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SHORT-RUN REAL EXCHANGE RATE DYNAMICS*

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The short-run dynamics of German mark and US dollar real exchange rates are investigated for a panel of 19 OECD economies in a vector error correction framework for the 1973–96 period. The novel persistence profiles approach of Pesaran and Shin ('Cointegration and Speed of Convergence to Equilibrium', *Journal of Econometrics*, Vol. 71, (1996), pp. 117–143) indicates that the effect of system-wide shocks declines rapidly initially but decays slowly thereafter. It yields an average of just one year for the half-life of such shocks but some seven years before they fully dissipate. These half-life estimates are just one-quarter of the consensus estimates. Our results are consistent with non-linear adjustment and with monetary factors being the main source of real exchange rate volatility.

1 Introduction

Interest has recently shifted from testing long-run purchasing power parity (PPP) or real exchange rate stationarity to trying to measure the speed of adjustment to the equilibrium relation. The rate of mean reversion after a shock is central to the PPP puzzle which Rogoff (1996, p. 664) poses as: 'How is it possible to reconcile the extremely high short term volatility of real exchange rates with the glacial rate (15 per cent per annum) at which deviations from PPP seem to die out?' Rogoff gives the consensus estimates of the half-life of shocks as 3–5 years for the industrialized countries.[1] Much of the existing literature on short-run PPP dynamics follows the structural vector autoregressive (VAR) approach of Clarida and Gali (1994). In this tradition, orthogonalized impulse response functions are employed to

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[1]Papell (1997) notes that such estimates need to be treated with caution since they are typically estimated with some bias from a simple AR(1) process. Note also that some panel approaches have produced smaller half-life estimates of 1–3 years. See Coakley and Fuertes (1997), MacDonald (1996) and Papell (1997).

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investigate the impact of shocks to individual variables, be they real or monetary, supply side or demand side shocks. One weakness of this literature is that the impulse response functions are not uniquely identified.

This paper starts from the presumption that long-run PPP holds for the post-Bretton Woods era[2] and has a different focus. It is concerned with the speed of adjustment of real exchange rates to system-wide rather than variable specific shocks. The original contribution lies in its new estimates of the speed of adjustment to such shocks. We find that, while these shocks have persistent effects, the short-run dynamics are rather different from those depicted in much of the existing literature. The impact of system-wide shocks is investigated by means of the persistence profiles approach developed by Pesaran and Shin (1996). This focuses on analysis of the effect of system-wide shocks on equilibrium relations within a cointegration framework.[3]

The persistence profile estimates for our OECD economies show an average total life of shocks of 6 and 8 years, respectively, for both DM and US$ series. Assuming a linear adjustment process, these estimates do not seem out of line with the consensus half-life estimates in the literature and give the impression that real exchange rates mean revert very slowly due to their long tails. More critically, the persistence profiles indicate rapid mean reversion in the early months after the initial impact but slow adjustment thereafter. The half-life of shocks is only just over 1 year on average for both DM and US$ series. This combination of rapid and slow mean reversion may well reflect a non-linear adjustment mechanism.[4] These new half-life estimates contrast sharply with the consensus estimates and are easier to reconcile with standard monetary exchange rate models.

The paper is organized as follows. Section 2 outlines the persistence profiles approach in a cointegrating VAR framework. Section 3 presents and analyses our empirical results, while a final section concludes.

2 Persistence Profiles and Short-run Dynamics[5]

As a framework for analysing the short-run dynamics of equilibrium relations, consider the following m-dimensional, unrestricted $\mathrm{VAR}(L)$ model:

[2]For recent panel tests supporting PPP see, for example, Boyd and Smith (1999), MacDonald (1996), Papell (1997), Taylor and Sarno (1998) and Coakley and Fuertes (1997, 2000a). Note, however, that Engel (1996) and O'Connell (1998) find contrary evidence.

[3]See Pesaran and Smith (1997) for a readable introduction to impulse response functions and persistence profiles analysis and Pesaran and Smith (1998) for a more formal overview.

[4]This finding is in line with some new estimates of half-lives provided by Obstfeld and Taylor (1997) and Coakley and Fuertes (1998a) from threshold autoregressive (TAR) models of real exchange rates and by Taylor *et al.* (2000) from smooth transition autoregressive (STAR) models.

