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Abstract

The real exchange rate is an important measure of the relative strength of an economy. Given long-term productivity differentials between countries, Harrod-Balassa-Samuelson effects suggest that stronger economies will experience real exchange rate appreciations and vice versa. How long can these effects last? Using a novel dataset and trend tests robust to pre-testing for the order of integration, we examine the path of ten ultra-long real exchange rates relative to Sterling with data commencing in the 13th century. Whilst we show Sterling commonly presents a trend appreciation from the 16th century to the 19th century, a striking trend depreciation occurred throughout 20th century with some evidence of prior decline. Further analysis reveals that real exchange rates are cointegrated with productivity differential proxies over much of the last millennia, suggesting the UK's current productivity decline is more entrenched and persistent than previously thought.

Keywords: Real exchange rate; Sterling; Harrod-Balassa-Samuelson effects; Robust trends; Cointegration.

1. Introduction

This paper examines the duration of trends in real exchange rates and why those trends may exist? Using ultra-long series, in some cases dating back to the 13th century, we seek to assess the characteristics of exchange rates and whether HBS (Harrod, 1933; Balassa, 1964; Samuelson, 1964) effects can be detected even prior to the industrial revolution during the late medieval period or early modern age. HBS effects occur given countries have persistent and differing productivity levels, leading to long-run secular movements in the real exchange rate and therefore, their identification can tell us about the changing relative 'strength' of economies over the last millennia. Following trend estimation, and to examine further evidence for HBS effects, we assess whether real exchange rates present an ultra-long association with productivity differentials

In terms of real exchange rates, much the extant literature has typically examined whether purchasing power parity (PPP) holds (see Itskhoki, 2021; Vo and Vo, 2023) and discussed why deviations may occur (see Berka *et al.*, 2018; Lambertini and Proebsting, 2023). Indeed, in an attempt to find empirical evidence for PPP, several papers in the extant literature have examined the path of real exchange rates over a very long period or in some historical context.[1] For example, recently, Papell and Prodan (2020) use the exchange rates 16 OECD countries (relative to the US dollar) over the period 1870 to 2013 and show that long-run purchasing power parity (PPP) – implying that the real exchange is mean reverting - cannot be rejected for nine countries including France and the UK.[2] Likewise, Taylor (2002) using an unbalanced sample of 20 US dollar exchange rates with the earliest series commencing in 1850, shows that long-run PPP is 'supported in all cases with allowance for deterministic trends' (Taylor, 2002, p.140). Drawing on even longer

[1] In the original conception of purchasing power parity (PPP), Cassel (1918) viewed deviations of the exchange rate from relative prices as limited.
[2] This is at the 5 percent level of significance. At 10 percent, 11 countries cannot reject the null of long-run PPP.

time series, Lothian and Taylor (1996, 2000) examine the Franc-Sterling and US dollar-Sterling rates from 1790 to 1990, asserting that long-run PPP is a 'good first approximation' of real exchange rate behaviour (see Lothian and Taylor, 2000, p.764).[3]

As noted by Taylor (2002) above, evidence for long-run PPP is further enhanced when we allow for reversion around deterministic trends. Indeed, in their reply to Cuddington and Liang (2000), Lothian and Taylor (2000) suggest that the linear time trend added to their model of the US dollar-Sterling series is statistically significant, indicates a trend depreciation of Sterling over almost 200 years of around 0.16 percent p.a., and implies an increased speed of mean reversion with a lower half-life of approximately 3 years.[4] In explaining the economic rationale for the trend, they note that over long time spans, some shocks to the real exchange rate are likely contain a permanent component and particularly so in the UK-US context stating: 'Over the sample period considered, Britain underwent the first industrial revolution and became the world's leading industrial and economic power before declining relatively during the twentieth century. During the same period, the United States transformed itself from an exclusively rural economy to take Great Britain's mantle as the leading international economic power. It therefore seems reasonable to suppose that real effects such as the Harrod-Balassa-Samuelson (HBS) would have made themselves felt over the sample' (*ibid*, p.762).

HBS effects can be explained given a country receiving a positive productivity shock in its tradable sector, will see nominal wages not only rise in the traded sector (reflecting the increased marginal product of labour) but, assuming labour is perfectly mobile, also in the non-tradable

[3] Earlier work that employs a long period perspective includes Rogers (1994: Sterling-US dollar rate, 1859-1992), Grilli and Kaminsky (1991: Sterling-US dollar rate, 1885-1986), Kim (1990: five US dollar exchange rates, 1900-1987) and Abuaf and Jorian (1990: eight US dollar exchange rates, 1901-1972).

[4] Lothian and Taylor (2000) estimate an AR(1) process for the Sterling-US dollar series. Without a linear time trend, the AR(1) coefficient is 0.887, with the trend the same coefficient becomes 0.807, hence producing the increased speed of mean reversion and lower half-life in the latter case.

sector. To pay these higher wages, the price level of non-tradables will have to rise. Of course, the exchange rate only adjusts to changes in the tradable sector, leaving a higher relative price level of non-tradables in the higher productivity country. In other words, the real exchange rate would (allowing the higher productivity country to be the numeraire) appreciate and given persistent productivity differentials, present an upward trend.

Whilst testing for a trend (either linear or non-linear) in the real exchange rate is one approach, others exist to assess the prevalence of HBS effects. For example, Lothian and Taylor (2008) include productivity differentials (with productivity proxied by the ratio of real GDP to population) within an ESTAR model of the real exchange rate to capture HBS effects. They find significant HBS effects for the Sterling-US dollar but not for the Franc-Sterling rate from 1820-2001. Chong *et al*. (2012) use a panel of 21 OECD countries over a much shorter period (1973 to 2008) but with a higher, quarterly frequency than some other studies.[5] Again, using GDP per capita to measure productivity, they show within a group-mean panel dynamic OLS framework that HBS effects are significant in the cointegrating vector. Finally, Bordo *et al.* (2017), returning to the use of long span data (i.e., 14 countries from 1880-1987), suggest that the modern, rather than the typically considered traditional form, of the HBS effect is empirically supported. Theoretically, this modern form assumes that if PPP (even in the long run) does not hold in the tradable sector, then productivity differentials in this same sector will not only affect the real exchange via relative price levels in the non-tradable sector (the traditional HBS effect) but also through the terms-of-trade (i.e., the relative price of tradables).[6] This additional component could either lessen the HBS effect, as suggested by Bordo *et al*. (2017), or increase the impact. As a rationale for the latter, Corsetti *et al*. (2008) argue that productivity shocks have wealth effects in their country of origin,

[5] Much of the previously cited work uses an annual frequency given availability over longer time spans.
[6] Bordo *et al.* (2017) use specialization in production or monopolistic competition to perturb PPP.

increasing demand for local goods above supply and consequently, affecting the relative price level and the real exchange rate.[7]

A question less asked in the literature, at least explicitly, is how long HBS effects exert a particular directional pressure on the real exchange rate? The answer to such a question potentially reveals how past economic activity translates through history to the present; an intergenerational policy issue. Although evidence is relatively scarce, Lothian and Devereux (2020) examine an 'ultra-long' data series for the Guilder-Sterling real exchange rate from 1590 to 2018.[8] Using a Bai and Perron (1998, 2003a and 2003b) structural break approach, they show that the real exchange rate is mean reverting and presents a positive intercept break in 1797, followed by a negative intercept break in 1931. Whilst Lothian and Devereux do not explicitly examine a deterministic trend component, they suggest the 135-year higher level of Sterling could be the result of HBS effects due to the higher price of non-tradables over this period caused by the effect of the industrial revolution in the UK.[9,10] Moreover, recent work focussing on England (see Bouscasse *et al.*, 2021) suggests that productivity growth commenced in the 17th century and accelerated in the 19th century, providing further motivation to examine the potential existence of HBS effects prior to the industrial revolution.

To investigate further, this paper compiles new ultra-long data series for 10 real exchange rates, relative to Sterling, in some cases stretching back till 1264 (France and Italy) and running

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[7] As discussed in Berka *et al.* (2018), non-productivity shocks such as a labour market wedge, can also result in deviations from PPP given that higher labour costs lead to an appreciation of the real exchange rate. Work such as Gubler and Sax (2014) also focus on technological change.
[8] Froot *et al*. (2019) in their examination of the law of one price between seven commodities common to England and Holland, obtain an even longer nominal Shilling-Guilder exchange rate back to the mid-to-late 14th century.
[9] Papell and Prodan (2006) examining 16 US dollar real exchange rates (1870-1998) and using unit root tests, include a trend and allow for up to two changes in the intercept. They find that allowing for structural change additional countries to reject the unit root null.
[10] To support this conjecture, Lothian and Devereux (2020) show that London rents rise noticeably from around the beginning of the 19th century.

annually and continuously until 2020.[11] These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556). Having such long series, with consequently greater degrees of freedom, gives more opportunity for a reasonable identification of time series properties; particularly as trends, given they exist, are typically small relative to exchange rate volatility. Of course, as Froot *et al*. (2019) discuss, measurement error, as one moves backward through history is ‘intuitively’ of greater concern; however, given we will be most concerned with low frequency characteristics, this is perhaps less of an issue.[12] More importantly, the relatively low quality of the per capita income in the pre-industrial period is countered by the improved quality of the consumer price data as we go back in time back because the quality and the variety of goods were hardly changing over time as we should expect from endogenous growth models.

In the spirit of Obstfeld (1993) and Lothian and Taylor (2008), we provide a model which shows that presence of time-dependent HBS effects can result not only in the real exchange rate containing a trend, but that both the intercept and the trend may be subject to structural breaks.[13] Consequently, to examine the longevity of HBS effects, we employ trend tests that allow for multiple breaks whilst being robust to whether the real exchange rate presents a unit root. This latter feature is particularly important given debate over the order of integration of the real exchange rate and the effect that the result of any pre-test has on subsequent estimation of time series properties (see Harvey *et al*., 2010, 2017). Strikingly, it is shown that breaks in trend indicate

[11] Other recent work using very long series of data includes Karaman *et al*. (2020) examining the determinants of European monetary stability with regressions over the period 1500-1910 and Rogoff *et al*. (2022) assessing the time series properties (including structural breaks) of long maturity real interest rates since 1311.

[12] Moreover, Froot *et al*. (2019) suggest that measurement error does not play an important role in their results noting, “within-country variation of prices of different goods is not large compared to across-country variation of prices of the same good. Put differently, all the different series yield very similar results, across goods and across centuries” (p.5).

[13] Some studies (see Lothian and Taylor, 2008) have included higher order terms (i.e., cubic) to model the trend. Such functions, empirically, can be observationally equivalent to allowing for appropriate breaks in a linear trend.

that HBS effects may last, in some cases, for several centuries. To assess further, and while bearing in mind potential data quality issues noted earlier, we obtain real GDP per capita from the Maddison Project Database, version 2020,[14] with long enough series available for the UK (commencing 1252), France (1280), Italy (1310), Sweden (1300), Netherlands (1348) and Portugal (1530) and analysis revealing an ultra-long association between real exchange rates and productivity differentials proxies. The paper continues in the following manner: Section 2 contains theoretical considerations, whilst Section 3 outlines the empirical methodology. Section 4 presents the new data, whilst Section 5 presents the empirical results and related discussion. Section 6 concludes.

2. Theoretical underpinnings

The real exchange rate, Q, can be defined as:

$$Q = \frac{SP}{P^*} \tag{1}$$

where S is the nominal exchange rate expressed as foreign currency per unit of domestic currency, P is the logarithm of the domestic price level and P^*, the foreign price level. For a model, we begin from the perspective of the home economy and assume symmetric equations for that of the foreign economy. Each economy has two sectors, one producing a tradable good (*T*) and the other, a non-tradable good (*N*). Allow *C*, the consumption basket, to embody consumer preferences using a Cobb-Douglas function:

$$C = \frac{C_N^{\gamma} C_T^{1-\gamma}}{\gamma^{\gamma}(1-\gamma)^{1-\gamma}} \tag{2}$$

[14] See Bolt and van Zanden (2020).

where C_N and C_T are the consumption of non-tradable and tradable goods, respectively, and γ is the expenditure share of non-tradables. It can be shown that (2) leads to the following geometric weighted price index:

$$P = P_N^{\gamma} P_T^{1-\gamma} \tag{3}$$

where P_N and P_T are the prices of non-tradable and tradable goods, respectively. As noted by Lothian and Taylor (2008), labour is perfectly mobile between sectors in the long-run and therefore:

$$\frac{W_N}{P} = \frac{W_T}{P} \tag{4}$$

with W_N and W_T representing the nominal wage level in the non-tradable and tradable sectors, respectively, and (4) embodying long-run real and nominal wage equality. Also, allowing L_N and L_T to be the marginal product of labour in each sector, and assuming that over the long-run, the nominal wage, W, is the marginal revenue product of labour then:

$$W = W_N = L_N P_N \tag{5a}$$

$$W = W_T = L_T P_T \tag{5b}$$

Substituting (5a) into (3) gives:

$$P = P_T \left(\frac{L_T}{L_N}\right)^{\gamma} \tag{6}$$

Given (6) holds for the foreign economy as well, and the law of one price holds among tradables, then Q in (1) can be expressed as:

$$Q = \frac{SP}{P^*} = \frac{(L_T/L_N)^{\gamma}}{(L_T^*/L_N^*)^{\gamma^*}} \tag{7}$$

Assuming $L_N = L_N^*$ and taking logarithms (represented by lowercase letters) of (7) results in:

$$q = \gamma l_T - \gamma^* l_T^* \tag{8}$$

Equation (8) conveys the HBS effect whereby relative improvements in (tradable good) productivity in the home economy, leads to a real appreciation of the home currency relative to the foreign currency. Of course, the static analysis thus far can be transferred into a dynamic context, with the long-run equilibrium real exchange rate, q_t:

$$q_t = \gamma(l_t - l_t^*) \tag{9}$$

with l_t and l_t^* representing appropriate productivity measures at time *t*, and for simplicity assuming that $\gamma = \gamma^*$. Of course, the wedge between home and foreign productivity could present various dynamics including a linear time trend where $\gamma(l_t - l_t^*) = \delta t$, assuming a constant rate of differential productivity growth (see Obstfeld, 1993) generated by technology shocks. However, it is plausible that the growth itself is time-varying and that we need to allow for trend breaks, or that a technology shock could produce an immediate level shift in the differential. Encompassing these deterministic possibilities, examining (9) could be carried out within the following regression framework:

$$q_t = \alpha + \beta t + \sum_{j=1}^{m} \theta_j DT_{jt}(B_j) + \sum_{j=1}^{m} \mu_j DU_{jt}(B_j) + u_t,\ t = 1, \dots, T \tag{10}$$

where B_j denotes the break dates, $DT_{jt}(B_j) = 1(t > B_j)(t - B_j)$, $DU_{jt}(B_j) = 1(t > B_j)$ and $1(.)$ represents an indicator function.

