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THE OVERREACTING BEHAVIOR OF REAL EXCHANGE RATE DYNAMICS

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This study reviews and discusses empirical evidence corroborating the existence of overreaction in the short-term responses of real exchange rates. The amplification of shock responses, albeit occurring over a short time period only, can delay and substantially prolong the time it takes for the real exchange rate to converge to parity. Interestingly, the findings of short-term amplified responses of the real exchange rate—with its subsequent reversal and gradual reversion toward the long-run equilibrium—appear compatible with the chartist-fundamentalist model of the foreign exchange market microstructure.

1 Introduction

The purchasing power parity (PPP) theory, which suggests that two countries' price levels are equal at equilibrium when expressed in a common currency unit, has served as a major building block for many models of exchange rate determination. Under the PPP theory, nominal disturbances have no permanent effects on the real exchange rate, as implied by long-run monetary neutrality. Although short-run departures from parity are commonly recognized, many economists continue to hold the view that PPP, as a long-run proposition, will prevail. The faith in PPP has been weakened by the recent floating-rate experience, nonetheless. PPP deviations, gauged by real exchange rates, are often observed to be large, volatile and highly persistent. Such dynamics appear for the most part unexplained by economic fundamentals.

It is difficult to reconcile the immense short-term volatility of the real exchange rate with its very slow rate of convergence to parity. Slowly evolving changes in economic fundamentals—such as changes in tastes and technology—may contribute to the slow convergence, but they are not volatile enough over the short term to account for the vast exchange rate volatility. Sticky-price models, à la Dornbusch's (1976) overshooting analysis, are often used to show how monetary shocks can bring about large, volatile deviations from PPP. The strong short-term correlation generally found between exchange rates and real exchange rates can be viewed as indirect evidence for price stickiness (Mussa, 1986). With sticky prices, an unexpected change in money supply will alter real cash balances and interest rates, thereby affecting the value of the domestic currency. As prices gradually adjust later in response to the monetary shock, it will lead to reverse movements in interest rates and hence the currency value. Along with the adjustment in the currency rate during this phase, PPP deviations will dwindle at a rate depending upon how fast prices can adjust. The process of reversion continues until a long-run equilibrium consistent with PPP is reached. Accordingly, the real exchange rate will be more persistent the more sluggish the price adjustment is.

Rogoff (1996) points out that if PPP deviations are really driven by sluggish price adjustment, the real exchange rate should be expected to converge at a much faster rate than what has typically been found. The empirical rate of convergence appears far too slow to be explained by price stickiness, however. This poses a serious challenge for the literature. No existing macroeconomic models can consistently explain both the vast short-term volatility and the "excessively" high persistence observed in the real exchange rate.

In looking beyond the influence of macroeconomic fundamentals, Taylor (1995) recognizes the possible role of microstructural factors—including the behavior of foreign exchange market agents—in generating short-term PPP deviations. For example, the rising importance of chartists in currency trading can extend and magnify the short-term impact of market shocks on exchange rate movements. Based on survey expectations data for major currencies, Frankel and Froot (1990) report that, "at short horizons, [traders] tend to forecast by extrapolating recent trends, while at long horizons they tend to forecast a return to a long-run equilibrium such as purchasing power parity" (p.183). If short-term exchange rate responses can be magnified by trend-following currency trading, they will impart similar behavior into real exchange rates, especially when operating under sticky prices.

This study reviews and discusses empirical evidence, which corroborates the existence of amplified short-term responses of real exchange rates. These amplified shock responses—which tend to magnify PPP deviations initially—can not only contribute to the short-term volatility of the real exchange rate but also prolong substantially the time it takes for the real exchange rate to converge to parity.

