Non-linear and non-symmetric exchange-rate adjustment:

New evidence from medium- and high-inflation countries

by

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Abstract

This paper analyses a model of non-linear exchange rate adjustment that extends the literature by allowing asymmetric responses to over- and under-valuations. Applying the model to Greece and Turkey, we find that adjustment is asymmetric and that exchange rates depend on the sign as well as the magnitude of deviations, being more responsive to over-valuations than under-valuations. Our findings support and extend the argument that non-linear models of exchange rate adjustment can help to overcome anomalies in exchange rate behaviour. They also suggest that exchange rate adjustment is non-linear in economies where fundamentals models work well.

Keywords: exchange rates, non-linearity, non-symmetry, PPP, Flexible-price Monetary Model (FPMM)

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1. INTRODUCTION

Whether theoretical monetary models of the exchange rate can explain past and predict future movements of actual exchange rates continues to remain the subject of intense debate. Since Meese and Rogoff (1983) reported that a number of fundamentals-based monetary models were outperformed by a simple random walk at horizons of up to one year, a large number of researchers have examined the empirical validity of the purchasing power parity (PPP) hypothesis and the flexible-price monetary model (FPMM). In its traditional form, PPP is the proposition that the equilibrium exchange rate equalises price levels across countries when measured in terms of a common currency, thus implying a constant real exchange rate. The FPMM models the exchange rate as the relative price of different currencies and so the relative supply of and demand for currencies is equalised at the equilibrium exchange rate.

A large empirical literature (summarised by Froot and Rogoff, 1995, Rogoff, 1996, Taylor, 1995, and Taylor and Sarno, 1998) has found that the long-run relationships implied by the PPP and FPMM models are not cointegrated (other than in periods of very high inflation, as argued by McNown and Wallace, 1989). This is what Taylor, Peel and Sarno (2001) refer to as the first puzzle in exchange rate behaviour[1]. The literature has also found that the short-run adjustment of the exchange rate to changes in fundamentals is surprisingly slow, with estimated half-lives of reversion to equilibrium ranging between three to five years. These implausibly low estimates of the speed of adjustment seem to be caused by high levels of volatility in exchange rates compared to underlying fundamentals. The failure of traditional theory to explain this disparity is what Taylor, Peel and Sarno (2001) define as the second puzzle in exchange

[1] Theoretical explanations for these negative findings include Balassa-Samuelson effects (Kravis and Lipsey, 1983; Bhagwati, 1984), wealth effects caused by accumulated current account imbalances (Krugman, 1990) and the distortionary effects of government spending (Froot and Rogoff, 1991, Rogoff, 1992,1996).

rate behaviour. Taken together, these puzzles cast doubt on our most familiar models of the exchange rate.

A number of authors have sought to resolve these puzzles by developing theoretical models with non-linear adjustment of the exchange rate, including Dixit (1989); Dumas (1992); Uppal (1993); Coleman (1995); Sercu et al. (1995); Shleifer and Vishny (1997); and O'Connell (1998). These models assume limits to arbitrage, through spatially separated markets with transaction costs or sunk costs. This leads to a non-linear model in which exchange rates only respond to larger movements in fundamentals, since the marginal cost of arbitraging differences between the exchange rate and fundamentals exceeds the marginal benefit for smaller deviations from fundamentals. These models suggest that the exchange rate puzzles may be a result of incorrectly using a linear framework rather than a symptom of deeper problems with PPP and FPMM models.

A number of empirical studies, including Obstfeld and Taylor (1997), Michael et al. (1997), Taylor and Peel (2000), Taylor, Peel and Sarno (2001) and Baum et al (2001) have found evidence of this[2]. These studies allow for the type of non-linear behaviour predicted by theory by using the threshold autoregressive (TAR) model (Tong, 1990) or the exponential smooth transition autoregressive (ESTAR) model (Granger and Teräsvirta, 1993), in which the behaviour of exchange rates differs between an "inner regime", where exchange rates are close to fundamentals and an "outer regime", where the gap between exchange rates and fundamentals is larger. The speed of adjustment is typically found to be stronger in the outer regime and exchange rates are often found to be a random walk in the inner regime.

These findings have restored some faith in models of exchange rates and opened up interesting new areas of research. However, these models are very restrictive.

[2] An exception is the study of Enders and Falk (1998) who, working within a non-linear unit tests framework, find limited evidence in support of PPP.

Although they allow exchange rate adjustment to vary with the size of the gap between exchange rates and fundamentals, they do not allow for other forms of non-linearity. In particular, they do not allow exchange rates to respond differently to under-valuations and over-valuations. But this type of asymmetry is quite plausible. Consider, for example, exchange rate intervention by a policymaker that assigns greater loss to employment being below the socially desirable level than to employment being too high (such a model has been analysed in a closed economy context by Cukierman and Gerlach, 2003). Such a policymaker may well be more responsive to exchange rate over-valuations than to under-valuations.

In this paper, we investigate asymmetric exchange rate adjustment. We use the quadratic logistic Smooth Transition Error Correction Model (QL-STECM), (see van Dijk et al., 2002). This is similar to the TAR and ESTAR models in that it allows the response of exchange rates to depend on the size of the deviation from fundamentals. However it goes beyond these models since it also allows the response of exchange rates to depend on the sign of the deviation from fundamentals and thus allows for different responses to under-valuations and over-valuations. This model therefore allows us to assess the importance of asymmetry in exchange rates.

We investigate the nominal exchange rates between the Greek Drachma and the Turkish Lira against the ECU (the Euro since 1999). We do this for two reasons. First, empirical research on non-linear exchange rate behaviour has focused almost exclusively on the G7 economies and has neglected moderate- and high-inflation economies. This may reflect the view that monetary models are "extremely useful in explaining exchange rates across countries with significant...inflation" (Rogoff 1999). Evidence of non-linear adjustment in moderate and high inflation environments, for which Greece and Turkey are representative examples, would suggest that non-linearity is pervasive. Second, it has been argued that the main focus of macroeconomic policy in both countries has been to attempt to maintain high levels of output[3]. Policymakers may therefore have exhibited the sort of asymmetric preferences that may lead to asymmetric exchange rate adjustment.

Our main finding is that exchange rates adjustment is asymmetric. We find, in common with other studies, that exchange rates respond more vigorously to larger deviations from fundamentals. But we also find that exchange rates are more responsive to over-valuations than to under-valuations.

The remainder of the paper is structured as follows: Section 2 discusses our methodology. Section 3 consists of our econometric analysis. In particular, Section 3.1 discusses the data. Section 3.2 presents the results of linear cointegration tests for the validity of PPP and the FPMM. Section 3.3 reports estimates of linear models of short-run exchange rate adjustment. Section 3.4 presents the results of non-linearity tests for exchange rate behaviour. Section 3.5 reports estimates of non-linear models of short-run exchange rate adjustment. Finally, Section 4 summarises and offers some concluding remarks.

2. METHODOLOGY

The standard linear model of exchange rate adjustment is

$$\Delta s_t = \beta(L)\,\Delta s_{t-1} + \gamma(L)\,\Delta s^*_t + \delta(s-s^*)_{t-1} + \varepsilon_t \qquad (1)$$

In (1), s is the log of the actual (observed) exchange rate, s^* is the log of the equilibrium exchange rate (the exchange rate consistent with macroeconomic fundamentals), $\beta(L)$ and $\gamma(L)$ are polynomials in the lag operator, L, ε is a white noise error term and Δ is the first difference operator. The mechanism through which the

[3] See, among others, Alogoskoufis (1995) and Kibritcioglu et al (2000) for Greece and Turkey respectively.

actual exchange rate converges to its equilibrium value is the error correction term $(s-s^*)_{t-1}$, which measures exchange rate misalignments. If this is statistically significant, there exists a long-run (cointegrating) relationship between exchange rates and macroeconomic fundamentals.

