STRUCTURAL BREAKS, CO-BREAKS AND REAL EXCHANGE RATES REGIMES

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ABSTRACT:

This paper examines the properties of the historical real exchange rate series constructed in Taylor (2002). We first split the sample into four differentiated intervals, namely, 1870-1918, 1919-1936, 1937-1972 and 1973-1996, which are almost coincident with the division carried out in the original paper. These periods have been selected by using methods that look for the number and location of structural breaks endogenously. We then analyse the persistence and the volatility of shocks in those periods. We find that periods dominated by a fixed exchange regime exhibit lower volatility, but greater persistence, than what is observed in periods with a flexible exchange regime. This leads us to conclude that the degree of persistence and volatility of the real exchange shocks are clearly altered by the decisions taken by the monetary authorities.

JEL Classification: C22; E43

KEYWORDS: Exchange Rate Regimes, Structural Breaks; Long data span; persistence; volatility

1. INTRODUCTION

The enormous literature devoted to real exchange rates (RER) has paid relatively little attention to the regime dependence issue. An important fraction has focused on the study of specific exchange rate regimes such as the Gold Stardard [Diebold, Husted and Rush (1991)], the interwar period [Michael, Nobay and Peel (199)] and especially, the post-Bretton Woods period [Abuaf and Jorion (1990), Taylor and Sarno (1998), Gadea, Montañés and Reyes (2004), Murray and Papell (2005a) and Choi, Mark and Sul (2006) among others). Furthermore, an other group of studies use long-horizon data span with the aim of increasing the power of unit root tests and, consequently, of better capturing the mean-reverting properties or real exchange rates [Frankel (1986), Abuaf and Jorion (1990), Lothian and Taylor (1996) and Taylor (2002)].

These studies have provided strong evidence in favor of Purchasing Power Parity (PPP) compliance, although with a high persistence. As Rogoff (1996) highlighted, they show a very *remarkable consensus* of 3-5 year half-life deviations from the equilibrium level[1]. Such lasting deviations are difficult to reconcile with the high short-run variability of real exchange rates, provoked by monetary factors, and they are the origin of the so-called PPP puzzle. However, as Rogoff (1996) pointed out, long-span data sets mix fixed and floating exchange rates regimes, which could exhibit different adjustment speeds and volatilities [Liang (1998)]. Nevertheless, a recent and

[1] It has recently been argued that the reported 3-5 year half lives are underestimated, implying that the puzzle is even bigger than was first thought. The works of Murray and Papell (2005b) and Lopez, Murray and Papell (2005b), among others, that re-examine the data of Lothian and Taylor (1996) and Taylor (2002), respectively show that traditional measures underestimate persistence maintaining the 'PPP puzzle'. As far as panel data are concerned, although they reduce the interval confidence, the lower bound is still too high to be explained by nominal rigidities (Murray and Papell, 2005a; Choi, Mark and Sul, 2006).

comprehensive work of Taylor (2002) did not find a significant change in mean-reversion speed across the four regimes that existed during the XX century.

In a framework of studies which considere long historical periods of real exchange rate behavior, it is plausible, as Sarno and Taylor (2002) and Taylor and Taylor(2004) argue, to think about the possibility of structural breaks. However, while the presence of structural breaks has been considered in unit root tests that deal with PPP compliance, studies about the stability parameters involved in real exchange rates are relatively scarce. Lothian and Taylor (1996), using a Chow test, do not find different behavior in the dollar-pound and French franc-pound real exchange rates as a result of the collapse of Breton Woods. Neither Lothian and McCharty (2002) find evidence of structural change for the case of the Irish pound against the German mark, the British pound and the dollar due to the transition to a floating regime during the seventies. However, the setting up of the European Monetary System provoked a break in the behavior of the Irish pound and the German Mark. The main inconvenience of this kind of studies, as well as of the paper of Taylor (2002), stem from the exogeneity of the number and location of break points.

Following a recent strand of literature that emphasizes the importance of a endogenous determination of structural breaks, Hegwood and Papell (1998) apply the technique developed by Bai and Perron (1998) that is able to determinate the structural breaks together with the rest of the parameters of the model[2]. They coined the term 'quasi purchasing power parity' for reflecting the change that real exchange rate equilibrium suffers in long periods. Furthermore, the structural breaks seem to be explained by autochthonous national features and not by international agreements.

[2] This method has also been applied by Sabate et al.(2003) for the Gold Standard period and by Gadea et al. (2004) for the European Union currencies. The results show the importance of considering structural breaks in the models that test the PPP hypothesis, and find changes related to specific changes in exchange rate regimes in the first case and to the creation of the euro in the second.

Finally, by considering different regimes in the sample, they obtain such reduced persistence in real exchange rates that doubts are cast on PPP puzzle validity.

Summarizing, there is no consensus eeither about the presence or not of structural breaks in long periods of real exchange rate evolution or the most suitable econometric techniques to select them. Also, it is not clear whether the origin of these possible changes, is institutional agreement or the national features of each country. In the first case, we could point to an international path of exchange rate regimes; in the second, we would have individual regimes due to the specific monetary conditions. Finally, neither is clear how the different regimes, institutional or not, affect the main statistical properties, namely, persistence and volatility.

