THE SPEED OF ADJUSTMENT TO PPP: IS THERE ANY PUZZLE?*

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1. Introduction

The logarithmic version of purchasing power parity is represented as:

$$e_t = k + p_t - p_t^* \qquad [1]$$

where e is the logarithm of the nominal exchange rate measured in units of currency A per unit of currency B, p is the logarithm of the price level in country A, p^* is the logarithm of the price level in country B and k is a constant term. It establishes that different national price levels, once converted into a common currency, should differ only by a constant term.

This relationship has been widely analysed at the empirical level[1]. The typical tools of analysis have been either unit root tests on the real exchange rate or cointegration tests among the variables entering [1]. There is large consensus on the validity of relative PPP when the period of analysis is a century or more.

The evidence is mixed when a shorter time period is analysed. For the recent floating exchange rate period, a first group of studies, mainly on the basis of unit root tests (for example, Adler and Lehman, 1983; Meese and Rogoff, 1988; Grilli and Kaminsky, 1991) cannot find any evidence in favour of PPP. More recent studies support a weak version of PPP[2] (for example, Cheung, Fung, Lai and Lo, 1995). A general feature of the latter group, is the use of system based tests of cointegration.

From the growing literature adopting a panel data framework of analysis, we don't have a clear cut evidence on the topic (for example, on the positive side, Wei and Parsley, 1995, and, on the negative one, Engel, Hendrickson and Rogers, 1997)[3].

An additional common finding in the literature is the slow speed of adjustment to equilibrium (half-life of deviations from PPP of about 4 years). This result jointly with the very high short run volatility of real exchange rates generates the so-called Rogoff's (1996) puzzle.

In the light of the frequent failure for PPP to hold in the post Bretton Woods period, some authors (Edison and Klovland 1987, Johansen and Juselius 1992, Juselius 1995, Sjoo 1995, Ott 1996, Apte, Sercu and Uppal 1996), adopt a less stringent approach allowing for the role of other economic variables in the short run dynamics of a vector autoregressive model or in a single equation with an error correction mechanism. The choice is, usually, to include asset market variables as monetary aggregates or interest rates. Most of these studies doesn't find evidence in favour of PPP, but supports a long run cointegrating relationship between the real exchange rate and the interest rate differential:

$$(e_t - p_t + p_t^*) = g(i_t - i_t^*) \qquad [2]$$

where i is the domestic interest rate and i^* is the foreign interest rate.

Equation [2] is one of the key relationships in the Dornbusch's (1976) sticky price model of exchange rate determination. It is based on the joint hypothesis that PPP holds in the long run, the uncovered interest parity (UIP) holds at all times and the expectations are rationally formed.

A relationship like [2] can also be obtained adopting the framework developed by Feenstra and Kendall (1997) in which a risk-averse profit maximising exporting firm adopts a pass-through behaviour and hedges the exchange rate uncertainty with transactions in the forward market. Within this set up, the assumption of complete pass-through behaviour generates a parity condition between the prices and the forward exchange rate. Utilising the covered interest parity (i.e. substituting the forward rate in place of the sum of the spot exchange rate and the interest rates differential) equation [2] immediately follows.

In this paper we focus on the post Bretton Woods period and use a cointegration analysis in order to analyse whether a relationship like [1] is accepted by the data for nine bilateral parities having Italy and Switzerland[4] as pivotal countries, and United States, Germany, United Kingdom and Japan as foreign countries. Our major contribution is the adoption of a more appropriate methodology to measure the speed of adjustment to PPP. It is the 'persistence profile', a system wide measure developed by Pesaran and Shin (1996). Differently from the standard approach, it does not require any strong exogeneity property of the variables involved

in PPP and provide information on the shape of the whole adjustment path. Our results reject the existence of a puzzle since the estimated speed of adjustment appears to be faster than in the previous literature.

The reliability of the inference based on the persistence profiles depends on the correct identification of the equilibrium relationship with respect to which the speed of adjustment is calculated. In the light of this, two other contributions of our paper are the adoption of a fully identifying cointegration analysis[5] and the use of a likelihood dominance criterion in order to select the optimal identifying structure for the cointegration space. In other terms, after testing for cointegration, we try to fully identify the cointegration space by imposing, on the cointegrating vectors, two competing sets of over-identifying constraints that are empirically tested: the first one allows for the restrictions suggested by [1], whereas the second is based on [2]. Adopting a dominance criterion we choose the former identification in most of the considered cases and conclude in favour of the PPP.

The rest of the paper is structured as follows. In section 2 we test for cointegration of a relationship like [1]. In section 3 we include also interest rates into the analysis to check whether they play a role either in the short or in the long run. In the fourth section, we estimate the speed of adjustment. Conclusion are contained in section 5[6].

2. Does the PPP hold alone?

To test the PPP as a long-run stationary relationship we adopt the full information maximum likelihood (FIML) cointegration approach developed by Johansen (1995)[7]. We start from a "small" VAR specification including price and exchange rate variables[8]. Given X_t^o (p_t, p_t^*, e_t), the vector error correction representation of the "small" VAR has the following form:

$$\Delta \mathbf{X}_t = \delta + \psi \mathbf{K}_t + \gamma D_t + \mathbf{\Phi}_1 \Delta \mathbf{X}_{t-1} + \mathbf{\Phi}_2 \Delta \mathbf{X}_{t-2} + \ldots\ldots + \mathbf{\Phi}_{p-1} \Delta \mathbf{X}_{t-p+1} + \Pi \mathbf{X}_{t-1} + \boldsymbol{\varepsilon}_t \quad [3]$$

where δ is a 3×1 vector of unrestricted intercept terms describing the presence of a drift in the level of the series, **K** is a 3×1 vector of seasonal dummies, D_t is an intervention dummy controlling for a break[9] located in 1982:3, ψ is a 3×3 matrix, γ is a 3×1 vector, the $\mathbf{\Phi_i}$'s are 3×3 matrices, $\varepsilon_t \sim$i.i.d. N(0,Ω) and, under the cointegration hypothesis, the 3×3 matrix Π can be factorised as $\Pi = \alpha\beta'$ where α and β are $3 \times r$ matrices of rank r£3. The matrix β contains the r cointegrating vectors, while the matrix α contains the so-called factor loadings characterising the short run adjustment to the equilibrium.

In testing for cointegration rank r we take into account that the critical values for the trace test depend on the specification of the deterministic part of the VAR. The tabulated values by Osterwald-Lenum (1992) are not suitable for a model allowing for an intervention dummy; hence, we simulate the asymptotic correct critical values[10]. Moreover, since with small samples (we have 76 observations) the empirical size of the test is greater than the theoretical one, biasing test toward finding too many cointegrating vectors (Reimers, 1992; Gregory, 1994), we obtain finite sample critical values adopting the Reimers (1992) correction to the asymptotic values. The corrected critical values are reported in table 1 with the results of the cointegration rank trace tests[11]. The evidence is in favour of one cointegrating relationship in the Italy/Us, Italy/Germany, Italy/Switzerland and Italy/UK cases; two cointegrating relationships are found in the Italy/Japan, Switzerland/Germany, Switzerland/UK and Switzerland/Japan cases, while in the Switzerland/Us case the hypothesis of cointegration rank equal to zero, against the alternative of 3, cannot be rejected (i.e. there is no cointegration).

The finding of cointegration is only a necessary condition for PPP; in addition proportionality and symmetry condition should be satisfied. The first step of the Johansen's approach doesn't allow any conclusion in terms of the nature of the estimated cointegrating vectors. In fact, they define only a basis of the cointegration space, empirically indistinguishable from another one obtained with an alternative factorisation of $\mathbf{\Pi}$ matrix. In other words, the estimated $\boldsymbol{\beta}$ matrix describes an exactly identified structure based on r^2 constraints that usually don't have an economic interpretation[12]. In order to verify whether the $\boldsymbol{\beta}_\mathbf{i}$'s (the rows of $\boldsymbol{\beta}'$) satisfy economically meaningful restrictions, Johansen (1995) suggested to define a full set of over-identifying restrictions that constrain all the $\boldsymbol{\beta}_\mathbf{i}$ vectors (i.e. the basis of the cointegration space) and can be empirically tested. These restrictions can be expressed in explicit form as:

$$\boldsymbol{\beta} = (\mathbf{H}_1 \boldsymbol{\varphi}_1, \ldots, \mathbf{H}_r \boldsymbol{\varphi}_r) \qquad [4]$$

where $\mathbf{H}_\mathbf{i}$ is a $n \times s_i$ matrix, n is the number of endogenous variables of the VAR, s_i is the number of free parameters in $\boldsymbol{\beta}_\mathbf{i}$ and $\boldsymbol{\varphi}_\mathbf{i}$ is a vector of parameters to be estimated.

Johansen proposes an LR test to check the empirical plausibility of these restrictions; if not rejected, then the standard necessary and sufficient rank conditions for identification have to be controlled.

In our framework, when the detected rank is equal to one, the $\mathbf{H}_\mathbf{1}$ matrix, imposing the constraints suggested by [1], is the following:

$$\mathbf{H}_1 = \begin{bmatrix} 1 \\ -1 \\ -1 \end{bmatrix} \qquad [5]$$

When the trace test suggests *r=2*, the identification of the cointegrating vectors is based on the previous $\mathbf{H_1}$ and a $\mathbf{H_2}$ matrix defined as:

$$\mathbf{H_2} = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}, \qquad [6]$$

imposing on the second β vector a zero value for the foreign price coefficient and leaving the other parameters unconstrained[13]. The identification analysis results (table 2) show that PPP is empirically rejected by the LR tests in all but three cases: Italy/Switzerland, Italy/Japan and Switzerland/Japan. It seems correct to conclude that in the post-Bretton Woods period, the evidence supporting the standard PPP hypothesis, in this "small" VAR framework, is quite weak and occurs only for currencies that float against each other.