2 Direct Evidence of Parity Reversion

In analyzing the mean-reverting property of real exchange rates, conventional unit root tests are known to be afflicted by low statistical power, leading to the widespread failure to find reversion toward PPP in early studies of the modern floating-rate period. The problem may be aggravated by the high volatility of floating exchange rates, making it difficult to detect parity reversion in the noisy data. Several approaches have

been advanced to overcome the power problem, nevertheless. They include the use of long-horizon data to extend the sample period (Diebold, Husted and Rush, 1991; Lothian and Taylor, 1996). This method involves using data from the pre-float period. Another approach uses cointegration tests with good power to explore the long-run relationship between exchange rates and relative prices (Cheung and Lai, 1993; Edison, Gagnon and Melick, 1997). An alternative approach considers pooling data across real exchange rates in panel unit root tests (Frankel and Rose, 1996; Oh, 1996; Papell, 1997), but the robustness of panel test results has been called into question (Engel, Hendrickson and Rogers, 1997; O'Connell, 1998; Taylor and Sarno, 1998). Still another approach is to use efficient unit root tests with optimal power. Cheung and Lai (1998) employ this direct approach and unveil significant evidence of PPP reversion without using either long-horizon or panel data. The direct approach is adopted here.

The data under study are monthly real exchange rates constructed from nominal exchange rates and consumer price indices. Specifically, the real exchange rates of four European countries—France (FR), Germany (GE), Italy (IT), and the United Kingdom (UK)—vis-à-vis the United States (US) are investigated. Taken from the International Monetary Fund's *International Financial Statistics* data CD-ROM, the data cover the sample period from April 1973 through December 1996. All the series of real exchange rates are expressed in logarithms, following the common practice in previous PPP studies.

Before analyzing the intertemporal path of adjustment, we first establish evidence of mean reversion for the individual series of real exchange rates. The efficient unit root test devised by Elliott, Rothenberg and Stock (1996) is carried out. These authors establish the asymptotic power envelope for unit root tests by analyzing the sequence of Neyman-Pearson tests of the null hypothesis H_0: $\rho = 1$ against the local alternative H_a: $\rho = 1 + \bar{c}/T$, where ρ is the largest autoregressive (AR) root in the AR($k + 1$) model, T is the sample size and $\bar{c} < 0$. Based on asymptotic power calculation, it is shown that a modified Dickey-Fuller test, called the DF-GLS test, can achieve significant power gains over standard unit root tests. The superior performance of the DF-GLS test is also supported by the Monte Carlo results reported by Stock (1994). Although the DF-GLS test shares very similar size properties as the ADF test, the former shows much better test power than the latter.

For a real exchange rate series, denoted by $\{y_t\}$, the DF-GLS test entails the following regression:

$$(1 - L)y_t^{\mu} = \phi_0 y_{t-1}^{\mu} + \Sigma_{j=2}^{k} \phi_j (1 - L) y_{t-j+1}^{\mu} + v_t \tag{1}$$

where L is the usual lag operator such that $Ly_t = y_{t-1}$; v_t is the random error term; and y_t^{μ}, the locally demeaned data process under the local alternative of $\bar{\rho} = 1 + \bar{c}/T$, is given by

$$y_t^{\mu} = y_t - \tilde{g}' z_t \tag{2}$$

Table 1: Results from the ADF and DF-GLS Unit Root Tests

Test	Series	k	Statistic	10% CV	5% CV
ADF	FR/US	4	-2.148	-2.864	-2.571
	GE/US	2	-1.868	-2.859	-2.566
	IT/US	2	-2.021	-2.859	-2.566
	UK/US	4	-2.337	-2.864	-2.571
DF-GLS	FR/US	4	-2.152**	-1.684	-2.000
	GE/US	2	-1.872*	-1.689	-2.006
	IT/US	2	-2.009**	-1.689	-2.006
	UK/US	4	-1.874*	-1.684	-2.000

Notes: The column beneath "k" gives the lag parameter chosen using the Akaike information criterion. Finite-sample critical values (CV) for the ADF test are obtained from Cheung and Lai (1995a) based on response surface estimation for $T = 285$. Finite-sample CVs for the DF-GLS test are from Cheung and Lai (1995b) for $T = 285$, as described by Eq. (4). Asymptotic CVs for the DF-GLS test are given, respectively, by -1.62 and -1.95 for the 10% and 5% significance levels. Statistical significance is indicated by a single asterisk (*) for the 10% level and a double asterisk (**) for the 5% level.