The present paper tries to shed further light on this debate and confirm the apparent non-neutrality of real exchange rates on monetary regimes. Taylor (2002), being aware of the importance of institutional agreements on the evolution of real exchange rates during the last century, split the sample exogenously into four regimes. However, the data may contain this information and, consequently, we propose to look for the breaks point endogenously[3]. So, we propose a pure structural change model that includes changes in the mean and in the autoregressive parameter and we apply the method of Bai and Perron (1998, 2003).

Our findings support the non stability of real exchange rate behavior for the 20 countries analyzed and locate shocks of a permanent nature. This is not a surprising result because the XX century has been witness to numerous institutional agreements. Both the abandonment of the Gold Standard and the collapse of the Bretton Woods system appears as structural break points, illustrating real exchange rate regimes with very different statistical properties –persistence and volatility-. This outcome is in sharp

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[3] Furthermore, as Reinhart and Rogoff (2004) have recently shown, the institutional regimes do not have to coincide with the true behavior of exchange rates.

contrast with that obtained by Taylor (2002). During the floating regimes we estimate half-lives under 2 years, with narrowly bounded confidence intervals. Nevertheless, the 'PPP puzzle' remains in fixed regimes where we find higher half-lives and wide confidence intervals that tend to infinite in their upper bound.

The remainder of the paper is organized as follows. Section 2 present the methodology and results to detect the number and location of structural breaks endogenously. The possibility of co-breaks and, consequently, a common international path is explored in Section 3. The main statistical properties of RER across regimes, persistence and volatility, are analyzed in Section 4. Finally, a last section of conclusions summarizes the main findings of the paper and compare our results with the obtained by Taylor (2002).

2. DETECTING STRUCTURAL BREAKS.

The aim of this section is to analyse the evolution of the behaviour of the exchange rates from a long-run perspective. We follow the seminal paper of Taylor (2002) because the data set employed in this paper considers a high volume of information about the evolution of the real exchange rates from a time perspective and the sample of countries is very broad. The time span varies for the different countries, but mostly considers information of more than a century (with a maximum of 127 annual observations). The countries in the data base are mostly the so-called industrialized countries, although it also includes 3 Latin American countries (Argentina, Brazil and Mexico). Finally, in order to obtain the real exchange rate, two different deflators are employed: Consumer price index and the GDP deflator.

The analysis carried out in Taylor (2002) involves the division of the evolution of the exchange rates to different regimes. However, the number and the length of these periods are imposed a priori without being verified by the data. Although the resulting regimes are linked to very well-known economic periods, we consider it more appropriate to determine these periods without imposing any a priori restriction, obtaining this information directly from the data by way of the application of powerful econometric techniques to detect and estimate the presence of structural breaks in the variables. In particular, and given that the work of Taylor (2002)[4] [5] offers evidence against the presence of a unit root in the real exchange rates of the currencies analysed, the use of the procedure recently proposed in Bai and Perron (1998[a,b], 2003) seems to a be an appropriate method because it offers a quite accurate method for determining the number of breaks and estimating the periods delimited by these breaks. Following these authors, we base our study on the estimation of the following linear model with m breaks:

$$q_{it} = \mu_{ij} + \rho_{ij}\, q_{it-1} + u_{it}, \qquad t= TB_{i\,j-1}, ..., TB_{ij} \qquad j=1, 2, ...m+1 \; i=1, 2,\ldots, 20 \qquad [1]$$

where q_{it} is the log of the real exchange rate of the i-th country in period t and TB_{ij} represents the moments when the breaks appear, with $T_0 = 0$ and $T_{m+1} = T$. Model [1] implies that we are assuming that the real exchange rates are following an AR(1) process. More complex models could have been considered but most of the papers devoted to the study of the real exchange rates choose this simple model as the most appropriate to capture the evolution of this variable, especially when working with yearly data. In any event, we should also note that the presence of autocorrelation in the

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4 Taylor (2002) concludes "that PPP has held in the long run over the twentieth century for my sample of 20 countries". Therefore, "it is no longer productive to devote further attention to the stationarity question". These assertions have recently been qualified by Papell and Prodan (2003) and Lopez et al (2005b) who, employing a similar database, cannot reject the unit root null hypothesis for all the real exchange rates. Their criticism lies in the size and power distortions caused by selecting the number of lags using a(n) LM statistic. In this regard, we should recall that Taylor (2002) cannot reject the presence of a unit root for the Japanese case.

5 It is now well-known that not considering structural breaks may induce the non rejection of the unit root null hypothesis (see Perron, 1989). By contrast, when these changes are taken into account, the evidence in favour of the PPP during the XX century is greater, as is proven in Papell and Prodan (2003).

perturbation of the model [1] should be interpreted as the inadequacy of the AR(1) and we have not been able to reject the non-autocorrelation null hypothesis for all the cases considered. Thus, the simple AR(1) seems to be sufficient to describe the behavior of the real exchange rates that we analyze in this paper. We should also note that we admit the presence of changes in all the parameters of the model, which allows for both changes in the intercept of model [1] and in its autoregressive parameter. Finally, we should recall that we have not included deterministic trends in the model, in the spirit of Cassel's PPP (see Papell and Prodan, 2003)[6].

