EXCHANGE RATE PEGGING: CREDIBILITY AND FUNDAMENTALS

Evidence from Greece

by

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The paper examines the predictability of a currency-peg collapse on the basis of the movements of underlying fundamentals and the observed behaviour of currency markets. Our findings suggest that the *ultimate* collapse of a peg is predictable if the peg is inconsistent with valid long-run macroeconomic relationships. A strong currency policy is confirmed to be helpful in terms of reducing inflation but its prolonged implementation reduces the peg's credibility. In the case-study examined, markets appear to have anticipated the peg's collapse almost a year before it occurred. However, the results show that the *exact* timing of the devaluation took markets by surprise. Even under conditions of devaluation expectations, central banks have the capacity to surprise the markets.

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1. INTRODUCTION

How predictable are currency devaluations? Do movements of macroeconomic fundamentals contain information regarding the future sustainability of a currency peg? What information concerning the credibility of an exchange rate target can we gather from the observed behaviour of currency markets? These questions seem increasingly relevant within the framework of a globalized and deregulated economic system where on the one hand market participants pursue international portfolio diversification; and, on the other, a number of national central banks set exchange rate targets in an effort to achieve price stability. The 1990s saw a number of currency-pegs collapses, some of which share common background macroeconomic similarities. Have markets drawn any conclusions from these events? And how are the latter relevant to monetary policy makers? In particular, do Central Banks, in view of past experience, determine their policies taking into consideration the behaviour of currency markets? In this paper we attempt to provide some insights regarding the link between fundamentals, market behaviour and central bank decisions based on a fairly recent currency-peg collapse, the one of the Greek drachma. In March 1998, the latter was devalued against the ECU by 14%. This event discontinued the strong drachma policy to which Greek monetary authorities were at the time officially committed. The policy, based on an "the-advantages-of-tying-one's-own-hands" logic (see Giavazzi and Pagano, 1988) and implemented consistently since 1988, effectively involved a rate of nominal depreciation of the drachma against the DM smaller than the inflation differential between Greece and Germany. Indeed, the 1990s saw a significant reduction in Greek inflation and interest rates on government bonds (see Figure 1). And yet, despite the progress achieved, the drachma was devalued. Was the devaluation predictable? We address the question from two points of view. First, we investigate whether the drachma/DM exchange rate is linked to macroeconomic fundamentals and, if it is, whether drachma's peg to the DM was consistent with their movements. Second, we investigate whether markets perceived the peg as credible, i.e. whether they anticipated the devaluation. The questions are interesting not least because at present a number of countries set exchange rate targets in an environment of increasing financial liberalisation. For example, it is now very likely that the countries which aspire to be involved in the next EU enlargement (e.g. transition economies) will be asked to participate in the ERM-II mechanism. The Greek experience may provide useful lessons concerning devaluation risks and markets' behaviour in the presence of fixed exchange rate policies. The remainder of the paper is structured follows: Section 2 presents a brief literature review. Section 3 presents the data and our econometric results. Section 4 discusses the results. Section 5 summarises and concludes.

2. EXCHANGE RATES AND FUNDAMENTALS: A BRIEF REVIEW

The question as to whether macroeconomic fundamentals influence the behaviour of the exchange rate is certainly not a new one. For example, Meese and Rogoff (1983a) suggest that the random walk model outperforms a range of fundamentals-based models of exchange rate determination at horizons of up to a year, an opinion still shared by many economists (see e.g. Frenkel and Rose, 1995, Rogoff 1999). On the other hand, the same authors (Meese and Rogoff, 1983b) find that the random walk model does not yield the minimum forecast errors when the forecast horizon is extended to periods beyond twelve months, a conclusion supported by Mark (1995) and Chinn and Meese (1995). This second group of findings suggest that if fundamentals play a role in exchange rate determination, this is more relevant in a long-run framework. To investigate the validity of this proposition economists set-off by testing whether the purchasing power parity (PPP) condition holds in the long-run (MacDonald, 1999). The reason is that PPP is a basic assumption in a number of fundamentals-based models of exchange rate determination where the exchange-rates regime is a floating one[1]; but also in a number of fundamentals-based models of speculative attacks against a fixed exchange rate[2]. If the PPP hypothesis were consistent with the data, PPP calculations would have significant diagnostic value in order to assess whether a currency is fundamentally over- or undervalued (Isard, 1995). Failure to establish the validity of the PPP would imply that PPP-based models may have to be reconsidered in the light of the empirical evidence. Since Frenkel (1981a) concluded that PPP had collapsed in the 1970s, a number of researchers have dealt with the issue, mainly using integration and cointegration econometric techniques. In this context, PPP tests have mainly taken two forms. Initially, researchers focused on the time series properties of real exchange rates. According to both versions of the PPP (absolute and relative), the real exchange rate, q_t, should be constant over time. If long-run PPP holds, q_t should be mean-reverting and described by equation (1) below:

$$q_t = \rho\ q_{t-1} + \varepsilon_t, \qquad 0 < \rho < 1 \qquad \varepsilon_t \sim N(0,\sigma^2) \qquad (1)$$

[1] For example in the flex-price monetary model (see e.g. Frenkel, 1976) or in sticky-price (overshooting) monetary models (see e.g. Dornbuch, 1976);
[2] See e.g. Krugman (1979) and Obstfeld (1986).

Researchers testing the existence of mean-reversion in real exchange rate series typically concluded that q_t is a random walk process (see e.g. Adler and Lehmann, 1983). However such tests were subsequently criticised on the basis of the low power of univariate unit root test (see e.g. Abuaf and Jorion, 1990); and for implicitly imposing potentially invalid restrictions (see below). As a result, research on PPP mainly focused on the investigation of the existence and the nature of a long-run relationship between nominal exchange rates and relative prices. This approach involves the application of cointegration techniques on an equation like (2):

$$s_t = \alpha + \beta_1 p_t + \beta_2 p^*_t + \varepsilon_t \qquad (2)$$

In (2), s_t is the logarithm of the nominal exchange rate, α is a constant term, p_t and p^*_t are the logarithms of the domestic and foreign price level respectively and ε_t is a white noise term. If s_t, p_t and p^*_t are integrated of order 1, any meaningful discussion on PPP pre-supposes that the residual term ε_t is stationary. If ε_t is indeed I(0), for the relative version of PPP to be valid in its so-called strong form, price changes in the two countries should exert a homogeneous and symmetric impact on the exchange rate, i.e. $\beta_1 = 1$ and $\beta_2 = -1$. In that case, equation (2) can also be written as equation (3) below:

$$s_t = \alpha + \beta \frac{p}{p^*} + \varepsilon_t, \qquad \beta = 1 \qquad (3)$$

PPP tests employing equation (1) implicitly impose the homogeneity and symmetry restrictions $\beta_1 = 1$, $\beta_2 = -1$[3]. PPP tests based on equations like (3) implicitly impose the symmetry restriction $\beta_1 = -\beta_2$. However, both restrictions are based on the assumption of zero transportation costs, identically-constructed price indexes in the home and foreign country and no measurement errors in the price series. Since these assumptions are usually not valid, the unfavourable to PPP results reached by studies using equations (1) and (3)[4] are often attributed to the implicit imposition of invalid restrictions. Taking into consideration factors like measurement errors and transactions costs, some form of PPP, the so-called weak-form PPP, is valid provided that the econometric analysis yields cointegration among the variables in equation (2), even if the estimated coefficients are not consistent with symmetry and homogeneity (see Taylor, 1988 and Cheung and Lai, 1993). PPP in its weak form only presupposes that $\beta_1 > 0$ and $\beta_2 < 0$. Most empirical studies conducted within this framework (see e.g. MacDonald, 1993, Kugler and Lenz, 1993, Cheung and Lai,1993), provide evidence suggesting that long-run PPP holds and real exchange rates are mean-reverting. However, the size of the reversion is suggested to be in the range of three to five years, far too slow to be consistent with traditional PPP analysis or the real exchange rate movements predicted by overshooting models (see e.g. Dornbuch, 1976). A number of explanations have been proposed to reconcile these findings with the PPP hypothesis. One is that the latter does not consider the effect of real shocks on the real exchange rate: it only accounts for monetary ones. However, this explanation is not satisfactory by itself given the high short-run volatility observed in the level of nominal exchange rates relative to the levels of the underlying fundamentals. Other

[3] In terms of logarithms, the real exchange rate q is given by an equation of the form $q_t = s_t + p_t - p^*_t$. If, measurement errors exist in the calculation of p_t and p^*_t, the observed price indexes are biased and the series on which unit root tests are performed are not in reality the actual real exchange rate series.
[4] See for example Corbae and Ouliaris (1988), Enders (1988), and Taylor (1988).

explanations (see Rogoff, 1996, 1999) highlight the role of frictions like transaction costs, impediments to trade, imperfect information and pricing-to-market commercial tactics. Such frictions create a buffer zone within which exchange rates respond in changes in goods prices slowly, whereas outside the band they adjust much faster. Studies taking into consideration such non-linear adjustments yield estimated speeds consistent with traditional PPP (see e.g. Michael *et al*, 1997 and O'Connell, 1998).