with $\hat{g}$ being the least squares coefficient estimated from regressing $\tilde{y}_t$ on $\tilde{z}_t$:

$$\tilde{y}_t = g'\tilde{z}_t + \epsilon_t \qquad (3)$$

for which $\tilde{y}_t = (y_1, (1 - \bar{\rho}L)y_2, ..., (1 - \bar{\rho}L)y_T)'$ and $\tilde{z}_t = (z_1, (1 - \bar{\rho}L)z_2, ..., (1 - \bar{\rho}L)z_T)'$. In general, $z_t = (1, t)$, allowing for a linear trend. No time trend is considered in our case here, so $z_t = 1$. The DF-GLS statistic is given by the conventional t-ratio, testing H_0: $\phi_0 = 0$ against H_a: $\phi_0 < 0$. The parameter, $\bar{c}$, which defines the local alternative through $\rho = 1 + \bar{c}/T$, is recommended to be set equal to -7 for the no-trend case.

Finite-sample size properties of the DF-GLS test have been explored by Cheung and Lai (1995b). Approximate finite-sample critical values (CV) can be computed from a response surface equation of a polynomial form:

$$CV_{T,k} = \eta_0 + \Sigma_{i=1}^{2}\eta_i(1/T)^i + \Sigma_{j=1}^{3}\xi_j(k/T)^j \qquad (4)$$

where $CV_{T,k}$ is the critical value estimate for a sample size T and lag k; and the relevant parameter values for $\{\eta_0, \eta_1, \eta_2, \xi_1, \xi_2$ and $\xi_3\}$ are tabulated by Cheung and Lai (1995).

Table 1 contains the statistical results obtained from the DF-GLS test. To facilitate comparison, results from the ADF test are reported as well. When a time trend was included, it was statistically insignificant in all the four cases. Accordingly, the results for the no-trend case are reported. For the choice of the lag parameter, k, data-dependent lag selection is implemented using the Akaike information criterion. In contrast to the standard ADF test, which consistently fails to identify stationarity, the test results from the efficient DF-GLS test indicate significant evidence in favor of PPP reversion in all the cases under examination. More specifically, the hypothesis of a unit root can be rejected in favor of stationary alternatives at either the 10% significance level in the GE/US and UK/US cases or the 5% significance level in the FR/US and IT/US cases. The results are consistent with those reported by Cheung and Lai (1998), who examined a shorter sample of real exchange rate data for FR/US, GE/US, and UK/US (but not for IT/US) and uncovered parity reversion in the data series using the DF-GLS test.

In analyzing long historical data, Culver and Papell (1995) and Perron and Vogelsang (1992) report that the behavior of real exchange rates can be characterized by trend-break models. Hegwood and Papell (1998) illustrate that the presence of structural breaks can cause a significant upward bias in the estimation of half-life persistence of PPP deviations in long-horizon data. As part of the preliminary data analysis, different trend-break unit root tests devised by Banerjee, Lumsdaine and Stock (1992)—henceforth BLS—were performed on the recent float data. The BLS tests involve the following regression:

$$(1 - L)y_t = \mu_0 + \mu_1 t + \mu_2 d_t(n) + \beta_0 y_{t-1} + \Sigma_{j=1}^{p}\beta_j(1 - L)y_{t-j} + \zeta_t \qquad (5)$$

where $d_t(n)$ is a dummy variable and ζ_t is the random error term. When a trend shift is allowed for at time n, $d_t(n) = (t - n)I(t > n)$, with $I(\cdot)$ being the indicator function. Alternatively, when a mean shift (or a break in the trend) is allowed for at time n, $d_t(n) = I(t > n)$. For the usual Dickey-Fuller test, $d_t(n) = 0$. A sequence of t-statistics for testing $\beta_0 = 0$, denoted by $\tau_{DF}(n)$, can be generated by varying n over the sample. BLS discuss different versions of the mean-shift or trend-shift sequential test. The minimal sequential test is applied in this analysis, and its test statistic is defined by