Having described our initial model, we should then test for the presence of breaks. The Bai-Perron procedure implies the estimation of the above equation considering that the break may appear in any period of the sample size. A Chow-type test is then defined in order to determine the existence of a first break, which coincides with the period where this Chow-type statistic reaches its maximum value. The existence of multiple breaks is analyzed by applying this procedure sequentially, combining with the repartition method described in Bai (1997). In order to determine the existence of breaks, we can use the UD_{max} and WD_{max} statistics, which test for the null hypothesis of no structural breaks versus the presence of an unknown number of breaks. Once we can offer evidence against the absence of breaks, then we can apply a sequential procedure in order to determine the most appropriate number of breaks. This approach is based on the sequential application of the sup $F_T(\ell+1|\ \ell)$ statistic, with the statistic $F_T(\ell+1|\ \ell)$ being defined as the difference between the sum of squared residuals obtained with ℓ breaks and the one obtained with $(\ell + 1)$ breaks. This sequential procedure starts by estimating a model with no breaks and, subsequently,

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6 Cuddington and Liang (2000) and Lothian and Taylor (2000) open the debate about the inclusion of deterministic trends in order to explain the behavior of the real exchange rates.

performs parameter-constancy tests for every subsample (those obtained by cutting off at the estimated breaks), adding a break to a subsample associated with a rejection using the test sup $F_T(\ell+1|\ \ell)$. This process is repeated by increasing ℓ sequentially until the test fails to reject the null hypothesis of no additional structural break[7].

Table 1 about here

The results that we have obtained from the application of this method are presented in Table 1. Inspection of this table allows some interesting insights. First, we should note that the evidence against the no breaks null hypothesis is quite clear. We can only not reject this hypothesis for the case of Japan[8], whilst for Argentina we can only reject it when a liberal 10% significance level is considered. We can also observe that the number of breaks selected by the Bai-Perron procedure differs for the different real exchange rates. A single break or a double break is considered for 14 real exchange rates, but we need 3 breaks when explaining the cases of Finland, Italy and the United Kingdom and 4 breaks are required to appropriately model the Brazilian real exchange rate.

A second interesting insight is obtained when the periods originated by structural breaks are analyzed, in that some evidence in favor of the existence of a common international pattern of behavior seems to emerge. This result is very interesting because this common pattern of behavior is not always supported in the literature. For example, Hegwood and Papell (1998) suggest the existence of country specific causes in order to explain the presence of breaks in the real exchange rates, although we should note that our sample is quite different to that considered in Hegwood and Papell (1998). Our

7 Following Bai and Perron (2003), we have considered a trimming parameter of 15%, although we have occasionally used 20%.

8 We should note that the rejection of the unit root null hypothesis is not supported much for the Japanese case and, therefore, this can be the cause of the lack of evidence against the no structural breaks null hypothesis.

results are more related to the number and the length of the breaks considered in Taylor. To see this, we should note that, according to our results, it can be appreciate that the abandoning of the Gold Standard and the collapse of the Bretton Woods parities system have influenced the evolution of our set of real exchange rates. Furthermore, we could also observe the existence of a third break. However, it is not easy to say where this break occurs. Both the mid 1930's and mid 1940's are possible candidates (Obstfeld et al, 2004).

Thus, our results could first point to the existence of an international common pattern of behavior, as in Taylor (2002), although they exhibit some discrepancies in the number and the length of the breaks. The next section is devoted to the possible confirmation of these initial results, placing special emphasis on the confirmation of the existence of this common international pattern of behavior in the exchange rates, as well as the more accurate estimation of the different regimes that characterized the evolution of this variable during the last century.

3. CO-EVOLUTION OF EXCHANGE RATE REGIMES BETWEEN COUNTRIES.

In the previous section we have found some evidence in favour of the existence of a common international pattern of behaviour for real exchange rates, in that the group of real exchange rates studied exhibits a high degree of coincidence with respect to the presence of structural breaks in their evolution. This result could help to reinforce the choice of the regimes that were a priori imposed in Taylor (2002) in order to describe the evolution of the exchange rates during the last century. The aim of this section is to confirm this initial idea. To that end, we propose to estimate a joint model and, then, to study whether a unique common international pattern of behaviour exists. The idea is to estimate the following model:

$$q_{i,t} = \mu_i + \rho_i q_{i,t-1} + \delta_{i,1} D_1 + \phi_{i,1} D_1 \cdot q_{i,t-1} + \ldots + \delta_{i,m} D_m + \phi_{i,m} D_m \cdot q_{i,t-1} + u_{i,t} \qquad [2]$$

where i=ar, au, be, …, uk, where the sample size covers the period 1892-1996. The dummy variable D_m takes the value 1 if t > TB_m and 0 othewise, with TB_m defining the period where the break appears. Instead of estimating 17 separate equations, we consider it more sensible to estimate a joint model that includes the 17 countries by way of the SUR methodology[9]. We should note that the joint estimation of the model allows us to take into account the existence of cross-sectional relationships of the shocks and, furthermore, it provides an appropriate framework in which to test for the presence of a unique international common pattern of behaviour.

The estimation of model [2] requires defining the different dummy variables appropriately. The previous section offers very useful information in this regard and, therefore, we have decided to employ it in order to estimate the joint model. Thus, the strategy that we have followed is to admit the presence of up to 4 breaks, with each of these breaks being related to a single value of the following four intervals (1914-1920), (1934-1939), (1944-1949) and (1968-1974). These intervals are clearly related to the results obtained from the application of the Bai-Perron methodology, whose results were presented and discussed above. Thus, we can consider the estimation of the periods where the breaks appear quasi-endogenous.