3. Empirical methodology

Initially, when thinking about a DGP for the logarithm of the real exchange rate, consider a break-free version of (10) with:

$$q_t = \alpha + \beta t + u_t,\ t = 1, \dots, T \tag{11}$$

$$u_t = \rho u_{t-1} + \varepsilon_t,\ t = 2, \dots, T \tag{12}$$

and ε_t following a stationary process. As discussed earlier, and given we are principally interested in examining $H_0: \beta = 0$, to circumvent any order of integration issues and the effect unit root pre-tests have on trend estimation we allow u_t to be $I(0)$ or $I(1)$ by assuming ρ lies in the region $-1 < \rho \leq 1$.[15] An appropriate approach in this context employs the $t_{\beta}^{RQF}(MU)$ statistics of Perron and Yabu (2009a) which, under the null, follows an asymptotic standard normal distribution.[16]

Given the ultra-long real exchange rate series to be examined, if trends exist due to differential productivity growth, it is possible that breaks in such series exist as well. A single-break counterpart to the Perron and Yabu (2009a) is provided by Perron and Yabu (2009b), whilst Kejriwal and Perron (2010) provide a sequential, multiple-break generalisation of the latter procedure. In particular, replacing (11) with (10), the initial step for Kejriwal and Perron (2010) is to employ the Perron and Yabu (2009b) test to assess the null of no breaks versus the one break in level/trend alternative. In the limit, the distribution of their *Exp-W* test statistic varies depending on whether u_t is $I(0)$ or $I(1)$. However, Perron and Yabu (2009b) also show that critical values (c.v.) at usual levels of significance are "nearly the same", recommending that a more conservative approach is adopted by using the maximum of the $I(0)$ or $I(1)$ c.v.[17]

As Kejriwal and Perron (2010) is a sequential process, if we do not reject the null of no breaks, the procedure finishes. On the other hand, if the null is rejected, then a second step conditional on the existence of at least one break (i.e., $k = 1$) is undertaken. This employs the test statistic $F_T(k+1|k)$, assessing whether more than one break is present by comparing the test

[15] Along with others, Harvey *et al.*, (2010) note that 'incorrect' order of integration test results lead to trend tests with severe size distortions or that are inefficient depending on whether the true DGP respectively contains a unit root or not.

[16] Other robust to order of integration trend tests exist including Skrobotov (2022), Harvey *et al.* (2007) and Bunzel and Vogelsang (2005). As Skrobotov (2022) notes, Perron and Yabu (2009a) and Harvey *et al.* (2007) are "uniformly the most powerful in terms of one-sided alternatives and uniformly most powerful unbiased in the case of the two-sided alternative."

[17] We set the trimming parameter $\pi = 0.1$.

statistic to c.v. also obtained from Kejriwal and Perron (2010). In our empirical work, we set three as the maximum number of breaks; theoretically, the sequential procedure could continue until there is no evidence for further breaks – however, Kejriwal and Perron note that the smaller subsamples resulting from allowing for too many breaks in finite samples would likely generate size distortions and tests with low power.

Of course, as noted in the introduction, there can be other approaches to testing for HBS effects rather than looking for trends and trend breaks in q_t. For example, if q_t and $(l_t - l_t^*)$ are $I(1)$, then (9) implies that they will be cointegrated.[18] To examine this possibility, and where data availability allows, we estimate the following FM-OLS cointegrating regression:

$$q_t = \varphi + \omega(y_t - y_t^*) + v_t \,, \qquad (13)$$

employing the Bartlett kernel and Newey-West fixed bandwidth for the long-run covariance estimate and constructing the Phillips-Ouliaris τ and z test statistics PO_τ and PO_z, respectively.

4. Data

We build a novel unbalanced panel dataset of 10 real exchange rates relative to Sterling with the earliest rates dating back to 1264, and running annually and continuously until 2020: France and Italy (commencing 1264), Sweden (1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556). Constructing such a dataset is a non-trivial task involving many sources and these are listed in the Appendix. Exchange rates were chosen on the basis of data availability till the present day. It's important to note that whilst we use contemporary country nomenclature to refer to currencies (e.g., France, Netherlands), the names, geographical boundaries and polity have often changed over the nine centuries spanned.

[18] At this juncture, it should be noted that unit root and trend-stationary with breaks are often observationally equivalent in finite samples.

For example, the Treaty of Verdun divided the Carolingian Empire in 846, and the resulting westernmost component, West Francia became the Kingdom of France in 987 with East Francia, the Kingdom of Germany. However, whilst France remained a unitary state, East Francia was subsumed within the Holy Roman Empire which itself was dissolved in 1805. Germany became a nation state in 1871 with the unification of many smaller German states. In such cases, and to obtain earlier exchange rates, data from a number of cities and city states (e.g., Munich, Antwerp, Florence, Milan and Venice) are employed.

[Insert Figure 1 around here]

Figure 1 contains plots of the log real exchange rate (q_t) for the ultra-long series. Perhaps the most notable feature the graphs, and common to all series, is the sharp drop and subsequent rebound from the early 1540s to the early 1550s. This is due to a concomitant movement in our measure of Sterling over the same period; analogously, Karaman *et al*. (2020) find a devaluation in England's monetary unit (in terms of silver) and relate this to the Great Debasement (1544-1551) when the mint's silver output increased approximately sevenfold (Palma, 2018). To ameliorate the effect of this common spike on trend tests, we additionally apply a simple linear interpolation from 1540-1555 and the resulting graphs can be found in Figure 2.

[Insert Figure 2 around here]

Eyeballing the series in Figures 1 and 2, several appear to contain a downward or no trend until the mid-to-late 1500s, a long upward trend thereafter until the 1800s, followed eventually by a downward trend over the late 19th century and 20th century. In the next section, we shall assess these observations more formally.

In subsequent analysis, we employ differential growth rates in real GDP per capita as a proxy for productivity (see Lothian and Taylor, 2008). The Maddison Project Database, version

2020, provides ultra-long, annual and almost continuous real GDP per capita data for six of our countries: the UK (commencing 1252), France (1280), Italy (1310), Sweden (1300), Netherlands (1348) and Portugal (1530) and the sources used by the Maddison Project are also listed in the Appendix. For other countries, the data is relatively incomplete and thus the bivariate analysis involves five Sterling real exchange rates and the attendant country real GDP per capita differentials.[19]

5. Empirical results

Table 1 and Table 2 present the robust to the order of integration $t_{\beta}^{RQF}(MU)$ test statistic, discussed in section 3, for each of the ultra-long series (i) without adjustment and (ii) with adjustment for the Great Debasement, respectively.

[Insert Table 1 and Table 2 around here]

Notably, the null of no trend cannot be rejected for any of the ultra-long series. At first sight, one might be tempted to conclude that HBS effects cannot be detected over these long time periods. However, given the earlier theoretical discussion and the eyeballing of the data, there is at least a *prima facie* case for multiple trend breaks which change both magnitude and sign over the centuries and are thus masked by a test which examines the presence of a single trend. To examine this and first employing the ultra-long series without any adjustment for the Great Debasement,

[19] For Spain, real GDP per capita data begins in 1285 but is then recorded each decade (with an additional value in 1820) until running continuously from 1850. For Belgium, data begins in 1500 and is generally recorded each 50 years (although with no value for 1800 and additional values for 1812 and 1820) until running continuously from 1846. For Germany, data also begins in 1500, is next recorded in 1600 before occurring every 50 years (with an additional value in 1820) until running continuously from 1850. For Austria, data begins in 1820 and is recorded each decade until running continuously from 1870. For Denmark, data runs continuously from 1820.

Table 3 provides the results for the Kejriwal and Perron (2010) sequential break procedure, which by construction, is also robust to the order of integration of series.

[Insert Table 3 around here]

In particular, Table 3 shows the results for each step of the sequential procedure, reporting the test statistic $F_T(k+1|k)$ and, given a rejection in favour of $k+1$ breaks, estimated break date(s) are also shown. Strikingly, at the 1% significance level, there is evidence of in favour of three breaks for all the ultra-long real exchange series providing the first statistical indication of multiple breaks that would otherwise remain latent. However, it is also notable that several of the estimated break dates occur in and around the England's Great Debasement in the mid-16th century. To circumvent any potential effects of this, the Kejriwal and Perron (2010) sequential break procedure is applied to the ultra-long series with an adjustment for the Great Debasement, and the results are disclosed in Table 4.

[Insert Table 4 around here]

It can be observed that there is evidence in favour of three breaks for all Great Debasement adjusted series – again, at the 1% significance level. Moreover, estimated break dates are now more evenly distributed throughout the long sample period, with the first break most frequently occurring in the 1500s. The second break dates provide more variability, being spread between the 16th century and 19th century, with a mean of 1736 – whilst the third break dates, are again more concentrated, occurring most often in the 20th century. Before assessing the historical relevance of these periods, it is useful to know the magnitude and sign of any break. To do so, Table 5 presents the fitted values at these minimum sum of squared residual (SSR) dates and Figure 3, provides a corresponding graphical representation.[20]

[20] In other words, the fitted values from (10).

[Insert Table 5 and Figure 3 around here]

Using these, four broad regimes for trends in ultra-long real exchange rates relative to Sterling can be elucidated. *The first regime*, typically lasts from the start of the series to the 1500s and roughly corresponds to the Late Medieval Period and aftermath; the trend sign in this regime is varied (although the majority are negative) and consequently conditional on the country counterparty. *The second regime,* which lasts on average 167 years, runs primarily during the Early Modern Period and presents almost exclusively, a positive trend. *The third regime*, lasts on average until 1893, and hence typically encompasses the Industrial Revolution; in this case, sign is country-dependent (although the majority are positive) whilst, *the fourth regime*, running predominantly through the 20th century and early 21st century, uniformly presents a negative trend during the period of relative UK decline.

To provide further historical interpretation, consider the path of some individual real exchange rates. For example, during the first regime and over the 261-year period from 1264 to 1525, the UK real exchange rate appreciated relative to the French currency (i.e., the livre tournois) with a trend rate of growth equivalent to 25 basis points annually. In part, this would seem due to the countries' different policies on debasement. Spufford (1988) notes that: *"[…] England and Aragon were both exempt from the waves of debasement that intermittently afflicted the nobility of France and Castille so badly. Successive English and Aragonese kings kept to a strong monetary policy throughout both the fourteenth century and the fifteenth, and never succumbed to the temptation to debase the coinage to pay for their wars. The English, at least, had a far more sophisticated and efficient tax system"* (p.316).[21]

[21] Other relevant work includes Bell *et al*. (2017) and Miskimin (1963, 1984).

The second regime, coinciding with the Early Modern Period or the Age of Exploration, saw seafaring European nations and, in particular Portugal, Spain, the Netherlands, France and the UK, explore outside the European sphere, establishing international trade routes and colonial empires. Importantly, the 200-year period between 1550 and 1750, saw the UK become the dominant naval power and relatedly, its real GDP per capita grew faster relative to most other countries in our sample, likely underpinning the almost ubiquitous positive movement in the real exchange rate over the same period. [22, 23]

[Insert Table 6 around here]

To illustrate the relative changes in real GDP per capita, Table 6 employs data from the Maddison Project Database, version 2020,[24] showing for example, that whilst French GDP per capita grew by 12.6% between 1550 and 1750, UK GDP per capita increased by a salient 63.4%. This corresponds with recent work by Bouscasse *et al*. (2021) showing that productivity growth in England began in 1600.

The third regime, coinciding with the Industrial Revolution, saw the UK become the richest country in our sample (and indeed, the World) by 1850. Moreover, Table 6 shows the UK GDP per capita grew by 181.1% between 1750 and 1900, outpacing many countries in our sample including the Netherlands (40.1%) and Portugal (-5.0%) and corresponding with the positive real exchange rate trend that can be seen in Figure 3 for those countries over the same period. It also corresponds with work such as Mokyr (2010) suggesting that the UK's technological progress at this juncture was spread across a range of industries, being enabled by innovative thinking

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[22] Rodgers (2004) notes that during the 18th and 19th centuries, the Royal Navy was the largest, best financed and most advanced maritime organisation globally.

[23] Growing even faster than the UK, Portuguese GDP per capita grew 64.1%. However, if we extend the focal period by 50 years (i.e., from 1550 to 1800) whilst UK GDP per capita grew by 102.1% over the 250 year period, Portuguese GDP per capita increased by only 9.6%, a figure primarily due to the collapse in economic growth in the latter half of the 17th century (see Costa *et al*., 2015).

[24] See Bolt and van Zanden (2020).

rendered into saleable products by technically advanced human capital. However, some countries closely matched the UK's economic performance, including Belgium (174.2%), France (173.3%) and Germany (184.2%). Indeed, from 1800 to 1900, while the UK's GDP per capita increased 127.2%, Germany's increased by 202.7%, supporting the negative trend in the German-UK real exchange rate during almost the entirety of the 19th century.

Finally, during the fourth regime, the UK's relative economic decline can be observed in Table 6, where every country outperforms the UK in GDP per capita growth over the 1900 and 2000, an occurrence that matches the common negative trend in all our ultra-long real exchange rates. Relatedly, the extant literature has commented on how Britain lost its technological leadership at the beginning of the 20th century, reinforcing the deterioration of the real exchange rate; for example, Nicholas (2014) argues that while relative technological decline for the UK was unavoidable given the catch-up by other industrial nations and some, such as the US, possessing "*larger markets which promoted demand-induced innovation, larger populations leading to a greater supply of human capital for invention, or deeper capital markets to promote investment*" (p.182), this decline was exacerbated by structural factors including low product market competition. Others such as Broadberry and Crafts (1992) additionally stress low human capital, restrictive worker practices and firm collusion during the interwar period, with decline still contemporaneously underpinned by issues such as poor management practice (Bloom and Van Reenen, 2007) and organisational inflexibility leading to inefficient use of new technology (see Bloom *et al*., 2012).[25]

As discussed previously, alongside trend analysis, HBS effects can be investigated via a bivariate approach; the theory in Section 2 suggesting that time-varying differential productivity

[25] Broadberry (1997) emphasises that, since the 1850s, UK relative labour productivity in manufacturing has remained approximately stationary and therefore explanations for the overall fall in labour productivity must look elsewhere.

growth may contribute towards trend breaks in real exchange rates over the very long-run. As shown above, descriptively, there seems to be a reasonable correspondence between real exchange rate trends and differential growth rates in real GDP per capita – the latter being a proxy for productivity, and as mentioned previously, used by work such as Lothian and Taylor (2008), who show significant HBS effects for the Sterling-US dollar but not for the Franc-Sterling rate over the period 1820-2001.

