A story of PPP persistence. The Spanish Peseta (1870-1998) against the Dollar and the Pound

María Dolores Gadea
Marcela Sabaté
José María Serrano
University of Zaragoza (Spain)

Correspondence
MARCELA SABATE
Departamento de Economía Aplicada
Facultad de Económicas
Gran Vía, 4
50005 ZARAGOZA
SPAIN
Tel.: 00 34 976761848 - Fax:00 34 976761840
e-mail:* *msabate@unizar.es

Abstract

This paper models the peseta/dollar and the peseta/pound real exchange rates (1870-1998) as two fractionally integrated processes. The evidence of mean-reversion supports the idea of the PPP as a long-run rule, despite the dominance of floating exchange rates in the series. As regards their reversion, real-sided factors cannot explain the long-lived deviations from parity. Conversely, the half-life deviations can be dramatically reduced if considering the peseta black market exchange rates during the Spanish autarky. We argue that the reduction obtained when using free black rates instead of the controlled ones proves the responsibility of adjustment costs, as suggested by Rogoff (1996), in the PPP puzzle.

Classification JEL: F31, N10, C22

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1. Introduction

This paper analyses the performance of the real exchange rates of the peseta against the dollar and the pound from the setting of the peseta as the Spanish national currency in 1870 to its disappearance into the Euro in 1998. Thus, this is a still new study within the kind of *long-horizon data sets* to which Rogoff (1996) referred when proposed his celebrated PPP puzzle. Then, he recognized to the procedure of using long span of data –as had done in Froot and Rogoff (1995)-, the advantage of increasing the power of unit root tests and, consequently, of better capturing the mean-reverting property of real exchange rates. In fact, the long-horizon studies surveyed by Rogoff (1996) provided generous evidence of mean–reverting performance for several currency real exchange rates, showing a very *remarkable consensus* on 3-5 year half-life deviations from PPP. But despite this evidence, two problems remained. First, studies of long span of data mixed years of fixed and floating exchange rates, with a dominant presence of the former. Thus, mean reversion might be simply reflecting the temporal dominance of the less variable fixed rates in the sample, what would impede the acceptance of the PPP as a general rule, specifically valid for periods of floating rates, like the opened in 1971 on which, precisely, focussed the PPP discussion at the moment. Beyond this problem, there was the issue of the coincident 3-5 year half-life deviations in the historical studies, that were too long to be explained by monetary factors, but found equally difficult to be explained by real factors, supply or demand sided. The end is well-known: Rogoff (1996) finally proposed to take adjustment costs into account as the only way to solve the Persistence puzzle.

In this framework, the analysis of the Spanish peseta (1870-1998) emerges with great possibilities. First, because the presence of floating exchange rates, as explained in section 2, is much longer than for currencies that belonged to the gold standard. The peseta never formally joined the gold standard and, in relation to the European Monetary System (EMS), the peseta was only tied to the pound from 1990 to September 1992. Moreover, the performance path of the Spanish economy along that century combined periods of cumulative backwardness with others of dramatic catching-up, as well as periods of external openness with others of even autarky. So, by taking the UK and the US as developed reference countries, the Spanish offers itself as a natural case of study to capture the potential effects on exchange rates of economic real and institutional factors.

To exploit this possibility, we analyse in section 3 the mean-reverting properties of the peseta/dollar and the peseta/pound real exchange rates during 1870-1998. Tests of stationarity and unit roots are carried out and half-lives of PPP deviations are calculated according to a median-unbiased procedure, in order to overcome the persistence underbias of least squares estimates. In our case, the results of applying stationarity and unit root tests are inconclusive, and furthermore, the length of parity deviations (of around 6-7 years) is clearly above the consensus range (3-5 years) referred in Rogoff (1996). Then, given the ambiguous test results plus the presence of very high long-lived deviations with wide confidence intervals, we wonder if *subtle forms of reversion* to parity –in the words of Diebold, Husted and Rush (1991)-, might be operative. Their idea was that the integer integrated techniques can be a quite crude method to model the dynamics of series with slow reversion, in that the power of standard unit root tests would be in this case seriously reduced. We study this possibility by using fractionally integrated analysis and find that both, the peseta-dollar and the peseta-pound exchange rates can be clearly characterized, even if not stationary, at least as mean-reverting series. That would be enough to speak about the PPP as an *anchor* (Rogoff ,1996) or a *good approximation* (Lothian and Taylor, 1996) for the long-run behaviour of the peseta real exchange rates.

Still, the issue of the persistence in parity deviations, even stressed when modelling with fractional methods, does not allows us to forget the PPP puzzle. For this reason, we study in section 4 if the persistence can be explained by real-side factors and its influence on the relative price of traded and non traded goods. On the supply side, long-lived deviations from PPP, might just reflect the succession of backwardness and catching-up episodes experienced by the Spanish economy along the years 1870-1998. On the demand side, the slow adjustment of the peseta real exchange rate to its parity might respond to the differences between the path of Spain's government expenditure and those of the USA and the UK along the period 1870-1998. We test for both possibilities in a cointegration framework and find neither supporting evidence for the Balassa-Samuelson effect nor for the influence of relative public expenditure on the real exchange behaviour. In other words, the PPP persistence remains, leading us to considerate that the puzzle might have not be born from the interference of real factors, but from market barriers, as the own Rogoff (1996) suggested.

Such an issue is considered in section 5, where we use impulse response functions obtained in a fractional framework to calculate the half-life deviations recursively, by enlarging yearly the initial sub-sample 1870-1900 until the whole period 1870-1998. As main result, we find that the longest deviations concentrate for both series, the dollar and the pound real exchange rates during the forties-fifties, when not surprisingly, strong interventions dominated the Spanish economy. From 1939 to 1959, during the so-called Franco's autarky, Spain reached the lowest levels of good market integration. Furthermore, a strict exchange rate control system was in force, the domestic price dynamics not having been taken into account in the few official devaluations that then took place. Interestingly, the availability of black market rates for the peseta/dollar and the peseta/pound during these years allows us to approximate the weight of the controls on the persistence exhibited by both series, since by considering the free black rates instead of the controlled ones, the half-life deviations of the whole sample 1870-1998 falls dramatically. This way, the paper would prove the responsibility of the market barriers suggested by Rogoff (1996) in the persistent deviations of the peseta/dollar and the peseta/pound real exchange rates from parity. This finding centres the concluding section that close the paper.

2.Monetary historical background and data

As mentioned earlier, the analysis of the PPP hypothesis is applied to the peseta/ dollar and the peseta/ pound exchange rate series covering the years 1870 to 1998. Thus, we are considering the period that runs from the generalization of the gold standard as the international financial system in the last quarter of the nineteenth century to the beginning of the third stage of Economic and Monetary Union (EMU), when the Spanish peseta lost its domestic exchange rate identity. An interesting point here is that the peseta did not maintain for long any exchange rate commitment with either the dollar or the pound.

In fact, until its entry into the EMU in 1998, Spain's involvement in international monetary projects came very late in the day, when at all. For example, the peseta was never formally a part of the gold standard; neither of the classical gold standard (up to 1913), nor of the revived gold exchange standard (in the post-First World War period). Afterwards, for two decades, during the forties and fifties of the twentieth century, Spain maintained strict exchange rate controls, the peseta being a late entrant, in 1961, in the international financial order created at Bretton Woods. Finally, the Spanish currency did not form part of the European Monetary System (EMS) until 1989, whereas the pound spent just two years inside the EMS, from 1990 to 1992. Thus, there is no doubt that autonomy was the dominant feature of the peseta monetary history against the dollar and the pound.

With regard to the data, Figures 1 and 2 show the evolution of the peseta real exchange rate against the dollar and pound, the series constructed by correcting the nominal exchange rate by relative prices (Spanish over US or British prices). Spanish and foreign price indexes are GDP deflators and until 1970 come from Prados (2003) and from Mitchell (2003a,b), respectively. From 1970 onwards, all the deflators come from OECD. As far as nominal exchange rate is concerned, note that for both currencies Figures 1 and 2 present two series, the differences lying in the years of Franco's autarky.

