Globalization, long memory, and real interest rate convergence: a historical perspective

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Received: 26 January 2021 / Accepted: 8 January 2022 / Published online: 10 February 2022
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Abstract

This paper investigates whether the real interest rate parity (RIRP) is valid during the three waves of globalizations that occurred in the last 150 years (1870–1914, 1944–1971, 1989 to the present). If any, these periods should favor RIRP, since globalization is a process where economies and financial markets become increasingly integrated into a global economic system. In contrast to the existing literature, we model the departures from RIRP as a long-term memory process and apply fractional integration methods on a sample of real interest rate differentials of seven developed countries: France, Germany, Holland, Italy, Japan, Spain, and the UK across the three globalization waves paired against the USA. We compute impulse response functions (IRF) to gain further insight into the memory characteristics of the RIRP differential processes and provide half-life estimates. We find that deviations from RIRP are mean reverting, providing robust evidence of real interest rate convergence during the three globalization waves. We shed further light on financial and commodity market integration during the three globalization waves by assessing the memory properties of uncovered interest rate parity (UIP) and relative purchasing power parity (PPP) differential processes. We find that deviations from relative PPP and UIP are not always mean-reverting processes. RIRP, relative PPP, and UIP hold simultaneously only in 7 out of 21 cases; RIRP and UIP hold in 11 out of 21 cases; RIRP hold without the support of relative PPP and UIP in 3 out of 21 cases. Thus, the evidence in favor of real interest rate convergence appears to be driven more by UIP than relative PPP. All these results are, to the authors knowledge, new to the literature.

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Keywords Globalization · Fractional integration · Real interest rate parity · Purchasing power parity · Uncovered interest rate parity · Half-life · Impulse response function

JEL Classification C22 · E43 · G15 · N20

1 Introduction

Real interest rate parity (RIRP) is one of the central tenets in international macroeconomics. Long-run convergence of real interest rates is a key assumption of the monetary approach to exchange rate determination (Mussa 1976; Frankel 1979) and represents one of the most relied upon indicators of financial globalization (Levich 2013) and of the degree of economic integration across countries (Phylaktis 1999; Obstfeld and Taylor 2003). On a policy level, RIRP has profound implications for the viability of an independent national monetary policy (Mark 1955; Chortareas et al. 2018).

Empirical studies utilizing the $I(0)/I(1)$ dichotomous approach often find the parity to not hold or to give rise to mixed results. See, for instance, Wu and Chen (1998), Awad and Goodwin (1998), Wu and Fountas (2000), Nakagawa (2002), Obstfeld and Taylor (2003), Holmes and Maghrebi (2004), (Ferreira and León-Ledesma (2007), Arghyrou et al. (2009), Dreger (2010), Cuestas and Harrison (2010), Güney and Hasanov (2014), Çorakcı et al. (2017), and Bahmani-Oskooee et al. (2019), among many others.

This paper expands the RIRP literature in two ways. First, in contrast to the $I(0)/I(1)$ prevailing literature, we model departures from RIRP as a long-term memory process and apply fractional integration methods on a sample of real interest rate differentials of seven countries: France, Germany, Holland, Italy, Japan, and the UK across the three globalization waves paired against the USA. In other words, we examine the parity by using the deviations from RIRP. Under this approach, unlike the conventional $I(0)/I(1)$ approach, reversion to parity and convergence of real interest rates can occur even if departures from RIRP may not be $I(0)$. That is, RIRP is valid if the real interest rate differentials display a mean reversion process. We complement the fractional integration results by computing the corresponding impulse response functions (IR) and half-lives (HL). Finally, we shed further light on financial market integration during the three globalization waves by assessing the memory properties of uncovered interest rate parity (UIP) and relative PPP differential processes, which we argue represents the economic fundamentals of RIRP (Marston 1995; MacDonald and Marsh 1999; Dreger 2010).¹

Second, in contrast to the previous literature, we investigate whether RIRP is valid during the three waves of globalization that occurred in the last 150 years (1870–1914, 1944–1971, 1989 to the present). If any, these periods should favor RIRP, since globalization is a process where economies and financial markets become increasingly

¹ Any work dealing with historical data must confront the problem of structural change. As emphasized by Dreger (2010), however, focusing on historical episodes, the structural break argument is less relevant.

integrated into a global economic system (Rodrik 1998; Obstfeld and Taylor 2003; Bordo et al. 2003).²

The rest of the paper is organized as follows. Section 2 offers a brief perspective on the economic factors and trends that have uniquely characterized each of the three globalization waves. Section 3 reviews the relevant theory and Sect. 4 briefly outlines the statistical model. Section 5 displays the test results of RIRP. Section 6 presents the tests of relative PPP and UIP. Section 7 offers concluding remarks. In Appendix A, we present the test results applied to real interest rates.

2 Three globalization waves

The last 150 years of economic history witnessed severe economic and financial crises such as the Long Depression (1873–1896), the German hyperinflation (1919–1923), the Great Depression (1929–1939) and, more recently, the Great Recession (2007–2009) and the COVID-19 recession (2020). These episodes stand out as events that transformed the world capital markets and left interest arbitrage differentials higher and more volatile than ever before (Obstfeld and Taylor 2003; Levich 2013; Rebucci et al. 2021).

The last 150 years also witnessed three waves of globalization of trade and finance (Piketty 2014; Palley 2018). Globalization, a complex phenomenon, reflects the interaction of many technological, cultural, economic, social, and environmental trends. Given this complexity, any attempt to give a satisfactory definition of globalization is doomed to fail. Rather, we identify the process of globalization as one where the economies and the financial markets become increasingly integrated into a global economic system (Rodrik 1998; Obstfeld and Taylor 2003; Bordo et al. 2003).

The first wave of globalization, called the first “golden age” of globalization (Zinkina et al. 2019), took place between 1870 and 1914 (Piketty 2014; Palley 2018), coming to an end with World War I. During this period, capital moved freely between countries and trade accelerated significantly (O’Rourke et al. 1994). The gold standard dominated this period, and a truly global market for capital emerged, to which the spread of the gold standard greatly contributed (Zinkina et al. 2019). The volume of global exports increased by nearly two orders of magnitude during 1800–1913 and a major part of this increase occurred during the “golden age” of globalization (Zinkina et al. 2019). In some respects, financial integration was more pronounced than it is today. International migration was certainly greater than it is today with roughly over 40 million people leaving Europe to seek their fortunes in the New World (Easterlin 1961; Baines et al. 1994; Hatton et al. 1994). In 1870, world trade as a share of GDP was 17.7 percent, rising to 29.1% in 1913 (Ortiz-Ospina and Roser 2017). Important drivers behind this wave of globalization were the new technologies of the era—steam engine, internal combustion engine, telegraph, electricity—that could bridge long geographical distances and the fact that many countries began to embrace liberal trade

² The previous literature has discussed RIRP in the context of historical episodes. Dreger (2010) and Obstfeld and Taylor (2003) test the validity of RIRP during four distinct monetary regimes, the gold standard (1890–1914), the interwar period (1921–1938), the Bretton Woods system (1950–1973), and the float (1974–2000). Their methodology, however, operates in the $I(0)/I(1)$ framework.

policy after years of protectionism. The UK led the world's economy with the British Empire as the epicenter of colonialism. The basis for the European free-trade system was the 1860 free-trade pact between the UK and France. Many other European countries subsequently aligned themselves with this free-trade system (Collier and Dollar 2002).

Following the outbreak of World War I, the world economy went from a globalized to an almost autarkic world order in the space of a few decades. Global trade collapsed and trended downward throughout the interwar period. The second wave of globalization began in 1944 and ended in 1971 (Piketty 2014; Palley 2018). In 1946, world trade as a share of GDP was 15.1%. In 1972, it stood at 25.2% (Ortiz-Ospina and Roser 2017). International regulations and organizations to support economic integration at the global level were created after World War II. During this period, the World Bank (IBRD) and the International Monetary Fund (IMF) were created to facilitate the smooth operation of the international exchange of goods, services, and assets. In addition, the General Agreement on Tariffs and Trade (GATT) began operation in 1948. GATT set the framework for several important steps toward increased global free trade, particularly via successive reductions in industrial tariffs. Cooperation among countries was based on the Bretton Woods Agreement of 1944. The USA became the leading economy in the world and the dollar became the monetary basis, or key currency, of the financial system.

From a political perspective, this era is dominated by the Cold War. The Bretton Woods system meant that nations fixed their currency exchange rate relative to the US dollar, which, in turn, fixed the dollar's parity to gold. In an important aspect, the post-World War II international economy was less open than the period prior to World War I. Before World War I, the international flow of capital had been free. The Bretton Woods system, in contrast, was based on governmental control of the international flow of capital. Important drivers of the second wave were the new technologies of the era (jet planes, television, communication satellites, container shipping) (Collier and Dollar 2002).

