ABSTRACT

INTEREST RATE DIFFERENTIALS AND PURCHASING POWER PARITY PUZZLE:
A NONLINEAR PERSPECTIVE

The aim of this paper is to construct a simple nonlinear model for the U.S. dollar - German mark (euro) real exchange rate. The central bank is shown to have an important role in the exchange rate system when the real exchange rate is measured by the CPI deflator. The monetary authority is assumed to follow a Taylor rule and to raise interest rate when the home currency is depreciated relative to its long-run PPP level.

We use a nonmonotonic logistic smooth transition model to investigate a nonlinear adjustment towards the parity level. We begin our empirical part by examining the statistical properties of the data. The results of our analysis suggest that when allowing for a non-linear alternative the real exchange rate seems to follow a globally stationary process. The nonlinear model considered allows the adjustment towards long-run equilibrium to be sudden as well as smooth. We found that the adjustment is sudden. This is probably due to official announcements of the central banks. Furthermore, there is a symmetric and relatively large reference band around the parity level. This implies that the weight put on the U.S dollar-German mark (euro) real exchange rate in the monetary policy rules is low.

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1. INTRODUCTION

The wide swings in the value of the U.S. dollar- Euro during the past three years have again rekindled interest in the search for understanding exchange rate movements. The aim of this paper is to present empirical evidence for the interrelations between the purchasing power parity (hereafter PPP), uncovered interest rate parity (UIP) and the terms structure of interest rates (TS). Our final goal is to construct a multivariate and nonlinear model for the U.S. dollar - German mark (euro) real exchange rate. The PPP is measured using CPI deflators.

We examine the U.S. dollar - German mark (euro) real exchange rate over the period 1975 to 2003. The data for multivariate model consists of observations from 1982 to 2003. The bilateral exchange rate between the United States and Germany has been chosen in order to fully understand the complicated interrelationship between USA and Europe. Because of the different economic structure of the Euro-area relative to Germany and also the different operational objectives of the European Central Bank relative to the Bundesbank, the Euro may, of course, have different properties compared to the German mark. Nevertheless, there is a good reason to believe that the study of the dollar-mark system is useful in terms of indicating how the exchange rate between two large economies may be determined in the market.

The keynote in our model is that the deviations from the PPP equilibrium are stationary. Although the stationarity assumption is theoretically plausible between two industrialized countries, it is a common finding in an international finance literature that real exchange rates contain a unit root for the recent floating period. If the real exchange rate after all is a non-stationary variable, standard inference is not valid and our findings can be questioned. Thus, we should consider a stationary assumption a bit more carefully.

While we build on the stationarity assumption, we also allow for the possibility that the real exchange rate follows non-linear time series process. A number of studies have utilized variants of threshold autoregressive models to capture nonlinear behavior of real exchange rate. Michael et al. (1997), Baum et al.(2001), Taylor et al. (2001) and Sollis et al.(2002) apply a smooth transition autoregression (STAR) models. They provide considerable evidence on nonlinear real exchange rates. Intuitively, nonlinearities in the real exchange rate imply that frictions in international trade result in bands within which relative international prices can fluctuate without a strong tendency to adjust towards the parity level.[1] Large price differentials converge towards transaction cost differentials, not towards absolute parity. Once nonlinearity is modeled, the speed of mean reversion is less inconsistent with the traditional arbitrage conditions.

PPP is, however, somewhat distinct from the pure law of one price concept applied to commodities. Our view is that PPP is examined more properly in the domain of monetary economics and macroeconomic theory of inflation. What we attempt to do in this paper is to show that the central bank has an important role in an exchange rate system when the real exchange rate is measured by the CPI deflator. The central bank is assumed to respond to deviations of the expected inflation and output from their desired levels, i.e. the central bank follows a Taylor rule. Several papers have added terms in exchange rates to otherwise standard Taylor rules (e.g. Clarida et al., 1998, Alexius, 2000, Engel and West, 2002). To isolate the effects of relative international consumer prices on the target function, we also use a Taylor rule with the real exchange rate explicitly included.

The monetary authority is assumed to raise interest rate when the home currency is depreciated relative to its long-run PPP level. Using uncovered interest rate parity (UIP) as an identifying restriction, we show that the real exchange rate today can be expressed as the expected sum of future monetary policy shocks. Transaction costs imply that the weight put on the real exchange rate in the monetary policy rule is rather low close to PPP. There is no consensus in a neighborhood of PPP as to whether the exchange rate is overvalued or undervalued in terms of the expected monetary shocks. A time-varying uncertainty premium separates expected exchange rate changes from the expected interest rate differential. Thus, when transactions costs are presents, there is scope for persistent deviations from PPP. Those deviations might be related to economic fundamentals but also bubbles and herding might temporarily send the exchange rate off on equilibrium paths that result in the appearance of slow convergence to the PPP equilibrium, as noted by Engel and Morley (2001).

In our empirical work, we do not primarily attempt to identify and trace through the effects of expected monetary shocks. Instead, we do aim to carefully elaborate the idea that the low weight put on the real exchange rate in the monetary policy rule in a neighborhood of PPP result in bands within which the deviations from PPP are possible. Based on the assumption of heterogeneous agents, smooth transition models are considered. Since the role of the central bank might be crucial,

[1] This is called Heckscher's commodity point approach. The approach begins with Heckscher's (1916) insight that international transaction costs should create some scope for deviations from PPP level.

it is desirable to allow the adjustment towards long-run equilibrium to be discrete as well as smooth. We use a nonmonotonic second-order logistic smooth transition autoregressive model (LSTAR2) to investigate nonlinear mean reversion. This a special case of a three-regime switching regression in which the two outer regimes are equal.

We begin our empirical part by examining the statistical properties of the data. Using a traditional augmented Dickey- Fuller (ADF) unit root test and Eklund F-test (EF), i.e. the joint linearity and unit root test against the second- order LSTAR, we find a weak evidence against the null hypothesis of a unit root for our data measures. For the real exchange rate series, we cannot find evidence against a unit root hypothesis. Enders and Granger (1998), and Bec et al. (2001). have shown that standard unit-root tests have low power in the presence of nonlinear adjustment. Although power simulations in Eklund (2003) show some gain in power compared to the standard ADF test, also EF has low power discriminating a random walk from a stationary nonlinear process. Thus, we proceed with the stationarity assumption to find out whether the possible nonstationary of the real exchange rate is due to local unit root behavior.

Local unit root behavior is consistent with the findings that the real exchange rates series fluctuate without a strong tendency to return to the equilibrium level when they wander in a neighborhood of it. We find using a second-order STAR that the most of the observations belong to this regime. Large deviations from the parity level, however, return relatively quickly towards parity level. This may lead to somewhat different conclusion concerning nonstationarity . The results of our analysis seems to suggest that when allowing for a non-linear alternative the real exchange rate may follow a globally stationary process and standard inference should hold.

Finally, we employ a multivariate nonlinear model for the U.S. dollar - German mark (euro) real exchange rate. We found that the adjustment is sudden and symmetric around the parity level. To find strong support to the symmetry assumption, we have to include a temporary exchange rate premium in the estimation vector. The sudden and symmetric adjustment is probably due to coordinated responses of central banks to the large deviations from PPP. Furthermore, the band around the PPP equilibrium is relatively large. This implies that the bilateral dollar-mark exchange rate seems to fluctuate within a wide and flexible reference band.

2. PURCHASING POWER PARITY PUZZLE

The real exchange rate is measured using CPI deflators. We define purchasing power parity as follows:

$$q_t \equiv ppp_t = p_t - p_t^* - s_t . \qquad (2.1)$$

where s_t is the nominal exchange rate, p_t the domestic price index, p_t^* the foreign price index. Hence, the deviation of the nominal exchange rate from the underlying PPP fundamental is in fact the real exchange rate (q_t).

Sticky prices implies that the nominal exchange rate can move without producing an immediate proportional response in relative domestic prices. A typical half life is, however, found to be from three to five years.[2] That is too slow adjustment to be explained by nominal rigidities. If the sources of PPP disturbances are real in nature, we would argue this will have a more permanent effect on the real exchange rate. More permanent shocks also imply that the real exchange rate might be a non-stationary variable. But existing models based on real shocks cannot account for short-term exchange rate volatility. Hence, we have a purchasing power parity puzzle (Rogoff ,1996).

Engel's (1999) results suggest that consumers' prices for tradables goods behave very much in the same way as non-tradables consumer prices. This implies that there might be a band of inaction around the parity level within which arbitrage does not yield an expected net gain. This results in scope for persistent deviations from PPP. In most of theoretical models (e.g. Dumas,1992; Secru et al.,1995; Obstfeld and Rogoff, 2000) proportional or "iceberg" transport costs create a narrow and symmetric band within which the marginal cost of arbitrage exceeds the marginal benefits. As pointed out by Obstfeld and Taylor (1997), advances in the theory of investment under uncertainty imply, that a band of no-arbitrage should be interpreted as a resulting not only from concrete

[2] Kilian and Zha (2002) conduct a survey among international economist regarding their views on real exchange rate half-lives. The consensus half-life was found to be around four years.

shipping costs and trade barriers, but also from sunk costs of international arbitrage.[3] This will widen the band within which price differentials can fluctuate before arbitrage commences.