3. The augmented system: allowing for interest rate differential

Given the previous not encouraging evidence and following recent studies (Johansen and Juselius 1992, Juselius 1995, Sjoo 1995), we perform the analysis also within VAR models enlarged to include domestic and foreign interest rates. The aim of this is to take into account the potential long and short run interaction between good markets and asset markets; in this framework the testable cointegrated relationships could be both [1] and [2]. In both cases the interest rates are allowed to play a role in the short run. The specified VAR has the same form as in [3], but now $X_t \, º \, (p_t , p_t^* , e_t , i_t , i_t^*)$.

The cointegration rank trace tests[14] (table 3) show the existence of two cointegrating vectors for the bilateral case Italy/Switzerland, while four cointegrating vectors are found for Italy/Japan, Italy/UK, Switzerland/Japan and Switzerland/UK. In all the previous cases the test gives the same results both at the 5% and at the 10% significance level. The Switzerland/Us, Switzerland/Germany, Italy/Germany and Italy/Us cases show a mixed evidence: in the first case the test concludes in favour of the null hypothesis of no-cointegration at the 5% level and suggests the existence of one vector at the 10% level. Given the low power of these tests in small samples, Dickey and Rossana (1994) suggest the use of the 10% level critical value. On these basis, we conclude in favour of rank one. Adopting the same strategy, we choose rank two

for the Switzerland/Germany case and three for the Italy/Germany case. Finally, for the last case four vectors are found.

In our "large VAR" the plausible equilibrium relationships are both [1] and [2]. For this reason, there are two competing sets of overidentifying restrictions. The first overidentifying structure is characterised by the following $\mathbf{H_1}$ matrix:

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad [7]$$

imposing the symmetry and proportionality restrictions implied by PPP; the second over-identifying set of constraints has the following $\mathbf{H_1}$ matrix:

$$\mathbf{H}_1^{\mathrm{b}} = \begin{bmatrix} 1 & 0 \\ -1 & 0 \\ -1 & 0 \\ 0 & 1 \\ 0 & -1 \end{bmatrix} \qquad [8]$$

that forces the first cointegrating vector to follow [2][15].

In case the competing over-identifying structures satisfy the generic and empirical identification conditions, we discriminate between them using the Likelihood Dominance Criterion by Pollak and Wales (1991). The idea is that, given two non-nested hypothesis regarding the specification of the cointegration space, one can select the dominant one by simply comparing their associated adjusted likelihood values[16]. The results of the over-identification tests and of the application of Pollack and Wales criterion are reported in table 4. In six out of nine cases (Italy/Us, Italy/Germany, Italy/Switzerland, Italy/Japan, Switzerland/Japan, Switzerland/UK) we find that a set of constraints based on PPP (matrix H_1 in [7]) formally over-identifies the cointegration space and satisfies the empirical LR tests.

In two out of these cases favourable to PPP (Italy/Germany and Italy/Switzerland), also a relationship like [2] can validly represent an equilibrium relationship. However, it seems dominated by [1] on the basis of the dominance criterion. For the remaining four cases, various attempts have failed to find a set of overidentifiyng restrictions implying [2]. The set of restrictions reported in the Appendix only exactly identifies the cointegration space. As a consequence we cannot use the dominance criterion.

For the bilateral cases of Switzerland/Us, Switzerland/Germany and Italy/UK both PPP and the relationship described in [2] are rejected.

In summary, we found evidence in favour of PPP in six out of nine cases for the post-Bretton Woods period. These results are in line with the evidence obtained by some other recent studies that allow for the possibility of one or more structural breaks either in the context of a univariate analysis of the real exchange rate (Perron and Vogelsang 1992; Enders and Lee, 1997; Wu 1997) or in a multivariate framework (Jorion and Sweeney, 1996). Differently from these studies we allow the interest rates to influence the short run dynamics[17].

4. The speed of adjustment to PPP: is there any puzzle ?

Rogoff (1996) compares two results commonly obtained by the empirical literature: the very high short run volatility of real exchange rates and the very low estimated speed of adjustment to PPP. The former stylised fact is usually explained on the basis of monetary and financial shocks. Under this condition and in presence of sticky prices we don't expect to find a very fast adjustment to equilibrium; however, the estimated consensus speed of adjustment (half-life of three to five years; Froot and Rogoff , 1995) is too slow to be explained by nominal rigidities. For this reason, part of the literature on PPP advocates real shocks to productivity and/or preferences as essential elements in the explanation of the latter stylised fact (Rogoff, 1996).

Our proposed solution to the puzzle points to a different direction: we argue that the previous empirical evidence could be "biased" by the choice of the methodology adopted in order to measure the speed of adjustment to PPP. Most of the existing empirical studies extracts information on it looking at the size of the estimated factor loading in the context of a single equation error correction approach to cointegration. The measures obtained in this way are not fully satisfying for two main reasons: firstly they are obtained within a framework in which all the (potentially important) system wide short run interactions among the variables involved in PPP during the adjustment process are omitted. This approach is correct only if the right hand side variables in the estimated equation are strongly exogeneous. Secondly, all the synthetic measures, as the median lag, do not provide any information on the whole shape of the adjustment path and represent sufficient statistics only when this one can be described by a straight line.

In the light of the previous remarks, in order to measure the speed of adjustment to PPP, we compute the scaled persistence profiles[18] proposed by Lee and Pesaran (1993) and Pesaran and Shin (1996). Within this approach no assumption is required with respect to the exogeneity *status* of the variables; moreover, this measure describe the full dynamics of the adjustment over

the selected simulation horizon. The persistence profiles are derived from the FIML estimation of the vector error correction models used in the previous section for the cointegration analysis and provide the time evolution of the responses of a cointegrated equilibrium relationship to system-wide shocks; differently from the impulse response functions, they are unique and do not depend on the specifically defined shocks orthogonalization procedure.

The scaled s-th element of the persistence profile matrix associated to the i-th cointegrating relationship, $\boldsymbol{\beta}_i$, is[19]:

$$\mathbf{P}_i(s) \equiv \boldsymbol{\beta}_i' \mathbf{H}_s \boldsymbol{\Omega} \mathbf{H}_s' \boldsymbol{\beta}_i / \boldsymbol{\beta}_i' \boldsymbol{\Omega} \boldsymbol{\beta}_i \qquad \text{for } \mathbf{s}=0,1,2,....,40$$

where *s* represents the horizon at which we evaluate the persistence of a shock occurred in *t-s*, $\boldsymbol{\Omega}$ is the variance/covariance matrix of the VAR innovations and the $n \times n$ matrix $\mathbf{H}_s$ describes a linear combination of the matrices containing the autoregressive coefficients of the original VAR.

As underlined by Pesaran and Shin, the indications coming from the estimation of the persistence profiles make sense only if the cointegration space has been previously identified in a proper way in order to isolate one or more economically meaningful relationships. This fact underlines the importance of the accurate identification exercise we performed in section 3.

The graphs contained in figures 2a to 2c present the persistence profiles for the bilateral cases for which PPP holds in the previous section[20]. As a matter of comparison we also report in figures 3a to 3c the univariate impulse response functions (henceforth IRFs) of the real exchange rates to respect to its own shocks[21]. The evidence coming from the graphs of the persistence profiles clearly points to a speed of adjustment which is higher than the one usually[22] obtained in the empirical literature: half life, defined as the number of quarters needed in order to absorb half the initial unit shock, is never larger than 7 quarters and in four out of six cases (Italy/Germany, Italy/Switzerland, Italy/Japan, Switzerland/Japan) the median lag is less than one year. Corresponding to this half life homogeneity, a relevant heterogeneity arises if we look at the longer horizons behaviour: for example, in the Italy/Germany case, 90% of the shock disappears approximately after 18 quarters, while in the Italy/Switzerland and Switzerland/Japan cases this adjustment takes place within only five or six quarters. This highlights the importance of having information on the whole shape of the adjustment path and not only a synthetic measure of it. On the other side, the comparison with the IRFs confirms that for all cases the half life suggested by graphs 4a to 4c is at least two times the one described by the persistence profiles. The discrepancy is particularly evident for the Italy/Germany (13 quarters on the IRF basis, one quarter as for the persistence profile), the Italy/Switzerland (10

quarters against 3) and the Italy/Japan case (10 against 2). Also the longer horizons adjustments appear to be very sticky, since they occur from 13 (Switzerland/UK case) to 42 quarters (Italy/Germany case) in order to dissipate the 90% of the shock. Anyway, all the univariate IRFs converge to zero, but the period required for convergence is so long (in some cases more than 60 quarters) that it is not surprising that the usual unit root tests do not provide any support to PPP.

In relation to whole shape of the adjustment, another feature is the frequent lack of monotonicity of the persistence profiles. In three cases (Italy/Us, Switzerland/Japan and Switzerland/UK), the plotted profile starts increasing for some quarters after the shock and then it monotonically decreases up to the final adjustment. This inverted U-shape is obtained also by Pesaran and Shin (1996) and, with a different approach, by Clarida and Gali (1994)[23]. A possible explanation refers to the overshooting of the nominal exchange rate in the context of a sticky-price environment. Another rationale (Pesaran and Shin 1996) lies in the J-effect characterising the adjustment path of the current account in presence of monetary shocks. Moreover, the lack of monotonicity seems to generate a kind of cyclical adjustment in Italy/Germany and Italy/Japan cases, where the persistence profile is characterised by frequent and persistent inversions of its slope. Non-linearity in the adjustment of the real exchange rates could be the determinant of such a cyclical behaviour as showed also by Obstfeld and Taylor (1997).

