The long memory story of ex post real interest rates. Can it be supported?

Ioannis A. Venetis[a,1], Agustin Duarte[b], Ivan Paya[c,*]

[a]Centre of Planning and Economic Research (KEPE), Hippokratous 22, 106 80 Athens, Greece

[b]Departamento Economia Aplicada y Politica Economica, University of Alicante, 03080 Alicante

[c]Departamento Fundamento Analisis Economico, University of Alicante, 03080 Alicante

April 2004

Abstract

This papers finds evidence of fractional integration for a number of monthly ex post real interest rate series using the GPH semiparametric estimator on data from fourteen European countries and the US. However, we pose empirical questions on certain time series requirements that emerge from fractional integration and we find that they do not hold pointing to "spurious" long memory and casting doubts with respect to the theoretical origins of long memory in our sample. Common stochastic trends expressed as the sum of stationary past errors do not seem appropriate as an explanation of real interest rate covariation.

Keywords: Real interest rate; Long memory, Fractional Integration

JEL classification: C22, E40, F41.

[1]A large part of this work was conducted when the author was a research associate in Cardiff Business School. The author gratefully acknowledges financial support from ESRC grant *L/138/25/1004*.

[*]**Corresponding author.** E-mail address: ivanpaya@merlin.fae.ua.es. Tel: +34 965 909303

1 Introduction

The relationship between real interest rates across countries is of central importance to the understanding of open macroeconomic models and to the application of economic policies. Real interest rate parity (RIP) is an essential assumption in most open-macroeconomic models. In few words, this assumption states that rates of interest for similar assets in two different countries must be equal once they have been adjusted by their respective expected inflation rates. The policy implication of this assumption is straightforward. In a context where goods and capitals flow freely and real interest rates are settle in the international markets, individual countries will find their scope for stabilization policies very limited. In other words, the scope of economic policies over real economic variables depends to a great extent on the degree to which international real interest rates can influence domestic monetary policy.

Empirical investigation on real interest rates equalization does not yield a clear-cut conclusion. Early studies (see e.g. Mark, 1985; Cumby and Mishkin, 1986; Fraser and Taylor, 1990; Dutton, 1993 and Edison and Pauls, 1993) mostly rejected the real interest rate hypothesis using regression analysis. More recent attempts include the application of cointegration techniques although the results are also inconclusive. Some studies find little evidence in favor of parity (see Throop, 1994), while others find positive results for RIP (see Goodwin and Grennes, 1994; Fountas and Wu, 1999). Additional recent research using panel estimations find increasing evidence that the real rate hypothesis could hold for most of western developed countries (see, e.g. Gagnon and Unferth, 1995; Wu and Chen, 2001). In a different set-up Evans and Lewis (1995) allow the data to follow a non-linear process and their results are supportive of the parity relationship. In sum, from the 1980's, empirical evidence is showing a change in trend from less to more supportive tests on RIP. These results may reflect, on the one hand, the evolution over the last twenty years towards a more integrated international financial market, and, on the other hand, the implementation of new developments in econometrics.

A preliminary step in assessing the most appropriate technique to test the hypothesis of real interest rate parity (RIP) is to examine whether real interest rates are stationary or not. Now, it is widely accepted that classical regression techniques may become invalid if applied to non-stationary variables. More recently, it has become standard practice to pursue different modelling strategies when real interest rates are either stationary or non-stationary. For instance, stationary real interest rates can be best modelled in levels, while first differences are strongly recommended when interest rates are non-stationary. Also, in the case that real interest rates are stationary, as economic theory would predict[2], the application of cointegration technique might be misleading.

Testing for stationarity of real interest rates is essential to explore the proposition that real rates are equal

[2]The prediction that real interest rates are stationary is consistent with Lucas-type-consumption-based asset pricing models. See Lucas (1978) and Hansen and Singleton, (1983).

across countries. The familiar Fisher (1930) equation, which postulates a rationale for the long-run relationship between nominal interest rates and expected inflation, is usually the link between this proposition and its empirical application. An essential requirement for this long-run relationship to hold is that ex ante real rate of interest -that is, the difference between the nominal rate and expected inflation- should be mean reverting. However, empirical evidence does not give much support to the mean reverting property of real interest rates. In the literature some studies find evidence on existence of unit roots (Rose, 1988). Using cointegration technique, some researchers have pointed out that the nominal rate and realized inflation are non-stationary processes and cointegrated. However, the estimated slope coefficients are considerable different from one, as economic theory would require (Hodrick, 1987; Mishkin, 1992; McCallum, 1994).

Given that, it has been widely acknowledged that unit root tests have low power to distinguish between near unit root and unit root processes, our intention in this paper is to use new techniques to test for stationarity of real interest rates. We aim to provide international evidence, thus we employ monthly ex post real interest rates for fourteen European countries and the United States from the mid-1970 till early 1999. In particular, we will allow the possibility of a fractional unit root to form the empirical basis of the real interest parity hypothesis. A clear advantage of fractional integration analysis over conventional unit root tests is that the former permits a wider range of mean-reverting behavior. We concentrate on two related issues. One issue is whether real interest rates are best described as fractionally integrated processes. The second issue directly originates from the restrictions imposed on long run co-behavior of series that possess fractional unit roots. Moreover, the second issue is directly related to the question of convergence of real interest rates (real rate equalization). According to the literature, there are a number of reasons to explain departures from the long-run constant equilibrium value of RIP. For example, price stickiness in good markets (Dornbusch, 1976; Mussa 1984), existence of time-varying risk premium in foreign exchange markets (Domowitz and Hakkio, 1985; Niewland, et al., 1998), and policy behavior (McCallum, 1994; Christensen, 2000). Hence, the response of economies to common shocks may differ and the equalization of real rates is not instantaneous but takes time to adjust following idiosynchratic time lag adjustment structures. Indeed, although we find strong evidence that real interest rates are mean reverting long memory processes we also find that bilateral real interest rate differentials would become stationary once a certain time lag adjustment is taken into account. This is true even though the fractional orders of integration for each rate are different.