As we said, after the Spanish Civil War (1936-1939), exchange rate controls extended for two decades, until 1959. Between 1939 and 1947 the official exchange rate applied by the regulating body (*Instituto Español de Moneda Extranjera*, IEME) was 10.95 pesetas/dollar. This regime remained unaltered until 1948, when a multiple exchange rate system was introduced with 9 import rates and 15 export rates. The new rates, all revised in a depreciating direction, remained unchanged until 1951, when the strangulation of the balance of payments imposed another revision. This latter year all the rates were again devaluated and, moreover, exporters were allowed to negotiate a determined percentage of the foreign currency (with this percentage depending on the product exported) on the Madrid stock market, where higher rates per dollar were paid. The following revision did not come until the devaluation of 1957, when the exchange rate was unified at 42 pesetas/dollar[1]. Series q^1 would reflect the path of the peseta/dollar and peseta/pound real exchange rates calculated by Serrano and Asensio (1997) by taking into account the different official rates applied to different commodity transactions and also considering the part of each operation negotiated freely on the domestic stock market until 1957. This latter year a new devaluation, aimed at rationalizing the exchange rate policy, unified the official rate for any kind of transactions by devaluating until 42 pesetas/dollar and 121 pesetas/pound. Nevertheless, this devaluation soon proved incapable of containing the pressure exerted by the importers and incapable of stimulating exports. For this reason, a system of surcharges and refunds came into operation, supposing, in practice, the maintenance of the mentioned system of multiple exchange rates. In any event, this devaluation could not avoid the situation whereby, two years later, in June 1959, Spain was on the verge of declaring an international suspension of payments. It was precisely at that time, under the threat of an external crisis, that General Franco's regime fixed an exchange rate of 60 pesetas/dollar and 168 pesetas/pound, approaching to the rate that the peseta was averaging on the free markets of Tangier, New York and Zurich in 1959. The evolution of black market real exchange rates from 1940 on is shown by the series q^2, to which we will refer again in section 5. Until then, the analysis will focus on the exchange rate actually in force for the Spanish economic agents, that is to say, on the series q^2.

With regards to the sources, nominal exchange rates come from Martín Aceña (1989) for the period 1870-1935 and from Eguidazu (1978) for the years 1936-1939.

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[1] Reflecting its changes against the dollar, the minimum official rate for the pound was 40 pesetas until 1943 and 44 pesetas/pound until 1948. Afterwards, the minimum rate was around 30 pesetas/pound till the devaluation of 1957, which set the rate at 117pesetas/pound.

The data for the period 1940-1959 come from Serrano and Asensio (1997) for the series q^1 and from Eguidazu (1978) and Ros Hombravella et al. (1978) for the series q^2. Since 1959, the data of nominal exchange rates can be found in the *Boletín Estadístico* (Statistic Bulletin) from the *Banco de España*.

3. Individual analysis of real exchange rates

Under PPP, the nominal exchange rate of a currency equals the ratio of domestic and foreign price levels. In its logarithmic form, the PPP hypothesis may be written:

$$e_t = p_t - p^*_t \qquad [1]$$

e denoting nominal exchange rate (pesetas per dollar or pound), *p* are the domestic (Spanish) prices and *p** the foreign prices. Then, the real exchange rate being expressed as:

$$q_t = e_t - p_t + p^*_t \qquad [2]$$

when testing for the presence of unit roots or stationarity in q_t series, we are testing for deviations from or adjustment to the PPP value.

The results of applying different unit root tests [the ADF of Dickey and Fuller (1981), the PP of Phillips-Perron (1988) and the MZ-GLS of Ng and Perron (2001)] and the KPSS test of stationarity of Kwiatkowski et. al (1992), are presented in Tables 1 and 2 for the dollar and the pound, respectively[2]. For both currencies the analysis of unit roots and stationarity show ambiguous results, which depend on the test or significance level selected. Thus, we cannot conclusively reject the null of stationarity nor the null of a unit root. Moreover, when measuring the persistence of deviations from parity we find that such deviations have very high half-lives.

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2 These Tables also present the results of analysing the integration order of the variables included in the different models. In no case can we reject the existence of a unit root in the series.

As it is well-known, the half-life of deviations from the PPP parity refers to the number of periods for deviations to be corrected by one half. One generalized way to estimate half-lives started from the habitual Dickey-Fuller expression, that in its Augmented (ADF) form is:

$$q_t = \mu + \alpha q_{t-1} + \sum_{i=1}^{k} \psi_i \Delta q_{t-i} + \varepsilon_t \qquad [3]$$

where α is the autoregressive or parameter associated with persistence. Then, the Half-Life (*HL*) was calculated as $HL = \frac{\ln(0.5)}{\ln(\hat{\alpha})}$ [4].

Despite its popularity, this estimating method of persistence has some problems as highlighted in Murray and Papell (2002). First, the usual least square estimation of a parameter exhibits down ward bias in finite samples. Second, if the order of the AutoRegressive (AR) is superior to 1, then the calculus of half-life from the expression [4] is not appropriated. To overcome the first problem, Andrews and Chen (1994) propose the approximately median-unbiased method to estimate the AR parameters in ADF regressions, with the additional advantage of offering asymptotic confidence intervals valid for the AR(1) or AR (p) cases. The second problem can be overcome by calculating the half-lives directly from the Impulse Response Function (*IRF*), as recommended by Andrews and Chen (1993) themselves.

The results of estimating the persistence parameter α by both procedures, least square and median-unbiased methods, are displayed in Table 3, where the values of half-lives calculated in different ways are also reported. By using the standard least-square estimation half-lives are of 5.53 years for the dollar and 4.82 for the pound when calculated from the α parameter; of 6.02 and 5.31 years respectively, when calculated from the IRF. However, if we apply the median-unbiased method, the half-life increases to 7.17 and 6.02 with the α parameter, and to 7.94 and 6.60 with the *IRF*, pushing the peseta real exchange rates far away from the 3-5 year half-life *consensus* of Rogoff (1996). Furthermore, with this procedure, the confidence intervals of the half-lives are very wide. The upper bound is about 30 years for the pound and ∞ for the dollar, which in practice provides no useful information on the point estimates, being even consistent, in the dollar case, with the presence of a unit root in the series.

In short, we cannot always reject the stationarity nor the unit root hypotheses. Moreover, when rejected, the use of the median-unbiased method increases the half-life of deviations, at the time that widen the confidence intervals. Put together, the inconclusive test results and the presence of high long-lived deviations suggest that the integer I(0)-I(1) approach might not be *subtle* enough to model the peseta real exchange rate behavior[3]. Thus, if it seems to have been the case the peseta exhibited a very slow reversion to parity, then, the integer tests I(0) versus I(1) could not capture this reversion, offering instead the ambiguous results that we obtained. To check this possibility we decided to apply a more flexible paradigm, the so-called long-memory or AutoRegressive Fractionally Integrated Moving Average (ARFIMA) models[4].

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3 A possibility also supported when applying to the real exchange rate the Hodrick-Prescott filter. The series exhibit the typical path of a long-memory process, with fluctuations especially pronounced in some periods.

4 Although most empirical work over PPP is based on the integer differencing paradigm, recent papers has showed that the fractional integration offers a more realistic framework for the modelling of the exchange rate behaviour. Examples of fractionally integrated evidence in favour of the PPP are those of Diebold, Husted and Rush (1991) for the gold standard; Cheung and Lai (1993) for the period 1914-1989; Chou and Shih (1997) for Asian counties in 1965-1992; Cheung and Lai (2000) for developing countries in 1973-1994; Achy (2003) for middle income countries in 1973-1998, and Holmes (2002) for less developed countries in 1973-2001.

The ARFIMA models extend the dichotomy I(1) versus I(0) and permit stationary and non-persistence alternatives if real exchange rate is a I(d) process with $d \in [0,1/2]$, or non-stationary but non-persistence alternatives if $d \in [1/2,1]$. In any case, a shock does not persist indefinitely but disappears, giving the series its mean-reverting behaviour. Thus, an ARFIMA model (p, d, q) can be defined as follows:

$$\phi(L)(1-L)^d (y_t - \mu) = \theta(L)\varepsilon_t \qquad [5]$$

where $\phi(L) = 1 - \sum_{j=1}^{p} \phi_j L^j$ and $\theta(L) = 1 + \sum_{j=1}^{q} \theta_j L^j$ are polynomials of lags of order p and q respectively, whose roots lie outside the unit circle and u_t is iid (0, σ^2). By contrast with a covariance-stationary and short memory process, whose autocorrelation function $\sum_{j=-\infty}^{\infty} |\rho(j)| < \infty$ is absolutely summable and decays at an exponential rate $\rho_j \approx c^j$ (with *c* constant and $|c| < 1$), if *d*>0 the autocorrelation functions decay at a slower hyperbolic rate $\rho_j \approx j^{(2d-1)}$. This implies that $\sum_{j=-\infty}^{\infty} |\rho(j)| = \infty$, and the spectral density function will not be bounded in the zero frequency. If $0 < d < 0.5$, the series is stationary with finite variance and long memory; if $0.5 \leq d < 1$ the series is not stationary, with infinite variance and permanent memory, but registering a mean reversion; finally, if $d \geq 1$ the series does not revert to its mean[5].

From among the different methods used to estimate ARFIMA models and the parameter *d*, we decided to use those of Geweke and Porter-Hudak (GPH) (1983) and the Gaussian of Robinson (GSP) (1995) estimates, both of the semiparametric type in the frequency domain, and the exact maximum likelihood method in a full parametric

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5 A complete review of the concepts of fractional integration in economic series can be found in Baillie (1996).

approach developed by Sowel (MLE) (1992). All methods have specific problems and bias; so, the use of various is a guarantee for robustness[6].