To investigate the linkage between ultra-long real exchange rates and differential productivity growth, we first plot in Figure 4, the log real exchange rate (q_t) and the quantity ($l_t - l_t^*$) from (9) proxied by ($y_t - y_t^*$), where y_t represents log real GDP per capita using available data from The Maddison Project Database, version 2020.

[Insert Figure 4 around here]

Once again, eyeballing the plots in Figure 4 reinforces the view that there may be an association between the two series with, typically, concomitant long positive movements (often over hundreds of years) followed by steep declines towards the end of the sample.

[Insert Tables 7 and 8 around here]

Of course, for a bivariate analysis we need to establish that both series have a common order of integration. Unit root test results in Table 7 confirm that all series can be adjudged as presenting a unit root and therefore, Table 8 provides the estimation of FM-OLS cointegrating regression (13). Strikingly, for all countries, the log real exchange rate (q_t) and country differentials in real (log) GDP per capita ($y_t - y_t^*$) are found to be cointegrated over many centuries – for example, in the case of the France-UK rate, over 738 years.

6. Conclusions

Constructing a novel dataset of ten ultra-long real exchange rates covering much of the last millennia and using trend tests that are robust to the usual pre-testing, we show Sterling's behaviour can be characterised by four regimes with typically strong trend-appreciation during the early modern period, balanced by a striking trend-depreciation over the period, either prior to, or since the beginning of the 20th century. It is notable that some trends last for several centuries.

To explain trend duration and direction, we turn to Harrod-Balassa-Samuelson (HBS) effects – the long-term movements in real exchange rates due to persistent and differing country productivity levels. These have been evidenced in the extant literature (e.g., Lothian and Taylor, 2008; Bordo *et al.*, 2017; Lothian and Devereux, 2020) and our analysis extends this work both backwards, in terms of sample period, and scope, by considering a much wider range of real exchange rates; our results showing that HBS effects prevailed not only in the industrial age but also in the pre-industrial period.

Given theory, we examine the bivariate relation between real exchange rates and country GDP per capita differentials (the latter as a proxy for productivity differences) showing that even over a period extending back to the 13th century, the series are cointegrated. This finding, coupled with the trend-depreciation in Sterling towards the end of the sample, implies that the UK's current productivity decline is an entrenched and persistent phenomena.

Finally, this work is a useful counterpart to other recent studies employing ultra-long economic and financial series. In particular, Karaman *et al*. (2020) examine the determinants of European monetary stability, with regressions over the period 1500-1910 showing annual changes in the value of a monetary unit are associated with taxation, war and political factors. Moreover, Rogoff *et al*. (2022) assess the time series properties of long maturity real interest rates, finding a negative trend from the 14th century. Complementing this work, we show that ultra-long real exchange rates from 1264 present multiple trends and are closely associated with country-level productivity differentials.

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Tables and Figures

Table 1: Tests for a trend in the (log) real exchange rates for ultra-long series

Notes: ***, **, * denotes rejection at the 1%, 5% and 10% significance level, respectively. Data begins as early as 1264, without any adjustment for the Great Debasement. The $t_{\beta}^{RQF}(MU)$ test statistic, under the null of no trend in (11), follows an asymptotic standard normal distribution. Additionally, the 'Growth rate (%)' is the estimate of β in (11) multiplied by 100 with '95% c.i.,' the accompanying 95 percent confidence interval. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

Country	$t_{\beta}^{RQF}(MU)$	Growth rate (%)	95% c.i.
Austria	0.4290	0.13	-0.48, 0.74
Belgium	-0.4545	-0.19	-1.02, 0.63
Denmark	0.3383	-0.07	-0.69, 0.55
France	1.0122	0.24	-0.22, 0.69
Germany	0.3209	0.09	-0.47, 0.65
Italy	0.0546	0.02	-0.57, 0.60
Netherlands	0.0197	0.01	-0.66, 0.67
Portugal	-0.0900	-0.03	-0.64, 0.59
Spain	0.3113	0.12	-0.64, 0.88
Sweden	-0.3119	-0.05	-0.40, 0.29

Table 2: Tests for a trend in the (log) real exchange rates for ultra-long series (Great Debasement adjusted)

Notes: ***, **, * denotes rejection at the 1%, 5% and 10% significance level, respectively. Data begins as early as 1264, with adjustment for the Great Debasement. The $t_{\beta}^{RQF}(MU)$ test statistic, under the null of no trend in (11), follows an asymptotic standard normal distribution. Additionally, the 'Growth rate (%)' is the estimate of β in (11) multiplied by 100 with '95% c.i.,' the accompanying 95 percent confidence interval. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

Country	$t_{\beta}^{RQF}(MU)$	Growth rate (%)	95% c.i.
Austria	0.445	0.13	-0.45, 0.72
Belgium	-0.6774	-0.19	-0.74, 0.36
Denmark	0.3383	0.06	-0.30, 0.42
France	1.4507	0.24	-0.08, 0.56
Germany	0.3611	0.09	-0.41, 0.59
Italy	0.0635	0.02	-0.49, 0.52
Netherlands	0.033	0.01	-0.39, 0.40
Portugal	-0.09	-0.03	-0.64, 0.59
Spain	0.4651	0.12	-0.39, 0.63
Sweden	-1.5708	-0.06	-0.13, 0.01

Table 3: Sequential tests for multiple breaks in level/trend - (log) real exchange rates for ultra-long series

Notes: ***, **, * denotes rejection at the 1%, 5% and 10% level, respectively. Data begins as early as 1264, without any adjustment for the Great Debasement. The $F_t(k+1|k)$ test statistic provides a test of $k+1$ against k breaks using critical values obtained from Kejriwal and Perron (2010). Conditional on $k+1$ breaks being detected, the break dates (i.e., $B_1, .., B_k$) are estimated by minimising the global sum of squared residuals (SSR) from an OLS estimation of (10) across all candidate dates. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

Country		$F_t(k+1\|k)$	Estimated break date(s)		
Austria	$F_T(1\|0)$	144.2769***	1914		
	$F_T(2\|1)$	68.6683***	1858	1946	
	$F_T(3\|2)$	41.5677***	1541	1551	1947
Belgium	$F_T(1\|0)$	179.9214***	1923		
	$F_T(2\|1)$	69.9221***	1588	1924	
	$F_T(3\|2)$	46.8183***	1538	1551	1923
Denmark	$F_T(1\|0)$	73.8019***	1820		
	$F_T(2\|1)$	14.8976***	1756	1820	
	$F_T(3\|2)$	3.3826**	1722	1809	1965
France	$F_T(1\|0)$	122.0337***	1717		
	$F_T(2\|1)$	27.6047***	1541	1718	
	$F_T(3\|2)$	13.8119***	1538	1551	1718
Germany	$F_T(1\|0)$	184.8064***	1764		
	$F_T(2\|1)$	58.6091***	1541	1842	
	$F_T(3\|2)$	11.3809***	1541	1551	1850
Italy	$F_T(1\|0)$	104.518***	1672		
	$F_T(2\|1)$	30.8712***	1527	1738	
	$F_T(3\|2)$	95.5852***	1537	1551	1747
Netherlands	$F_T(1\|0)$	154.4005***	1575		
	$F_T(2\|1)$	35.8904***	1553	1931	
	$F_T(3\|2)$	49.4843***	1551	1623	1931
Portugal	$F_T(1\|0)$	21.2379***	1660		
	$F_T(2\|1)$	32.9266***	1551	1750	
	$F_T(3\|2)$	16.8172***	1552	1662	1836

Spain	$F_T(1\|0)$	12.636***	1655		
	$F_T(2\|1)$	47.1896***	1552	1666	
	$F_T(3\|2)$	86.2347***	1552	1656	1964
Sweden	$F_T(1\|0)$	25.537***	1610		
	$F_T(2\|1)$	40.6529***	1541	1849	
	$F_T(3\|2)$	34.2185***	1562	1575	1815

Table 4: Sequential tests for multiple breaks in level/trend - (log) real exchange rates for ultra-long series (Great Debasement adjusted)

Notes: ***, **, * denotes rejection at the 1%, 5% and 10% significance level, respectively. Data begins as early as 1264, with adjustment for the Great Debasement. Notes: ***, **, * denotes rejection at the 1%, 5% and 10% level, respectively. The $F_t(k+1|k)$ test statistic provides a test of $k+1$ against k breaks using critical values obtained from Kejriwal and Perron (2010). Conditional on $k+1$ breaks being detected, the break dates (i.e., $B_1, .., B_k$) are estimated by minimising the global sum of squared residuals (SSR) from an OLS estimation of (10) across all candidate dates. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

Country		$F_t(k+1\|k)$	Estimated break date(s)		
Austria	$F_T(1\|0)$	34.0727***	1914		
	$F_T(2\|1)$	26.0249***	1858	1946	
	$F_T(3\|2)$	28.8357***	1628	1858	1946
Belgium	$F_T(1\|0)$	207.1553***	1923		
	$F_T(2\|1)$	87.4492***	1909	1952	
	$F_T(3\|2)$	80.0113***	1527	1906	1953
Denmark	$F_T(1\|0)$	73.8019***	1820		
	$F_T(2\|1)$	14.8976***	1756	1820	
	$F_T(3\|2)$	3.3826**	1722	1809	1965
France	$F_T(1\|0)$	35.8356***	1717		
	$F_T(2\|1)$	20.5135***	1525	1718	
	$F_T(3\|2)$	13.1149***	1525	1718	1798
Germany	$F_T(1\|0)$	45.8324***	1764		
	$F_T(2\|1)$	42.3059***	1593	1850	
	$F_T(3\|2)$	23.2458***	1593	1809	1916
Italy	$F_T(1\|0)$	35.4878***	1690		
	$F_T(2\|1)$	22.5459***	1527	1751	
	$F_T(3\|2)$	16.2952***	1378	1527	1751
Netherlands	$F_T(1\|0)$	20.3865***	1931		
	$F_T(2\|1)$	53.5466***	1575	1931	
	$F_T(3\|2)$	4.1902**	1581	1799	1931

Portugal	$F_T(1\|0)$	8.2438***	1660		
	$F_T(2\|1)$	32.9266***	1567	1740	
	$F_T(3\|2)$	21.048***	1580	1687	1914
Spain	$F_T(1\|0)$	39.3297***	1655		
	$F_T(2\|1)$	47.7989***	1587	1721	
	$F_T(3\|2)$	42.1722***	1589	1668	1945
Sweden	$F_T(1\|0)$	28.7586***	1809		
	$F_T(2\|1)$	16.1392***	1562	1849	
	$F_T(3\|2)$	9.3145***	1562	1575	1815

Table 5: Fitted values of (10) at minimum SSR dates - (log) real exchange rates for ultra-long series (Great Debasement adjusted)

Notes: Data begins as early as 1264, with adjustment for the Great Debasement. Results represent fitted values of the parameters of (10) at the minimum SSR dates (i.e., the estimated break dates) shown in Table 4. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

Country	α	β	θ_1	θ_2	θ_3	μ_1	μ_2	μ_3
Austria	-0.46	0.0013	0.0035	0.0137	-0.0262	0.32	-0.94	-1.12
Belgium	1.55	0.0063	-0.0028	-0.0659	0.0601	-0.51	0.04	0.12
Denmark	1.99	0.0056	-0.0102	0.002	-0.0022	0.17	0.49	-0.29
France	-1.26	0.0023	0.0048	-0.0104	-0.0005	-0.53	0.48	0.39
Germany	-0.06	-0.0040	0.0098	-0.0165	0.0060	0.31	0.41	0.59
Italy	-0.00	-0.0019	0.0015	0.0092	-0.0147	0.60	-0.52	0.01
Netherlands	0.26	-0.0108	0.0119	0.0014	-0.0090	0.82	0.23	-0.19
Portugal	0.27	-0.0178	0.0329	-0.0141	-0.0097	0.09	0.26	0.14
Spain	-0.46	-0.0139	0.0295	-0.0125	-0.0142	-0.02	0.37	0.20
Sweden	3.45	-0.0020	-0.1483	0.1514	-0.0064	0.01	1.82	0.27

Table 6: Real GDP per capita in 2011$

Notes: Source is Maddison Project Database, version 2020.

Country	1550	1600	1650	1700	1750	1800	1850	1900	1950	2000
Austria	-	-	-	-	-	-	2630	4594	5907	34796
Belgium	2410	2533	2303	2192	2169	-	2944	5947	8706	33720
Denmark	-	-	-	-	-	-	2817	4809	11067	39021
France	1490	1610	1659	1694	1677	-	2546	4584	8266	33410
Germany	-	1286	1511	1497	1674	1572	2276	4758	6186	33367
Italy	2527	2404	2465	2604	2703	2404	2611	3264	5582	32717
Netherlands	2884	4270	4316	3377	3777	4184	3779	5306	9558	37900
Portugal	1331	1258	1322	1572	2184	1459	1470	2075	3325	23372
Spain	-	-	-	-	-	-	1706	2676	3464	26995
Sweden	1655	1258	1388	2002	1554	1366	1715	3320	10742	34203
UK	1654	1691	1446	2412	2702	3343	4332	7594	11061	31946

Table 7: Unit root tests - (log) real exchange rates for ultra-long series (Great Debasement adjusted) and country differentials in real (log) GDP per capita

Notes: In Table 7 and for both the ADF and the Ng and Perron (2001) MZ_{α} and MZ_t tests, we allow for an intercept and employ SIC and modified AIC to determine the appropriate lag length k respectively. The Ng and Perron tests are obtained on the basis of GLS detrending. In terms of sample period, cases stretch back till 1280 (France) and run annually and continuously until 2020. These include Sweden (commencing 1300), Italy (1310), and Netherlands and Portugal (1500).

