centuries was foreign-owned, bought by European investors in search for financial safety.

hold for long samples, and confirmed for a range of test specifications. Global real rates exhibit a gentle but firm downward trend (averaging about 1.6% every 100 years), and the evidence appears strongly consistent with the hypothesis that they are stationary around this trend. We confirm this result in both global and individual country data, using both ADF-GLS tests and the Zivot and Andrews (1992) test (the latter allows for structural breaks, though for reasons discussed below we prefer a Chow test approach). Nevertheless, we can replicate the non-stationarity findings for the much shorter data series available for short-maturity rates even when a trend is included. However, we show also that evidence from long-maturity corporate and mortgage rate trends over time suggests that trend stationarity may be present more broadly in long-maturity debt.

Second, we analyze the data for structural breaks: we start our analysis with a Chow test (Chow, 1960), focusing on a list of widely-discussed break dates (including 1981), consistent with traditional and recent scholarship. Among our posited structural breaks, only the Black Death, as well as 1557 - the onset of a largely forgotten global financial crisis that "shook Europe to its foundation" (Ehrenberg, 1928, 114) and saw the concurrent default of three of the largest sovereign creditors, together with the impending collapse of perhaps the largest "systemically important financial institution" ever to have existed - are consistent statistically significant periods of structural change on both the global and country levels. The two key 20th century inflections - 1914 and 1981 - are also confirmed for a variety of subseries, but overall appear to be less significant than one would expect. If instead of the Chow test we implement the Bai and Perron (1998) test, in which the break dates are not pre-specified, again there are no confirmed breaks after 1921 except in the case of the United States and Japan, which of course have the shortest sample. We suggest in our interpretation that structural breaks can be plausibly linked to major combined political-economic shocks - but not per se to institutional innovations (such as the founding of central banks), sovereign default events alone, or FX regime changes - and generally fail to be associated with major monetary policy events (whether defined institutionally, or measured by dramatic shifts in policy).

Third, on the basis of our new data, we also present results on real rate persistence; when real rates drop, how long do they take to return to trend? Existing knowledge about long-run trends in the persistence of financial variables is thus far limited to scattered evidence on FX rates, and deviations from purchasing power parity. We present the first analysis on real interest rate persistence over the very long-run: the estimated half life deviations from the long run trend hover around two years over the first six centuries of the sample, showing a moderate upwards trend. Notably, therefore the early modern period (defined in this paper as the full period prior to the Congress of Vienna, 1300-1815) already experienced adjustment speeds in capital markets comparable to those during the late classical gold standard. After 1914, we find that half-lives for global real rates meaningfully increased, before moderating once more since the 1980s. Such evidence can contextualize the persistence of contemporary deviations of global real rates from deeper trends, and indicates that while mean-reversion from the sharp falls in global real rates induced by the GFC and the pandemic shock is to be expected, directionally the historical trend does suggests that a downward tendency in global rates will continue to prevail: the expected mean-reversion takes places in the context of a falling trend, rather than in the context of a pre-GFC, or pre-1980s constant value.

Finally, part five assesses the historical and economic context of our statistical results on real interest rate dynamics, in particular evaluating the very long-run relationship with demographics and productivity: we take advantage of the significant advances over recent years in the reconstruction of output growth series by economic historians (e.g. Broadberry and Fouquet (2015); Ridolfi and Nuvolari (2021); Pfister (2021); de la Escosura et al. (2022)) to examine recent literature on the relationship between real interest rates and both output growth and population growth in long-run perspective: the evidence leads us to question arguments that posit that the "recent" decline in global real rates can be primarily explained by productivity or demographic trends (as opposed, for example, to the way that rising global wealth might affect rates of time preference or the intertemporal elasticity of substitution). It appears that for most of history, real interest rates have trended counter to growth and demographics. Taken together, these results suggests far stronger continuities in the present, post-GFC international financial system with (pre-)20th century dynamics than commonly acknowledged: not just regarding the downward trend in real interest rates, but also regarding levels of variation and adjustment speeds. There has indeed been an especially sharp downward shift of real interest rates in the period after the global financial crisis. As we discuss in section 5.2 however, there have been four previous long-lasting eras of below trend rates (generally periods of relative geopolitical calm, though there are additional factors, all of which we briefly assess). What we do know if that all of them ended, and, based on our evidence, the leading theory has to be that the present one will correct from having declined more than long-run trend would imply, but that the direction will secularly continue to be downwards.

2 Literature

Repeatedly, the macroeconomic literature has posited that, on purely *a priori* grounds, real interest rates should exhibit mean-reverting, stationary properties. Foundational asset pricing frameworks such as the Black-Scholes relationship or the consumption CAPM assume constancy in real interest rates. Standard finance no arbitrage models of the yield curve also assume that interest rates are stationary (see survey in Neely and Rapach (2008) and discussion in, e.g., Bauer and Rudebusch (2020)).

However, finance and macro literature actually employing statistical tests has had significant difficulties to confirm these a priori assumptions. Instead, they have overwhelmingly been unable to reject non-stationarity in real interest rates.

The earliest contributions focused on nominal rates: Fama (1975) instigated an intense (and broadly dismissive) debate after suggesting U.S. short-maturity bill rates during 1953-1971 were able to predict subsequent inflation rates. Shiller and Siegel (1977) investigated both long-maturity and short-maturity U.K. interest rates' correlation to price level changes - but were not interested in statistical analysis.[4] Afterwards, attempting a falsification of expectations theory based on data over 1890-1979, Mankiw and Miron (1986) narratively defined four interest rate regimes and suggested that the founding of the Federal Reserve led to a key change in the statistical properties of short-maturity nominal interest rates: over 1915-1979, short-maturity U.S. nominal interest rates followed a random walk; while they exhibited predictable behavior prior to this inflection date.

A key contribution was made afterwards by Rose (1988), who showed that U.S. real interest rates appear to exhibit a unit root: on the annual level, Rose used an ADF test for high-grade long-maturity corporate bonds, and for short-term commercial paper rates, for two periods spanning 1892-1970, and 1901-1950.[5] Later, Garcia and Perron (1996) analyzed a regime-based framework of U.S. ex post real bill rates between 1961-1986: the authors conclude that U.S. real rates exhibit an "essentially random" process, with shifting means and variances in the post-war period, though without a clear predictable patterns. Ang and Bekaert (2002) studied U.S., German, and U.K. bill rates and propose that a regime-switching framework best captures out-of-sample observations in post-war real rate data. Later, Rapach and Weber (2004) used an extended unit-root test (Ng and Perron, 2001) to re-examine Rose (1988)'s findings - and confirmed non-stationarity across a range of countries. While the clear majority of existing studies tested short-maturity (real) interest rates, Nelson and Plosser (1982), Perron (1989) and Rapach and Weber (2004) tested U.S. long-maturity bond yields, albeit for a relatively short time period (1900-1973, or in the case of Rapach and Weber (2004) for 12 countries beyond the U.S.).[6]

More recently, Hamilton et al. (2016) identified real interest rate non-stationarity, partly using policy and discount rates on the basis of U.S. annual data beginning in the late 19th century, and quarterly data beginning in the 1940s. On the annual level, the authors splice together short-maturity rates for various periods, relying on a mix of commercial paper rates and Fed fund rates for the U.S., for instance. A heteroskedasticity-robust test and KPSS results lead the authors to reject the null hypothesis of stationarity for U.S. short-maturity data at the 5% level: Bai-Perron break tests identify 1915 and 1921 as U.S. break dates. The only systematic relationship in the authors' identification relates to a stationary gap between U.S. and world real rates.

Negro et al. (2019) investigated both global long-maturity and short-maturity interest rates over

[4]Specifically, Shiller and Siegel (1977) use British consol yields between 1727-1973 for long-maturity rates, and they splice a variety of 3-month bill rates together for short-maturity rates, beginning with Overend and Gurney corporate rates over 1824-1844, followed by bank bill rates over 1845-1938. Despite the paper title, the authors do not estimate real rates directly. Siegel (1992) relies on the same nominal basis, adding inflation data to construct ex post measures for the U.S. and U.K.

[5]We replicate Rose (1988)'s approach in the appendix (tables A.6.1-A.6.3 and discussion), applying both alternative inflation expectation approaches, for his two periods, and two assets, but also allowing for a trend: we fail to reject a unit root for all 16 variations of the data. Rose sourced historical U.S. inflation and interest rate data primarily from Friedman and Schwartz (1963) and Nelson and Plosser (1982). Non-U.S. data in his analysis begins in 1957 and is tested for quarterly three-month bill rate equivalents, using IMF International Financial Statistics. An earlier paper by Huizinga and Mishkin (1986) used short-maturity U.S. data between 1916-1927, and 1953-1984, and posited a higher degree of (ex post) real rate stability, but did not utilize any ADF, KPSS or related tests.

[6]Perron (1989) relies on Nelson and Plosser (1982)'s data, who fail to reject non-stationarity for U.S. nominal bond yields over 1900-1970.

1870-2016, finding evidence that interest rates "fluctuated between 1% and 2% for about a century, rose in the late seventies, and [have] been on a steady decline ever since". The authors rely empirically on the "Jorda-Schularick-Taylor" (JST) database - but do not undertake formal statistical tests, are not interested in linking the contribution to literature on general interest rate properties, and omit longer periods of time.[7] Finally, Bauer and Rudebusch (2020) propose a time-varying model with shifting endpoints, which the authors suggest is superior in accounting for evidence of the "extreme" interest rate persistence. For ten-year treasury yields on a quarterly basis between 1971:IV - 2018:I, the authors reject stationarity using ADF and Phillips-Perron (PP) tests.

In sum, despite strong a priori propositions that real interest rates should exhibit stationary properties, existing literature has overwhelmingly failed to reject non-stationarity of real interest rates - however, such literature has focused almost exclusively on a combination of spliced short-maturity nominal interest rate data fusing different issuers, and realized inflation rates. And despite multiple studies using timespans of more than one century, their statistical power to examine "secular" properties still remains limited, as evidenced by the apparent spuriousness of the previously-highlighted 1970s and 1980s "inflections". In sum, substantial extensions of sample spans and a clear focus on long-maturity rates could in fundamental ways increase both the quantitative power and qualitative insights in the context of the extensive debates on real interest rate properties: we turn to precisely such an exercise in the next section.

3 Long-run empirics in global real rates, 1311 - 2021

3.1 A short history of short-maturity debt

To understand existing biases arising from the predominant use of short-maturity data in the literature on the properties of real interest rates, it is useful to revisit the historical background to the evolution of short-maturity and long-maturity debt instruments. We have observed in the literature review above that the vast majority of existing contributions assessing real interest rate properties and trends focus on multi-decade empirics for short-maturity yields, or even outright policy discount rates - often splicing short-term discount and market rates. Perhaps one reason for this existing bias is an implicit assumption that an "expectations hypothesis" (EH) holds - which systematically links all maturities in nominal and real terms across the term structure: under this framework, long-maturity yields are a direct function of continuously rolled-over short-maturity assets, and cannot therefore unduly deviate from the properties of the short end. In other words, expected excess returns on long-maturity bonds over short-maturity bonds are constant over time. In other words, "the risk premium of an n-period bond over a one-period bond (the term premium)

[7]As we shall discuss in section 3, the JST database only covers long rates from 1870-2016 and, because it draws quite heavily on Homer and Sylla (2005), it inherits many of the its limitations, as discussed in Schmelzing (2022), which uses primary sources, fills in many significant gaps, and covers a much longer time span.

is constant" (Campbell et al., 2018, chapter 8.2).[8]

Of course, another major reason for the focus on short-rates is that, in modern times, the (very) short rate is the variable controlled by central banks. This does not, however, seem like a compelling reason to prefer it to longer rates that have a much bigger impact on the economy. An additional reason - though rarely stated explicitly - is the exclusive focus of recent "natural" interest rate debates on the short-maturity component, typically referencing Wicksell (1936). For instance, in Laubach and Williams (2003) the natural rate is treated as identical to the real short-term rate: "Since Wicksell (1936), the natural rate of interest—the real short-term interest rate consistent with output equaling its natural rate and constant inflation—has played a central role in macroeconomic and monetary theory". Later in the same paper (p.1064), the authors justify the focus on short-term rates with the inability to measure long-term inflation expectations adequately. This is of course an important issue, and we return to it in section 3.[9]

Hence, however plausible the choice to focus on short-maturities might at first appear - such a choice crucially restricts the insights that can be derived: as is known among economic historians, *consolidated* sovereign short-maturity assets are historically a decidedly "recent" innovation, only emerging as a regularly-used public debt instrument towards the end of the 19th century. Accordingly, this time is also the point of inception for existing "long-run" investigations on general real rate trends (Hamilton et al., 2016; Negro et al., 2019). Short-maturity datapoints, in other words, were only able to cover a fraction of the full empirical horizon of advanced economy "sovereign debt" trends: hence, the statistical power of short-maturity tests has thus far been dramatically limited. However, even key reference works, meanwhile, never reconstructed long-maturity data points on a sufficiently high frequency to allow extensions into longer samples. Only recently, with a new contribution (Schmelzing, 2022) did such an extension become possible for the first time - a development we exploit in the subsequent sections. But before we turn to the new data, we provide the key historical context.

On the back of a decree by its Grand Council in 1262 (the *ligatio pecuniae*), the Venetian Republic is typically regarded as the first issuer of perpetual consolidated sovereign debt, following a century of experimentation across Northern Italian city states with creditor consortia, who were given multi-year rights on tax revenues in return for emergency loans. Swiftly, other city states and cities

[8]Mainstream affine term structure models equally assume that the expectations hypothesis (EH) holds, see (Campbell et al., 2018, chapter 8.3).

[9]It is perhaps relevant to observe that Wicksell (1936)'s definition of the concept not once specifically referred to a short-maturity dimension - in fact, the author has in mind an economy-wide average concept encompassing both short-maturity and long-maturity interest rates. Refer to Wicksell's following original summaries: "At any moment and in every economic situation there is a certain level of the average rate of interest which is such that the general level of prices has no tendency to move either upwards or downwards. This we call the normal rate of interest. Its magnitude is determined by the current level of the natural capital rate, and rises and falls with it". And further: " No statistics of the natural rate of interest are available. A precise investigation would necessitate an ad hoc enquiry, and for the past this is as good as impossible. While the general trend of the natural rate over decades can be observed in the movement of the money rate itself, of the banks' discount rates and the so called open-market rate, and of the prices of debentures and Government securities, this can be done only by assuming that on the average the two rates of interest are equal to one another" (Wicksell, 1936, 120, 168).

north of the Alps follow suit and consolidate individual *rentes* contracts, so that over the course of the 14th century, a pan-European public debt market in long-maturity assets emerges – see Epstein (2000); Munro (2003); Eichengreen et al. (2021). And the evidence suggests that these markets were remarkably liquid: for instance, by the mid-15th century, the face value of annually-traded long-maturity public debt in hubs such as Venice reaches up to 29% of GDP.[10]

By contrast - though they often contracted ad hoc personal loans on short-maturity bases - no sovereign on either side of the Alps issued consolidated short-maturity debt prior to the 18th century. The only outlier may be said to have been the Spanish Empire, which famously experimented with large-scale *Asientos* issuance during the 16th century. The monarchy sharply distinguished this type of issuance from the regular long-maturity *Juros* instruments (on which it never defaulted), and even at the peak of Spanish geopolitical dominance towards the end of the century, the *Asientos* share in total public debt stood at less than 10% (Alvarez-Nogal, 2014).

Regular German, French, Italian, and Dutch consolidated debt issuance up until the 19th century takes place exclusively in long-maturity assets, even though instruments with maturities of more than five years are occasionally referred to as "bills".[11] In the U.S., the first Treasury Bill is auctioned no earlier than December 1929 (Garbade, 2008). Prior, several war financing operations see the intermittent issuance of U.S. "certificates of indebtedness", which are documented to have been sold over 1812-5, in 1837, during the Mexican War, in 1857, during the Civil War, and in 1907. For many decades in between these dates, the U.S. government does not issue any negotiable short-maturity debt instruments. [12] The situation is analogous in the other leading advanced economies: first U.K. Treasury bill issuance is legislated in the 1877 Treasury Bills Act, but only formalized in detail in 1889, and only issued in meaningful volumes from the First World War (BoE, 1963). In dynamics

[10]A figure based on Mueller (1997)'s information that in 1434, 300,000 ducats of *Monte Vecchio* debt in face value traded in Venice.

[11]For the United Provinces, which "occasionally" issued unfunded short-maturity *obligaties* during the 16th century, which "tended to become long-term" debt since they were continually rolled-over, see Gelderblom and Jonker (2011); for the key Italian States, see Gelabert (1999); Pezzolo (2008). The Crown of Aragon similarly issues *Prestechs* under Alfons V. and John II., which - while de jure short-term debts - are de facto long-maturity obligations, see Kuechler (1983). For the Holy Roman Empire - whose debt over centuries is synonymous with the Habsburg personal domain ("Erblande") debts, as there is no centralized Treasury - Huber (1893) provides a breakdown, by debt type, for the year 1564: 69.9% of Ferdinand I.'s debt is recorded in the form of long-maturity "pledge" debts, de facto consolidated and transferable instruments, with 30% in "unverwiesene" Schulden - often, but not exclusively short-maturity de jure unconsolidated claims (many of which are treated by all involved sides as de facto long-maturity negotiable claims. Much of sovereign debt activity in the Holy Roman Empire takes place on the level of the sovereign prince-electors: for detailed financial studies, including debt breakdowns, see for instance Schirmer (2006) for Saxony; France intensifies short-maturity issuance under Colbert and during the 18th century, erecting facilities such as the designated *Caisse d'emprunts*: such developments lag regular annuities issuance by two centuries, however, and do not assume comparable volumes, regularity, and credibility, see Velde (2008). Schmelzing (2022) provides further documentation and an overview of more than 330 short-maturity issuances between 1312-1850, together with nominal interest rates, and more detailed discussion.