Arbitrage will be heavy once it is profitable enough to outweigh the cost of arbitrage, but inside a band of inaction the pass-through from the nominal exchange rate to import prices depends on several issues.[4] If cost are, for example, partly dominated in home currency, then foreign suppliers are not thoroughly foreign. Swings in currency value may also change either the elasticity of import demand or foreign suppliers' own elasticity if they face domestic competition. Furthermore, when markets are segmented and the price elasticities of demand are not constant, a monopolistically competitive firm's optimal pricing behavior can create a wedge between common currency prices of the same good destined to different market.

According to Goldberg and Knetter (1997) there appears to be price discrimination between European and American markets, dictated by distinct competitive conditions in those markets. We proceed with the assumption that the United States and Germany (Euro-zone) are countries with distinct competitive conditions but similar deterministic productivity trends, i.e. there is an identical production technology in the long-run. Furthermore, as discussed in De Grauwe and Grimaldi (2001, pp. xx), to the extent that the observed real exchange rate changes fall within the transaction costs band, they cannot be associated with long-term divergent developments in productivity or other real variables. In the next Chapter we pursue an idea that the persistent nominal exchange rate deviations from PPP are due to the transactions costs and expected monetary policy shocks.

3. A GENERAL MONETARY REAL EXCHANGE RATE MODEL

The central bank is assumed to respond to deviations of expected inflation ($E_t \pi_{t+1}$) and output (y_t) from their desired levels, i.e. monetary policy rules in the foreign and home countries follow a Taylor rule. We consider a Taylor rule with the real exchange rate explicitly included.

$$i_t = \psi_q q_t + \psi_\pi (E_t \pi_{t+1} - \bar{\pi}) + \psi_y (y_t - \bar{y}_t) . \tag{3.1}$$

[3] O'Connell and Wei (1997), provide a continuous time model on goods arbitrage that highlights the relative importance of proportional and fixed cost of transactions. Their evidence indicates that the fixed component is dominant.
[4] Note, that our approach is the partial equilibrium standpoint of a firm that takes the exchange rate as given, i.e. we exclude the discussion on the source of shock.

where i_t denotes the difference in between home and foreign short-term interest rate ($i_t - i_t^*$). Also other standard variables in (3.1) denote differentials. We assume both countries have same monetary policy parameters and parameter restrictions are $\psi_\pi > 1$ and $\psi_y > 0$. E_t denotes expectations on the basis of time t information. We omit constant and trend terms here and throughout.

q_t is the real exchange rate. To keep things as simple as possible, we further assume that the real exchange rate is only included in the home country monetary policy rule. Consistently with a Taylor rule, short-run interest rate will respond to the deviations of the nominal exchange rate from the underlying PPP target. The real exchange rate parameter is assumed to obey $\psi_q > 1$. With $\psi_q > 1$, the monetary authority raises interest rate when the real exchange rate is depreciated relative to its long-run PPP level. We use $\psi_q q_t$ to isolate the effects of international consumer prices on standard target variables. The weight put on the real exchange rate in the monetary policy rule is positively related to the pass-through from the exchange rate to import prices and openness of economy.

We can write one-period uncovered interest rate parity (UIP) as follows:

$$i_t = E_t s_{t+1} - s_t + \rho_t , \qquad (3.2)$$

where ρ_t is the deviation from rational expectations UIP. It can be interpreted as a risk premium or an expectational error.[5] This implies that short-run UIP does not hold.

Subtract the expected inflation from both sides of (3.2) and use the definition of q_t. Then UIP in the real exchange rate is obtained as:

$$E_t q_{t+1} - q_t = i_t - E_t \pi_{t+1} - \rho_t . \qquad (3.3)$$

[5] We label ρ_t as an uncertainty premium.

Let υ_t denote an expected monetary policy respond to the deviations of standard domestic target variables in Equation (3.1). Using (3.3) to substitute out for i_t on the left hand side of (3.1), we get

$$q_t = \delta E_t q_{t+1} - \delta\ \upsilon_t + \delta\rho_t \tag{3.4}$$

where $\delta = \dfrac{1}{1+\psi_q}$, $0 < \delta < 1$. Forward iterate on equation (3.4) k-period ahead to find k-period solution for q_t. The solution exists if we have terminal condition which guarantees that q_t is non-explosive. Since we assume δ <1, the exchange rate is expected ultimately converge to the constant $\overline{q} = 0$.

If prices are set n-period ahead of time at the expected level, the solution can be written as:

$$q_t = \left(-\sum_{j=0}^{k} \delta^j \left(E_t \upsilon_{t+j} - E_{t-n} \upsilon_{t+j} \right) + \sum_{j=0}^{k} \delta^j E_t \rho_{t+j} \right) \tag{3.5}$$

Today's real exchange rate is expressed as the expected partial sum of future monetary policy shocks and uncertainty premiums.[6] Even if the weight, ψ_q, put on the real exchange rate in the monetary policy rule was small, introducing ψ_q in the model restricts the real exchange rate volatility that can be accounted for expected shocks. Increasing the weight decreases always the variance of the real exchange rate.

An expected contractive monetary policy shock appreciates both the real and nominal exchange rates in the short run.[7] Since δ_t <1, it leaves long-run deviations from PPP unaffected. Consider also a positive uncertainty shock that arises an uncertainty premium. A real depreciation will result, i.e. the opposite effect to the expected monetary policy shock. In the next subsection we examine

[6] Mark and Moh (2003) find evidence, that high expected future interest rate differentials are associated with a strong dollar today. The capitalization is, however, incomplete. This is consistent with delayed overshooting (as in Eichenbaum and Evans, 1995).
[7] Empirical findings support the conclusion that there is no instantaneous and complete adjustment to interest rate news. The adjustment appears to occur gradually (Engel and Hamilton,1990) or as delayed overshooting (Eichenbaum and Evans,1995).

the effects of future monetary policy shocks and uncertainty premiums on the real exchange rate in a little more detail.

3.1. THE EFFECTS OF SHOCKS ON REAL EXCHANGE RATE

The empirical evidence overwhelmingly rejects models that explain movements of the major exchange rates by the movements of underlying fundamentals. A well known stylized fact about nominal exchange rates among low-inflation advanced countries like the United States and Germany (Euro zone) is that their time series follow approximately random walks. Then the knowledge that eventually the exchange rate should adjust to PPP will have little significance. Indeed, there appears to be evidence that the large fluctuations in the nominal exchange rate cannot be interpret as movements towards fundamental PPP equilibrium, but instead as movements generated traders' behavior in the foreign exchange rate market that is not related to PPP (see Engel and Morley, 2001; Juselius and MacDonald, 2001). The extended Taylor rule (Eq.3.1) implies that fundamentally the time t real exchange rate depends on current and all future values of υ_t and ρ_t. Our fundamental model is not, however, contradictory to stylized facts discussed above, since the model is able to generate random walk time series behavior.

The involved intuition for the connection between the model and random walk behavior is easily clarified considering the central bank's monetary policy rule. The central bank is expected to extend only to monetary shocks that do not threaten its primary objective of low deviations of domestic target variables from their desired levels. The domestic target variables are less susceptible to shocks induced by exchange rate fluctuations if the international trade share on gross domestic product is small and/or the pass-trough from the nominal exchange rate to import prices is low. Frictions in international trade implies that the weight put on the real exchange rate in the monetary policy rule might be lower close to PPP than far away from it.

The low value of ψ_q parameter in a neighborhood of PPP indicates that more weight is being placed on expectations of fundamentals far into further (see Engel and West, 2004). It is difficult to forecast the distant future values of fundamentals. Thus, we argue that agents' lack of conviction regarding their expectations plays an important role in explaining a random walk behavior. There is no consensus in a neighborhood of PPP as to whether the exchange rate is overvalued or undervalued in terms of the expected monetary shocks. Thus, the exchange rate might be driven in part by noise (see Shleifer and Summers, 1990). This implies that there is a large uncertainty premium and the exchange rate may follow a process arbitrarily close to a random walk in a neighborhood of PPP although monetary fundamentals do not follow random walks.[8] Note also that an uncertainty premium in UIP does not necessary imply that agents include irrational factors in their expectation functions. Equivalently, ρ_t might be fundamental shock to the exchange rate causing uncertainty and at the same the appropriate correlation between the forward premium and the rational expectations of the future spot rate.

Kilian and Taylor (2003) state that rational market agents take stronger positions against exchange rates levels far away from the latent PPP equilibrium. In their model a consensus is gradually build among market agents that the exchange rate is misaligned as the exchange rate moves away from the equilibrium. This ensures a non-linear mean reversion of the exchange rate towards the PPP equilibrium. The expected variance of the real exchange rate is restricted by heterogeneous market agents expectations and the real exchange rate time series follows a random walk process only in a neighborhood of PPP.