The results of applying these methods to estimate the memory parameter appear in Table 4. By using non-parametric methods, when τ=0.5, 0.6 or 0.7 for GPH or with the equivalent truncation lags for GPS, the long-memory parameter d lies between 0.6-0.9. The results are more unstable by using the EML method. In this case, by selecting the best model ARFIMA (1, d, 0) according to the AIC criterion, the d parameter decreases to 0.54 for the dollar and to 0.30 for the pound. However, this discrepancy is not surprising since in the ARFIMA model short memory components are included in the AR coefficient; moreover, the model without autoregressive terms tends to increase d, since capture the short-run behavior of the series[7]. To sum up, the peseta/dollar and the peseta/pound real exchange rates show themselves as two fractional integrated processes that, even if not stationary, can be characterised as mean reverting processes.

To check the robustness of the option, we test for I(1) and I(0) hypotheses against the I(d) alternative. First, we apply the extension of the Dickey-Fuller test recently carried out by Dolado et al. (FI-DF) (2002) for the null of $d=1$ against the alternative $d<1$[8]. For both series the null is rejected. Second, we test for the null of $d=0$ against the alternative of $d>0$ by using the test carried out by Lobato and Robinson (1998). For both series, the test reject again the null. So, according to the Table 5, we can maintain the modelling of the peseta exchange rate behaviour against the dollar and the pound as two fractionally integrated processes.

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6 Smith et. al (1997) have a broad study comparing bias between semiparametric and ML methods.

7 See Agiakloglou et al. (1992).

8 We use several estimations of d obtained by the GPH, GSP and EML.

In order to obtain a persistent measure in this fractional context, we derive the *IRF* from expression [7]. To ensure that the Wold representation exists and the MA(∞) expression can be obtained, previously we have taken differences:

$$(1-L)^{d-1}(1-L)Y_t = \phi(L)^{-1}\theta(L)\varepsilon_t = A(L)\varepsilon_t \qquad [6]$$

$$(1-L)Y_t = (1-L)^{1-d}\phi(L)^{-1}\theta(L)\varepsilon_t = (1-L)^{\delta}A(L)\varepsilon_t$$

where $(1-L)^{\delta} = 1+\pi_1(\delta)L+\pi_2(\delta)L^2+...$ and the IRF is:

$$I(h) = \sum_{i=1}^{h}\pi_i(\delta)a_{h-i} \text{ with } \pi_i(\delta) = \frac{\Gamma(i-\delta)}{\Gamma(i+1)\Gamma(-\delta)} \qquad [7]$$

Finally, to estimate the effect on the level of the series Y_t, we calculate the cumulative impulse-response function:

$$CI(h) = \sum_{h=1}^{\infty}\ \sum_{i=1}^{h}\pi_i(\delta)a_{h-i} \qquad [8]$$

or equivalently $CI(h) = \sum_{h=1}^{\infty}\ \sum_{i=1}^{h}\pi_i(-d)a_{h-i}$

From this expression, we can estimate the *IRF* evolution and the half-life deviations. Previously, all parameters of the ARFIMA model have been jointly estimated by using the parametric approach of Sowel (1992), the best model being selected with the AIC criterion. The results are shown in Figures 3 and 4 , where the I(d) model is compared with the alternatives I(1) and I(0). With the ARIMA(1,1,0) la evolution of IRF shows a initial overshooting reaction and a permanent effect of the shock. With the ARMA(2,0,0) the *IRF* decay quickly offering a half-life, as has been reported earlier, of 7.94 and 6.60 years for the dollar and the pound, respectively. Finally, the path of *IRF* from the ARFIMA model selected is decreasing but more slowly, with a half life of 12.89 years for the dollar and 7.79 years for the pound.

To sum up, by using fractional integration methods the peseta/dollar and the peseta/pound real exchange rates can be adequately modelled as two mean-reverting series. This is enough to speak about the PPP as an *anchor* (Rogoff,1996) or a *good approximation* (Lothian and Taylor, 1996) for the long-run behaviour of real exchange rates. Nonetheless, the half-life deviations obtained by modelling the real exchange rate series as fractionally integrated processes are, as expected, too high. So, the question of the PPP puzzle remains open.

4. Introducing real factors

Given the long period under examination, it is natural to wonder about the possible influence of real-side factors on the long-lived deviations from the equilibrium detected in the peseta real exchange rate series. On the supply side, the long-lived deviations might answered to the presence of a Balassa-Samuelson effect. The Balassa-Samuelson theory starts by supposing that wages in traded good sectors evolve according to productivity, traded good sectors being those which concentrate the highest increases in productivity. It is also assumed that within a country wages equalize across sectors, so that the non-traded good sectors end up translating the salary increases originated by the increases in productivity of traded good sectors into relative increases of non-traded prices over traded good prices. As a consequence, considering that arbitrage only works for traded goods, if prices used in PPP tests include both kind of goods –which is the case with deflators-, we should take into account the possibility that a productivity bias is being introduced. This bias would be of real appreciation for the currency of the country of faster productivity growth and of real depreciation for the country of comparatively slower growth.

On the demand side, the same effect can be registered due to differentials in government expenditures. Public spending has a marked bias towards services consumption and, since this is a consumption of basically non-traded goods, government expenditures might increase the relative price of non-traded over traded goods. Here again, considering that arbitrage only works for traded goods, we should take into account the possibility that a bias is being introduced, acting in the same way as the productivity one.

For dealing with these issues we need a multivariate method such as the co-integration procedure proposed by Phillips and Hansen (1990) PH, and the JH approach of Johansen (1988, 1995). First of all, we apply these tests before introducing real factors with the aim to compare the previous results in the univariate context. Starting from the logarithmic specification:

$$s_t = \beta_0 - \beta_1 p_t + \beta_2 p*_t \qquad [9]$$

the long-run compliance of PPP requires $\beta_1= -\beta_2$ (the symmetry condition) and $\beta_1= \beta_2=1$ (the proportionality condition). Things do not change substantially when we apply cointegration tests to identify any long-run equilibrium between nominal exchange rate and relative prices. As is shown in Tables 6 and 7, the hypotheses of symmetry and proportionality always hold when applying the PH method. However, coherently with the ambiguous results from modelling the real exchange rates in an integer framework, there has not been found any co-integration relationship, either in the case of the dollar or the pound. Similar results are obtained through the Johansen approach, as Tables 8 and 9 report.

Thus, we first analyse if this lack of cointegration can be explained by the Balassa-Samuelson effect. To test for this possibility we introduce real per capita incomes as a proxy of productivity increases in [9]:

$$s_t = \alpha - \beta pp*_t - \gamma yy*_t \qquad [10]$$

where yy^* denotes the relative real per capita income of Spain over the foreign country, and pp^* the relative prices[9].

The introduction of incomes per head does not improve the adjustment of the model in a significant way. Although the compliance of proportionality remains with the two methods used, there is not clear evidence of cointegration between nominal exchange rate and prices, neither for the dollar nor for the pound. Only for the dollar when applying the JH procedure, a weak cointegration relationship emerges, even though the half-life of deviations from the mean keeps very high.

With regard to the public spending, this factor is considered in a similar way that the relative income:

$$s_t = \alpha - \beta pp^*_t - \delta gg^*_t \qquad [11]$$

where gg^* denotes the relative public spending in percentage of GDP of domestic and foreign countries.

By introducing this variable, there is evidence of cointegration for the dollar, although for the pound the null of non cointegration can only be rejected by the JH method at 10% level of significance. In any case, the half-live of deviations remained in both cases, for the dollar and the pound at a very high level, as occurs when the jointly effect of income and public spending is considered. Consequently, the lack of strong co-integration evidence for the pound and the long-lived deviations of the peseta/dollar and the peseta/pound real exchange rates from the equilibrium, led us to considerate that the Persistence puzzle might have born, not from the interference of economic real-side factors, but from market barriers such as the own Rogoff (1996)suggested.

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9 Although not reported here, the results are very similar when considering disaggregated incomes and prices.

5. Institutional factors: the persistence over periods

Looking for the responsibility of market barriers in the long-lived deviations of the peseta real exchange rate from parity, a useful mean might be the formal identification of the periods with highest persistence. To this end, we estimate the persistence in a recursive way. We start with a initial sample size of 30 observations, enlarged yearly until cover the full period, to estimate at each time point the corresponding best ARFIMA models (0,d,0), (1,0,0) or (1,d,0). Then, we derivate the IRF in accordance to the expression [8] and calculate the corresponding half-lives,

Figures 5 and 7 show the results from having estimated the autoregressive and *d* parameters according to this procedure. For the dollar, we can see how the fractional parameter is not significant before the outbreak of the Spanish civil war (1936). As far as the pound is concerned, the long-memory parameter does neither start to be significant until the thirties. Conversely, both *d* memory parameters increase systematically in the forties, during the first decade of the so-called Franco's autarky. In a different way, the same fact is reflected in the Figures 6 and 8, where the recursively computing of the half-live of parity deviations shows a sustained growth for both currencies from 1935 until 1950, in sharp contrast with what happened during the previous years, for most of which the peseta was a floating currency in a context of noticeable market integration. Such a contrast made us wonder about the degree of persistence for different periods according to the chronology of changes in the Spanish external policies.