The third wave of globalization began in 1989 and continues to the present (Piketty 2014; Palley 2018). Total trade stood at 59.4% of global GDP in 2018 as compared to 27.3% in 1970, while total exports stood at 30.1% of global GDP as compared to 13.6% in 1970 (World Bank, <https://data.worldbank.org/indicator>). Foreign direct investment increased twice as fast as trade. Microprocessors, personal computers, the Internet, and mobile phones are the technological drivers of the third wave. The USA lost its leading role in the world economy, which has become multi-polar, triangularized by the USA, the European Union, and China (Collier and Dollar 2002).

3 Relevant theory

The (ex-ante) RIRP hypothesis emerges by assuming the validity of UIP, (ex-ante) relative PPP, and the Fisher conditions. UIP proposes that the nominal interest rate differential equals the expected rate of depreciation of the exchange rate. That is,

$$i_t - i_t^* = E_t(s_{t+1} - s_t), \quad (1)$$

where i_t and i_t^* denote the domestic and foreign nominal interest rates, respectively, and s_t denotes the logarithm of the nominal exchange rate expressed as units of domestic currency per unit of foreign currency at time t . E_t is the expectation operator conditioned on information available at time t . A superscript $*$ refers to foreign variables.

The ex-ante relative PPP links the expected rate of depreciation of the exchange rate to the expected inflation rate differential:

$$(s_{t+1} - s_t) = E_t(\pi_{t+1} - \pi_{t+1}^*), \quad (2)$$

where $\pi_{t+1} = p_{t+1} - p_t$ represents the domestic inflation rate at time $t + 1$ and p_t equals the logarithm of the domestic price level at time t . The two parity conditions imply:

$$i_t - E_t\pi_{t+1} = i_t^* - E_t\pi_{t+1}^*. \quad (3)$$

If the Fisher equation is valid for both the domestic and foreign country, then Eq. (3) implies the equality of the ex-ante real interest rate across countries

$$E_tr_{t+1} = E_tr_{t+1}^*, \quad (4)$$

where r_{t+1} and r_{t+1}^* denote the domestic and foreign real interest rates at time $t + 1$. Equation (4) gives the ex-ante RIRP condition, or stated differently, the hypothesis that the ex-ante real interest rates equalize across countries. Under the assumption of rational expectations, the difference between the ex-post-real interest rates (i.e., the real interest rate differential) equals a random error term related to the inflation forecast error. That is,

$$\text{RIRD}_{t+1} = r_{t+1} - r_{t+1}^* = \varsigma_{t+1}, \quad (5)$$

where ς_{t+1} represents the inflation expectation forecasting error that is serially uncorrelated with zero mean and RIRD_{t+1} denotes the real interest rate differential at time $t + 1$. Equation (5) provides the basis for the main empirical analysis. Thus, we test the validity of RIRP in the long run by examining whether real interest rate differentials are mean reverting. We work with the ex-post rather than the ex-ante real interest rate, where the latter is defined as the nominal interest rate minus the expected inflation, while the former is the nominal rate minus the actual inflation. This is because the ex-ante real rate is not directly observable because expected inflation is not directly observable. Following the work of Rose (1988), by assuming that inflation expectation errors are stationary, the empirical literature, in general, tests the integration properties of the ex-post real rate as proxy for the ex-ante real rate.

4 Methodology

There exist several methods for estimating and testing the fractional integration (or memory) parameter d , including parametric (i.e., maximum likelihood) and semiparametric methods (e.g., Sowell 1992; Shimotsu and Phillips 2005). In this paper, we use the parametric approach developed by Robinson (1994). This procedure uses the Lagrange multiplier (LM) principle and the Whittle function in the frequency domain (Dahlhaus 1989). Robinson (1994) tests the null hypothesis $H_0 : d = d_0$ for any real value of d_0 in a model given by:

$$y_t = \beta_0 + \beta_1 t + x_t, \quad t = 1, \dots, T, \quad (6)$$

where y_t is the observed series, β_0 and β_1 are the coefficients corresponding, respectively, to the intercept and linear time trend (Gil-Alana and Robinson 1997), and x_t is an $I(d)$ process. The latter process is defined as follows:

$$(1 - L)^d x_t = u_t, \quad t = 1, \dots, T, \quad (7)$$

where u_t is an $I(0)$ process (defined as a covariance stationary process with spectral density function that is positive and finite at all frequencies). This includes for u_t the white-noise case and the ARMA process of the form:

$$\Phi_p(L)u_t = \Theta_q(L)\varepsilon_t, \quad (8)$$

where all the roots of the AR polynomial $\Phi_p(L)$ and the MA polynomial $\Theta_q(L)$ are outside the unit circle and L is the lag operator ($L x_t = x_{t-1}$) (Box and Jenkins 1970).

The unknown fractional integration parameter d is estimated together with β_0 and β_1 . Using Eqs. (7) and (8) together, we can identify x_t as an autoregressive fractionally integrated moving-average process, ARFIMA(p, d, q). The fractional differencing operator $(1 - L)^d$ is defined by the binomial series expansion

$$\begin{aligned} (1 - L)^d &= \sum_{k=0}^{\infty} \frac{\Gamma(k - d)}{\Gamma(k + 1) \Gamma(-d)} L^k \\ &= 1 - dL - \frac{1}{2}d(1 - d)L^2 - \frac{1}{6}d(1 - d)(2 - d)L^3 \dots, \end{aligned} \quad (9)$$

where $\Gamma(\cdot)$ denotes the gamma function. Thus, as k increases, the fractional differencing operator $(1 - L)^d$ provides an infinite-order lag polynomial with declining weights. The interpretation of the fractional differencing operator $(1 - L)^d$ is that x_t exhibits an infinite lag-order dependence. If the invertibility condition ($d > -1$) is satisfied, the above process admits an infinite-order moving-average (MA) representation such that

$$x_t = \sum_{k=0}^{\infty} a_k u_{t-k}, \quad (10)$$

where

$$a_k = \frac{\Gamma(k+d)}{\Gamma(k+1)\Gamma(d)}. \quad (11)$$

The parameter d governs the long-run dynamics of x_t . Baillie (1996) and Hosking (1981, 1996), among others, detail several cases depending upon the value of d . The fractional parameter can take on any real value and enables one to distinguish between unit-root and mean reversion, and within mean reversion, long memory, short memory, and anti-persistence.³ In particular, following Hosking (1981), if $-0.5 < d < 1$, x_t (and, thus, y_t) is a mean-reverting process, but effects of shocks die out at a much slower hyperbolic (rather than geometric) rate, contrary to what happens if $d = 1$ (unit-root case) or $d > 1$ (explosive case) when they persist forever. If $d = 0$, $x_t = u_t$ and x_t is a short-memory process corresponding to stationary and invertible ARMA modeling. In such case, the effect of shocks to u_t on x_t decays geometrically. If $0 < d < 0.5$, then the autocorrelation function (ACT) declines hyperbolically toward zero and x_t is stationary and possesses long memory, while if $0.5 \leq d < 1$, x_t is mean reverting but not (covariance) stationary. Lastly, if $-1 < d < 0$, x_t is referred to as anti-persistent by Mandelbrot (1977), because the spectral density function is dominated by high-frequency components.⁴ Anti-persistence does not appear as a frequent phenomenon in economics. Memory properties mostly focus on persistence. Anti-persistence, however, does appear in a few papers such as Caporale et al. (2020) and Dimitrova et al. (2019). Persistent and anti-persistent processes are two well-known examples of time-series models with hyperbolic decay. In the time domain, persistent processes show a positive long-range dependence between the observations, while anti-persistent often reverse direction and exhibit strong negative autocorrelations. Anti-persistent dynamics may be related to tighter and suboptimal policies: deviations are followed by a rapid over-correction, which produces oscillations around target values.

³ In the context of fractional integration, considerable disagreement exists about the definition of mean reversion. This is mostly apparent when the fractional integration coefficient is in the range $[0.5, 1)$. See Clark and Coggin (2011). For example, Robinson (2003, p. 20), refers to the region with d in the range $[0.5, 1)$ as nonstationary but mean reverting. Thus, although non-stationary, such series behave “like” stationary series in one interesting aspect: they are mean-reverting. In such series, shocks eventually dissipate, and the resulting behavior of the series is called mean reverting regardless of whether the data are stationary or not. This definition of mean reversion implies that the MA coefficients associated with the lags of the series decay. In contrast, Phillips et al. (1999, p. 34) argue against this characterization of mean reversion and indicate that a process with the fractional integration coefficient in the range $[0.5, 1)$ is non-stationary and non-mean reverting, although their impulse response function decays to zero. Thus, in this case, shocks are not persistent. We implicitly adopt the Robinson (2003) definition, implying that (1) a test for fractional integration can be used to determine the existence of mean reversion and (2) mean reversion occurs when the fractional integration parameter d is smaller than 1. This definition is not necessarily restricted to the mean but implies that any shock to the series eventually dies out and the series reverts to its long-term evolution that might be a constant or a long-run trend.