A quite different picture it does emerge from the univariate IRFs: the starting J-effect still remains in the same bilateral cases as before, but there is a weak evidence of it also in the Italy/Japan case. From the other side, having totally disregarded the structure of instantaneous and lagged correlations among prices and exchange rates, the univariate IRFs are not able to find the cyclical behaviour affecting the medium-long run evolution of the persistence profiles. The IRFs show an absolutely monotonic evolution throughout the simulation horizon.

5. Conclusions

In this paper we focused on the Rogoff's puzzle on the very high short run volatility of real exchange rates and the very low estimated speed of adjustment to PPP. At first we tried to identify in a proper way the PPP as a cointegrated equilibrium relationship for a set of nine bilateral cases having Italy and Switzerland as pivotal countries. Starting from the FIML estimates of a set of error correction VAR models, on the basis of a dominance criterion we concluded in favour of PPP in six out of the nine cases. As a second step of the analysis we measured the speed of adjustment to PPP. Differently from most of the previous attempts,

mainly based on single equation approaches and on synthetic persistence measures, we adopted the Pesaran and Shin (1996) persistence profiles. We did not find any evidence in favour of the puzzle: shocks to PPP are relatively quickly absorbed and the median lag never exceeds seven quarters. Some cyclical and non linear patterns make the adjustment a bit more sticky in some cases; their interpretation, besides the quite simple hypothesis of the existing literature (J-effect, overshooting mechanism), can trigger future research.

Appendix 1: The data and their univariate properties

All variables are quarterly sampled for the period 74:1, 92:4. For all countries we use consumer price indexes (P_i) and three-month treasury bills interest rates (I_i)[24]; the exchange rates are spot bilateral rates (E_{ij}) with two pivotal currencies: the Italian Lira and the Swiss Franc[25]. Prices and exchange rates are in logarithms.

Within the Johansen approach, preliminary tests of unit root are not necessary if one has strong a-priori that the analysed variables have at most one unit root. If one is uncertain about the existence of a second unit root in the level of the series, then unit root tests should be performed. The latter is our case, mainly because of the price variables.

The results of the Augmented Dickey-Fuller (ADF) test are presented in tables 5 and 6. The testing strategy is the general to particular one in terms of the treatment of deterministic nuisance parameters (Perron 1988). Given the small sample size we use the finite-sample critical values tailored to different lag orders in the 'augmented' part of the ADF test calculated from the response surface analysis of Cheung and Lai (1995).

There is clear evidence pointing to the presence of one unit root in all variables. Moreover, for all price series but the Swiss one, the null of a second unit root cannot be rejected at a significance level of 5% (ADF test on first differences, table 6). However, the presence of a second unit root is rejected if we adopt the SM (Schmidt and Phillips) test or the PP (Phillips and Perron) test. This ambiguity is common in the literature on unit root tests on price series[26]. However, the evidence in favour of a second unit root might be due to a structural break. In fact, there is some evidence of a break in the price series toward the end of 1982. (more precisely the series show a broken drift, with a slope change). This might be interpreted as a consequence of the beginning of a period of lower inflation in the industrialised economies, due to the stabilisation after the two oil price crisis[27].

The results of the unit root tests performed on the first difference of each price series (table 6, AO-ADF) after controlling for the structural break[28] reject the I(2) hypothesis for all price series with the exception of those for Germany and the Us. Nonetheless, we decide in favour of the existence of a single unit root for all series on the basis of results obtained in the cointegration analysis. In fact, further evidence arises from the eigenvalues of the companion matrix of the "large" models (table 7): after imposing the cointegrating rank detected without the allowance for the structural break, there seem to be more unit roots than suggested by cointegration analysis.

Appendix 2

Here we show the $\mathbf{H_i}$ matrices defining the overidentifying constraints imposed on the cointegrating vectors estimated within the "large" VAR models. For each bilateral case, two competing sets of restrictions are reported: the first one (superscript a) is based on equation [1], the second one (superscript b) on equation [2].

Italy/US case:

$$\mathbf{H}_1^a = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^a = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_3^a = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_4^a = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}$$

Italy/Germany case:

$$\mathbf{H}_1^a = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^a = \begin{bmatrix} 0 & 0 & 0 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ 0 & 0 & 0 \end{bmatrix} \qquad \mathbf{H}_3^a = \begin{bmatrix} 0 & 0 & 0 \\ 1 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix}$$

$$\mathbf{H}_1^b = \begin{bmatrix} 1 & 0 \\ -1 & 1 \\ -1 & 0 \\ 0 & 1 \\ 0 & -1 \end{bmatrix} \qquad \mathbf{H}_2^b = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_3^b = \begin{bmatrix} 0 & 0 \\ 0 & 0 \\ 1 & 0 \\ 0 & 1 \\ 0 & 0 \end{bmatrix}$$

Italy/Switzerland case:

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{a}} = \begin{bmatrix} 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$$

$$\mathbf{H}_1^{\mathrm{b}} = \begin{bmatrix} 1 & 0 \\ -1 & 0 \\ -1 & 0 \\ 0 & 1 \\ 0 & -1 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{b}} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix}$$

Italy/Japan case:

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{a}} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_3^{\mathrm{a}} = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_4^{\mathrm{a}} = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}$$

Italy/UK case:

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{a}} = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_3^{\mathrm{a}} = \begin{bmatrix} 0 & 0 \\ 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_4^{\mathrm{a}} = \begin{bmatrix} 0 & 0 \\ 0 & 0 \\ 1 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}$$

Switzerland/US case:

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_1^{\mathrm{b}} = \begin{bmatrix} 1 & 0 \\ -1 & 0 \\ -1 & 0 \\ 0 & 1 \\ 0 & -1 \end{bmatrix}$$

Switzerland/Germany case:

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{a}} = \begin{bmatrix} 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$$

$$\mathbf{H}_1^{\mathrm{b}} = \begin{bmatrix} 1 & 0 \\ -1 & 1 \\ -1 & 0 \\ 0 & 1 \\ 0 & -1 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{b}} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix}$$

Switzerland/Japan case:

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{a}} = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_3^{\mathrm{a}} = \begin{bmatrix} 0 & 0 \\ 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_4^{\mathrm{a}} = \begin{bmatrix} 0 & 0 \\ 0 & 0 \\ 1 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}$$

Switzerland/UK case

$$\mathbf{H}_1^{\mathrm{a}} = \begin{bmatrix} 1 \\ -1 \\ -1 \\ 0 \\ 0 \end{bmatrix} \qquad \mathbf{H}_2^{\mathrm{a}} = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_3^{\mathrm{a}} = \begin{bmatrix} 0 & 0 \\ 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \end{bmatrix} \qquad \mathbf{H}_4^{\mathrm{a}} = \begin{bmatrix} 0 & 0 \\ 0 & 0 \\ 1 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}$$

Tables

TABLE 1: Cointegration rank tests: "small" VAR models

	Null Hypoth	Statist.	Simulated Critical Value 5%	Simulated Critical Value 10%	Rank
	r ≤ 0	45.51	29.16	26.40	
Italy/US (2)	r ≤ 1	9.18	10.38	8.44	1
	r ≤ 2	0.70	#### *	#### *	
	r ≤ 0	34.26	29.16	26.40	
Italy/Germany (2)	r ≤ 1	9.46	10.38	8.44	1
	r ≤ 2	0.35	#### *	#### *	
	r ≤ 0	35.84	30.56	27.67	
Italy/Switzerland (3)	r ≤ 1	6.72	10.88	8.85	1
	r ≤ 2	0.11	#### *	#### *	
	r ≤ 0	61.89	30.56	27.67	
Italy/Japan (3)	r ≤ 1	15.95	10.88	8.85	2
	r ≤ 2	3.01	#### *	#### *	
	r ≤ 0	49.00	30.56	27.67	
Italy/UK (3)	r ≤ 1	10.16	10.88	8.85	1
	r ≤ 2	0.86	#### *	#### *	
	r ≤ 0	25.04	30.56	27.67	
Switzerland/US (3)	r ≤ 1	5.02	10.88	8.85	0
	r ≤ 2	0.04	#### *	#### *	
	r ≤ 0	36.74	29.16	26.40	
Switzerland/Germany (2)	r ≤ 1	12.84	10.38	8.44	2
	r ≤ 2	0.02	#### *	#### *	
	r ≤ 0	47.53	29.16	26.40	
Switzerland/Japan (2)	r ≤ 1	21.07	10.38	8.44	2
	r ≤ 2	0.73	#### *	#### *	
	r ≤ 0	45.75	30.56	27.67	
Switzerland/UK (3)	r ≤ 1	20.53	10.88	8.85	2
	r ≤ 2	0.17	#### *	#### *	

Notes:
- lag order of the VAR in brackets
- the critical values are obtained by simulation with the package DisCo.
- for all cases the adopted specification includes an unrestricted constant.

* DisCo can perform the simulation only if the number of unrestricted deterministic components in the model is at most equal to the minimum between the number of common trends (n-r) and the number of endogenous variables in the VAR(n). In these cases the minimum is n-r=1 and we can't obtain the critical values for a VAR with two unrestricted components (the intercept term and the break dummy).

TABLE 2: LR tests on overidentifying restrictions: “small” VAR models

	Chi-square test	**p-value**	**Inference**
Italy/US	23.32 (2)	0.0000086	rejected
Italy/Germany	6.81 (2)	0.033	rejected
Italy/Switzerland	2.10 (2)	0.350	not rejected
Italy/Japan	2.94 (1)	0.086	not rejected
Italy/UK	19.32 (2)	0.000064	rejected
Switzerland/Germany	8.82 (2)	0.012	rejected
Switzerland/Japan	0.93 (1)	0.333	not rejected
Switzerland/UK	5.13 (1)	0.02	rejected

Notes:
- we reject the null hypothesis when the p-value is less than 0.05
- degrees of freedom in brackets.