The equilibrium exchange rate is determined by

$$s^*_t = \pi' z_t \tag{2}$$

where z_t is a (k×1) vector of macroeconomic fundamentals relevant to exchange rate determination. We consider two alternative models of the equilibrium exchange rate. In the case of PPP, we have

$$s^*_t = \alpha + \beta_1 \, p_t + \beta_2 p^*_t \tag{2a}$$

where p_t and p^*_t denote the log of domestic and foreign price level respectively. The absolute form of PPP postulates $\alpha = 0$ and $\beta_1 = -\beta_2 = 1$. Measurement errors in price levels may violate this assumption and result in weak-form PPP, which only requires $\beta_l > 0$ and $\beta_2 < 0$ (see Taylor, 1988). Relative PPP allows for a non-zero constant. For the FPMM we assume (see e.g. Taylor and Peel, 2000).

$$s^*_t = \alpha + \beta_1 \, (m - m^*)_t + \beta_2 \, (y - y^*)_t \tag{2b}$$

where m and m^* respectively denote the log of domestic foreign money supply; y and y^* respectively denote the log of domestic and foreign real output. Money neutrality implies $\beta_1 = 1$ and $\beta_2 < 0$.

Equation (1) can be estimated using two alternative methodologies: The first is to apply cointegration techniques to equation (2), obtain estimates of the vector of parameters π' and proceed to replace s^* in equation (1) with its fitted values $\hat{s}^*$ obtained from (2). In that case, (1) becomes (3) below where $\hat{s}^* = \hat{\pi}' z_t$:

$$\Delta s_t = \beta(L)\, \Delta s_{t-1} + \gamma(L)\, \Delta\, \hat{s}^*_t + \delta(s - \hat{s}^*)_{t-1} + \varepsilon_t \tag{3}$$

The second alternative is to substitute (2) into (1) and estimate the resulting equation described by (4):

$$\Delta s_t = \beta(L)\, \Delta s_{t-1} + \gamma(L)\, \Delta\, (\pi' z_t) + \delta(s - \pi' z)_{t-1} + \varepsilon_t \tag{4}$$

We use the first methodology because it requires estimation of a smaller number of parameters, an important consideration when estimating non-linear models using relatively short samples.

The linearity assumption in (1) can be tested using the procedure described in Saikonnen and Luukkonen (1988), Luukkonen et al (1988), Granger and Teräsvirta (1993) and Teräsvirta (1994). To implement this test we estimate

$$(s-\hat{s}^*)_t = \gamma_{00} + \sum_{j=1}^{\phi}\{\gamma_{0j}(s-\hat{s}^*)_{t-j} + \gamma_{1j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)_{t-d} + \gamma_{2j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)^2_{t-d} + \gamma_{3j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)^3_{t-d}\} + \gamma_4(s-\hat{s}^*)^2_{t-d} + \gamma_5(s-\hat{s}^*)^3_{t-d} + v(t) \tag{5}$$

where $(s - \hat{s}^*)$ is the estimated deviation from equilibrium obtained from (2), d is the delay parameter of the transition function to be used and $v(t) \sim niid\,(0,\sigma^2)$. Linearity implies the null hypothesis H_0: $[\gamma_{1j} = \gamma_{2j} = \gamma_{3j} = \gamma_4 = \gamma_5 = 0]$ for all $j \in (1,2...\phi)$. This can be tested using an LM-type test. Having determined ϕ through inspection of the partial autocorrelation function[4], (5) can be estimated for all plausible values of the delay parameter d. The correct value of d is that which yields the largest value of the test statistic.

If we reject linearity, the second stage in our methodology is to estimate a non-linear model of exchange rate adjustment. We do this using the Quadratic Logistic Smooth Transition Error Correction Model (QL-STECM), specified as follows:

$$\Delta s_t = \theta_t M_{It} + (1-\theta_t)\, M_{Ot} + \varepsilon_t \tag{6}$$

$$M_{It} = \beta_{I1}(L)\, \Delta s_{t-1} + \gamma_{I1}(L)\, \Delta \hat{s}^* + \delta_I (s-\hat{s}^*)_{t-1} + \varepsilon_t \tag{7}$$

$$M_{Ot} = \beta_{O1}(L)\, \Delta s_{t-1} + \gamma_{O1}(L)\, \Delta \hat{s}^* + \delta_O (s-\hat{s}^*)_{t-1} + \varepsilon_t \tag{8}$$

$$\theta_t = pr\,\{\, \tau^L \le (s-\hat{s}^*)_{t-d} \le \tau^U \,\} = 1 - \frac{1}{1+e^{-\sigma[(s-\hat{s}^*)_{t-d}-\tau^L][(s-\hat{s}^*)_{t-d}-\tau^U]}} \tag{9}$$

Equation (6) models exchange rate changes as a weighted average of the linear models M_I and M_O, where M_I represents the inner regime and M_O the outer regime. Equations (7) and (8) describe M_I and M_O as linear error-correction models, similar to (1). Equation (9) specifies the regime weight θ as the probability that the transition variable $(s-\hat{s}^*)_{t-d}$ lies within the "regime boundaries" τ^L and τ^U, where the probability is described using a quadratic logistic function and we expect $\tau^L < 0$ and $\tau^U > 0$. Exchange rates are mainly determined by M_I (the inner regime) when the exchange rate is close to its fundamental value and mainly by M_O (outer regime) in periods of significant exchange rate misalignment, with σ denoting the speed of transition between the two regimes.

The speed of adjustment of the exchange rate differs between regimes if $\delta_I \ne \delta_O$. If $\delta_I = 0$ and $\delta_O < 0$, the exchange rate only adjusts towards its fundamental

[4] Granger and Teräsvirta (1993) and Teräsvirta (1994) advise against choosing ϕ using an information

value in the outer regime, evolving as a random walk in the inner regime. In the case where $\tau^U + \tau^L = 0$, the model is in effect equivalent to the ESTAR model since the speed of adjustment depends only on the size of the deviation of exchange rates from fundamentals. If $\tau^U + \tau^L \neq 0$, the model is more general than the ESTAR model since the speed of adjustment depends both on the size and on the sign of the deviation from fundamentals. In particular, if $\tau^U + \tau^L > 0$, the exchange rate responds more vigorously to under-valuations, while $\tau^U + \tau^L < 0$ indicates a stronger response to over-valuations[5].

3. DATA AND EMPIRICAL RESULTS

3.1. Data

We use quarterly data on the exchange rates of the Greek Drachma and Turkish Lira against the ECU. For our PPP model of exchange rate fundamentals, we use data on producer (wholesale) prices. For our FPMM model, we use a narrow M1 definition of the nominal money supply and data on real GDP[6]. Data availability limits us to analysing the period 1982(1)-2000(4) for Greece and 1986(1)-2001(3) for Turkey where we model fundamentals using PPP and 1980(1)-2000(4) for Greece and 1987(1)-2001(3) for Turkey in the case of FPMM fundamentals.

Preliminary analysis for the order of integration of the individual variables, using both the augmented Dickey-Fuller (ADF) tests (Dickey and Fuller, 1979) and the semi-parametric Phillips-Perron tests (Phillips and Perron, 1988 and Perron, 1988),

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criteria such as the Akaike since this may induce a downward bias.