The rest of the paper is structured as follows. Section 2 briefly presents some preliminary concepts on RIP and discusses recent evidence of long memory in real interest rates. We then make an asumption regarding time lag adjustment in RIP and the restrictions imposed by the presence of long memory. We finish the section with a brief presentation of the semiparametric estimators of the long memory parameter that we will adopt in our empirical testing. Section 3 reports the estimates of the long memory parameter for all individual series as well as for a large number of contemporaneous or lagged pairwise differences and discusses our findings.

Section 4 concludes.

2 Estimation Methodology

2.1 Preliminary concepts

The Fisher equation has been extensively used as the link between the long-run behavior of real interest rates and the empirical application. Formally,

$$r_t = i_t - \pi_t^e \tag{1}$$

$$r_t^* = i_t^* - \pi_t^{*e} \tag{2}$$

where i_t denotes the nominal interest rate, r_t the real interest rate, and π_t^e is expected inflation; asterisks denote foreign variables. Equation (1) and (2) simply define the ex ante real interest rate - as the nominal rate less the expected inflation rate over the same period. Two major approaches have been predominantly used to check the Fisher equations. One is to test for unit root in r_t, and the other is to explore cointegration between i_t and π_t^e.

Theory predicts that domestic and foreign real interest rates should equal. This equality of domestic and foreign ex ante real interest rates is obtained from two parity conditions. The uncovered interest parity (UIP) representing market equilibrium in capital markets, and ex ante purchasing power parity (PPP) representing the equilibrium in international goods markets. According to this

$$r_t = r_t^* \tag{3}$$

Equation (3) is the real interest parity condition (RIP), which says that in the long-run, for this condition to hold, real interest rate differential should be equal to zero. In empirical terms, RIP would hold in its strictest form if real interest rate differential is a stationary process, that is, an I (d) process with $d = 0$[3]. Again, two approaches have been mostly taken to test condition (3). One is to examine the order of integration of the real interest rate differential. The second one looks for cointegration between the real rates.

Standard unit root tests on real interest rates are often ambiguous in the sense that some (or sometimes) rates appear to be non-stationary or borderline stationary or when treating a large number of countries a subset

[3] If the nominal exchange rate is integrated of order one $I(1)$, then the uncovered interest parity condition would imply that the nominal interest rate differential is stationary. If we assume that inflation is a stationary process, then the stationarity of the real interest rate differential follows by definition.

of them is found to be nonstationary (see Kugler and Neusser, 1993; Pain and Thomas, 1997; Fountas and Wu, 1999). On theoretical grounds, it is difficult to reconcile non-stationarity with real interest rates (see e.g. Rose, 1988; Garcia and Perron, 1996), and on empirical grounds, unit root tests are found to have low power against either long memory alternatives (see, e.g. Diebold and Rudebusch, 1991) or deterministic trends with breaks (see e.g. Perron, 1989; Zivot and Andrews, 1992). On the basis of these results, the theoretical and empirical validity of unit root tests on real interest rate series appears to be rather distant, casting doubts on the practical use of cointegration methods involving ex post real interest rates[4].

Recently, Lai (1997) and Phillips (1998) provide evidence, based on semiparametric estimators, that ex ante and ex post U.S real interest rates are fractionally integrated. Tsay (2000) employs an ARFIMA model and provides further evidence that the U.S real interest rate can be described as an $I(d)$ process. The fractional unit root can reconcile the statistical "discontinuity" associated with the stringent $I(0)/I(1)$ dichotomy.

2.2 Time lag adjustment

Let aside the stationarity (or not) of some rates, different fractional integration orders create complexities that in bivariate analysis could be insurmountable. A necessary condition for fractional cointegration is that at least two elements of the underlying vector series share the same maximal order of integration (see Davidson, 2001). Given the difficulty to theoretically justify long memory in real interest rates, let us assume that the presence of long memory in our series is spurious and attributable to either structural breaks or cyclical components. Also let us assume that (a) some of the series share the same unobserved component(s) responsible for the observed real rate dynamics and (b) there may exist time or phase disparities in these unobserved components. Time disparities correspond to the existence of structural breaks at separate dates whereas phase disparities correspond to underlying cycles being out-of-phase for different countries. Then, we expect that if y_t appears to be $I(\hat{d}_y)$ and x_t appears to be $I(\hat{d}_x)$ there would be a certain lag l such that $(x_t - y_{t-l}) \sim I(a)$, $a = \min\{d_y, d_x\} - b$, $b > 0$. When $a = 0$ (short memory stationarity) then at the particular lag l there is full elimination of time or phase disparities of the components while for $l = 0$ the components are synchronous. The distinction with the known cointegration paradigm is that (i) $d_y \neq d_x$ and (ii) the reduction of the maximal order of intergration to a (or $b > 0$) is materialized only at certain lag(s) l. Such behavior is not justifiable if the underlying series where truly generated as $I(d)$ processes. Recently, Levy (2002) has examined the frequency domain implications of cointegration between two $I(1)$ time series in terms of the zero-frequency squared coherence, phase, and gain. Nielsen (2004) generalizes to the $I(d)$ case. Both authors show that the phase-shift (the frequency domain equivalent of time-delay) between them will equal zero.