⁴ Alternatively, when $d = 0$, high-frequency and low-frequency contributions are equal, resulting in an uncorrelated time series. When $d > 0$, and as d gets larger, the low-frequency contributions increasingly dominate over the high-frequency ones.

5 Test results for RIRP

Data on real interest rates come from the online appendix to the Bank of England Staff Working Paper No 845 version 1.2, which reconstructs global real interest rates on an annual basis going back to the fourteenth century for France, Holland, Italy, Japan, Spain, the UK, and the USA (Schmelzing 2020).

We define the real interest rate, r_t , as the difference between the nominal interest rate, i_t , and the rate of inflation, π_t , and take the USA as the reference country. Then, the real interest differential, $\text{RIRD}_t^{\text{US}}$, between a given country and the USA is defined as the difference between the real interest rate of a given country, $r_t = i_t - \pi_t$, and the real interest rate of the USA, $r_t^{\text{US}} = i_t^{\text{US}} - \pi_t^{\text{US}}$, and $\text{RIRD}_t^{\text{US}} = r_t - r_t^{\text{US}}$. We estimate the fractional integration parameter over the three globalization periods from 1870 to 1914, 1944 to 1971, and 1989 to 2018.

Following Robinson (1994), we test the null hypothesis $H_0 : d = d_0$ in Eqs. (6) and (7) for d_0 values from -2 to 2 with 0.01 increments and choose the values of d_0 that produce the lowest value of the LM statistic $\hat{r}$. The functional form of $\hat{r}$ is rather complex and is not reported here. See Gil-Alana and Robinson (1997) for details. Robinson (1994) shows that under certain mild regularity conditions, the LM statistic $\hat{r}$ converges in distribution to the standard normal.

Table 1 displays the estimates of d obtained with the Whittle function under three standard deterministic cases in Eq. (6) examined in the literature, that is, the case of no deterministic terms ($\beta_0 = \beta_1 = 0$), an intercept only (β_0 unknown and $\beta_1 = 0$), and an intercept with a linear trend (β_0 and β_1 unknown) under the assumption that u_t in Eq. (7) is a white-noise process. We highlight in bold the significant models according to the deterministic terms in Eq. (6). In parenthesis, we report the 95% confidence bands of non-rejection values of d .

The specification with no deterministic terms proves adequate in the first wave for Germany, Holland, and Japan, and in the second and third waves for Spain and the UK, while the specification with the intercept only is adequate for Spain in the first wave, France, Germany, and Italy in the second wave, and Italy and Japan in the third wave. The linear trend specification, in contrast, is required for France, Italy, and the UK in the first wave, Holland and Japan in the second wave, and France, Germany, and Holland in the third wave.

The issue of time trend or drift and convergence deserves a comment. The issue is not so much relevant if $d > 0$. For instance, if $d = 1$, and the time trend is significant, then Eqs. (6) and (7) taken jointly collapse to

$$(1 - L)y_t = \beta_1 + u_t, \quad t \geq 1, \quad (12)$$

which implies that the first difference of the real interest rate differentials is a random walk with a drift, if u_t is a white-noise process. That is, the real interest rates differ by the drift and random noise. If $d = 0$ and the time trend is significant, then the two real interest rates diverge, but shocks will have only transitory effects. That is, the dynamics of the differentials will return to its original trend. While there is divergence from equilibrium, mean reversion to the long-run trend will take place. In a world with changing policies and price stickiness this outcome is not totally unlikely, as

Table 1 Estimates of d during the three globalization waves: RIRP model (US-based differentials)

Country-US	Wave	No regressors	An intercept	A linear time trend
FRANCE	1870–1914	– 0.01 (– 0.16, 0.25)	– 0.01 (– 0.16, 0.23)	– 0.22 (– 0.43 , 0.13)
	1944–1971	0.98 (0.59, 1.54)	0.79 (0.46 , 1.74)	0.74 (0.36, 1.92)
	1989–2018	0.16 (– 0.15, 0.50)	0.12 (– 0.10, 0.36)	– 0.01 (– 0.25 , 0.36)
GERMANY	1870–1914	0.04 (– 0.18 , 0.41)	0.04 (– 0.15, 0.32)	– 0.40 (– 0.62, 0.02)
	1944–1971	– 0.46 (– 0.56, – 0.28)	– 1.01 (– 1.14 , – 0.63)	– 0.98 (– 1.19, – 0.64)
	1989–2018	0.13 (– 0.12, 0.42)	0.13 (– 0.11, 0.41)	0.00 (– 0.12 , 0.30)
HOLLAND	1870–1914	0.01 (– 0.18 , 0.26)	0.01 (– 0.18, 0.29)	0.00 (– 0.20, 0.29)
	1944–1971	– 0.70 (– 0.81, – 0.02)	– 0.62 (– 0.88, – 0.02)	– 0.67 (– 0.96 , – 0.01)
	1989–2018	0.36 (0.05, 0.78)	0.34 (0.05, 0.75)	0.47 (0.16 , 0.82)
ITALY	1870–1914	0.00 (– 0.14, 0.26)	0.03 (– 0.14, 0.29)	– 0.04 (– 0.23 , 0.26)
	1944–1971	0.05 (– 0.23, 0.46)	0.04 (– 0.21 , 0.52)	1.12 (– 0.07, 1.47)
	1989–2018	0.46 (0.24, 0.80)	0.42 (0.21 , 0.69)	0.46 (0.26, 0.71)
JAPAN	1870–1914	– 0.08 (– 0.31 , 0.37)	– 0.08 (– 0.34, 0.37)	– 0.14 (– 0.42, 0.36)
	1944–1971	0.13 (– 0.08, 0.48)	0.12 (– 0.07, 0.42)	0.04 (– 0.19 , 0.38)
	1989–2018	0.19 (– 0.22, 0.73)	0.17 (– 0.18 , 0.71)	0.17 (– 0.19, 0.70)
SPAIN	1870–1914	0.27 (0.07, 0.56)	0.22 (0.04 , 0.53)	0.14 (– 0.11, 0.53)
	1944–1971	0.06 (– 0.16 , 0.49)	0.06 (– 0.20, 0.49)	0.02 (– 0.34, 0.50)
	1989–2018	0.38 (0.18 , 0.65)	0.37 (0.18, 0.63)	0.42 (0.23, 0.66)
UK	1870–1914	0.00 (– 0.16, 0.22)	0.00 (– 0.15, 0.20)	– 0.16 (– 0.34 , 0.09)
	1944–1971	– 0.03 (– 0.31 , 0.43)	– 0.03 (– 0.30, 0.43)	– 0.02 (– 0.29, 0.45)
	1989–2018	0.47 (0.17 , 0.94)	0.50 (0.17, 0.97)	0.46 (– 0.01, 0.96)

The table reports in bold the significant results on the basis of the deterministic terms. In parenthesis, the 95% confidence band of non-rejection values of d using (Robinson 1994) parametric approach

convergence may not be constant through time. In these first two cases, we may attribute the real interest rate differentials to differences in risk or a risk premium between the two countries. Finally, if d is a non-integer value, Eqs. (6) and (7) can be re-cast as

$$(1 - L)^d y_t = \beta_0 \tilde{I}_t + \beta_1 \tilde{I}_t + u_t, \quad (13)$$

where $\tilde{I}_t = (1 - L)^d 1$, and $\tilde{I}_t = (1 - L)^d t$ both now tend to 0 as t approaches infinity.

Cochrane (2001) argued that variations in risk premiums caused the most variation in asset prices and interest rates in the new view of financial markets. In other words, the variability of risk premiums vastly outweighs the variability of interest rates. Alvarez et al. (2009) develop a general equilibrium monetary model of interest rates and exchange rates with endogenously segmented markets, where exchange rates follow near random-walk processes and interest rate differentials emerge as highly variable and persistent. Thus, this model corresponds to the observation in Cochrane (2001) that risk premiums caused most of the variations in interest rates.

From Table 1, we observe that in 20 out of 21 cases the real interest rate differentials exhibit evidence of mean reversion (i.e., d is significantly below 1 in all cases), indicating that shocks are expected to be transitory and return to their original long-term projections and, thus, supporting the convergence hypothesis of global interest rates.⁵ Only 19, however, are stationary mean-reverting processes. France in the second wave displays non-stationary mean-reverting (i.e., $0.5 \leq d < 1$) evidence. Germany in the second wave exhibits a unit-root MA(1) process as the null of $d = -1$ cannot be rejected. Stationary long memory (i.e., $0 < d < 0.5$) is present in the first wave for Spain and in the third wave for Italy, Holland, Spain, and the UK. This indicates that convergence in these cases takes more time. Short memory (i.e., $d = 0$), on the other hand, appears in the first wave for Germany, Holland, Japan, France, Italy, and the UK, in the second wave for Italy, Japan, Spain, and the UK, and in the third wave for France, Germany, and Japan. For Holland, the estimate of d in the second wave is in the range $-1 < d < 0$. This indicates a strongly anti-persistent process (Debowksi 2011; Bondon and Palma 2006; Hosking 1981). Debowksi (2011) shows that we can generalize ARFIMA processes so that we obtain stationary and invertible processes even for anti-persistent processes with $-1 < d < 0$, and Palma (2007) shows that under regularity conditions, ARFIMA processes are stationary and invertible even with $d > -1$.