As said above, the Spanish economy enjoyed acceptable international links before 1936. The context changed radically during Franco's autarky, the sub-period that ran from 1939 to 1959. During these two decades, until the *Plan de Estabilización y Liberalización* (Plan of Stabilization and Liberization) of 1959 approved a project of gradual openness, all kind of controls (including exchange rate controls) dominated the Spanish market. As a result of the Plan of 1959, the external convertibility of the peseta was restored in 1960 and the Spanish currency entered the Bretton Woods agreement in 1961. The peseta remained linked to the dollar until 1974 when a new period of floating opened. From then until its disappearance in 1998, the peseta floated against the dollar and also, if we exclude the brief British presence in the European Monetary System (1990-1992), against the pound. In order to consider these changes, we present the impulse response functions and the half-lives corresponding to the periods 1870-1935, 1870-1960, 1870-1974 and 1870-1998 in Figures 9 and 10.

To start with, both Figures show the impressive gains in persistence when the first period 1870-1935 is enlarged until 1870-1959 to include the autarky. If the years of autarky are considered, the half-life of deviations increase strongly, by around 100% for both currencies. Interestingly, these coincident gains in persistence correspond to a period of effective closure to the world, when the Spanish ratio of openness (as percentage of exports and imports over the GDP) decreased dramatically from an average of 22% in 1870-1935 to 8.5% in 1940-1959. In fact, these years concentrate the highest increase of persistence in 1870-1998, what could be explained by the presence of rigid exchange rate controls that, in a context of extreme protectionism, did not take into account the divergent domestic price behaviour in the few official rate revisions that took place during the autarky.

Conversely, relative prices do seem to have influenced the behaviour of the peseta in the foreign black markets, since if we substitute the exchange rates in force during the autarky in Spain, by the value of the peseta in such black markets, the half-life of deviations in the whole period 1870-1998 falls dramatically from 12.5 to 8.8 years for the dollar and from 7.8 to 3.5 years for the pound. Only twenty years out of a total of hundred thirty are able to reduce the half-life by 30% in the former case and 55 % in the latter one. This finding provides direct empirical evidence on the role played by relative prices in the black exchange rate markets of the peseta against the dollar and the pound[10]. Therefore, in an indirect way, the difference in half-lives when considering the black market or controlled exchange rates, serves to perfectly illustrate the responsibility of market interventions for impeding the accomplishment of the PPP hypothesis[11]. En última instancia, the dramatic reduction of half-lives when considering the black market rates, even if not enough to solve the puzzle, serves at least to support the Rogoff's (1996) idea about its institutional barrier origins.

As regards the rest of the period, the most remarkable issue lies in the different behaviour of the peseta/dollar and the peseta/pound real exchange rates. The half-life of deviations, although in a less pronounced way, continued to grow for the case of the dollar: from 8.6 years in 1870-1960 to 10.5 and 12.5 years in 1870-1974 and 1870-

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10 In line with the PPP model proposed by Culbertson (1975) to explain the behaviour of exchange rates in black markets.

11 An effect not captured in Taylor (2002), who finds a surprising reduced half-life of deviations coinciding with the Spanish autarky (1939-1959). Despite the fact that capital and trade barriers were then much higher than in previous years, the estimated half-life of the deviations of the peseta/dollar exchange rate from parity is much less long. The reason for such a shocking finding comes precisely from his use not of actual but of black market peseta/dollar exchange rates, which out of any control, reflected the forces (prices) operating in the market. The same remark is valid for López, Murray and Papell (2003). Their use of black rates would explain why when we consider the black peseta/dollar exchange rate, we obtain a reduction in the half-life deviations from 12.9 to 8.9 years that so much approaches to the 9.4 years obtained by them for the period 1880-1998. For this very reason, the use of the actual instead of the black market exchange rates would reinforce their idea that the PPP puzzle is *even more dramatic* than initially thought.

1998, respectively. By contrast, the half-life diminished, although slightly, in the case of the pound: from 9.2 years in 1870-1960 to 8.7 and 7.8 years in 1870-1974 and 1870-1998, respectively. This finding would fit in perfectly with the results of Papell and Theodoridis (2001), who in a framework of panel unit root tests dominated by European countries, show how the evidence in favour of PPP from 1973 onwards is stronger when take European rather than non-European numeraire currencies. Here, as suggested by Papell and Theodoridis (2001), the lesser variability of the peseta against the pound and the higher trade integration with the UK could contribute to explain the difference in the fractional parameter when considering the years 1974-1998. The dollar variability was much higher during these years (twenty times fold in terms of standard deviation) than the registered by the pound. Moreover, the reduction of persistence in the case of the pound could also be reflecting that the good arbitrage worked more effectively between Spain and the UK after the entrance of the latter in the European Community in 1973, with whom the former had signed a significant agreement in 1970, and above all, after the entrance of Spain in the European Union in 1986.

8. Conclusions

This paper studies the performance of the peseta/dollar and the peseta/pound exchange rates during the period 1870-1998. We start with an analysis of stationarity and unit roots and obtain inconclusive results. Moreover, we find long-lived deviations from parity; specially if the down bias of the least squares method is corrected by applying the median-unbiased procedure of Andrews and Chen (1994). Such long-lived deviations lead us to wonder about the possibility that both series exhibit slow mean reversion, not properly captured by integer analysis, this being the reason of our inconclusive results. The intuition is confirmed when finding that both real exchange rates can be successfully modelled by using fractionally integrated analysis. This itself serves to illustrate the well-known low power of integer tests to reject unit roots against mean-reverting but long-memory processes. More importantly, the finding means that the dominance of floating exchange rates in the series of the peseta/dollar and the peseta/pound rates does not impede that both of them behave as mean-reverting variables, supporting the idea of the PPP as an *anchor* (Rogoff,1996) or a *good approximation* (Lothian and Taylor, 1996) for the long-run behaviour of real exchange rates.

With regard to the slow reversion, we test in a cointegration framework for the possibility that real-sided factors might have influenced the behaviour of the peseta exchange rates, but found that neither the Balassa-Samuelson effect nor the relative government spending reveal significant in causing the long-lasting deviations from PPP. The result, again coincident with the mixed evidence in favour of real-sided factors reported in Rogoff (1996), drives us to consider the interference of market barriers as the cause for persistent deviations. To that end, we use impulse response functions in a fractional framework to calculate the half-life deviations recursively, by enlarging yearly the initial sub-sample 1870-1900 until the whole period 1870-1998. The analysis shows that the long-memory parameter increase systematically in the forties for both currencies, when not surprisingly, the Spanish external links experienced a radical change. From 1939 to 1959, during the so-called Franco's autarky, the Spanish closure to the exterior permitted the implementation of a rigid exchange rate control system that was managed independently from the path of relative prices. This could explain why if the period 1870-1935 is enlarged until 1870-1959 to include the autarky, the half-life of deviations increased strongly, by around 100% for both currencies. For the same reason that if we consider the black market instead of the controlled exchange rates of the peseta, the half-lives of deviations reduce dramatically in the whole period 1870-1998, from 12.5 to 8.8 years for the dollar and from 7.8 to 3.5 for the pound. To the extent that black rates escaped to any control, the noticeably difference between half-lives serves to perfectly illustrate the responsibility of institutional barriers, as suggested by Rogoff (1996), in the slow adjustments of real exchange rates to parity. Therefore, if not entirely, the paper would solve part of the PPP puzzle personalized in the stories of peseta/dollar and the peseta/pound exchange rates.

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Table 1: ***Unit roots and stationarity tests (dollar)***

With constant and trend	ADF	PP	MZt-GLS	KPSS (η_τ)
s	-2.14	-1.73	-1.67	0.29**
pp*	-0,87	-0.75	-0.40	0.33**
yy**	-1.79	-1.47	-1.53	0.26**
gg*	-3.80*	-3.02	-3.74*	0.17*
q=s-p+p*	-3.27*	-2.86	-3.22*	0.12**
With constant	**ADF**	**PP**	**MZt-GLS**	**KPSS (η_μ)**
s	0.21	0.47	1.00	1.26**
pp*	1.83	2.26	3.94	1.23**
yy**	-2.08*	-1.80	-1.19	0.39
gg*	-2.35	-2.95*	-2.41*	0.55*
q=s-p+p*	-3.20*	-2.80	-3.19**	0.30

Notes: ** Significant at the 1% level and * significant at the 5% level. The critical values for ADF and PP test are in McKinnon (1996). The number of lags of ADF has been selected in accordance with the method of Ng and Perron (1995)¸ and in MZ_t-GLS by SBIC criterion. In the PP test, Bartlett´s window has been used as a kernel estimator, choosing the bandwidth in the PP and KPSS test by the Newey and West method (1994).