2. ECONOMETRIC METHODOLOGY, DATA AND RESULTS

In this section we apply Johansen's (1988) multivariate cointegration methodology to investigate the validity of the PPP hypothesis regarding the drachma/DM exchange rate for the period of floating exchange rates (1975-1998)[5]. Since the Johansen methodology is widely used in empirical research we will not present it in detail here. The DM was selected to be the reference currency in our analysis as Germany and the countries which used to participate in the ERM before the launch of euro (hence the drachma values of their currencies were determined by the level of the drachma/DM rate) are, by far, the most important trading partner of Greece. Our data source is the International Financial Statistics Databank provided by Datastream. The analysis is in terms of end-of-period spot exchange rate series[6]. The price levels used are wholesale

[5] Tests regarding the validity of the PPP for various drachma exchange rates have also been conducted by Brissimis and Leventakis (1984), Karfakis and Moschos (1989) and more recently by Dockery and Georgellis (1994), Bahmani-Oskooee (1995) and Voumvaki *et al* (1998). The results reached by these studies, which employ alternative methodologies, are mixed.

[6] The International Financial Statistics does not provide national currency/DM exchange rates whereas Datastream does not provide a data series for the drachma/DM exchange rate extending back to 1975. As a result, we had to calculate the cross-rate between the drachma and the DM using the drachma/dollar and DM/dollar series.

price indexes (line 63)[7]. Our analysis is based on quarterly data and covers the period 1975-1 to 1998-4, a total of 96 observations, covering a period of 24 years.

We first investigate the order of integration of the variables involved. The ADF tests reported in Table 1 suggest that the logarithms of the drachma/DM exchange rate and the German WPI (s_t and p^*_t respectively) are integrated of order one whereas their first differences are stationary[8]. Regarding the Greek WPI, the relevant unit root test suggests that Δp_t is an I(1) series. However, this result is rather unconvincing as the existence of a trend in the rate of growth of prices is very difficult to interpret[9]. A careful look at the graph of Δp_t in the Appendix suggests that the non-rejection of the unit root hypothesis may well be the result of structural breaks in the series. The issue of structural breaks and how to test for them is quite extended to be discussed here (see Maddala and Kim, 1998). Fortunately, as far as the Greek economy is concerned, the existence of structural monetary breaks is a well-established fact and their timing is also more or less known (see e.g. Alogoskoufis and Philippopoulos, 1992 and Alogoskoufis, 1995). If the breaks in the series are known it is relatively simple to adjust the ADF test by including dummy variables to ensure that there are as many regressors as there are deterministic components in the data generating process (Harris, 1995, p. 40). Hence, we have added to the relevant ADF equation two

[7] Wholesale price indexes are preferred to consumer price indexes on the grounds that they contain a much higher proportion of traded goods. As a result, biases due to the Balassa-Samuelson effect (according to which PPP is violated because non-traded goods are cheaper in terms of a common currency in countries with lower per-capita income) are less pronounced.

[8] The lag-length of the ADF tests is such as to ensure that the residual term in the ADF equation is a white noise process. Given that the analysis which follows is based on quarterly data, strictly speaking, one should also test for the existence of seasonal unit roots. However, it is well known that seasonal unit roots are not encountered very often in macroeconomic series which can typically be described as $I(1)$ with a deterministic seasonal pattern superimposed (Osborn, 1993, p. 300). As a result, and due to space limitations, seasonal unit roots tests are not presented here.