Table 2 about here

Figure 1 about here

For all the possible combinations, we use a likelihood ratio statistic in order to test for the significance of the *m-th* break versus the existence of m-1 breaks. The results obtained from the application of this strategy are summarized in Table 2. It can be seen

9 We have excluded Brazil and Japan from this first panel due to lack of data. Nevertheless, they have been included in a second, less extensive panel (1386 observations), with the GDP deflator as the price index with conclusions that are analogous to those presented here.

that, if we consider a unique break, the likelihood ratio statistic takes the highest value for the year 1918. This result can be easily understood if we take into account that the Bai-Perron methodology chose this period for 8 of the 17 real exchange rates. Similarly, when the presence of two breaks is considered, then the likelihood function attains its maximum when the breaks are located at periods 1918 and 1972. Again, we reject the existence of just 1 break is rejected by the LR statistic. Consequently, we pose the presence of a third break. Now, the inclusion of a third breaks implies the likelihood to attains its maximum when the breaks appear in the periods 1918, 1936 and 1972, which are closely related to those used in Taylor (2002), although they are not coincident with them. This sequential process ends when we include a fourth break. In this case, the maximum value of the likelihood is quite close to that obtained for the case where 3 breaks are considered. Therefore, the LR statistic takes a value close to 0 and, so, we should conclude that the presence of 4 breaks is not supported by the data set employed.

[10] *Table 2 about here*

Thus, the results obtained from the estimation of the joint model offer evidence in favour of the existence of a common international pattern of behaviour for our international set of real exchange rates. This pattern is similar to that employed in Taylor (2002), in the sense that 4 different regimes of exchange rates seem to be necessary to explain the evolution of this variable from a long-run perspective, although it is not fully coincident. The first regime consists of the 27 initial years of the sample (1892-1918) and is obviously ruled by the Gold Standard. The second regime is slightly shorter, covering the period (1919-1936), and is related to the interwar period. This a period dominated by the flexibility of exchange rates as well as by an attempt to recover the credibility and stability that existed during the extinguished Gold Standard. The

10 They are different combinations of locations of the four breaks, some significant and others not.

third regime (1937-1972) includes the II World War years and those years when the exchange rates are determined by the Bretton Woods system. Finally, the collapse of Bretton Woods determines the beginning of the fourth, and last period (1973-1996).

Once the different periods in the evolution of the exchange rates during the last century have been defined, we consider it interesting to analyze them in order to describe their more relevant characteristics. In particular, the determination of the different degrees of persistence in the shocks for each of these four periods could offer some important insights for understanding the evolution of exchange rates. This is the aim of the next section.

4. STATISTICAL PROPERTIES THROUGH REGIMES.

Two main conclusions arise from the previous sections. First, we have endogenously detected breaks in the RER for most of the countries analyzed, and they match with different exchange rate regimes. Second, the structural changes are similar across countries and we can establish a common international path in exchange rate regimes. In this section, we return to the univariate analysis and investigate crucial properties of RER across regimes, namely, persistence and volatility. There is an abundance of empirical evidence that the volatility of real exchange rates varies across nominal regimes, being higher in floating than in fixed regimes. Nevertheless, and despite the great academic effort in the PPP area, there is no similar consensus about the influence of monetary arrangements in the persistence properties of RER, the literature in this field being relatively scarce.

4.1 Persistence

With the aim of studying the persistence of RER, we use the half-life (HL), probably the most popular persistence measure in PPP literature, which is defined as the number of periods that a shock needs to vanish by 50 percent. Rogoff (1996) pointed out the consensus of about 3-5 year half-lives of deviations of PPP equilibrium found in empirical studies using long-horizon data sets. Starting from these results, this author described the well-known puzzle, namely, the difficulty to explain simultaneously the high volatility of real exchange rates, provoked by monetary factors, and their persistent deviations from equilibrium. Nevertheless, this finding has been obtained with long-run studies that mix different types of exchange rate regimes. Although the PPP hypothesis should hold in the long run in both regimes, fixed and flexible, it is plausible to think that the adjustment speed and, as a result, the persistence of RER, could change across regimes.

On the other hand, it has recently been argued that the reported 3-5 year half lives are underestimated, implying that the puzzle is even bigger than was first thought. This conclusion is based on the work of Andrews (1993) and Andrews and Cheng (1994) who show that traditional measures underestimate persistence because they rely on estimates of the AR parameter, which is known to be downwardly biased in finite samples and more so if they have deterministic components[11]. For example, in a Dickey-Fuller regression with only a intercept as the deterministic term and a sample size of 40, the LS estimate of the autoregressive parameter of 0.87 has a very high bias, around 10%. In other words, the median unbiased coefficient would be 0.97 with

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[11] Sowell (1990) shows that the finite sample distribution of the LS estimates of the AR parameter is very downwardly biased when the process has long memory. The possibility of modeling RER as a fractional integrated process is explored in Mayoral et al. (2006). In this case, we have discarded this approach due to the reduced sample size of each regime.

dramatic consequences for the degree of persistence of RER shocks, worsening the "*purchasing power parity puzzle*" of Rogoff (1996).

Following this approach, Murray and Papell (2002) reported new HL estimates based on median unbiased estimators, as well as confidence intervals. Although most point estimates fall within the 3-5 year range, the upper bounds are too high to rule out the failure of the PPP. Furthermore, lower bounds are too large to be compatible with models based on nominal rigidities[12].