Country		ADF	MZ_{α}	MZ_t	k(ADF/ MZ)
France	q_t	-1.78	0.28	0.32	7/16
	$y_t - y_t^*$	-2.34	-1.96	-0.97	2/9
Italy	q_t	-1.77	-1.55	-0.86	7/9
	$y_t - y_t^*$	-2.34	-0.09	-0.07	2/9
Netherlands	q_t	-2.23	-3.51	-1.33	2/9
	$y_t - y_t^*$	-2.08	-5.23	-1.61	2/8
Portugal	q_t	-1.48	-0.57	-0.47	3/10
	$y_t - y_t^*$	-1.32	-1.75	-0.94	3/6
Sweden	q_t	-5.70***	-2.21	-0.88	1/7
	q_t'	-2.86**	-0.70	-0.38	6/7
	$y_t - y_t^*$	-2.03	-3.21	-1.24	3/13

Table 8: Fully modified least squares (FMOLS) cointegrating regression

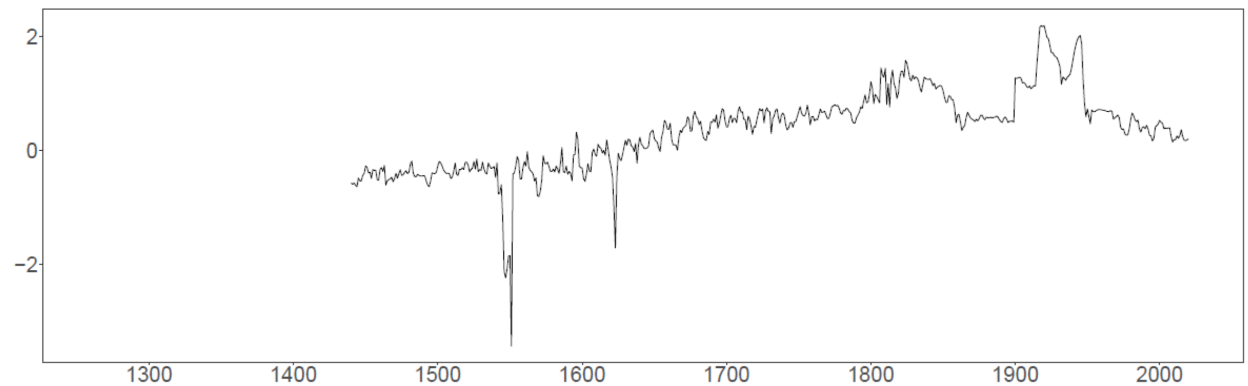
Notes: Table 8 shows the FM-OLS estimation of $q_t = \varphi + \omega(y_t - y_t^*) + v_t$, employing the Bartlett kernel and Newey-West fixed bandwidth for the long-run covariance estimate. PO_τ and PO_z are the Phillips-Ouliaris τ and z test statistics respectively. When intercept dummies are included this refers to the inclusion of identified level dummies from the earlier break tests, as deterministic regressors in the FM-OLS regression. In terms of sample period, cases stretch back till 1280 (France) and run annually and continuously until 2020. These include Sweden (commencing 1300), Italy (1310), and Netherlands and Portugal (1500).

Country	φ	ω	Incl. intercept dummies	PO_τ	PO_z
France	-0.37	2.15	N	-5.38***	-54.81***
	-0.74	0.75	Y	-6.70***	-81.45***
Italy	0.58	0.74	N	-3.98***	-30.43***
	0.03	0.20	Y	-3.97***	-31.02***
Netherlands	0.58	0.70	N	-5.07***	-47.61***
	0.00	0.41	Y	-8.11***	-112.81***
Portugal	-0.14	0.96	N	-2.96	-16.12
	-0.81	0.24	Y	-4.87***	-44.19***
Sweden	2.92	0.12	N	-5.83***	-60.44***
	3.17	0.47	Y	-9.48***	-154.68***

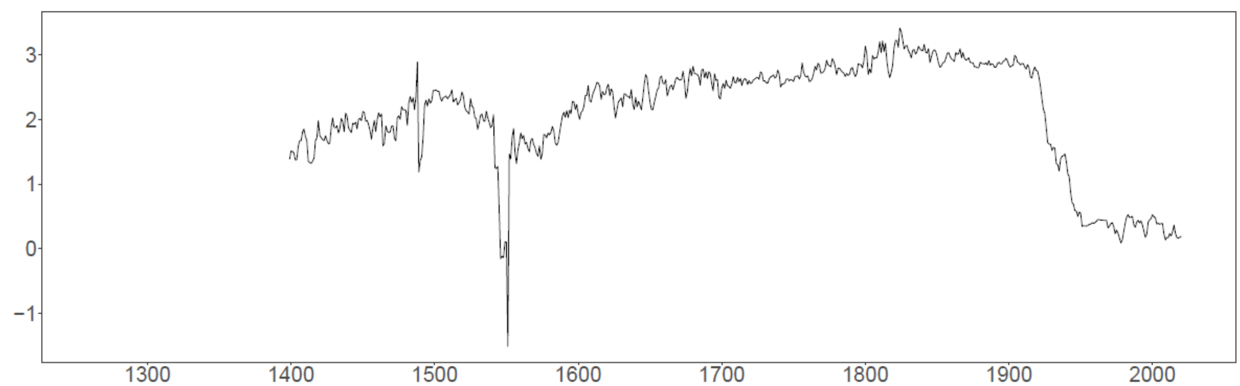
Figure 1: (Log) real exchange rates for ultra-long series (1264 – 2020)

Notes: Figure 1 contains plots of the log real exchange rate (q_t) for the ultra-long series and without any adjustment for the Great Debasement. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