[5]This section draws on Lee and Pesaran (1993) and Pesaran and Shin (1996).

$$x_t = \sum_{i=1}^{L} \Phi_i x_{t-i} + c_0 + c_1 t + u_t \qquad t = 1, 2, \ldots, T \tag{1}$$

where x_t is an $m \times 1$ vector of jointly determined dependent variables, c_0 and c_1 are $m \times 1$ vectors of unknown coefficients, Φ_i $(i = 1, 2, \ldots, L)$ are unknown $m \times m$ matrices of autoregressive parameters, and u_t is an unobserved vector of shocks or innovations satisfying the assumptions $E(u_t) = 0$, $E(u_t u_t') = \Sigma\ \forall t$, $E(u_t u_{t-j}') = 0\ \forall j \neq 0$, with Σ an $m \times m$ constant positive-definite matrix. The system in (1) can be reparameterized in vector error correction (VEC) form as follows:

$$\begin{aligned} \Delta x_t &= \sum_{i=1}^{L-1} \Gamma_i \Delta x_{t-i} - \Pi x_{t-1} + c_0 + c_1 t + u_t \qquad t = 1, 2, \ldots, T \\ \Pi &= I_m - \sum_{i=1}^{L} \Phi_i \qquad \Gamma_i = -\sum_{j=i+1}^{L} \Phi_j \qquad i = 1, 2, \ldots, L-1 \end{aligned} \tag{2}$$

Assuming x_t to be first-difference stationary, if $\operatorname{rank}(\Pi) = r < m$, the $m \times m$ matrix Π can be expressed as $\Pi = \alpha\beta'$ where α and β are $m \times r$ matrices of full column rank. In this context β is the cointegrating matrix and $z_t = \beta' x_t$ is an $r \times 1$ long-run equilibrium vector representing r cointegrating relations.

To analyse the response of the cointegrating or equilibrium relations z_t to particular shocks, one can use a simple adaptation of the orthogonalized impulse response approach based on the Cholesky decomposition of the covariance matrix of innovations ($\Sigma = TT'$). This has the drawback of being sensitive to the ordering of the variables in the cointegrating VAR and to the choice of the matrix T. Alternatively, one can use the variance-based persistence profiles of Lee and Pesaran (1993) and Pesaran and Shin (1996) which measure the effect over time of system-wide shocks on the equilibrium relations. The attractive feature of these time profiles is that they are uniquely identified since their estimation does not require prior orthogonalization of the vector of shocks.

Pesaran and Shin (1996) propose the following unscaled measure of persistence to analyse the effect of system-wide shocks on cointegrating or equilibrium relations:

$$H_z(n) \equiv V(z_{t+n} | I_{t-1}) - V(z_{t+n-1} | I_{t-1}) \qquad n = 0, 1, 2, \ldots \tag{3}$$

where $V(z_{t+n} | I_{t-1})$ is the variance of z_{t+n} conditional on the information set I_{t-1}. In the context of (2) this is given by

$$H_z(n) \equiv \beta' B_n \Sigma B_n' \beta \qquad n = 0, 1, 2, \ldots \tag{3$'$}$$

where $B_j = \sum_{i=0}^{j} A_i$ $(j = 0, 1, 2, \ldots)$ and A_i are the $m \times m$ coefficient matrices of the MA(∞) representation of Δx_t. The matrices B_j satisfy the following recursive relation with the matrices Φ_j of the VAR(L) model:

$B_j = \Phi_1 B_{j-1} + \Phi_2 B_{j-2} + \ldots + \Phi_L B_{j-L}$ for $j = 1, 2, \ldots, \infty$, with $B_0 = I_m$ and $B_j = 0$ for $j < 0$. The diagonal elements of the matrix $H_z(n)$ are the unscaled persistence profiles or system-wide impulse responses of the cointegrating relations $z_t = \beta' x_t$. The (scaled) persistence profile of the jth cointegrating relation $z_j = \beta_j' x_t$ is then given by

$$h_{z_j(n)} = \frac{H_{z_j(n)}}{H_{z_j(0)}} = \frac{\beta_j' B_n \Sigma B_n' \beta_j}{\beta_j' \Sigma \beta_j} \qquad n = 0, 1, 2, \ldots \tag{4}$$

which by construction has unit value at the time of impact when $n = 0$. It should tend to zero as $n \to \infty$ if β_j' is indeed a cointegrating vector. Pesaran and Shin (1996) show that the maximum likelihood (ML) estimates of these persistence profiles are $\sqrt{T}$ consistent with a limiting normal distribution.