Table 2: ***Unit roots and stationarity tests (pound)***

With constant and trend	ADF	PP	MZt-GLS	KPSS (η_τ)
s	-2.05	-1.81	-1.62	0.24**
pp*	-1.53	-1.60	-0.78	0.29**
yy**	-2.20	-2.07	-2.23	0.17*
gg*	-2.80	-2.72	-2.84	0.23**
q=s-p+p*	-3.33	-2.77	-3.23*	0.09
With constant	**ADF**	**PP**	**MZt-GLS**	**KPSS (η_τ)**
s	-0.06	0.18	0.95	1.24**
pp*	0.74	0.88	1.97	1.21**
yy**	-1.75	-1.75	-1.02	0.50*
gg*	-2.85	-2.78	2.02*	0.44
q=s-p+p*	-3.28*	-2.75	-2.64**	0.19

Notes: See Table 1.

Table 3: Half-life and persistence

Parameter persistence α	α_{ls}	CI(95%)	α_{mu}	CI(95%)
dollar	0.882	(0.864, 0.901)	0.908	(0.833, 0.992)
pound	0.866	(0.842, 0.891)	0.891	(0.812, 0.978)

Half lives	HL_{ls}	CI(95%)	HL_{irf}^{ls}	CI(95%)	HL_{mu}	CI(95%)	HL_{irf}^{mu}	CI(95%)
dollar	5.53	(4.74, 6.65)	6.02	(4.86, 7.26)	7.17	(3.78, 93.70)	7.94	(4.26, ∞)
pound	4.82	(4.03, 6.01)	5.31	(4.51, 6.60)	6.02	(3.33, 30.72)	6.60	(3.81, 33.20)

Notes: We start of the model $q_t = \mu + \alpha q_{t-1} + \sum_{i=1}^{k} \psi_i \Delta q_{t-i} + \varepsilon_t$. The persistence parameter α is estimated by lest-square (ls) and the median unbiased (mu) method proposal by Andrews (1993) and Andrews and Chen (1994). The number of lags, k, selected by the Ng and Perron (1995) method is 2 in both cases. The half-life is calculated from the ls and mu method and estimating the correspondence AR(2) impulse-response function (irf).

Table 4: Estimation of the long-memory parameter (*d*)

	GPH			EML-Arfima				GSP m=T/n		
	τ=0.5	τ=0.6	τ=0.7	(0,d,0)	(1,d,0)	(0,d,1)	(1,d,1)	n=11	n=7	n=4
USA	0.76	0.73	0.90	1.10	0.54	0.83	0.46	0.71	0.65	0.86
	(0.001)	(0.000)	(0.000)	(0.300)	(0.021	(0.130)	(0.052)	(0.000)	(0.000)	(0.000)
AR					0.60		0.61			
					(0.001)		(0.017)			
MA						0.34	0.14			
						(0.001)	(0.250)			
AIC				-213.13	-223.00	-219.70	-222.15			
GB	0.74	0.62	0.82	1.05	0.30	0.85	0.13	0.92	0.68	0.86
	(0.001)	(0.000)	(0.000)	(0.553)	(0.000)	(0.197)	(0.001)	(0.000)	(0.000)	(0.000)
AR					0.75		0.81			
					(0.000)	0.26	(0.000)			
MA						(0.024)	0.13			
							(0.388)			
AIC				-266.75	-274.15	-268.83	-267.67			

Notes: p-value in brackets.

The fractionality integrated parameter *d* is estimated by the Geweke, Porter Hudak (1983) method, GPH, using ordinary least squares (OLS) as the negative slope of a regression of the following expression:

$$\ln\left\{I(w_j)\right\} = \beta_0 - \beta_1 \ln\left\{4\sin^2(w_j/2)\right\} + \eta_j$$

where the spectral density function $f_u(w)$ has been substituted by the sample periodogram, evaluated in a harmonic frequencies band $w_j = 2\pi j/T$, j=1,...m close to 0, where m= T^{τ}, with τ= 0.5, 0.6, and 0.7. The errors standard are calculated by the theoretical asymptotic variance of $\varepsilon_t = \pi^2/6$.

The exact maximum likelihood method –EML- proposed by Sowell (1992) stars of a general fractionally integrated time series model for $Y_t \sim I(d)$:

$\phi(L)(1-L)^d Y_t = \theta(L)\varepsilon_t$ and supposing that $Y_t \sim N(0,\Sigma)$ the probability density function is:

$f(Y_T,\Sigma) = (2\pi)^{-T/2}\left|\Sigma\right|^{-1/2}\exp\left\{-1/2Y'_T\Sigma^{-1}Y_T\right\}$ and under stationarity the autocovariance function in term of the parameters of the model is calculating by:

$$\gamma(s) = \frac{1}{2\pi}\int_0^{2\pi} f_Y(\lambda)e^{i\lambda s}d\lambda$$

Finally, we can maximize the follow expression:

This method is under valid for –0.5>d>0.5. If it suspect d>0.5 the series should be previously differenced.

$$\log L(d,\phi,\theta,\sigma_\varepsilon^2) = -\frac{T}{2}\log(2\pi) - \frac{1}{2}\log\left|\Sigma\right| - \frac{1}{2}Y'\Sigma Y$$

The semiparametric gaussian estimator of Robinson (1995) and Robinson and Henry (1999) define the periodogram $I(\lambda) = \frac{1}{2\pi n}\left|\sum_{t=1}^{n} x_t e^{it\lambda}\right|^2$ and estimating H by:

$\hat{H} = \arg\min_{\Delta_1 \le h \le \Delta_2} R(h)$ where $0 < \Delta_1 < \Delta_2 < 1$ and

$R(h) = \log\left\{\frac{1}{m}\sum_{j=1}^{m}\frac{I(\lambda_j)}{\lambda_j^{1-2h}}\right\} - (2h-1)\frac{1}{m}\sum_{j=1}^{m}\log\lambda_j$ and where *d*=H-1/2.

Table 5: Test for fractionally integrated alternatives

	DF-FI			LR non-parametric test		
	$\hat{d}$	$t(\hat{d})$	$Q(4)$	m=20	m=30	m=40
dollar				18.32**	45.67**	47.30**
	0.54	-2.73**	1.24			
	0.65	-2.60**	1.32			
	0.71	-2.52**	1.36			
	0.73	-2.50**	1.37			
	0.76	-2.46**	1.39			
	0.86	-2.34**	1.47			
	0.90	-2.29**	1.50			
pound				5.35*	10.25**	14.81**
	0.30	-2.63**	0.88			
	0.62	-1.96*	1.05			
	0.68	-1.85+	1.06			
	0.74	-1.75+	1.07			
	0.82	-1.65+	1.07			
	0.86	-1.55	1.07			
	0.92	-1.46	1.08			

Notes: +, *, ** significant at the 10%, 5% and 1% levels respectively

The test proposed by Dolado, Gonzalo and Mayoral (2002), DF-FI, is carried out parting from the following regression

$$\Delta^{d_0} y_t = \phi \Delta^{d_1} y_{t-1} + \sum_{i=1}^{p} \xi \Delta y_{t-i} + \varepsilon_t$$

where the hypothesis of interest is d_0=1 and d_1<1, and the series have been filtered using the binomial expansion of the operator $(1\text{-L})^d$:

$$(1-L)^d = \sum_{j=0}^{\infty} \pi_j(d) L^j \text{ where } \pi_j(d) = \frac{\Gamma(j-d)}{\Gamma(-d)\Gamma(j+1)}$$ and $\Gamma(\bullet)$ is a gamma function.

Dolado et. al (2001) show that $t(d_1)$ is an OLS consistent estimator of ϕ, and when $d_1 \in [0,0.5)$ the test has a non-standard distribution under the null that is tabulated for various sample sizes and a wide range of values of *d*. And for $d_1 \in [0.5,1)$ it follows a Gaussian law.

$Q(k)$ Ljung-Box test calculated for the first k-autocorrelations of residuals.

The LR test proposed by Lobato and Robinson (1998) tests Ho: *d*=0, against the alternative *d*>0. The LR test is distributed as a $\chi^2(1)$.

Table 6: Cointegration analysis (Phillips-Hansen method), dollar

Model	$\hat{\beta}(t-ratio)\ H_O : \beta = 1$			Wald test	CRADF
	pp*	yy*	gg*		
s, pp*	1.01(0.000)			0.167(0.683)	-3.22
s, pp*, yy*	1.00(0.000)	-0.14(0291))		0.011(0.917)	-3.32
s, yy*, gg*	0.85(0.000)		-0.38(0.000)	18.124(0.000)	-4.17+
s, pp*, gg*, yy*	0.86(0.000)	-0.01(0.931)	–0.34(0.000)	20.043(0.000)	-4.17+

Notes: p-value in brackets; +, *, ** significant at the 10%, 5% and 1% levels respectively. The long-run variance has been estimated using a Bartlet's window with size=4. The CRADF tests for the existence of a unit root on the residuals; the critical values can be consulted in McKinnon (1991). The number of lags has been selected in accordance with the method of Ng and Perron (1995). We have introduced two dummies for First and Second World War in the models with government spending.