Even though individual market agents may strongly believe the exchange rate to be misaligned, they hardly have the power or resources to broke the trend. If the exchange rate persistently disconnects from its fundamental PPP value, market agents may also lose their confidence in the usefulness of PPP trading rule.[9] Then an unanticipated monetary policy shocks might be necessary to broke a trend. Sarno and Taylor (2001) suggest that even sterilized interventions might be useful once the exchange rate has moved a long away from the fundamental equilibrium. Publicly announced interventions can now be seen as fulfilling a coordinated role in that they encourage individual market agents to enter market at the same time. This implies homogeneous respond to large deviations from the PPP equilibrium.

4. EMPIRICAL NONLINEAR REAL EXCHANGE RATE MODEL

We do not primarily attempt to identify the effects of shocks. We do, however, empirically aim to carefully elaborate the idea that the low weight put on the real exchange rate result in wide band within which the real exchange rate time series follows a random walk process. Exchange rate

[8] The combined effect , $E_t\rho_t - E_t\upsilon_t$, is close to zero or even positive.

levels far away from the PPP equilibrium are less likely to be persistent. Thus, the real exchange rate is a stationary variable which follows a nonlinear process. Uncertainty regarding expectations implies that a transition between two regimes might be discrete as well as smooth.

4.1. TEMPORARY SHOCKS

Most of the time the real exchange rate fluctuates within a band around the PPP equilibrium. The size of a band is determined by market agents' expectations on future monetary shocks. Inside the band the exchange rate follows random walk and the best traders can do is to forecast no change. The fluctuation around the parity level might be interrupted by rare periods of turbulence, when the real exchange rate, under influence of a succession of random shocks in the right direction, crosses over to the other implicit boundary. At such time, the deviations are expected to be temporary and to adjust towards parity.

We can relate interest rates of different maturity based on the expectations model of the term structure by assuming that a longer rate (k) is a weighted average of current and expected rates of shorter interest rates

$$i_{t,k} = \frac{1}{k}\sum_{j=0}^{k-1} E_t(i_{t+j}) + \vartheta_t \, .$$

Using UIP it holds at k-period horizon that

$$i_{t,k} = \frac{1}{k}\sum_{j=1}^{k} E_t \Delta s_{t+j} + \sum_{j=0}^{k-1} E_t \rho_{t+j} + \vartheta_t$$

Uncovered interest rate parity combined with term structure of interest rates implies that the time t k-period nominal interest rate differential is an average of current and all future expected short-term exchange rate movements. ρ_t is an UIP risk premium. ϑ_t contains term-premium differentials but also risk that is traditionally related to the conditional covariance of the asset with a benchmark portfolio return. Assuming the variance of the UIP premium is larger than that of the term premium

[9] Cheung and Chinn (2001) found out that at six month horizon 81 % of traders view PPP as irrelevant. At the very long horizon only 40% of traders agree that PPP has some influence.

(and other risk components), then large fluctuations in a time-varying premium ($\rho+\vartheta$) are induced by the large changes in UIP .[10]

An uncertainty premium is high around the latent PPP equilibrium. Only large deviations are expected to adjust towards PPP. If there is a larger than average deviation from PPP, this should induce large and temporary fluctuations in a residual of the following Engle-Granger cointegration regression:

$$i_{t,k} = \alpha + i_{t,s} + \phi_t ,$$

where ϕ_t is a stationary residual. The expected change in the exchange rate does not appear in the TS cointegration relationship since on average over full sample period it will have an expected value of zero. However, the non-zero expected exchange rate change far away from the PPP equilibrium will be captured in the short-run dynamic. Thus, a stationary residual in above cointegration equation mimics the shock to the exchange rate which reduces uncertainty premium.

4.2 THRESHOLD METHODOLOGY

When using a nonlinear methodology for the analyzing price convergence, a discrete threshold methodology is found to be appropriate in one good world. Obstfeld and Taylor (1997), for example, use a self-exciting autoregressive (SETAR) model where the reversion is towards the edge of the band. They identify reasonable convergence speeds for disagregated tradable goods baskets.[11] Outside this simple analytical structure, the specification of fixed thresholds becomes problematic. Moreover, when the real exchange rate is measured using price indices made up of goods prices each with a different size of international arbitrage costs, one would expect adjustment of the overall real exchange rate to be smooth rather than discontinuous, as noted by Taylor et al (2001). Teräsvirta (1994) show, in turn, that time aggregation is also likely to result in smooth regime changes rather than discrete as long as heterogeneous agents do not act simultaneously.

[10] Bekaert *et. al.* (2002) find that the relative term premiums are not economically important in a dollar-mark system. They also find that the variability of a combined risk premium is mostly accounted for by variation in UIP risk premium.

[11] Obstfeld and Taylor (1997) use data measured relatively to the US after 1980.

An alternative characterization of the discrete nonlinear adjustment is provided by smooth transition autoregressive models (STAR). Here, in contrast with discrete SETAR model, regime changes occur gradually and are determined with a smooth function, which need only be continuos and non-decreasing. The assumption of heterogeneous agents and the need for symmetry in the response to positive and negative deviations from PPP leads empirical studies to the exponential STAR (ESTAR) model.

$$G(\gamma, c, z_t) = 1 - \exp\{-\gamma(z_t - c)^2\} \qquad \gamma > 0. \tag{4.5}$$

The transition function goes from zero to one as z_t, the transition variable, increases. The slope parameter γ indicates how rapid the transition from zero to unity is as a function of z_t. Finally, c is the location parameter, which determines where the transition occurs. Michael *et al.*(1997), Baum, *et al.*(2001) and Taylor *et al.* (2001), among others, apply ESTAR model and find support for the nonlinear representation.[12]

The transition function in ESTAR model is symmetric about c and $G(\gamma, c, z_t) \rightarrow 1$ for $z_t \rightarrow \pm\infty$. A drawback of the exponential transition function is that for either $\gamma \rightarrow 0$ or $\gamma \rightarrow \infty$, the transition function collapses to a constant and the model becomes linear in both cases. Hence, the ESTAR model does not nest a SETAR model as a special case, as noted by van Dijk et al. (2000). An exponential transition function is a suitable transition function if we assume non-linearity in the model is only due to symmetric and proportional transaction costs. As discussed above we also consider other sources of non-linearity. Thus, we use a second-order logistic function (LSTAR2).

$$G(\gamma, c, z_t) = (1 + \exp\{-\gamma(z_t - c_1)(z_t - c_2\})^{-1}, \tag{4.6}$$

A nonlinear mean reversion does not need to be a symmetric.[13] Furthermore, it is desirable to allow the adjustment towards long-run equilibrium to be discrete as well as smooth, since the central bank might encourage market agents to enter market at the same time. If $\gamma \rightarrow 0$, the model becomes

[12] Michale, et al. (1997) use long time series spanning 1791-1992 and also post-war time series for bilateral U.S. real exchange rates. Baum et al. (2001) examine bilateral U.S dollar CPI and WPI proxies over the post Bretton -Wood period. The data set in Taylor, et al. (2001) comprises several bilateral real exchange rates against U.S. dollar during the post-Bretton-Wood period.

[13] Sollis et al. (2002) show that there is stronger mean reversion when foreign currency is overvalued against the U.S dollar than when it is undervalued by the same proportionate amount.

linear, whereas if $\gamma \to \infty$ and $c_1 \neq c_2$, the function $G(\gamma, c, z_t)$ is equal to 1 for $z_t < c_1$ and $z_t > c_2$, and equal to 0 in between. Thus, LSTAR2 model nest a SETAR model as a special case.

4.3. THE MODEL

Presumably the weight put on the real exchange is low in the monetary policy rules followed by the Bundesbank (the ECB) and the FED. Thus, large uncertainty regarding expected shocks implies the real exchange rate follows random walk process close to PPP. We use a simple AR(1) structure to mimic this property. Our theoretical model omits the direct effect of the present value of monetary policy shocks on the real exchange rate. Monetary policy shocks may have, however, an independent and direct effect on the real exchange rate.[14] Thus, we also include a short term real interest rate differential in the model. ϕ_t is a temporary unobserved shock to real exchange rate, after which the real exchange crosses over to implicit boundaries. We cannot say much about the contribution to ϕ_t, since it is not observed by us. Our nonlinear model for the real exchange rate is basically a two regime model:

$$q_t = \begin{cases} \alpha_{1,2} q_{t-1} + \alpha_{2,2}(r_{t,s} - r^*_{t,s}) + \alpha_{3,2}\phi \;\; if \;\; q > c \\ \alpha_{1,1} q_{t-1} + \alpha_{2,1}(r_{t,s} - r^*_{t,s}) + \alpha_{3,1}\phi \;\; if \;\; |q| \le c \\ \alpha_{1,2} q_{t-1} + \alpha_{2,2}(r_{t,s} - r^*_{t,s}) + \alpha_{3,2}\phi \;\; if \;\; q < c \end{cases} \tag{4.4}$$

q_t is centered on zero and corresponds to the deviation of the nominal exchange rate from the underlying PPP fundamental. The hypothesis are as follows: $\alpha_{1,2} < \alpha_{1,1}$ (adjustment towards parity), $\alpha_{2,1}, \alpha_{2,2} > 0$ (real interest rates), $\alpha_{3,1} = 0$ and $\alpha_{3,2} > 0$ (temporary exchange rate component). At the level c a stochastic trend in q_t is broken.