The persistence profiles can be given different interpretations. First, the unscaled persistence profiles can be regarded as the variance of the revision in the n-step-ahead forecast of z_t, i.e. $H_z(n) = V(v_{t+n})$, where $v_{t+n} = E(z_{t+n}|I_t) - E(z_{t+n}|I_{t-1})$. Alternatively, when the system has just one cointegrating vector, β', there is a formal correspondence between the persistence profile and the impulse response function of the cointegrating relation. More specifically, Pesaran and Shin (1996) show that $h_z(n) = \delta_n^2$, where δ_n denotes the impulse response function of the cointegrating relation to a unit composite shock $\varepsilon_t = \beta' u_t$.[6]

The persistence profiles approach can readily be applied to analyse the short-run dynamics of the PPP relation in the wake of a shock. Define the real exchange rate q_t as the nominal exchange rate s_t minus a price differential $p_t - p_t^*$:

$$q_t = s_t - (p_t - p_t^*) \tag{5}$$

where all variables are in natural logarithms and s_t is defined as units of domestic currency per unit of foreign currency. The stochastic process q_t can be viewed as deviations from long-run PPP equilibrium. We formulate a VEC model as in (2) with $x_t = (s_t, p_t, p_t^*)'$ to investigate real exchange rate dynamics.

3 Data and Empirical Results

Our monthly data were taken from Datastream and span the 23 year period 1973m7–1996m6. We employ end-of-month spot bilateral exchange rates with both the German mark and the US dollar as *numéraires* and, as price series, the monthly wholesale price index with 1990 as base year.

[6]Yet another way of motivating persistence profiles is by means of the generalized impulse response analysis of Koop *et al.* (1996).

Our panel consists of 19 OECD countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the UK and the USA. This facilitates the definition of 18 DM-based real exchange rates of which all but four are European. The predominance of European countries in this context justifies our presumption of mean reversion. This is partly due to geographical proximity and trade links,[7] and partly to the effects of the EU Single Market and European Monetary Union (EMU) programmes, including the hybrid nature of nominal exchange rate regimes within Europe post-1973. For comparative purposes, we also analyse the dynamic behaviour of US$ real exchange rates.

3.1 Cointegration Analysis

The augmented Dickey–Fuller, Phillips–Perron and Kwiatkowski *et al.* (1992) univariate unit root tests on the series in both levels and first differences indicate that nominal exchange rates and prices can be considered as I(1) variables.[8] Next a VEC model is formulated with two alternative specifications for the deterministic part: one with intercepts only and one with both intercepts and trends. The intercepts are unrestricted in both formulations to allow for linear trends in the data given the upward-trending nature of the price series. In the second specification the trends are restricted so that the cointegrating space includes a linear time trend.[9] The lag order in each model is selected using the Akaike and Schwarz model selection criteria and a sequence of likelihood ratio (LR) statistics.

The Johansen (1988, 1991) procedure applied to both VEC specifications produced evidence favouring cointegration between the nominal exchange rate, domestic prices and foreign prices with a cointegrating rank r of 1 or 2 in all countries. Table 1 reports the DM series results.[10] Given the finite sample problems associated with the Johansen trace test and since economic theory predicts one equilibrium (PPP) relation in the above system, the rank restriction $r = 1$ is imposed in all cases.[11] Following Pesaran and Smith (1998), the choice between the two deterministic

[7]On these issues see Engel *et al.* (1997).

[8]Detailed results are available from the authors on request. Some attention has focused recently on price series as realizations of I(2) processes. See, for example, Juselius (1997). Alternatively Granger *et al.* (1997) propose a class of non-linear growth processes which might characterize some price series.

[9]Gonzalo and Lee (1998) stress the importance of correctly specifying the deterministic terms.

[10]The analysis was carried out using Microfit 4.0. See Pesaran and Pesaran (1997).