[5] Asymmetry can also be captured by the logistic smooth transition autoregressive (LSTAR) model (Granger and Teräsvirta, 1993) or the M-TAR model (Enders and Dibooglu, 2001). However these models have a single threshold, giving an "upper" regime and a "lower" regime. We feel that the QL-STECM is more useful for modelling the exchange rate since the QL-STECM, along with most theoretical models of non-linear exchange rate adjustment, postulates the existence of an inner and an outer band (rather than a single threshold) for the adjustment of the exchange rate. Finally, non-linear exchange rate adjustment can be modelled using a three-regime QL-STECM model. Such a model would account not only for asymmetric regime bands but also for different speeds of adjustment within the inner-regime, over-valuation rates below the lower threshold of the inner-regime, and under-valuation rates exceeding the upper band of the inner-regime. Estimates of this type of model were not successful.

[6] In the case of Greece, this variable was not available on a quarterly basis, hence we used as a proxy the volume of real industrial production.

suggest that all variables are I(1). Neither these tests, nor subsequent estimates of cointegrating relationships are affected by non-linearities (Michael, et al, 1997).

3.2. Linear cointegration tests

As a first step towards testing exchange rate behaviour we examine the properties of Greek and Turkish real exchange rates against the Euro (the ECU prior to 1999).[7] A visual inspection of the two series in Figure 1 suggests the existence of a unit root for the Drachma/Euro rate and mean-reverting behaviour for the Lira/Euro one. The ADF and Phillips-Perron unit root tests presented in Table 1 confirm this. This informal evidence would tend to reject PPP for Greece but not for Turkey.

This impression is confirmed by more formal statistical evidence. Table 2 reports estimates of the PPP and the FPMM equations (2a and 2b) for Greece and Turkey, together with ADF and Phillips-Perron tests for cointegration. For Greece, neither test rejects the null of no cointegration at the 5% level. This is consistent with the findings of Karfakis and Moschos (1989). For Turkey, both tests reject non-cointegration of the FPMM, while the results for PPP are mixed. On balance, and taking into account the stationarity of the Turkish real exchange rate, we feel that both models for Turkey are cointegrated. This is consistent with the findings of McNown and Wallace (1989) for four other high inflation countries and suggests that their finding that fundamentals' models are cointegrated where inflation is high can be extended to other countries.

Figures 2(a) and 2(b) show the exchange rate misalignments implied by these estimates. Misalignments for PPP and FPMM estimates are similar, although somewhat higher for the FPMM. The Greek drachma appears to have been overvalued against the ECU during the first half of the 1980s; undervalued between 1986-94; and increasingly overvalued during 1995-2000, thus confirming that the strong-drachma policy followed by Greek authorities during that period resulted in currency over-valuation. The devaluations of 1983(1) and 1998(4) appear to have corrected over-valuation. The devaluation of 1985(4) appears to have achieved its goal of undervaluing an exchange rate that was previously close to equilibrium[8]. Our estimates also suggest that the drachma was overvalued by 3-5% on entry to the ERM. The Turkish experience appears to have been one of relatively short, alternating periods of over-valuation and under-valuation, with each period typically lasting between three to four years. There were very violent changes in 1994 and 2001, which accompanied the two recent financial crises experienced by Turkey[9]. Misalignment of the Lira is generally much more volatile than misalignment of the Drachma.

3.3. Linear Error Correction Models

Table 3 presents estimates of the linear error correction equation in (3), where we present estimates of parsimonious models obtained using a general-to-specific specification search on a baseline model using twelve lags of all variables. For Greece, the PPP model passes all mis-specification tests, although the FPMM model has some residual non-normality. For Turkey, both models pass the mis-specification tests[10].

Since the Granger representation theorem states that any cointegrated relationship can be represented as an error-correction model, we can use the significance of error-correction terms as an additional test of cointegration (for a detailed discussion on this point, see Madalla and Kim (1998), section 6.3; however we

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[7] The real exchange rate is defined as the product of the nominal exchange rate by the ratio of foreign (European) to domestic price levels. An increase (reduction) in the value of the real exchange rate denotes a real depreciation (appreciation) of the domestic currency against the Euro.

[8] For a detailed discussion of monetary strategies in Greece during the post-1974 period, see Mourmouras and Arghyrou (2000).

[9] For a detailed discussion of the Turkish financial crises, see OECD, 2002.

[10] These models include dummies for periods of particular turbulence: 1983Q1, 1985Q4 and 1998Q1 for Greece; and 1994Q1, 1994Q2 and 2000Q1 for Turkey. These dummies improve some mis-specification tests but do not change the qualitative nature of the results.

should also note that these tests will be invalid if the relationship between exchange rates and fundamentals is non-linear). We note that the estimated coefficients on the error-correction terms are insignificant for Greece but significant for Turkey. This is consistent with the results in Tables 1 and 2. The speed of adjustment to equilibrium is also much higher in Turkey. In summary, Greece exhibits both puzzles identified by Taylor, Peel and Sarno (2001) since fundamentals models are not cointegrated and the speed of adjustment is implausibly low. In Turkey, by contrast, fundamental exchange rate relationships are cointegrated and the speed of adjustment is higher.

3.4. Linearity tests

Table 4 presents our tests of linearity, using (5). Inspection of the partial autocorrelation functions in Figures 3(a) and 3(b) reveals that $\phi=1$. We calculated the test statistics for 8 values of the delay parameter, *d*, reporting the test statistic for that value which maximises the test (Granger and Teräsvirta, 1993, and Teräsvirta, 1994). The null hypothesis of linear adjustment is rejected at the 5% level in both countries (even though linear models appeared to work well in the case of Turkey). We therefore conclude that the relationship between exchange rates and fundamentals is non-linear and proceed to estimate a non-linear model.

3.5. Non-linear Error Correction Models

Table 5 presents the estimates of the QL-STECM models described by equations (6) to (9). The reported equations are again obtained using a general-to-specific specification search. The three main results emerging from Table 5 are the following: First, exchange rate adjustment is asymmetric. The null hypotheses H_0: $\tau^L + \tau^U = 0$ is clearly rejected against the two-sided alternative H_0: $\tau^L + \tau^U \neq 0$ for the FPMM model in Greece and the PPP model in Turkey. The only case where the restriction of symmetry passes is the FPMM equation for Turkey, but even in that case, the absolute value of the point estimate for the upper band threshold is noticeably higher than that of the lower threshold. These findings suggest that using symmetric models, such as the ESTAR, may be misleading.

Second, exchange rates are more sensitive to over-valuations than under-valuations. For Greece, the regime thresholds are estimated to be (–3%, 4.5%) for PPP and (-4%, 10%) for the FPMM. For Turkey they are (-3.5%, 5%) for PPP and (-6.5%, 10%) for the FPMM. The upper threshold is larger than the (absolute value of the) lower in every case. The null hypothesis H_0: $\tau^L + \tau^U = 0$ is rejected against the one-sided alternative H_0: $\tau^L + \tau^U > 0$ for the FPMM model in Greece and the PPP model in Turkey at the 95% level. It is also rejected for the Greek PPP model at the 93% level.

Third, our estimates are consistent with the existing literature since the adjustment of exchange rates towards fundamentals is stronger when exchange misalignment is more pronounced[11]. In the case of Greece, the error-correction term is insignificant in the inner regime but significant and relatively large in the outer regime. Exchange rates are therefore a random walk in the inner regime. This is also true of Turkey for the FPMM. In the case of PPP, exchange rates adjust towards equilibrium in both regimes, although the speed of adjustment is higher in the outer regime. We continue to find that adjustment is stronger in Turkey than Greece and is stronger with for PPP than the FPMM. If a significant error-correction term in the outer regime indicates cointegration between the exchange rate and macroeconomic fundamentals, then for both countries our estimates provide significant evidence in favour of the PPP and FPMM in their non-linear versions.