5. DATA

The sample consists of monthly observations from the January 1975 to the June 2003. A multivariate model is estimated using a shorter time period 1982:10-2003:6. The source of the data has been the OECD's Main Economic Indicators. Interest rates set of variables includes the German (euro zone) long 10 year bond yield, the US long 10 year bond yield, the German (euro zone) 3

[14] The share of exchange rate variance due to monetary policy shocks is found to be between near zero to over half. (Clarida and Gali, 1994, Eichenbaum and Evans, 1995, Kim and Roubini,2000, and Faust and Rogers, 2003).

month treasury bill rate and the US 3 month treasury bill rate. The exchange rate set of variables is defined by

$ppp_t = p_t - p_t^* - s_t$, where

p_t = the German (Euro), or 'home', consumer price index,

p_t^* = the US, or 'foreign', consumer price index,

s_t^* = the spot exchange rate, defined as $/DM(euro),

All the variables except interest rates are in logarithmic forms. Two alternative measures of expected inflation are considered. The first alternative is a eighteen-months (18 lags and 18 leads) centered moving-average of CPI inflation rates (a real short-term interest rate R18). The another measure is based on twelve months changes in the CPI index (a real short-term interest rate Rrat).

A temporary exchange rate premium, ϕ_t in (4.4), is labeled as *temp*. To control for extraordinary shocks in term premium, the following dummy specification is used over the period 1982:10-2003:6

Temp = No dummies

Tempg = German reunification dummy 19xx:6-199x:x

Tempgt = Tempg + trend

Tempget = Tempg + Euro trend dummy 1999:1-2003:6.

German reunification dummy implies the sharp increase in short term interest rates during the time period of German reunification. Euro trend dummy is included since we allow for the possibility that it took time before the European Central Bank establish its credibility.

The data is represented as line graphs in Appendix 1. The persistence of the nominal exchange rate deviations from the PPP equilibrium is apparent in Figure A.1 It appears that the persistent does not arises as the results of a single period, such as the large swings in the mid of 80's. Figure A.2 show our temporary exchange rate estimates. There is a quite remarkable co-movement between the large deviations from PPP and our temporary exchange rate estimates especially in middle of eighties and again in the middle of nineties[15]. Finally, Figure A. 3 demonstrates the large variation in real short term interest rate differential in the beginning of nineties. Most of time the real short term interest rate differential is relatively stable variable.

6. TRENDING PROPERTIES OF THE DATA

One difficulty often presented in empirical analysis of economic time series is the determination of the order of integratedness of a series. Theoretically we can classify variables exhibiting a high degree of time persistence (insignificant mean-reversion) as nonstationary I(1) variables and variables exhibiting a significant tendency to mean reversion as stationary I(0). In practice many variables are borderline cases such that distinguishing between a strongly autoregressive I(0) or I(1) process is far from easy.

As a preliminary exercise, we use augmented Dickey-Fuller (thereafter ADF) to determine the degree of integration of the time series. The lag length of the ADF unit root test has been chosen using the sequential rule suggested by Hall (1994). This is shown to be the most efficient way to define the lag length of ADF test.[16] Because the frequency of our data is monthly, the testing was started with 12 lags. The results in Table 1 and 2 are only weakly supportive for the hypothesis that the real exchange rate, the real interest rate differentials and the temporary exchange rate premiums are stationary variables. However, rejection of the null hypothesis of a unit root at the 10 % level of significance is found for all variables.[17] The real exchange rate variable is a stationary variable only at the 10% significance level without drift over the shorter time period 1982:10 –2003:6. Putting the real exchange series into first-difference form did appear to induce stationarity.[18]

[15] Findings in Meridith and Ma (200x) are consistent with our time-varying exchange rate uncertainty estimates…
[16] See the discussion, for example, in Ng and Perron (1995).
[17] R18 estimate is based on eighteen-months centered moving-average of CPI inflation rates. Important MA components in the structure of the series implies that a number of nuisance parameters is needed in the estimation. Since we furthermore lose one effective observation for each extra lag introduced, the Dickey-Fuller approach may have lower power when MA terms exists than if the errors were i.i.d.

Table 1. Trending properties of PPP variable

75:1-2003:6	obs:	349		82:10-2003:6	obs:	249	
PPP				PPP			
ADF(t),10	sigma			ADF(t),1	sigma		
drift	0,31	-2,45	*	drift	0,32	-1,82	
no drift	0,31	-2,21	**	no drift	0,32	-1,78	*
ADF(t),1				ADF(t),1,dif			
drift	0,30	-2,05		drift	-0,73	-11,45	***
no drift	0,30	-1,98	**	no drift	-0,75	-10,03	***
Eklund(F)				Eklund(F)			
drift	0,37	1,37		drift	0,36	1,35	
no drift	0,37	1,15		no drift	0,36	1,14	
* = rejection of the null hypothesis at the 10 % level of significance							
**= 5%							
***=1%							

Table 2. Trending properties of explanatory variables.

82.10-2003.6	obs:	249					
R18				Rrat			
ADF(t),5	sigma			ADF(t),1	sigma		
drift	0,41	-1,99	*	drift	0,3	-2,12	*
no drift	0,41	-2,11	**	no drift	-0,02	-2,11	**
Eklund(F)				Eklund(F)			
drift	0,49	2,52	*	drift	0,43	2,48	*
no drift	0,49	2,20	*	no drift	0,43	2,17	*
Temp				Tempgt			
ADF(t),1	sigma			ADF(t),1	sigma		
drift	-0,05	-2,56	*	drift	-0,09	-3,23	**
no drift	-0,05	-2,56	**	no drift	-0,07	-3,24	**
Eklund(F)				Eklund(F)			
drift	-0,2	2,05		drift	-0,19	2,58	*
no drift	-0,2	1,71		no drift	-0,19	2,15	*
Tempg				Tempget			
ADF(t),1	sigma			ADF(t),1	sigma		
drift	-0,08	-2,57	*	drift	-0,05	-2,65	*
no drift	-0,05	-2,54	**	no drift	-0,06	-2,65	***
Eklund(F)				Eklund(F)			
drift	-0,18	2,02		drift	-0,19	2,27	
no drift	-0,18	1,68		no drift	-0,19	1,89	
* = rejection of the null hypothesis at the 10 % level of significance							
**= 5%							
***=1%							

[18] The results did not change if a linear trend was included in the Dickey-Fuller regression.

We use the Eklund-F test for the joint testing of linearity and unit root hypothesis against the second-order logistic STAR (Eklund,2003). The artificial regression is as follows:

$$y_t = \delta_1 \Delta y_{t-1} + \delta_2 (\Delta y_{t-1})^2 + \delta_3 (\Delta y_{t-1})^3 + \phi_1\ y_{t-1}\Delta y_{t-1} + \phi_2 y_{t-1}(\Delta y_{t-1})^2 + \alpha + \rho y_{t-1} + \varepsilon_t \tag{6.1}$$

A joint test of linearity and the unit root hypothesis amounts to testing the hypothesis $H_{01} : \delta_2 = \delta_3 = \phi_1 = \phi_2 = \alpha = 0, \rho = 1$ in (6.1). The null hypothesis is that the true data generating process is a random walk. Excluding $\alpha = 0$ from H_{01} results in another null hypothesis H_{02} that allows for a unit root process with drift component. Since under the null hypothesis the real exchange rate follows a unit root process, the null hypothesis complicates the testing procedure analogously to the way in which the distribution of a Dickey-Fuller statistic cannot be assumed to be Students's t. Monte Carlo study of the critical values, size and power properties of the EF is provided by Eklund (2003).[19]

The rejection of the null hypothesis provides evidence of stationary but nonlinear time series.[20] The results in Table 1 and 2 indicate that the random walk without drift can be rejected for both real interest rate estimates and one out of the four temporary exchange rate premiums. The results do not give any support to the conclusion that the real exchange rate follows a stationary and nonlinear process.

The ADF tests lack power against stationary PPP alternative over the post-Bretton-Wood time period. Taylor *et al.* (2001) pointed out that the unit root behavior does not necessarily imply that no long-run equilibrium exists. The failure to reject a unit root may indicate, conversely, that the most of time real exchange rates have been in a neighborhood of the long-run equilibrium level. This is because real exchange rates behave more like unit root processes the closer they are to long-run equilibrium. We call this feature of real exchange rate time series as local unit root behavior.