[11]The result $r = 2$ is consistent with the Greenslade *et al.* (1998) finding that treating all variables as endogenous has the effect of generally overestimating the true number of cointegrating vectors for an adequate VAR lag length.

Table 1
Johansen Cointegration Results for DM Series

		Cointegrating rank		Deterministic specification			Over-identifying restrictions
		Trace test $r=0/r\geq 1$ $r\leq 1/r\geq 2$ $r\leq 2/r=3$		SBC for VEC model[c]		Co-trending restrictions[d]	Normalized $\beta'=(1\ \beta_1\ \beta_2)$ H_0: $\beta_1=-1, \beta_2=1$
	L[a]	NT[b]	T	NT	T	LR test (p value)	LR test (p value)[e]
AT	4	55.224* 23.573* 5.074	61.539* 26.779* 5.707	**2626.9**	2625.6	3.0 (0.083)	11.618 (0.003) (NT)
AU	3	55.953* 22.675* 6.642	75.547* 34.253* 10.543	3003.5	**3004.7**	8.0 (0.005)	9.051 (0.011) (T)
BG	2	39.940* 18.407* 1.436	60.682* 28.284* 8.815	**2905.6**	2903.3	1.0 (0.317)	2.871 (0.238) (NT)
CN	2	44.060* 11.001 5.026	71.963* 24.738** 5.302	**2746.7**	2745.4	3.2 (0.074)	33.805 (0.000) (NT)
DK	3	42.188* 13.108 2.087	50.031* 20.948 7.043	**3001.0**	2998.2	0.0 (0.997)	17.462 (0.000) (NT)
FN	2	73.074* 21.512* 5.734	81.668* 27.819* 11.923	**2729.7**	2728.0	2.2 (0.138)	9.493 (0.010) (NT)
FR	2	53.423* 20.788* 4.907	63.327* 29.694* 12.884*	**2717.0**	2714.7	1.0 (0.317)	3.365 (0.186) (NT)
GR	3	53.859* 22.792* 8.030	65.914* 25.707 9.768	2617.7	**2619.5**	9.0 (0.003)	25.873 (0.000) (T)
IR	4	40.495* 12.819 3.306	57.768* 30.092* 8.301	**2785.9**	2783.1	0.0 (0.997)	22.241 (0.000) (NT)
IT	4	35.219* 13.299 5.132	39.533** 16.464 6.964	**2735.6**	2733.3	1.0 (0.317)	4.731 (0.094) (NT)
JP	4	37.550* 17.961* 5.328	49.557* 20.720 7.985	2726.5	**2728.3**	9.2 (0.002)	12.139 (0.002) (T)

TABLE 1 (*Continued*)
JOHANSEN COINTEGRATION RESULTS FOR DM SERIES

		Cointegrating rank		*Deterministic specification*			*Over-identifying restrictions*
		Trace test $r = 0/r \geq 1$ $r \leq 1/r \geq 2$ $r \leq 2/r = 3$		*SBC for VEC model*[c]		*Co-trending restrictions*[d]	*Normalized* $\beta' = (1\ \beta_1\ \beta_2)$ H_0: $\beta_1 = -1, \beta_2 = 1$
	L[a]	NT[b]	T	NT	T	*LR test (p value)*	*LR test (p value)*[e]
NH	2	86.177*	95.085*				
		27.712*	28.504*	3177.2	**3178.4**	8.0 (0.005)	41.302 (0.000)
		7.464	7.533				(T)
NW	2	50.347*	63.667*				
		13.896	20.439	2885.2	**2885.7**	6.8 (0.009)	2.394 (0.302)
		5.754	5.950				(T)
SD	2	54.275*	72.238*				
		26.780*	28.853*	2854.4	**2859.5**	16.0 (0.000)	3.732 (0.155)
		5.333	7.067				(T)
SP	4	42.848*	59.438*				
		20.716*	26.515*	2725.7	**2728.3**	10.8 (0.001)	3.582 (0.167)
		7.821	10.797				(T)
SW	2	58.505*	62.293*				
		29.712*	33.424*	**2979.5**	2976.7	0.0 (0.997)	5.420 (0.067)
		5.229	7.102				(NT)
UK	3	78.727*	100.749*				
		21.142*	42.871*	**2795.6**	2792.9	0.2 (0.655)	5.890 (0.053)
		4.971	5.449				(NT)
US	2	60.580*	80.878*				
		18.483*	32.966*	2695.8	**2695.9**	5.8 (0.016)	1.174 (0.556)
		3.989	4.936				(T)