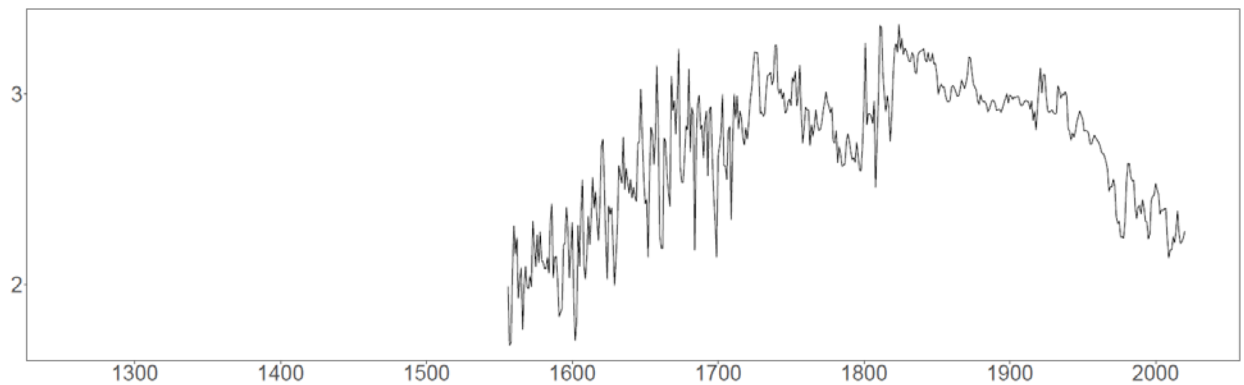
Austria



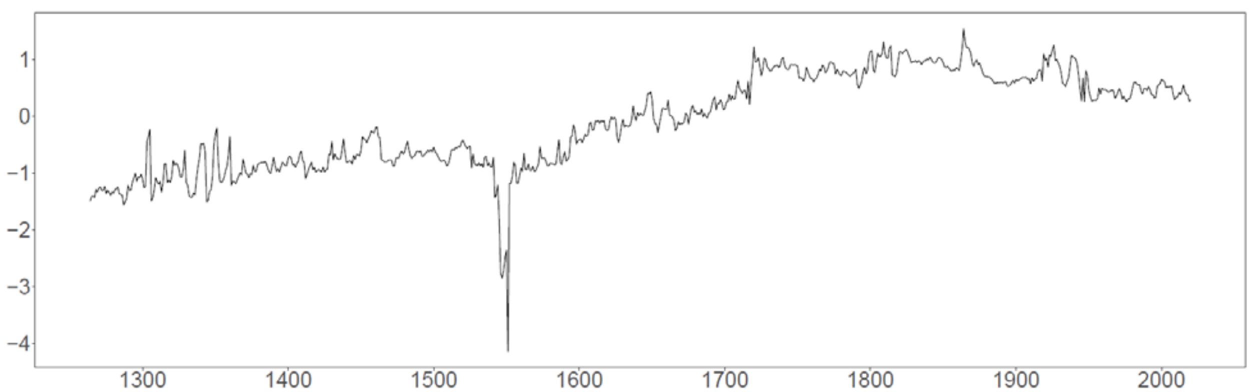
Belgium



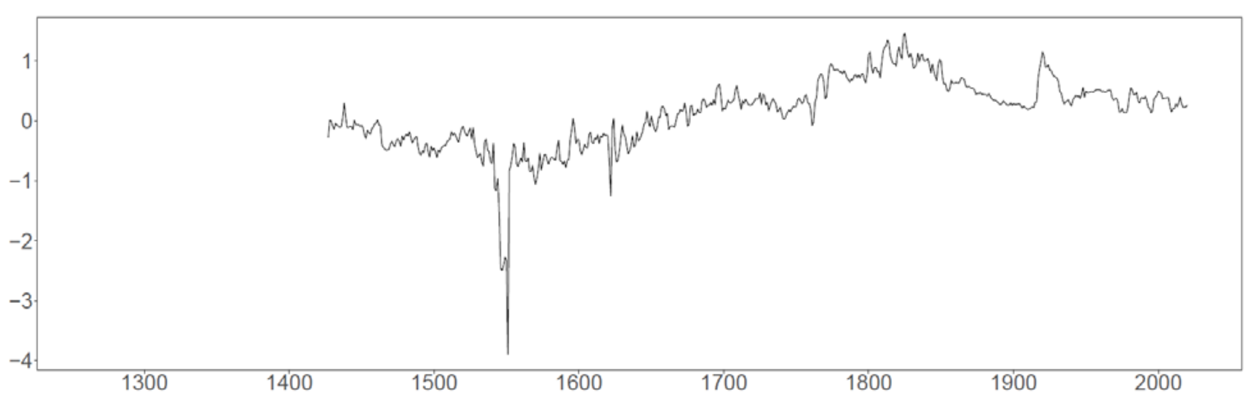
Denmark



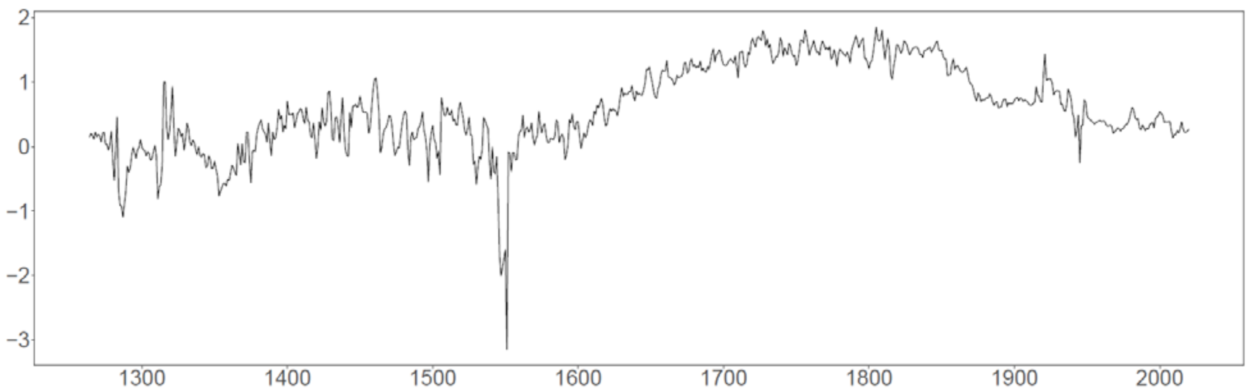
France



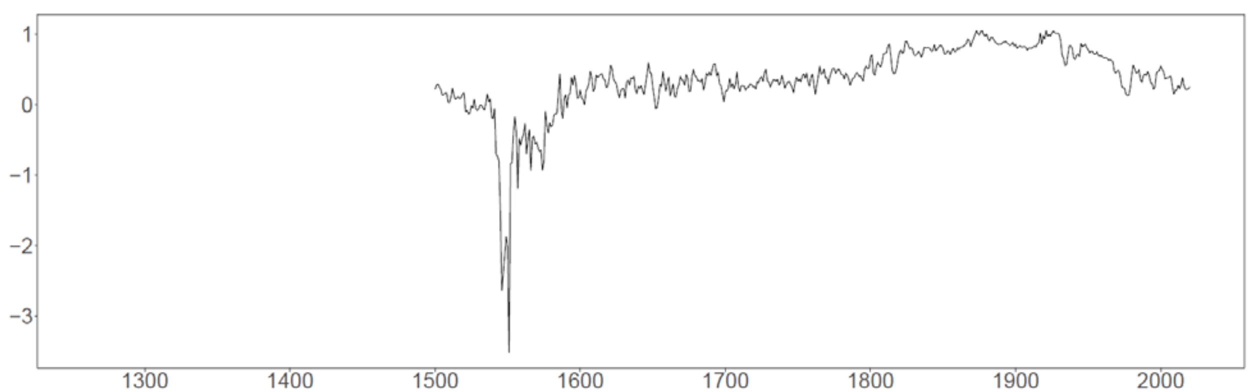
Germany



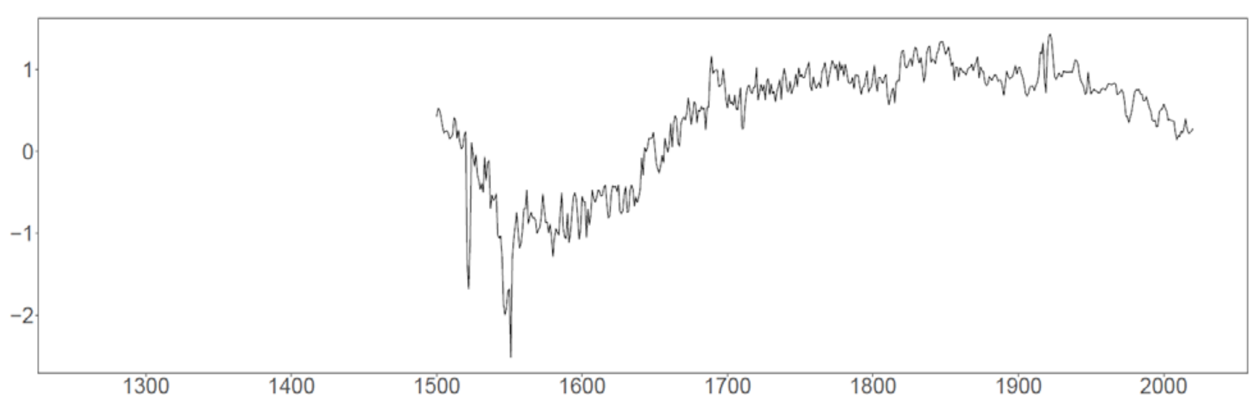
Italy



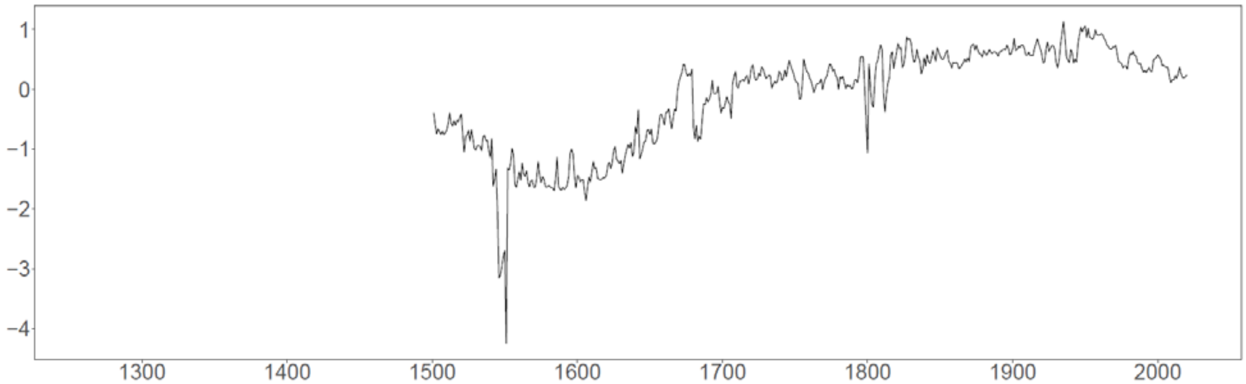
Netherlands



Portugal



Spain



Sweden

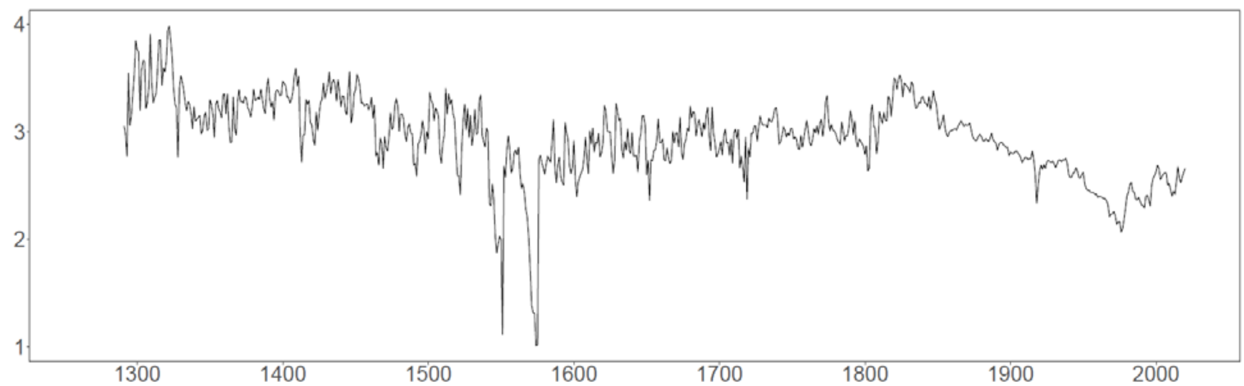
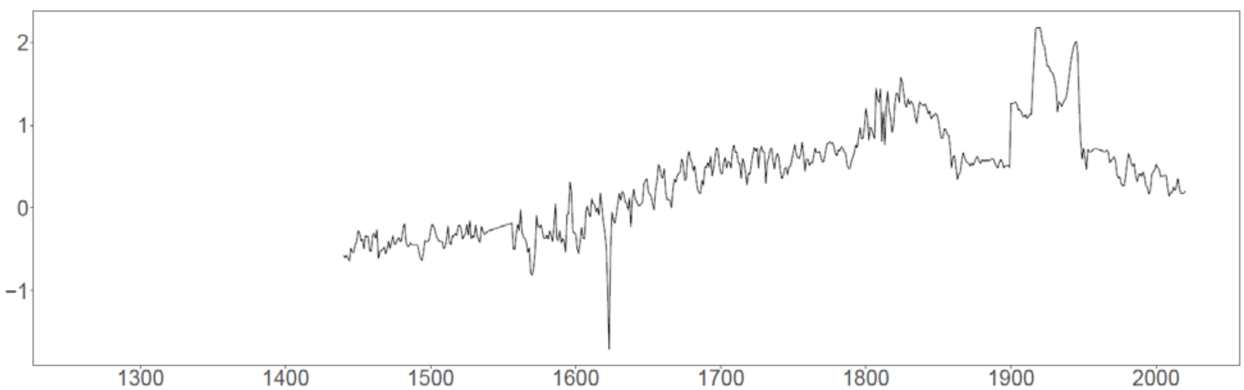


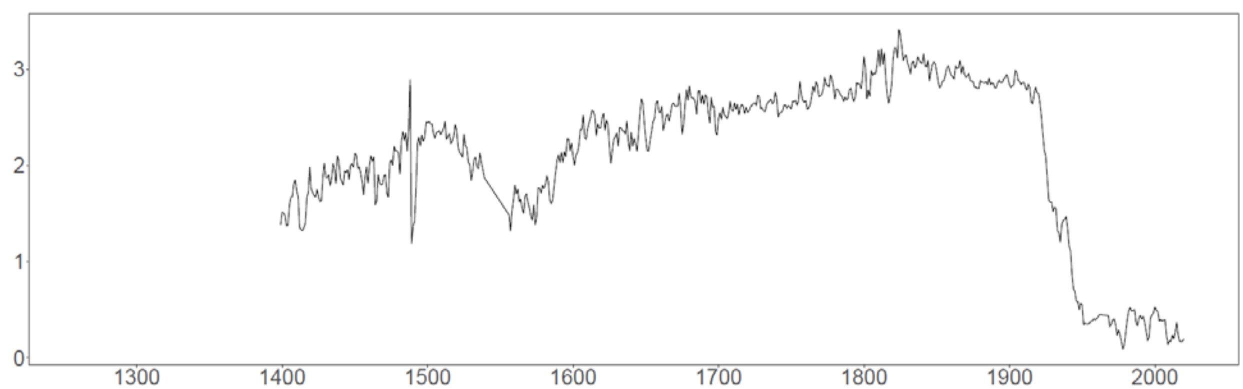
Figure 2: (Log) real exchange rates with Great Debasement adjustment (1264 – 2020)

Notes: Figure 2 contains plots of the log real exchange rate (q_t) for the ultra-long series and with adjustment for the Great Debasement. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