[4]There is a large number of studies dealing with real interest parity or with real interest rate equalization across countries. The vast majority of them is based on cointegration techniques involving $I(1)$ variables.

2.3 Semiparametric estimators

In this section we briefly describe the semiparametric estimators that will be employed in subsequent empirical analysis. The Geweke and Porter-Hudak (GPH, 1983) estimate of d is simply

$$\hat{d} = \frac{\sum_{j=1}^{m} (x_j - \bar{x}) \log I_{T,j}}{\sum_{j=1}^{m} (x_j - \bar{x})^2} \tag{4}$$

where $I_j = \frac{1}{2\pi T} \left| \sum_{t=1}^{T} y_t e^{i\lambda_j t} \right|^2$, $j = 1, ..., m$, $i = \sqrt{-1}$, denotes the first m periodogram ordinates, $\lambda_j = 2\pi j/T$ is the j^{th} Fourier frequency, m is an integer less than T such that $m = o(T^{4/5})$, $x_j = -2\log\lambda_j$. $d > 0$ indicates long memory, $0.5 \leq d < 1$ indicates nonstationarity but mean reversion in the sense that movements of the underlying series y_t are independent of initial conditions. In the case where y_t is described by an $ARFIMA(p, d, q)$ model, shocks will not infinitely persist. Robinson (1995) showed that under appropriate regularity conditions $m^{1/2}(\hat{d} - d) \longrightarrow_d N(0, \pi^2/24)$ and also that the standard error of $\hat{d}$ can be consistently estimated by the residual standard error[5]. The semiparametric nature of the estimator becomes apparent since exact knowledge of the underlying short run dynamics is not required. One of the main advantages of the GPH procedure is the simplicity and consistency of the estimator. Moreover, it is obtained without restrictions on the distribution of the data generating mechanism. However, there is a trade off between bias and variance when choosing m, the number of periodogram ordinates to be employed in (4). GPH used $T^{1/2}$ but Hurvich et al. (1998) showed that $m = o(T^{4/5})$ is mean square optimal. Thus, in all our following applications, we employ $m = [T^{0.7}]$.

A modification to the standard GPH estimator can be obtained if we use the tapered periodogram instead of the standard one. The tapered[6] GPH estimate employs $\log I_{T,j}^{TA}$ instead of $\log I_{T,j}$ where $I_j^{TA} = \frac{1}{2\pi H} \left| \sum_{t=1}^{T} w_t y_t e^{i\lambda_j t} \right|^2$ with $H = \sum_{t=1}^{T} w_t^2$ and w_t defines the taper function. Hurvich and Ray (1995) and Velasco (1999) showed that the tapered version of GPH gives better results in the case of non-stationary long memory processes with $d > 0.5$ and Velasco (1999) proves consistency and asymptotic normality. Recently, Sibbertsen (2003) provides simulation evidence on the bias reduction power of the $\hat{d}_{TA}$ against non-stationarities that arise from particular types of trends in the data. His results cover slowly decaying trends and a single change in mean model. Sibbertsen and Venetis (2003), SV from now on, develop a statistic based on the difference between $\hat{d}$ and $\hat{d}_{TA}$ that is able to test for the presence of a wide range of trends in the data, including threshold and smooth transition changes in the mean. The test statistic has the form $m^{1/2}(\hat{d} - \hat{d}_{TA})^2$ and its asymptotic distribution is calculated by bootstrap methods since it depends asymptotically on an unknown parameter (the

[5]Notice that (4) is a least squares estimate from an equation relating the periodogram to x_j. See Robinson (1995) for details.

[6]For a detailed discussion on tapering see Bloomfield (1976, p. 80-94).

true d value).

3 Empirical analysis

3.1 The data

Our data set consists of monthly observations from 1973 to 1999 of short run interest rates (3-month money rates) and consumer price indices (CPI: various consumer goods and services, 1995=100) for 14 European countries and United States. The data appendix details the construction of ex post real interest rates and summarizes the countries and dates considered. Figure 1 illustrates the constructed ex post real interest rate series.

3.2 Empirical results

Both the non-tapered and tapered versions of the GPH estimator are applied on our ex-post real interest rate series and the Sibbertsen and Venetis (2003) statistic on their difference is calculated. The results are presented in table 1. None of the series could be characterized as $I(1)$. In accordance with Tsay's (2000) result all series appear persistent enough to deviate from the typical $I(0)$ case but are far from being characterized as random walks. All series are mean reverting since in all cases $\hat{d}, \hat{d}_{TA} \in [0, 1)$. Focusing on the $\hat{d}_{TA}$ estimates, four series (US[7], Ir, Fr and Fi) appear to be non-stationary with $\hat{d}_{TA} > 0.5$ while Germany is borderline stationary. In several cases (Au, Fi, Ger, Nl) the SV test statistic rejects the null hypothesis of equality between d and d_{TA} suggesting the presence of "trends" in the series.