$$\tau_{DF}^{min} \equiv \min_{r \le n \le T-r} \tau_{DF}(n) \qquad (6)$$

for the sample size, T, and a trimming parameter, r. Following BLS, r is set equal to the integer part of $.15T$. According to the BLS test results (not reported here), in no case could significant evidence be found to support the relevance of trend-break models in explaining the real exchange rate dynamics over the recent float, regardless of whether mean-shift or trend-shift models were entertained.

3 Analyzing the Adjustment Process Toward Parity

Although the findings of no unit root in the real exchange rate confirm the long-run convergence to parity, they offer no specific information on the dynamic process of adjustment itself. The question is, How do deviations from PPP behave over the short or medium run? Such information may bear upon the issue concerning the slow rate of parity convergence. To obtain the relevant information, impulse response analysis can be used.

Given that the real exchange rate is shown to be stationary, its dynamics can be captured in general by an autoregressive moving-average (ARMA) model as follows:

$$B(L)y_t = D(L)u_t \qquad (7)$$

where $B(L) = 1 - b_1L - \ldots - b_pL^p$; $D(L) = 1 + d_1L + \ldots + d_qL^q$; all roots of $B(L)$ and $D(L)$ are stable; and u_t is the white-noise innovation term. The persistence of the process over different time horizons can be analyzed by studying the moving-average representation for y_t: $y_t = C(L)u_t$ with

$$C(L) = B^{-1}(L)D(L) \qquad (8)$$

where $C(L) = 1 + c(1)L + c(2)L^2 + \ldots + c(j)L^j + \ldots$ Consider a unit shock to the process. The impact of a unit innovation at time t on the level of y at time $t+j$ is given by $c(j)$. This $c(j)$ measure—referred to as the impulse response—summarizes the basic information concerning persistence over all time spans up to infinite after the initial shock. For a stationary process, which contains no unit root, the infinite impulse response is $c(\infty) = 0$. A stationary process thus has zero long-run persistence. Over horizons much shorter than infinity, on the other hand, $c(j) \neq 0$ and sizable persistence can still exist over the short or medium run.

Instead of studying the entire sequence of $c(j)$, $j = 1, 2, \ldots$, a simple summary measure of persistence typically employed in the PPP literature is the half-life, which indicates how long it takes for the impact of a unit shock on the real exchange rate to dissipate by half. By definition, the half-life, denoted by ℓ_h, is given by $c(\ell_h) = ½$. Since discrete-time data are analyzed, and the half-life does not have to be exactly an integer number, the approximate value of ℓ_h will be calculated using a simple interpolation method when $c(j) < c(\ell_h) = ½ < c(j+1)$ for some j.

4 Empirical Findings of Overreaction in Initial Responses

The adjustment dynamics of real exchange rates in response to a shock to parity are examined through the sequence of cumulative impulse responses, $c(j)$. The DF-GLS test applied earlier is based on approximating AR models. Using the GLS estimates of the fitted AR models, impulse response functions are constructed for the individual series of real exchange rates.

Table 2 reports estimates of up to the first 120 cumulative impulse responses, which cover a time span of 10 years for monthly data. The results reveal the existence