[12]For the U.S. history, see in particular Hollander (1919): prior to the first irregular issuance of short-maturity certificates in 1812, the Treasury under Alexander Hamilton actually relies on bank overdrafts and loans to satisfy the short-maturity spectrum.

similar to U.S. certificates, between 1707 and 1853, various issuances of U.K. "Exchequer Bills" take place; in effect, revenue stream conversions by the Bank of England, with maturities typically of one year and with peak issuance during wartime (Richards, 1936). Prior to the founding of the Bank of England, "tallies" are regularly issued - intended as short-maturity IOUs funded upon anticipated tax revenues: de facto, the instruments are negotiable, but like Dutch "obligaties" they are in practice assuming effective long-maturity, through automatic roll-overs upon expiry at the prevailing long-maturity rate.

Overall, historically, long-maturity debt thus not only constitutes the "original" sovereign debt - it is the sole historically-consistent basis enabling a multi-century assessment of the time-series properties of sovereign interest rates. In this sense, a consistent and longer long-maturity time series allows much higher statistical power, particularly for example to discriminate between unit root and near unit root hypotheses. Short-maturity interest rate series are by construction far more restricted (a limitation amplified further for *policy* rates, where time series are even shorter). At the same time, we stress that we do not claim that our results will speak directly to the multitude of existing studies focusing on short-maturity rates: different time-series properties may affect different parts of the term structure.

3.2 Long-maturity data

We turn to the new data that now allows a critical expansion of interest rate samples, and potentially new insights into statistical interest rate properties. We draw upon Schmelzing (2022), who constructed a comprehensive sample of advanced economy long-maturity liquid interest rates since the "birth of sovereign debt" in the Renaissance. The series covers over 80% of advanced economy GDP on an annual level, across eight countries, and incorporates both marketable and relevant personal loan transactions. The inception year of 1311 for the series is not accidental: from this date onwards, both voluntary long-maturity nominal data points *and* suitable price data is available, allowing the calculation of inflation-adjusted, real interest rates.[13]

On the inflation side, our headline series utilize mainly the reference urban price series constructed by Allen (2001), as well as subsequent country-level improvements (e.g. Alvarez-Nogal and de la Escosura (2013) for Spain; Ridolfi (2019) for France). All data points on the inflation side refer to realized inflation rates of representative consumption baskets for workers denominated in silver equivalents - there exists as of yet no convincing approach to assess inflation expectations in the early modern period, but lagged measures of realized inflation rates are widely regarded as a close approximation of inflation expectations, including in the literature of particular relevance to our results.[14] We opt for a construction of an ex ante real rate measure, and follow the approach in

[13]A version of key long-term interest rate and inflation data we use was posted by Schmelzing via the Bank of England in 2020 (https://www.bankofengland.co.uk/working-paper/2020/eight-centuries-of-global-real-interest-rates-r-g-and-the-suprasecular-decline-1311-2018), and earlier long-dated price inflation data can also be found in Reinhart and Rogoff (2009).

[14]Recall that most previous studies - including Rose (1988) - treat realized past inflation as a proxy for inflation

Homer and Sylla (2005), who deflate the nominal rate in year t with a seven-year lagged inflation measure to approximate inflation expectations (over t-7 to t-1, not including the current year t), using progressive weights. We also smooth a few extreme inflation outliers, mainly associated with the 20th century, post-fixed exchange rate system, to avoid the severe measurement issues for these years.[15] Given its relevance, in the appendix we test two alternative approaches to construct inflation expectations, adopting methodologies presented in Eichengreen (2015) and Hamilton et al. (2016). The confirmation of our ADF-GLS results in these variations suggest that the particular deflator approach does not crucially change the properties we posit below. Of course, a virtually unlimited menu of alternative expectations constructions are possible, and the issue merits investigation, but the ones we use are well established in the literature and certainly form a plausible benchmark.

The effective average maturity of the assets over time closely approximates modern ten-year long-maturity Treasury bonds, and draws upon a wide sample of primary, printed primary and secondary sources. A large sample of non-voluntary sovereign debt operations - including intra-governmental lending, de facto confiscations, war indemnities, and office sales - undertaken by all advanced economies in the early modern period is separately identified and excluded from headline series. On average, the dataset covers more than four long-maturity data points per country-year between 1311-1870 on the consolidated sovereign debt side alone, in total over 12,000 long-maturity data points pre-1870.[16] It thus represents a significant empirical leap over previous reference works such as Homer and Sylla (2005). Focusing only on the most recent period of 1870-2016, a related undertaking is represented by "JST" (Jorda et al., 2017). As noted earlier, a substantial share of observations in "JST" ultimately traces back to Homer and Sylla (2005), with its associated gaps and inconsistent definitions. Besides covering a much shorter time span than the Schmelzing (2022) data, there are issues of consistency for example where JST/Homer and Sylla fuse yields for both medium- and long-maturity sovereign instruments to create a "long-term bond yield"; moreover data for key countries is not observed directly, but based on estimated risk

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expectations. A range of robustness tests on the inflation side are presented in Schmelzing (2022), including on local currency denominations, "elite" consumption baskets, or long-run country-level inflation series, available via https://scholar.harvard.edu/pfschmelzing/jmp-dataset.

[15]In particular, we follow Homer and Sylla (2005) by attaching a 33% weight to the t-1 CPI annual figure, and decreasing this weight successively over t-2 to t-7. For the severe inflation shocks, we interpolate the last "pre- shock" CPI figure to the first "post-shock" CPI figure for the hyperinflation years in Germany (1916-1923), Spain (1937-1939), Italy (1943-1945), and Japan (1944-1947).

[16]No assets with a contracted maturity below two years are included at any point in the dataset. The all-time average effective maturity of non-perpetual assets in the sample is stated as 11.3 years: for full details, see in particular the appendix of Schmelzing (2022), available via https://scholar.harvard.edu/pfschmelzing/jmp-dataset. Schmelzing (2022) also attempts to isolate a "safe" interest rate over time, which splices successive country-level long-maturity real rates and traces the ex ante contemporary "dominant" economy, a series which aligns with secondary literature and features not a single default event on principal over seven centuries. Other "safety" features associated with the interest rate data are tested in (ibid.), including the frequency of debasement operations, military and geopolitical power of the issuing sovereign, and the absolute economic size (output) of the issuing sovereign. While of importance for the eventual identification of positive real rate drivers, neither the existence or the particular level of risk premia over time represent relevant issues for our endeavors in this paper.

premia. See Schmelzing (2022), whose newer data work takes advantage of recent advances in economic history to address these limitations, and is based largely on primary sources.[17]

Our data therefore traces the only consistent and high frequency definition of sovereign real interest rates over centuries, allowing us to take full advantage of the significantly higher statistical power associated with a multi-century sample length. Figure 1 displays the resulting headline global real rate series on this basis, together with its structural trend over time: this GDP-weighted basis, weighting each of the eight economies according to its rolling GDP share, is our default basis to report "global" results, unless otherwise noted (this basis is abbreviated "Global GW" in the following).

The blue structural trend line already suggests visually that the series appears to exhibit a downward trend over time: apparently, the oft-invoked low interest rate environment of recent decades therefore has deep historical roots. Specifically, the series displays a linear downward trend slope that declines by just over 1.6 basis points (1/100th of one percent - thus amounting to a decline of 1.6% every century) per annum over the observation period (1311-2021). But the next section will proceed beyond "visual" impressions and confirm statistical properties in more detail.

4 Statistical Properties

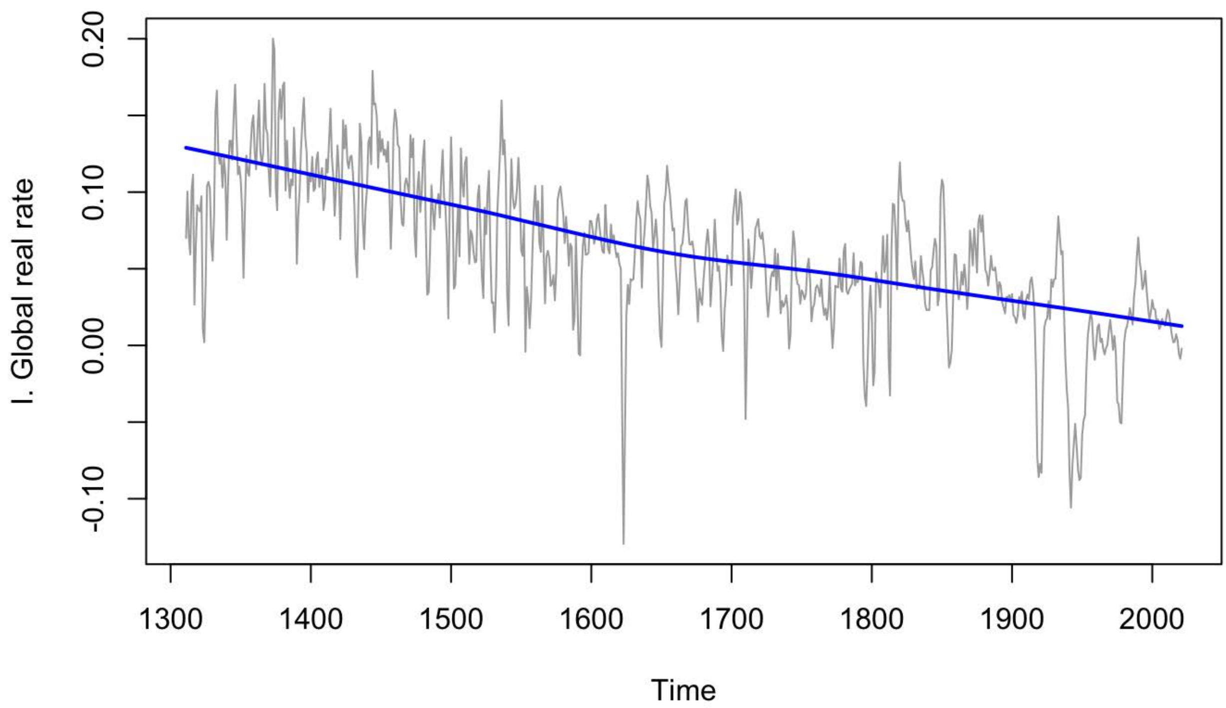
We undertake an Augmented Dickey-Fuller test (ADF-GLS), as modified by Elliott-Rothenberg-Stock (Elliott et al., 1996), and also test our series for structural breaks over a set of historically well-known break dates following Chow (1960). Finally, we conclude this section with an investigation of long-run persistence (half-lives) of global real rates. A range of alternative specifications of the ADF-GLS (including a specification assuming no time trend) and related robustness tests (including results from a Monte-Carlo simulation, and a Bai-Perron test) are to be found in the appendix, broadly confirming our results.[18] Not least, we present the evidence for short-maturity real rates, using the shorter horizons available - we are able to essentially corroborate previous studies using the same approach that we utilize to demonstrate long-maturity trend-stationarity over long horizons.[19]

[17] For instance, in Jorda et al. (2017) German, Japanese, and Italian government long-maturity bond yields prior to 1913 are based on an unpublished estimate of risk premia over British consols; later periods for Italy splice together evidence from "medium-term" and "long-term" observations. Similarly, the source for Dutch and German yields between 1880-1913 in "JST" is given as Flandreau and Zumer (2000) - but these authors in turn provide as their source for German and Dutch yields Homer and Sylla (2005) - France over 1870-1947 is also fully based on the latter.

[18] See Appendix table A.2 for ADF-GLS without time trend. A companion paper, provisionally titled "Uncovered Interest Rate Parity Across Long Term Bond Markets in the Very Long Run", is set to examine very long-run properties of the separate nominal rate and inflation components in detail.

[19] Rose (1988) uses short-maturity U.S. commercial paper rates, among other rates - we can replicate his failure to reject the unit root for all 8 possible variations on short-maturity real rates (in addition to replicating all other failures to reject), using multiple inflation approaches, see appendix tables A.6.2 and A.6.3.

Figure 1: Headline global real rates, and trend, 1311-2021



Notes: Data based on Schmelzing (2022)'s GDP-weighted global real rate data. Long-maturity ex ante basis, deflated by using seven-year progressively-lagged inflation, excluding current year *t*. Structural trend (in blue) displayed as lowess function: the lowess (locally weighted average) function fits a non-parametric model, which is the "best fit" given the time series data.

4.1 ADF-GLS

Per table 1, we find unambiguous evidence in favor of trend stationarity for global real rates when tested via the ADF-GLS test. The null hypothesis of a unit root can be rejected at the 5 percent significance level for the headline real series (using GDP-weights in any given year, "Global GW" - this is the default basis on which we report all "global" results unless otherwise noted), for all headline real rate country series, and for an arithmetically-weighted global series ("Global AW"), which gives equal weight to each of the eight constituent countries in any given year, regardless of absolute size of GDP. Specifically, we also note that the evidence in favor of trend stationarity is confirmed for U.S. and U.K. country-level data, with which some researchers might be more familiar: the fact that the results hold up for these variations, and a variety of different inflation and weighting bases, gives us a high degree of confidence in the robustness of the above methodology. Specifically, the fact that we reject a unit root using both the ADF-GLS test specification without a time trend as well as the specification with a time trend makes us confident that these results are not a function of the particular model used: the results, instead, appear to be firmly a function of the extended sample length, the fact that we exclusively focus on long-maturity rates, or a combination of both factors.[20]

Further, we test a range of alternative asset price series separately for the more recent period in the appendix (section 1.5), including testing private real long-maturity rates, and real long sample policy rates. The results broadly confirm our propositions above, and appear to indicate that the ability to reject non-stationarity appears to extend to other fixed income series once long samples are employed. The intuition that our powerful sample length (rather than idiosyncrasies of our data sources, maturities, or geographies) are driving our results is also suggested by Monte Carlo simulations where we generate time series with the same properties (mean, trend, variance and serial correlation) as our global real rate series and, by varying the sample length, we demonstrate that the power of the ADF-GLS test significantly falls when sample sizes shorten. That is, researchers working with samples of progressively smaller sizes are significantly less likely to reject non-stationarity (a more detailed discussion is available in appendix section 1.6).

[20]We detail the results for ADF-GLS without time trend in appendix table A.2: specifically, we fail to reject stationarity in this exercise for countries for which we subsequently identify structural breaks (when allowing but not imposing breaks with Bai-Perron, such as for Spain).

Table 1: ADF-GLS Test with Time Trend			
Real Rate Series			
Region	Number of lags	ADF-GLS test statistic	Optimal lag
Global GW	3	-6.420	MAIC
	2	-6.776	Seq, SIC
	1	-9.499	
Global AW	3	-6.549	
	2	-6.751	Seq, SIC, MAIC
	1	-9.119	
Italy	3	-9.937	MAIC
	2	-10.886	Seq, SIC
	1	-14.422	
UK	3	-6.763	
	2	-6.792	Seq, SIC, MAIC
	1	-8.956	
Dutch	3	-8.812	Seq, MAIC
	2	-10.356	SIC
	1	-13.192	
France	3	-6.017	MAIC
	2	-6.456	Seq, SIC
	1	-8.308	
Germany	3	-8.244	
	2	-8.184	Seq, SIC, MAIC
	1	-11.577	
Spain	3	-5.722	
	2	-5.742	Seq, SIC, MAIC
	1	-7.846	
U.S.	3	-5.420	
	2	-6.011	
	1	-6.188	Seq, SIC, MAIC
Japan	3	-4.034	
	2	-3.953	Seq, SIC, MAIC
	1	-5.258	

Note: the table reports the ADF-GLS test statistic for several choices of the number of lags (with a maximum of three 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. "Global GW" = GDP-weighted global real rate basis; "Global AW" = arithmetically-weighted global real rate basis, taking equal weights for each of the eight countries.

4.2 Zivot-Andrews test

Next, we report the test by Zivot and Andrews (1992), which tests for non-stationarity while being robust to the presence of structural breaks in the trend. A nice feature of the test is that it also estimates the most likely break date, in case a break exists. However, we are inclined not to overinterpret these dates since the test fits the data to (at most) one break, whereas in the Chow tests below, or the Bai-Perron tests reported in the appendix, we allow for multiple breaks. Per table 2, the unit root is rejected once more for both global real rates, as well as for all country-series. The most likely break date is reported for all respective series; for completeness, we briefly discuss the historical context of these days in appendix section 1.3. The next section 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 real interest rates over time.[21]

Table 2: Zivot-Andrews Test Results

	Test statistic	Break date
Global Real GW	-9.727	1937
Global Real AW	-8.948	1803
Italy Real	-12.581	1916
UK Real	-10.164	1553
Dutch Real	-9.497	1437
France Real	-8.466	1937
Germany Real	-10.769	1628
Spain Real	-8.235	1814
US Real	-6.610	1850
Japan Real	-4.701	1938

Note: the table reports the Zivot and Andrews (1992)'s 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. When the test rejects the unit root, the column labeled "break date" reports the estimated break date. "Global GW" = GDP-weighted global real rate basis; "Global AW" = arithmetically-weighted global real rate basis.

4.3 Chow Tests

We present Chow tests in table 3, choosing break dates in years where the existing literature would suggest or recognize high probabilities of expecting a trend break. The Chow test is implemented

[21] With that in mind, we interpret the full historical-economic context for the dates in table A.2 in appendix section 1.3. We provide a full Bai-Perron structural break test in the appendix (see table A.1 and the discussion there).

in a model that includes a constant and a deterministic trend (as a fraction of the total sample size), using Newey and West's HAC standard errors with a maximum lag length of 4. Again, as with all other exercises in this main body, we use the progressively-lagged realized inflation measure to construct real rates, excluding the current year t, following the methodology in Homer and Sylla (2005).[22]

We choose five annual dates (1349, 1557, 1694, 1914, 1981) to perform the Chow test: these dates are chosen on the basis of long-standing debates in the economic and financial history literature - literature that suggests to us a high degree of ex ante plausibility for each of these dates, as they all represent events of deep structural changes.