To check whether the data could support both our assumption made in earlier section, we conducted the following experiment. Let x_t and y_t denote the ex post real interest rate for two different countries. Using a) German and U.S rates for y_t (the leading countries), and b) other European countries for x_t, we estimate the fractional order of integration of the rate differential $x_t - y_{t-l}$ for all available pairs and for $l = 0, ..., 24$. Then, we search for estimates of parameter a that are statistically insignificant. In table 2 we report differentials that appeared to be $I(0)$ and the corresponding lag l where the reduction of the order of integration was achieved. The final choice of lag l was based on the minimum a_{TA} estimate although in the majority of cases there was an interval of l satisfying $a = 0$ (Figures 2 and 3 illustrate this observation). The aforementioned procedure was computational straightforward and in table 2 we also report estimates of a that were not necessarily zero but quantitatively much less than $\min\{d_y, d_x\}$.

Some results are noteworthy Although all rates appear to be $I(d)$, with some series falling into the non-stationary but mean reverting zone, the order of integration for their pairwise lagged differences depends crucially on the lag value l. That means that real interest rate differentials exhibit different mean-revertion

[7] Our results regarding the U.S rates are consistent with previous findings by Lai (1997), Phillips (1998) and Tsay (2000).

speeds depending on the lag employed for the "leading" country. Such a behavior is not compatible with $I(d)$ processes. Either the highest order of integration of the two series should prevail or if cointegration between real rates - with nonstationary or stationary regressors[8] - existed, it should hold irrespective of the lag value l (let aside the cases where d estimates are unequal and could be attributed to small sample errors).

Notice that 7 out of 13 European pairs involving the German rate result in $I(0)$ series and the France - Germany pair produces $\hat{a}_{TA} = 0.261$ when the initial d_{TA} estimates for Germany and France produced a minimum order of integration $\min\{\hat{d}_x, \hat{d}_y\} = 0.499$. Similar results were obtained for 7 pairs involving European rates against the U.S rate although it seems that lag l has been increased.

The results of Table 2 create a number of important questions in the area of economic policy. We should recall that convergence of bilateral real rates of interest is intimately connected to the issue of effectiveness of stabilization policies either in the short or the long run. In this sense, Table 2 reports the time lag of the process of adjustment, and in particular, as has been already explained, the methodology used to measure it is based on an estimate of the number of lags necessary to reach stationary real interest rate differentials.

Another result reported in Table 2 is that the speed of adjustment of the European countries is higher with respect to Germany rather than the US. In other words, European countries show that the process of adjustment towards US real rates is relatively slow. However, UK and Portugal display similar time lag adjustment toward either Germany or the US. It is also worth noting that Italy and Spain do not achieved a stationary real interest rate differential with Germany at any lag whilst they do so with respect to the US at a low speed though.

Within this general result, there are some aspects worth mentioning. In the literature, European countries participating in EU have been divided into three groups according to their stance regarding membership of the EU and the associated exchange rate regime[9]. Group 1, is the so-called core EMU group, includes Belgium France and the Netherlands. These countries have always participated in the EMU project from the outset. Group 2 includes Greece, Italy, Spain, Portugal and the UK. All these latter countries joined the EU during the 1970s and 1980s and their currencies have had varied experiences as ERM members. As it is well known, Italy and the UK were ejected from the ERM while Portugal and Spain remained in the ERM with a different band of exchange rate fluctuation. Group 3 incorporates the most recent new entrants to the EU, Austria, Finland and Sweden. On this basis, Group 1 and Group 3 countries show a reasonable speed of adjustment. However, Group 2 displays mixed results[10]. It is also interesting to highlight the fast speed of adjustment between neighboring countries like Belgium and Holland, Ireland and the UK, and, Denmark and Germany[11].

---

[8] See Davidson (2002) and Robinson and Yajima (2002) for a treatment of fractional cointegration with nonstationary or stationary variables respectively.

[9] Similar groups have been identified in the business cycle literature (see e.g. Flaig et al. 2003; Holmes, 2000)

[10] Flaig, et al. (2003) have measured the synchronization of business cycles in the frequency domain in Europe. As a result, they showed that those countries included in our Groups 1 and 2, have higher co-movements with the European business cycle than countries in Group 3.

[11] Similar result have been found in terms of business cycle synchronization. Wynne and Koo (2000) find that long standing

Finally, inspection of Table 2 suggests that there is not a clear conclusion on the influence of US rates over the European ones. This issue is linked to a relevant question, that is, whether the autonomy of European Central Bank decisions over monetary policies could be preserved over time. On this topic, further research is needed.

4 Conclusions

Our findings shed new light for future studies concerning the statistical properties of ex post real interest rates as well as studies on the convergence of real interest rates. Although the rates appear to have long memory there is evidence that this property is spurious with our results corroborating the existence of either structural breaks or "cyclical" movements in ex post real rates that could exhibit significant time or phase differences. Our evidence suggests that none of the 14 monthly European ex post real rates nor the U.S rate for the post-1975 pre-1999 period could be characterized as $I(1)$.

We have also examined the real interest parity condition between major European countries with Germany and the US. Two major conclusions are drawn. First, stationarity of real interest rates differentials is not independent of the time lag structure. Second, most European countries show higher speed of real rates equalization with Germany rather than the US.

members of the EU have highly synchronized cycles. Moreover, the business cycle in large EU countries, and in particular in the UK, also tend to be correlated with the US (see also, Duarte and Holden, 2003).