TABLE 3: Cointegration rank tests: “large” VAR models

	Null Hypothesis	Statistic	Simulated Critic Value 5%	Simulated Critical Value 10%	Rank
	r ≤ 0	133.25	81.40	76.73	
	r ≤ 1	64.15	54.62	50.63	
Italy/US (2)	r ≤ 2	28.53	30.99	28.05	4
	r ≤ 3	11.92	11.03	8.97	
	r ≤ 4	0.21	#### *	#### *	
	r ≤ 0	127.69	97.48	91.88	
	r ≤ 1	71.42	65.41	60.63	
Italy/Germany (4)	r ≤ 2	32.94	37.10	32.59	3
	r ≤ 3	7.08	13.21	10.74	
	r ≤ 4	1.26	#### *	#### *	
	r ≤ 0	129.24	97.48	91.88	
	r ≤ 1	72.26	65.41	60.63	
Italy/Switzerland (4)	r ≤ 2	32.25	37.10	32.59	2
	r ≤ 3	7.03	13.21	10.74	
	r ≤ 4	0.00	#### *	#### *	
	r ≤ 0	132.54	97.48	91.88	
	r ≤ 1	80.44	65.41	60.63	
Italy/Japan (4)	r ≤ 2	46.48	37.10	32.59	4
	r ≤ 3	19.10	13.21	10.74	
	r ≤ 4	7.36	#### *	#### *	
	r ≤ 0	126.19	81.40	76.73	
	r ≤ 1	77.27	54.62	50.63	
Italy/UK (3)	r ≤ 2	35.94	30.99	28.05	4
	r ≤ 3	14.53	11.03	8.97	
	r ≤ 4	1.34	#### *	#### *	
	r ≤ 0	93.67	97.48	91.88	
	r ≤ 1	43.91	65.41	60.63	
Switzerland/US (4)	r ≤ 2	23.16	37.10	32.59	1
	r ≤ 3	8.08	13.21	10.74	
	r ≤ 4	0.02	#### *	#### *	
	r ≤ 0	95.08	81.40	76.73	
	r ≤ 1	53.17	54.62	50.63	
Switzerl/Germ (2)	r ≤ 2	24.72	30.99	28.05	2
	r ≤ 3	5.76	11.03	8.97	
	r ≤ 4	0.98	#### *	#### *	
	r ≤ 0	106.23	81.40	76.73	
	r ≤ 1	67.25	54.62	50.63	
Switzerl/Japan (2)	r ≤ 2	37.26	30.99	28.05	4
	r ≤ 3	14.49	11.03	8.97	
	r ≤ 4	0.40	#### *	#### *	
	r ≤ 0	111.84	81.40	76.73	
	r ≤ 1	62.13	54.62	50.63	
Switzerland/UK (2)	r ≤ 2	33.47	30.99	28.05	4
	r ≤ 3	14.52	11.03	8.97	
	r ≤ 4	0.35	#### *	#### *	

Notes:- lag order of the VAR in brackets
- for all cases the adopted specification includes an unrestricted constant.

TABLE 4: Empirical identification of the cointegration space : Likelihood Ratio tests

	Relationship defined by H_1 matrix [a]	Likelihood of the restricted model	Degrees of freedom	Test statistic	p-value	Inference	Dominant overiden-tifying structure
Italy / US	[7]	1386.61	1	χ^2 =1.92	0.170	not rej.	[7]
	[8]	1387.57 [b]	--	--	--	not rej.	
Italy / Germany	[7]	1485.04	2	χ^2 =5.49	0.064	not rej.	[7]
	[8]	1484.60	3	χ^2 =6.37	0.095	not rej.	
Italy / Switzerland	[7]	1456.45	3	χ^2 =3.25	0.350	not rej.	[7]
	[8]	1455.91	3	χ^2 =4.33	0.230	not rej.	
Italy /Japan	[7]	1378.91	1	χ^2 =2.89	0.090	not rej.	[7]
	[8]	1380.35 [b]	--	--	--	not rej.	
Italy / UK	[7]	1333.80	1	χ^2 =13.9	0.000	rejected	
	[8]	1340.77 [b]	--	--	--	not rej.	[7]
Switzerland / US	[7]	1311.51	4	χ^2 =40.3	0.000	rejected	
	[8]	1324.28	3	χ^2 =14.7	0.002	rejected	
Switzerland / Germany	[7]	1388.09	3	χ^2 =12.1	0.007	rejected	
	[8]	1388.12	3	χ^2 =12.1	0.007	rejected	
Switzerland / Japan	[7]	1309.60	1	χ^2 =0.87	0.350	not rej.	[7]
	[8]	1310.03 [b]	--	--	--	not rej.	
Switzerland / UK	[7]	1309.56	1	χ^2 =2.24	0.134	not rej.	[7]
	[8]	1310.68 [b]	--	--	--	not rej.	

Notes:
- [a] [7] means that $\mathbf{H_1}$ matrix is described by [7]; the same for [8]
- [b] in these cases it does not exist any set of constraints containing [8] that overidentifies the cointegration space; all the plausible restrictions structures produce only exact identification and they don't need to be empirically tested.
- [c] we perform the tests also with different alternative specifications of $\mathbf{H_j}$, $j=2,..r$: the results are qualitatively the same.

TABLE 5: Unit root tests

Series	ADF test	Critic. values at 5% level*
p_{ita}	-2.61 (2)	-2.89
p_{us}	-1.59 (3)	-2.88
p_{ger}	-0.44 (4)	-2.87
p_{swi}	0.78 (3)	-2.88
p_{uk}	-2.50 (5)	-2.86
p_{jap}	-2.03 (4)	-2.87
e_{itus}	-2.14 (0)	-2.91
e_{gerus}	-1.17 (0)	-2.91
e_{itager}	-2.36 (0)	-2.91
e_{itajap}	-2.81 (0)	-2.91
e_{itauk}	-2.08 (0)	-2.91
e_{itaswi}	-2.51 (0)	-2.91
e_{swius}	-1.71 (1)	-2.90
e_{swiger}	-2.85 (1)	-2.90
e_{swiuk}	-2.09 (0)	-2.91
e_{swijap}	-1.72 (1)	-2.90
i_{ita}	-1.93 (0)	-2.91
i_{us}	-0.47 (2)	-2.89
i_{ger}	-1.59 (1)	-2.90
i_{swi}	-1.76 (0)	-2.91
i_{uk}	-2.41 (0)	-2.91
i_{jap}	0.13 (1)	-2.90

Notes:
-the adopted specification is the one containing a costant .in brackets the number of lags included
* simulated by Cheung and Lai (1995)

TABLE 6: Unit root tests

	ADF test *	PP test	SM test	AO-ADF test **
Δp_{ita}	-2.50 (1)	-3.40 (5)	-3.2 (5)	-5.64 (0)
Δp_{us}	-2.00 (2)	-3.27 (5)	-3.1 (5)	-2.88 (2)
Δp_{ger}	-1.79 (3)	-5.46 (5)	-5.54 (5)	-1.96 (3)
Δp_{swi}	-3.84 (2)	-5.27 (5)	-5.22 (5)	-----
Δp_{uk}	-1.99 (3)	-5.22 (5)	-5.21 (5)	-3.82 (3)
Δp_{jap}	-2.66 (4)	-6.45 (5)	-6.79 (5)	-4.48 (1)

Notes:
* for the critical values see table 1.
** 5% critical value for a break fraction $0.4<\lambda<0.6$ is -3.35 (table 4 in Perron, 1990)

TABLE 7: Eigenvalues of companion matrix Italy/US and Italy/Germany

Italy/US (without step dummy)	Italy/Germany (without step dummy)
± 0.99	1.01
± 0.97	0.97
± 0.61	± 0.93
0.38	± 0.62
± 0.21	0.30
0.006	0.21
	0.09
	0.06
Detected Rank	
3	2