We have corrected the bias by using the procedure of Andrews (1993) and Andrews and Cheng (1994), and we have specified the following regression for each of the four regimes

$$q_t = \mu + \rho q_{t-1} + \sum_{i=1}^{k} \phi_i \Delta q_{t-i} + u_t \qquad [3]$$

where the value of k is selected in accordance with a general-to-specific criterion with $k_{\max} = 3$, and the half-life[13] defined as

$$HL = \frac{\ln 0.5}{\ln \rho} \qquad [4]$$

Table 3 about here

The results of median unbiased estimates of half-lives are displayed in Table 3, and allow us to obtain two main findings[14]. First, the second and fourth regimes (1919-1936 and 1973-1996), associated with flexible nominal exchange rates, show lower

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12 Other attempts, such as Lopez et al. (2003, 2004), Cashin and McDermott (2003), Rossi (2003), Caporale et al. (2005) and Murray and Papell (2005) obtain similar results.

13 Cheung and Lai (2000b) show that parity adjustments are non monotonous when the process is more complex than AR(1), and the use of impulse response functions (IRF) is more suitable than HL to measure persistence. However, the fact that the HL is a scalar measure has the advantage of simplicity and comparability across countries.

14 Lopez et al (2005b), analyzing the same sample (only for the sixteen industrialized countries) and considering one regime, estimate a HL median of 11.34 years with a confidence interval at 95% of [3.02, ∞). Hegwood and Papell (1998), including a structural break, obtain significant reductions in HL for long-span samples.

shock persistence. Nevertheless, there are significant differences between both regimes, the post-Breton Woods period being slightly more persistent.

More specifically, during the inter-war regime shocks disappear after 1 year at most for 5 of the 19 real exchange rates (for 11 RER if we enlarge the time to 1 year and a half). The median value of this regime –reported in the last row of Table 3- is slightly over 1 year, and the confidence interval (at 90% level of significance) is very low, decreasing the lower bound of the implicit interval of 3-5 years underlying the consensus highlighted by Rogoff (1996). The estimated half-life of the post-Bretton Woods regime has a median of under two years but the confidence interval, although bounded, is wider[15]. Its lower bound is still less than a year but the upper bound reaches a value of 14.5 years[16]. Moreover, Argentina, Finland, France, Mexico and Switzerland show closely bounded intervals. In any case, the point estimation of persistence in both regimes is compatible with models based on nominal rigidities. There is a 50% probability that the shocks vanish at a maximum of 1.3 and 1.9 years, for the second and fourth regime respectively.

The second main finding is that the intermediate regimes, first and third, reflecting usually fixed exchange rates, exhibit significantly more persistence. The first regime, corresponding to the Gold Standard period, 1870-1918, has a median half-life

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15 Notice that the dispersion of estimated half-lives in this regime is markedly inferior, –half- of the dispersion in the second regime (excluding Germany). This is presumably because the context is more homogeneous in the first floating regime while, in the second, countries faced more different environments. Cashin and McDermott (2004), for a sample of 90 countries in the post-Bretton Woods era find different variances in the median HL between industrialized and developing countries but not across fixed and floating regimes.

Furthermore, Cheung and Lai (2000a) study a post-Bretton Woods monthly sample and show the heterogeneity of persistence for the 94 RER considered. The persistence tends to be higher in developing countries than in the industrialized ones, the opposite conclusion to that obtained by Cashin and McDermott (2004), although these authors use unbiased methods.

16 Murray and Papell (2002) study a quarterly sample for the post-Bretton Woods era and 20 industrialized countries and obtained median unbiased HL of 2.39 years with a confidence interval at 95% of $[0.74, \infty)$. When HL are calculated from the impulse response functions, the value is 3.07 and the interval is $[1.24, \infty)$.. Cashin and McDermott (2003), with effective real exchange rate, analyze a similar updated and monthly sample and report a median half-life, through IRF of 59 months, around 5 years.

still within the *consensus*: 4.3 years, but with unbounded intervals. For individuals countries, the shocks suffered by the RER of Argentina, Germany, Sweden and Switzerland, in the first regime, and Germany, Japan, Holland and Switzerland in the third, seem to have a permanent effect, hardly reconcilable with the PPP hypothesis. Furthermore, both in the first and the third regimes, the upper bound of the confidence intervals for most half-lives is infinite. This is not the case of Belgium and Denmark, in the first regime, and Finland and France in the third, because then show reduced estimated half-lives at both the lower bound and the upper bound.

Summarizing, our findings reveal that half-lives in flexible exchange rate regimes range between 1 and 2 years (which corresponds to a mean-reversion speed of 30 and 42% per year, respectively). With respect to the confidence intervals in these periods, they are narrowly bounded, disentangling the PPP puzzle. Contrarily, fixed exchange rate regimes exhibit more persistent shocks (with a mean-reversion between 8 and 15%) and have an infinite upper bound in their confidence intervals, supporting the PPP puzzle.

These conclusions contrast with those obtained by Taylor (2002) who, using an exogenous partition of the sample and LS estimates, finds the "*provocative*" and "*remarkable*" result that "*the deterministic aspects of persistence of PPP deviations have been fairly uniform in the international economy*" despite the significant institutional changes during a century. Nevertheless, our findings are very similar to those obtained recently by Sarno and Valente (2006) using non-linear techniques. Their disaggregate analysis indicates that long-run PPP holds in any type of monetary regime. The difference lies in what drives the adjustment to restore deviations from equilibrium: the relative prices in fixed exchange rate regimes or the nominal exchange rate in floating regimes. In addition, they find, like us, much more persistent deviations in fixed than in flexible regimes. Consequently, in a similar framework of sticky prices, a floating exchange rate has more capacity to return the system to equilibrium after shocks. Moreover, monetary arrangements do not seem to have been successful in reducing the inflation persistence[17].