Data appendix

The following ex post real interest rate was employed in all calculations: $r_t^3 = i_t^3 - \pi_t^3$, where i_t^3 is the short run interest rate (annualized 3-month money market rates) and ex post inflation is constructed as: $\pi_t^3 = 400[\ln p_{t+3} - \ln p_t]$ with p_t the consumer price index (CPI, various consumption good and services, 1985=100). The German rate includes ex-GDR from 1991 onwards. All real rate series were seasonally adjusted after regressing them on monthly dummy variables.

Series: r_t^3. Sample period and number of observations (obs)		
Country (Symbol)	Period	(obs)
Austria (Au)	1980/01 - 1998/12	228
Belgium (Be)	1975/01 - 1998/12	288
Denmark (Dk)	1975/01 - 1999/03	291
Finland (Fi)	1981/01 - 1998/12	216
France (Fr)	1975/01 - 1998/12	288
Germany (Ger)	1975/01 - 1998/12	288
Greece (Gre)	1980/05 - 1999/03	227
Ireland (Ir)	1989/01 - 1998/12	120
Italy (It)	1975/01 - 1998/12	288
Netherlands (Nl)	1975/01 - 1998/12	288
Portugal (Pt)	1975/01 - 1998/12	288
Spain (Sp)	1977/01 - 1998/12	264
Sweden (Sw)	1987/01 - 1999/03	147
U.K (UK)	1975/05 - 1999/03	287
U.S.A (US)	1975/01 - 1999/07	295

References

Bloomfield, P., 1976. Fourier analysis of time series: An introduction. New York: Wiley

Christensen, M., 2000. Uncovered Interest Parity and Policy Behavior: New Evidence. Economics Letters 69, 81-87

Cumby, R. E., Mishkin, F., 1986. The International Linkage of Real Interest Rates: The European-US connection. Journal of International Money and Finance 5, 5-24

Davidson, J., 2001. Convergence to stochastic integrals with fractionally integrated integrator processes: Theory, and application to fractional cointegration analysis. Cardiff Business School Working paper

Davidson, J., 2002. A model of fractional cointegration, and tests for cointegration using the bootstrap. Journal of Econometrics 110, 187-212

Diebold, F.X., Rudebusch, G.D., 1991. On the power of Dickey Fuller tests. Economics Letters 35, 155-160

Domowitz, I., Hakkio, C.S., 1985. Conditional variance and the risk premium in the foreign exchange market. Journal of International Economics, 19, 47-66

Dornbusch, R., 1976. Expectation and Exchange Rate Dynamics. Journal of Political Economy 84, 1161-76

Duarte, A., Holden, K., 2003. The business cycles in the G-7 economies. International Journal of Forecasting 19, 685-700

Dutton, M.M., 1993. Real interest rates parity: New measures and tests. Journal of International Money and Finance 12, 62-67

Edison, H.J.,.Pauls D.B., 1993. A re-assessment of the relationship between real exchange rates and real interest rates: 1974-1990. Journal of Monetary Economics 31, 165-188.

Flaig, G., Sturm, J., Woitek, U., 2003. Synchronization of National Business Cycles in Europe?. Proceedings of the 4th Eurostat colloquium on Modern Tools for Business Cycle Analysis, Luxemburg, October

Fraser, P., Taylor, M.P, 1990. Some Efficient Tests of International Real Interest Rate Parity. Applied Economics. 22 (8), 1083-92

Fountas, S., Wu, J.-L., 1999. Testing for real interest rate convergence in European countries. Scottish Journal of Political Economy 46 (2), 158-174

Gagnon , J.E., Unferth, M.D., 1995. Is there a world real interest rate?. Journal of International Money and Finance 14 (6), 845-855

Garcia, R. and Perron, P., 1996. An analysis of the real interest rate under regime shifts. Review of Economic and Statistics 78 1, 111 125

Geweke, J., Porter-Hudak, S., 1983. The Estimation and Application of Long Memory Time Series Models. Journal of Time Series Analysis 4, 221-238

Goodwin, B.K, Grunnes, T., 1994. Real Interest Rate Equalization and the Integration of International Financial Markets. Journal of International Money and Finance 13, 107-124

Hansen, L.P., Singleton, K.J., 1983. Stochastic consumption, risk aversion and the temporal behavior of asset returns. Journal of Political Economy 91 (2), 249-265

Hodrick, R. J., 1987. The empirircal evidence on the efficiency of forward and futures foreign exchange markets. Harwood Academic Publishers, Chur

Holmes, M. J., 2000. Exchange Rate Policy and Economic Convergence in the European Union. Research Paper 2000/3, Loughborough University

Hurvich, C.M., Ray, B., 1995. Estimation of the memory parameter for nonstationary or noninvertible fractionally integrated processes. Journal of Time Series Analysis 16, 17-42

Hurvich, C.M., Deo, R., Brodsky, J., 1998. The mean squared error of Geweke and Porter-Hudak's estimator of the memory parameter of a long memory time series. Journal of Time Series Analysis 19 (1), 19-46

Kugler, P., Neusser, K., 1993. International real interest rate equalization: A multivariate time-series approach. Journal of Applied Econometrics 8 (2), 163 174

Lai, K.S., 1997. Long term persistence in the real interest rate: Some evidence of a fractional unit root. International Journal of Finance and Economics 2, 225-235

Levy, D., 2002. Cointegration in frequency domain. Journal of Time Series Analysis 23, 333-339