FIGURE 1a: Real exchange rates

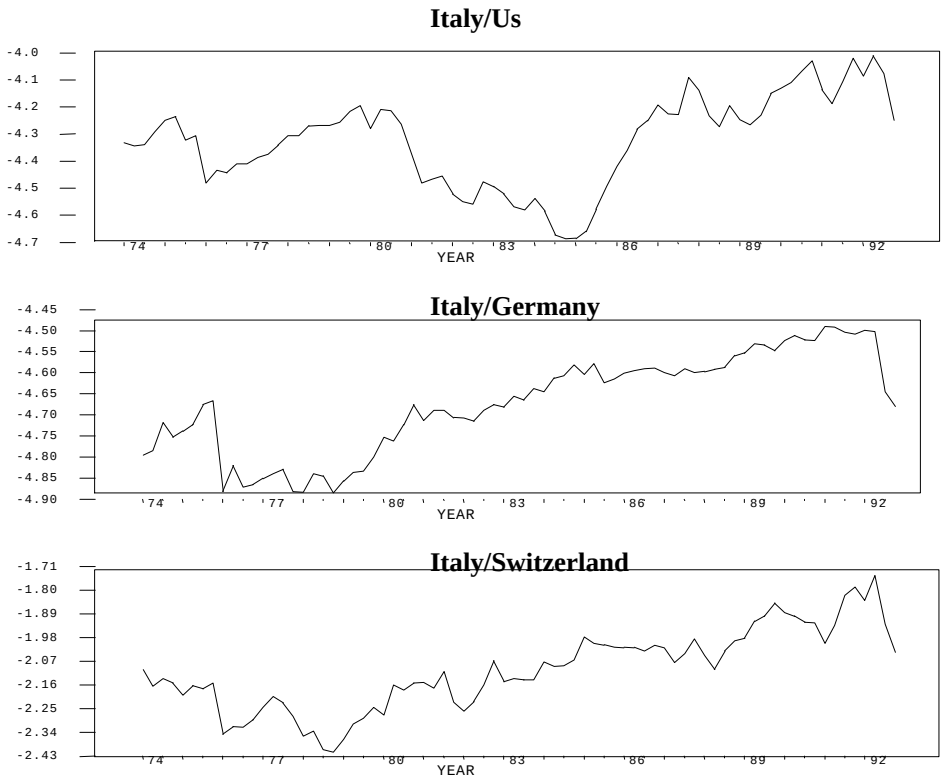


FIGURE 1b: Real exchange rates

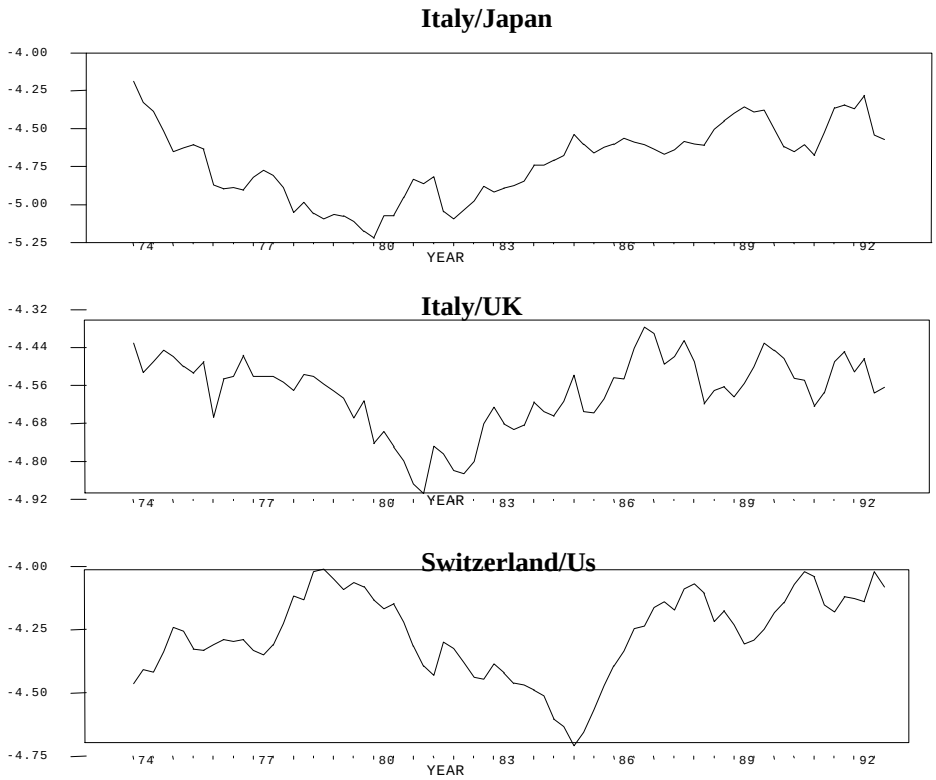


FIGURE 1c: Real exchange rates

A

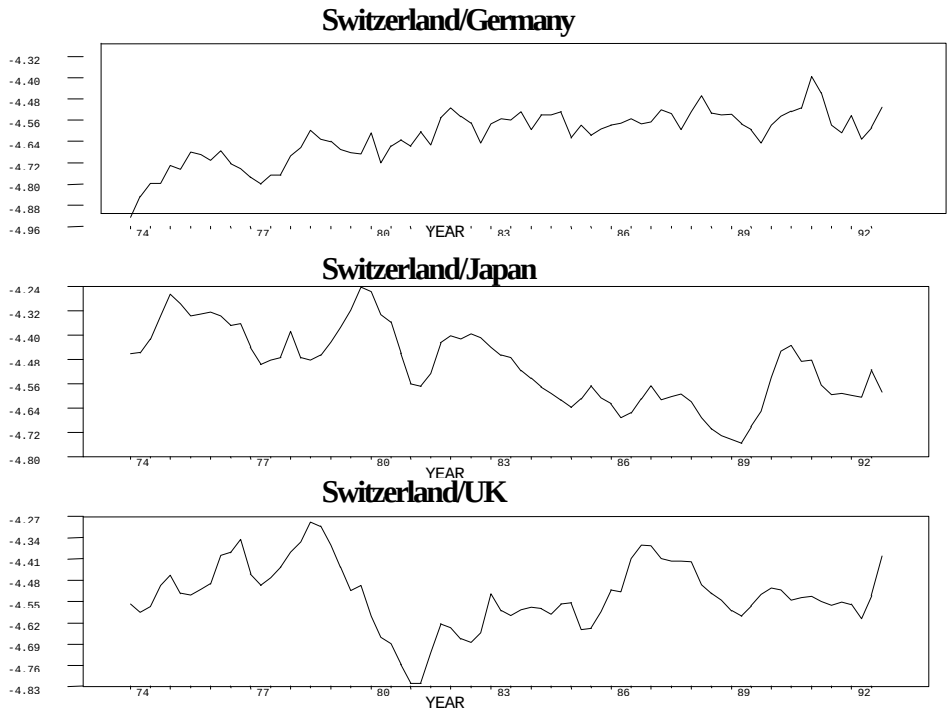
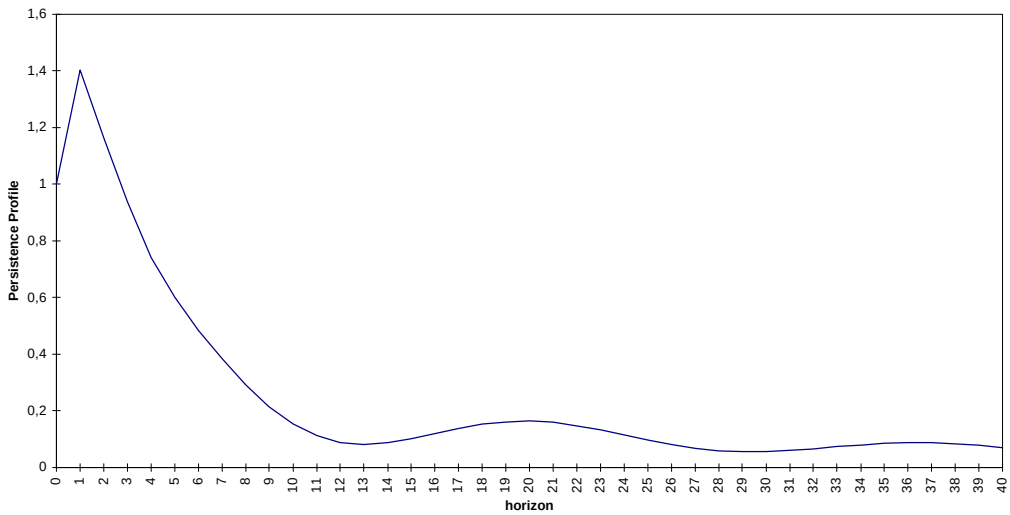


FIGURE 2a: Persistence Profiles

Italy/Us



Italy/Germany

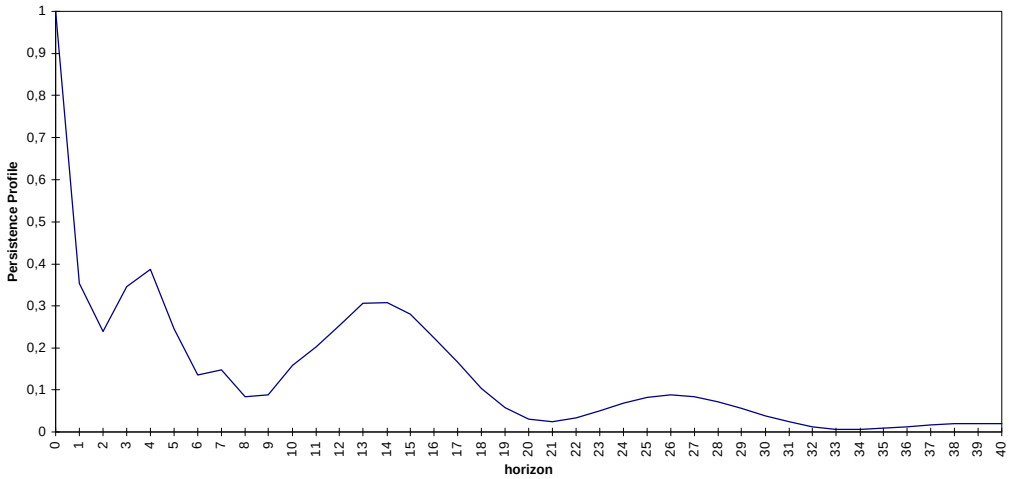
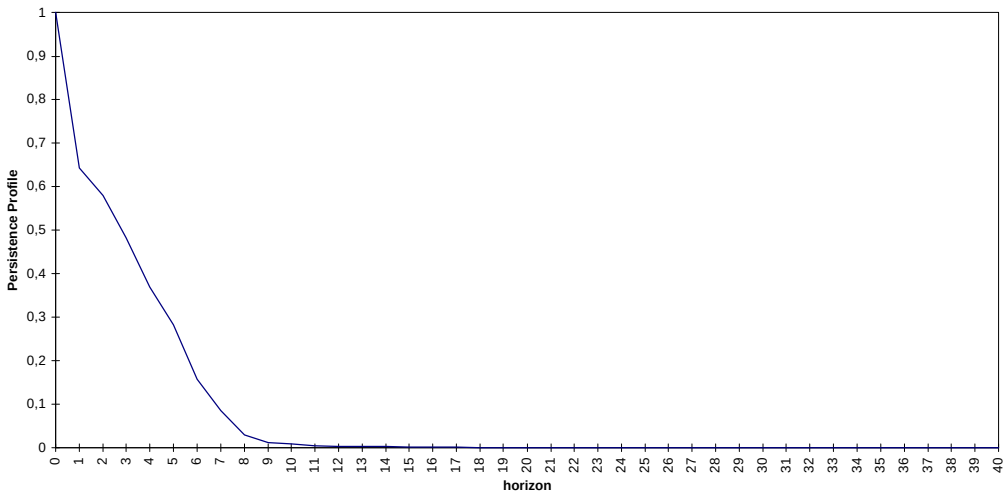