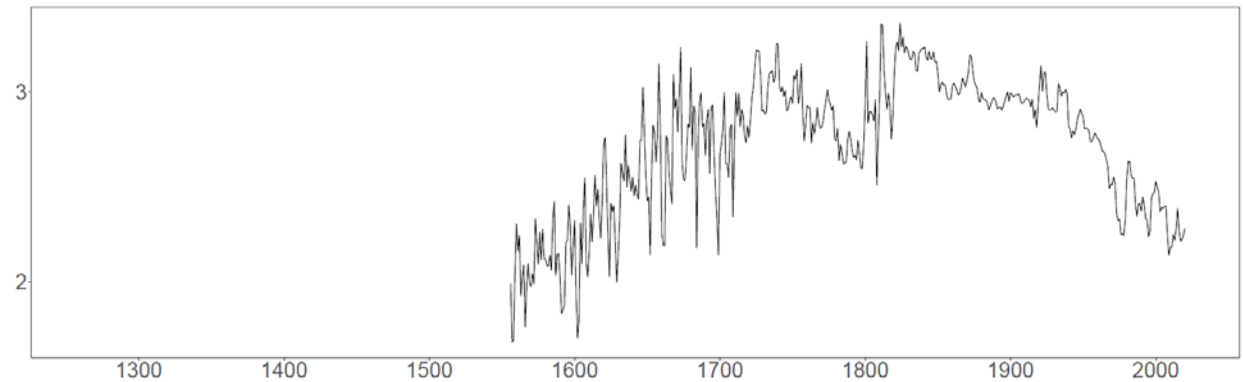
Austria



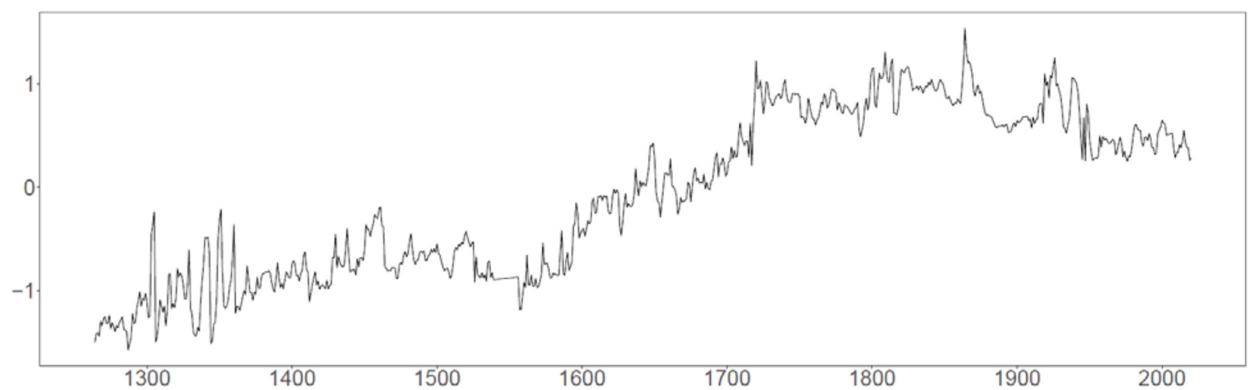
Belgium



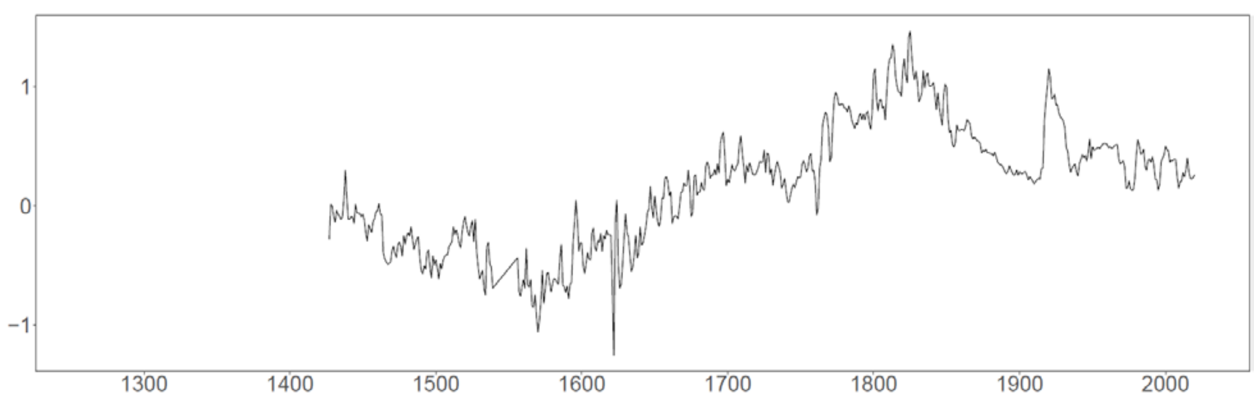
Denmark



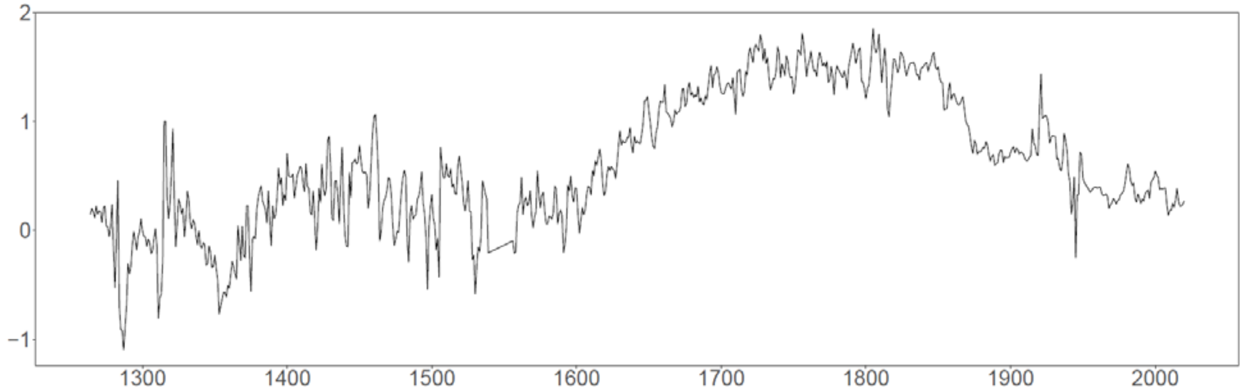
France



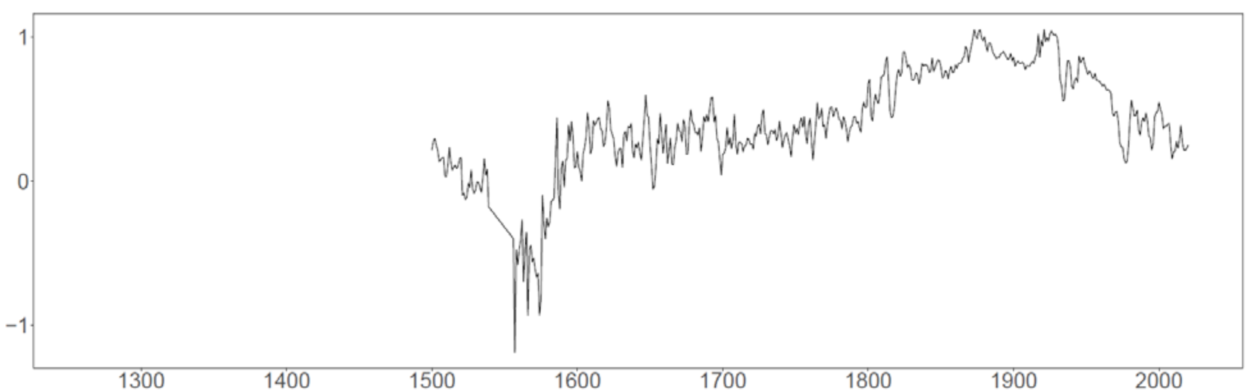
Germany



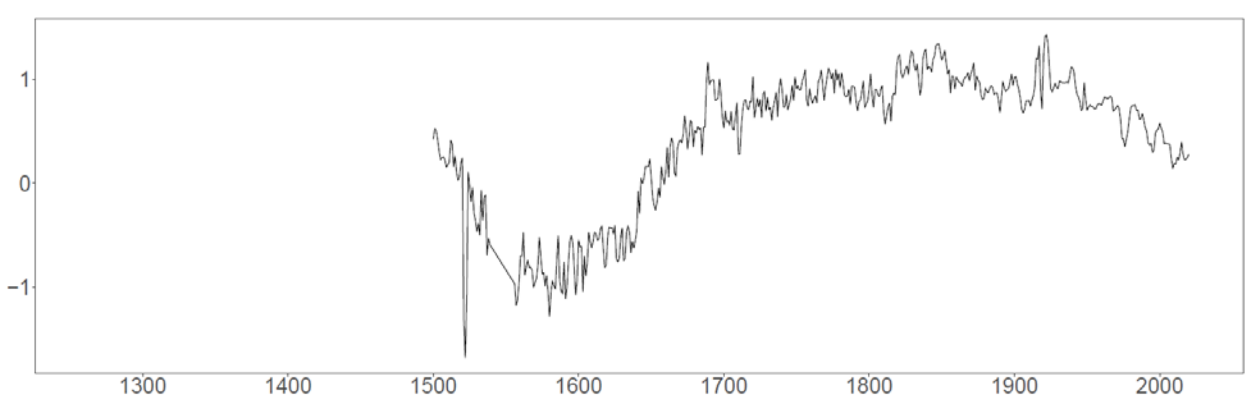
Italy



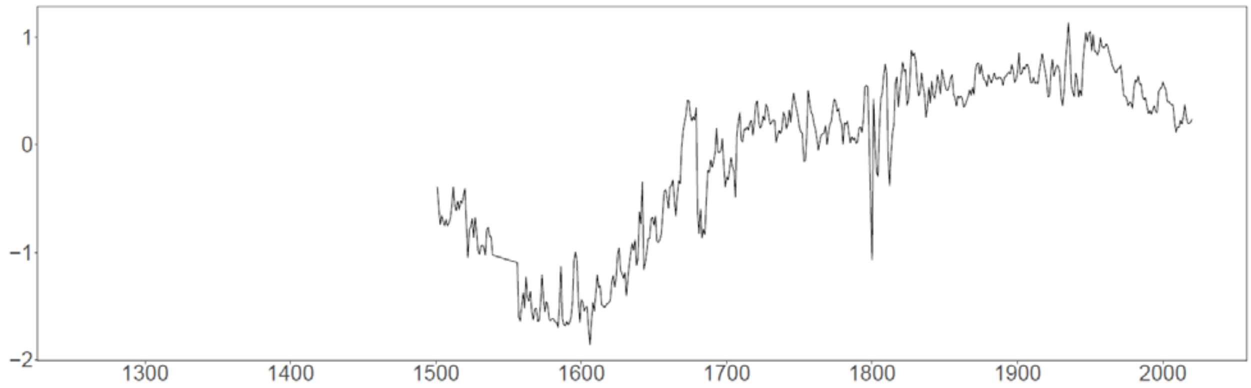
Netherlands



Portugal



Spain



Sweden

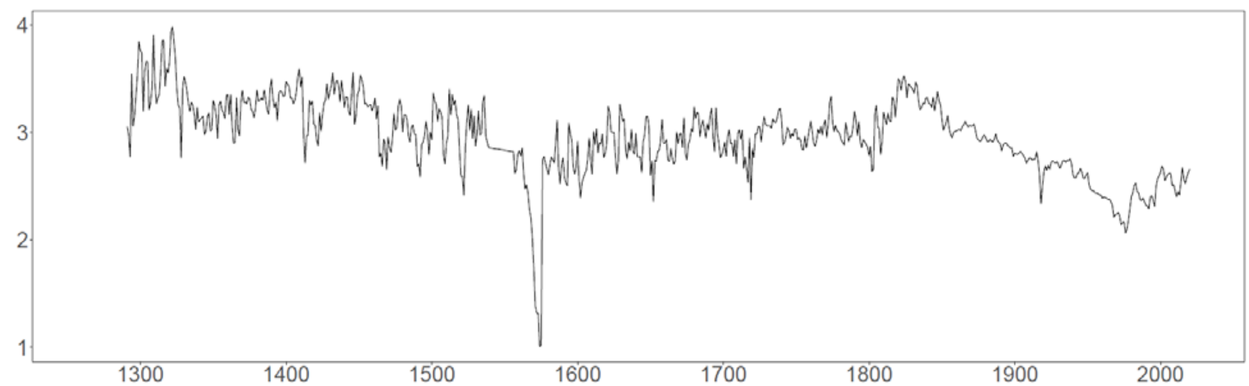
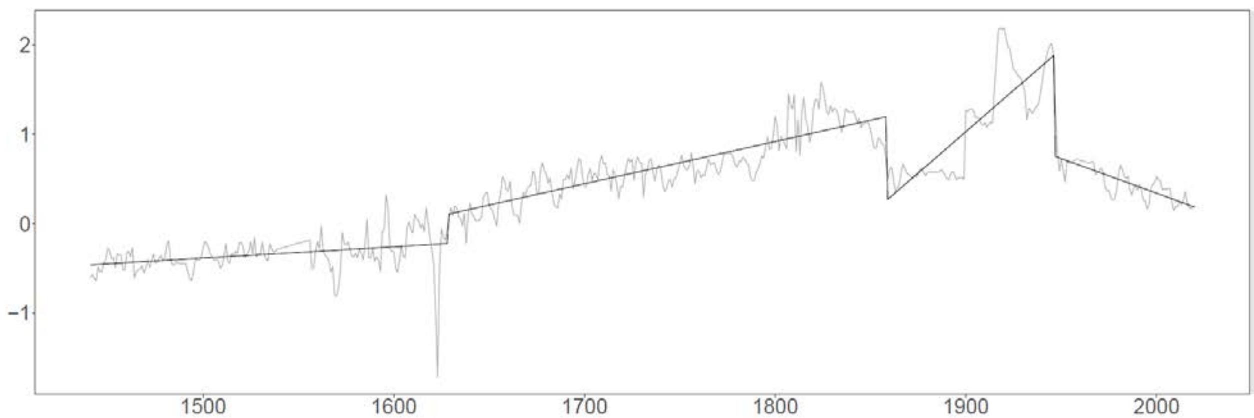


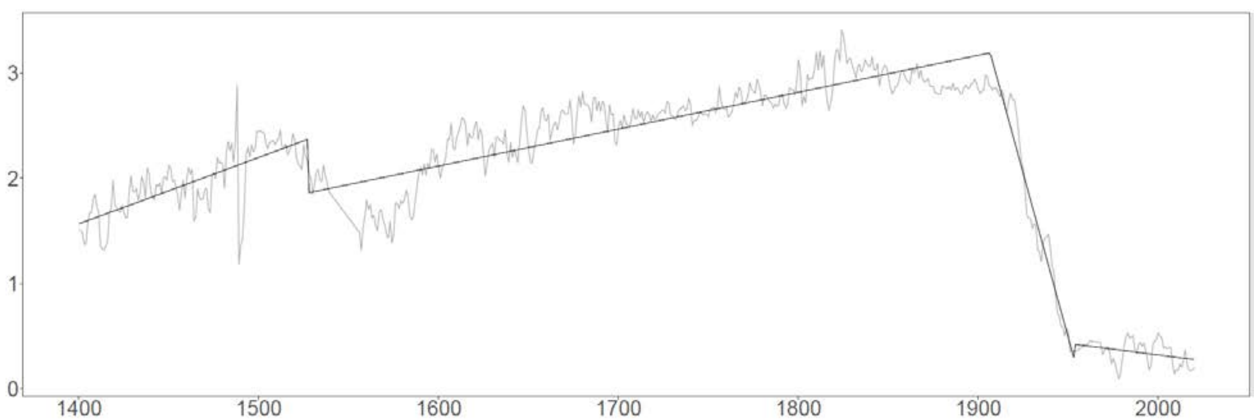
Figure 3: (Log) real exchange rates for ultra-long series (Great Debasement adjusted) with fitted broken trend

Notes: Figure 3 contains plots of the log real exchange rate (q_t) for the ultra-long series, with adjustment for the Great Debasement and with the fitted broken trend. Table 5 presents the fitted values of the parameters of (10) at the minimum SSR dates (i.e., the estimated break dates) and consequently Figure 3, provides a corresponding graphical representation. In terms of sample period, cases stretch back till 1264 (France and Italy) and run annually and continuously until 2020. These include Sweden (commencing 1291), Belgium (1399), Germany (1427), Austria (1440), Netherlands and Portugal (1500), Spain (1501) and Denmark (1556).