Germany is the only case of a non-invertible process (Box and Jenkins 1970). The unit-root MA(1) problem can arise in many modeling contexts, especially if a time series exhibits trend (Davis and Song 2011; Anderson and Takemura 1986). One explanation for this phenomenon is that detrending or demeaning often involves the application of a high-pass filter to the time series. In particular, the filter diminishes or obliterates any power in the time series at low frequencies (including the zero frequency). Consequently, the detrended or demeaned data will have a spectrum with

⁵ Note, however, that the confidence bands of these estimates are quite large. This problem is related to the small sample size of the three waves. See Diebold and Rudebusch (1991) on this issue. Our interpretation of the results focuses more on point estimates, the impulse response functions, and the half-lives than confidence intervals, recognizing at the same time the imprecision inherent to the small sample size.

zero power at frequency zero, which can only be fitted with ARMA process that has a unit root in the MA component (Davis and Song 2011).

Several factors may explain the lack of international integration of Germany during the second wave. One factor, however, is likely to be the time period itself. It is well known that after War World II, the economy of Germany was in shambles. Germany was under the Allied occupation from 1945 to 1949 and then divided into East and West Germany in 1949. The abolition of price controls in 1948 ended repressed inflation and the introduction of fiscal reforms and tax cuts opened the door to the German economic miracle. For this reason, we also re-estimate the fractional model over the restricted period 1948–1971. The specification with no deterministic terms proves adequate. The estimate of d is -0.46 with a 95% confidence band $(-0.63 - 0.12)$. This result suggests that anti-persistence and mean reversion characterize the memory property of Germany during the second wave provided we choose a sample that exclude the immediate post-WWII years. We observe, however, that the unit-root MA(1) problem is not present in the case of Japan and Italy; two countries that after World War II had ended were also devastated. Thus, including the restricted sample for Germany, the evidence in favor of mean reversion of the real interest rate differentials is quite striking, as all estimates of d are less than unity, and some are very much less than unity.

The results in Table 1, however, illustrate also the heterogeneous and unstable dynamics of the departures from RIRP across the three waves, the so-called cross-over phenomenon, which may also be due to the small sample sizes used in these subsamples. For Spain, the dynamics shifts from long memory in the first wave to short memory in the second wave and returns to long memory in the third wave. For France, the fractional parameter shifts from short memory in the first wave to non-stationary long memory in the second and reverts to short memory in the third. Germany shifts from short memory in the first wave, to anti-persistence (in the restricted sample) in the second wave, and to short memory in the third wave. For Holland, the fractional parameter shifts from short memory in the first wave, to anti-persistence in the second wave, and to long-memory in the third wave. Italy and the UK display short memory in the first and second wave, and persistence in the third. Japan is the only series that displays short memory in all three waves. In this sense Japan appears as the most integrated economy in all three globalizations.

During the first wave of globalization, the UK started to dominate the world: geographically, through the establishment of the British Empire; technologically, through the innovations of the industrial revolution; and financially, through the hegemonic role of the pound sterling during the gold standard (Spahn 2001). For these reasons, we conduct an additional analysis of the first wave and consider the case where the UK is the reference country, that is, $RIRD_t^{UK} = r_t - r_t^{UK}$. The results, presented in Table 2, do not show significant differences, except in the case of Spain which changes from long-memory to short-memory processes. In all cases, we cannot reject the hypothesis of mean reversion of the real interest rate differentials. As for the deterministic components, no deterministic terms are required in the case of Japan, whereas an intercept is required for Italy, Germany, and France, and a linear time trend for the remaining cases of Holland, Spain, and the USA. Interestingly, RIRP appears quite resilient to the global financial crisis and the low real interest rate environment after the Great

Table 2 Estimates of d during the first globalization wave (1870–1914): RIRP model (UK-based differentials)

Country	No terms	An intercept	A linear time trend
FRANCE	– 0.16 (– 0.36, 0.18)	– 0.15 (– 0.38 , 0.16)	– 0.13 (– 0.35, 0.17)
GERMANY	– 0.31 (– 0.64, 0.09)	– 0.27 (– 0.49 , 0.08)	– 0.25 (– 0.47, 0.12)
HOLLAND	0.04 (– 0.12, 0.28)	0.04 (– 0.11, 0.27)	– 0.07 (– 0.26 , 0.21)
ITALY	– 0.07 (– 0.22, 0.21)	– 0.06 (– 0.27 , 0.20)	– 0.11 (– 0.31, 0.18)
JAPAN	– 0.11 (– 0.38 , 0.46)	– 0.11 (– 0.42, 0.47)	– 0.12 (– 0.43, 0.47)
SPAIN	0.42 (0.26, 0.65)	0.33 (0.19, 0.56)	0.17 (– 0.04 , 0.50)
USA	– 0.00 (– 0.16, 0.24)	– 0.00 (– 0.15, 0.20)	– 0.16 (– 0.34 , 0.09)

See Table 1

Recession and to hold irrespectively of the nominal exchange rate regimes (the gold standard, the Bretton Woods system, and the current managed float).

We further examine the properties of the real interest rate deviations from RIRP using impulse response analysis (Campbell and Mankiw 1987; Watson 1986; Diebold and Rudebusch 1991; Aloy et al. 2011). Impulse response functions provide useful information about the effect on k -period ahead real interest rate differentials following a one-time change in the innovations u_t , holding everything else constant. The impulse responses are affected by the magnitude of d . The higher the value of d is, the higher the responses are (Aloy et al. 2011). Figure 1 plots the impulse response functions (IRF) of the real interest rate differential with respect to the USA and their associated 95% confidence bands for lags up to $k = 20$. We report in the case of Germany in the second wave the impulse response computed from the estimate obtained over the period 1948–1971.

From Table 1, mean reversion appears to be the main feature in all cases, consistent with the parameter estimates of d , although the effect of a shock displays different patterns across globalization waves. Faster mean reversion toward a long-run equilibrium is observed in the short-memory cases (specifically, all the countries except Spain in the first wave, all the countries except France, Germany, and Holland in the second wave, and France, Germany, and Japan in the third wave). France in the second wave shows a slow declining pattern, which is consistent with the non-stationarity but mean reversion of the series, as apparent also from the widening of the confidence bands as k increases. Anti-persistence is apparent in the cases of Holland and Germany in the second wave from the quick damped oscillation of the stationary process. Slower mean reversion is observed in the long-memory cases (specifically, Spain in the first wave, France in the second, and Holland, Italy, Spain, and the UK in the third). Furthermore, in the long-memory cases, a positive shock exerts an immediate positive effect on the real interest rate differential, with the magnitude remaining positive even after 20 years and the effect disappearing at a hyperbolic rate. Conversely, in the short-memory cases, a positive shock exerts a negative effect, but this effect dies out almost immediately. The anti-persistence cases, on the other hand, indicate that the pattern of damped oscillations dies out after 2–3 years.

Fig. 1 Impulse response functions of real interest rate differentials in the three waves with respect to the USA. *Note:* The graphs display the IRF together with the 95-percent confidence band based on the results of the models in Table 1 with preferred deterministic terms

Lastly, we compute the half-life, which measures how long it takes for a unit shock to dissipate by half. For ARMA models, an analytical expression for the half-life can be derived; for example, it is well known that the half-life of the AR(1) model $y_t = \varphi y_{t-1} + \varepsilon_t$ is given by $h = -\log(2)/\log(\varphi)$. For the ARFIMA model, however, the half-life remains difficult to compute. Following Aloy et al. (2011), we approach the problem by linearly interpolating the impulse response functions. That is, we estimate the half-life as follows. If k is such that $\text{IRF}[k] \geq 0.5 \geq \text{IRF}[k + 1]$, then the

Fig. 1 continued

linear approximation for the half-life estimates is given by

$$h = (0.5 - (k + 1)\text{IRF}[k] + k\text{IRF}[k + 1]) / (\text{IRF}[k + 1] - \text{IRF}[k]). \quad (14)$$

See Aloy et al. (2011) for details.

No half-life measure is able to convey the informational content of an impulse response since it is only a summary statistic. Hence, it is possible that cases exist where any half-life measure will not be able to convey properly and completely the content of the underlying impulse response. Nevertheless, the half-life measure has the advantage that it can readily be interpreted in time units and provides an alternative way to cast the real interest rate convergence (Chortareas and Kapetanios 2013).