Although power simulations in Eklund (2003) show some gain in power compared to the standard ADF test, also EF has low power discriminating a random walk from theoretically meaningful stationary and nonlinear PPP alternative at sample size available for the tests. Thus, we build an

[19] According to Eklund (2003) the size of the test is distorted when the value of δ_1 is close to –1 or 1. The size of the ADF test is distorted only when δ_1 is close to 1.
[20] We use critical values for 250 observations.

univariate model for the U.S. dollar- German mark (euro) real exchange rate to find out whether local unit root behavior dominates during the time periods 1975:1-2003:6 and/or 1982:10-2003:6.

6.1 UNIVARIATE MODEL FOR REAL EXCHANGE RATE

We begin our analysis by assuming that the real exchange rate series is stationary. We also assume linearity, but we prepare to consider the possibility that the real exchange rate cannot be adequately characterized by a linear autoregressive model. Our alternative model to the linear model is the smooth autoregressive model (STAR)

$$q_t = \phi' x_t + \theta' x_t G(\gamma, c, z_t) + \mu_t \,, \tag{6.2}$$

$x_t = (1, q_{t-1}, ..., q_{t-p})'$, $\phi = (\phi_0, \phi_1, ..., \phi_p)'$, $\theta = (\theta_0, \theta_1, ..., \theta_p)'$, and $\mu_t \sim nid(0, \sigma^2)$. G is a transition function.

The first step is to specify a linear AR(p) model for the real exchange rate series to serve as our null hypothesis.[21] The order of autoregression, p=2, is chosen through inspection of partial autocorrelation function.[22] Table 3 presents the results of testing for nonlinearity using the third order artificial regression suggested by Luukkonen (1998). The delay length, d, is varied in order to provide the strongest probability of non-linearity.[23]

Table 3. Linearity against non-linearity. Full sample.

	p-Values of Tests of Linearity against STR , Transition Variable Assumed Known.											
	LAG											
PPP	1	2	3	4	5	6	7	8	9	10	11	12
F	0,59	0,67	0,60	0,43	0,31	0,16	0,2	0,11	0,02	0,03	0,09	0,09
F(4)									0,05	0,04		
F(3)									0,01	0,02		
F(2)									0,91	0,90		

[21] All the pre-test results presented here are for full sample. The results for shorter sample (1982:10-2003:6) are qualitatively alike.
[22] We prefer partial autocorrelation function since the use of information criteria may bias our estimates toward low values and this may lead to a false rejection of linearity as discussed in Teräsvirta (1994).
[23] We use ordinary F-test, since, as found out by Teräsvirta and Granger (1993), an F-approximation works much better in the small sample size than LM test with the asymptotic χ^2 distribution.

The null of linearity is strongly rejected when p= 2 and d = 9 or 10. We use both d-values to find out whether our results are sensitive to this choice. The choice between transition functions is based on test sequence suggested by Teräsvirta (1994).[24] This supports either a LSTAR2 or an ESTAR function. We will use a LSTAR2 model. The final choice between them is done based on a discussion in the Chapter 4.[25]

We subsequently found that the simplifying restriction $\phi_2 = 1 - \phi_1$ could not be rejected at standard significance levels for any of the estimates. Furthermore, a nonlinear part of autoregressive parameters provide always negative values. This gives the model a local unit root when G = 0. Table 4 shows the estimated results.

Table 4. Nonlinear model for real exchange rate. Bold values correspond to a significant value at the 5% level.

Full		drift	Gamma	Low	High	ϕ (1)	θ (N)	Share
PPP_9	Drift	**-0,01**	**51,20**	**-0,27**	**0,23**	**1,34**	**-0,09**	18 %
	No drift	-	-	-	-	-	-	
PPP_10	Drift	**-0,01**	**21,40**	**-0,26**	**0,20**	**1,34**	**-0,07**	23 %
	No drift	-	-	-	-	-	-	
82:10		drift	Gamma	Low	High	ϕ (1)	θ (N)	
PPP_9	Drift	0,00	-	-	-	-	-	
	No drift	-	1,98	**-0,33**	**0,24**	**1,32**	**-0,11**	8 %
PPP_10	Drift	0,00	-	-	-	-	-	
	No drift	-	**90,60**	**-0,25**	**0,20**	**1,34**	**-0,08**	20 %

A natural question is whether the rejection of the linear model is perhaps driven by a single episode in the data. Probably this is not the case here since there are quite many observations outside the band.[26] Gamma is adjusted by $\sigma^2(q)$ which is the sample variance of q_t.[27] Since $\gamma \to \infty$ then $G(\gamma, c, z_t) \to 0$ for $C1 \le z_t \le C2$; and for other values $G(\gamma, c, z_t) \to 1$.

The real exchange rate time series seems to follow local unit root behavior when the transition variable is not large. Most of the observations are inside the band. The failure to reject a unit root

[24] The p-values for the whole sequence of test are given only if the general linearity test (F) lies below 0.05.
[25] Using ESTAR model, results indicate…
[26] The figures in the last column are the percentages of observations below c1 and above c2 respectively.

seems to indicate that the real exchange rate has been most of the time relatively close to equilibrium, rather than the real exchange rate is not globally stationary process. Thus, our results indicate the presence of nonlinear adjustment towards parity level.

Although we do not explicitly prove our stationary assumption, the empirical results shown in Table 4 reveal under the assumption of stationarity that the estimated transition parameters are strongly significantly different from zero, which itself indicates a correct specification. Combining this evidence with the strong theoretical and empirical evidence for nonlinear real exchange rates, makes a stationarity assumption as an interesting alternative to a non-stationarity assumption based on linear unit root tests.[28] Our model is very close to switching regression model.[29] This contrasts with the results that one would expect by considering the ESTAR transition function, as several authors have assumed. Notably, this character might imply that the central bank(s) has a role in determining exchange rates as discussed in Chapter 4.

The differences of positive boundary values between models based on PPP_9 or PPP_10 can be explained by the sharp dollar appreciation in the middle of 80's. Some of the estimated bands are also found to be asymmetric.[30] This is because the real exchange rate wander outside the quasi-symmetric band longer time in the middle of 90's than in the middle of 80's. We also find that the order of delay parameter is nine or ten. Although large d values are typical findings also in ESTAR models, they are problematic if we do not have any clear economic intuition which supports large d values.[31] In the next Chapter we introduce temporary nominal exchange rate shocks as an explanation for asymmetry and large d-values.

[27] As discussed in Teräsvirta (1998) Gamma is not a scale-free parameter as its value depends on the magnitude of the values of transition variable z_t.

[28] Taylor et al.(2001) generates real exchange rates time series using a nonlinearly mean-reverting ESTAR model. They show that there is 70 percent probability that their data set (288 observations) would never allow them to reject the unit root hypothesis using standard univariate unit root tests.

[29] See similar finding in Bec *et al.* (2002). They use a three-regime LSTAR model with the symmetry restriction.

[30] Note that we included a drift term to capture difficulties in identifying the long run equilibrium level of the real exchange rate.

[31] Taylor et al.(2001) is an exception (d =1). As an example for large d value, see Baum et al. (2001).

7. MULTIVARIATE MODEL FOR REAL EXCHANGE RATE

We use a shorter sample for a multivariate model.[32] Multivariate model begins at 1982:10, the date of possible regime shift in U.S. monetary policy.[33] We estimate the model in (4.4). However, only the inclusion of PPP_2 was able to "clean" the residual.

The linearity test results are shown in Table 6 bellow. The null of linearity is strongly rejected when PPP_2 is used as a transition variable.[34] The other strong rejection occurs when PPP_9 or PPP_10 are assumed to be the transition variable. Since the PPP_2 provides us more accurate parameter estimates in a following analysis, it is assumed that the data has generated by such a LSTR2 model.[35]

Table 6. Linearity against non-linearity.

	p-Values of Tests of Linearity against STR		
	LAG		
PPP	2	9	10
F	0,02	0,01	0,02
F(4)	0,95	0,1	0,1
F(3)	0,01	0,01	0,01
F(2)	0,07	0,54	0,51

7.1. RESULTS

Assuming the linearity hypothesis is rejected, we can estimate the LSTR2 model for the real exchange rate series. After removing the redundant variables the final results are given in Table 7 bellow.

[32] The advent of euro did not necessarily change drastically the weight put on the bilateral dollar-mark (euro) in monetary policy rules. This is because, by definition, the weight of this bilateral exchange rate has increased in monetary policy rules and also because stable EMU currencies before the advent of euro.
[33] From 1979 to 1982, the Federal Reserve (FED) targeted non-borrowed reserves. Thereafter, the FED has been followed a monetary policy rule, which practically amounts to an interest rate targeting policy.
[34] Lagged values of PPP from 1 to 12 were again a priori regarded as potential transition variables.
[35] PPP_9 and PPP_10 estimates are similar to univariate model but not always statistically significant.