Per table 3, the Chow test results indicate that for the GDP-weighted global real rate ("Global GW") we find empirical evidence in favor of breaks in the trend or the mean in 1349 or 1557. For the equal-weighted global real rate ("Global AW", not separately shown here), we find empirical evidence of breaks in either the trend or the mean in the years 1349, 1557, or 1981.[23]

[22] An alternative deflator approach, using seven-year equally-lagged inflation including current year t, in the spirit of Eichengreen (2015), is presented in the appendix, section 1.2, table A.4.

[23] On the country level, the empirical evidence points towards the existence of a break at the following times: for the Italian real rate in 1694 in either the trend or the mean; for the UK real rate in 1349 and 1557 in either the trend or the mean; for the French real rate in all the dates in either the trend or the mean; for the German real rate in 1981 in either the trend or the mean (and marginally in 1557); for the Spanish real rate in 1694 and 1914 in either the trend or the mean , and in the mean in 1557 as well; for the Dutch real rate in 1981 in either the trend or the mean; for the US real rate in 1981 in either the trend or the mean as well. There is no evidence of breaks in the Japanese real rate.

Table 3: Chow Test Results - Global Real Rate GW - progressively-lagged inflation						
	Coefficient	Std error	t-statistic	P-value	95% Confidence interval	
trendbreak 1349	1.649739	.3034669	5.44	0.000	1.053923	2.245555
trendbreak 1557	-.1732261	.0720339	-2.40	0.016	-.3146549	-.0317974
trendbrak 1694	-.026058	.0671763	-0.39	0.698	-.1579495	.1058334
trendbreak 1914	-.1531424	.3122241	0.49	0.624	-.459867	-.7661518
trendbreak 1981	.4594748	.4091601	1.12	0.262	-.3438552	1.262805
meanbreak 1349	-.0890055	.0141602	-6.29	0.000	-.1168072	.0612038
meanbreak 1557	.0690446	.0275366	2.51	0.012	-.0149802	.123109
meanbreak 1694	-.0229655	.0350931	0.65	0.513	-.045935	.091866
meanbreak 1914	-.0869152	.2854017	-0.30	0.761	-.6472625	.4734321
meanbreak 1981	-.4911894	.3866844	-1.27	0.204	-1.250392	.2680126
trend	-.6291718	.2674622	-2.35	0.019	-1.154297	-.1040463
constant	.6330557	.2629259	2.41	0.016	.1168366	1.149275

Note: 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 HAC procedure with a maximum lag length of 4. The coefficient associated with "trendbreak 1349" denotes the difference of the estimated trend coefficients before and after a break in 1349 (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 1349. Similarly, "meanbreak 1349" refers to a break in the mean in 1349.

We observe that the years 1349 and 1557 are confirmed as breaks - a result broadly echoed in the country level results, but on the other hand, we find that the years 1694, 1914, and 1981 cannot be confirmed on the "Global GW" Chow basis: all of which are years where existing literature has posited strong claims that would expect one to find such breaks. While finding outright strong regularities for these confirmations and rejections, respectively, may not be straightforward, the most promising approach remains a better understanding of the exact historical and economic context over time, in the context of existing research. The following list thus asks: what is the particular logic of selecting these dates? And what do our results suggest for the economic-financial context of the five break years?

- **1349.** This year was chosen against the backdrop of a long literature positing a major financial trend break associated with the Black Death (Epstein, 2000; Pamuk, 2007; Clark, 2016). The death of one-third to one-half of the population is said to have created a boost in capital-per-capita, and a substantial increase in real wages resulting from labor scarcity. As we detail below, we confirm the epochal role of the Black Death on credit markets: however, the directional context pre- and post-1349 appears more idiosyncratic than recognized thus far.

- **1557.** While little progress on our understanding on the subject has been made since an authoritative study appeared in German in 1896 (Ehrenberg (1896), only partially translated via Ehrenberg (1928)) - which regarded the initial wave of shocks over 1557-62 as events "that shook the finance and trade of Europe to its foundation" - a variety of case-study and piecemeal literature appearing since confirmed that an epochal crisis afflicted the international economy during the second half of the 16th century. Indications are that the "triple sovereign default of 1557-8" (France, Spain, and the States General) could have been either a consequence or the actual inception point of a very deep-seated reversal in financial markets lasting at least to the early 17th century.[24] In direct consequence of the sovereign defaults, by far the largest bank of its day - perhaps the most significant "systemically important financial institution" ever to have existed - the Fuggers, slid to the brink of default. Narrowly escaping themselves, a seemingly unending string of international merchant and merchant-bank defaults can afterwards be traced in the respective sources. Recurring chaos at the largest financial fairs over decades - from Seville over Lyon, to Rome - is reported in equal measure (Ehrenberg, 1928; Lapeyre, 1955; Delumeau, 1959; Kindleberger, 1998).[25] We chose this date mainly to test for a major early modern standalone sovereign default-cum-private financial crisis event: our affirmative results on a break date here indicate that secularly, though financial turmoil on its own may still not be a sufficient driver of "regime changes" - at least when accompanied by significant political volatility (the French Wars of Religion; the near-uninterrupted Habsburg military campaigns in the 16th century) these combined forces can culminate into deep structural change.

- **1694.** This year was chosen against the backdrop of the prominent "North Weingast" thesis, which associated the Glorious Revolution with a revolution in "credible commitments" (North and Weingast, 1989). It appears particularly relevant that the authors used (nominal) interest rate data to prove their seminal thesis. Yet, this event is widely rejected in our Chow tests, on both the global levels and all country levels. Of course, the original statement of the North-Weingast thesis received extensive criticism over the years (Sussman and Yafeh, 2006) - but has not been contextualized in the setting of a comprehensive (real) interest rate dataset covering centuries. The failure to identify 1694 with a structural break is part of more comprehensive evidence that fails to associate all other central bank inceptions with structural breaks, too.[26]

[24]The Netherlands are not a formally recognized independent nation state until the Peace of Westphalia 1648: de facto, however, the States are characterized by a high degree of autonomy, with the Receiver General of the States General, and the individual provinces, issuing their own debt secured by land and other taxes, the *bede* Fritschy (2017, chapter 2). Ehrenberg (1928, 113f.) puts the Fugger exposure alone at 600,000fl., which the Receiver General refused to honor in 1557: it is not specified in the source whether the debt was denominated fully in Dutch guilders or Ducats. Hauser (1930) dates the French default to September 1, 1557, with other sources dating the default to the year 1558.

[25]Ehrenberg (1928, 114ff.) details the uncertainty over the Fugger's survival in 1557-8, and also presents balance sheets of the bank: he indicates that the Fuggers had formally written off more than 620,000fl from the Dutch and Spanish defaults by 1563, against a balance sheet size of 5.6M fl. Ehrenberg regards 1557 as the beginning of the end of the Fugger empire. Significant parts of his work remain untranslated; the contemporary account of Hauser (1930) relies mainly on him.

[26]The same failure to associate central bank inceptions with structural breaks is observed when undertaking Bai-Perron tests, see appendix table A.1.

- **1914.** This year was chosen against the backdrop of the strong narrative of an institutional inflection arising from the founding of the Federal Reserve (Barsky et al., 1988; Bernstein et al., 2010), coupled with the major monetary inflection related to the (de facto) departure from the centuries-old bullion standard. In addition, the economies deal with the geopolitical shock of the First World War - the first industrial war far outstripping the human and financial costs of earlier inter-state conflicts. And indeed, across our series, the years in and around the First World War constitute strong (though not fully consistent) breaks across global and country-levels: ironically, however, the U.S. is not one of these. As in 1694, we therefore fail to confirm any obvious monetary channel influencing real rates.

- **1981.** This year was chosen against the backdrop of prevalent narratives - intensifying after 2008 - of a key inflection point in advanced economies and financial markets during the early 1980s, though there is no clear consensus on a single driver of this alleged inflection (Rachel and Summers, 2019; Mian et al., 2021; Goodhart and Pradhan, 2021). Antedating the 2008 crisis, a sizable literature found evidence of inflation and interest rate "regime changes" during the late 1970s or early 1980s (Garcia and Perron, 1996; Ang and Bekaert, 2002; Neely and Rapach, 2008) - coupled with sharply declining general macroeconomic volatility (Bernanke, 2004) - invoking monetary policy dynamics as the driving causal force.

4.4 Persistence

We measure persistence and half-life next. To the extent that the existing literature has investigated whether the post-1980s behavior of global real rates is "transitory" - whether rates will eventually return to "a more normal level" (Laubach and Williams, 2016), long-run estimates of such adjustment speeds can provide an indication of how quickly global real rates typically return to their structural trend after shocks. As per table 4, which uses our Chow test dates to delineate regimes, we observe that the half-life of real interest rate deviations for the entire sample period ranges between 1 and 10 years across individual regimes, with one sharp outlier in the case of France. Breaking down this figure for sub-periods reveals that half-lives stood below two years during the fourteenth and fifteenth centuries, a figure that moderately rises over the remainder of the bullion regime era, after which it sharply increases over the late gold standard and early gold-exchange standard period, only to moderate again towards the end of the 20th century, to levels slightly (but not sharply) above long-run averages seen during the early modern period. This general curve is equally visible if alternative weighting methods are employed ("Global AW"), and if country-level trends are investigated separately. The "safe asset providers" over time - Italy pre-1600, the U.K. after 1557, the U.S. in the 20th century - appear to record consistently faster adjustment speeds. It is important to note that our relatively short half lives (around two years) do not contradict the fact that we need a relatively large sample in order to have sufficient power to reject non-stationarity, as we show in our Monte Carlo simulations in the appendix (section 1.6).[27]

[27] In fact, our Monte Carlo simulations, tailored to our estimated global real rate series, show that a researcher would reject non-stationarity only 77 percent of the times in a sample of 100 observations, and only 37 percent of the times in

In terms of levels, our data suggests that early modern persistence in global real rates appears to be broadly comparable to respective persistence in real exchange rates - but generally higher than identified persistence in real exchange rates after 1914, from which half-lives of 3-5 years have traditionally been posited (Rogoff, 1996; Lothian and Taylor, 1996; Itskhoki, 2021). [28] A decomposition of aggregate persistence measures at the country level does not reveal notable cross-country differences: while aggregation biases similar to those posited for real exchange rate persistence measurements are therefore not apparent.[29]

a sample of 50 observations, while he/she would reject it with probability almost equal to one in our sample of 711 observations. See appendix section 1.6 for details.

[28]The uncertainty around our estimated half-lives is much lower than the substantial uncertainty surrounding estimates of the half-lives of real exchange rates, due to the high persistence of the latter.

[29]We intend to discuss nominal and inflation persistence breakdowns separately in a companion paper provisionally titled "Uncovered Interest Rate Parity Across Long Term Bond Markets in the Very Long Run".

Table 4: Half-Lives of Real Rates, by sub-samples

	Period	Half-life	Confidence interval	
Global Real GW	1311-1349	1.2068	0.8219	1.7101
	1350-1557	1.4674	1.2171	1.7141
	1558-2021	2.7947	2.3746	3.7538
Global Real AW	1311-1349	1.4057	1.0381	1.8739
	1350-1557	1.4684	1.2141	1.7878
	1558-1981	2.9377	2.3892	5.1246
	1982-2021	3.1947	1.7151	8.9091
Italy Real	1311-1694	1.3487	1.1968	1.5028
	1695-2021	2.9033	2.3573	4.1928
U.K. Real	1311-1349	1.4533	1.1059	1.8816
	1350-1557	1.4208	1.1253	1.7764
	1558-2021	1.6769	1.4417	1.9141
Dutch Real	1367-1981	1.4007	1.2499	1.5634
	1982-2021	3.5355	2.0216	7.6561
France Real	1311-1349	0.6484	0.4825	1.0242
	1350-1557	1.4689	1.2555	1.7063
	1558-1694	1.9536	1.3281	3.0483
	1695-1914	2.3102	1.9086	2.7919
	1915-1981	7.9032	3.6772	38.5262
	1982-2021	6.1940	1.7688	N.A.
Germany Real	1311-1981	2.3343	2.0484	2.6317
	1982-2021	3.1061	1.4449	10.1295
Spain Real	1332-1557	1.4091	1.1292	1.7348
	1558-1694	1.9593	1.3370	3.0222
	1695-1914	1.9299	1.6413	2.7374
	1915-2021	5.3997	3.4453	8.8593
U.S. Real	1791-1981	3.2676	2.5545	4.2219
	1982-2021	0.7855	0.5515	1.2806
Japan Real	1870-2021	2.8812	2.1358	5.4815

Note: The half-life is estimated as the first horizon at which the impulse response equals one-half of the initial impact effect. The impulse response is estimated using a deterministically detrended linear autoregressive model (with the number of lags chosen by the Bayesian Information Criterion) using Kilian (1999)'s bootstrap; the point estimate is based on the median unbiased response while the confidence interval is based on the 5th and 95th quantiles of the bootstrap distribution. The half-lives are separately estimated in each sub-sample identified by the statistically significant break dates (according to the Chow test). "Global GW" = GDP-weighted global real rate basis; "Global AW" = arithmetically-weighted global real rate basis. NA denotes situations where the half-life is infinity.

Generally, our persistence levels are consistent with related financial evidence from earlier studies, but they also suggest intriguing differences. Volckart and Wolf (2006), focusing on FX bullion content deviations between Flanders, Lübeck, and Prussia over 1385-1450 to assess money market integration found that arbitrage dynamics generated a half-life of eight months for the correction of silver price deviations between Flanders-Lübeck, and a half-life of 21 months for deviations between Lübeck-Prussia. Volckart and Wolf (2006) further point towards a secular trend towards sharply lower half-lives, with the classical gold standard period reporting FX arbitrage speeds (half-lives) of six days between London and New York in demand bill markets.

On the price side, several studies have investigated long-run trends in price dispersion and persistence: most recently, Federico et al. (2021) presented new evidence on both price co-movement, and price convergence in European grain markets over 1350-1910. Trend-wise, the authors find a slow but steady improvement of market efficiency and integration, reflected by increasing co-movement of agricultural prices and fewer price deviations, including intra-national deviations - dating the onset of price convergence in advanced economies to the fifteenth century.[30]

Against this evidence, our *levels* of persistence generally fall in line with the lower end of the range of estimates for related financial variables - however, we are unable to confirm any evidence of secularly faster adjustment speeds. On the contrary, real rate persistence on our basis appears to exhibit a wave-like trend over the early modern era, with some evidence of a "hump shape" over the 17th century. A sharp upwards break in half-lives then coincides with the end of the fixed regime in 1914, followed by a moderation after 1981 (to a present level above the early 20th century). This finding is relevant in the context of existing literature that associated the eve of the First World War - coinciding closely with the founding of the Federal Reserve and the departure from fixed exchange rate regimes - with a decline in financial volatility and seasonality in interest rates (Barsky et al. (1988); Bernstein et al. (2010)): while the management of seasonality, and general predictability in financial markets may both have improved, this did not necessarily imply faster adjustment speeds. Overall, the figures indeed suggest that, at least until the end of 2021, post-GFC global real rates have "under-corrected" relative to historical adjustments (given that they remained visibly below structural trend levels and persistence confidence bands after well over a decade as of today) - however, as of yet this multi-year deviation does not register as a structural break: and it cannot necessarily be expected to do so even in the event of a further downward entrenchment of rates, given the rare occurrence of structural breaks we identified (occurring on average every one-and-a-half centuries). An ultimate "normalization" to the structural downward trending level would therefore be the default expectation on the basis of our data.

To summarize: this section suggests that testing long horizon samples of real interest rates strongly

[30]Previously, Froot et al. (1995), Persson (1999) and Bateman (2011) have calculated persistence for European grain markets over time, with the former positing half-lives of 2.8 - 6.2 years for Dutch and English food price deviations (law of one price (LOR) basis) over a 700-year period, with no pronounced change in the modern or float periods; and Persson (1999) finding 1.5 - 3 years in the early modern period, with a secular trend towards faster adjustments. Using European wheat prices, Bateman (2011) found lower half-lives, on average 6.34 months, over the period of 1550-1750, with no pronounced trend.

confirms their stationary properties around a downward deterministic trend - in contrast to existing literature that focused on short-maturity rates, using short samples. Our results are robust to a range of specifications, including alternative deflator approaches and the existence of potential structural breaks, all of which are described in detail in the appendix. Importantly, we also present ADF-GLS tests for short sample and short-maturity series in the appendix (appendix tables A.6.1, A.6.2, and A.6.3): we can replicate results in previous literature via this exercise, demonstrating that it is indeed impossible to reject a unit root using such shorter horizons (keeping other attributes constant). Our results on the persistence are novel, too, with the finding of a slight increase in half-lives in global real rates presenting somewhat of a challenge in the context of existing literature. Taken together, these results rationalize why the existing consensus on real interest rate properties has arisen: but they equally demonstrate that previous results do not necessarily hold once a higher statistical power is achieved by using longer samples - albeit we emphasize that if the term premium is non-stationary, it is perfectly possible for the long-term rate to be stationary and the short-rate not, at least over a prolonged period. The long-maturity real interest rate appears to exhibit quite distinct properties when compared to other financial variables - notably real exchange rates (RER) (Itskhoki, 2021).

5 The continuity of real rate trends, Twentieth Century breaks, and low interest rate eras over time

From the discussion of statistical properties, we now proceed to an attempt at broader interpretations of historical patterns and possible economic drivers. How do the statistical results accord with economic and historical evidence, and to what extent do structural breaks and general trends fit the empirics of other macroeconomic variables?