Our findings for Greece support and extend the argument that non-linear models of exchange rate adjustment can help to overcome the exchange rate puzzles of Taylor, Peel and Sarno (2001). Our findings for Turkey extend this argument even further, since they provide evidence of non-linear exchange rate adjustment even in an economy where fundamentals model work well, suggesting that non-linear adjustment is more pervasive than previously thought.

We also note that these estimates are superior to those of the linear models reported in Table 3. The regression standard errors are all smaller than those of the linear models reported in Table 3. In three out of four cases the reduction in the standard error is large. We also note that all models pass the mis-specification tests, in contrast to estimates of the linear models reported in Table 3. The PPP model again outperforms the FPMM model, the difference being especially marked in the case of Turkey. Finally, in each case, the estimate of the σ parameter implies a moderate speed of transitions between regimes. This tends to argue against models such as the TAR or MTAR, which postulates abrupt change between the two regimes[12].

Figures 4(a) and 4(b) plot currency misalignment against the estimated thresholds. In the case of Greece, the PPP model suggests exchange rates were in the inner regime for most of this period. The FPMM model gives more weight to the outer regime, which was dominant in 1985-1990 and 1997-2000. Both models suggest the devaluations of 1983Q1 and 1998Q1, restored the drachma's exchange rate from the outer regime to the inner regime. By contrast, the devaluation of 1985Q4, moved the Drachma from the inner regime to the outer regime, achieving the intended deliberate under-valuation of the currency. The estimates also show that the effectively fixed exchange rate policy followed by Greece between 1995 and 1997, led to currency over-valuation taking values beyond, according to the FPMM, or very close, according to the PPP, the level of over-valuation the government was willing to accept. Finally, both

[11] These models include the crisis dummies mentioned in footnote 10, which ensures that our non-linear findings apply to the whole of our samples and do not simply pick up the influence of these one-off events.

models suggest that at the end of 2000, drachma's misalignment against the Euro was hovering around the lower threshold, with the PPP suggesting an over-valuation rate just within the band, the FPMM just outside. Joining the Euro with an overvalued exchange rate implies that EMU participation came at a premium for Greece. What our estimated bands suggest is that this premium was chosen in a way that would render it just affordable from the Greek authorities' point of view.

In the case of Turkey, both models suggest a more frequent alternation between the two regimes, especially during 1986-1995. However the inner regime was dominant from 1996 to mid-2000. The two models provide a consistent picture and suggest that incidences of over-valuation beyond the band's lower limit are roughly as frequent as incidences of under-valuation beyond the upper threshold of the band. Finally, it seems that the devaluations of 1994 and 2001 restored the Lira to the inner regime.

Our findings relating to non-symmetric exchange rate adjustment have two implications for exchange rate policy in Greece and Turkey in this period. First, the finding that the absolute value of the upper threshold is higher than the lower implies that policymakers in both countries were more tolerant of under-valuation than over-valuation. Second, our findings imply that both countries were correct not to committing themselves to an exchange rate target against the ECU, for example, by participating in the ERM prior to its reform in late 1992. Given systematically higher inflation rates than the ERM countries over this period, a policy of shadowing the ECU very closely would have induced substantial currency over-valuation, leading to conflict between domestic and exchange rate policy goals and to probable speculative attack. The experience of Greece between 1995-1998 tends to confirm this analysis. During that period, the Drachma's rate against the German mark (and the ECU) was effectively leading, as is apparent from Figure 4(a), to significant currency over-valuation. This

[12] However, in all cases σ is imprecisely estimated as the likelihood function is very insensitive to this parameter (see the detailed discussion on this point in van Dijk et al., 2002).

was not consistent with domestic policy objectives. This inconsistency did not go unnoticed. The drachma was subject to a speculative attack in November 1997 and was finally officially devalued a few months later in March 1998. Following this devaluation, the drachma did join the ERM.

4. SUMMARY AND CONCLUDING REMARKS

In recent years, a number of authors have tested the empirical validity of a class of models for the exchange rate that predict non-linear adjustment towards an equilibrium determined by macroeconomic fundamentals. By validating these models, recent research has achieved a significant contribution in resolving two long-standing "puzzles" in the literature on exchange rates, the first relating to the long-run validity of monetary models of the exchange rate, the second to the excessive volatility of exchange rates relative to that of macroeconomic fundamentals.

We have argued that these models are restrictive. In particular, they do not allow for the possibility that exchange rate adjustment is asymmetric. We have proposed an alternative model of non-linear exchange rate adjustment that captures the features of existing models but also allows for asymmetric adjustment. Applying our model to Greece and Turkey, we find clear evidence of asymmetric exchange rate adjustment.

Our work can be extended and refined in several ways. It would be helpful to develop a formal model of non-linear exchange rate behaviour, perhaps drawing on the recent literature on non-linear policy rules, in order to provide a clearer theoretical grounding for our work. This might also suggest a way to examine the effects of institutional reforms, such as central bank independence, on exchange rate behaviour. A further empirical extension would be to investigate the degree to which movements of the individual variables involved in the analysis (exchange rates and macro fundamentals) contribute towards the non-linear behaviour of the exchange rates. Finally, an interesting empirical application would be to compare the in- and out-of sample forecasting capacity of the non-linear models estimated here against the forecasts obtained by certain recently developed econometric models which use the information content of the term structure of forward exchange rates (see Clarida et al, 2002). Such a comparison would be particularly interesting in applications relating to medium- and long-term exchange rate prediction.

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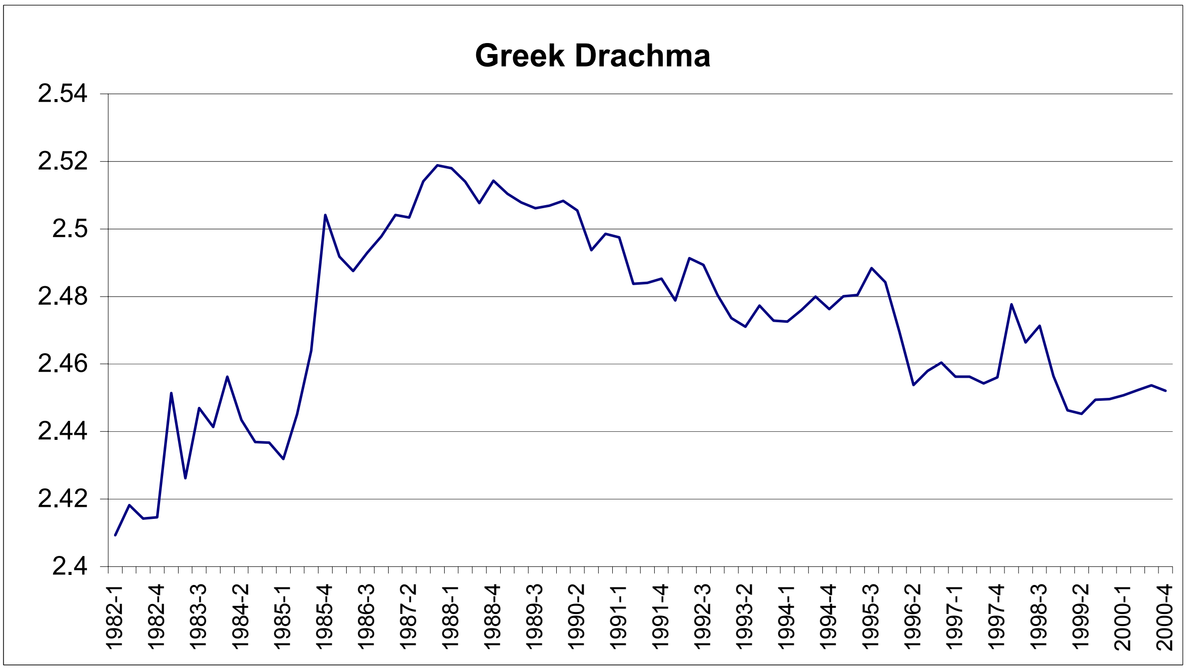
Figure 1: Real exchange rate against the euro