FIGURE 2b: Persistence Profiles

Italy/Switzerland



Italy/Japan

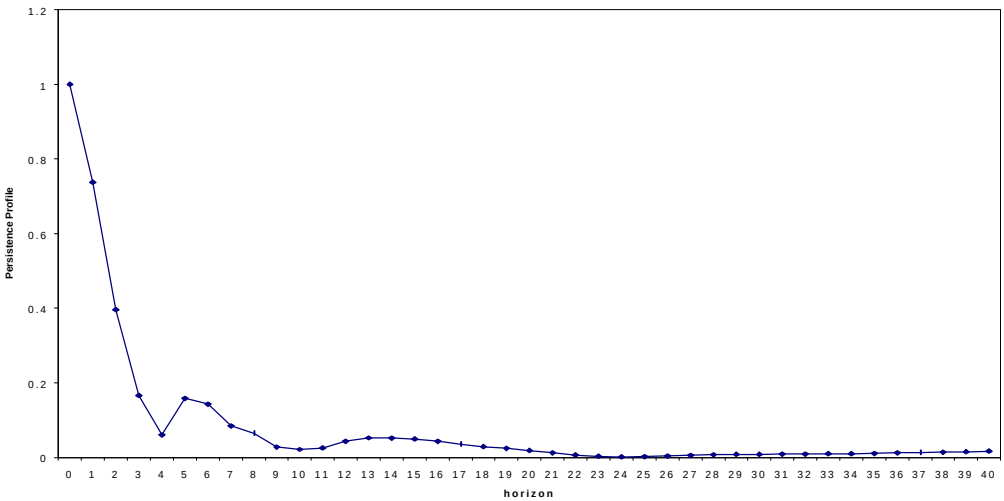
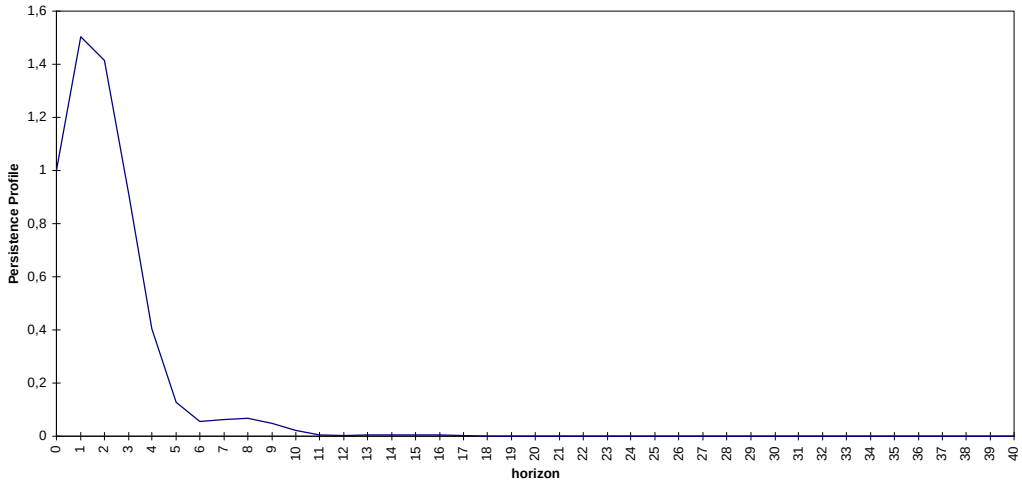


FIGURE 2c: Persistence Profiles

Switzerland/Japan



Switzerland/UK

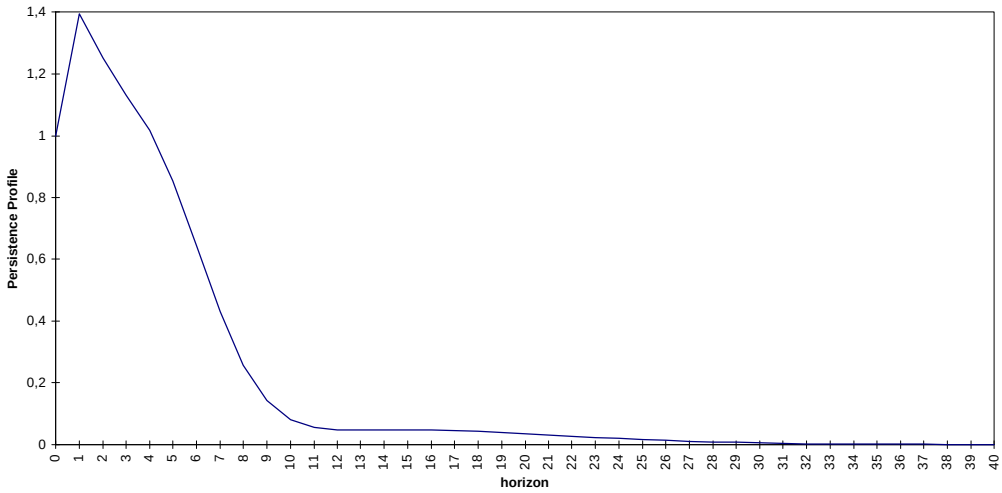
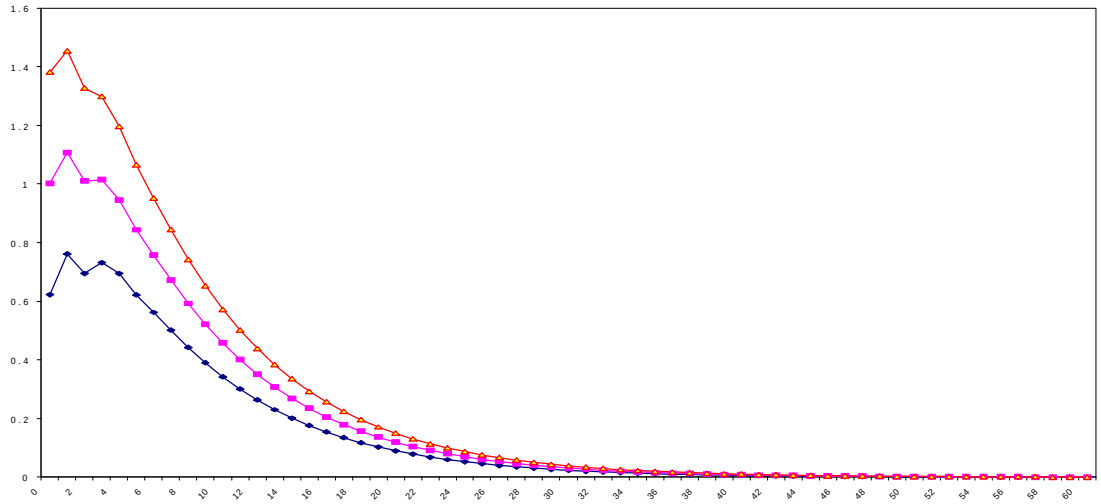


FIGURE 3a: Impulse Response Functions

Italy/Us



Italy/Germany

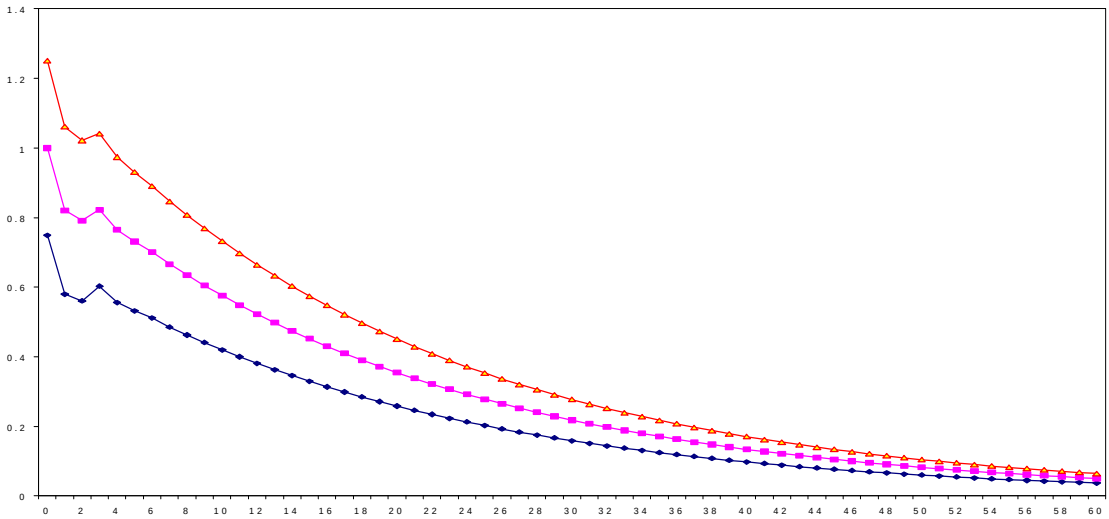
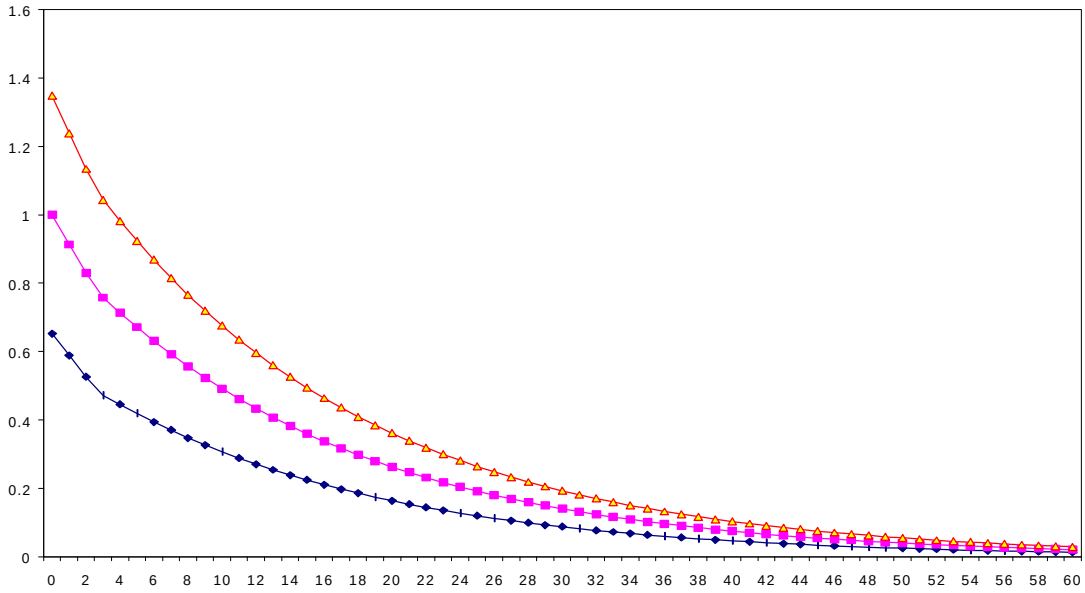