Table 7. Non-linear regression 1982:10 – 2003:6

Variable name	coeff. estimate	std.deviation	t-value
GAMMA	22,000	1,03	21,30
C1	-0,193	0,01	-19,40
C2	0,204	0,02	10,10
Constant	-0,003	0,00	-1,59
PPP_1	1,270	0,06	20,70
PPP_2	-0,264	0,06	-4,22
R18	0,002	0,00	2,76
Nonlinear parameters: PPP_2			
PPP_1	-0,104	0,03	-3,56
Temp	0,015	0,01	2,84

$R^2 = 0.98$, $\hat{\sigma} = 0{,}02$, $\hat{\sigma}/\hat{\sigma}_l = 0.96$, sk = 0.36, k =3.25

Normality test: Chi^2(2) = 5.11 [0.08]

$\hat{\sigma}$ is a residual and $\hat{\sigma}_L$ denotes the residual standard deviation of the corresponding linear model. The residual standard deviation of the nonlinear model is about 96 per cent of that of the corresponding linear model. This high correlation results because most of the time the real exchange rate fluctuations fall within the discrete borders. The residual from above regression is graphed in Figure 1 together with the residual from the linear model.

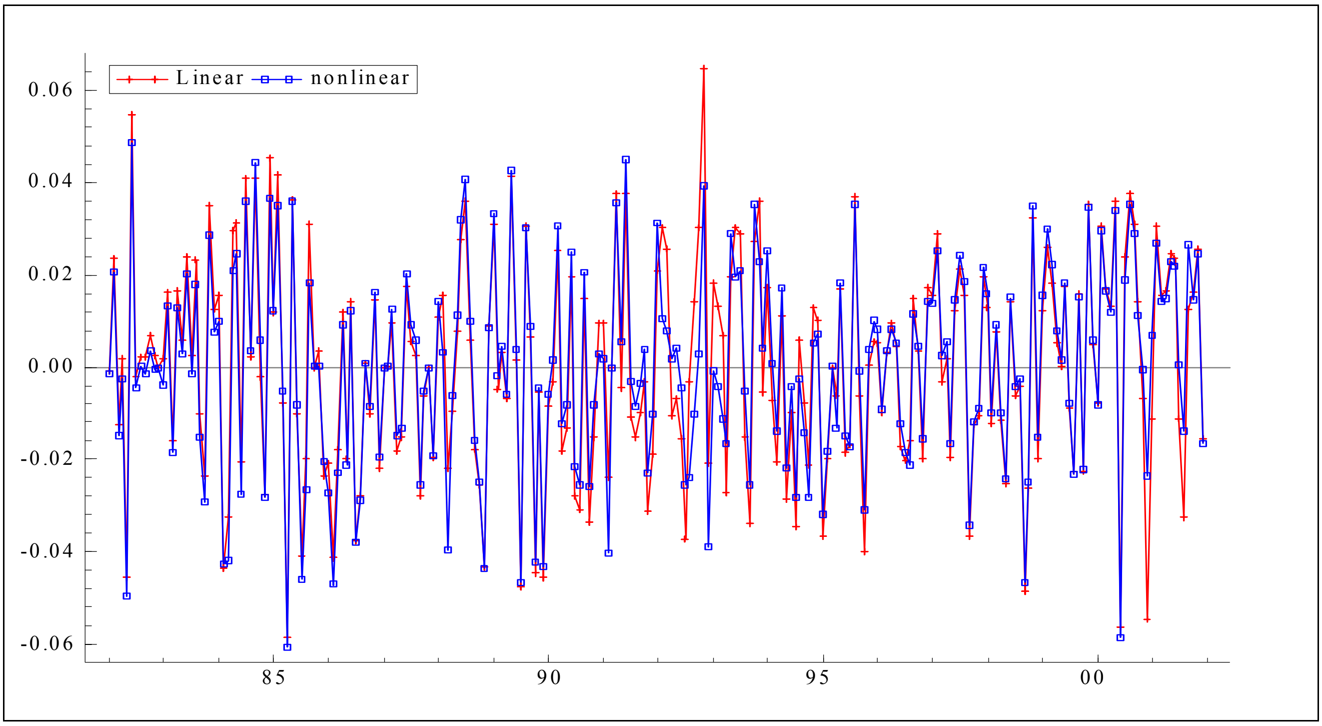

Figure 1 Residuals from linear model and from the STR model

In our view, our model captures, however, two essential features of the data: nonlinear deviations from PPP, and a discrete and symmetric band around PPP. The main difference in the fits of the two models is due to different characterization of the large fluctuations. Figure 2 shows that the transition function obtains value one especially during the dollar appreciation period in the eighties and again during the dollar depreciation period in the middle of nineties. The band around PPP is now almost symmetric +/- 0,2 and Gamma parameter is large implying a discrete regime change (see Figures 2 and 3 bellow)

We cannot impose a temporary exchange rate premium in the linear part of the model but it enters the model non-linearly. This is consistent with the discussion in Chapter 3. The model derived also shows that most of the movements in the real exchange rate is accounted for by its own past. However, the real short-term interest rate differential is also statistically significant in explaining movements in the real exchange rate. The adjustment towards parity level is relatively rapid if the boundaries are broken.[36] Note also that the order of delay parameter is now only two implying that adjustment towards parity begins almost immediately after the boundaries are broken.

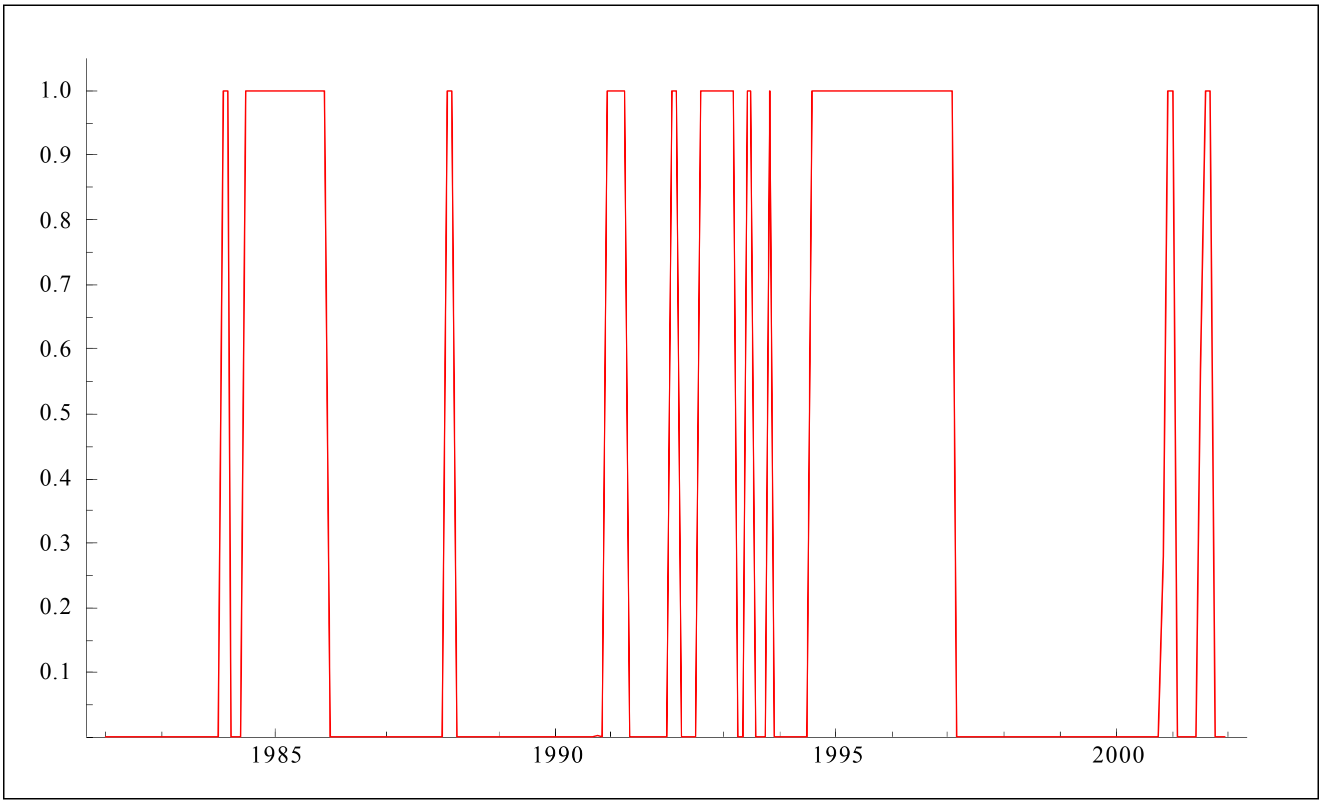

Figure 2. Values of the transition function.

[36] Parameter estimates are, by definition, inaccurate, since Temp variable mimics an unobserved shock.