Notes:
AT, Australia; AU, Austria; BG, Belgium; CN, Canada; DK, Denmark; FN, Finland; FR, France; GR, Greece; IR, Ireland; IT, Italy; JP, Japan; NH, Netherlands; NW, Norway; SD, Sweden; SP, Spain; SW, Switzerland; UK, United Kingdom; US, United States.

[a] Maximum lag order of the underlying VAR is selected using the Akaike information criterion, the SBC and a sequence of LR statistics.
[b] NT indicates unrestricted intercepts and no trends and T unrestricted intercepts and restricted trends.
[c] SBC for a VEC with $r = 1$ and L as indicated above; maximum value given in bold.
[d] The null is that the cointegrating relations contain no deterministic linear trends; p values from $\chi^2(1)$.
[e] p values from $\chi^2(2)$ distribution.
* Significant at the 5 per cent level; **significant at the 10 per cent level.

specifications considered is based on both the Schwarz Bayesian criterion (SBC) and LR tests of the co-trending hypothesis or absence of a deterministic linear trend in the cointegrating relation. Table 1 shows that both methods give the same qualitative results. The VEC model without trend yields the highest SBC value in 10 countries, which is corroborated by the LR test failing to reject the co-trending null. The VEC model with intercepts only is adopted for these countries. The remaining eight countries, Austria, Greece, Japan, the Netherlands, Norway, Spain, Sweden and the USA, favour the inclusion of a trend.[12]

To impose structure on the Johansen (normalized) cointegrating vector $(1\ \beta_1\ \beta_2)'$, we implement an LR test of the PPP over-identifying restrictions $\beta_2 = -\beta_1 = 1$. The results in Table 1 show that the null cannot be rejected in 10 out of 18 countries at the 5 per cent level and in a further two countries at the 1 per cent level.[13] The remaining cases might be explained by possible price measurement bias and the relatively short span of the data which is only some 23 years.[14] Bearing these reservations in mind, we proceed with an analysis based on the theory-restricted cointegrating vector.

The error correction parameters $\alpha = (\alpha_1, \alpha_2, \alpha_3)'$ are explored while recognizing that, in a system approach, their interpretation in terms of adjustment speeds is not straightforward.[15] Separate tests on each α_j $(j = 1, 2, 3)$ in the restricted VEC system indicate statistically significant feedback coefficients in the nominal exchange rate, domestic price and foreign price equations in 10, nine and six countries, respectively, for the DM series. This suggests that all three variables play a role to varying extents in adjustment to PPP equilibrium and this may reflect the impact of convergence policies in the run up to EMU. The results for the US$ series show a different pattern with significant feedbacks in zero, six and 15 cases, respectively, for the same three variables. The lack of feedback of PPP disequilibria on the nominal exchange rate may relate to the sharp swings in US$ exchange rates during the 1980s alluded to by Lothian (1998) or to the dollar regime shifts as analysed by Evans and Lewis (1995). Equally, the seemingly strong responsiveness of US prices to PPP disequilibria may reflect the pervasive impact of US monetary policy.

[12]The US$ series supported the trend specification in seven cases.

[13]However, the weak version of PPP, given by the restriction $(1\ -\alpha\ \alpha)'$, is accepted in virtually all countries.

[14]The rejections might also relate to the fact that all the variables (s_t, p_t, p_t^*) in the VEC system are jointly endogenous.

[15]Rossana (1998) points to two problems in this respect. One relates to statistically arbitrary normalizations for cointegrating matrices having rank greater than 1, which does not apply in our case since a unique cointegrating vector is assumed. The other is that, in the case of VARs with high order lags, the adjustment matrix will contain the effects of autoregressive parameters of unobservable shocks as well as the adjustment speed.