Table 7: Cointegration analysis (Phillips-Hansen method), pound

Model	$\hat{\beta}(t-ratio)\ H_O : \beta = 1$			Wald test	CRADF
	pp*	yy*	gg*		
s, pp*	1.01(0.000)			0.100(0.751)	-3.30
s, pp*, yy*	0.97(0.000)	0.26(0.116))		0.563(0.453)	-3.22
s, yy*, gg*	0.97(0.000)		-0.18(0.004)	0.911(0.340)	-3.48
s, pp*, gg*, yy*	0.90(0.000)	0.35(0.027)	–0.21(0.001)	5.180(0.023)	-3.58

Notes: See Table 4.

Table 8: Cointegration analysis (Johansen method), dollar

Model	VAR(k)	Cointegration rank		Standardized	α estimated	half life	Proportionality
		max.eigenv.	trace	coefficients β'			hypothesis
s, pp*	2	12.94	15.34	(1, -0.96)	(-0.09, 0.05¡4)	7.35	0.42(0.519)
s, pp*, yy*	1	14.03	21.39*	(1, -0.89, 0.24)	(-0.06, 0.04)	16.98	1.94(0.164)
s, pp*, gg*	1	16.07*	26.71**	(1, -0.86, 0.25)	(-0.07, 0.05)	9.55	4.78(0.029)
s, pp*, yy*, gg*	1	20.66*	31.44**	(1, -0.90, 0.49, 0.16)	(-0.03, 0.05)	22.76	2,36(0.124)

Notes: ** Significant at the 95% level; * Significant at the 90% level; p-value in brackets. The order of VAR(k) has been selected in accordance with the information criterion and the diagnosis of the residuals omitted for reasons of space. The hypothesis tested is whether the vector β_1 (1, -1) belongs to the cointegration space, using the likelihood ratio test distributed as a $\chi^2(1)$. Income and government spending are introduced as an I(1) exogenous variables. Deterministic terms such as the constant have been introduced unrestricted.

Table 9: Cointegration analysis (Johansen method), pound

Model	VAR(k)	Cointegration rank		Standardized	α estimated	half life	Proportionality
		max.eigenv.	trace	coefficients β'			hypothesis
s, pp*	2	11.71	12.17	(1, -1.00)	(-0.10, 0.04)	6.58	0.00(0.954)
s, pp*, yy*	1	8.70	12.55	(1, -0.92, -0.30)	(-0.02, 0.06)	34.31	0.46(0.500)
s, pp*, gg*	1	15.73	21.15*	(1, -0.97, 0.41)	(-0.10, 0.03)	6.58	0.18(0.671)
s, pp*, yy*, gg*	1	15.76	22.14	(1, 0.98, 0.07, 0.042)	(-0.03, 0.04)	22.76	0.05(0.827)

Notes: See table 6.

Figure 1: Evolution of real exchange rate peseta/dollar

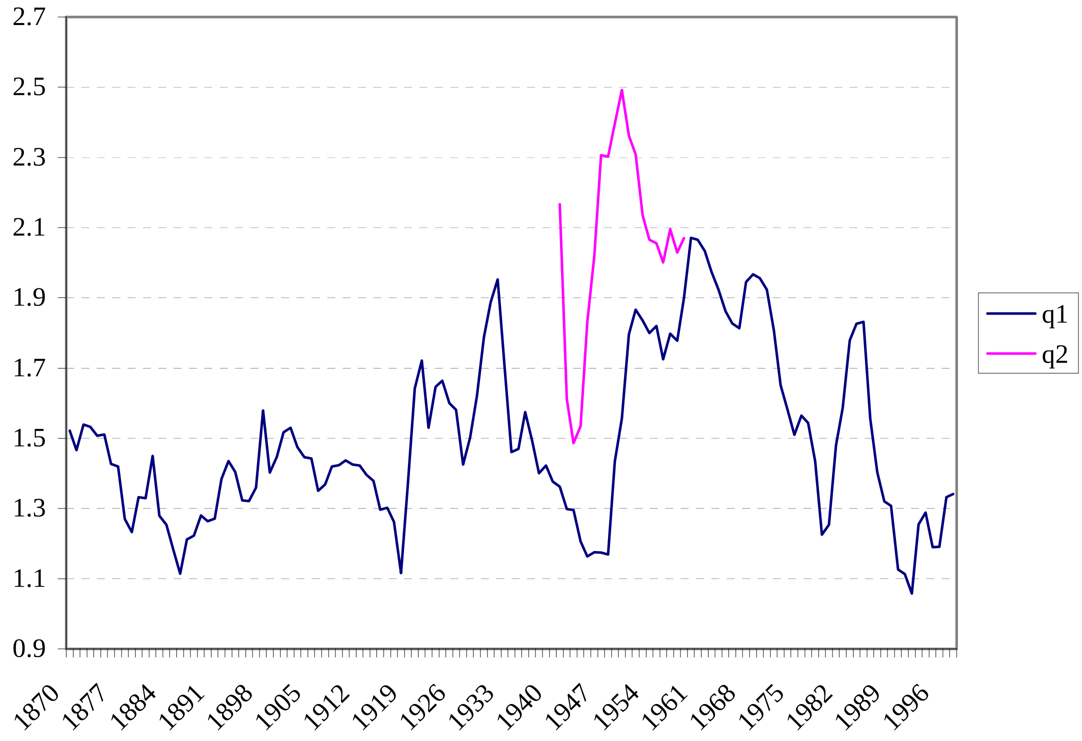

Figure 2: Evolution of real exchange rate peseta/pound

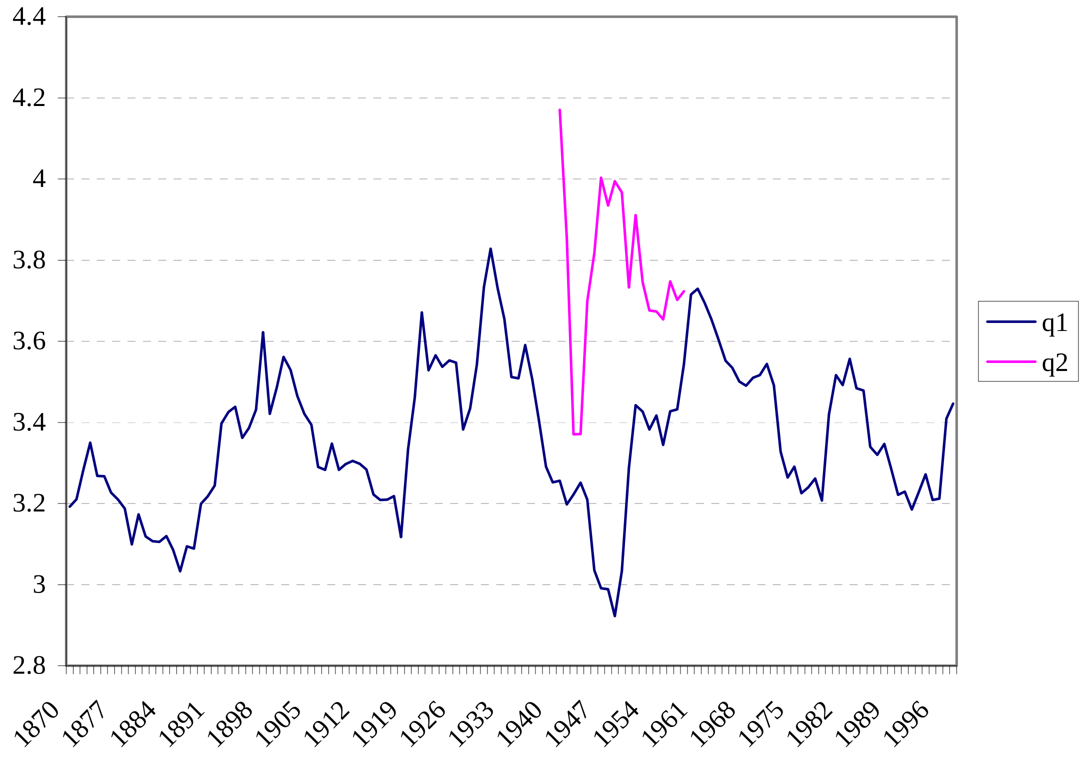

Figure 3: Impulse response functions of real exchanghe rate (dollar)