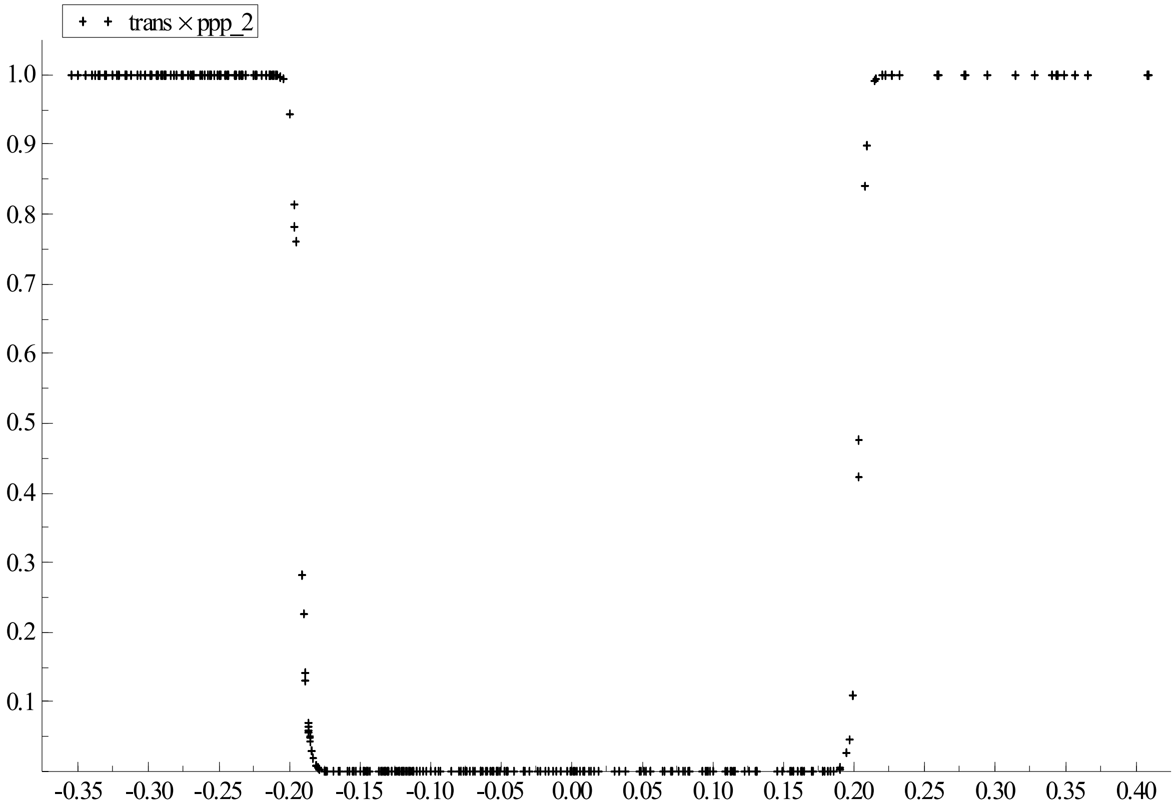

Figure 3 The value of transition variable (PPP_2) and transition function $G(\gamma, c, z_t)$

Residual diagnostic tests are reported in Table 8 bellow.

Table 8. Residual diagnostic tests.

	P-values of the LM test of no Error Autocorrelation against AR(p) and MA(p) Error Process, and the LM- test of no Autoregressive Condotional Heteroskedasticity against ARCH(p) in (x). Maximum Lag p					
Test	1	2	3	4	5	6
No error autocorrelation	0,37	0,39	0,18	0,31	0,29	0,38
No ARCH	0,80	0,86	0,95	0,92	0,98	0,99

Tests of no error autocorrelation are Lagrange multiplier test statistics for the first order up to sixth-order serial correlation in the residual, as discussed in Eithreim and Teräsvirta (1996). A traditional ARCH LM test is used for autoregressive conditional heteroskedasticity. Results of the LM test of no error autocorrelation do not indicate autocorrelation nor is there any evidence of ARCH. Thus, the LSTR2 model appears to provide an acceptable representation for the adjustment process towards PPP. Finally, a real exchange rate variable with the band is graphed in Figure 4 bellow.

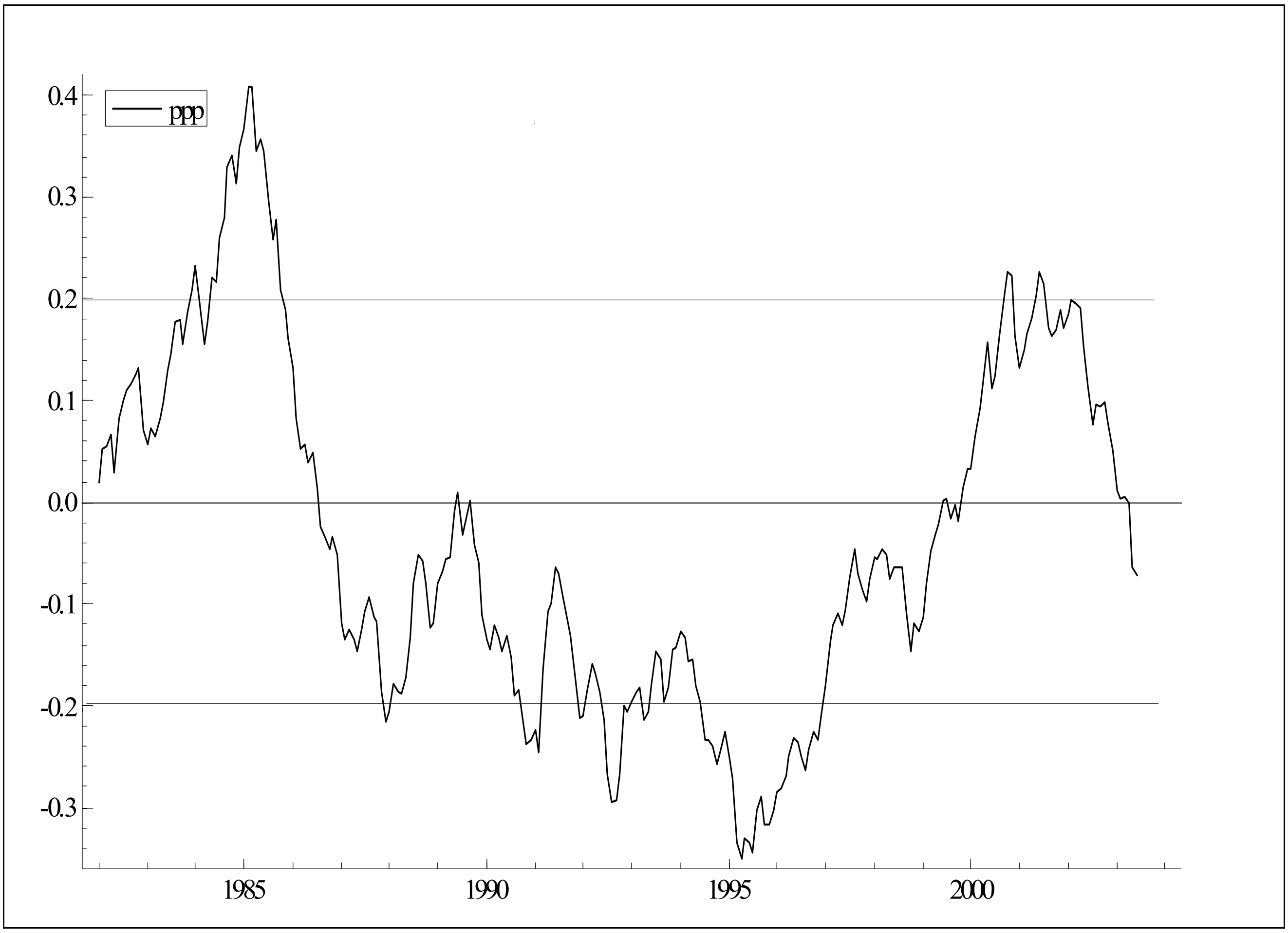

Figure 4 Real exchange rate time series and the boundaries of a band.

It is also crucial to find out whether the relationship between variables is dependent on the data measures we use. Table 9 presents a summary of the results based on alternative data measures. The results do not vary considerably with the different data measures. The values of Gamma variable vary relative much, but the shape of transition function does not change substantially, since the values of Gamma variable are large enough to provide a switching regression model. The exception is the result based on the specification including a trend in a Temp variable.

Table 9. Alternative data measures.

Gamma	20,20	1,46	0,36	28,28	1,71	0,36	30,80
High	0,20	0,20	0,16	0,20	0,20	0,18	0,20
Low	-0,19	-0,19	-0,15	-0,19	-0,19	-0,15	-0,19
	Rrat Temp	Rrat Tempg	Rrat Tempgt	Rrat Tempget	R18 Tempg	R18 Tempgt	R18 Tempget

7.2. DISCUSSION

The band around PPP is wide. This implies that the weight put on the real exchange rate is quite low.[37] Our model is also very close to switching regression model. This might indicate that the near the lower and upper boundaries the expectations on future monetary reactions are rapidly increased, i.e. traders have homogeneous expectations near the boundaries. This is in all probability due to official announcements by the central bank(s).