Table 2: Time-Profiles of $c(j)$ Estimates for Real Exchange Rates

j	FR/US	GE/US	IT/US	UK/US
0	1.000	1.000	1.000	1.000
1	1.263	1.277	1.337	1.339
2	1.282	**1.336**	1.431	1.358
3	1.319	1.329	**1.437**	1.365
4	**1.364**	1.304	1.412	**1.391**
5	1.333	1.273	1.376	1.371
6	1.307	1.241	1.338	1.334
7	1.282	1.209	1.298	1.301
8	1.245	1.178	1.260	1.267
9	1.206	1.148	1.222	1.232
10	1.168	1.119	1.185	1.196
11	1.131	1.090	1.150	1.162
12	1.093	1.062	1.116	1.129
13	1.056	1.035	1.082	1.096
14	1.021	1.008	1.050	1.065
15	0.987	0.982	1.018	1.034
20	0.830	0.862	0.874	0.893
25	0.699	0.757	0.751	0.771
30	0.588	0.665	0.645	0.666
35	0.495	0.584	0.554	0.575
40	0.416	0.513	0.476	0.497
50	0.295	0.395	0.351	0.370
60	0.209	0.305	0.259	0.276
70	0.148	0.235	0.191	0.206
80	0.105	0.181	0.141	0.154
90	0.074	0.140	0.104	0.115
100	0.052	0.108	0.077	0.085
110	0.037	0.083	0.056	0.064
120	0.026	0.064	0.042	0.047

Notes: Columns 2-5 give the cumulative impulse responses of individual real exchange rate series at different time horizons, j, following a unit shock to parity. For each real exchange rate series, the number in boldface shows where the value of $c(j)$ peaks after the shock and the corresponding size of the shock amplification.

of non-monotonicity in the process of convergence to parity. The cumulative impulse responses are not a monotonic function of the adjustment horizon. Although the eventual decay of $c(j)$ toward zero affirms the existence of parity reversion, the real exchange rate initially tends to overreact to the shock in the sense that it continues its momentum to move farther apart from its long-run equilibrium level. Consequently, the PPP deviation tends to magnify first before diminishing (see Figure 1). Such non-monotonic responses of the real exchange rate—albeit they occur over merely a short time period—can delay and substantially prolong the process of convergence. In the cases studies here, it takes at least 15 months to offset the impact of the amplified responses and for $c(j)$ to simply return back to the unity level. The results here are similar to those obtained by Cheung and Lai (1999) from ARMA models, supporting the robustness of the impulse response results with respect to model specifications.

The presence of short-term overreaction in shock responses leaves the real exchange rate to adjust to, in effect, a more-than-unit shock during the subsequent reversion phase. For a given rate of decay of PPP deviations, the half-life persistence for the real exchange rate can vary widely, depending on the size of the short-term overreaction. If the size of the overreaction is considerable, the calculated half-life will show up to be long even when PPP deviations can die out at a relatively fast speed. This can be illustrated using a numerical example. Suppose that the real exchange rate reverts at the same speed as price adjustment, say, at a speed of $\ell_h = 2$ years. This value of ℓ_h implies that PPP deviations will dampen out at a rate of about 29.3 percent per year. Suppose also that there is shock amplification by a factor of 1.5 during the first month in response to a monetary shock. Under this situation, it can be shown that it will take about 1.2 year just to reverse the impact of the initial overreaction. Moreover, the situation will produce a ℓ_h estimate of roughly 3.3 years, much higher than the corresponding ℓ_h value of 2 years in the absence of the initial shock amplification. Table 3 illustrates a positive relationship between the size of the shock amplification and its contribution (in percentage) to the observed half-life persistence. In the case in which the amplification factor equals 1.5, the short-term overreaction can explain in excess of one-third of the observed persistence.

The half-life estimates are computed from our actual data on real exchange rates. Confidence intervals of the estimates are presented as well. For a given horizon, j, $c(j)$ is a nonlinear function of the ARMA parameter vector, $\zeta = (b_1, ..., b_p, d_1, ..., d_q)$. Using the delta method, standard errors for $c(j)$ can be estimated from

$$\mathrm{Var}(c(j)) = \nabla c'(j)\Omega\nabla c(j) \qquad (9)$$

where $\nabla c(j)$ is $\partial c(j)/\partial\zeta$ and Ω is the variance-covariance matrix of ζ (Campbell and Mankiw, 1987). The sample standard errors will be used to construct confidence intervals for $c(j)$, from which confidence intervals for ℓ_h estimates are derived. It should be noted that these asymptotic estimates of standard errors do not correct for any finite-sample bias and can understate the level of potential imprecision associated