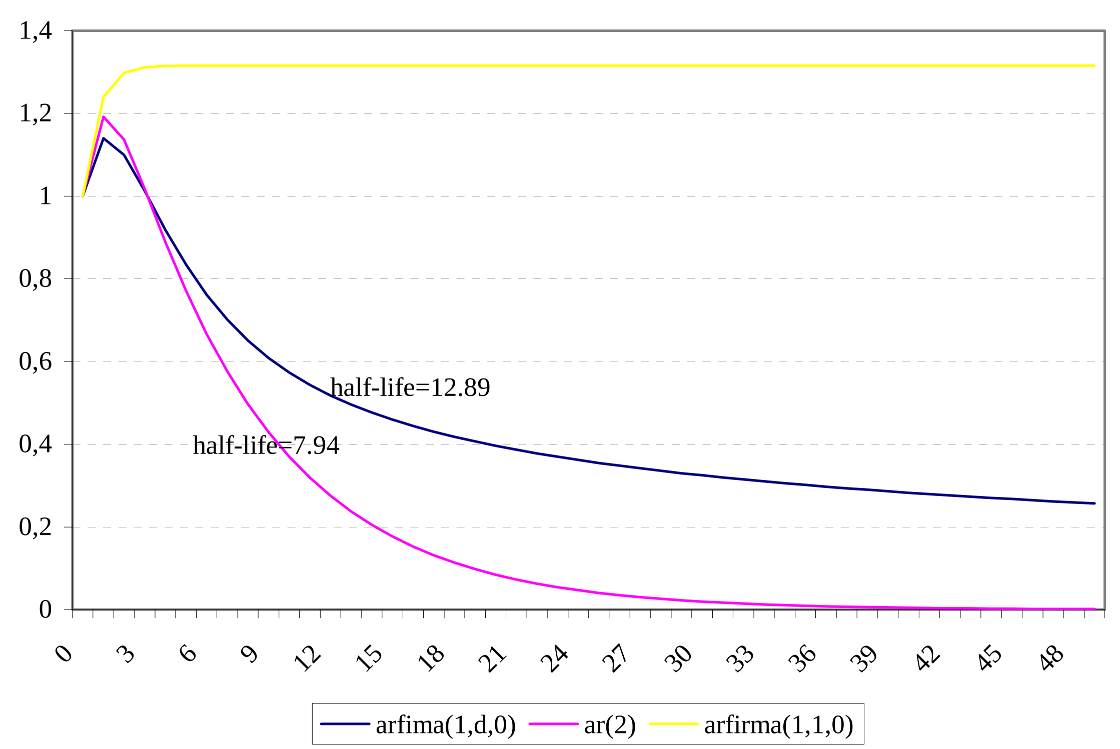

Figure 4: Impulse response functions of real exchanghe rate (pound)

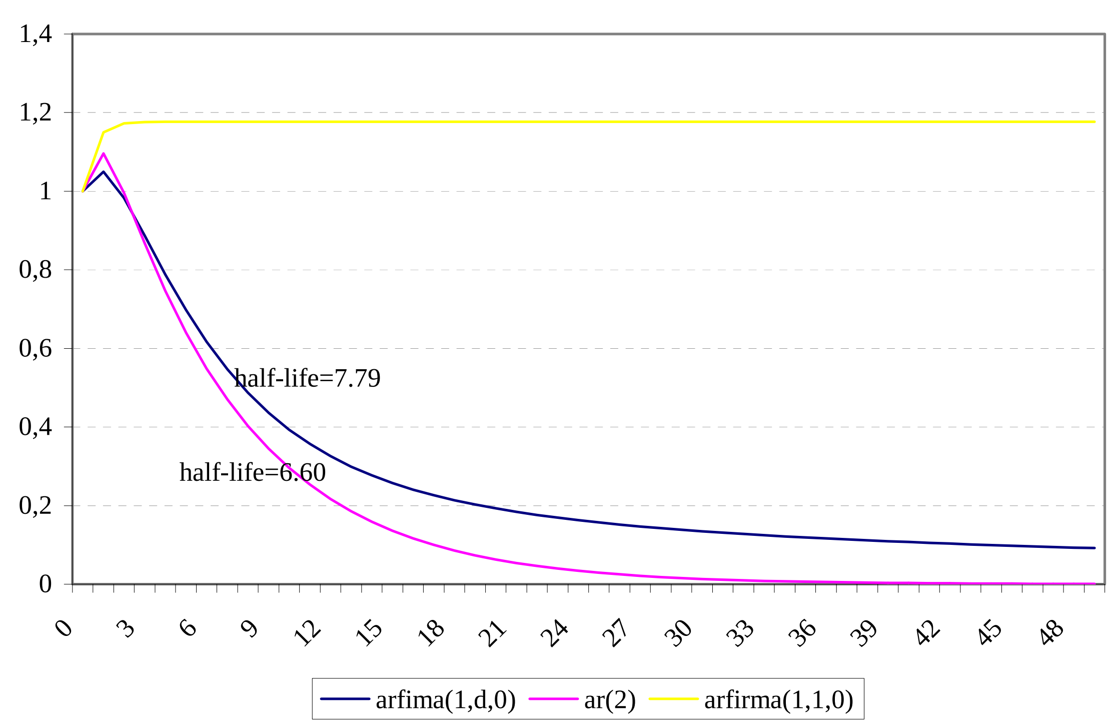

Figure 5: Recursive estimation of ARFIMA model (dollar)

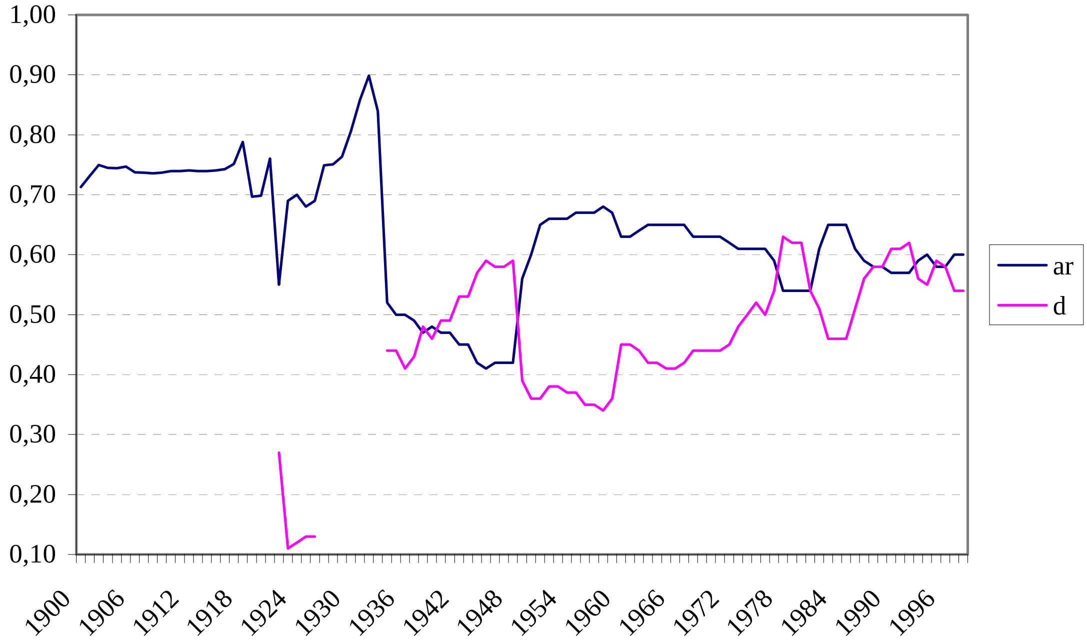

Figure 6: Recursive computing of half-life from a ARFIMA model (dollar)

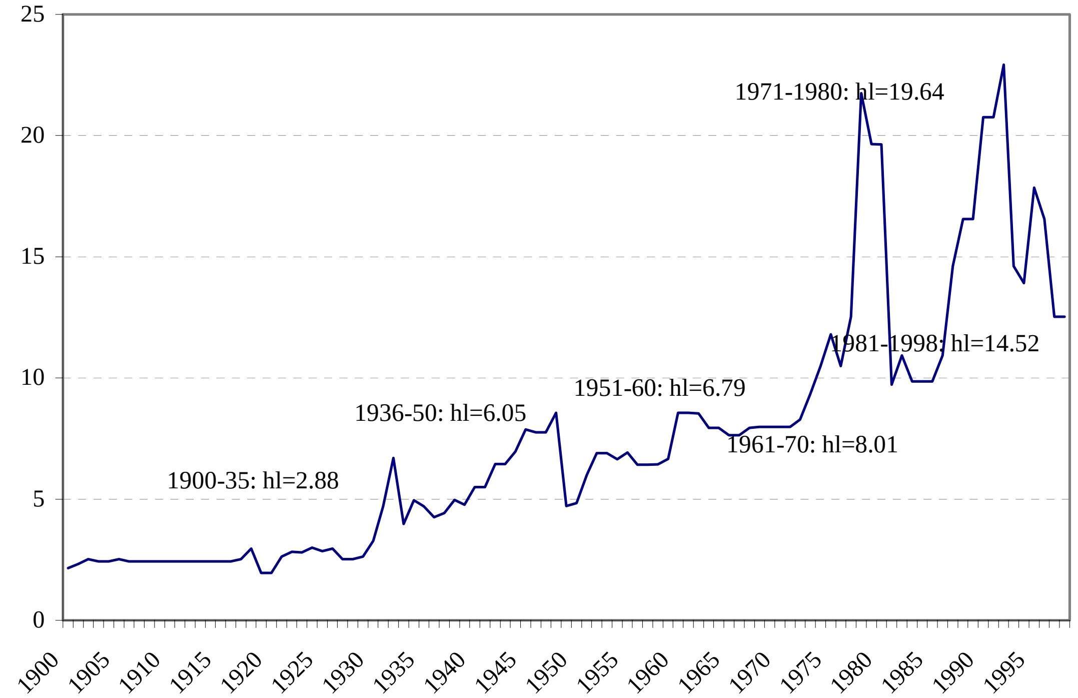

Figure 7: Recursive estimation of ARFIMA model (pound)