FIGURE 3b: Impulse Response Functions

Italy/Switzerland



Italy/Japan

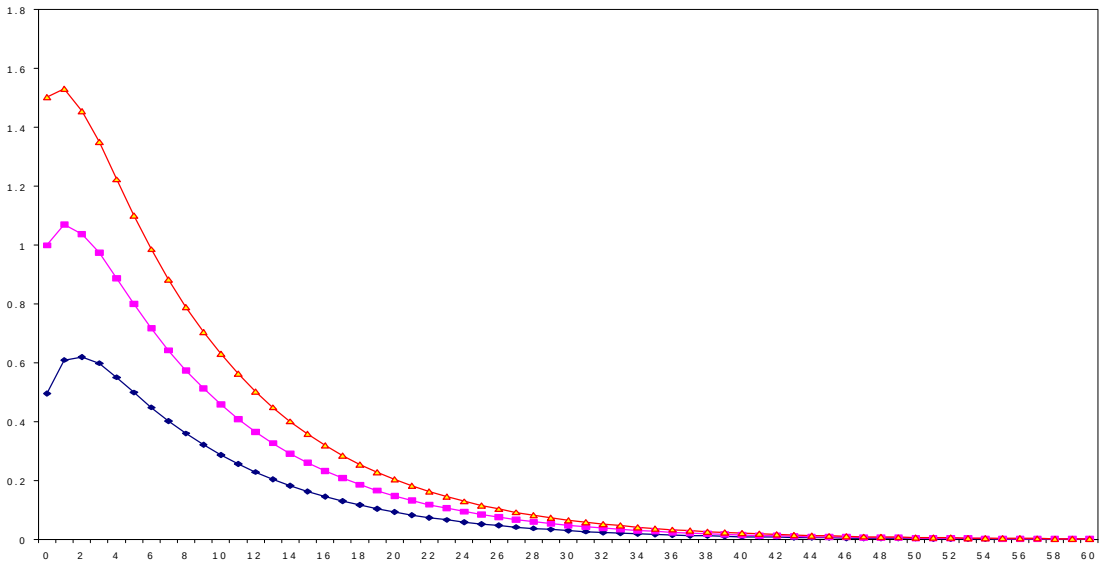
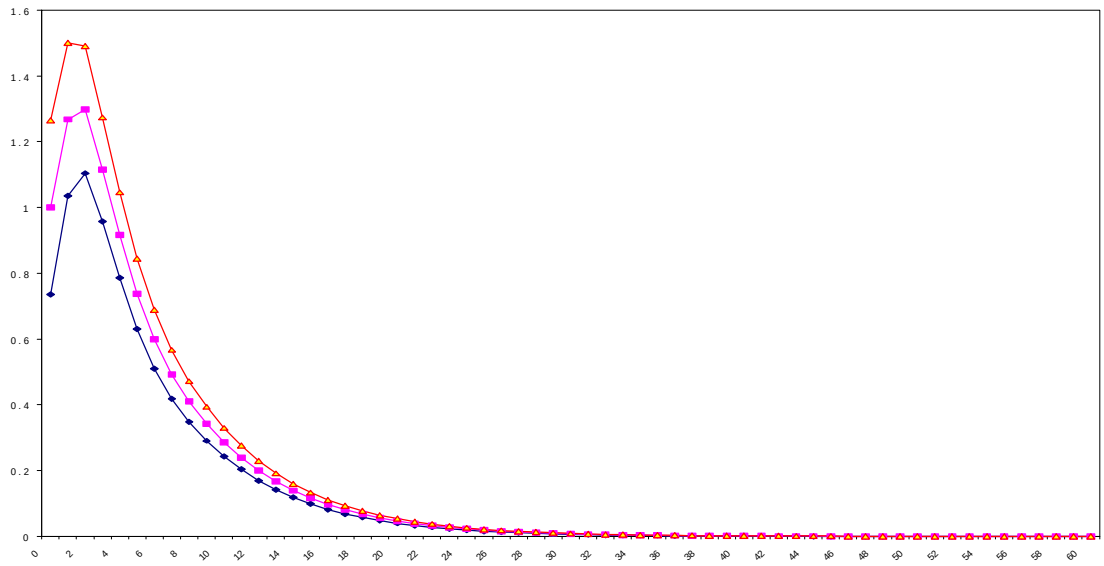


FIGURE 3c: Impulse Response Functions

Switzerland/Japan



Switzerland/UK

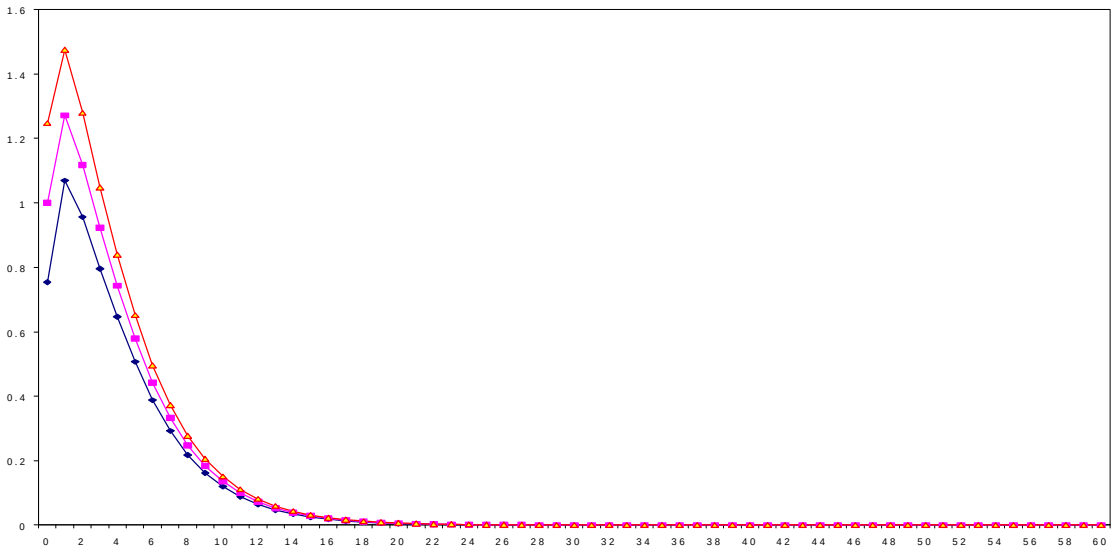
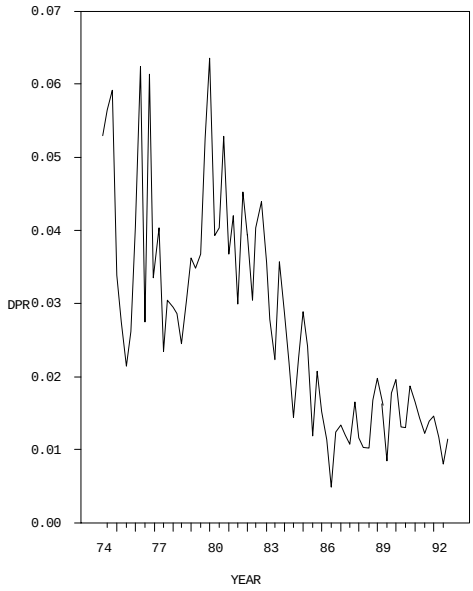


FIGURE 4a: First differences of price series

Italy



USA

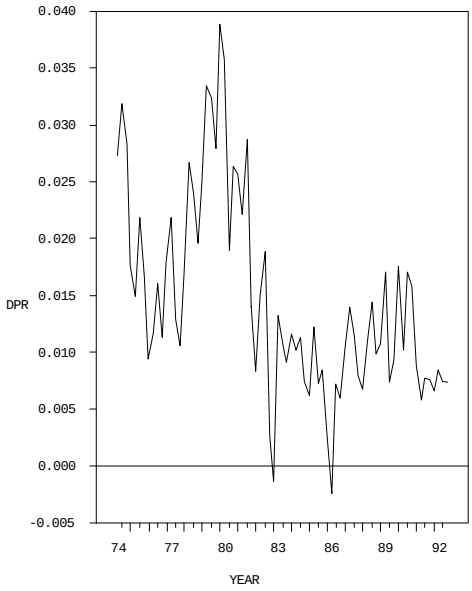
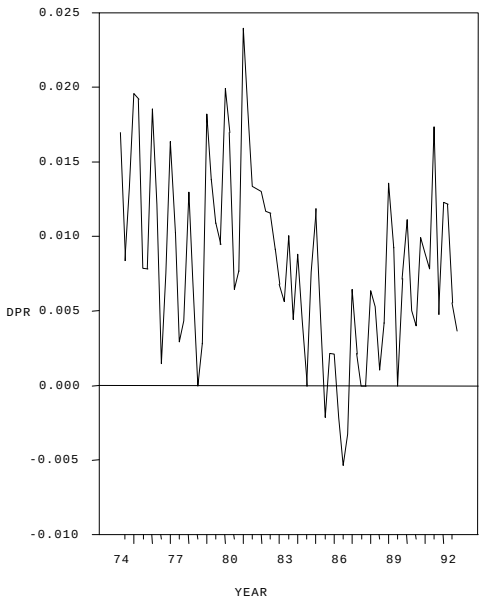