[9] For example, a non-stationary inflation rate would imply a non-stationary real interest rate. It would be very difficult to sustain such a conclusion as a steady-state equilibrium situation, unless one makes the (also highly unlikely) assumption that the nominal interest rate is also non-stationary.

dummies: the first, takes the value of 1 for the period 1979-1 to 1985-1 and 0 otherwise, capturing the effect of the particularly expansionary monetary policies followed by Greece during that period. The second, takes the value of 1 for the period following 1993-1 to 1998-4 and 0 otherwise, capturing the effect of the change in the policy regime towards price stabilisation. The adjusted ADF equation suggests that the Δp_t series is an I(0) process with structural breaks.

We then test for the number of long-run relationships using Johansen's (1988) cointegration methodology. To that end, we estimate a number of VAR models (including a constant restricted to the cointegration space and unrestricted, deterministic, centred seasonal dummies)[10], making use of different lag structures. Table 2, panel A presents the values of the Schwarz, Hannan-Quinn and Akaike information criteria associated with each lag structure, together with a number of system mispecification tests. Two out of the three information criteria suggest that a VAR structure involving two lags is the most appropriate specification whereas the third suggests a lag-length of 1. For the latter, all diagnostic tests suggest system mispecification. Hence, we select a VAR structure with two lags. However, for this lag order (as for all lag structures in Table 2, panel A), there is a problem of residuals' non-normality. To acquire Gaussian residuals, we investigated the statistical significance of dummy variables corresponding to one-off exogenous monetary shocks in Greece and in Germany. Out of these, five incidents proved statistically significant: the drachma devaluation of January 1983 (D1983-1); the drachma devaluation of March 1998 (D1998-1); and three major jumps in the DM/US dollar

[10] Elimination of the constant term from the cointegration space would be equivalent to testing the validity of only the absolute version of PPP. However, the cointegration results which follow are not affected in case the constant is not restricted to the cointegration space. Seasonal dummies are centred

exchange rate (D1980-2, D1985-3, D1985-4). Adding these dummies as unrestricted variables (not entering the cointegration space) yields a well-specified system and individual equations with Gaussian errors (see Table 2, panel B)[11].

Panel A in Table 3 presents the results of the cointegration tests. Both the rank (λ_{max}) and the trace (λ_{trace}) statistics suggest the existence of one cointegrating vector. Table 3, panel B sheds some additional light regarding its nature. First, three LR-statistics are presented testing whether each of the three variables enter the cointegration space. All of them appear to do so. The fact that s_t enters the cointegrating vector confirms, combined with a statistically significant constant term, that some form of relative PPP is valid. We then present the results of the weak exogeneity LR tests proposed by Johansen and Juselius (1992). These consist of testing zero restrictions on the elements of the α matrix (i.e. the matrix of coefficients of the speed of adjustment to long-run equilibrium) embedded in the estimated Vector Error Correction Model (VECM). The results suggest that only p^*_t is weakly exogenous to the system. This confirms that PPP is not a model of exchange rate determination *per se*. It is a long-run equilibrium condition between two endogenous variables, s_t and p_t where long-run causality runs both ways. In addition, Greece's own price level changes does not determine in any way foreign inflation (a reasonable conclusion given that Greece is a

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to ensure that they sum to zero over time and thus they do not affect the underlying asymptotic distributions upon which the tests for cointegration rank depend (see Harris, 1995, p. 81).

[11] Acquiring Gaussian errors by means of adding dummy variables other than the centred seasonal ones may sometimes by costly as the extra dummies will affect the underlying distribution of cointegrating rank statistics. In that case the power of the cointegration tests is reduced and the published critical values will only be indicative (see Harris 1995, p. 81). However, in our case there are two factors which make us believe that our results are robust. First, even without these dummies, the nature of the results of the cointegration analysis which follows remain the same (the results of the cointegration tests without the dummies are not included here due to space limitations but are available by the author upon request). Second, even when the dummies are included in the system, the existence of one cointegrating vector is statistically significant at the 1% level; whereas the values of the trace and maximal eigenvalue testing for different cointegration ranks are way apart the critical values (see Table 3, panel A).

small country). After imposing the identified valid restriction $\alpha_{p*} = 0$, we arrive at an estimated cointegrating vector involving coefficients significantly different to those implied by the proportionality and symmetry restrictions (Table 3, panel C). Finally, given that p^*_t is weakly exogenous, one can safely assume that in terms of a Vector Error Correction Model (VECM), the equation for Δp^*_t contains no information about the long-run cointegration vector since the cointegration relationships do not enter into this equation. It is therefore valid to condition on p^*_t and proceed with a partial version of the VECM where only the equations for Δs_t and Δp_t are estimated (see Harris, 1995 p. 98). The estimated VECM equations can then be considered as the processes describing short-run dynamics[12].