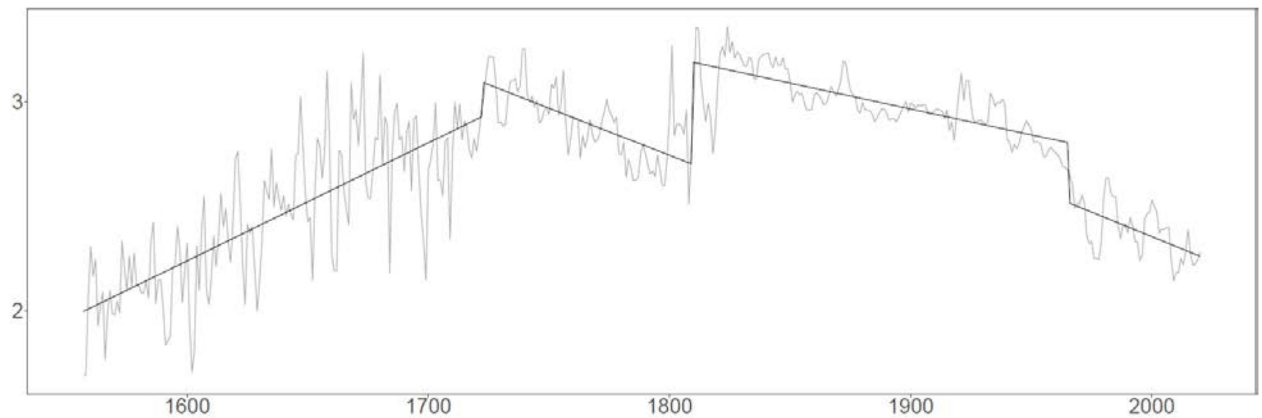
Austria



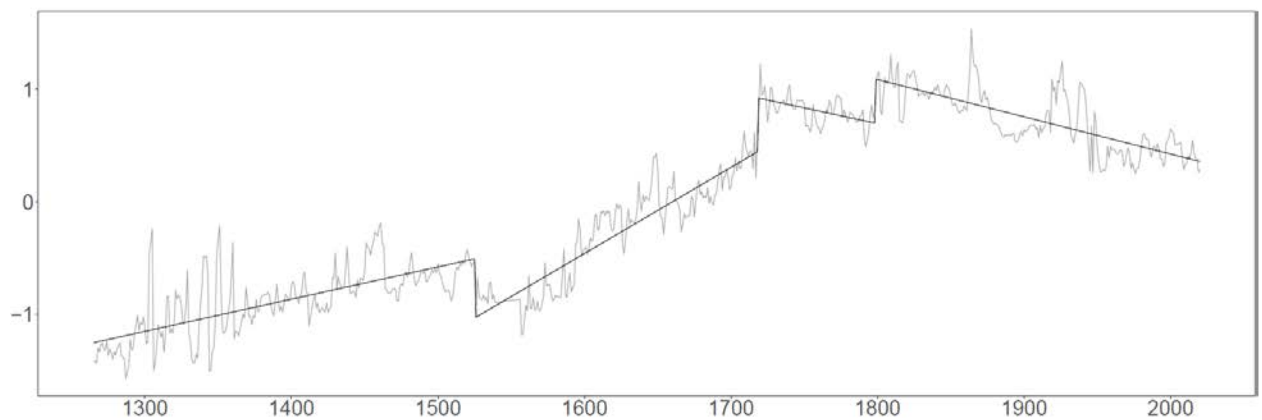
Belgium



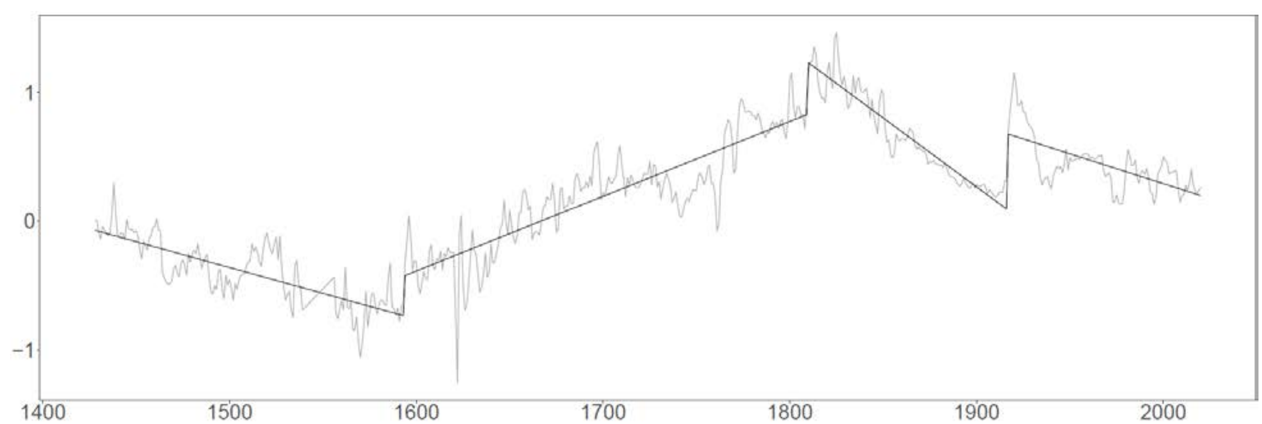
Denmark



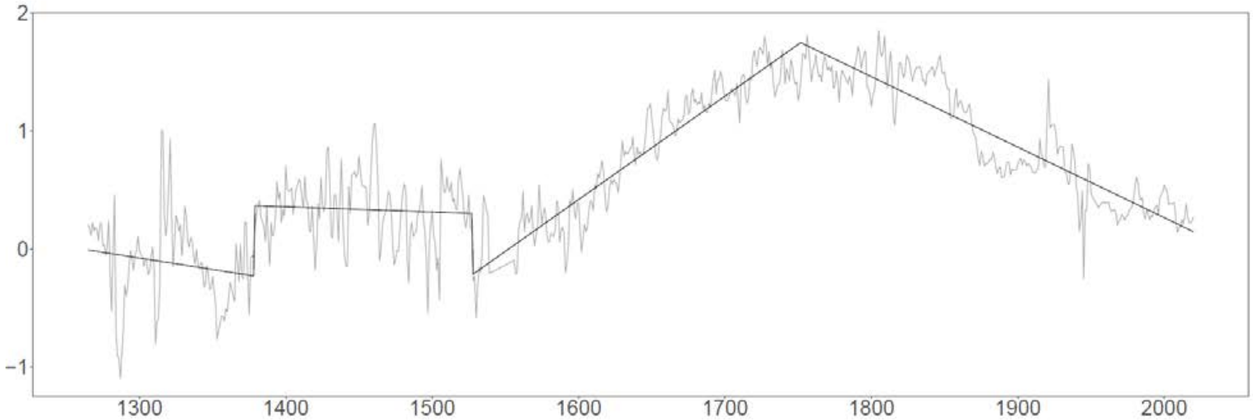
France



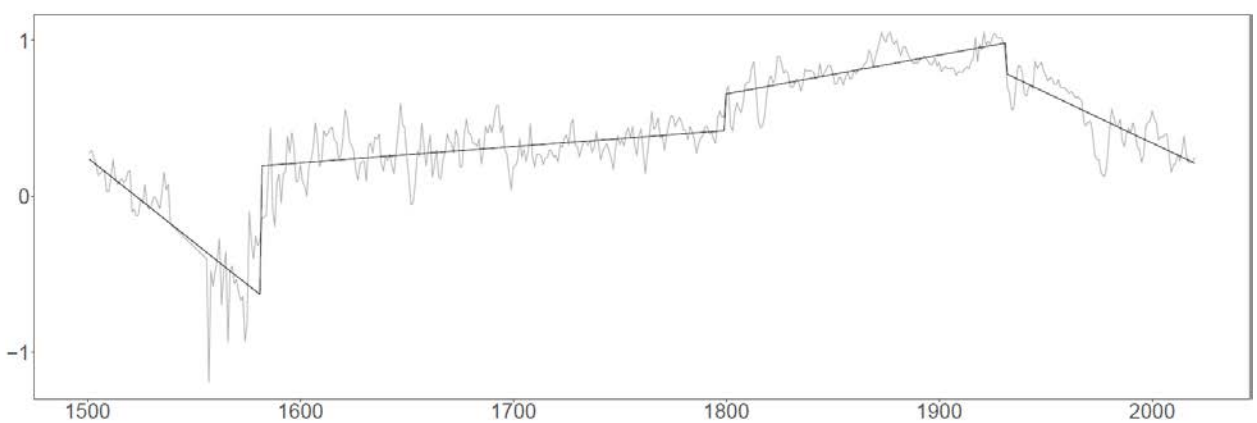
Germany



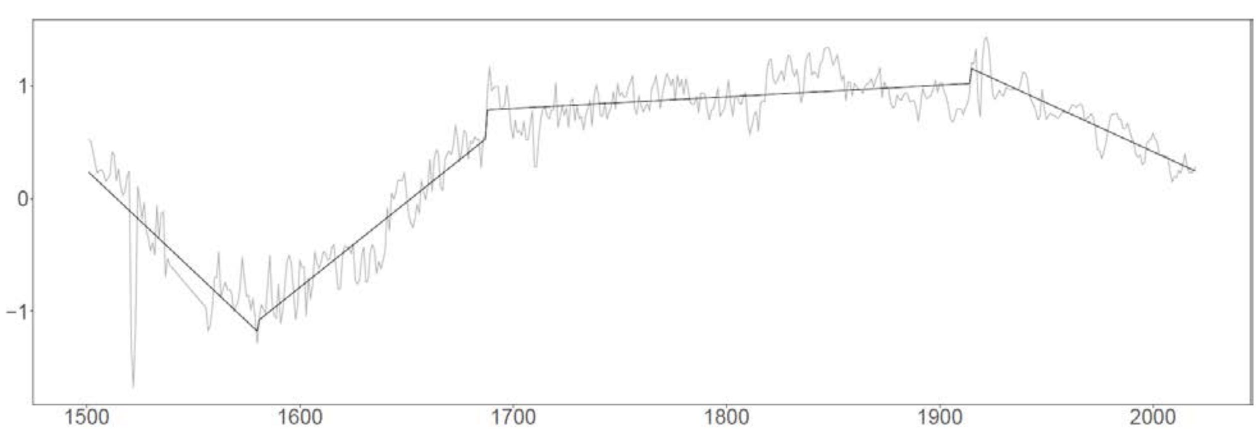
Italy



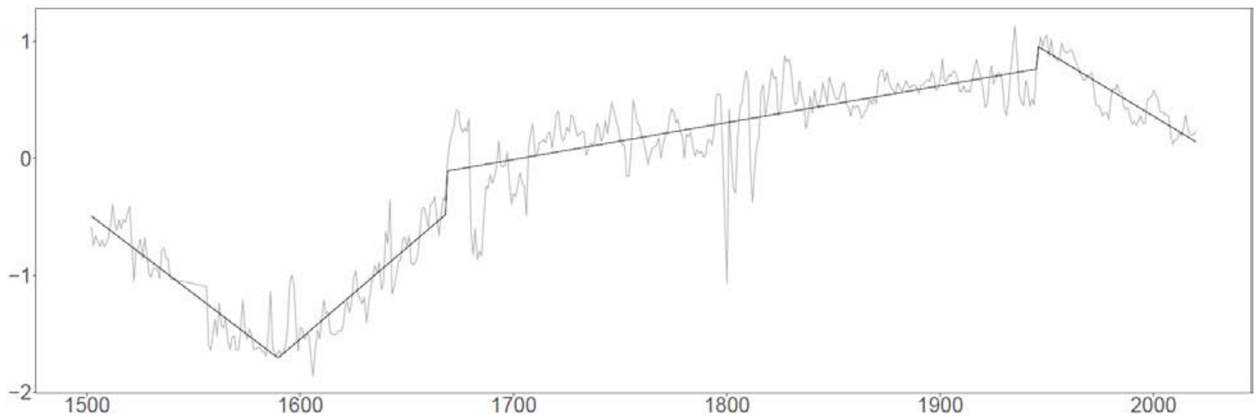
Netherlands



Portugal



Spain



Sweden

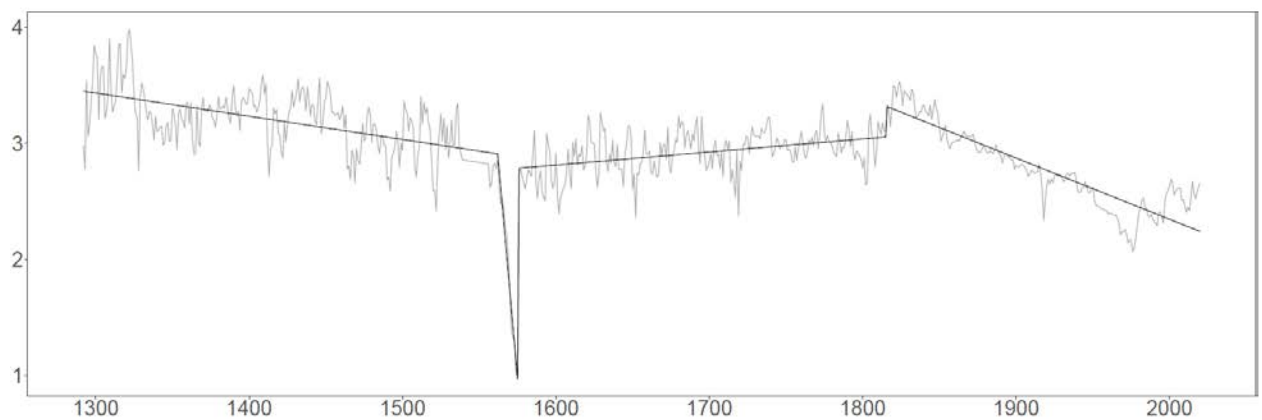
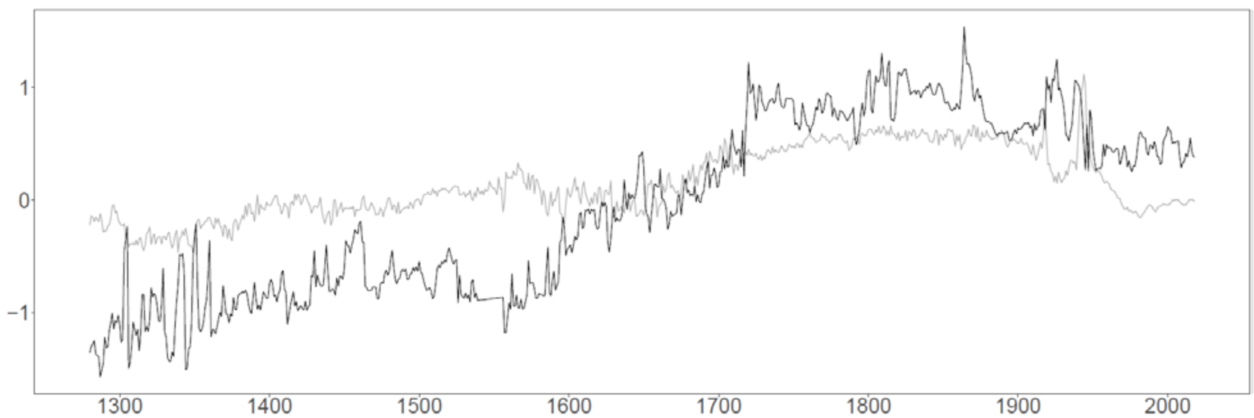


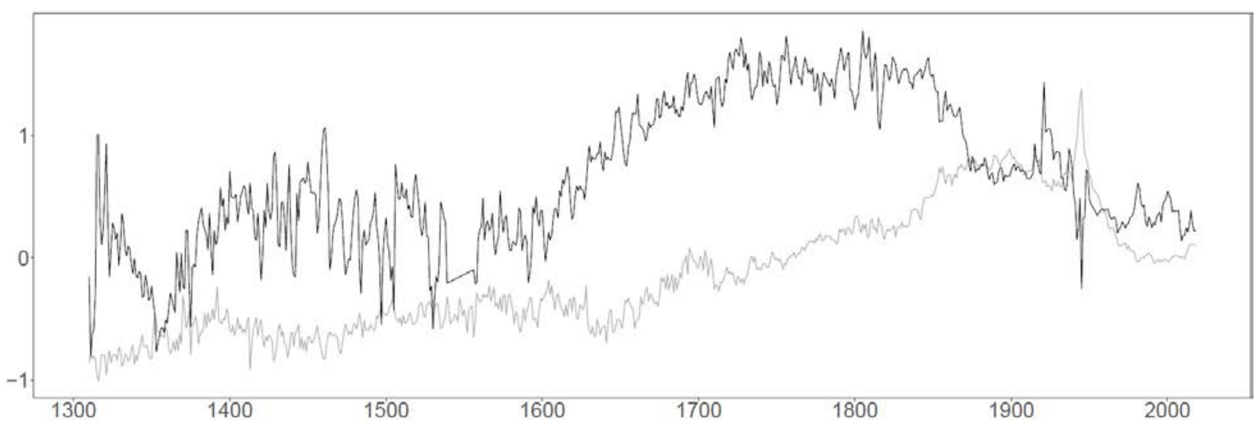
Figure 4: (Log) real exchange rates for ultra-long series (Great Debasement adjusted) and country differentials in real (log) GDP per capita

Notes: Figure 4 contains plots of the log real exchange rate (q_t) for the ultra-long series, with adjustment for the Great Debasement, alongside the quantity $(l_t - l_t^*)$ from (9) proxied by $(y_t - y_t^*)$, where y_t represents log real GDP per capita. In terms of sample period, cases stretch back till 1280 (France) and run annually and continuously until 2020. These include Sweden (commencing 1300), Italy (1310), and Netherlands and Portugal (1500).