5.1 The 1914 break, and assessing the (alleged) 1981 inflection

We concentrate here on a discussion of the much-discussed significance of "recent" trend breaks in global real rates, particularly in 1914 and the 1980s. We recall that in our (unbalanced) panel above, we found limited evidence of both 1914 and 1981 as breaks - neither assumed the constancy of 1349 or 1557 as relevant inflection points.

In tables 5.1 and 5.2. we now test for structural breaks in a balanced panel, to take account of the fact that the U.S. and Japanese series begin notably later than the European observations (in 1790 and 1870, respectively). In this exercise, we initiate all series in 1870 and end in 2016: below, the results are shown for the GDP-weighted global real series ("Global GW"), and for the arithmetically-weighted global real series ("Global AW"). We test the 20th century break points (1914 and 1981) once more in this set-up via the Chow test.

Once we focus on this balanced panel, both 1914 and 1981 emerge as a more relevant break date, though still not as quite a unanimous one. Specifically, for Global AW, we now find breaks in both 1914 and 1981. For Germany, France, Holland, and the U.S., we now record 1981 as a trend break.[31]

Table 5.1: Chow Test Results - balanced panel Global GW - progressively-lagged inflation

Regressor	Coefficient	Std error	t-statistic	p-value	95% Confidence interval	
trendbreak 1914	-.1450128	.0756641	-1.92	0.057	-.2945955	-.0045699
trendbreak 1981	.0678054	.090859	0.75	0.457	-.1118166	.2474274
meanbreak 1914	.0619555	.0423217	1.46	0.145	-.0217116	.1456225
meanbreak 1981	-.1069321	.0725561	-1.47	0.143	-.2503705	.0365063
trend	-.1028904	.0637047	-1.62	0.109	-.2288303	-.0230494
constant	.1153798	.0595556	1.94	0.055	-.0023576	.2331171

Note: 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 HAC estimator with a maximum lag length of 4. 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.

Overall, the idea that existing narratives around both the 1914 and the 1980s "interest rate inflections" are a function of a restricted sample length that existing research had thus far limited itself to, remains consistent with our data. We recall: besides the fact that our "imposed" Zivot-Andrews test did not identify any 1980s breaks, we also point to the fact that a Bai-Perron exercise allowing for multiple breaks only identifies a single structural break across all series after the year 1917, in the case of the U.S. (1983, see appendix table A.1). Our intuition therefore remains that narratives of a "recent", a 1980s inflection point in global real rates, are partly a function of too short a sample length that existing research was able to utilize (overall, we interpret our results to be a function of three factors: our approach of allowing for a trend in the data, of our much longer sample length, and of the fact that we focus on long-maturity rates).

[31] Given space, we do not report country results here. In addition to the above, we also find a break in the mean for the U.K. in 1914, for U.S., Italy, Spain, and France also in 1914. For Japan we find no break.

Table 5.2: Chow Test Results - balanced panel Global AW - progressively-lagged inflation						
Regressor	Coefficient	Std error	t-statistic	p-value	95% Confidence interval	
trendbreak 1914	-.1947143	.0790839	-2.46	0.015	-.3510578	-.0383709
trendbreak 1981	.1678928	.0946711	1.77	0.078	-.0192654	.355051
meanbreak 1914	-.0832306	.0480721	1.73	0.086	-.0118046	.1782658
meanbreak 1981	-.1669644	.0721482	-2.31	0.022	-.3095965	-.0243323
trend	-.1364817	.0583981	-2.34	0.021	-.2519307	-.0210327
constant	.1463777	.0541504	2.70	0.008	.039326	.2534294

Note: 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 HAC estimator with a maximum lag length of 4. 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.

5.2 Low real rate eras over time

An alternative to pinpointing particular break points is to identify secular departures from the structural (downward) trend of real rates in the context of historical evidence and literature - in other words identifying narrative "low real interest rate periods". We undertake such an approach in addition to the econometric approaches, cognizant of the qualifications such judgments must necessarily involve. We stress that we view these episodes as akin to "long-lived transitory shocks", not corresponding to break dates in the spirit of Chow and Bai-Perron: importantly, at the conclusion of these respective eras, the structural trend in each case returned to the long-run regularities. Put differently, the crucial observation is that these low rate eras all ended.

The resulting periods are visualized in Figure 2 as shaded areas. They include the following four episodes:

- **The (pre-) Black Death era (1311-1353)**. A long consensual literature viewed the most devastating demographic shock in recorded history as a key inflection point in financial markets (Homer and Sylla, 2005; Pamuk, 2007; Clark, 2016): and indeed, our Chow test indicated 1349 as a relevant break date, for the global series and for country-levels. In representative fashion, Stephen Epstein posited that "the Black Death saw a major change of trend in European interest rates which set in motion a gradual decline in the real cost of capital that lasted up to the eighteenth century. . . the fall in the expected rates of return and cost of capital for individuals was nearly as impressive" (Epstein, 2000, 61f.). However, our evidence suggests that the pre-Black Death era saw comparatively *low* real interest rates - only the early 1500s once more returned to rates prevailing two centuries prior. While we cannot be conclusive, we note that recent scholarship associates the years immediately preceding the Black Death with the protracted "Fourteenth Century Crisis": stagnant productivity growth and incomes, seigniorial and rural impoverishment, indeed "international commercial recession" - a context against which some literature even asserted that the Black Death led to a "relief" from highly adverse structural trends.[32]

- **Post-Bullion Famine (1483-1541)**. We discussed the intense money shortages associated with the "Bullion famine" dynamics in the context of our 15th century structural breaks above. Towards the end of the century, the re-capture of the Balkan mines by European armies, and the ability to resume coining output familiar from 14th century levels led to reversal, and a boom in advanced economy money supplies, with the classic account in Day (1978) still providing a plausible framework. More recent discussions include Munro (2003); Aerts (2006).

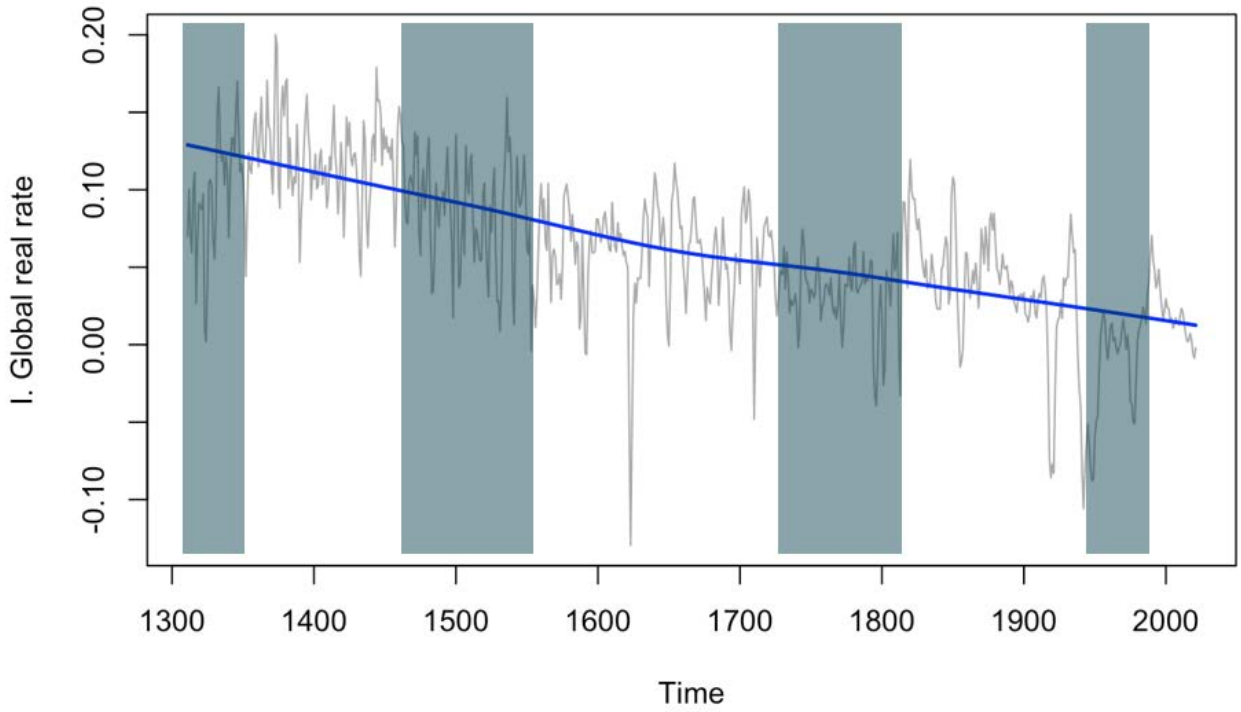
- **The 1732-1810 boom**. This second period broadly spans the time from the South Sea Bubble and the French "Law" Crash, to the Congress of Vienna. It is a period of - relative - geopolitical calm (to the 1790s), demographic and economic advancement, and institutional security. Evidence also abounds for the existence of a prolonged capital market boom spanning several decades, leaving traces in both sharply rising sovereign debt/GDP figures, and a boom in private credit (Neal, 2015, chapter 6), Hoffman et al. (2019).

- **The great FX transition era, 1937-1985**. Beginning with the prominent 1937 break, this period broadly spans the Bretton Woods regime and the early years of the successive global float (Eichengreen, 2019; Ilzetzki et al., 2022). In geopolitical terms, it almost exactly covers the Cold War era (covering the unprecedented reduction in actual inter-state military conflicts after the first use of nuclear weapons). It ends amid the onset of the great moderation, the introduction of inflation targeting by advanced economy central banks, and the dissolution phase of the Soviet Union.

Are there any regularities that would unite these four eras? Without developing formal models or claiming to be conclusive, we note, for one, that all four periods are marked by relative geopolitical calm: not necessarily in terms of conflict incidence, but certainly in terms of inter-state conflict *intensity*. This is even more apparent when we isolate the "Dominant" economy in respective eras, in other words the equivalent "safe asset provider": Northern Italy, prior to the arrival of the marauding "mercenary companies" from the mid-14th century, but after the peak of the Guelph-Ghibelline clashes, enjoys a phase of relative political idleness in the early 1300s (Caferro, 1998); Habsburg Spain under Charles V is almost uninterruptedly engaged in the struggle against the rival Valois, and takes an uncompromising stance against Protestant dissidents in the Germany prior to the Peace of Augsburg - however, these are low-intensity conflicts, nowhere near

[32] According to Campbell (2005), "the marginal productivity of male agricultural labour fell by 60 per cent between 1209 and 1309 and was lower between 1270 and 1329 than at any other time between 1209 and 1869...it was the massive and fortuitous loss of life in the Black Death of 1348-9 that took the demographic tension out of the equation and thereby broke the deadlock on the land".

Figure 2: Global real rates, structural trend, and low-rate eras, 1311-2021



Notes: Global real rates displayed as nominal long-maturity rates deflated by progressively-lagged seven-year realized inflation. The four low-rate episodes (shaded) are 1311-1353, 1483-1541, 1732-1810, and 1937-1985. For details, see text above. Periods are narratively defined.

comparable with the internal religious wars in France, or the geopolitical and religious escalations of the 17th century (Parker, 2019); for most of the eighteenth century, and only interrupted by the reluctant participation in the War of the Austrian Succession, the safe asset provider Britain under the dominating influence of the Duke of Newcastle retreated militarily into a "pose of studied insularity" (Scott, 2002). And of course, the second half of the 20th century - while marked by the intense ideological competition against communism - saw actual conflict intensity reach unprecedentedly low levels: perhaps not since the War of Saint Sabas in 1256 did the leading economy experience a longer stretch of abstention from direct inter-state conflict. While these periodizations can be underpinned by more quantitative evidence - including reference conflict chronologies by political scientists - as we reiterate later in this paper, all this does not imply that we see geopolitical risk premia as the driving force of secular real rate trends per se (or, more broadly, necessarily subscribe to a "rare disaster" framework that would be applicable here).[33]

While we are able to go well beyond anecdotal associations - geopolitical factors clearly appear to be a regular occurrence immediately before, after, and during our interest rate breaks - we are cautious to propose sweeping generalizations about the drivers of the structural breaks over time here. Whereas politico-financial crises appear to be a *necessary* component for a structural real

[33]Two such reference conflict chronologies include Clodfelter (2017) and Levy (1983), from the latter of which we can derive long-run statistical evidence on conflict intensity.

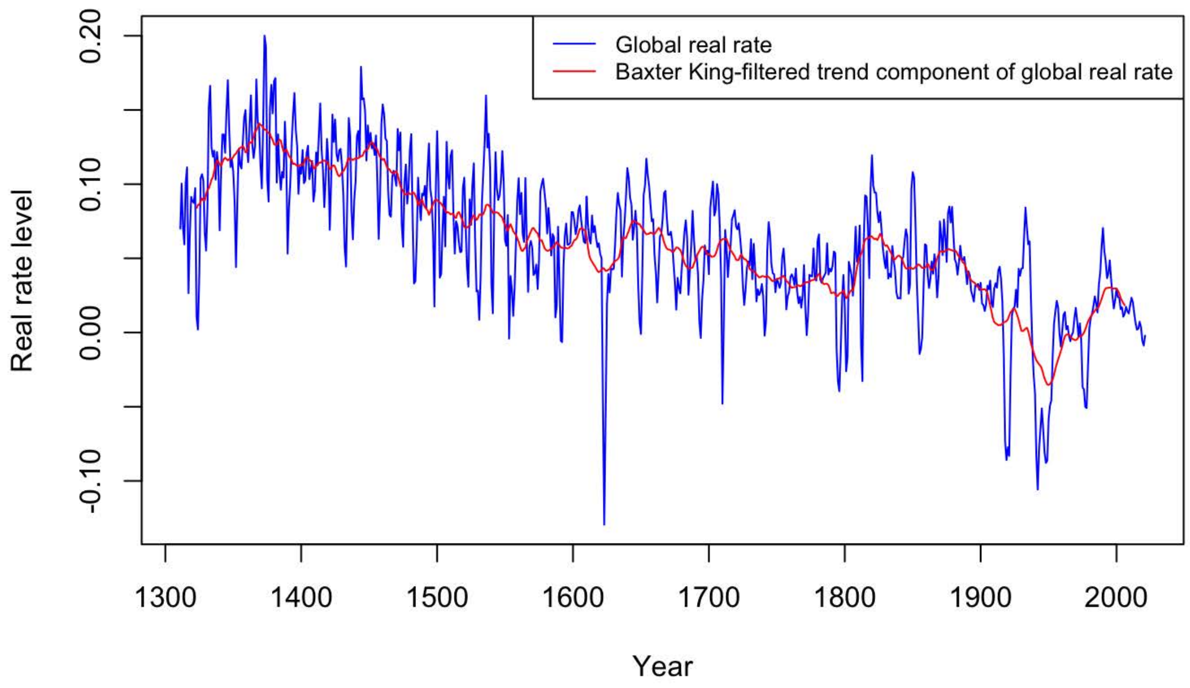
interest rate break, they appear to be far from *sufficient*. What appears to be certain, at least, is that the evidence for a primary role for "monetary regime changes", in the form of either policy rule shifts or key institutional events - long a focal point of economic and econometric literature - appears to be surprisingly poorly supported in the long context. As flagged, not least, the foundational dates of central banks (both de facto and de jure), such as 1609 in Holland, 1694 in Britain, 1876 in Germany, 1893 in Italy, or 1914 in the U.S. all fall outside of our structural break samples.

6 Real interest rates in the context of growth and demographic variables

The relationship between real interest rates and other macroeconomic variables has been of key interest in the literature for many years, but very long-run empirical investigations were hampered for the same reasons that prevented the study of very long-run interest rate trends: a lack of satisfactory data. Fortunately, over the most recent years in particular, significant advances have been made in the improvement of long-run output and population growth series - two variables of particular interest in the real interest rate literature. These advances allow us to relate our new evidence on real interest rates more granularly to output growth trends, and population growth over centuries. Table 6 reports simple period averages of our three key macro variables over time. All figures are based on an identically-weighted eight country sample. The early phase of the demographic sample still contains a non-negligible share of (decadal- or semi-centennial-level) interpolations - but this does not influence subsequent statements or calculated averages. It is evident from the table already that real interest rates and both population growth on the one hand, and real output growth rate on the other, move in divergent directions secularly. Afterwards, we discuss both variables in detail and undertake a few more formal exercises involving the use of a Baxter-King bandpass filter that removes short- and medium-run fluctuations (shorter than 100 years) from all variables. Figure 3 displays the effect of the Baxter-King filter, using the example of our global real rate sample. Beginning with long-run output growth trends, recent years have seen fundamental updates to the seminal work of Maddison (2010). While "big assumptions" (mainly involving interpolation) still have to be made in the underlying construction, new series have significantly improved granularity. New multi-century output reconstructions have been undertaken for Northern Italy (Malanima, 2011), France (Ridolfi and Nuvolari, 2021), Spain (de la Escosura et al., 2022), Germany (Pfister, 2021), Holland (van Zanden and van Leeuwen, 2012), and the U.K. (Broadberry and Fouquet, 2015). These typically rely on (wage-based) demand-side constructions, or more granular reconstructions of sectoral outputs.[34] The new developments are fortunate as they now allow a nuanced comparison both on the country-level and on the aggregate

[34] In "demand-side approaches", the GDP series are constructed on the basis of (combinations of urban and rural) real wage data - assuming that they reflect aggregate income trends when combined with estimates of urbanization rates, and of trends in days worked: the series for France (Ridolfi and Nuvolari, 2021), Germany (Pfister, 2021), and Northern Italy Malanima (2011) fall into this group. English and Dutch data (van Zanden and van Leeuwen, 2012; Broadberry and Fouquet, 2015), meanwhile, rely on sectoral output reconstructions. For a recent alternative approach using wealth data to construct Italian GDP benchmarks, see van Zanden and Felice (2022).