4.2 VOLATILITY TRHOUGH REGIMES

There is a generalized belief that real exchange rates are more volatile during floating regimes due to the high variability of nominal exchange rates, and the results of Taylor (2002) analyzing four regimes over a century confirm this idea[18]. In this paper, we analyze volatility thought the estimation of a GARCH (Generalized Autoregressive Conditional Heteroskedasticity) model that allows us to relax the assumption of constant variance and introduce the concept of conditional variance, which moves in accordance with the previous information. Supposing that real exchange rate follows an AR(1)-GARCH(1,1) model, we obtain the following expression:

$$\begin{aligned} q_t &= \mu + \rho q_{t-1} + \sum \phi_i \Delta q_{t-1} + u_t \\ \sigma_t^2 &= \alpha + \beta u_{t-1}^2 + \gamma \sigma_{t-1}^2 + \varepsilon_t \end{aligned} \qquad [4]$$

where the RER is represented as an autoregressive process with an error term that has a conditional variance, with can be modeled in function of a constant, the volatility of the previous period –the ARCH term- and the last period's forecast of the conditional variance –the GARH term-. Then, the estimation of conditional variance gives a measure of the evolution of volatility during each period.

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17 In this regard, Alogoskoufis and Smith (1991), Alogoskoufis (1992) and Obstfeld (1995) have related the monetary autonomy afforded by floating regimes to the inertia of inflation series in these periods.

18 A recompilation of empirical literature in favor of this hypothesis can be bound in Frankel and Rose (1995). More recently we can see the works of Hasan and Wallace (1996) and Liang (1998) among others.

Table 3 reports the mean of conditional variance by periods and countries. We can see that this measure of volatility is low during the first and third regimes, and increases significantly in the second and fourth. This is a stylized fact that holds in most countries with the only exceptions of Canada during the second regime and Sweden and the United Kingdom during the third. However, we observe a very different behavior in the South American countries. Their path of volatility looks like a staircase, so that the mean of conditional variance increases as the regime moves forward, reaching a maximum in the last[19]. For this reason, we have excluded these countries in the calculation of the mean because they could bias our interpretation. Notice that this exclusion is especially relevant in the two last regimes.

Table 4 about here

The volatility during the Gold Standard, characterized by strict rules, was the most reduced of the century analyzed. The mean of conditional variance of the Bretton Woods regime is at a similar level. At the opposite extreme, is the volatility of floating regimes, that corresponding to the inter-war period being slightly higher than the post-Bretton Woods period. In short, our results are not very surprising with respect to the existing literature.

A more interesting issue at this point, is to relate the persistence of the shock with the volatility of RER[20]. Cashin and McDermott (2004) find a negative and statistically significant relation between half-lives and volatility across different monetary regimes. However, Kanas and Genius (2005) show an association between stationarity and low volatility regimes of real exchange rate and, vice versa, between

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19 Hausmann et al. (2004) have recently pointed out that the volatility is around three times higher in developing countries than in industrialized countries.

20 Hasan and Wallace ((1996) find that the volatility of RER is higher in floating regimes than in fixed-rate periods. They also warn that if a series is non-stationary, calculations of volatility based on the levels of the variables would be misleading because the variance increases with the time span of the data.

nonstationarity and high volatility periods. Finally, the work of Cheung and Lai (2000a) study the cross-sectional relationship between RER persistence and other variables, obtaining a negative relation with inflation and positive with public spending.

We have studied the relation between persistence and volatility in the four regimes through a simple Spearman's rank correlation coefficient, applied to half-lives and the mean of conditional variance. The calculation of the coefficient has been carried out only for industrialized countries and we have obtained a value of –0.80 ($p = 0.20$)[21] (see last rows of Tables 3 and 4). Thus, there is a strong and negative association between the two variables. To sum up, a higher volatility of real exchange rate is compatible with a faster adjustment speed to parity.

6. CONCLUSIONS.

This paper has been devoted to the analysis of the time properties of the historical real exchange rate series constructed in Taylor (2002). By applying the methods proposed in Bai and Perron (1998, 2003), we have offered evidence in favor of the non-stability of these real exchange rates during our sample period. Our results lead us to admit the existence of four clearly differentiated periods 1870-1918, 1919-1936, 1937-1972 and 1973-1996. This partition is almost coincident with that employed in Taylor (2002) and, consequently, this first result would only confirm the division carried out in that paper.

However, the new partition is relevant when we analyse the persistence of the real exchange rates during the four periods of time considered. The study of the persistence for each of the intervals shows that the persistence is lower for those periods

[21] The coefficient is only significant at the 20% significance level, but we only have two degrees of freedom. If we build the series with 105 observations, the correlation coefficient decreases slightly but the significance increases ($p < 0.001$).

dominated by a system of flexible exchange rate, whilst the adjustment to the parities are clearly slower for those periods with a fixed system of exchange rates. Moreover, we obtain that the half-life of a shock is lower than 2 years for flexible exchange rate regimes with very narrow confidence intervals, whilst this measure increases for the periods with a fixed exchange rate regime where the upper bound tends to infinity. These results contrast with some previous works, such as Lothian and McCarthy (2002), Parsley and Popper (2002) or Cheung and Lai (2000a), which suggest a lower persistence in fixed regimes. Additionally, if we relate the half-life of the shock to the volatility of the exchange rates, we observe a negative correlation, in the sense that a greater volatility of shocks is related to a lower degree of persistence.