Table 3 reports the half-lives for real interest rate differentials with respect to the US real interest rate. According to the results in Table 3, in the first wave of globalization, the half-life ranges between one and one-and-a-half years; in the second wave, the half-life ranges between 1 and 2 years, with the exception of France with a 5-year-and-8-month half-life and Holland with a 2-year-and-6-month half-life; finally, in the third wave, the half-life ranges between 1 year and 6 months and 2 years, with the exception of Germany with a 6-month half-life. It, thus, appears that with few exceptions, the

Table 3 Half-lives estimates (in years) of bilateral real interest rate differentials with respect to the USA

Country	Waves	Half-life
FRANCE	1st wave	1.049
	2nd wave	5.733
	3rd wave	1.525
GERMANY	1st wave	1.520
	2nd wave	1.248
	3rd wave	0.500
HOLLAND	1st wave	1.505
	2nd wave	2.515
	3rd wave	1.943
ITALY	1st wave	1.480
	2nd wave	1.520
	3rd wave	1.862
JAPAN	1st wave	1.463
	2nd wave	1.520
	3rd wave	1.602
SPAIN	1st wave	1.358
	2nd wave	1.595
	3rd wave	1.806
UK	1st wave	1.431
	2nd wave	1.485
	3rd wave	1.943

The half-life is calculated by using a linear interpolation as follows: if k is such that $\text{IRF}[k] \geq 0.5 \geq \text{IRF}[k+1]$ then the linear approximation for the half-life estimates is given by $h = (0.5 - \text{IRF}[k]) / (\text{IRF}[k] - \text{IRF}[k+1])$

speed of mean reversion is high, indicating that real interest rate differentials tend to be short lived, with the half-lives of shocks lower in the first and the second waves (under fixed exchange rates) than in the third wave, probably due to higher price flexibility before WWII (Dreger 2010).

6 Test results for relative PPP and UIP

The finding of mean reversion in RIRP during the three waves of globalization, as highlighted by our findings, invites further inquiry into what are its driving and restraining sources. The RIRP condition, as noted previously, combines two cornerstones in international economics, UIP, and ex-ante relative PPP. See Marston (1995) and MacDonald and Marsh (1999) for details. In this regard, the mean reversion of departures from relative PPP and UIP parities can serve as an additional indicator, respectively, for commodity markets convergence and financial markets integration (Dreger 2010) and

may help to explain the driving forces of mean reversion in RIRP. This section provides this additional information.

The data on exchange rates and prices come from the Macrohistory Database (Jordà et al. 2017), a comprehensive macro-financial panel dataset of 17 countries spanning 1870–2017. We updated the data to 2018 using the World Bank's World Development Indicators.

Table 4 displays tests of relative PPP during the three globalization periods. The Whittle estimates of d are displayed under the three standard deterministic cases along with the 95% confidence bands. Equation (2) provides the basis for the empirical test of relative PPP. That is, we test the validity of the relative PPP in the long run by examining whether the difference between depreciation of the exchange rate ($s_{t+1} - s_t$) and inflation differential ($\pi_{t+1} - \pi_{t+1}^*$) is mean reverting.⁶ We observe that in the first wave, the linear trend specification is required in all cases; in the second wave, the no regressors specification proves adequate for France, Italy, Spain, and the UK, while the linear trend specification is required for Germany, Holland, and Japan; in the third wave, on the other hand, the linear trend specification is required in all cases.

Evidence of mean reversion of relative PPP differentials (i.e., $d < 1$) does not emerge in all cases. Short memory (i.e., $d = 0$) is present in the cases of France and Italy in the second wave, while stationary long memory (i.e., $0 < d < 0.5$) and non-stationary long memory (i.e., $0.5 < d < 1$) appear in no cases. Anti-persistence ($-1 < d < 0$) is displayed by Holland, Italy, Japan, Spain, and the UK in the third wave. The preponderance of the findings, however, indicates that $d = -1$, which signals the presence of an MA(1) with a unit-root representation. Estimation of the model for Germany over the restricted period 1948–1971 produces an estimate of d of -0.96 with a 95% confidence band ($-1.19 - 0.55$) under the linear trend specification. This result suggests that the memory property of PPP differentials for Germany during the second wave is still characterized by a unit-root MA(1) representation. A puzzle has emerged in the literature with respect to PPP (Rogoff 1996; Lothian and Taylor 1996; Frankel 1979; Taylor et al. 2001), which refers to the very slow rates of mean reversion in the deviations of the real exchange rate from its PPP implied value. The majority of our findings indicate that the puzzle of slow mean-reverting behavior of the PPP differentials occurs in 14 out of 21 cases, including the three globalization waves.

Table 5 displays tests of UIP for the three globalization periods.⁷ Equation (1) provides the basis for the empirical test. That is, we test the validity of UIP in the long run by examining whether the difference between the nominal interest rate differential ($i_t - i_t^*$) and the rate of depreciation of the exchange rate ($s_{t+1} - s_t$) is mean reverting.

⁶ Tests of the PPP using $I(d)$ methods are confined to its absolute form. Some authors have employed fractional cointegration tests between nominal exchange rates and relative prices (Diebold et al. 1991; Cheung and Lai 1993; Baillie and Bollerslev 1994; Masih and Masih 2004; Triki and Maktouf 2015, among many others). Others have focused on the properties of real exchange rate (Holmes 2001, 2002, Gil-Alana 2000). For a recent survey of the literature, see Sarno and Taylor (2002).

⁷ Uncovered interest parity has generally been rejected in empirical studies. Even more puzzling, many studies have found that currencies offering higher interest rates tend to appreciate against currencies yielding lower interest rates, creating the forward premium anomaly (Meese and Rogoff 1983; Froot and Thaler 1990; Alexius 2001; Chinn and Meredith 2004, among others). For a recent survey of the literature, see Sarno and Taylor (2002).

Table 4 Estimates of $\hat{d}$ during the three globalization waves: relative PPP model (US-based differentials)

Country	Wave	No regressors	An intercept	A linear time trend
FRANCE	1870–1914	-1.02 (-1.09, -0.90)	-0.94 (-1.07, -0.73)	-1.10 (-1.26, -0.84)
	1944–1971	-0.21 (-0.57, 0.30)	-0.19 (-0.52, 0.33)	-0.29 (-0.67, 0.90)
	1989–2018	-0.75 (-0.87, -0.54)	-0.81 (-1.02, -0.53)	-0.99 (-1.20, -0.69)
GERMANY	1870–1914	-1.04 (-1.13, -0.86)	-0.94 (-1.07, -0.74)	-1.25 (-1.39, -1.01)
	1944–1971	-1.40 (-1.53, -1.15)	-1.48 (-1.71, -1.12)	-1.83 (-2.04, -1.39)
	1989–2018	-1.03 (-1.15, -0.82)	-1.11 (-1.34, -0.74)	-1.19 (-1.45, -0.79)
HOLLAND	1870–1914	-1.07 (-1.18, -0.90)	-0.95 (-1.10, -0.73)	-1.19 (-1.44, -0.85)
	1944–1971	-1.27 (-1.51, -1.37)	-0.95 (-1.14, -0.59)	-1.13 (-1.38, -0.69)
	1989–2018	-0.89 (-1.03, -0.22)	-0.49 (-0.68, -0.19)	-0.64 (-0.86, -0.28)
ITALY	1870–1914	-1.03 (-1.12, -0.85)	-0.95 (-1.09, -0.72)	-1.16 (-1.36, -0.86)
	1944–1971	0.01 (-0.27, 0.43)	0.01 (-0.25, 0.49)	1.14 (-0.11, 1.64)
	1989–2018	-0.44 (-0.65, -0.09)	-0.43 (-0.66, -0.10)	-0.58 (-0.85, -0.17)
JAPAN	1870–1914	-1.25 (-1.42, -0.61)	-0.92 (-1.09, -0.59)	-0.92 (-1.10, -0.55)
	1944–1971	-1.01 (-1.12, -0.79)	-1.04 (-1.23, -0.71)	-1.54 (-1.93, -0.97)
	1989–2018	-0.64 (-0.88, -0.17)	-0.69 (-0.95, -0.18)	-0.49 (-1.28, -0.20)
SPAIN	1870–1914	-1.01 (-1.08, -0.90)	-0.90 (-1.02, -0.72)	-0.92 (-1.05, -0.71)
	1944–1971	-1.09 (-1.54, -0.52)	-1.10 (-1.47, -0.52)	-1.12 (-1.47, -0.53)
	1989–2018	-0.56 (-0.74, -0.27)	-0.56 (-0.76, -0.27)	-0.67 (-0.88, -0.33)
UK	1870–1914	-1.08 (-1.16, -0.96)	-1.06 (-1.18, -0.88)	-1.18 (-1.33, -0.97)
	1944–1971	-1.07 (-1.36, -0.54)	-1.05 (-1.37, -0.52)	-1.09 (-1.41, -0.51)
	1989–2018	-0.54 (-1.18, -0.10)	-0.43 (-0.63, -0.11)	-0.56 (-0.82, -0.17)

See Table 1

Table 5 Estimates of d during the three globalization waves: UIP model (US-based differentials)