We observe in the VEC system for each country that at least one of the three feedback coefficients is statistically significant, which implies Granger causality between some of the variables to sustain long-run PPP. We verify this by using an LR statistic to test the null that the three coefficients are jointly insignificant. For the DM series this null is rejected in 15 cases at even the 1 per cent level and in the remaining three (Australia, Finland and Ireland) at the 10 per cent level.[16] These results support our restricted VEC model relative to a VAR in first differences.

3.2 Persistence Profiles

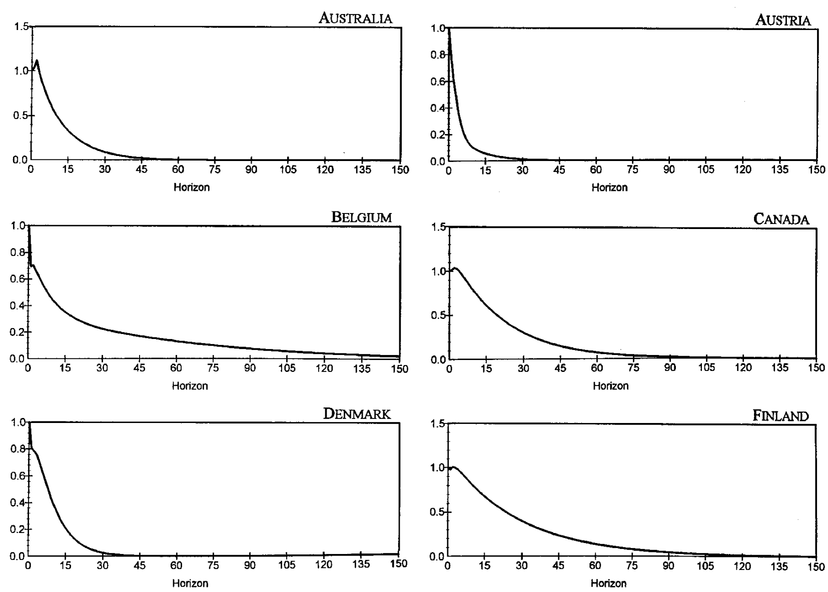
The persistence profiles of the PPP cointegrating relation are computed from the ML estimates of the restricted VEC model for each country. Figure 1 shows these for the DM series. The persistence profiles point estimates show that the effects of system-wide shocks decline approximately monotonically, absent a few overshoots and kinks, and eventually die out in line with the stationarity of real exchange rates. For instance, mean reversion resumes following a small initial overshoot in Australia, Canada, Finland, Sweden, Switzerland and the UK.

The mean total life of these shocks is under 6 years for the DM series and just in excess of 8 years for the US$ series as can be seen in Table 2. At first glance this suggests persistence in real exchange rates, especially for the US$ series. However, this is to overlook the rapidity of the initial adjustment phase in which most of the effect of the shock is dissipated. From the persistence profiles we can calculate the empirical half-life of system-wide shocks as the time elapsed for half of their initial impact to disappear. These estimates have two merits. First they are obtained via ML estimation in a system-based cointegrating framework which captures the interaction of shocks to the separate components of real exchange rates, whereas the standard approach is based on an AR(1) regression. Moreover, the VEC lag order chosen by information criteria and a sequence of LR tests allows richer dynamics. Second, our definition furbishes a conservative measure of half-lives in the case of overshooting since the latter delays the onset of the mean reversion process. The half-life estimates are reported in Table 2.[17]

For the DM series, an average of half of the initial impact of the system-wide shock has disappeared after 10 months which is only 15 per cent of the average total life. What is remarkable for the DM series is that

[16]The results for the US$ are qualitatively similar, with marginal rejections for Belgium, France, Italy and the Netherlands.