Finally, if we take into account the different degree of persistence and volatility of the shocks during the four intervals of time considered, our results suggest the non neutrality of the monetary regimes with respect to the behaviour of the real exchange rates. Thus, we should conclude that the degree of persistence and volatility of the real exchange shocks are clearly altered by the decisions taken by the monetary authorities.

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Tables and Figures

Table 1: Bai – Perron procedure.

	T	UD max	WD max			Sup $F_T(l+1\|l)$				Break dates			
			α=0.10	α=0.05	α=0.01	l=1	l=2	l=3	l=4	T_1	T_2	T_3	T_4
AR	113	6.258	11.645*	12.270		6.331	6.397	4.691	7.298	1974 #			
AU	127	13.518**		13.518**	13.518	3.255	5.219	1.111	1.111	1916			
BE	117	68.198***		80.229**	86.272***	98.547***	10.752	4.627	6.233	1918	1935		
BR	102	15.925***		21.286**	23.709***	11.816**	4.361	20.246***		1913	1932	1952	1975
CA	127	9.763		15.506**	17.271***	8.137	11.258	2.160		1976			
DE	117	20.417***		24.019**	25.828***	5.432	5.715	2.326	2.102	1970			
FI	116	68.769***		68.769**	68.769***	43.567***	19.929***	5.203	1.456	1901	1918	1939	
FR	117	10.536*		12.986**	14.165	2.535	5.878	0.548	0.709	1913			
GE	117	8.597		13.404**	14.930	11.954**	3.646	9.153	0.950	1939	1970		
IT	117	20.992***		23.545**	25.487***	12.113**	11.166*			1922	1945	1962 #	
JA	105	6.716	7.634	7.901		8.288	3.573	3.399	3.399				
ME	111	12.015**		23.557**	26.381***	18.174***	8.558	9.621	6.335	1918	1980		
NE	127	13.539**		15.928**	17.127***	4.757	1.801	1.253	2.192	1970			
NO	127	23.567***		28.816 **	31.117***	17.829***	3.329			1918	1969		
PO	107	15.483***		24.700**	27.512***	3.779	3.427	2.458		1920			
SP	117	15.495***		21.331**	23.091***	12.049**	9.710			1918	1970		
SW	117	28.415***		28.415**	28.749***	14.377**	7.015	7.015	7.015	1918	1935		
SZ	105	10.875*		16.379**	18.342***	13.031**	8.309	1.771		1918	1970		
UK	127	13.121**		18.671**	20.796***	22.582***	12.337*	1.071		1918 #	1948	1976	

*, **, *** indicates significance at the 10, 5 and 1% levels, respectively.
# indicates a significant break date at the 10% level.

Figure 1: First break point model: log likelihood.

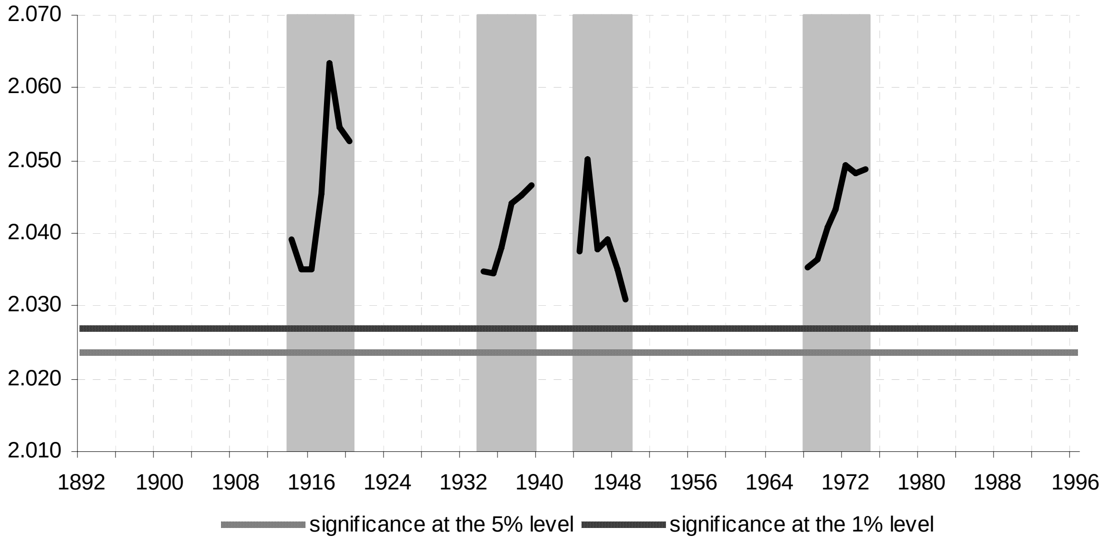

Table 2: Likelihood ratio.