Table 3: The Impact of Overreacting Responses on Half-Life Persistence

Amplification Factor	Proportion of ℓ_h Explained
1.0	0.0%
1.1	15.3%
1.2	23.4%
1.3	29.6%
1.4	34.4%
1.5	38.5%
1.6	41.9%
1.7	44.6%
1.8	47.1%
1.9	49.2%
2.0	51.0%
2.1	52.6%
2.2	54.1%

Notes: In the benchmark case of no short-term overreaction, the amplification factor equals 1.0. The second column gives the computed proportion of the observed half-life attributable to the overreaction. The computation assumes that the overreaction occurs during the first month only.

Table 4: Half-Life Estimates

Variable	FR/US	GE/US	IT/US	UK/US
ℓ_h	2.89	3.41	3.20	3.31
L_{95}	1.36	1.44	1.40	1.41
L_{90}	1.50	1.61	1.56	1.57
U_{90}	5.38	6.85	6.09	6.56
U_{95}	5.79	7.41	6.56	7.07

Notes: The column "ℓ_h" provides the point estimates of the half-life adjustment speed (in years). $[L_{90}, U_{90}]$ represents the 90% confidence interval for ℓ_h; whereas, $[L_{95}, U_{95}]$ gives the 95% confidence interval for ℓ_h.

with half-life estimation. Berkowitz and Kilian (1998) and Kilian (1998) recently advocate the use of bootstrapping methods to evaluate sampling uncertainty. For example, the distribution of the innovation term can be approximated by the empirical distribution of the estimated residual using resampling (with replacement) techniques.

Table 4 gives the half-life persistence estimates for the individual real exchange rate series. The point estimates of ℓ_h range from 2.9 to 3.4 years and yield an average of about 3.2 years. As noted by Rogoff (1996), these half-life estimates seem too long to be explained by price stickiness because they suggest—according to Eq. (10) below—a very slow rate of convergence of about 19 percent per year on average.

Pesaran and Shin (1996) investigate the speed of convergence to PPP under a multivariate cointegration framework. Unlike the univariate time series method considered here, these authors analyze the equilibrium relations between exchange rates, prices and interest rates in the case of the UK. Specifically, the persistence profiles of both the PPP relation and the UIP (uncovered interest-rate parity) relation are estimated simultaneously. Their point persistence estimates show that the estimated rate of convergence to PPP is very slow, while the convergence to UIP seems rather fast. Interestingly, these authors also observe that the persistence profile for the PPP relation is hump-shaped such that a shock to PPP tends to magnify first before reverting.

Reversion speeds have often been computed indirectly from half-life estimates under an implicit assumption of monotonic convergence at a constant speed (AS): the speed is computed from ℓ_h as

$$AS = 1 - \exp[\ln(½)/\ell_h]. \tag{10}$$

Such an assumption is not valid in general situations, nonetheless. In particular, the convergence can be not monotonic. Non-monotonic convergence may arise when undershooting occurs before reverting back to parity. The convergence may also come in the form of oscillating dynamics. In short, the simple half-life measure unsatisfactorily ignores the actual structure of the adjustment process. In our case, shock responses are found to amplify before dissipating, leading to non-monotonic dynamics. Such short-term amplification of shock responses can delay and protract the adjustment process, leading to the slow convergence to parity. The non-monotonicity confounds the half-life measure to give distorting estimates of the adjustment speed.