Country	Wave	No regressors	An intercept	A linear time trend
FRANCE	1870–1914	0.88 (0.63, 1.19)	0.84 (0.61, 1.14)	0.84 (0.61, 1.14)
	1944–1971	0.79 (0.59, 1.08)	0.71 (0.48, 1.05)	0.71 (0.35, 1.05)
	1989–2018	0.29 (0.00, 0.84)	0.29 (0.00, 0.66)	0.17 (– 0.17, 0.62)
GERMANY	1870–1914	0.91 (0.69, 1.20)	0.78 (0.60, 1.07)	0.76 (0.55, 1.07)
	1944–1971	0.60 (0.20, 1.20)	0.95 (0.13, 1.43)	0.98 (0.28, 1.38)
	1989–2018	0.25 (0.05, 0.65)	0.31 (0.07, 0.63)	0.03 (– 0.27, 0.52)
HOLLAND	1870–1914	1.07 (0.84, 1.38)	1.04 (0.83, 1.33)	1.04 (0.83, 1.33)
	1944–1971	– 0.16 (– 0.33, 0.32)	– 0.27 (– 0.68, 0.32)	– 0.20 (– 0.62, 0.39)
	1989–2018	0.29 (0.03, 0.82)	0.33 (0.04, 0.71)	0.14 (– 0.18, 0.63)
ITALY	1870–1914	0.68 (0.52, 0.92)	0.78 (0.55, 1.25)	0.82 (0.59, 1.26)
	1944–1971	0.78 (0.47, 1.19)	0.58 (0.24, 1.18)	0.65 (0.13, 1.18)
	1989–2018	0.43 (0.18, 0.78)	0.34 (0.14, 0.59)	0.29 (0.08, 0.59)
JAPAN	1870–1914	0.81 (0.62, 1.10)	0.70 (0.54, 0.97)	0.69 (0.50, 0.97)
	1944–1971	0.58 (0.40, 0.82)	0.58 (0.42, 0.81)	0.58 (0.42, 0.81)
	1989–2018	0.64 (0.46, 0.86)	0.64 (0.43, 0.97)	0.62 (0.38, 0.97)
SPAIN	1870–1914	1.17 (0.98, 1.44)	1.17 (0.95, 1.48)	1.17 (0.95, 1.47)
	1944–1971	0.73 (0.46, 1.05)	0.37 (0.12, 0.81)	0.28 (– 0.30, 0.82)
	1989–2018	0.48 (0.20, 0.87)	0.40 (0.17, 0.69)	0.39 (0.16, 0.70)
UK	1870–1914	0.90 (0.69, 1.19)	0.98 (0.79, 1.21)	0.98 (0.83, 1.21)
	1944–1971	0.07 (– 0.10, 1.06)	0.11 (– 0.22, 0.90)	– 0.18 (– 0.99, 0.92)
	1989–2018	1.00 (0.48, 0.71)	0.88 (0.46, 1.56)	0.88 (0.51, 1.58)

See Table 1

The Whittle estimates of d are displayed under the three standard deterministic cases along with the 95% confidence bands.

The specification with no deterministic terms proves adequate in the first and second waves for France and for Italy and Japan in the second wave. The specification with only the intercept is adequate for Germany, Holland, and Spain in the first wave, Germany and Holland in the second wave, and Japan and the UK in the third wave. The linear trend specification is required in all remaining cases, that is, in the first wave for Italy and the UK, in the second wave for Japan and Spain, and in the third wave for France, Germany, Holland, Italy, and Spain.

Unit-root behavior is evident in case of Holland, Spain, and the UK in the first wave and Germany in the second wave. The estimation of the model for Germany over the restricted period 1948–1971, however, produces, under the preferred specification of the intercept only, an estimate of d of 0.01, which implies short-memory behavior. We should note, however, the wide 95% confidence band (– 1.09 1.23). The remaining estimates provide evidence of mean reversion and, thus, show that the UIP hypothesis generally holds in the three globalization waves. Short memory ($d = 0$) is present

in the second wave in the cases of Holland and Spain and in the third wave in the cases of France, Germany, and Holland. Stationary long memory ($0 < d < 0.5$) is present in the third wave in the cases of Italy and Spain. Non-stationary long memory ($0.5 \leq d < 1$) defines the remaining cases.

We gain a final perspective by summarizing the driving and restraining sources of RIRP. Table 6 contains a compact summary of the results. Three results, in particular, should be highlighted. First, the three parities hold simultaneously in the long run in 7 out of 21 cases: France, Italy, and Spain in the second wave, and in the third wave Holland, Italy, Japan, and the UK. Second, RIRP and UIP hold in 11 out of 21 cases: France, Germany, Italy, and Japan in the first wave, Germany, Holland, Japan, and the UK in the second wave, and France, Germany, and Spain in the third wave. Third, RIRP holds without the support of relative PPP and UIP in three cases: Holland, Spain, and the UK in the first wave. Thus, in the final analysis, the evidence in favor of real interest rate convergence appears to be driven more by UIP than relative PPP.

Table 6 Summary of the mean reverting parity outcomes

Country	Wave	RIRP	Relative PPP	UIP
FRANCE	1870–1914	Y	N	Y
	1944–1971	Y	Y	Y
	1989–2018	Y	N	Y
GERMANY	1870–1914	Y	N	Y
	1944–1971	Y	N	Y
	1989–2018	Y	N	Y
HOLLAND	1870–1914	Y	N	N
	1944–1971	Y	N	Y
	1989–2018	Y	Y	Y
ITALY	1870–1914	Y	N	Y
	1944–1971	Y	Y	Y
	1989–2018	Y	Y	Y
JAPAN	1870–1914	Y	N	Y
	1944–1971	Y	N	Y
	1989–2018	Y	Y	Y
SPAIN	1870–1914	Y	N	N
	1944–1971	Y	Y	Y
	1989–2018	Y	N	Y
UK	1870–1914	Y	N	N
	1944–1971	Y	N	Y
	1989–2018	Y	Y	Y

Y means that the parity outcome is mean reversion; N means that the parity outcome is not mean reversion. The outcomes for Germany are based on the restricted sample

7 Concluding comments

Did globalization integrate the financial markets into the world capital market? We examine the extent to which three waves of globalization affected the departures from RIRP. This paper provides first-time evidence on the memory properties of real interest rate differentials during three waves of globalization, 1870–1914, 1944–1971, and 1989–2018. We apply fractional integration methods, a more flexible approach than the standard ARMA framework based on the classical dichotomy of $I(0)/I(1)$, since it allows the integration parameter to assume fractional values and, thus, capture a much wider range of stochastic dynamics.

The test results for RIRP indicate that real interest rate differentials converge: short memory is present in 13 out of 21 cases (6 in the first wave, 4 in the second, and 3 in the third), long memory occurs in 6 out of 21 cases (1 in the first wave, 1 in the second wave, and 4 in the third), with anti-persistence occurring in 2 cases (in the second wave). Interestingly, RIRP appears quite resilient to the global financial crisis and the low real interest rate environment after the Great Recession and to hold irrespectively of the nominal exchange rate regimes (the gold standard, the Bretton Woods system, and the current managed float). These findings are further confirmed by the impulse response functions and the half-lives computations, a relatively new and potentially fruitful area that endows fractional integration models with richer and more realistic dynamic patterns. The results from the UK-based differentials for the first wave exhibit even stronger and more robust results in favor of RIRP, as they indicate short-memory outcomes in all cases, including Spain. The policy implication of these results is quite important, as they suggest that the monetary authorities may find it difficult to affect independently real interest rates.

We shed further light on financial market integration during the three globalization waves by assessing the memory properties of UIP and relative PPP differential processes, defined as the economic fundamentals of RIRP. The test results for UIP indicate that mean reversion holds in 18 out of 21 cases. The unit-root hypothesis cannot be rejected in 3 cases. Short memory is present in 5 cases, stationary long memory is present in 2 cases, and non-stationary long memory is present in 11 cases. The latter finding may be a reason why conventional unit-root tests do reject UIP. Finally, evidence in favor of relative PPP is scant. We find short memory only in 2 of 21 cases (France and Italy in the second wave) and anti-persistence in 5 out of 21 cases (Spain in the second wave; Holland, Italy, Japan, and the UK in the third wave). The preponderance of the test results, 14 out of 21 cases, indicates that $d = -1$, which signals the presence of an MA(1) with a unit-root representation. In such cases, deviations from relative PPP show no tendency to hold in the long run during the three globalization periods.

As a final contribution, the findings in this paper open up an interesting econometric issue. That is, the issue of unit-root MA(1) collapses of the fractional integration model. We believe that investigating further this issue could provide an important agenda and a fruitful endeavor for future research.

Acknowledgements We acknowledge the helpful comments of two anonymous referees. Remaining errors reflect our mistakes.