Break dates				Log likelihood	Likelihood Ratio
D_1	D_2	D_3	D_4		
				1999.48	
1918				2063.35	127.74***
1918	1936			2170.65	214.60***
1918	1939			2177.38	228.06***
1918	1945			2139.38	152.05***
1918	1972			2126.47	126.23***
1918	1936	1972		2227.90	114.50***
1918	1945	1972		2185.99	93.23***
1918	1935	1944	1970	2231.96	42.60

, * Indicates significance at the 5 and 1% levels, respectively.

Table 3: Median-unbiased half-lives.

	1870-1918		1919-1936		1937-1972		1973-1996	
	HL_{MU}	90% CI	HL_{MU}	90% CI	HL_{MU}	90% CI	HL_{MU}	90% CI
AR	∞	[6.58–∞)	1.79	[0.85–10.62]	2.08	[0.93–23.97]	1.51	[0.72–6.13]
AU	2.72	[1.17–∞)	1.09	[0.52–3.13]	10.63	[2.42–∞)	2.92	[1.24–∞)
BE	1.17	[0.57–3.57]	5.06	[1.68–∞)	3.89	[1.48–∞)	2.18	[0.97–42.71]
BR			0.59	[0.19–1.34]	8.38	[2.23–∞)	8.04	[2.19–∞)
CA	4.97	[1.66–∞)	0.51	[0–1.15]	2.89	[1.23–∞)	1.86	[0.88–12.52]
DE	0.50	[0–1.12]	2.05	[0.92–21.11]	6.63	[2.00–∞)	2.13	[0.95–30.19]
FI	4.74	[1.61–∞)	0.64	[0.24–1.48]	1.34	[0.65–4.74]	1.56	[0.74–6.41]
FR	2.42	[1.06–∞)	1.23	[0.60–3.94]	1.03	[0.49–2.81]	1.70	[0.81–8.74]
GE	∞	[5.32–∞)	∞	[4.93–∞)	∞	[6.97–∞)	1.91	[0.90–13.82]
IT	3.40	[1.34–∞)	1.29	[0.63–4.39]	2.43	[1.06–∞)	1.76	[0.83–9.34]
JA	∞	[279.15–∞)	5.16	[1.82–∞)	∞	[4.70–∞)	2.28	[1.06–126.05]
ME	33.43	[19.10–∞)	2.60	[1.12–∞)	2.56	[1.11–∞)	1.51	[0.72–6.11]
NE	5.07	[1.68–∞)	3.54	[1.38–∞)	∞	[3.88–∞)	1.73	[0.82–8.83]
NO	2.91	[1.23–∞)	1.33	[0.65–4.70]	16.61	[2.77–∞)	2.15	[0.96–33.97]
PO	2.20	[0.98–47.46]	1.21	[0.59–3.81]	331.3	[3.60–∞)	4.28	[1.51–∞)
SP	3.79	[1.45–∞)	1.27	[0.62–4.21]	7.73	[2.15–∞)	2.40	[1.10–∞)
SW	∞	[3.91–∞)	0.97	[0.47–2.58]	26.06	[15.12–∞)	2.07	[0.93–23.12]
SZ	∞	[18.81–∞)	2.82	[1.20–∞)	∞	[4.33–∞)	1.63	[0.78–7.68]
UK	1.73	[0.82–8.89]	0.84	[0.39–2.07]	17.86	[2.82–∞)	1.93	[0.91–14.51]
Median	4.27	[1.53–∞)	1.29	[0.63–4.39]	8.38	[2.23–∞)	1.93	[0.91–14.51]

Notes: CI indicates confidence intervals; they are calculated with the Andrews (1993) critical point for T = 40.
Not all series star in 1870 (see Table-1). Japan in the only country where the GDP deflator has been used in the construction of RER; this series ends in 1989.
In the first regime (1870-1918), some dummy variables (additive or multiplicative) have been added, when they have been significant, trying to capture the possible exceptionality stemming from the FWW. In these cases this implies that the bias correction only corresponds to the period 1870-1913. In Argentina, Australia, Canada, Denmark, France and Japan it was not necessary to use dummy variables.

Table 4: Mean conditional variance series.

	1870-1918	1919-1936	1937-1972	1973-1996
Argentina	0.0060	0.0087	0.0406	0.3885
Australia	0.0019	0.0140	0.0019	0.0060
Belgium	0.0128	0.0218	0.0041	0.0093
Brazil		0.0140	0.0752	0.0839
Canada	0.0011	0.0006	0.0009 #	0.0019
Denmark	0.0037	0.0076	0.0032 #	0.0094
Finland	0.0044	0.0118	0.0092 #	0.0103
France	0.0037	0.0069	0.0021	0.0088
Germany	0.0006	0.0076	0.0018	0.0113
Italy	0.0016	0.0235	0.0018 #	0.0082
Japan	0.0053	0.0079	0.0025 #	0.0169
Mexico	0.0122	0.0123	0.0152 #	0.0609
Netherlands	0.0021	0.0137	0.0046	0.0087
Norway	0.0035	0.0090	0.0056	0.0083
Portugal	0.0093	0.0225	0.0054	0.0100
Spain	0.0047	0.0152	0.0049	0.0162
Sweden	0.0012	0.0066	0.0061	0.0090
Switzerland	0.0017	0.0078	0.0018	0.0130
United Kingdom	0.0007	0.0050	0.0057	0.0139
Mean	0.0043	0.0114	0.0101	0.0366
Mean*	0.0036	0.0113	0.0039	0.0101

* Indicate the exclusión of the three South-American countries.
# Indicate the exclusion of the period 1937-1951.