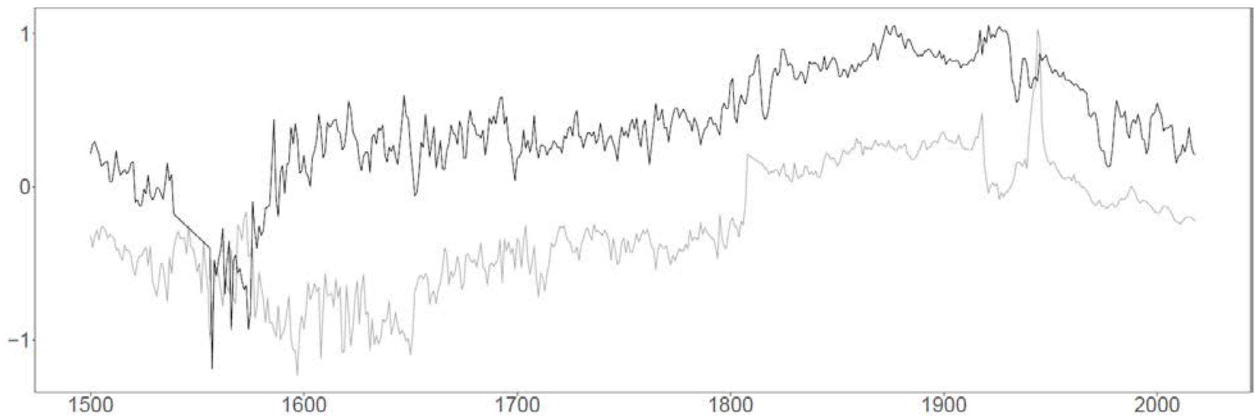
France (from 1280)



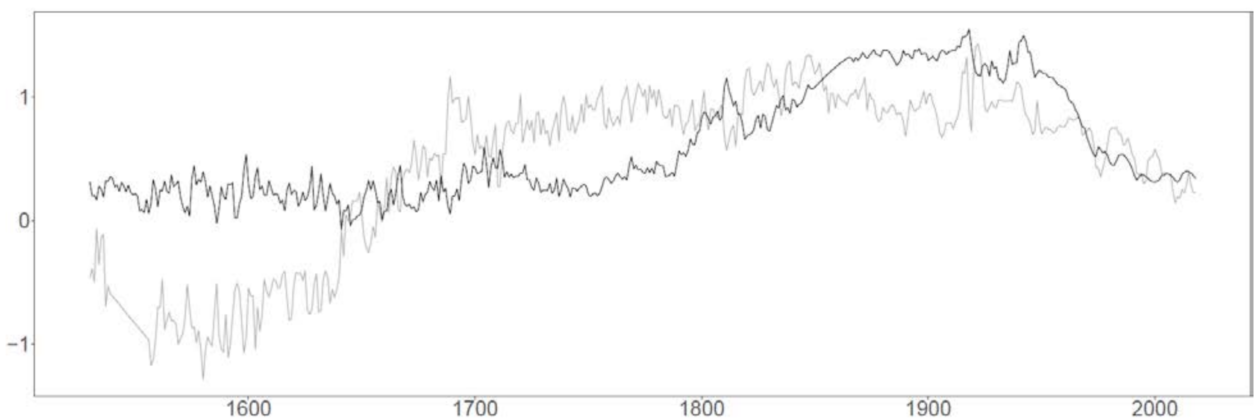
Italy (from 1310)



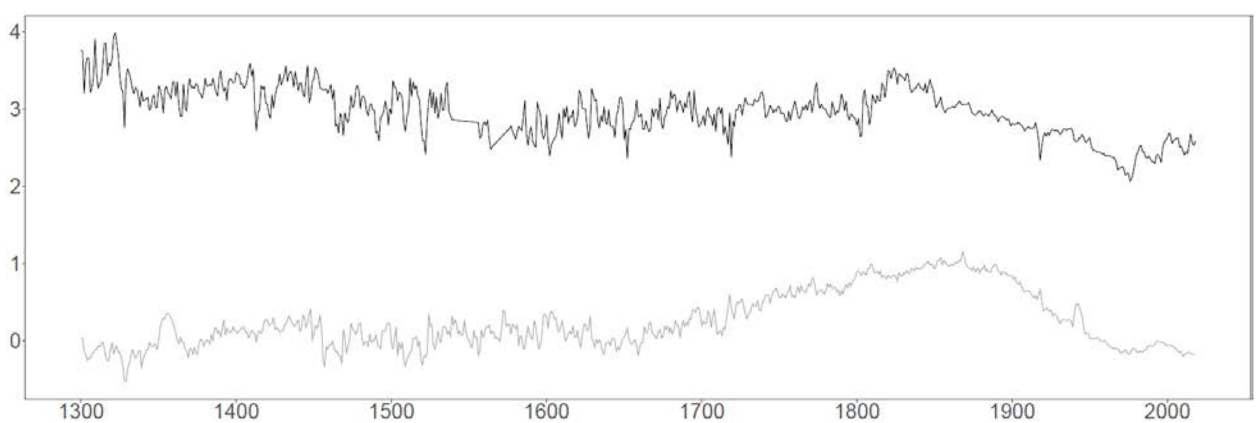
Netherlands (from 1500)



Portugal (from 1500)



Sweden (from 1300)



Appendix – Data Sources

EXCHANGE RATES

Vis-à-vis GBP

Austria
1440-1776. Allen, R.C. (2001), Great divergence in European wages and prices from the Middle Ages to the First World War. *Explorations in Economic History*, 38, 411-447. Data available from Bob Allen Research Pages, Nuffield College, University of Oxford, Oxford, originally viewed 20 July 2014, www.nuffield.ox.ac.uk/people/sites/allen-research-pages/; **1776-1862**. Denzel, Markus A. (2010), *Handbook of World Exchange Rates, 1590–1914*. Routledge, pp. 26-28, 261-62; **1863-1912**. Bank of Greece, Bulgarian National Bank, National Bank of Romania, Oesterreichische Nationalbank (2014), *South-Eastern European Monetary and Economic Statistics from the Nineteenth Century to World War II*, Athens, Sofia, Bucharest, Vienna, available from: https://www.bankofgreece.gr/en/publications-and-research/research/research-networks/seemhn/the-seemhn-data-volume and South-Eastern European Monetary and Economic Statistics from the Nineteenth Century to World War II - Oesterreichische Nationalbank (OeNB), viewed originally viewed 9 February 2015; **1913**. Denzel, Markus A. (2010), *Handbook of World Exchange Rates, 1590–1914*. Routledge; **1914-1946**. Statistical Yearbook Germany (various issues), and League of Nations, (various issues), Monthly Statistical Bulletin, Geneva: League of Nations. **1947-2020**. IMF, International Financial Statistics.

Belgium
1399-1831. Korthals Altes, W.L. (1996), Van Pond Hollands tot Nederlandse Gulden, NEHA, Amsterdam; and Allen, R.C. (2001), Great divergence in European wages and prices from the Middle Ages to the First World War. *Explorations in Economic History*, 38, 411-447. Data available from Bob Allen Research Pages, Nuffield College, University of Oxford, Oxford, originally viewed 20 July 2014, www.nuffield.ox.ac.uk/people/sites/allen-research-pages/; **1831-1913**. Denzel, Markus A. (2010) *Handbook of World Exchange Rates*, *1590–1914*. Routledge, pp. 20-23, Table 1.1.1. **1914- 1919**. League of Nations (various issues), Monthly Statistical Bulletin, Geneva: League of Nations. **1920-1974**. Banque Nationale de Belgique (1994) *Statistiques Economiques Belges 1980-1990, Tableaux*, Brussels, pp. 10-11, 146. **1974-2020**. IMF, International Financial Statistics

Denmark
1556-1817. The exchange rate is computed as the Swedish-Danish exchange rate converted to GBP using the Swedish exchange rate from Rodney Edvinsson (2010), Foreign exchange rates in Sweden 1658–1803. In Rodney Edvinsson, Tor Jacobson and Daniel Waldenström (eds), *Historical Monetary and Financial Statistics for Sweden, Volume I: Exchange Rates, Prices, and Wages, 1277–2008* pp. 238-290. Sveriges Riksbank and Ekerlids Forlag AB; and Håkan Lobell

(2010), Foreign exchange rates 1804–1914. In Rodney Edvinsson, Tor Jacobson and Daniel Waldenström (eds), *Historical Monetary and Financial Statistics for Sweden, Volume I: Exchange Rates, Prices, and Wages, 1277–2008*, pp. 291-339. Sveriges Riksbank and Ekerlids Forlag AB; **1816-1980**. Kim Abilgdren (2013) Stress scenarios from the tails of historical distributions of macro-financial risk factors in Denmark Danmarks Nationalbank Working Papers No. 86, November 2013; **1981-2020**. Danmarks Nationalbank Statistikbank (ND) Copenhagen, last viewed 27 March 2023. http://nationalbanken.statistikbank.dk/statbank5a/SelectTable/Omrade0.asp?PLanguage=1.

France

1259-1619. Alen, R.C. (2001), Great divergence in European wages and prices from the Middle Ages to the First World War. *Explorations in Economic History*, vol. 38, pp. 411-447. Data available from Bob Allen Research Pages, Nuffield College, University of Oxford, Oxford, www.nuffield.ox.ac.uk/people/sites/allen-research-pages/; **1620-1934**. B.R. Mitchell (1988), *British Historical Statistics*, Cambridge University Press, Cambridge, UK, pp. 700-703, Table 22; **1935-1946**. League of Nations, (various issues), Monthly Statistical Bulletin, Geneva: League of Nations; **1947-2020**. IMF, *International Financial Statistics*, various issues.

Germany

1427-1619. Alen, R.C. (2001), Great divergence in European wages and prices from the Middle Ages to the First World War. *Explorations in Economic History*, vol. 38, pp. 411-447. Data available from Bob Allen Research Pages, Nuffield College, University of Oxford, Oxford, www.nuffield.ox.ac.uk/people/sites/allen-research-pages/; **1620-1934**. B.R. Mitchell (1988), *British Historical Statistics*, Cambridge University Press, Cambridge, UK, pp. 700-703, Table 22; **1935-1946**. League of Nations, (various issues), Monthly Statistical Bulletin, Geneva: League of Nations; **1947-2020**. IMF, *International Financial Statistics*, various issues.

Italy

1252-1914. Average of the city states before unification in 1861 and Italy, thereafter, from Malanima, Paolo (2002), L'economia italiana: dalla crescita medievale alla crescita contemporanea. Bologna: Società editrice il Mulino, and Allen, R.C. (2001), Great divergence in European wages and prices from the Middle Ages to the First World War, *Explorations in Economic History*, 38, 411-447. Data available from Bob Allen Research Pages, Nuffield College, University of Oxford, Oxford, originally viewed 20 July 2014, www.nuffield.ox.ac.uk/people/sites/allen-research-pages/; **1914-1929.** Mood, James R. (1930), *Handbook of Foreign Currency and Exchange*. US Department of Commerce; **1930-1949**. K. Anderson and V. Pinilla (2017), Annual Database of Global Wine Markets, 1835 to 2018, Wine Economics Research Centre, University of Adelaide, August (available in Excel files at www.adelaide.edu.au/wine-econ/databases), updated January 2020. **1950-2020**. IMF, *International Financial Statistics*, various issues.

Netherlands

1500-1745. Metz, R., 1990: Geld, Währung und Preisentwicklung. Der Niederrheinraum im europäischen Vergleich: 1350 – 1800. Frankfurt/Main. Fritz Knapp Verlag; **1746-1913**. Markus A. Denzel (2010) *Handbook of World Exchange Rates, 1590-1914*, Ashgate, Surrey, England, pp. 533-534; **1913-1949**. K. Anderson and V. Pinilla (2017), Annual Database of Global Wine

Markets, 1835 to 2018, Wine Economics Research Centre, University of Adelaide, August (available in Excel files at www.adelaide.edu.au/wine-econ/databases), updated January 2020; **1950-2020**. IMF, *International Financial Statistics*, various issues.

Portugal

1500-1775. Costa, L., N. Palma and J. Reis (2015), The great escape? The contribution of the empire to Portugal's economic growth, 1500-1800. *European Review of Economic History*, 19(1), 1-22. See, for data, http://pwr-portugal.ics.ul.pt/?page_id=56. **1776-1890**. Markus A. Denzel (2010) *Handbook of World Exchange Rates, 1590-1914*, Ashgate, Surrey, England, pp. 32-36, Table 1.1.3; **1891-1949**. Nuno Valerio (ed.) (2001), *Portuguese Historical Statistics*, 2 vols, Instituto Nacional de Estatistica, Lisbon, p 745; **1950-2020**. IMF, *International Financial Statistics*, various issues.

Spain

1501-1791. Alen, R.C. (2001), Great divergence in European wages and prices from the Middle Ages to the First World War. *Explorations in Economic History*, vol. 38, pp. 411-447. Data available from Bob Allen Research Pages, Nuffield College, University of Oxford, Oxford, www.nuffield.ox.ac.uk/people/sites/allen-research-pages/; **1792-1820**. Markus A. Denzel (2010) *Handbook of world exchange rates, 1590-1914,* Ashgate, Surrey, England, pp. 30-33, Table 1.1.3. **1821-1950**. Albert Carreras and Xavier Tafunell (eds.) *Estadisticas historicas de Espana Siglos XIX-XX*, pp. 702-705, Table 9.19. **1950-2020.** IMF, *International Financial Statistics*, various issues.

Sweden

1291-1547. Rodney Edvinsson, Bo Franzén and Johan Söderberg (2010), Swedish Payment Systems 995–1534. In Rodney Edvinsson, Tor Jacobson and Daniel Waldenström (eds), *Historical Monetary and Financial Statistics for Sweden, Volume I: Exchange Rates, Prices, and Wages, 1277–2008* pp. 67-132. Sveriges Riksbank and Ekerlids Forlag AB, Table A3.1; **1548-1803**. Rodney Edvinsson (2010), Foreign exchange rates in Sweden 1658-1803. In Rodney Edvinsson, Tor Jacobson and Daniel Waldenström (eds), *Historical Monetary and Financial Statistics for Sweden, Volume I: Exchange Rates, Prices, and Wages, 1277–2008*, pp. 238-290. Sveriges Riksbank and Ekerlids Forlag AB; **1804-1912**. Hakan Lobell (2010) Foreign exchange rates in Sweden 1804-1914. In Rodney Edvinsson, Tor Jacobson and Daniel Waldenström (eds), *Historical Monetary and Financial Statistics for Sweden, Volume I: Exchange Rates, Prices, and Wages, 1277–2008*, pp. 291-339. Sveriges Riksbank and Ekerlids Forlag AB; **1913-1992**. Jan Bohlin (2010), From appreciation to depreciation – the exchange rate of the Swedish krona, 1913-2008. In Rodney Edvinsson, Tor Jacobson and Daniel Waldenström (eds), *Historical Monetary and Financial Statistics for Sweden, Volume I: Exchange Rates, Prices, and Wages, 1277–2008*, pp. 340-411. Sveriges Riksbank and Ekerlids Forlag AB; **1993-2020**. IMF, *International Financial Statistics*, various issues.

UK

Not needed since all the currencies are against GBP.

CONSUMER PRICES

Austria

1440-1800. (Vienna) Allen, R.C. (2001), Great divergence in European wages and prices from the Middle Ages to the First World War', *Explorations in Economic History*, vol. 38, pp. 411-447. Data available from Bob Allen Research Pages, Nuffield College, University of Oxford, Oxford, originally viewed 27 July 2014, www.nuffield.ox.ac.uk/people/sites/allen-research-pages/; **1800-1967.** Statistik Austria (1979), *Geschichte und Ergebnisse der zentralen amtlichen Statistik in Osterreich 1829-1979, Beitrage zur Österreichischen Statistik,* Vol.1, Heft 550, Vienna; **1967-2020**. IMF, *International Financial Statistics*, various issues.

Belgium

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