FIGURE 4b: First differences of price series

Germany



Switzerland

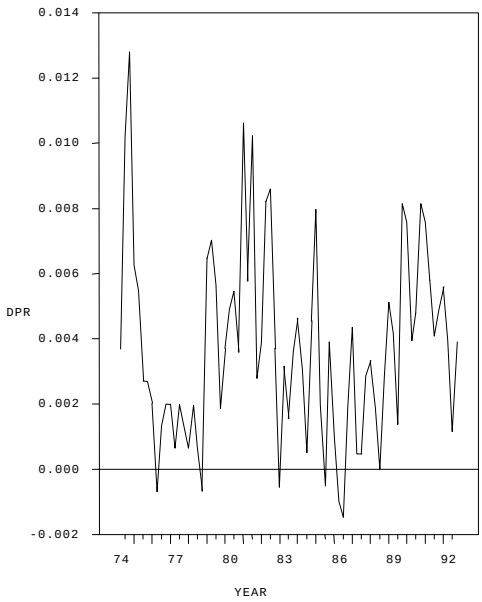
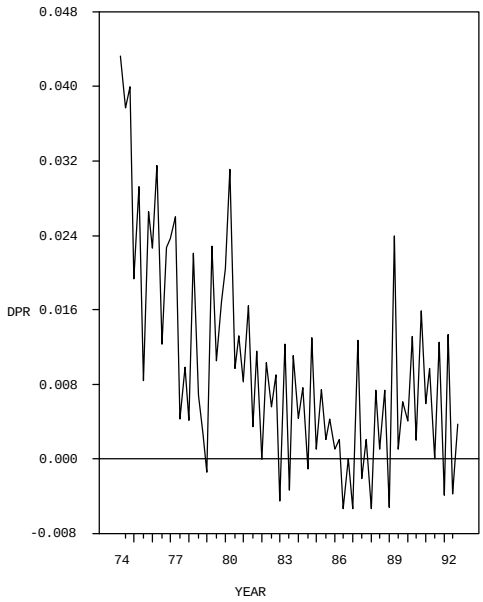
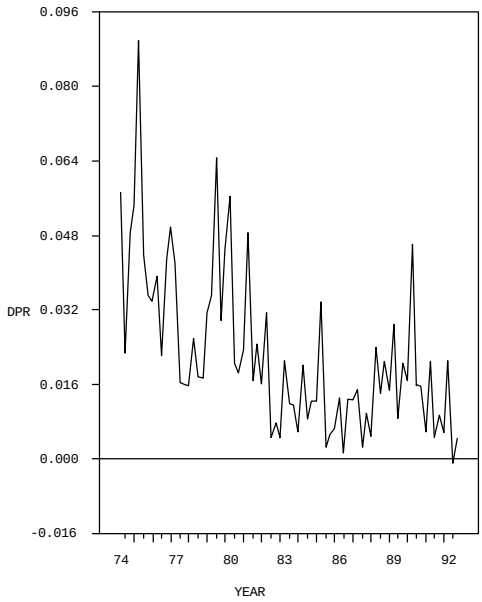


FIGURE 4c: First differences of price series

Japan



UK



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Notes

* We wish to thank Gianni Amisano and Carlo Favero for helpful comments. The usual disclaimer applies.

[1] For a detailed survey see Froot and Rogoff (1995)

[2] There is cointegration among the variables in [1], but symmetry and proportionality restrictions are generally rejected. The latter characteristic is attributed to measurement errors in price series (Taylor, 1988). Some evidence in favour of both cointegration and those homogeneity restrictions can be found in Edison, Gagnon and Melick, 1997.

[3] The panel approach is usually justified in terms of the gain in power deriving from having a larger variability in the sample. However, the actual gain is obtained when the speed of adjustment is the same for different real exchange rates (Liu and Maddala, 1996). In addition, O'Connell (1998) shows that the rejection of the unit root in the panel studies may be due to the failure to account for cross-sectional dependence and Taylor and Sarno (1998) introduce new tests to solve the problem of the frequent rejection of the null of non-stationarity when only one of the series is stationary.

[4] Over the considered period the former country is a member of EMS, the latter is not; therefore, we have the possibility to consider different categories of bilateral exchange rates: member/member, member/non-member, non-member/member, non-member/non-member.

[5] Full identification is defined as the simultaneous satisfaction of the algebraic condition for generic identification and the empirical non-rejection of the over-identifying restrictions on the cointegrating vectors (Johansen 1995).

[6] Appendix 1 contains a description of the data and their univariate properties.

[7] This approach is more efficient than the various single equation methods unless there is a unique cointegrating vector and the variables appearing on the right hand side are weakly exogenous. Results supporting this approach in terms of asymptotic and finite sample properties can be found in Phillips (1991) and Gonzalo (1994). For partially different results see Haug (1996).

[8] The nine bilateral real exchange rates resulting by combination of the considered currencies are plotted in Figures 1a to 1c.

[9] See Appendix 1.

[10] We utilise DisCo by Johansen and Nielsen (1993), version 1.4 of 1997. The simulation was performed with 10.000 iterations and the number of the discretizations of the Brownian motions, representing the asymptotic non standard theoretical distribution of the test, has been set at 600.

[11] The adopted test strategy is the one suggested by Pantula (1989) starting with the null hypothesis of rank zero and sequentially testing increasing rank orders. All the results are obtained with the package MALCOLM (Mosconi, 1998).

[12] Of course, the identification problem arises only if the cointegration rank is strictly greater than one.

[13] With respect to the second cointegrating vector we don't have strong theoretical a-priori. Hence, different alternative specifications of H_2 are plausible (three in this specific case). However, for our ten bilateral cases, the results of the LR tests are not sensitive to the choice of H_2.

[14] The critical values are obtained as in the previous section.

[15] Of course, each of the two alternative sets of restrictions is defined by *r* matrices of type **H**. For all the bilateral cases, the $\mathbf{H_i}$ matrices, $i=1,...r$, are reported in the Appendix 2.

[16] More precisely the Dominance Criterion acts as follow :
hypothesis H_1 is preferred to hypothesis H_2 if $L_2-L_1 <[\chi^2(n_2+1)- \chi^2(n_1+1)]/2$
H_2 is preferred to H_1 if $L_2-L_1 >[\chi^2(n_2 - n_1 + 1)- \chi^2(1)]/2$ the criterion is indecisive if $[\chi^2(n_2 - n_1 + 1)- \chi^2(1)]/2 > L_2-L_1 > [\chi^2(n_2+1)- \chi^2(n_1+1)]/2$ where n1 and n2 are respectively the degrees of freedom

related to H1 and H2 restrictions and L1 and L2 are the values of the constrained model likelihood functions.

[17] The FIML Johansen estimates are obtained by a multi-step concentration of the likelihood function of the system with respect to different blocks of parameters and the long run coefficients estimates are function of the short run ones.

[18] They are computed over a forty quarter horizon starting from our restricted VECM models (i.e. constrained to satisfy the PPP).

[19] We do not consider here the persistence profiles between cross terms $\boldsymbol{\beta}'_i \mathbf{X}_t$ and $\boldsymbol{\beta}'_i \mathbf{X}_t$ **,** so that we look only at the main diagonal elements of the persistence profile matrix.

[20] The persistence profiles have been obtained by a program written in Rats (version 4.2).

[21] The IRFs and their 90% confidence bounds have been obtained by a Monte-Carlo experiment with 10.000 iterations. They cover a 60 quarters simulation period.

[22] Exceptions to the common finding of a slow speed of adjustment can be also found in Obstfeld and Taylor (1997) and Hegwood and Papell (1998) that, respectively, allow for non linear adjustment and for structural breaks. By means of a (uniequational) regime-switching model for the exchange rate Klaassen (1999) shows that deviations from PPP became shorter lived after 1987; however this finding applies only to European countries (and not to Japan) and no measures of the speed of adjustment are provided in the paper.

[23] In the context of real exchange rate impulse response analysis.

[24] For Japan we use a 3-month interest rate which has been constructed with the rate on the gensaki market up to 1979Q2 and with the CD interest rate after that date. In this way we obtain a regulation-free interest rate variable. We have also utilised the interest rate on 60-day treasury bills, which shows the effects of high regulation during the first part of the period. The results are robust to the choice of the variable.

[25] All data have been obtained from Datastream.

[26] For example, for the Italian case, Hamilton (1994), using quarterly data for the period 73-89, concludes in favour of only one unit root. On the other hand, Paruolo (1993) cannot reject the presence of two unit roots for the same variable on the period 70-91.

[27] This evidence might imply a break also in the real exchange rate around mid-80's. Results in line with this conclusion can be found in Jorion and Sweeney (1996), Enders and Lee (1997), Wu (1997).

[28] The test is performed adopting the additive outlier specification (Perron, 1990) to model the shift of the constant with a dummy variable of the type: $D_t = \begin{cases} 0 & \text{if } t < 82:3 \\ 1 & \text{if } t \geq 82:3 \end{cases}$. In Figures 4a to 4c we report the plots of the first differences of the series.