Figure 3: Unfiltered global interest real rates, and Baxter-King filtered real interest rates, 1311-2021.



Notes: Figure displays the progressively-lagged global real rate series (red), unfiltered. And a Baxter-King filtered version of the same series, applying a minimum and maximum oscillation period of 6 and 32 observations.

level of our real interest rate trends, and associated trends in output. To arrive at aggregate real output growth rates, we GDP-weighted the country constituent output growth rates based on the aggregate total GDP share for the benchmark years (creating GDP shares based on Maddison's benchmark GDP figures, and interpolating the calculated benchmarks): Figure 4 displays the resulting series over the period 1311-2021, together with our respective global real rate series. We report the long-run trend in these variables, extracted using Baxter and King's bandpass filter tailored to retain only fluctuations at frequencies longer than 100 years. We observe modest - almost static - output dynamics over the first half of the series, with aggregate growth driven mainly by (moderate) levels of population growth prior to the 18th century, though with some more vigorous dynamics on the regional level: individual growth spurts do take place in the early modern period, with Dutch real per capita incomes rising by 82% between 1500-1600, for instance, and a more general 29% real per capita output increase for our entire sample over 1500-1700. A sharp acceleration is detectable across all country levels during the 18th and 19th centuries, giving rise to the familiar Industrial Revolution dynamics. During the second half of the 20th century, a deceleration is notable, with current (post-2008) rates of growth meaningfully below trend lines: yet, absolute rates of growth in the early 21st century remain highly elevated compared to very long-run averages.

Table 6: Macro variables: averages by period (in percent per annum)

	1311-2021	1311-1500	1500-1800	1800-1914	1914-2021
Real aggregate output growth rate	0.93	-0.03	0.41	2.04	2.88
Aggregate population growth rate	0.30	-0.21	0.24	0.91	0.000
Real interest rate	6.02	10.85	5.61	4.61	0.19

Note: The table reports the average value for three variables real aggregate output growth rate; aggregate population growth rate; real interest rate. All three variables are based on an identically-weighted sample of our eight advanced economies over time. For full underlying output and population sources, see text and notes to figures 3 and 4.

Crucially - and as is clearly visible from the chart itself - the output growth series overall displays a monotonic *upwards* trend.[35] The positive correlation between output growth and real rates appears to be a rather unique phenomenon of the very most recent part of the data series, starting in the second half of the 20th century. Generally speaking, therefore, real interest rates and aggregate output growth rates exhibit trend dynamics in directly *opposite* directions over the secular term.[36] Taken together, these observations imply that output growth trends and real interest rate trends do not have a correlation in the direction commonly assumed. In fact, over extended periods of time over the long-run, both measures in fact go in opposite directions.

Demographic factors have been among the most frequently-invoked factors in the assessment of real interest rate dynamics in recent decades. Lunsford and West (2019) for instance consider demographic variables as the most plausible explanatory variables compared to growth and productivity variables - in particular, the authors find statistically significant roles for the share of 40-64-year olds in the population, as well as the dependency ratio, over the period 1890-2016, relying on U.S. data.[37] Gagnon et al. (2021) find that "demographic factors account for much, if not all of the actual permanent decline in the equilibrium real rate [since the 1980s]" - highlighting in particular falling fertility rates and mortality in the 1960s and 1970s as causal variables. But what does the (very) long-run data suggest for demographic context? A first-order reconstruction of population growth rates since the Renaissance can equally utilize Maddison (2010)'s data. But advances in output growth empirics since then have been mirrored by advances in demographic data for advanced economies, often undertaken by the same authors. New population estimates - from which we derive aggregate population growth rates over the long-run - are equally provided by Malanima (2011), van Zanden and van Leeuwen (2012), Broadberry and Fouquet (2015), Ridolfi (2019, building on Dupaquier (1988)), Pfister (2021), and de la Escosura et al. (2022).

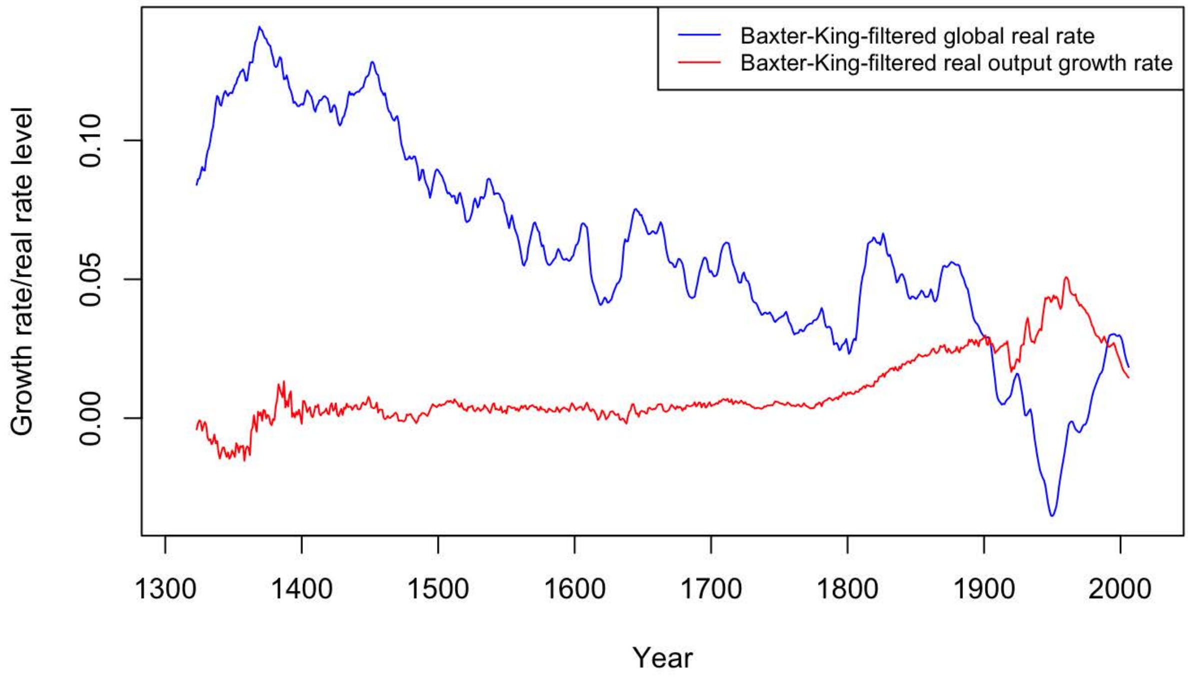
Figure 5 now displays long-run fluctuations in the aggregate population growth rate; as before, we

[35]This upwards trend is readily confirmed visually, but can also be obtained via standard tests such as Mann-Kendall (Mann, 1945; Kendall, 1955).

[36]Of course, historically, countries that have accumulated very high debt will often experience extended periods of slower growth and higher real interest rates, as suggested by Reinhart et al. (2012), and now the central conclusion of a large literature as surveyed in Abbas et al. (2019). This would not, however, explain the long-term opposing trends.

[37]In Lunsford and West (2019), the dependency ratio is positively correlated with real rates, while the share of 40-64-year olds is negatively correlated; the authors find no statistically significant role for life expectancy.

Figure 4: A comparison of Baxter-King filtered real aggregate output growth rates and real interest rates, 1311-2021.



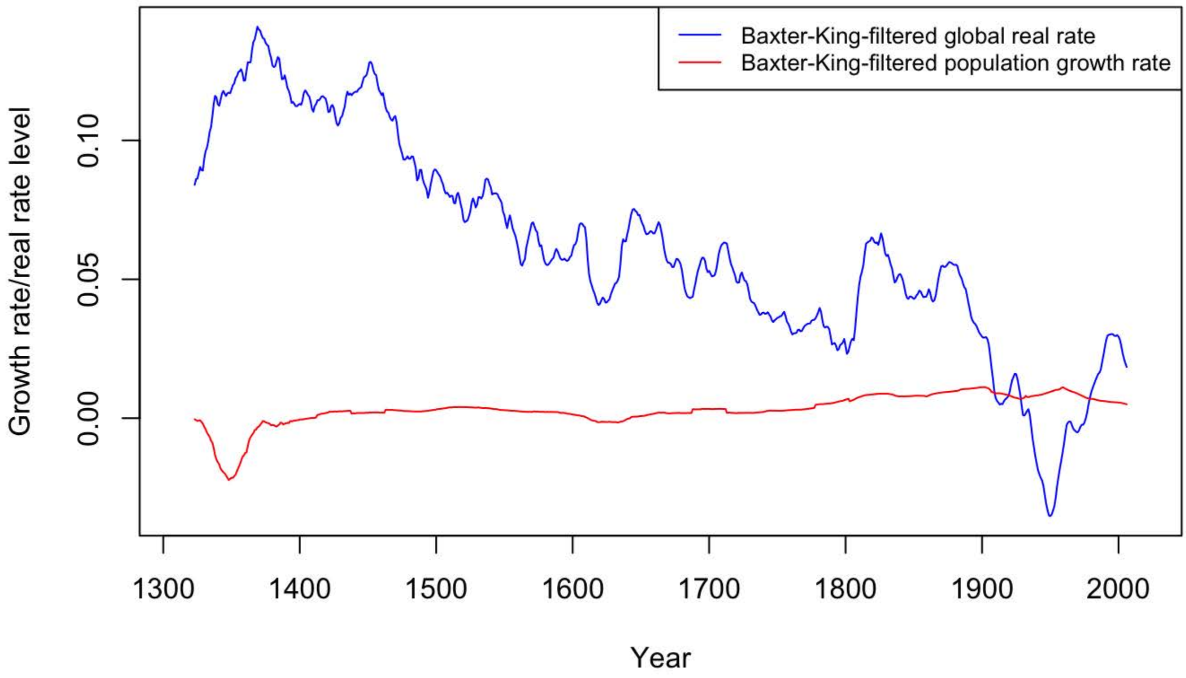
Notes: The aggregate output growth rate covers identical eight country sample, with country constituent series weighted according to GDP-weights of the sum of country outputs. Sources include Malanima (2011) for Northern Italy; Pfister (2021) for Germany; de la Escosura et al. (2022) for Spain; Ridolfi and Nuvolari (2021) for France; van Zanden and van Leeuwen (2012) for Holland; Broadberry and Fouquet (2015) for the U.K. Blue line displays the (filtered) long-run trend.

extract fluctuations of 100 years and more using Baxter and King's bandpass filter. The series covers the sample 1311-2021 for the identical weighted country-sample that we use for the global real rate series. As is visible from the chart, and can be confirmed by the more sophisticated statistical exercise below, aggregate population growth rates and real interest rates have structurally trended in *opposite* directions, except for the most recent decades. During major population growth shocks, the two series are positively correlated, such as during sharp plunge in aggregate population growth is observed during the Black Death of 1348-1351, the most dramatic demographic "rare disaster" striking advanced economies since the Renaissance - a period coinciding with a sharp plunge in global real rates.[38]

The finding that a secular upwards trend in population growth coincides with a downwards trend in real rates is not necessarily inconsistent with some recent literature - even though most contributions assume a positive correlation between the two variables (Rachel and Smith, 2016). But Carvalho et al. (2016) stress that population growth impacts (short-maturity) real rates through

[38]Of course, we are studying here global safe borrowing rates - this does not necessarily imply that the rate of return on, say, land moved in the same way.

Figure 5: A comparison of Baxter-King filtered aggregate population growth rates and real interest rates, 1311-2021.



Notes: The aggregate population growth rate covers identical eight country sample, with country constituent series weighted according to GDP-weights of the sum of country outputs. Sources include Malanima (2011) for Northern Italy; Pfister (2021) for Germany; de la Escosura et al. (2022) for Spain ("compromise estimate"); Dupaquier (1988) and Ridolfi (2019) for France; van Zanden and van Leeuwen (2012) for Holland; Broadberry and Fouquet (2015) for the U.K. Blue line displays the (filtered) long-run trend.

two divergent channels: a "supply effect" implying a rising (falling) population growth rate lowers (raises) capital per-worker, thus putting upwards (downwards) pressure on real rates; on the opposite, a (lagged) "demand effect" in which a rise (fall) in population growth should eventually lead to a lower (higher) dependency ratio, which generates downwards (upwards) pressure on real rates. Both channels are of course posited in the context of modern welfare states and the existence of substantial retirement periods: it is not obvious ex ante whether and to what extent prior to the introduction of modern welfare states and prolonged life expectancies, the balance between the two channels changes significantly.[39] However, note that even Carvalho et al. (2016) posit that the forces generating a positive correlation between population growth and real rates dominate: therefore, similar to the evidence from output dynamics, the long-run evidence from our new data

[39]Empirically, economy-wide life expectancy at birth for advanced economies accelerates sharply from the mid-19th century, but shows no decisive trend prior to this (thus also displaying an apparent disconnect with real interest rate dynamics). For a comprehensive recent empirical contribution to elite lifespan dynamics see Cummins (2017): though his sample over-weights English and Welsh data, his finding that European "elite" longevity increased sharply around two particular points, around 1400 and again around 1650, seems somewhat more consistent with our real rate evidence than data from either life expectancy at birth or population.

appears to run counter to recent causal demographic propositions of real rate dynamics.

Table 7 now reports the estimated correlation between the long-run filtered component of global real rates versus those of aggregate output and population growth. The table confirms a strong negative and statistically significant correlation.

Table 7: Global Real Rates Vs. Aggregate Real Output and Population Growth

	Correlation	95% CI	t-statistic	p-value
Baxter-King-filtered Long-run Component of Aggregate Population Growth	-0.6425	(-.803, -0.482)	-7.85	0.0000
Baxter-King-filtered Long-run Component of Aggregate Real Output Growth	-0.7399	(-0.818, -0.661)	-18.50	0.000

Note: The table reports the correlation of each series in the first column with Baxter and King's bandpass filtered component of the global real rate. The bandpass filter is tailored to retain fluctuations larger than 100 years. The t-statistic is obtained using Newey and West standard errors with a maximum of 4 lags.

In sum, the long-run evidence puts doubts on both productivity-centered and demographics-centered explanations of the post-1980s trend fall in real interest rates, as represented for instance by Gordon (2016), Holston et al. (2017), or Goodhart and Pradhan (2021). Several contributions have already voiced skepticism about productivity channels, including Hamilton et al. (2016) and Lunsford and West (2019) - we can confirm these from a much longer-term perspective. But importantly, we can qualitatively go beyond these existing facts, and find strong evidence in favor a long-run *negative* prediction of aggregate population growth rates and aggregate real output growth rates for real rates.[40] The very most recent decades - where for both variables the correlation with real rates is positive - in fact appear to be a unique exception to the structural rule.

Since demographics and productivity therefore cannot explain the 700-year trend decline in long-term real interests, is there a compelling alternative explanation? This important question is, unfortunately, beyond the scope of this paper: we can only offer a couple of suggestions for future research. Stefanski and Trew (2022), using the evidence in Schmelzing (2020) as motivation, offer a theoretical explanation based on trend rising social patience using a dynamic heterogenous agent model of fertility, in which more patient economic and social groups have more progeny. Changes in war are another possibility. Post-war Germany and Japan both (partially) defaulted outright on their pre-Second World War debt, Germany inflated away its First World War debt. Nevertheless, over time, the odds that any of the eight countries in our data set will outright default on any portion of their debt may have faded over time. Relatedly, although fluctuations are massive, it is almost surely the case that trend liquidity in major government bond markets has been positive over the centuries and continues to be.[41] Whatever the answer(s), we can only say that it needs to

[40]This assertion can be underpinned through a Granger causality set-up: results available from the authors.

[41]One would of course expect substantial increases in market liquidity over our sample length. But to what extent

be consistent with the evidence here on the fact that long-term real interest rates series appear to be trend stationary: and regardless of the underlying cause, the analysis does suggest that when real interest rates fall far below the (declining) long-term trend, there is a high chance of mean reversion unless of course there has been a rare structural break.

7 Conclusion

The sharp drop in the real interest rate in the 21st century, particularly in the years after the global financial crisis, has been arguably the most important macroeconomic development in modern times. To what extent should the fall be regarded as permanent (or even continuing), and to what extent can it be regarded as temporary, with expected eventual reversion to a longer-term trend. This paper has offered a novel approach to this problem by taking advantage of recent advances in quantitative financial history to try to get a better handle on the econometric properties of real interest rates than has previously been possible. Importantly, we focus on long-maturity sovereign debt (comparable to modern 10-year Treasury bonds) over eight centuries, from 1311-2021. This is in contrast to earlier research that has mainly been occupied with short rates, and uniformly with much shorter time periods. Importantly, prior to this paper, virtually all of the literature has been unable to reject the hypothesis of a unit root in real interest rates. The results here, based on much longer data sets, and on eight countries, are markedly different.

Overall, we do not find compelling evidence of a break in the long-run trend in global real interest rates since the 16th century, including across wars and changes in monetary regimes, although 1914 and 1981 show up in multiple variations of individual series. Not only do we strongly reject unit roots in almost every case (if one allows for a trend and sometimes without allowing for a trend), we also find that little robust evidence of any recent structural break, and indeed most of very long-run series yield only a couple significant breaks with any consistency. We find no evidence for several prominent "global" dates where existing literature would lead one to expect such breaks with high confidence: multiple advanced economy series have in fact not seen robust structural breaks since the 16th century, and the same is true when isolating the "modern" safe long-dated assets, including U.S. Treasury bonds (from 1787) and U.K. Treasury bonds. Importantly, this does not seem to be a manifestation of how we apply our (very standard) methodology, as we are able to replicate the non-stationarity results for short rates found in earlier studies over shorter time periods. This, in turn, suggests that it is a premium for short-term debt, or a liquidity premium

this development is driven by financial broadening rather than deepening is not established. Interestingly, in a number of polities at at the relative forefront of innovation and market depth, market turnover in the early modern period is highly elevated: as expounded above, Mueller (1997) reports a turnover figure for Venetian *Monte Vecchio* sovereign debt in the year 1434 equivalent to over 28% of Venetian nominal GDP for the year. Hammel (1988), who assembled one of the most nuanced housing market datasets for multiple centuries (1284-1700) in a leading Imperial city found turnover rates of 3.4% - 6.8% of the housing stock in a given year, overwhelmingly debt-financed (the U.S. turnover equivalent in 2003-2018 stood at 5.4% for single family homes). At least for the U.K., evidence also suggests that liquidity factors may have been a more relevant factor for equity markets, rather than for bond markets (Campbell et al., 2018). As of now, we lack high-frequency bid-ask spread data which would be one variable able to provide particularly nuanced insights.

for short term debt, rather than the real interest rate itself, that is non-stationary, at least this is an issue to be explored. The notion that our results are so different from the literature has to do with some anomalous feature of long-term sovereign bonds (not short-term debt) seems unlikely, given that in the appendix, we show that we can equally reject the unit root hypothesis for long-maturity mortgage and corporate real rates, even though these series are much shorter than our headline sovereign rate data.