[17]These half-lives need to be treated with caution. Although they share the same rationale as the conventional AR(1) based estimates, they are derived from different underlying models.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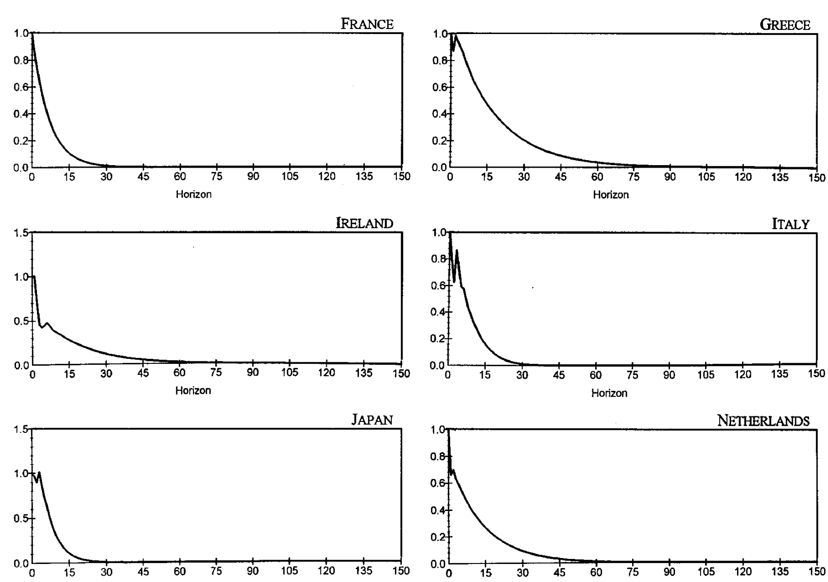
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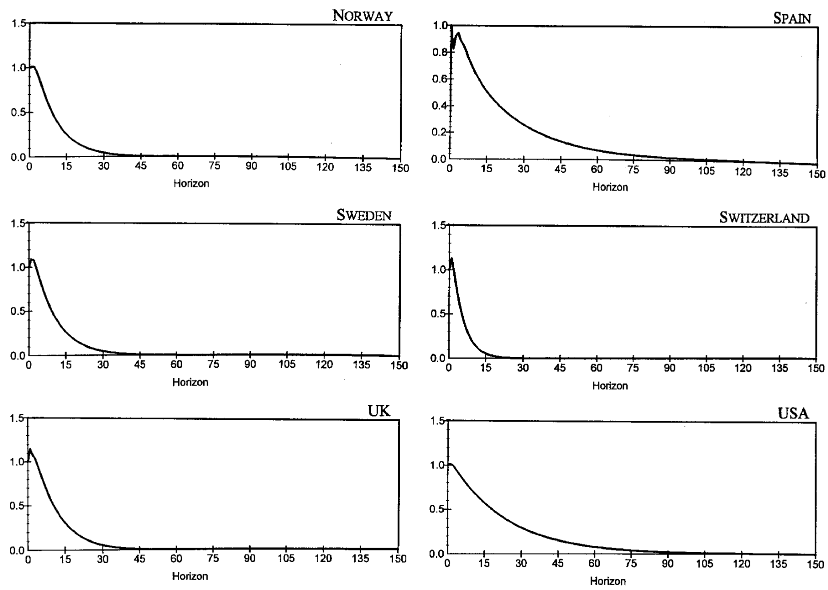

FIG. 1 Persistence Profiles of Real Exchange Rates (DM Series)

TABLE 2
PERSISTENCE OF SYSTEM-WIDE SHOCKS TO REAL EXCHANGE RATES

	DM series		*US$ series*	
	Half-life	*Total life*	*Half-life*	*Total life*
AT	10	49	7	48
AU	3	34	11	127
BG	8	165	23	138
CN	19	103	9	124
DK	8	39	19	84
FN	23	134	13	76
FR	5	33	15	86
GE	—	—	18	103
GR	14	84	10	79
IR	7	72	14	86
IT	7	32	18	165
JP	8	30	15	104
NH	6	61	19	165
NW	10	46	13	68
SD	9	44	18	76
SP	15	106	17	103
SW	5	24	14	75
UK	11	45	10	52
US	18	108	—	—
Averages	10.3	67.2	14.6	97.7

Note: Figures are in months.

only five countries yield half-life estimates in excess of 1 year: Canada, Finland, Greece, Spain and the USA. Most of the long-standing EU member countries have below-average half-lives. The shortest half-lives of all are recorded by Austria, France, the Netherlands and Switzerland, which are geographically adjacent to Germany. The estimates for the US$-denominated real exchange rates yield a longer average half-life of some 14.6 months. In all but five cases the half-lives of shocks are longer for the dollar series. The cases of Canada and Australia could be explained by reference to trade links but it is more difficult to find commonalities among the other three, Finland, Greece and the UK.