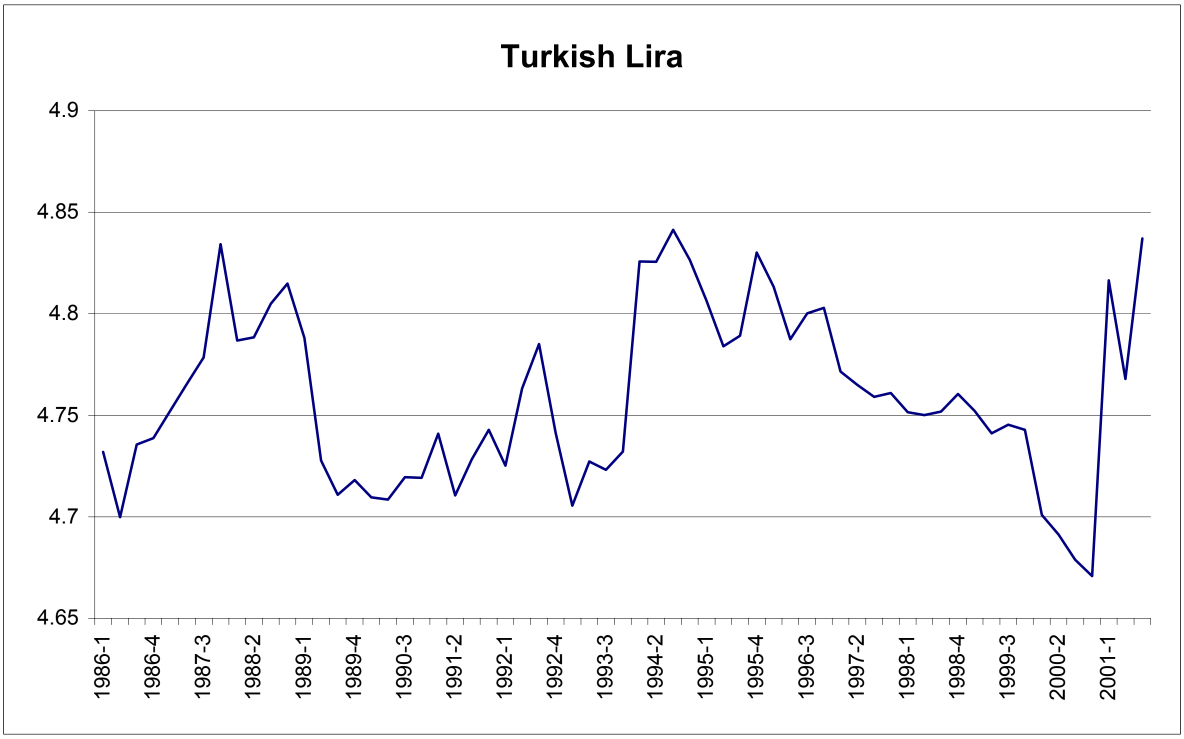


Table 1

Unit root tests - Real exchange rates

GREECE		TURKEY	
ADF	Phillips-Perron	ADF	Phillips-Perron
-2.437 [-2.900]	-2.391 [-2.900]	-2.989 [-2.908]	-3.032 [-2.908]

95% critical values in square brackets

NOTES: The lag-length of the reported ADF tests is the lowest lag for which the ADF equation does not present serial correlation (0 for both countries). The reported Philips-Perron tests and the accompanying critical values are obtained by E-views. To determine the truncation lag of the Phillips-Perron test, E-views uses the Newey-West automatic truncation lag selection function.

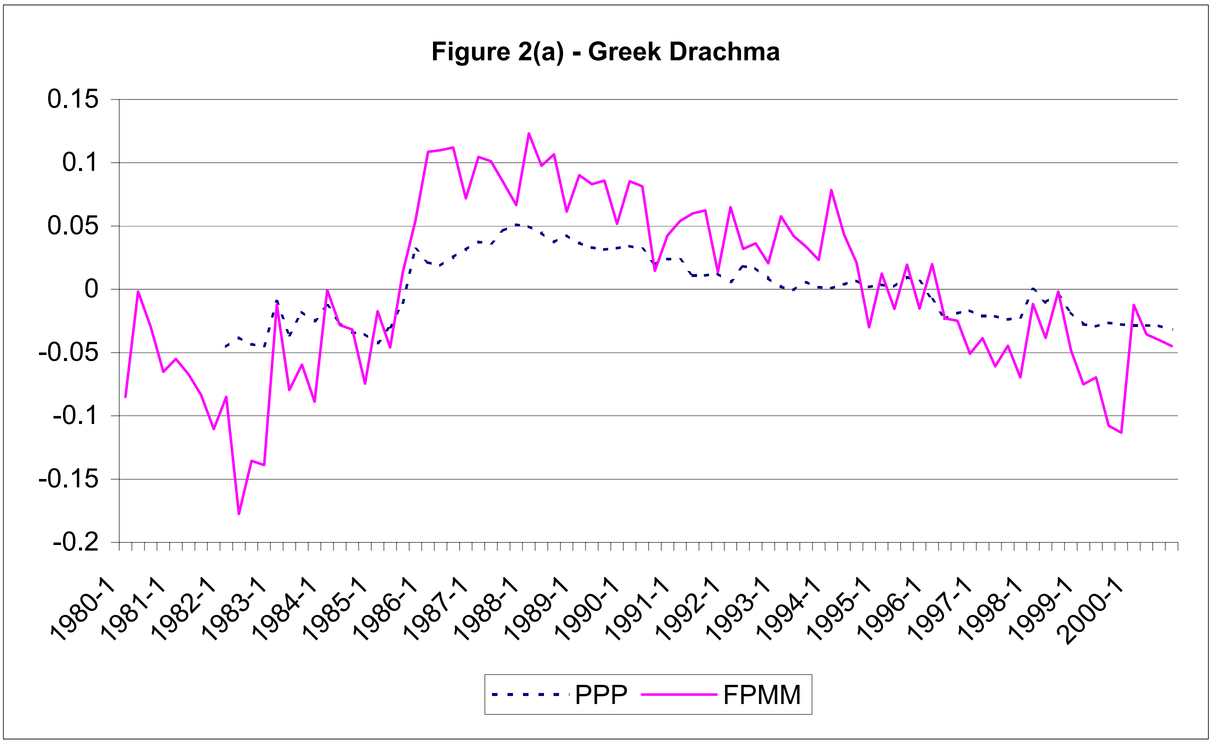
Table 2

Cointegrating regressions

	GREECE		TURKEY	
	PPP	FPMM	PPP	FPMM
Equation Sample period	(2a) 1982(1)-2000(4)	(2b) 1980(1)-2000(4)	(2a) 1986(1)-2001(3)	(2b) 1987(1)-2001(3)
α	1.88 (0.56)	0.91 (0.58)	5.00 (1.28)	5.83 (0.23)
β_1	0.96 (0.06)	0.80 (0.12)	1.00 (0.02)	1.10 (0.01)
β_2	-0.66 (0.33)	-1.11 (0.57)	-1.13 (0.66)	-0.23 (0.12)
R^2	0.99	0.93	0.99	0.99
Durbin-Watson	0.15	0.33	0.59	0.96
ADF [95% CV]	-1.933 [-1.945]	-1.406 [-1.944]	-3.041 [-1.946]	-3.647[-1.947]
Phillips-Perron Z($t_{\hat{b}}$) [95% CV]	-1.874 [-3.856]	-2.481 [-3.845]	-3.157 [-3.881]	-4.249 [-3.891]