Lucas, R.E., 1978. Asset prices in an exchnage economy. Econometrica 46 (6), 1429-1445

Mark, N.C., 1985. Some evidence on the international equality of real interest rates. Journal of International Money and Finance 4(2), 189-208

McCallum, B. T., 1994. A reconsideration of the uncovered interest parity relationship. Journal of Monetary Economics 33, 105-132

Mussa, M., 1984. The Theory of Exchange Rate Determination. In: Bilson, J.F.O., Marston, R.C., (Eds.) Exchange Rate Theory and Practice. NBER Conference Report, Chicago: Chicago

University Press

Mishkin, F.,1992. Is the Fisher Effect for Real? A Re-examination of the Relationship Between Inflation and Interest Rates. Journal of Monetary Economics 30, 195-215

Nielsen, M.O., 2004. Spectral analysis of fractionally cointegrated systems. Economics Letters 83 (2), 225-231

Nieuwland, F.G.M.C., Verschoor, W.F.C., Wolff, C.C.P., 1998. EMS exchange rate expectation and time-varying risk premia. Economics Letters 60, 351-355

Pain, D., Thomas, R., 1997. Real interest rate linkages: Testing for common trends and cycles. Bank of England Working Paper

Perron, P., 1989. The great crash, the oil price shock and the unit root hypothesis. Econometrica 57, 1361-1401

Phillips, P.C.B., 1998. Econometric Analysis of Fisher's Equation. Yale University, Cowles Foundation Discussion Paper 1180

Robinson, P.M., 1995. Log-periodogram regression of time series with long range dependence. Annals of Statistics, 23, 1048-1072

Robinson, P.M., Yajima, Y., 2002. Determination of cointegrating rank in fractional systems. Journal of Econometrics, 106, 217-241

Rose, A.K., 1988. Is the real interest stable?. Journal of Finance 43 5, 1095-1112

Sibbertsen, P., 2003. Log-Periodogram estimation of the memory parameter of a long-memory process under trend. Statistics and Probability Letters 61, 261-268

Sibbertsen, P., Venetis I.A., 2003. Distinguishing between long-range dependence and deterministic trends. Universitat Dortmund, Technical Report 16/2003, SFB 475.

Throop, A., 1994. International financial market integration and linkages of national interest rates. Federal Reserve Bank of San Francisco, Economic Review, No. 3.

Tsay, W.J., 2000. The long memory story of the real interest rate. Economics Letters, 67, 325-330

Velasco, C., 1999. Non-stationary log-periodogram regression. Journal of Econometrics, 91, 325-371

Wu, J-L., Chen, S-L 2001. Mean reversion of interest rates in the Eurocurrency market. Oxford Bulletin of Economics and Statistics 63 (4), 459-473.

Wynne, M. A. and Koo, J., 2000. Business Cycles under Monetary Union: A comparison of the EU and US. Economica, 67, 347-374

Zivot, E., Andrews, D.W.K., 1992. Further evidence on the great crash, the oil price shock and the unit root hypothesis. Journal of Business and Economic Statistics 10 (3), 251-270