Regarding historical triggers of such structural breaks - factors prompting generational "regime changes" - many of the key inflection points are associated with deep combined politico-financial crises, such as 1557 on the global level (against the backdrop of the "Trinity default" coupled with political turmoil in major economies), the Spanish political-cum-fiscal breakdown of 1814, associated with double state default and the restoration of the monarchy. More broadly, however, financial stress - while repeatedly a prominent component - does not appear as a *sufficient* cause to prompt such historical breaks. On the other hand, severe geopolitical stress by itself appears sufficient, at least in several sub-periods and geographies, to trigger structural breaks at times - as, for instance, during the 15th century. What appears generally surprising in the context of prior literature is that monetary factors alone constitute much less of an obvious source of radical change in very long samples: our approach raised doubts on multiple obvious monetary "inflections" including those previously associated with the establishment of central banks (such as 1694 or 1914), or major policy shifts - notably associated with 1981.

We are aware of the debate in the econometrics literature on the difficulties in distinguishing trend breaks from unit roots. However, the glaring downward trend in our 700-year real interest rates series is so consistent and clear it is hard to imagine not allowing for a trend, and in fact, and even for the series where we detect a trend break, the stationarity results still hold without allowing for one. Of course, further research is needed, and a Bayesian approach can obviously change the break results significantly, though we are doubtful it would radically change our novel finding that long-term real interest rates are stationary. We are also keenly aware that extrapolating our negative trend line another two centuries even at the glacial pace of 1.6 basis points per annum, would imply that the global trend real interest rate proceeds linearly, and ultimately reaches levels well below zero. A more likely possibility, perhaps, is that it will asymptote to a lower bound (positive or negative), though we do not see evidence of that yet and leave this issue to further study.

Finally, the recent literature has given a great deal of attention to demographics and productivity as drivers of today's real interest rates. Our research in no way rules this out, since it is possible there has been a modern structural break that has not yet lasted long enough to be statistically significant in the 700-year time series. Nevertheless, again making use of substantial recent advances in quantitative economic history, we can employ much-improved long-term demographic and real output data, which over most of our sample have been rising, not falling, yet the sovereign long-term real interest rate has trended down consistently. Hence, although we absolutely cannot conclude that the post-2008 epoch of seemingly very low real rates will eventually unwind back to a (gently declining) trend, we would argue that our data gives a valuable new perspective that may have implications for understanding a broad range of modern macroeconomic issues.

References

Abbas, Ali S., Alex Pienkowski, and Kenneth S. Rogoff. 2019. *Sovereign Debt: A Guide for Economists and Practitioners*. Oxford University Press.

Aerts, Erik. 2006. The European monetary famine of the late Middle Ages and the Bank of San Giorgio in Genoa. In *La Casa di San Giorgio: il potere del credito*, edited by Felloni, Giuseppe, 27–62.

Allen, Robert. 2001. The Great Divergence in European Wages and Prices from the Middle Ages to the First World War. *Explorations in Economic History* 38(4).

Alvarez-Nogal, Carlos. 2014. Monetary and Financial Innovation in the Spanish Empire: Lights and Shadows. In *Explaining Monetary and Financial Innovation*, edited by Bernholz, Peter, and Roland Vaubel, 301–321. Springer.

Alvarez-Nogal, Carlos, and Prados de la Escosura. 2013. The Rise and Fall of Spain (1270-1850). *Economic History Review* 66(1): 1–37.

Ang, Andrew, and Geert Bekaert. 2002. Regime Switches in Interest Rates. *Journal of Business and Economic Statistics* 20(2): 163–182.

Bai, Jushan, and Pierre Perron. 1998. Estimating and Testing Linear Models with Multiple Structural Breaks. *Econometrica* 66(1): 47–78.

Barsky, Robert, N. Gregory Mankiw, Jeffrey Miron, and David Weill. 1988. The Worldwide Change in the Behavior of Interest Rates and Prices in 1914. *European Economic Review* 32: 1123–1154.

Bateman, Victoria N. 2011. The evolution of markets in early modern Europe, 1350–1800: a study of wheat prices. *Journal of Economic History* 64(2): 447–471.

Bauer, Michael D., and Glenn D. Rudebusch. 2020. Interest Rates under Falling Stars. *American Economic Review* 110(5): 1316–1354.

Bernanke, Ben. 2004. The Great Moderation. *Speech at the Eastern Economics Association* .

Bernanke, Ben. 2005. The Global Savings Glut and the US current account deficit. *Sandridge Lecture, Virginia Association of Economists* .

Bernstein, Asaf, Eric Hughson, and Marc Weidenmier. 2010. Identifying the effects of a lender of last resort on financial markets: lessons from the founding of the Fed. *Journal of Financial Economics* 98(1): 40–53.

Blanchard, Olivier. 2022. *Fiscal Policy under Low Interest Rates*. MIT Press (forthcoming).

BoE. 1963. The Treasury Bill. In *Bank of England Quarterly Bulletin, Q3-1963*, 186–193.

Bolton, J.L., and Francesco Guidi Bruscoli. 2008. When Did Antwerp Replace Bruges as the Commercial and Financial Centre of NorthWestern Europe? The Evidence of the Borromei Ledger for 1438. *Economic History Review* 61(2): 360–379.

Broadberry, Stephen, and Roger Fouquet. 2015. Seven centuries of European economic growth and decline. *Journal of Economic Perspectives* 29(4): 227–244.

Caferro, William. 1998. *Mercenary companies and the decline of Siena*. Johns Hopkins University Press.

Campbell, Bruce M.S. 2005. The Agrarian Problem in the Early Fourteenth Century. *Past and Present* 188: 3–70.

Campbell, Gareth, John D. Turner, and Qing Ye. 2018. The liquidity of the London capital markets, 1825-1870. *Journal of Economic History* 71(3): 823–853.

Campbell, John Y. 2018. *Financial Decisions and Markets. A Course in Asset Pricing*. Princeton University Press.

Carvalho, Carlos, Andrea Ferrero, and Fernanda Nechio. 2016. Demographics and real interest rates: Inspecting the mechanism. *European Economic Review* 88: 208–226.

Chitu, Livia, Barry Eichengreen, and Arnauld Mehl. 2014. When did the dollar overtake sterling as the leading international currency? Evidence from the bond markets. *Journal of Development Economics* 111(4).

Chow, Gregory C. 1960. Tests of Equality Between Sets of Coefficients in Two Linear Regressions. *Econometrica* 28(3): 591–605.

Clark, Gregory. 2016. Microbes and Markets: Was the Black Death an Economic Revolution? *Journal of Demographic Economics* 82: 139–165.

Clodfelter, Micheal. 2017. *Warfare and Armed Conflicts: A Statistical Encyclopedia of Casualty and Other Figures, 1492-2015 (4th ed.)*. McFarland.

Comin, Francisco. 2012. Default, rescheduling and inflation: public debt crises in Spain during the 19th and 20th centuries. *Revista de Historia Economica* 30(3): 353–390.

Cummins, Neil. 2017. Lifespans of the European Elite, 800-1800. *Journal of Economic History* 77(2): 406–439.

Day, John. 1978. The Great Bullion Famine of the Fifteenth Century. *Past and Present* 79(2): 3–54.

de la Escosura, Prados, Carlos Alvarez-Nogal, and Carlos Santiago-Caballero. 2022. Growth recurring in preindustrial Spain? *Cliometrica* 16(2): 215–241.

de Moor, Tine, and Jaco Zuijderduijn. 2013. Spending, saving, or investing? Risk management in sixteenth-century Dutch households. *Economic History Review* 66(1): 38–56.

Delumeau, Jean. 1959. *Vie économique et sociale de Rome dans la seconde moitié du XVIe siècle. Tome second*. Bibliothèque des Ecoles françaises d'Athènes et de Rome.

Dimsdale, Nicholas, and Ryland Thomas. 2016. A Millennium of Macroeconomic Data, Version 3.1. *Bank of England Data Repository* .

Dupaquier, Jacques. 1988. *Histoire de la population française, 4 vols.* Presses Universitaires de France.

Ehrenberg, Richard. 1896. *Das Zeitalter der Fugger: Geldkapital und Creditverkehr im 16. Jahrhundert (2 Vols.)*. Gustav Fischer, Jena.

Ehrenberg, Richard. 1928. *Capital and Finance in the Age of the Renaissance. A Study of the Fuggers and their Connections.* Harcourt, Brace, and Row.

Eichengreen, Barry. 2015. Secular stagnation, the long view. *American Economic Review: Papers and Proceedings* 105(5): 66–70.

Eichengreen, Barry, Asmaa El-Ganainy, Rui Esteves, and Kris Mitchener. 2021. *In Defense of Public Debt*. Oxford University Press.

Eichengreen, Barry J. 2019. *Globalizing Capital*. Oxford University Press.

Elliott, Graham, Thomas J. Rothenberg, and James H. Stock. 1996. Efficient Tests for an Autoregressive Unit Root . *Econometrica* 64(4): 813–836.

Epstein, Stephen. 2000. *Freedom and Growth: The Rise of states and markets in Europe, 1300-1750*. Routledge.

Fama, Eugene. 1975. Short-Term Interest Rates as Predictors of Inflation. *American Economic Review* 65(3): 269–282.

Federico, Giovanni, Max-Stephan Schulze, and Oliver Volckart. 2021. European Goods Market Integration in the Very Long Run: From the Black Death to the First World War. *Journal of Economic History* 81(1): 276–308.

Flandreau, Marc, and Frederic Zumer. 2000. *The Making of Global Finance, 1880-1913*. OECD Publishing.

Frankel, Jeffrey A. 1986. International Capital Mobility and Crowding-Out in the U.S. Economy: Imperfect Integration of Financial Markets or of Goods Markets. In *How open is the U.S. economy?*, edited by Hafer, Rik W., 33–67.

Fratianni, Michele, and Franco Spinelli. 1997. *A Monetary History of Italy*. Cambridge University Press.

Friedman, Milton, and Anna J. Schwartz. 1963. *A Monetary History of the United States, 1867-1960*. Princeton University Press.

Fritschy, Wantje. 2017. *Public Finance of the Dutch Republic in Comparative Perspective: The Viability of an Early Modern Federal State (1570s-1795)*. Brill.

Froot, Kenneth A., Michael Kim, and Kenneth S. Rogoff. 1995. The Law of One Price over 700 years. *NBER Working Paper* .

Gagnon, Etienne, David Johannsen, and David Lopez-Salido. 2021. Understanding the New Normal: The Role of Demographics. *IMF Economic Review* 69(2): 357–390.

Garbade, Kenneth D. 2008. Why the U.S. Treasury Began Auctioning Treasury Bills in 1929. *Federal Reserve Bank of New York Review* 14(1): 1–17.

Garcia, Rene, and Pierre Perron. 1996. An Analysis of the Real Interest Rate Under Regime Shifts. *Review of Economics and Statistics* 78(1): 111–125.

Gelabert, Juan. 1999. The Italian States in the Early Modern Period. In *The Rise of the Fiscal State in Europe, c.1200-1815*, edited by Bonney, Richard, 417–442. Oxford University Press.

Gelderblom, Oscar, and Joost Jonker. 2011. Public Finance and Economic Growth: The Case of Holland in the Seventeenth Century. *Journal of Economic History* 71(1): 1–39.

Goodhart, Charles, and Manoj Pradhan. 2021. *The Great Demographic Reversal*. Springer.

Gordon, Robert. 2016. *The Rise and Fall of American Growth*. Princeton University Press.

Hamilton, James D., Jan Hatzius, Ethan Harris, and Kenneth D. West. 2016. The Equilibrium Real Rate: Past, Present, and Future. *IMF Economic Review* 64(4): 66–707.

Hammel, Rolf. 1988. Häusermarkt und wirtschaftliche Wechsellagen in Lübeck von 1284 bis 1700. *Hansische Geschichtsblätter* 106: 41–107.

Harris, Max. 2021. *Monetary War and Peace. London, Washington, Paris, and the Tripartite Agreement of 1936*. Cambridge University Press.

Hauser, Henri. 1930. The European Financial Crisis of 1559. *Journal of Economic and Business History* (2): 241–255.

Hoffman, Philip T., Gilles Postel-Vinay, and Jean-Laurent Rosenthal. 2019. *Dark Matter Credit*. Princeton University Press.

Hollander, Jacob H. 1919. *War Borrowing. A Study of Treasury Certificates of Indebtedness of the United States*. Macmillan Company.

Holston, Kathryn, Thomas Laubach, and John Williams. 2017. Measuring the Natural Rate of Interest: international trends and determinants. *Journal of International Economics* 108(S1): 59–75.

Homer, Sidney, and Richard Sylla. 2005. *A History of Interest Rates*. University of Rutgers Press.

Huber, Alfons. 1893. Studien über die finanziellen Verhältnisse Oesterreichs unter Ferdinand I. *Mitteilungen des Instituts fuer Oesterreichische Geschichtsforschung* 181–247.

Huizinga, John, and Frederic Mishkin. 1986. Monetary Policy Regime Shifts and the Unusual Behavior of Real Interest Rates. *Carnegie-Rochester Conference on Public Policy, April 1985* .

Ilzetzki, Ethan, Carmen Reinhart, and Kenneth Rogoff. 2022. Rethinking Exchange Rate Regimes. *Handbook of Macroeconomics* .

Itskhoki, Oleg. 2021. The Story of the Real Exchange Rate. *Annual Review of Economics* 13: 423–455.

Jorda, Oscar, Moritz Schularick, and Alan M. Taylor. 2017. Macrofinancial History and the New Business Cycle Facts. In *NBER Macroeconomics Annual 2016*, edited by Eichenbaum, Martin, and Jonathan A. Parker. University of Chicago Press.

Karaman, K. Kinvanc, Sevket Pamuk, and Secil Yildirim-Karaman. 2020. Money and monetary stability in Europe, 1300–1914. *Journal of Monetary Economics* 115(11): 279–300.

Kendall, M.G. 1955. Rank Correlation Methods .

Kilian, Lutz. 1999. Exchange rates and monetary fundamentals: what do we learn from long-horizon regressions? *Journal of Applied Econometrics* 14(5): 491–510.

Kindleberger, Charles P. 1998. Economic and Financial Crises and Transformations in Sixteenth-Century Europe. *Princeton Essays in International Economics* (208).

Kuechler, Winfried. 1983. *Die Finanzen der Krone Aragon waehrend des 15. Jahrhunderts (Alfons V. und Johann II.)*. Aschendorffsche Verlagsbuchhandlung.

Lapeyre, Henri. 1955. *Une famille de marchands - les Ruiz.* Librairie Armand Colin.

Laubach, Thomas, and John Williams. 2003. Measuring the Natural Rate of Interest. *Review of Economics and Statistics* 85(4): 1063–1070.

Laubach, Thomas, and John C. Williams. 2016. Measuring the Natural Rate of Interest Redux. *Business Economics* 51(2): 57–67.

Levy, Jack S. 1983. *War in the Modern Great Power System 1495-1975*. University Press of Kentucky.

Lothian, James R., and Mark P. Taylor. 1996. Real Exchange Rate Behavior: The Recent Float from the Perspective of the Past Two Centuries. *Journal of Political Economy* 105(3): 488–509.

Lunsford, Kurt G., and Kenneth West. 2019. Some Evidence on Secular Drivers of the Real Rate. *American Economic Review: Macroeconomics* 11(4): 113–139.

Maddison, Angus. 2010. *The World Economy - A Millennial Perspective*. OECD.

Malanima, Paolo. 2011. The long decline of a leading economy: GDP in central and northern Italy, 1300-1913. *European Review of Economic History* 15(2): 169–219.

Mankiw, N. Gregory, and Jeffrey Miron. 1986. The Changing Behavior of the Term Structure of Interest Rates. *Quarterly Journal of Economics* 101(2): 211–228.

Mann, H.B. 1945. Nonparametric tests against trend. *Econometrica* 13: 245–259.

Mian, Atif, Amir Sufi, and Ludwig Straub. 2021. What Explains the Decline in r*? Rising Income Inequality Versus Demographic Shifts. *Chicago BFI Working Paper* 1–48.

Mouré, Kenneth. 2002. *The Gold Standard Illusion: France, the Bank of France, and the International Gold Standard, 1914-1939*. Oxford University Press.

Mueller, Reinhold. 1997. *The Venetian money market: banks, panics, and the public debt, 1200-1500*. Johns Hopkins University Press.

Mulho, Anthony. 1971. *Florentine public finances in the Early Renaissance, 1400-1433*. Harvard University Press.

Munro, John H. 2003. The Monetary Origins of the 'Price Revolution' : South German Silver Mining, Merchant-Banking, and Venetian Commerce, 1470-1540' 1–60.