5 A Direct Measure of Adjustment Speed

A direct measure of the adjustment speed, which is more informative than the half-life measure, can be obtained from the cumulative impulse response function. Unlike the half-life measure, which is single-valued, the alternative measure is a function of time, giving the actual rate at which the impact of a shock die out along the entire path of adjustment. More specifically, the adjustment speed of the real exchange rate at time

Table 5: Time-Profiles of the Adjustment Speed (Per Month) of Real Exchange Rates

j	FR/US	GE/US	IT/US	UK/US
0	-26.29%	-27.71%	-33.65%	-33.94%
1	-1.49%	-4.61%	-7.05%	-1.39%
2	-2.91%	**0.49%**	-0.45%	-0.52%
3	-3.40%	1.93%	**1.75%**	-1.86%
4	**2.30%**	2.38%	2.53%	**1.42%**
5	1.94%	2.51%	2.82%	2.67%
6	1.93%	2.55%	2.93%	2.51%
7	2.89%	2.56%	2.97%	2.58%
8	3.13%	2.56%	2.99%	2.80%
9	3.09%	2.57%	3.00%	2.86%
10	3.24%	2.57%	2.99%	2.86%
11	3.33%	2.57%	2.99%	2.87%
12	3.32%	2.57%	3.00%	2.88%
13	3.35%	2.57%	2.99%	2.89%
14	3.38%	2.56%	3.00%	2.88%
15	3.39%	2.58%	3.00%	2.89%
20	3.40%	2.57%	3.00%	2.89%
25	3.39%	2.58%	3.00%	2.89%
30	3.40%	2.56%	3.01%	2.90%
35	3.40%	2.57%	3.00%	2.89%
40	3.41%	2.58%	2.99%	2.90%
45	3.40%	2.56%	3.01%	2.89%
50	3.40%	2.58%	2.99%	2.89%
60	3.41%	2.59%	2.98%	2.90%
70	3.39%	2.56%	2.99%	2.87%
80	3.44%	2.54%	2.98%	2.93%
90	3.38%	2.57%	3.08%	2.88%
100	3.35%	2.60%	3.00%	2.93%
110	3.50%	2.53%	3.01%	2.98%
120	3.44%	2.50%	3.02%	2.95%

Notes: Columns 2-5 present the adjustment speed (AS_j) estimates for individual real exchange rate series at different time horizons, j, subsequent to a shock to parity. For each real exchange rate series, the number in boldface indicates when the shock amplification ends and when the PPP deviation begins to die out.

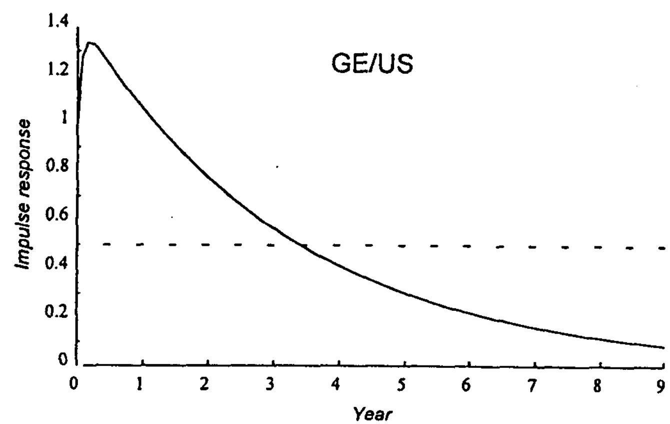

Figure 1. A sample plot (the GE/US case) of the dynamic responses of the real exchange rate to a unit shock.

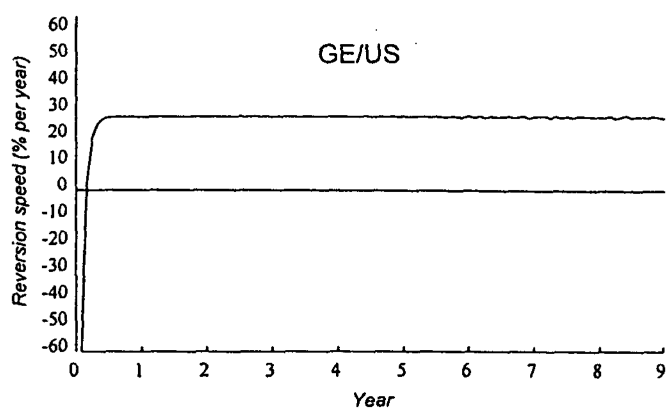

Figure 2. A sample plot (the GE/US case) of the different speeds of real exchange rate adjustment over time.