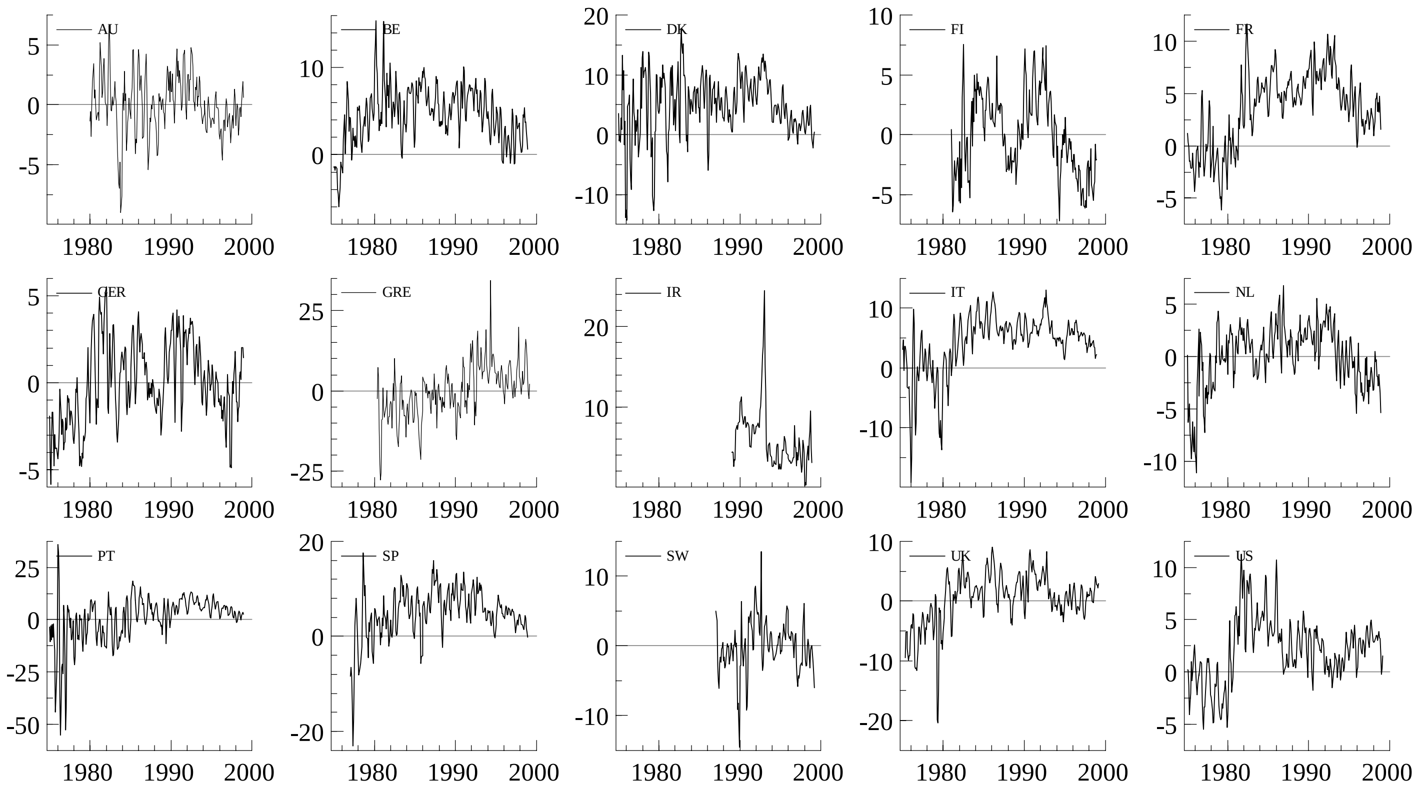

Figure 1: Ex post real interest rates. See data appendix table for details on labels.

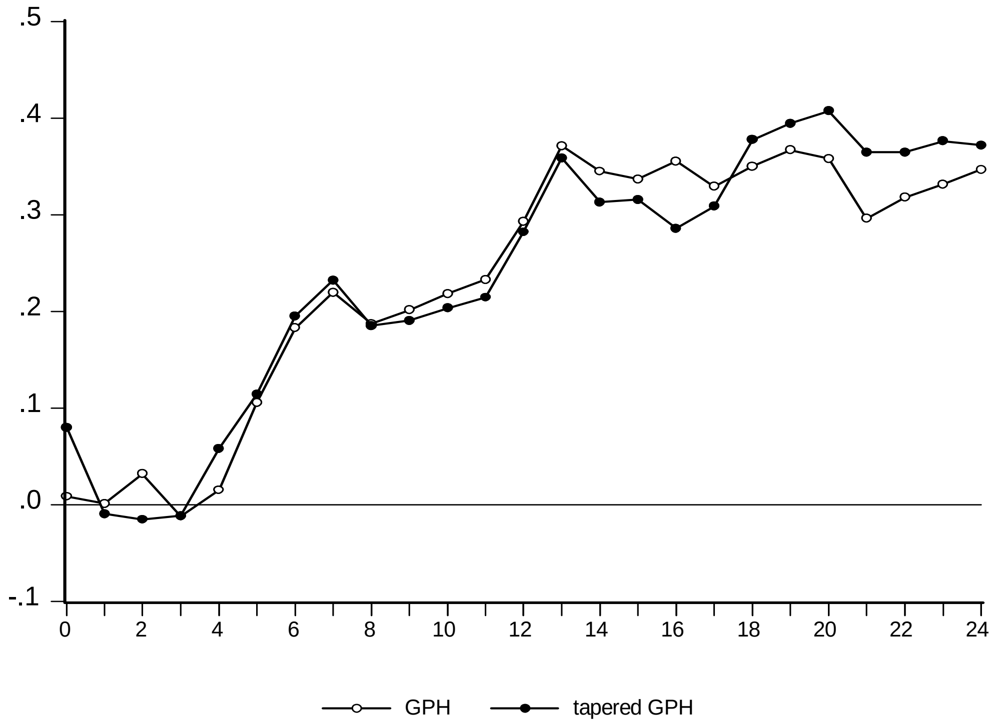

Figure 2: $\hat{d}$ and $\hat{d}_{TA}$ estimates of the constructed series $x_t - y_{t-l}$ where $l = 0, ..., 24$ (horizontal axis) and x_t, y_t denote the Austrian and German ex-post real interest rates respectively. The zero line is also displayed.