Munro, John H. 2010. The Coinages and Monetary Policies of Henry VIII (r. 1509-1547): Contrasts between Defensive and Aggressive Debasements. *University of Toronto Working Paper* (417): 1–79.

Münkler, Herfried. 2017. *Der Dreißigjährige Krieg: Europäische Katastrophe, deutsches Trauma 1618–1648*. Rowohlt Verlag.

Neal, Larry. 2015. *A Concise History of International Finance*. Cambridge University Press.

Neely, Christopher J., and David Rapach. 2008. Real Interest Rate Persistence: Evidence and Implications. *Federal Reserve Bank of St. Louis Review* .

Negro, Marco Del, Domenico Giannone, Marc Giannoni, and Andrea Tambalotti. 2019. Global trends in interest rates. *Journal of International Economics* 118: 248–262.

Nelson, Charles R., and Charles I. Plosser. 1982. Trends and Random Walks in Macroeconomic Time Series. *Journal of Monetary Economics* 10: 139–162.

Ng, Serena, and Pierre Perron. 2001. LAG Length Selection and the Construction of Unit Root Tests with Good Size and Power. *Journal of Econometrics* 69(6): 1519–1554.

North, Douglass C., and Barry R. Weingast. 1989. Constitutions and commitment: the evolution of institutions governing public choice in seventeenth-century England. *Journal of Economic History* 49(4): 803–832.

Obstfeld, Maurice, Jay Shambaugh, and Alan Taylor. 2005. The Trilemma in History: tradeoffs among exchange rates, monetary policies, and capital mobility. *Review of Statistics and Economics* 87(3): 423–438.

Pamuk, Sevket. 2007. The Black Death and the origins of the 'Great Divergence' across Europe, 1300–1600. *European Review of Economic History* 11(3): 289–317.

Parker, Geoffrey. 2019. *Emperor. A New Life of Charles V.* Yale University Press.

Perron, Pierre. 1989. The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis. *Econometrica* 57(6): 1361–1401.

Persson, Karl Gunnar. 1999. *Grain Markets in Europe, 1500-1900: Integration and Deregulation*. Cambridge University Press.

Pezzolo, Luciano. 2003. The Venetian Government Debt, 1350-1650. In *Urban public debts, urban government and the market for annuities in Western Europe (14th-18th centuries)*, edited by Boone, Marc, Karel Davids, and Paul Janssens, 61–74.

Pezzolo, Luciano. 2008. Government Debts and Credit Markets in Renaissance Italy. In *Government Debts and Financial Markets in Europe*, edited by Caselli, Fausto Piola, 17–32. Pickering and Chatto.

Pfister, Ulrich. 2021. Economic growth in Germany, 1500–1850. *University of Muenster Historical Seminar Working Paper* 1–42.

Pousada, Rafael Vallejo. 2015. Las vicisitudes de la deuda publica en un periodo turbulento: de canga arguelles a mon, 1808-1850. In *Historia de la deuda publica en Espana (siglos XVI-XXI)*, edited by Nogal, Carlos Alvarez, and Francisco Comin Comin, 159–186. Instituto de Estudios Fiscales.

Rachel, Lukasz, and Thomas D. Smith. 2016. Secular drivers of the global real interest rate. *Bank of England Staff Working Paper* (571).

Rachel, Lukasz, and Lawrence S. Summers. 2019. On Falling Neutral Real Rates, Fiscal Policy, and the Risk of Secular Stagnation. *Brookings Papers on Economic Activity* (1): 1–54.

Rapach, David E., and Christian E. Weber. 2004. Are real interest rates really nonstationary? New evidence from tests with good size and power. *Journal of Macroeconomics* 26(3): 409–430.

Reinhart, Carmen, and Kenneth Rogoff. 2009. *This Time is Different: Eight Centuries of Financial Folly*. Princeton University Press.

Reinhart, Carmen M., Vincent Reinhart, and Kenneth S. Rogoff. 2012. Public Debt Overhangs: Advanced-Economy Episodes since 1800. *Journal of Economic Perspectives* 26(3): 69–86.

Richards, R.D. 1936. The Exchequer Bill in the History of English Government Finance. *Economic Journal* 3(11): 193–211.

Ridolfi, Leonardo. 2019. Six Centuries of Real Wages in France from Louis IX to Napoleon III: 1250–1860. *Journal of Economic History* 79(3): 589–627.

Ridolfi, Leonardo, and Alessandro Nuvolari. 2021. L'histoire immobile? A reappraisal of French economic growth using the demand-side approach, 1280–1850. *Explorations in Economic History* 25(3): 402–428.

Rogoff, Kenneth S. 1996. The Purchasing Power Parity Puzzle. *Journal of Economic Literature* 34(2): 647–668.

Rose, Andrew K. 1988. Is the Real Interest Rate Stable? *Journal of Finance* 43(5): 1095–1112.

Schirmer, Uwe. 2006. *Kursachsische Staatsfinanzen (1456-1656) : Strukturen, Verfassung, Funktionseliten*. Franz Steiner.

Schmelzing, Paul. 2022. Eight centuries of global real rates, R-G, and the 'suprasecular' decline, 1311-2018. *mimeo* .

Scott, H.M. 2002. Britain's Emergence as a European Power, 1688-1815. In *A Companion to Eighteenth Century Britain*, edited by Dickinson, H.T., 431–446. Routledge.

Shiller, Robert J. 2015. *Irrational Exuberance. Revised and expanded Third edition*. Princeton University Press.

Shiller, Robert J., and Jeremy J. Siegel. 1977. The Gibson Paradox and Historical Movements in Real Interest Rates. *Journal of Political Economy* 85(5): 891–908.

Shizume, Masato. 2017. A History of the Bank of Japan. *WINPEC Working Paper Series* (E1719): 1–30.

Siegel, Jeremy. 1992. The Real Rate of Interest from 1800 to 1990. *Journal of Monetary Economics* 29: 227–252.

Simal, Juan Luis. 2017. «Strange Means of Governing»: The Spanish Restoration in European Perspective (1813–1820). *Journal of Modern European History* 15(2): 197–220.

Stasavage, David. 2011. *States of Credit: Size, Power, and the Development of European Polities*. Princeton University Pres.

Stefanski, Radek, and Alex Trew. 2022. Selection, Patience, and the Interest Rate. *University of Glasgow Working Paper* .

Sussman, Nathan, and Yishay Yafeh. 2006. Institutional reforms, financial development and sovereign debt: Britain 1690–1790. *Journal of Economic History* 66(4): 906–935.

van der Wee, Herman. 1963. *The Growth of the Antwerp market and the European Economy*. Louvain Publications Universitaires.

van Zanden, Jan Luiten, and Emanuele Felice. 2022. Benchmarking the Middle Ages: fifteenth century Tuscany in European perspective. *European Review of Economic History* .

van Zanden, Jan Luiten, and Bas van Leeuwen. 2012. Persistent but not Consistent: the growth of national income in Holland, 1347-1807. *Explorations in Economic History* 49(2): 119–130.

Vaughan, Richard. 1970. *Philip the Good. The Apogee of Burgundy*. Longmans.

Velde, Francois. 2008. French Public Finance between 1683 and 1726. In *Government debts and financial markets in Europe*, edited by Caselli, Fausto Piola, 135–166. Pickering and Chatto.

Volckart, Oliver, and Nicolas Wolf. 2006. Estimating Financial Integration in the Middle Ages: What Can We Learn from a TAR Model? *Journal of Economic History* 66(1): 122–139.

Wallis, Joseph John. 2005. Constitutions, Corporations, and Corruption: American States and Constitutional Change, 1842 to 1852. *Journal of Economic History* 65(1): 211–256.

Wicksell, Knut. 1936. *Interest and Prices (trans. of 1898 ed.)*.

Zivot, Eric, and Donald W.K. Andrews. 1992. Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis. *Journal of Business Economic Statistics* 10(3): 251–270.

APPENDIX

1. Robustness

We present robustness checks and a variety of additional results following from our benchmark results in the main body of the text. We begin by presenting the results from a Bai-Perron structural break test, following Bai and Perron (1998), reported in table A.1.

Table A.2 then follows by testing the data in using the ADF-GLS test in a specification *without* a time-trend.

Tables A.3 and A.4 proceed with variations on the deflator used to calculate real interest rates: recall that our exercises in the main text follow the methodology in Homer and Sylla (2005), and use a progressively lagged seven-year realized inflation basis, not including the current year, in order to infer inflation expectations. Table A.3 alternatively follows the approach in Hamilton et al. (2016), and estimates inflation expectations via an autoregressive model: per table A.3 we can confirm all our ADF-GLS results in this approach, too. Second, we follow Eichengreen (2015), who uses an equal-weighted seven-year realized inflation basis, including the current year, to deflate U.S. nominal yields in order to obtain real rates. For all except one series (Global GW), we can equally confirm the ADF-GLS results from the main body.

Afterwards, we show in particular that we can fully replicate the failure to reject unit roots in real rates when we apply our two alternative inflation expectations approaches for Rose (1988)'s shorter sample data (tables A.6.1-A.6.3). We confirm this failure to reject a unit root uniformly at the 1% level for all 16 possible real rate variations, including for his long-maturity corporate bond rate basis, as well as for his short-maturity commercial paper rate basis. This underscores that our headline results do not appear to be a function of how we specifically combine the individual nominal rate and inflation series.

.1 Bai Perron break dates, and ADF-GLS without time trend

To begin, Table A.1 displays Bai-Perron structural breaks in the global real rate series, following Bai and Perron (1998). The Bai-Perron test - here the test assumes a deterministic trend of the underlying series, given our ADF-GLS results - does not "impose" structural breaks: none, one, or multiple structural breaks are possible, up to a maximum of five breaks. The results yield an optimal partition of six periods for global real rates (GW and AW), and equally six periods for all country series, with most showing a break in the first half of the 15th century, the mid-16th century, the second half of the 17th century, the early 19th century, and during the First World War. The U.K., France and Spain display a break during the early or late 18th century, respectively, and only the U.S. and Japan display breaks after the First World War.

Like the breaks obtained via the Chow and Zivot-Andrews procedures in the main body, the break dates thus obtained retain a high degree of plausibility. For instance, the French and U.K. break occur during crucial phases of the Hundred Years' War; the U.K. does not record a break during the Napoleonic Wars, during which it is the only country not to experience war on its own soil; the U.S. does not record a break during World War One, when it is the last entrant and only peripherally involved.

Table A.1: Bai-Perron Test Results

Series	N. breaks	Estimated break dates
Global Real GW	5	1416 1524 1630 1806 1915
Global Real AW	5	1428 1552 1670 1806 1915
Italy Real	5	1443 1557 1696 1802 1915
UK Real	5	1416 1552 1658 1775 1892
Dutch Real	5	1465 1564 1663 1802 1921
France Real	5	1423 1565 1672 1793 1915
Germany Real	5	1417 1524 1630 1809 1915
Spain Real	5	1436 1594 1711 1814 1917
US Real	5	1825 1869 1907 1941 1983
Japan Real	5	1891 1926 1952 1974 1996

Note: the table reports the results of the sequential Bai and Perron's test (Bai and Perron, 1998) for an unbalanced panel that uses the period of 1311-2021. The maximum number of break points is 5 for all series.

Table A.2: ADF-GLS Test without Time Trend			
Real Rate Series			
Region	Number of lags	ADF-GLS test	Optimal lag
Global GW	3	-5.064	MAIC
	2	-5.411	Seq, SIC
	1	-7.735	
Global AW	3	-3.924	MAIC
	2	-4.134	Seq, SIC
	1	-5.848	
Italy	3	-9.480	MAIC
	2	-10.443	Seq, SIC
	1	-13.920	
UK	3	-3.006	
	2	-3.102	Seq, SIC, MAIC
	1	-4.281	
Dutch	3	-6.971	Seq, SIC, MAIC
	2	-8.385	
	1	-10.977	
France	3	-5.534	MAIC
	2	-5.963	Seq, SIC
	1	-7.721	
Germany	3	-8.129	
	2	-8.078	Seq, SIC, MAIC
	1	-11.444	
Spain	3	-3.090	
	2	-3.155	Seq, SIC, MAIC
	1	-4.477	
U.S.	3	-4.910	
	2	-5.507	
	1	-5.730	Seq, SIC, MAIC
Japan	3	-3.256	
	2	-3.241	Seq, SIC, MAIC
	1	-4.389	

Note: the table reports the test statistic for several number of lags (for a maximum of three lags). The regression includes a constant. For all series except for U.S. and Japan, the critical values at 1, 5, 10 percent significance levels are the following for all observations: -2.58 (1%); -1.95 (5%); -1.62 (10%). For the U.S., the critical values are -2.582 ; -1.95 ; -1.619. For Japan, they are -2.593 ; -1.95 ; -1.613. "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("seq"), the 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.

.2 Variations of inflation expectations

Hamilton et al. (2016) use an autoregressive inflation expectations approach. We replicate their approach for our multi-century data in table A.3 below. We calculate one-period-ahead inflation expectations following the same procedure as in Hamilton et al. That is, the evolution of inflation is recursively estimated in rolling windows using an autoregressive model with one lag; then, the estimated model is used to predict one-year-ahead inflation in an out-of-sample fashion. The rolling window estimation procedure guards against instabilities by using only past observations in the latest window of data. We use a window size equal to 30 years. As shown, we are able to reject a unit root with this inflation basis, too, for both the global (GW and AW), and all country series.

Second, Eichengreen (2015) uses a seven-year equal-lagged realized inflation approach to deflate nominal yields, and includes current-year inflation (year t). We replicate his approach in table A.4 below. As shown, we are equally able to reject a unit root with this inflation basis, for all series, including Global AW - with the only exception of the Global GW series.

Table A.3: ADF-GLS Test for Hamilton inflation basis			
Real rate Series			
Region	Number of lags	ADF-GLS test	Critical values: 1%
5%	10%	Optimal lag	
Global GW	3	-6.137	Seq, MAIC, SIC
	2	-7.041	
	1	-9.338	
Global AW	3	-6.549	
	2	-6.751	Seq, SIC, MAIC
	1	-9.199	
Italy	3	-9.158	MAIC
	2	-10.013	Seq, SIC
	1	-13.428	
UK	3	-11.775	MAIC
	2	-13.740	
	1	-17.098	SIC, Seq (0)
Dutch	3	-3.663	Seq, MAIC
	2	-3.912	SIC
	1	-4.496	
France	3	-4.791	MAIC
	2	-5.143	Seq, SIC
	1	-6.385	
Germany	3	-5.117	Seq, MAIC, SIC
	2	-6.033	
	1	-9.357	
Spain	3	-6.684	MAIC
	2	-7.264	Seq, SIC
	1	-10.144	
U.S.	3	-4.159	
	2	-5.899	
	1	-6.968	SIC, MAIC, Seq
Japan	3	-3.633	
	2	-3.681	Seq, MAIC
	1	-4.840	SIC

Note: the table reports the test statistic for several number of lags (for a maximum of three lags). The regression includes a constant and a time trend. For all series except for Japan, the critical values at 1, 5, 10 percent significance levels are the following for all observations: -3.48 (1%); -2.89 (5%); -2.57 (10%). For Japan, the critical values are -3.518 ; -2.978 ; -2.688. "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("seq"), the 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.

Table A.4: ADF-GLS Test for Eichengreen inflation basis			
Real rate Series			
Region	Number of lags	ADF-GLS test	Critical values: 1%
5%	10%	Optimal lag	
Global GW	3	-2.505	
	2	-2.480	Seq, MAIC, SIC
	1	-4.488	
Global AW	3	-3.365	Seq, MAIC, SIC
	2	-3.405	
	1	-6.120	
Italy	3	-3.163	
	2	-3.109	Seq, MAIC, SIC
	1	-5.749	
UK	3	-10.069	Seq, SIC
	2	-9.100	MAIC
	1	-15.852	
Dutch	3	-9.617	
	2	-10.232	Seq, SIC, MAIC
	1	-18.259	
France	3	-5.585	
	2	-5.617	Seq, MAIC, SIC
	1	-10.220	
Germany	3	-11.135	Seq, SIC
	2	-9.651	MAIC
	1	-18.133	
Spain	3	-6.946	Seq
	2	-6.603	SIC, MAIC
	1	-10.719	
U.S.	3	-6.964	Seq
	2	-6.719	SIC, MAIC
	1	-10.070	
Japan	3	-4.153	
	2	-3.983	Seq, SIC, MAIC
	1	-6.973	

Note: the table reports the test statistic for several number of lags (for a maximum of three lags). The regression includes a constant and a time trend. For all series except for Japan, the critical values at 1, 5, 10 percent significance levels are the following for all observations: -3.48 (1%); -2.89 (5%); -2.57 (10%). For Japan, the critical values are -3.518 ; -2.978 ; -2.688. "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("seq"), the 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.

.3 Context of Zivot-Andrews test dates and inflation variation for Chow test

We here discuss the historical context of the Zivot-Andrews break dates reported in table 2 in the main body: while the break dates indeed have historical plausibility, we reiterate that we have a strong conviction in our choice of the five Chow test dates based on our read of the long-run historical context and literature.