More generally, the shorter total and half-lives of system-wide shocks for the DM series compared with the US$ series is to be expected for several reasons. One is the preponderance in our panel of European countries which are geographically adjacent and have strong trade links. Engel *et al.* (1997) have stressed the importance of geographical proximity in a different context. Reinforcing this are the convergence policies pursued within the EU both as part of the 1992 Single Market programme and in the run up to EMU. Finally the apparent lack of adjustment of nominal dollar exchange rates (especially during the 1980s) may be yet another factor contributing to longer half-lives of shocks for US$ real exchange rates.

The average half-life estimates of about 1 year obtained from the persistence profiles analysis contrast very sharply with the consensus estimates of 3–5 years. However, they are in line with those obtained from the persistence profiles of a modified definition of relative PPP provided by Garratt *et al.* (1998) in the context of a UK macroeconometric model. While Garratt *et al.* focus on the overall length of their persistence profiles, their results indicate a half-life (as we define it) of just under 1 year. Our empirical half-lives are also consistent with the recent evidence provided by some non-linear adjustment models.[18]

An attractive property of these non-linear models, mainly TAR or STAR parameterizations, is that they can parsimoniously capture a stationary process whose speed of adjustment is not constant but proportional to the size of the deviations from equilibrium. A rationale for such models is that market frictions or transaction costs can make arbitrage unprofitable in a band around equilibrium where, as a consequence, the real exchange rate may exhibit persistent or random walk behaviour. This sluggish band would be consistent with the long tails observed in the persistence profiles. Additionally, large positive or negative deviations from equilibrium are likely to trigger arbitrage activities, thereby leading to a band where deviations from equilibrium die out relatively quickly. The latter may explain the rapid adjustment section in

[18]See, for example, Michael *et al.* (1997), Obstfeld and Taylor (1997), Taylor *et al.* (2000) and Coakley and Fuertes (1998a).

the persistence profiles. In accordance with these non-linear models and despite an inner band of potentially protracted deviations from parity, real exchange rates exhibit overall (or global) stationarity.[19] In this context, the half-lives of some 12–15 months from TAR models of real exchange rates provided by Obstfeld and Taylor (1997) and Coakley and Fuertes (1998a) are in line with our persistence profile findings.[20]

Finally, if our estimates are accepted, they may shed some light on the PPP puzzle by offering a reconciliation of short-run volatility with apparent persistence. Since we establish that the major impact of system-wide shocks is rapidly eroded, we can infer that this provides indirect evidence of monetary factors as the main source of real exchange rate volatility in the post-1973 era.[21] In this respect they are consistent with monetary models of the nominal exchange rate.

4 Conclusions

In this paper we analyse the short-run dynamics of DM and US$ real exchange rates for a panel of 19 OECD economies post-1973 in the context of a cointegrating VAR model. We use the novel persistence profiles approach of Lee and Pesaran (1993) and Pesaran and Shin (1996) to analyse the effect of system-wide shocks on the long-run PPP equilibrium relation. The results indicate that the effect of such shocks is persistent but transitory, with average total lives of some 6 and 8 years, respectively, for the DM and US$ series.

However, real exchange rates mean revert rapidly in the months following system-wide shocks and adjust slowly thereafter. The persistence profiles approach facilitates calculation of the half-life of shocks as the time elapsed before half of their initial impact disappears. On this basis, our results indicate a half-life of only 1 year on average. The DM series display more rapid mean reversion than the US$ series with average half-lives of 10.3 against 14.6 months, reflecting the weighting of our sample towards European countries. Although these figures are in line with some recent estimates from non-linear approaches, they are just one-quarter of the consensus 3–5 year estimates in the literature. Our results may be interpreted as providing indirect evidence of non-linearities in real exchange rate adjustment and as pointing to monetary factors as the main source of real exchange rate volatility.

[19]Further indirect evidence of non-linear real exchange rate adjustment is provided by the results of the Bierens (1997) non-parametric cointegration test in Coakley and Fuertes (1998b).

[20]See Coakley and Fuertes (2000b) for new bootstrap tests of asymmetries/non-linearities and Coakley *et al.* (2000) for an application to unemployment.

[21]Rogers (1998) provides new evidence on the importance of monetary shocks using a structural VAR model.

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