NOTES: Numbers in parentheses are standard errors. The lag-length of the reported ADF tests is the lowest lag for which the ADF equation does not present serial correlation (for Greece, 0 for the PPP and 3 for the FPMM equations; for Turkey 0 for PPP and 3 for the FPMM equations). The reported Philips-Perron tests are obtained by E-views. To determine the truncation lag of the Phillips-Perron test, E-views uses the Newey-West automatic truncation lag selection function. Critical values for the Phillips-Perron cointegration test ($t_{\hat{b}}$) are obtained from MacKinnon (1991).

Figure 2: Currency overvaluation relative to macroeconomic fundamentals



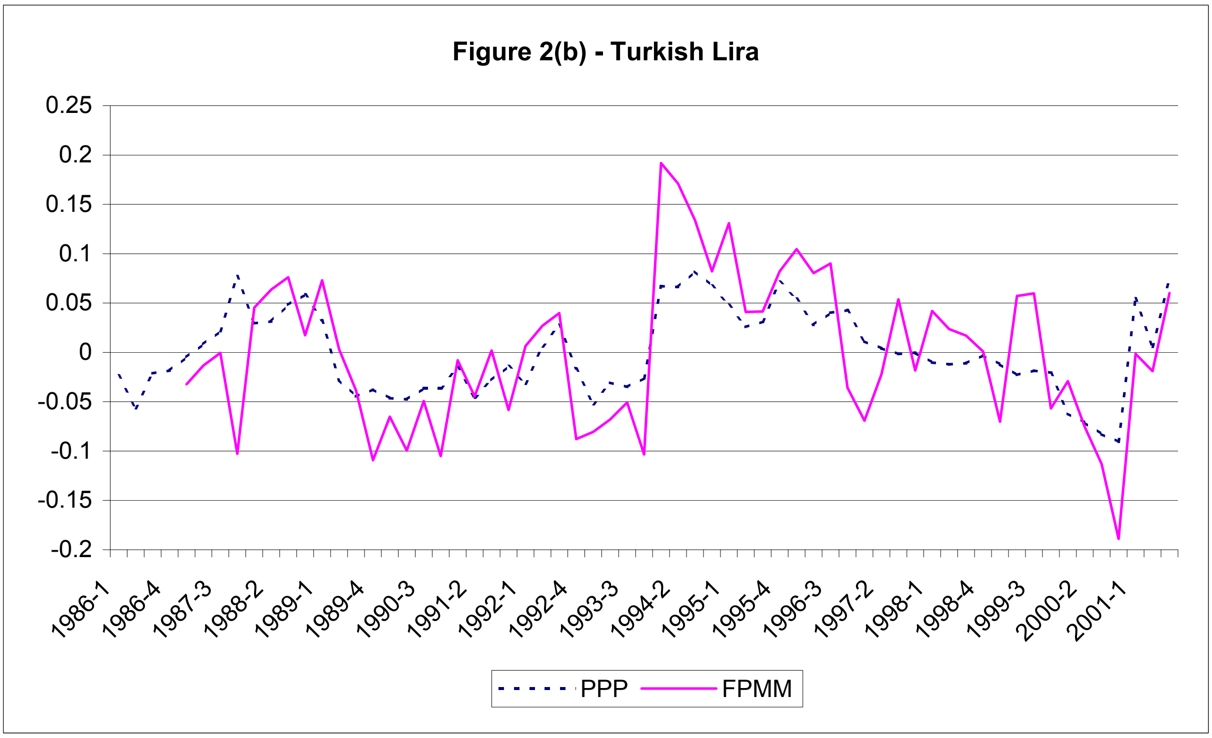


Table 3

Linear Error Correction Models

	GREECE		TURKEY	
	PPP	FPMM	PPP	FPMM
Sample period	1982(1)-2000(4)	1980(1)-2000(4)	1986(1)-2001(3)	1987(1)-2001(3)
constant	-0.0005 (0.002)	0.007 (0.001)	-0.017 (0.014)	-0.014 (0.016)
Δs_{t-1}				0.169 (0.108)
Δs_{t-2}			0.300 (0.089)	0.355 (0.114)
Δs_{t-4}		0.163 (0.084)		0.203 (0.117)
Δs_{t-7}	0.126 (0.057)		0.265 (0.092)	0.398 (0.116)
Δs^*_t	0.802 (0.137)		0.653 (0.174)	
$(s-s^*)_{t-1}$	-0.032 (0.035)	-0.015 (0.018)	-0.187 (0.094)	-0.161 (0.065)
D1	0.047 (0.009)	0.080 (0.011)	0.108 (0.021)	0.130 (0.024)
D2	0.024 (0.007)	0.020 (0.011)	0.054 (0.025)	0.121 (0.028)
D3			0.152 (0.020)	0.152 (0.025)
R^2	0.752	0.449	0.78	0.70
Std Error	0.0067	0.0108	0.019	0.023
RSS	0.0027	0.0087	0.017	0.025
DW	1.74	2.10	1.66	1.71
F ar	0.45	0.13	0.06	0.52
F arch	0.67	0.45	0.43	0.46
χ^2 norm	0.30	0.00	0.39	0.22
F het	0.30	0.14	0.98	0.63
RESET	0.45	0.39	0.19	0.26

NOTE: Numbers in parentheses are standard errors. D1, D2 and D3 are intercept dummies denoting periods of major currency crises. For Greece, D1 and D2 are defined for 1985(4) and 1998(3); For Turkey, D1, D2 and D3 are defined in 1994(1), 1994(2) and 2001(1) respectively. F ar is the Lagrange Multiplier F test for residual serial correlation of up to fifth order. F arch is the fourth order Autoregressive Conditional Heteroskedasticity F test. χ^2 normality is a Chi-square test for normality. F het is an F test for heteroskedasticity. The numbers reported for these tests are p-values.

Table 4

Linearity tests

	ϕ	*d*	*LM*	*p-value*
Greece				
PPP	1	4	2.547	0.036
FPMM	1	4	2.328	0.028
Turkey				
PPP	1	4	10.826	0.000
FPMM	1	5	2.894	0.023

NOTES:

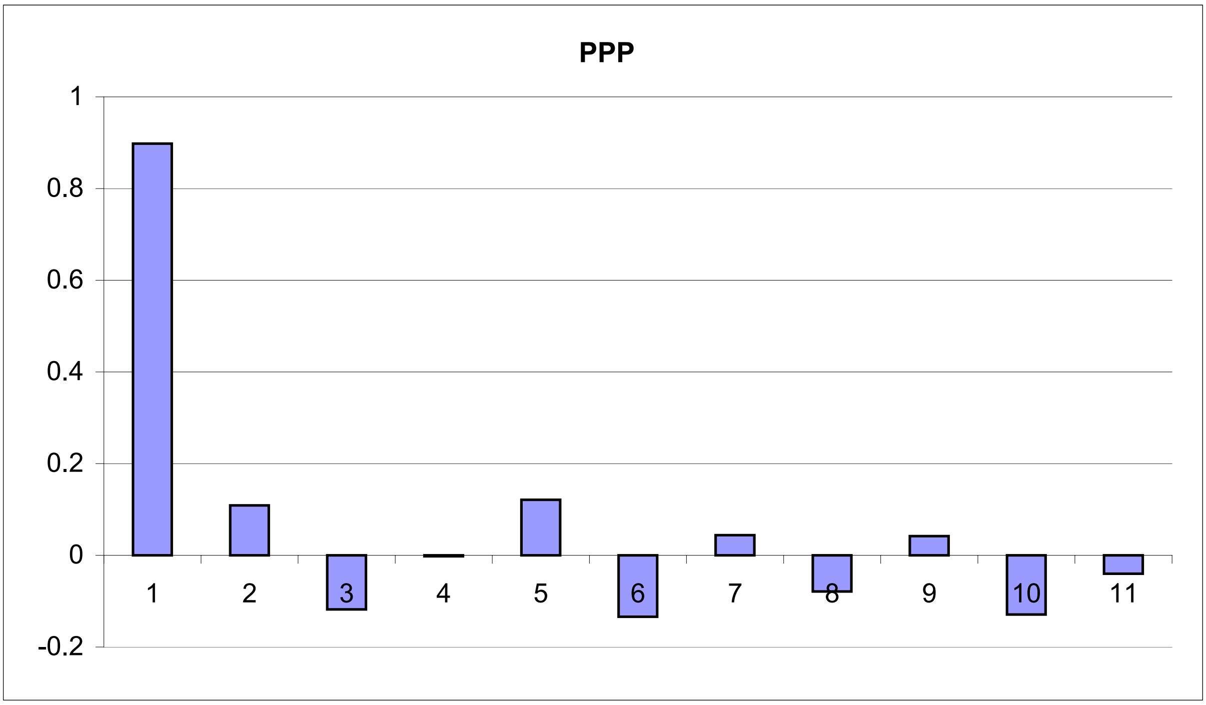
ϕ is the order of the autoregressive component and *d* the order of the delay parameter in the artificial regression (5).

$$(s-\hat{s}^*)_t = \gamma_{00} + \sum_{j=1}^{\phi}\{\gamma_{0j}(s-\hat{s}^*)_{t-j} + \gamma_{1j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)_{t-d} + \gamma_{2j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)^2_{t-d} + \gamma_{3j}(s-\hat{s}^*)_{t-j}(s-\hat{s}^*)^3_{t-d}\} + \gamma_4(s-\hat{s}^*)^2_{t-d} + \gamma_5(s-\hat{s}^*)^3_{t-d} + v(t) \quad (5)$$

The reported LM statistics are the estimated scores and the associated p-values are obtained from applying an LM F-test on equation (5) where the null is described by:

$H_0 = [\gamma_{2j} = \gamma_{2j} = \gamma_{3j} = \gamma_4 = \gamma_5 = 0]$

Figure 3(a) - Partial Autocorrelation Functions: Greek Drachma



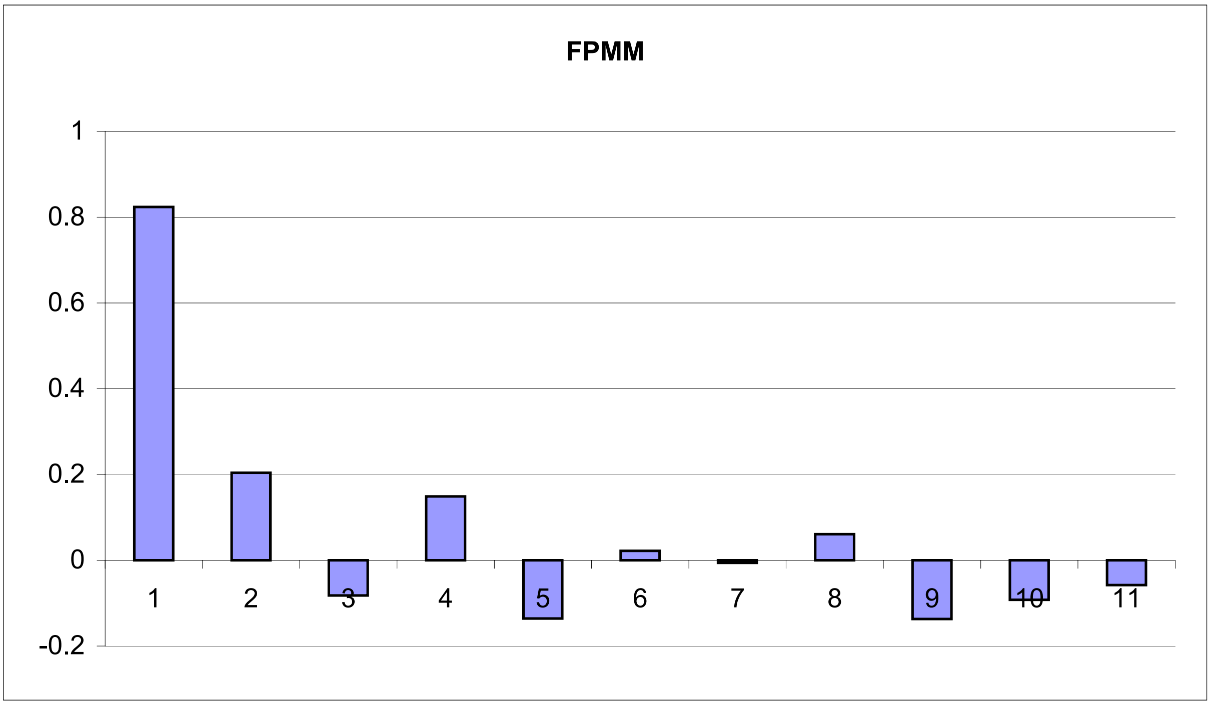
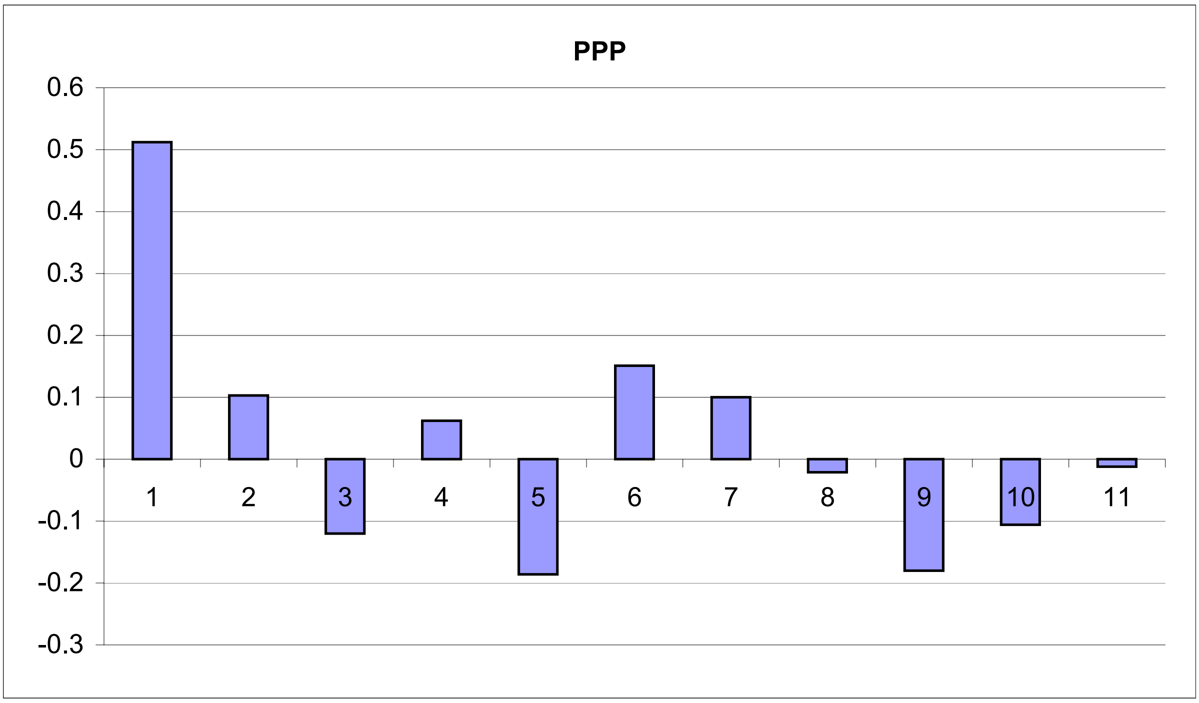