- Three of the Zivot-Andrews dates, including the headline global real series, are obtained for years immediately before the outbreak of the Second World War:
  - First, the interwar period is a period of general political volatility. Economically, it is identified with the transition from the U.K. to the U.S. as the global "safe asset provider", as well as the chaotic experimentation with the gold-exchange standard (Chitu et al., 2014), and major tail events such as the German hyperinflation and the Great Depression. The latter two events, however, are not specifically overlapping with real rate breaks.[42]
  - The global break appears to be driven by country-level breaks in or around 1937 in multiple countries, including France (1937), and Japan (1938), on the Zivot-Andrews basis. What does specific country-level context suggest for possible drivers of structural real rate breaks?
  - In French monetary history, 1937 equally represents a pivotal year, following the country's belated and rocky departure from the gold standard, the conclusion of the Tripartite Agreement, and the de facto nationalization of the Banque de France (Mouré, 2002; Harris, 2021).
  - In Japan, 1938 follows the outbreak of the Second Sino-Japanese War (Japan's de facto entering of the Second World War). Financial and monetary historians associate the years 1937-8 with the country's full-scale adoption of financial repression and debt monetization, including a loss of BoJ independence, and price and capital controls (Shizume, 2017).
- Germany's break date (1628) is set during the upheavals of the Thirty Years' War, and just after the unprecedented "Kipper- and Wipper" debasement crisis which plunged entire Central and Western Europe into a currency chaos. Politically, the year 1626 records several major geopolitical engagements in the country's North, with favorable outcomes for the Catholic side under Albrecht Wallenstein. While the full fiscal disintegration of the Holy Roman Empire and its constituent polities is typically dated to the late phase of the war (1635-1648), one can plausibly posit a combined politico-financial collapse as the driver for this break (Münkler, 2017).
- The year 1916 stands out in Italian monetary and political history. After the belated entry into the First World War in 1915, Italian public finances began to deteriorate sharply, with deficits reaching 60% of national income by 1918; by late 1916, the volume of government notes in circulation had already almost tripled compared to the pre-war volume (Fratianni and Spinelli, 1997, 108ff.).[43]
- The British break date (1553) immediately follows the Great Debasement begun by Henry VIII, the largest known currency debasement in the country's history (Munro, 2010; Karaman et al., 2020).

[42] As noted above, the German, Italian, and Japanese hyperinflation years are excluded from the Schmelzing (2022) data, but war years generally are included.

[43] As detailed above, we exclude the Italian 'hyperinflation' years of 1943-5, associated with the double invasion of Italy from the Nazis in the North ("Fall Achse") and Allies in the South ("Operation Avalanche"): economically, the years are associated with a major plunge in industrial output and exceptional deficit financing; the Lire exchange rate with the dollar which stood at 19 in the year 1940, was fixed at 100 in 1943 (Fratianni and Spinelli, 1997, 166f.). If these years were included, the year 1943 would mark the Zivot-Andrews break point.

- The Spanish break (1814) coincides with the liberation of Spain from Napoleonic occupation and royal restoration, following a disastrous inflation surge, a double sovereign default (first under the Cortes de Cadiz in 1812-3, then under Fernando VI in 1814), and institutional and economic breakdown (Comin, 2012; Pousada, 2015; Simal, 2017).
- In his classic study on Burgundy, Herman van der Wee called the years 1437-8 "the final blow" to the region's medieval economy (van der Wee, 1963, 61). Our Dutch break (1437) coincides with a period of deep state crisis in the Low Countries: popular revolts break out in that year in the major merchant hubs of Ghent and Bruges, both the English Crown and the Emperor Sigismund are engaged in military invasions to remove the Burgundian ruler Philip the Good from power, and in 1438, a full-scale trade war with the Hanse breaks out, halting Dutch Eastern maritime trade. On top of it all, a severe famine and extreme cold during 1437-8 aggravated widespread death and subsequent epidemics, as well as contributing to a further surge in inflation rates (reaching more than 36% for the year 1437) Bolton and Bruscoli (2008). "The perspective that emerges is one of turbulence and disorder, of danger and crisis, in the history of the Burgundian state", in the words of Vaughan (1970).
- The U.S. break (1850) coincides with major institutional reform in public finance and corporate governance, as reconstructed by Wallis (2005). Wallis argues that the period of 1842 to 1852 - following recurrent state defaults by 1840 - significantly strengthened secure private property rights in the U.S. Particular macroeconomic shocks, including the financial crises of 1838-40, and later 1857, are somewhat removed from our structural break, as is the major geopolitical shock of 1861-5.

To us, this list underscores our read of the historical context discussed in the main body, namely that major political-cum-economic upheavals are plausible contenders to explain structural breaks in real interest rates over time: it appears that political shocks can be accompanied by monetary shocks, but that the latter do not typically cause a break on their own. In particular, we do not read the evidence as necessarily suggesting that major FX transitions are themselves a sufficient factor in real interest rate inflections: we note that individual countries which departed the gold standard well before 1914 (namely Spain) still appear to have shown no trend break at this earlier date (Obstfeld et al., 2005). Neither do FX transitions without accompanying political escalations result in positive Chow identifications when tested (e.g. 1971).

We now present the Chow test familiar from the main body, but applying an equal-lagged inflation basis - analogous to the Eichengreen (2015) approach (which includes current year inflation) discussed for table A.4. We observe that all the main results from the progressively-lagged approach are confirmed, in particular the relevance of the 1349 and 1914 break points.

Table A.5: Chow Test Results - Global Real Rate GW - equal-lagged inflation

Year	Coefficient	Std error	t-statistic	p-value	95%	Confidence interval
trendbreak 1349	.2992621	.3542793	0.84	0.399	-.3963187	.9948429
trendbreak 1557	-.0040828	.0131019	-0.31	0.755	-.0298066	.021641
trendbrak 1694	.0165211	.0162744	1.02	0.310	-.0154316	.0484739
trendbreak 1914	-.1537574	.0781571	-1.97	0.050	-.3072086	-.0003063
trendbreak 1981	.3392345	.3263694	1.04	0.299	-.301549	.9800179
meanbreak 1349	-.0023604	.0120145	-0.20	0.844	-.0259492	.0212284
meanbreak 1557	.0027016	.005392	0.50	0.616	-.0078848	.013288
meanbreak 1694	-.0099685	.0089481	-1.11	0.266	-.0275369	.0075999
meanbreak 1914	.1332954	.0687009	1.94	0.053	-.0015899	.2681806
meanbreak 1981	-.3078964	.3157292	-0.98	0.330	-.9277891	.3119964
trend	-.1948497	.2854704	-0.68	0.495	-.7553332	.3656338
constant	.1810158	.2802269	0.65	0.519	-.3691728	.7312044

Note: The Chow test is implemented using Newey–West standard errors with a maximum lag length of 4. The coefficient associated with "trendbreak 1981" denotes the difference of the estimated trend coefficients before and after a break in 1981 (and 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.

.4 Replicating Rose (1988) with two inflation expectation variations

Tables A.6.1-A.6.3 replicate Rose (1988)'s results for U.S. shorter sample unit root tests, including long-maturity corporate bond yields (CB) and short-maturity commercial paper rates (CP), using his two alternative observation periods, 1892-1970 and 1901-1950 , and using annual nominal (table A.6.1) and real rate (tables A.6.2 and A.6.3) data. We use the same data for both nominal rates and inflation, sourcing from Friedman and Schwartz (1963) and Nelson and Plosser (1982): the latter data is taken from Peter C.B. Phillips data page (http://korora.econ.yale.edu/phillips/data/npenp.dat). Table A.6.2 uses the progressively-lagged inflation expectation approach as in Homer and Sylla (2005); table A.6.3 uses the equal-lagged approach as in Eichengreen (2015)).

We observe that once we apply our two alternative inflation expectation approaches to his underlying data - which for the underlying price data is based on a GNP deflator series, and alternatively a CPI index, both of which we test - and construct equivalent real rate samples we are unable to reject a unit root for all progressively-lagged real rate variations (table A.6.2), as well as for equal-lagged variations (table A.6.3), with no exceptions. In particular, we emphasize that we replicate Rose's results when allowing for a deterministic trend in ADF-GLS.

Rates	N. of lags	ADF-GLS test statistic	Critical values 1%	5%	10%	Optimal lag
Table A.6.1: Rose (1988) replication						
Nominal rate series						
1892-1970						
Corporate bond yields	3	-0.793	-3.660	-3.097	-2.803	
	2	-0.523	-3.660	-3.097	-2.803	
	1	-0.246	-3.660	-3.097	-2.803	Seq, SIC, MAIC
Commercial paper rates	3	-1.429	-3.660	-3.097	-2.803	Seq, SIC, MAIC
	2	-0.785	-3.660	-3.097	-2.803	
	1	-1.763	-3.660	-3.097	-2.803	
1901-1950						
Corporate bond yields	3	-1.494	-3.770	-3.190	-2.890	
	2	-1.212	-3.770	-3.190	-2.890	
	1	-1.355	-3.770	-3.190	-2.890	Seq, SIC, MAIC
Corporate bond yields	3	-2.164	-3.770	-3.190	-2.890	
	2	-1.766	-3.770	-3.190	-2.890	MAIC
	1	-2.405	-3.770	-3.190	-2.890	Seq(0), SIC

Note: The table reports the test statistic for several number of lags (for a maximum of three lags). The regression includes a constant and a time trend. "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("seq"), the SIC, or the Modified Information Criterion (MAIC). Seq(0) denotes cases where the sequential lag length procedure selects zero lags. The test rejects when the test statistic is negative and larger (in absolute value) than the critical value.

Table A.6.2: Rose (1988) replication						
Real rate series - progressively lagged inflation						
Rates	N. of lags	ADF-GLS test statistic	Critical values 1%	5%	10%	Optimal lag
GNP Deflator						
1892-1970						
Corporate bond yields	3	-2.669	-3.660	-3.097	-2.803	
	2	-2.757	-3.660	-3.097	-2.803	
	1	-3.026	-3.660	-3.097	-2.803	Seq, SIC, MAIC
Commercial paper rates	3	-2.452	-3.660	-3.097	-2.803	Seq
	2	-2.048	-3.660	-3.097	-2.803	MAIC
	1	-2.570	-3.660	-3.097	-2.803	SIC
1901-1950						
Corporate bond yields	3	-2.257	-3.770	-3.190	-2.890	
	2	-2.340	-3.770	-3.190	-2.890	
	1	-2.614	-3.770	-3.190	-2.890	Seq, SIC, MAIC
Commercial paper rates	3	-2.261	-3.770	-3.190	-2.890	Seq, SIC, MAIC
	2	-2.084	-3.770	-3.190	-2.890	
	1	-2.410	-3.770	-3.190	-2.890	
CPI Index						
1892-1970						
Corporate bond yields	3	-2.751	-3.660	-3.097	-2.803	
	2	-2.655	-3.660	-3.097	-2.803	Seq, MAIC
	1	-3.533	-3.660	-3.097	-2.803	SIC
Commercial paper rates	3	-2.571	-3.660	-3.097	-2.803	Seq, SIC
	2	-2.054	-3.660	-3.097	-2.803	MAIC
	1	-2.873	-3.660	-3.097	-2.803	
1901-1950						
Corporate bond yields	3	-2.196	-3.770	-3.190	-2.890	
	2	-2.189	-3.770	-3.190	-2.890	MAIC
	1	-3.005	-3.770	-3.190	-2.890	Seq, SIC
Commercial paper rates	3	-2.246	-3.770	-3.190	-2.890	
	2	-2.029	-3.770	-3.190	-2.890	MAIC
	3	-2.677	-3.770	-3.190	-2.890	Seq(0), SIC

Note: The inflation approach follows Homer and Sylla (2005), using seven-year progressively lagged inflation (*t-6* to *t*, excluding the current-year inflation *t*). The GNP deflator is from Nelson and Plosser (1982); the CPI index is from Nelson and Plosser (1982). The table reports the test statistic for several number of lags (for a maximum of three lags). The regression includes a constant and a time trend. "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("Seq"), the SIC, or the Modified Information Criterion (MAIC). Seq(0) denotes cases where the sequential procedure selects zero lags. The test rejects when the test statistic is negative and larger (in absolute value) than the critical value.

Table A.6.3: Rose (1988) replication						
Real rate series - equal lagged inflation						
Rates	N. of lags	ADF-GLS test statistic	Critical values 1%	5%	10%	Optimal lag
GNP Deflator						
1892-1970						
Corporate bond yields	3	-2.303	-3.660	-3.097	-2.803	
	2	-2.220	-3.660	-3.097	-2.803	
	1	-2.078	-3.660	-3.097	-2.803	Seq, SIC, MAIC
Commercial paper rates	3	-2.080	-3.660	-3.097	-2.803	Seq
	2	-1.428	-3.660	-3.097	-2.803	
	1	-1.659	-3.660	-3.097	-2.803	SIC, MAIC
1901-1950						
Corporate bond yields	3	-2.214	-3.770	-3.190	-2.890	
	2	-2.130	-3.770	-3.190	-2.890	
	1	-1.945	-3.770	-3.190	-2.890	Seq, SIC, MAIC
Commercial paper rates	3	-2.135	-3.770	-3.190	-2.890	
	2	-1.886	-3.770	-3.190	-2.890	
	1	-1.969	-3.770	-3.190	-2.890	Seq, SIC, MAIC
CPI Index						
1892-1970						
Corporate bond yields	3	-2.657	-3.660	-3.097	-2.803	
	2	-2.438	-3.660	-3.097	-2.803	MAIC
	1	-2.793	-3.660	-3.097	-2.803	Seq, SIC
Commercial paper rates	3	-2.857	-3.660	-3.097	-2.803	Seq, SIC, MAIC
	2	-1.653	-3.660	-3.097	-2.803	
	1	-2.080	-3.660	-3.097	-2.803	
1901-1950						
Corporate bond yields	3	-2.196	-3.770	-3.190	-2.890	
	2	-2.137	-3.770	-3.190	-2.890	
	1	-2.339	-3.770	-3.190	-2.890	Seq, SIC, MAIC
Commercial paper rates	3	-2.486	-3.770	-3.190	-2.890	Seq
	2	-1.870	-3.770	-3.190	-2.890	
	3	-2.152	-3.770	-3.190	-2.890	SIC, MAIC

Note: The inflation approach follows Eichengreen (2015), using seven-year equal lagged inflation (t-7 to t-1, including the current-year inflation t. The GNP deflator is from Nelson and Plosser (1982); the CPI index is from Nelson and Plosser (1982). The table reports the test statistic for several number of lags (for a maximum of three lags). The regression includes a constant and a time trend. "Optimal lag" indicates the optimal number of lags according to the sequential procedure ("seq"), the 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.

.5 Beyond sovereign long-maturity rates

Table A.7 now tests the statistical properties of a range of alternative fixed income series over the period of 1695-2021, 1853/4-2016, and 1913-2021, respectively. Our intention is to assess whether our result of trend stationarity for global sovereign real rates, as discussed in the main text above, extends to other asset classes. At least for selected countries and asset classes, this question can be addressed.

We find that, indeed, trend stationarity is confirmed for a number of additional asset bases, using equivalent real ex post rates, including British corporate bond real rates (long-maturity), and U.K. real mortgage rates (long-maturity) at the 1% critical level. Trend stationarity is equally confirmed for Bank of England real policy rates (short-maturity). The shorter U.S. real interest rate sample as collected by Shiller (2015), covering long-maturity government bonds over 1913-2015, on the other hand, only fails to reject non-stationarity at the 10% level.

Overall, these results suggest that stationarity properties of real asset yields and returns are not restricted to sovereign assets, at least in the post-1695 period. This cast doubts on the assertion that interest rates should be modeled as non-stationary - that they, essentially, follow a random walk over the long run.

Table A.7: ADF-GLS - Other Rate Variations

Series	Test statistic (t-statistic for α_1)	Critical values 1%	5%	10%
1 - Shiller US Real Rate, 1879-2015	-2.6567	-3.46	-2.93	-2.64
2 - BOE Bank Real Series, 1694-2016	-10.431	-3.48	-2.89	-2.57
3 - UK Mortgage Real Series, 1853-2016	-4.2927	-3.46	-2.93	-2.64
4 - UK Corporate Bond Real Series, 1854-2016	-4.4664	-3.46	-2.93	-2.64

The estimated model is an autoregression with three lags, a constant and a deterministic trend. The test rejects a unit root when the t-statistic is negative and, in absolute value, larger than the critical value.

Notes: [1], Shiller real rate, uses series "U.S. long government bond rate" data in Shiller (2015), deflated by seven-year progressively lagged change in "consumer prince index" series, both over 1913-2015, and both available from the author's homepage (http://www.econ.yale.edu/ shiller/data.htm); [2-4] use British interest rate series in the BoE's Millennium dataset, via Dimsdale and Thomas (2016), with the following chronological (annual frequency) coverage: [2]: 1695-2016; [3]: 1853-2016; [4]: 1854-2016. Deflation is throughout undertaken via seven-year progressively lagged headline U.K. consumer price index change, via the same source. The corporate bond series combines industrial and railway debentures and is long-maturity. The BoE bank rate is defined as the minimum lending rate between 1695-1972, the minimum band 1 dealing rate for 1981-1997, the repo rate for 1997-2006, and the bank rate for 2006-2016.

.6 Monte Carlo simulations on the power of the ADF-GLS test for the global real rate

To investigate the use of long versus small sample sizes in testing for unit roots in our data, we performed a small Monte Carlo simulation exercise. We generated 5000 time series that have the same features as our (stationary) global real rate (such as the mean, the deterministic trend and the standard error): $y_t = 0.7437y_{t-1} + 0.31 - 0.0000436t + 0.02313\varepsilon_t$. For each generated series, we calculated: (a) the ADF-GLS test statistic for a sample of 711 observations (the same size as our full sample); (b) the ADF-GLS test statistic for the sub-sample including the first 100 observations; (c) the ADF-GLS test statistic for the sub-sample including the first 50 observations. We then report the number of times the ADF-GLS test rejects a unit root at the 5 percent significance level across Monte Carlo simulations.

The simulation results indicate that the ADF-GLS test statistic rejects a unit root 100 percent of the times in the full sample, whereas it does so only 77 percent of the time in the sample with 100 observations and 37 percent of the time in the sample with 50 observations. This Monte Carlo simulation confirms the lack of power for time series with a short span.

The average estimated half-life across Monte Carlo simulations is about 3 periods (as the random series are generated to share the same features as the observed data). Thus, our estimated half-life of 3 periods is not incompatible with the fact that the test would need a relatively large sample (such as our full sample of 711 observations) to have large power.