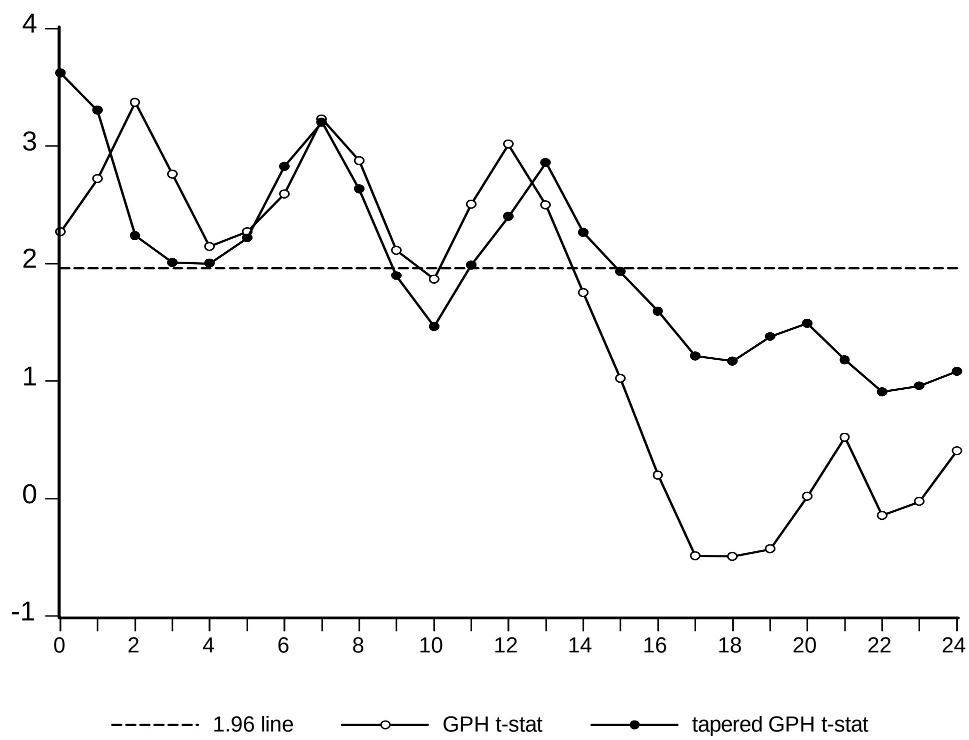

Figure 3: t-statistics for the $\hat{d}$ and $\hat{d}_{TA}$ estimates of the constructed series $x_t - y_{t-l}$ where $l = 0, ..., 24$ (horizontal axis) and x_t, y_t denote the Portuguese and Spanish ex-post real interest rates respectively. The dashed line is the 1.96 line.

Table 1: Semiparametric estimates of the long memory parameter of the ex post real rate series

Real rate	$\hat{d}$	$\hat{d}_{TA}$	$H_0 : d = d_{TA}$	90%	95%	99%
Au	0.695 (1.191)	0.347 (2.716)	**0.80**	0.20	0.32	0.53
Be	0.437 (4.398)	0.307 (2.833)	0.12	0.30	0.40	0.70
Dk	0.280 (3.583)	0.377 (3.827)	0.06	0.26	0.37	0.57
Fi	0.932 (1.585)	0.603 (6.371)	**0.70**	0.22	0.32	0.57
Fr	0.452 (4.250)	0.501 (5.454)	0.01	0.32	0.45	0.72
Ger	0.846 (1.751)	0.499 (5.560)	**0.86**	0.20	0.28	0.54
Gre	0.397 (3.452)	0.368 (3.728)	0.00	0.22	0.31	0.54
Ir	0.708 (7.055)	0.797 (7.361)	0.04	0.32	0.44	0.77
It	0.388 (4.259)	0.480 (4.626)	0.06	0.31	0.42	0.66
Nl	0.954 (1.967)	0.349 (2.662)	**2.64**	0.20	0.28	0.49
Pt	0.076 (0.784)	0.209 (2.023)	0.12	0.18	0.27	0.49
Sp	0.475 (4.271)	0.327 (2.868)	0.15	0.23	0.32	0.56
Sw	0.242 (2.516)	0.274 (2.273)	0.00	0.28	0.41	0.85
UK	0.505 (6.452)	0.413 (4.862)	0.06	0.19	0.27	0.54
US	0.526 (6.171)	0.580 (5.250)	0.02	0.24	0.33	0.61

Notes: Country codes are explained in the data appendix table. t statistics are reported in parentheses. $\hat{d}$ refers to the GPH estimate and $\hat{d}_{TA}$ to the GPH tapered estimate.

Table 2: Semiparametric estimates of the long memory parameter for the constructed series $x_t - y_{t-l}$

Leading country: Ger

x_t	y_{t-l}	$\hat{a}$	$\hat{a}_{TA}$	lag l
Au	Ger	0.001 (0.010)	-0.009 (-0.007)	1
Be	Ger	0.112 (1.104)	0.037 (0.364)	2
Dk	Ger	0.122 (1.129)	0.106 (0.980)	2
Fr	Ger	**0.203** (1.961)	**0.261** (2.720)	10
Nl	Ger	**0.321** (2.178)	0.188 (1.797)	3
Pt	Ger	0.119 (1.382)	0.117 (0.998)	19
Sw	Ger	0.089 (0.862)	0.005 (0.042)	7
Uk	Ger	**0.232** (2.744)	0.117 (0.891)	4

Other European pairs

x_t	y_{t-l}	$\hat{a}$	$\hat{a}_{TA}$	lag
Be	Nl	-0.005 (-0.054)	0.015 (0.146)	1
Ir	U.K	0.163 (1.422)	0.199 (1.760)	6

Leading Country: USA

x_t	y_{t-l}	$\hat{a}$	$\hat{a}_{TA}$	lag
Dk	U.S	0.009 (0.093)	0.083 (0.686)	19
Fr	U.S	**0.278** (1.994)	0.227 (1.816)	14
It	U.S	0.137 (1.345)	0.213 (1.617)	18
Pt	U.S	0.046 (0.440)	0.083 (0.618)	3
Sp	U.S	**0.348** (3.501)	0.175 (1.569)	20
Sw	U.S	0.207 (1.447)	0.159 (1.319)	23
Uk	U.S	**0.208** (2.710)	0.038 (0.389)	5

Notes : GPH, $\hat{a}$, and tapered GPH, $\hat{a}_{TA}$, estimates for the constructed series $x_t - y_{t-l}$. t-statistics are reported in parentheses. Bolded values denote statistical significance at the 5% level. Under the "lag l" column we report the value of l that produced the minimum $\hat{a}_{TA}$ estimates.