t is given by

$$AS_t = -[dC(t)/dt]/C(t) \tag{11}$$

which corresponds to the instantaneous (percentage) rate of decrease in the cumulative impulse response at time t. The AS_t measure is easy to interpret. When $AS_t > 0$, the real exchange rate is reverting toward parity at time t, and the magnitude of AS_t indicates the relevant speed. When $AS_t < 0$, on the other hand, the real exchange rate is moving further away from parity at time t at a speed of $|AS_t|$ per unit time. In this way, a researcher can compute and gauge both the direction and the speed at which the adjustment takes place at any time horizons after the initial shock.

Table 5 reports the time profile of adjustment speeds at various time horizons following a shock to parity. In every case, the adjustment speed starts from a negative value, showing the impact of the initial shock amplification (see also Figure 2). Reverting dynamics then take over quickly and then attain a steady positive speed toward parity. The AS_t estimates show that subsequent to the short-term amplified responses, real exchange rates converge at a rate of between 2.6 to 3.4 percent per month—an equivalent rate of between 31 to 41 percent per year—which is much faster than what half-life estimates have implied. Accordingly, the short-term amplified responses may create the appearance of slow reversion when measuring in terms of the half-life.

6 Concluding Remarks

The short-term adjustment of the real exchange rate has been found to be characterized by overreaction and amplified shock responses. Such dynamics can contribute to the large, volatile short-term PPP deviations. They can also delay and prolong the process of convergence to parity. Although the short-term overreacting dynamics may be viewed as overshooting behavior in the broad sense that reversion occurs only after persistent movements away from the long-run equilibrium, the dynamic adjustment pattern identified for the real exchange rate seems not compatible with the Dornbusch-type rational expectations models of overshooting. Specifically, short-term exchange rate overshooting under Dornbusch's (1976) model happens initially at the time of the shock only such that the maximal impact of the shock occurs contemporaneously. Following the shock, the real exchange rate reverts to its long-run value monotonically. This contrasts with our findings, in which the full impact of the shock is not felt immediately but until a few periods after the initial shock. It follows that the amplified responses observed in the real exchange rate cannot be explained by the conventional models of exchange rate overshooting.

On the other hand, the findings of short-term amplified responses of the real exchange rate—with its subsequent reversal and gradual reversion toward parity —seem consistent with the chartist-fundamentalist view of exchange rate dynamics, as identified in survey data on expectations of foreign exchange market participants (Allen and Taylor, 1990; Frankel and Froot, 1990, 1993). Chartists are market agents who like to follow recent trends and tend to have bandwagon expectations. Fundamentalists, in contrast, are market agents who base their forecasts on economic fundamentals and such forecasts tend to be regressive. Empirical findings from survey data generally suggest that exchange rate forecasts over short horizons are dominated by chartist analysis; whereas, exchange rate forecasts over long horizons are governed by fundamental analysis. Cheung and Wong (1999) explore the market practitioners' views on exchange rate dynamics and confirm the presence of bandwagon effects and short-term overreaction to news. To the extent that short-term currency trading is in large part spurred by bandwagon expectations, the effects of market shocks on real exchange rates will tend to amplify before dissipating.

The chartist-fundamentalist model also offers a possible explanation for persistent deviations from UIP, as noted by Eichenbaum and Evans (1995). In response to an expansionary monetary shock, for example, the domestic interest rate falls and the exchange rate (the dollar price of foreign currency) rises. The rise in the exchange rate will continue for a while after the initial shock because there are chartist traders who jump on the bandwagon, buying the foreign currency and causing further appreciation. When the rise in the exchange rate extends beyond the initial shock, the relative decrease in the U.S. interest rate will not be offset by an expected appreciation of the dollar, thereby giving rise to persistent expected excess returns and sustained deviations from UIP.

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