It is possible that the U.S. monetary policy rule may not have been closely following a Taylor rule in the beginning of our sample.[38] However, the interest rate differential rose in response to high U.S inflation and the dollar appreciated strongly in 1983. A large further appreciation in 1984 and early 1985 have frequently interpreted as a temporary speculative bubble.[39] This might imply a lack of policy co-ordination in the middle of eighties. However, the sharp appreciation of the US dollar in the first half of the eighties led the Federal Reserve and Treasury to commit itself to policy coordination as expressed in the Plaza Accord [40] and the Louvre Accord[41].

After the Louvre Accord the real exchange rate has been substantially crossed the informal and flexible reference ranges in the 7 years: 1990, 1992, 1994, 1995, 1996, 2000, and 2001. In three of the seven years in which the reference ranges was crossed, it took more than two months before an adjustment commenced: 1992 and 1994-1995. In 1990 and especially in 1992, the monetary policy was not consistent with moving exchange rate towards PPP (see Data Figures in Appendix). Thus, the German reunification and events related to the first years of European Exchange Rate Mechanism (ERM) crises in Germany and the conflict between the needs of domestic target variables and the bilateral dollar-mark exchange rate in the United States together induced forces

[37]The bilateral U.S. dollar -German mark (euro) real exchange rate is probably more important variable in the Bundesbank (the European Central Bank) monetary rule than in the Federal Reserve monetary rule. Clarida et al. (1998) estimate linear monetary policy reaction functions for Germany and Japan. They find that a one percent depreciation the mark relative to the dollar led Bundesbank to increase interest rates by five basis points. It has been very difficult to find such effect for the United States.

[38] Mäki-Fränti (2003) examined the parameter values of the Taylor rule and found the regime shift not until the mid-eighties, i.e. the Fed has reacted to the increased inflation by increasing also real interest rate not until the mid-eighties.

[39] Juselius and MacDonald (2002), for example, pointed out that interest rate differential is not a cause of dollar appreciation in middle of eighties.

[40] On September 22nd 1985, finance ministers from the United States, Japan, Great Britain, West Germany and France announced the Plaza Accord at the eponymous New York hotel. The plan turned out to be successful. By the end of 1987, the dollar had fallen by 54% against mark from its peak in February 1985.

[41] The Louvre Accord was hatched in 1987 to stabilize the dollar. Finally, heavy foreign exchange market interventions did stabilize the dollar again after sharp depreciation, although the dollar continued to fall immediately after the Louvre Accord.

that temporarily appreciated German mark more than average.[42] There was several exchange rate market intervention operations to move the dollar-mark exchange rate towards PPP, but they were not successful until 1993.[43]

The prolonged nature of the dollar depreciation in 1994 and early 1995 would seem to be unwarranted solely in terms of interest rate differentials.[44] Note also that during 1994-1996 the US economy grew at twice the rate observed in Germany. This did not prevent the dollar from depreciating vis-a-vis the German mark. The fiscal problems in several ERM member states increased the probability that these countries were not to qualify for the initial group of states forming a single currency. This increases the pressure within ERM which was intensified by the prospect of higher than expected wage settlement in Germany at the end of 1994. The downward pressure on the dollar against the mark was accentuated by expectations on less liquid cross exchange rates between the European currencies and the German mark. In the subsequent period the Mexican financial crisis caused many overseas investors to develop aversion to all North American assets, including dollar denominated assets. These two sources of uncertainty in an international financial system depreciated the dollar against the German mark although monetary policy were consistent with moving the exchange rate towards parity level. Finally, the stance of monetary policy and heavy interventions together broke the trend in 1995 and the dollar began steadily to appreciate.[45]

In 2000, the monetary policy stance did not support adjustment toward parity. Interventions and more supportive monetary policy in 2001 and 2002 bucked the trend. The dollar –mark exchange rate stabilized in a reference range, i.e. there were not substantial fluctuation outside the reference range. This is consistent with the fact that there were no random shocks in the right direction in an international financial system. It seems to be the case that only under of influence of shocks that are related to serious crisis in an international financial system the real exchange rate will cross the reference ranges.

[42] Note that our temporary exchange rate variable does not imply large adjustment expectations, although the real exchange rate fluctuates outside reference ranges over the half year. One possible explanation is that reunification offsets the adjustment expectations towards historical PPP.
[43] Here and throughout the source is a report: " Treasury and Federal Reserve Foreign Exchange Rate Operations" provided by Federal Reserve Bank of New York.
[44] Our temporary exchange rate variable does imply large adjustment expectations.
[45] On March 3, Treasury Secretary Rubin stated: " A strong dollar is in our national interest. That is why we have acted in the markets in concert with others."

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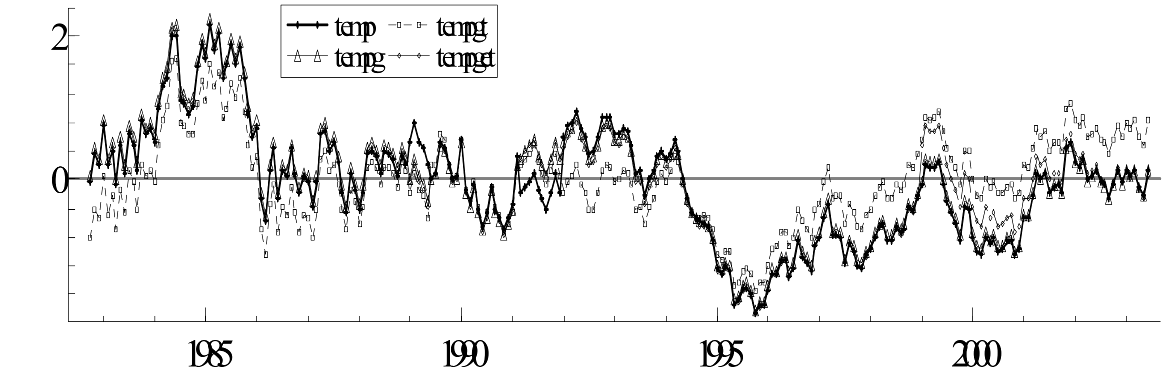
APPENDIX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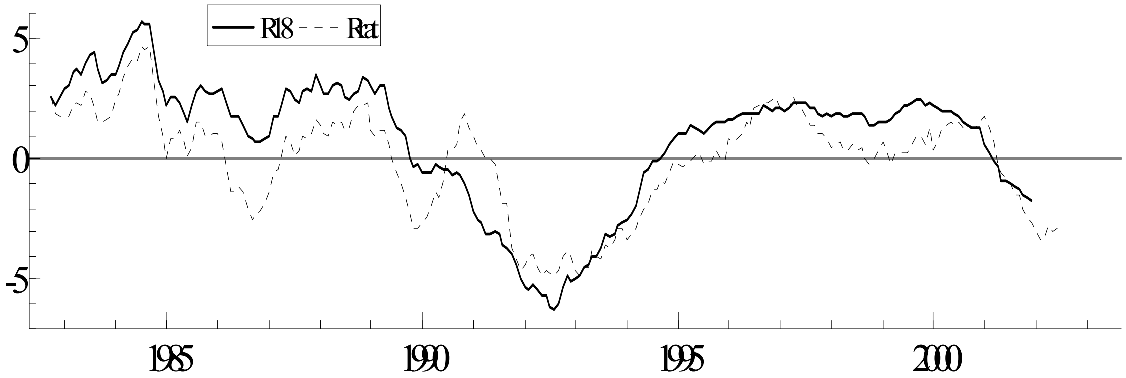
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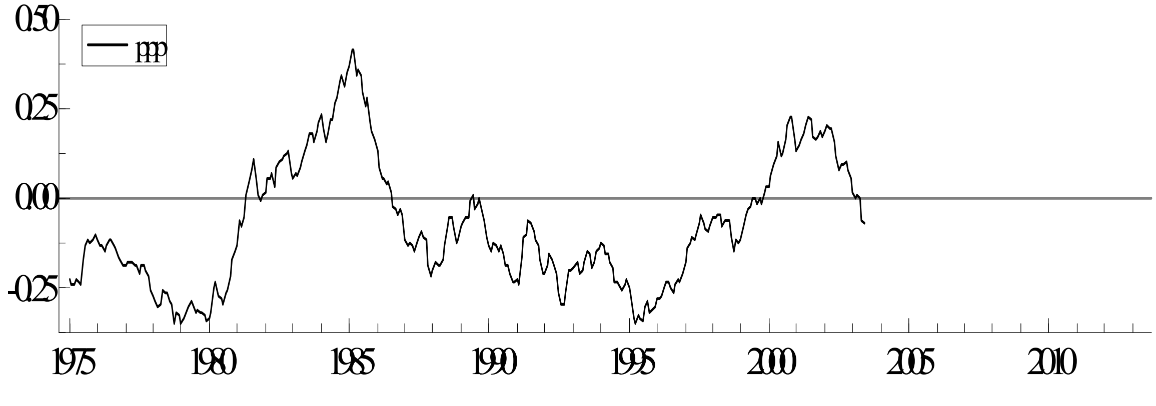
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