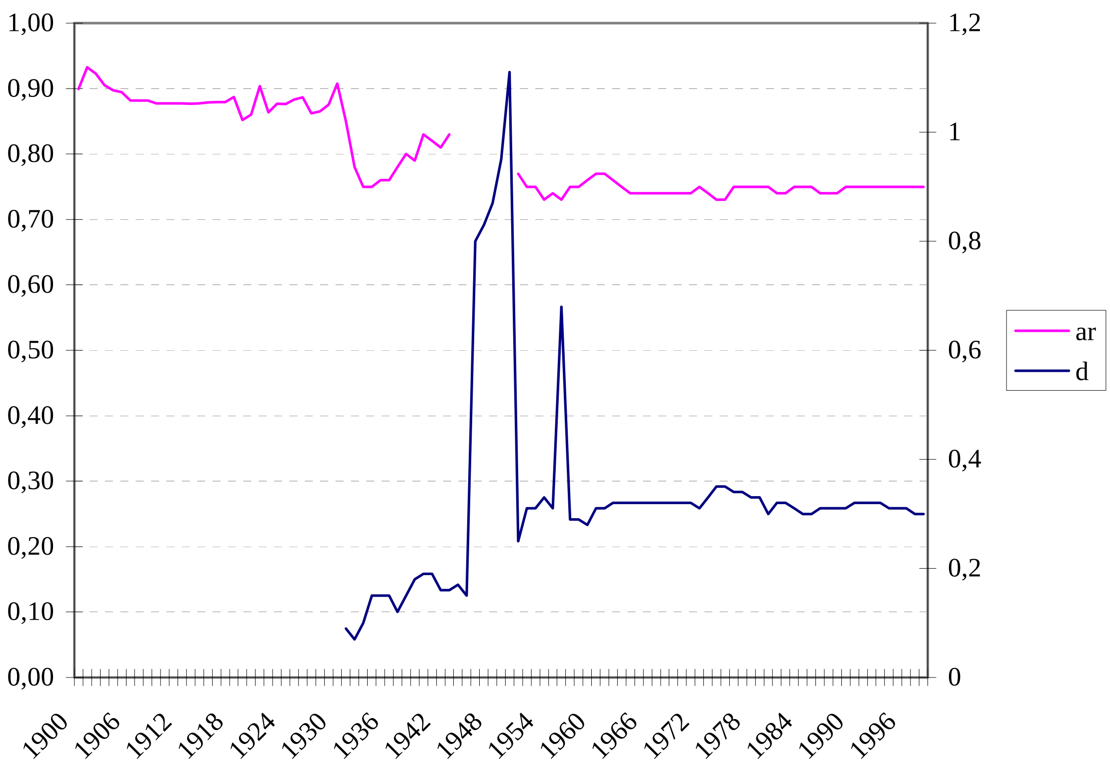

Figure8: Recursive computing of half-life from a ARFIMA model (pound)

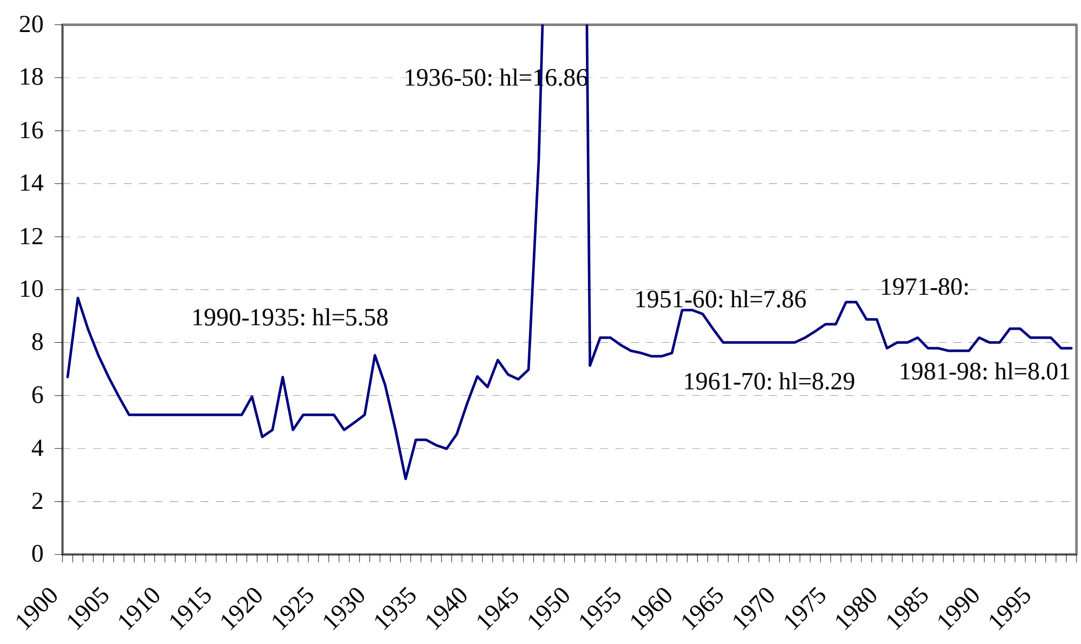

Figure 10: IRF from ARFIMA model for dollar real exchange rate in different periods

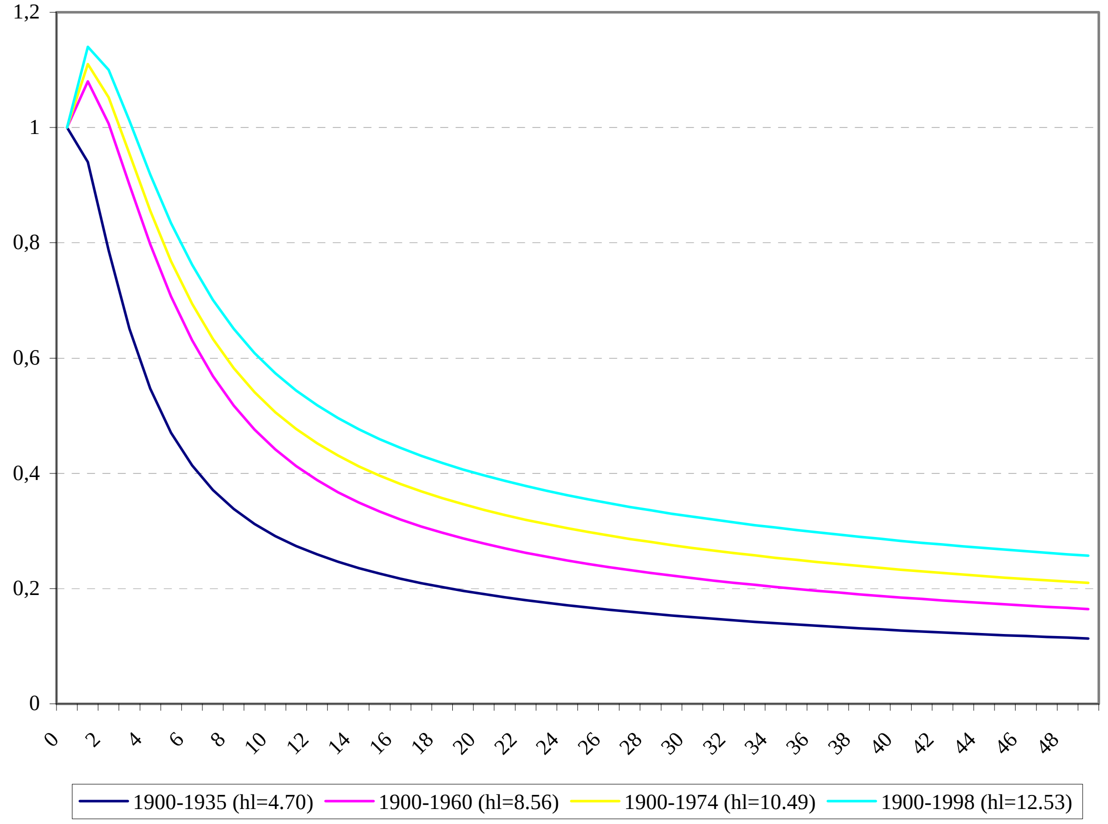

Figure 11: IRF from ARFIMA model for pound real exchange rate in different periods