Appendix

Tables 7 presents the estimates of d for the real interest model (RIR) under the three deterministic alternatives. We highlight in bold the significant models according to the deterministic terms in Eq. (6). In parenthesis, we report the 95% confidence bands of non-rejection values of d . The preponderance of the outcomes indicates short-memory, 16 out of 24 times. Waves 2 and 3 include six short-memory outcomes. The exceptions are France and the USA that record long-memory outcomes in wave 2 and Spain and Italy that record long-memory outcomes in wave 3. Wave 1 includes the most diversity with France recording an anti-persistence outcome and Holland, Spain, and the UK recording long-memory outcomes. We observe that with the exception of Spain in the second wave, no cases prove adequate for the specification with no deterministic terms. The intercept only specification is adequate for Germany, Japan, and the UK in the first wave, and France and Italy in the second wave. The linear trend is appropriate

Table 7 Estimates of d during the three globalization waves: RIR model

Country	Wave	No regressors	An intercept	A linear time trend
FRANCE	1870–1914	0.21 (0.06, 0.51)	0.04 (– 0.09, 0.22)	– 0.53 (– 0.74 , – 0.26)
	1944–1971	1.04 (0.61, 1.72)	0.79 (0.46 , 1.74)	0.74 (0.36, 1.92)
	1989–2018	0.65 (0.44, 0.91)	0.46 (0.34, 0.58)	– 0.11 (– 0.32 , 0.19)
GERMANY	1870–1914	– 0.24 (– 0.35, 0.09)	– 0.24 (– 0.43, 0.06)	– 0.40 (– 0.62 , 0.02)
	1944–1971	0.01 (– 0.13, 0.23)	0.01 (– 0.18, 0.27)	– 0.21 (– 0.42 , 0.11)
	1989–2018	0.66 (0.44, 1.02)	0.48 (0.34, 0.67)	– 0.09 (– 0.35 , 0.30)
HOLLAND	1870–1914	0.35 (0.15, 0.62)	0.26 (0.10 , 0.53)	0.13 (– 0.11, 0.51)
	1944–1971	0.30 (0.06, 0.50)	0.26 (0.06, 0.51)	0.11 (– 0.12 , 0.44)
	1989–2018	0.62 (0.38, 0.99)	0.45 (0.28, 0.62)	0.16 (– 0.16 , 0.62)
ITALY	1870–1914	0.04 (– 0.15, 0.37)	0.03 (– 0.14, 0.29)	– 0.09 (– 0.30 , 0.24)
	1944–1971	0.06 (– 0.22, 0.45)	0.06 (– 0.19 , 0.47)	1.14 (– 0.05, 1.29)
	1989–2018	0.63 (0.42, 0.88)	0.45 (0.31, 0.61)	0.40 (0.24 , 0.60)
JAPAN	1870–1914	– 0.09 (– 0.33, 0.47)	– 0.09 (– 0.40 , 0.46)	– 0.09 (– 0.41, 0.47)
	1944–1971	0.29 (0.11, 0.57)	0.27 (0.10, 0.51)	0.18 (– 0.02 , 0.47)
	1989–2018	0.48 (0.25, 0.79)	0.27 (0.19, 0.60)	– 0.14 (– 0.53 , 0.43)
SPAIN	1870–1914	0.57 (0.41, 0.78)	0.44 (0.29, 0.69)	0.31 (0.07 , 0.67)
	1944–1971	0.29 (– 0.37 , 1.03)	0.26 (– 0.28, 1.05)	0.14 (– 0.41, 1.05)
	1989–2018	0.56 (0.34, 0.84)	0.43 (0.27, 0.64)	0.43 (0.25 , 0.66)
UK	1870–1914	– 0.27 (– 0.03, 0.59)	0.18 (0.00 , 0.51)	0.11 (– 0.19, 0.50)
	1944–1971	0.51 (0.28, 1.21)	0.52 (0.30, 1.24)	0.19 (– 0.38 , 1.32)
	1989–2018	0.66 (0.44, 0.95)	0.58 (0.37, 0.88)	0.35 (– 0.12 , 0.86)
USA	1870–1914	0.23 (0.01, 0.47)	0.15 (0.00, 0.34)	– 0.18 (– 0.45 , 0.19)
	1944–1971	0.30 (0.07, 0.73)	0.29 (0.08, 0.69)	0.21 (0.09 , 0.64)
	1989–2018	0.57 (0.39, 0.75)	0.42 (0.28, 0.59)	– 0.11 (– 0.42 , 0.43)

for France, Germany, Italy, Spain, and the USA in the first wave, Germany, Holland, Japan, the UK, and the USA in the second wave, and all eight countries in the third wave. Evidence of mean reversion of the real interest rates (i.e., $d < 1$) emerges in all cases. Persistence or long memory (i.e., $0 < d < 1$) is evident for Holland, Spain, and the UK in the first wave, France and the USA in the second wave, and Italy and Spain in the third wave. Short memory (i.e., $d = 0$) is evident for Germany, Italy, Japan, and the USA in the first wave, Germany, Holland, Italy, Japan, Spain, and the UK in the second wave, France, Germany, Holland, Japan, the UK, and the USA in the third wave. France in the first wave is the only country that displays anti-persistent properties.

References

- Alexius A (2001) Uncovered interest parity revisited. *Rev Int Econ* 9:505–517
- Aloy M, Boutahar M, Gente K, Péguin-Feissolle A (2011) Purchasing power parity and the long memory properties of real exchange rates: does one size fit all? *Econ Model* 28:1279–1290
- Alvarez F, Atkeson A, Kehoe PJ (2009) Time-varying risk, interest rates, and exchange rates in general equilibrium. *Rev Econ Stud* 76:851–878
- Anderson TW, Takemura A (1986) Why do noninvertible estimated moving averages occur? *J Time Ser Anal* 7:235–254
- Arghyrou MG, Gregoriou A, Kontonikas A (2009) Do real interest rates converge? Evidence from the European Union. *J Int Finan Markets Inst Money* 19:447–460
- Awad MA, Goodwin BK (1998) Dynamic linkages among real interest rates in international capital markets. *J Int Money Financ* 17:881–907
- Bahmani-Oskooee M, Chang T, Elmi Z, Ranjbar O (2019) Real interest rate parity and Fourier quantile unit root test. *Bull Econ Res* 71:348–358
- Baillie RT (1996) Long-memory processes and fractional integration in econometrics. *J Econom* 73:5–59
- Baillie RT, Bollerslev T (1994) Cointegration, fractional cointegration and exchange rate dynamics. *J Finance* 49:737–745
- Baines D (1994) European labor markets, emigration and internal migration, 1850–1913. In: Hatton TJ, Williamson JG (eds) *Migration and the international labor market, 1850–1939*. Routledge, London, pp 35–54
- Bondon P, Palma W (2006) A class of antipersistent processes. *J Time Ser Anal* 28:261–273
- Bordo M, Taylor A, Williamson J (2003) Introduction to “Globalization in historical perspective”. In: *Globalization in historical perspective*. National Bureau of Economic Research, pp 1–10
- Box G, Jenkins G (1970) *Time series analysis: forecasting and control*. Holden Day, San Francisco
- Campbell JY, Mankiw NG (1987) Are output fluctuations transitory? *Quart J Econ* 102:857–888
- Caporale GM, Gil-Alana LA, Poza C (2020) Persistence, non-linearities and structural breaks in European stock market indices. *Q Rev Econ Finance* 77:50–61
- Cheung YW, Lai KS (1993) A fractional cointegration analysis, of purchasing power parity. *J Bus Econ Stat* 11:103–112
- Cheung YW, Lai KS (2001) Long memory and nonlinear mean reversion in Japanese yen-based real exchange rates. *J Int Money Financ* 20:115–132
- Chinn MD, Meredith G (2004) Monetary policy and long-horizon uncovered interest parity. *IMF Staff Pap* 51:409–430
- Chortareas G, Kapetanios G (2013) How puzzling is the PPP puzzle? An alternative half-life measure of convergence to PPP. *J Appl Economet* 25:435–457
- Chortareas G, Kapetanios G, Magkonis G (2018) Resuscitating real interest rate parity: new evidence from panels. *Eur J Finance* 24:1176–1189
- Christiano LJ, Eichenbaum M, Charles Evans C (1997) Sticky price and limited participation models: a comparison. *Eur Econ Rev* 4:1201–1249
- Clark S, Coggin T (2011) Are U.S. stock prices mean reverting? Some new tests using fractional integration models with overlapping data and structural breaks. *Empir Econ* 40:373–391