Figure 3(b) - Partial Autocorrelation Functions: Turkish Lira



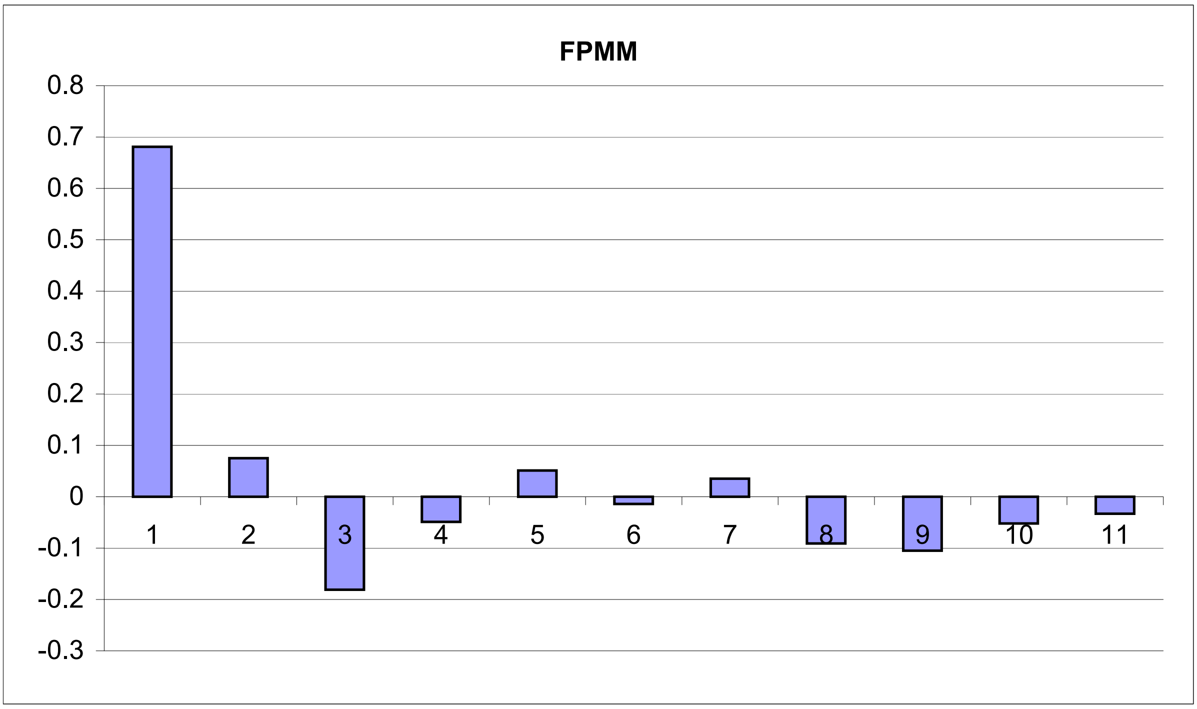


Table 5
Non-Linear Error Correction Models

	GREECE		TURKEY	
	PPP	FPMM	PPP	FPMM
Sample period	1982(1)-2000(4)	1980(1)-2000(4)	1986(1)-2001(3)	1987(1)-2001(3)
M1				
constant	-0.0009 (0.0016)	0.0007 (0.0026)	-0.061 (0.019)	0.033 (0.007)
Δs_{t-1}	0.115 (0.057)		0.531 (0.161)	0.242 (0.116)
Δs_{t-2}		0.321 (0.179)		
Δs_{t-4}		0.200 (0.081)		
Δs_{t-6}		0.364 (0.171)	0.432 (0.189)	
Δs_{t-7}				
Δs_{t-11}			0.242 (0.160)	
Δs^*_t	0.841 (0.144)		0.759 (0.182)	
$(s-s^*)_{t-1}$	-0.009 (0.039)	-0.0045 (0.026)	-0.283 (0.097)	0.065 (0.057)
M2				
constant	0.015 (0.003)	-0.0032 (0.0027)	-0.031 (0.020)	-0.010 (0.010)
Δs_{t-2}			0.449 (0.106)	0.503 (0.143)
Δs_{t-7}			0.403 (0.109)	0.547 (0.143)
Δs_{t-8}		0.719 (0.139)		
Δs_{t-11}		0.409 (0.182)		
Δs^*_t			0.848 (0.271)	
$(s-s^*)_{t-1}$	-0.203 (0.095)	-0.088 (0.026)	-0.469 (0.151)	-0.269 (0.096)
σ	10.001 (9.610)	25.000 (36.304)	10.373 (9.959)	8.174 (8.85)
τ^U	0.045 (0.0035)	0.098 (0.004)	0.051 (0.003)	0.099 (0.016)
τ^L	-0.031 (0.0018)	-0.039 (0.006)	-0.035 (0.002)	-0.064 (0.009)
D1	0.070 (0.008)	0.040 (0.012)	0.080 (0.022)	0.113 (0.025)
D2	0.025 (0.007)	0.026 (0.009)		0.165 (0.026)
D3			0.127 (0.020)	0.158 (0.025)
R^2	0.78	0.71	0.89	0.79
Std Error	0.0066	0.0088	0.0153	0.0208
RSS	0.0024	0.0045	0.0082	0.0182
DW	1.88	1.82	2.28	1.53
F ar	0.32	0.25	0.34	0.34
F arch	0.60	0.74	0.85	0.31
χ^2 norm	0.50	0.55	0.22	0.18
F het	0.62	0.55	0.99	0.54
F-Test H_0: $\tau^L + \tau^U = 0$ against H_1: $\tau^L + \tau^U \neq 0$ 95% critical values in square brackets	3.88 [5.28]	21.70 [5.28]	5.77 [5.42]	2.15 [5.42]
F-Test H_0: $\tau^L = \tau^U$ against H_1: $\tau^L < \tau^U$ 95% critical values in square brackets	3.88 [4.00]	21.70 [4.00]	5.77 [4.08]	2.15 [4.08]

NOTE: Numbers in parentheses are standard errors. D1, D2 and D3 are intercept dummies denoting periods of major currency crises. For Greece, D1 and D2 are defined for 1985(4) and 1998(3); For Turkey, D1, D2 and D3 are defined in 1994(1), 1994(2) and 2001(1) respectively. F ar is the Lagrange Multiplier F test for residual serial correlation of up to fifth order. F arch is the fourth order Autoregressive Conditional Heteroskedasticity F test. χ^2 normality is a Chi-square test for normality. F het is an F test for heteroskedasticity. The numbers reported for these tests are p-values.

Figure 4: Currency overvaluation relative to non-transaction bands