- Cochrane JH (2001) Asset pricing: revised edition. Princeton University Press, Princeton
- Collier P, Dollar D (2002) Globalization, growth and poverty: building an inclusive world economy. Oxford University Press, New York
- Çorakcı A, Emirmahmutoglu F, Omay T (2017) Re-examining the real interest rate parity hypothesis (RIPH) using panel unit root tests with asymmetry and cross-section dependence. *Empirica* 44:91–120
- Cuestas JC, Harrison B (2010) Further evidence on the real interest rate parity hypothesis in Central and Eastern European Countries: unit roots and nonlinearities. *Emerg Mark Financ Trade* 46:22–39
- Dahlhaus R (1989) Efficient parameter estimation for self-similar processes. *Ann Stat* 17:1749–1766
- Davis RA, Song L (2011) Unit roots in moving averages beyond first order. *Ann Stat* 39:3062–3091
- Debowski L (2011) On processes with hyperbolically decaying autocorrelations. *J Time Ser Anal* 32:580–584
- Dickey DA, Fuller WA (1979) Distribution of the estimators for autoregressive time series with a unit root. *J Am Stat Assoc* 74:427–431
- Diebold FX, Rudebusch GD (1991) Long memory and persistence in aggregate output. *J Monet Econ* 24:189–209
- Diebold E, Husted S, Rush M (1991) Real exchange rates under the gold standard. *J Polit Econ* 99:1252–1271
- Dimitrova V, Fernández-Martínez M, Sánchez-Granero MA, Trinidad Segovia JE (2019) Some comments on Bitcoin market (in)efficiency. *PLoS ONE* 14:e0219243
- Dreger C (2010) Does the nominal exchange rate regime affect the real interest parity condition? *N Am J Econ Finance* 21:274–285
- Easterlin R (1961) Influences in European overseas emigration before World War I. *Econ Dev Cult Change* 9:331–351
- Elliott G, Stock JH, Rothenberg TJ (1996) Efficient tests for an autoregressive unit root. *Econometrica* 64:813–836
- Ferreira AL, León-Ledesma MA (2007) Does the real interest parity hypothesis hold? Evidence for developed and emerging markets. *J Int Money Financ* 26:364–382
- Frankel J (1979) On the mark: a theory of floating exchange rates based on real interest differentials. *Am Econ Rev* 69:610–622
- Froot KA, Thaler RH (1990) Anomalies of foreign exchange. *J Econ Perspect* 4:179–192
- Gil-Alana LA (2000) Mean reversion in the real exchange rates. *Econ Lett* 69:285–288
- Gil-Alana LA, Robinson PM (1997) Testing of unit roots and other nonstationary hypotheses in macroeconomic time series. *J Econom* 80:241–268
- Güney PO, Hasanov M (2014) Real interest rate parity hypothesis in post-Soviet countries: evidence from unit root tests. *Econ Model* 36:120–129
- Hatton TJ, Williamson JG (1994) International migration 1850–1939: an economic survey. In: Hatton TJ, Williamson JG (eds) *Migration and the international labor market, 1850–1939*. Routledge, London, pp 3–32
- Holmes MJ (2001) New evidence on real exchange rate stationarity and purchasing power parity in less developed countries. *J Macroecon* 23:601–614
- Holmes MJ (2002) Purchasing power parity and the fractional integration of real exchange rate: new evidence for less developed countries. *J Econ Dev* 27:125–135
- Holmes MJ, Maghrebi N (2004) Asian real interest rates, nonlinear dynamics, and international parity. *Int Rev Econ Financ* 13:387–405
- Hosking JRM (1981) Fractional differencing. *Biometrika* 68:165–176
- Hosking JRM (1996) Asymptotic distributions of the sample mean, autocovariances, and autocorrelations of long memory time series. *J Econom* 73:261–284
- Im KS, Pesaran MH, Shin Y (2003) Testing for unit roots in heterogeneous panels. *J Econom* 115:53–74
- Jordà Ò, Schularick M, Taylor AM (2017) Macrofinancial history and the new business cycle fact. In: NBER Macroeconomics annual 2016, Vol. 31, edited by M. Eichenbaum and J. A. Parker, University of Chicago Press
- Lai KS (1997) Long-term persistence in the real interest rate: evidence of a fractional unit root. *Int J Financ Econ* 2:225–235
- Leviv RM (2013) Evidence on financial globalization and crises: interest rate parity. In: Caprio G (eds) *The evidence and impact of financial globalization*, vol 3. Elsevier, pp 417–427
- MacDonald R, Marsh IW (1999) Exchange rate modelling. Kluwer Academic, Boston
- Mandelbrot BB (1977) *Fractals: form, chance and dimension*. Free Press, New York

- Mark NC (1955) Some evidence on the international equality of real interest rates. *J Int Money Financ* 4:189–208
- Marston RC (1995) International financial integration: a study of interest differentials between the major industrial countries. Cambridge University Press, Cambridge
- Masih AMM, Masih R (2004) Fractional cointegration, low frequency dynamics and long-run purchasing power parity: an analysis of the Australian dollar over its recent float. *Appl Econ* 36:593–605
- Meese RA, Rogoff K (1983) Empirical exchange rate models of the seventies. Do they fit out of sample? *J Int Econ* 14:3–24
- Mishkin FS (1984) Are real interest rates equal across countries? An empirical investigation of international parity conditions. *J Finance* 39:1345–1357
- Mussa M (1976) The exchange rate, the balance of payments and monetary and fiscal policy under a regime of controlled floating. *Scand J Econ* 78:229–248
- Nakagawa H (2002) Real exchange rates and real interest rate differentials: implications of nonlinear adjustment in real exchange rates. *J Monet Econ* 49:629–649
- O'Rourke K, Williamson JG, Hatton TJ (1994) Mass migration, commodity market integration and real wage convergence: the late nineteenth-century Atlantic economy. In: Hatton TJ, Williamson JG (eds) *Migration and the international labor market, 1850–1939*. Routledge, London, pp 203–220
- Obstfeld M, Taylor A (2003) Globalization and capital markets. In: *Globalization in historical perspective*. National Bureau of Economic Research, pp 121–188
- Ortiz-Ospina E, and Roser M (2017) International trade. Published online at OurWorldInData.org
- Palley TI (2018) Three globalizations, not two: rethinking the history and economics of trade and globalization. *Eur J Econ Econ Policies Interven* 15:174–192
- Palma W (2007) Long memory time series: theory and methods. Wiley-Interscience, Hoboken
- Pesaran MH (2007) A simple panel unit root test in the presence of cross-section dependence. *J Appl Econom* 22:265–312
- Phillips PCB, Xiao Z (1999) A primer on unit root testing. In: McAleer M, Oxley L (eds) *Practical issues in cointegration analysis*. Blackwell Publishers, Oxford
- Phylaktis K (1999) Capital market integration in the Pacific Basin region: an impulse response analysis. *J Int Money Financ* 18:267–287
- Piketty T (2014) *Capital in the twenty-first century* (trans Goldhammer, A). Belknap Press of Harvard University Press, Cambridge
- Rebucci A, Hartley JS, Jimenez D (2021) An event study of COVID-19 central bank quantitative easing in advanced and emerging economies. NBER Working Paper 27339
- Robinson PM (1994) Efficient tests of nonstationary hypotheses. *J Am Stat Assoc* 89:1420–1437
- Robinson PM (2003) Time series with long memory. Oxford University Press, Oxford
- Rodrik D (1998) Symposium on globalization in perspective: an introduction. *J Econ Perspect* 12:3–8
- Rogoff K (1996) The purchasing power parity puzzle. *J Econ Literature* 34:647–668
- Rose A (1988) Is the real interest rate stable? *J Finance* 43:1095–1112
- Sarno L, Taylor MP (2002) Purchasing power parity and the real exchange rate. *IMF Staff Pap* 49:65–105
- Schmelzing P (2020) Eight centuries of global real interest rates, R-G, and the 'suprasecular' decline, 1311–2018. Bank of England Staff Working Paper 845
- Shimotsu K, Phillips PCB (2005) Exact local Whittle estimation of fractional integration. *Ann Stat* 33:1890–1933
- Smallwood A, Norrbin SC (2008) An encompassing test of real interest rate equalization. *Rev Int Econ* 16:114–126
- Sowell FB (1992) Maximum likelihood estimation of stationary univariate fractionally integrated time series models. *J Econom* 53:165–188
- Spahn HP (2001) The hegemony of pound sterling in the gold standard. In: *From gold to euro*. Springer, Berlin, Heidelberg, pp 97–125
- Sun Y, Phillips PCB (2004) Understanding the Fisher equation. *J Appl Econom* 19:869–886
- Taylor M, Sarno L (2001) Nonlinear adjustment in real exchange rates: towards a solution to the purchasing power parity puzzle. *Int Econ Rev* 42:1015–1042
- Triki MB, Maktouf S (2015) Purchasing power parity as a long-term memory process: evidence from some emerging countries. *Int J Emerg Mark* 10:711–725
- Tsay W-J (2000) Long memory story of the real interest rate. *Econ Lett* 67:325–330
- Watson MW (1986) Univariate detrending methods with stochastic trends. *J Monet Econ* 18:49–75
- Wu J-L, Chen S-L (1998) A Re-examination of real interest rate parity. *Can J Econ* 31:837–851

- Wu J-L, Fountas S (2000) 'Real interest rate parity under regime shifts and implication for monetary policy. *Manch Sch* 68:685–700
- Zinkina J, Christian D, Grinin L, Ilyin I, Andreev A, Aleshkovski I, Shulgin S (2019) The first “Golden Age” of globalization (1870–1914). A big history of globalization. World-systems evolution and global futures. Springer, Heidelberg, pp 195–224

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