4. DISCUSSION OF ECONOMETRIC RESULTS

Three main results emerge from our econometric analysis: First, PPP is a long-run equilibrium condition regarding the drachma/DM exchange rate from which deviations cannot be sustained in the long-run, unless real (demand or supply) shocks are taking place to justify them. In the case of Greece, relative PPP was violated for a ten-year period (1988-1997), a time-horizon which can be safely assumed to be prolonged. At the same time, there exists evidence (see Arghyrou, 2000), suggesting that Greece's *relative* competitiveness *vis-à-vis* its main trading partners was declining during the same period, i.e. no relative positive supply shocks were taking place to justify the real appreciation of the drachma. In this framework, the ultimate

[12] Due to space limitations, the estimated VECM equations are not presented here. However, it is necessary to mention that both the equation for Δs_t and Δp_t are well specified (none of the diagnostic tests suggests any form of mispecification). The VECM includes one lag for Δs_t, Δp_t and Δp^*_t (the latter however is not modelled since p^*_t is weakly exogenous to the system) and the (restricted) cointegrating vector estimated in Table 3, panel C, lagged for one period. It also includes three seasonal

collapse of the peg of the drachma to the DM was predictable as it was inconsistent with the underlying macroeconomic fundamentals, i.e. declining but yet existent inflation and productivity differentials *vis-à-vis* the EU average and a widening current account deficit caused by drachma's overvaluation[13].

Our second result is that the strong drachma helped Greece reduce its inflation. A strong-currency policy is by definition a policy where the level of the exchange rate acquires the character of an exogenous variable. This basically implies that in our cointegration vector, the only endogenous variable during the period of the application of the strong-drachma policy was the domestic price level. Normalising the cointegrating vector on p_t yields a rather significant coefficient for the level of the exchange rate (0.77). The restraining of the growth of the level of the exchange rate imposed by the strong drachma policy, in combination with the reduction in the other exogenous variable of the system, (German/EU inflation) can largely explain the significant reduction of Greek inflation between 1993 and 1997.

Our third and final result is that foreign exchange markets behaved in a way consistent with the hypothesis that they started anticipating a change in the parity of the drachma as early as mid-1997; but the exact timing of the devaluation took them by surprise. We base this argument on two tests. The first is the forecasts provided by the estimation of the VECM equations mentioned in section 3. Estimating the VECM[14] for the period 1975-2 to 1997-4 yields the forecast band for the change in

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dummies and the five unrestricted dummies earlier mentioned. These were added to ensure Gaussian error terms. Their exclusion does not affect the nature of the results.

[13] The current account deficit (in terms of % in Greek GDP) has as follows: 1994: 1%; 1995: 2.7%; 1996: 3.5%; 1997: 3.7%. Source: European Economy No 64, Statistical Appendix.

[14] The VECM system from which the forecasts presented in Figures 2 and 3 emerge is estimated on the basis of the (restricted) cointegration vector resulting from a cointegration analysis covering the period

the drachma's value against the DM which appears in Figure 2. The actual outcome in 1998-1 lies outside the band. This suggests that the devaluation, when it occurred, was "news" to the market. However, going a little back in time reveals a similar but slightly different picture. Estimating the VECM for the period 1975-2 to 1997-1 yields the set of forecasts which appear in Figure 3. The actual outcome of 1998-1 is once again outside the forecast band. However, one notices that the position of the forecast bands at the beginning of 1997 presents an upwards trend for the period 1997-2 to 1998-1. Given that the VECM is estimated using data which in 1997-1 was public information, one might interpret the shifting of the forecast band in a way suggesting that early in 1997 markets foresaw the possibility of an upwards adjustment in the value of the drachma in the following year or so. This conclusion is further supported by our second test regarding the credibility of the drachma's peg to the DM. This is a simple test suggested by De Grauwe (1997) and is essentially based on a comparison between the value of the exchange rate target set by a central bank (or the upper limit of a target zone) and the implicit forward rate as the latter is derived by the covered interest parity condition (CIP)[15]. The idea behind the test is quite simple: as the forward rate reveals the markets' risk-adjusted expectation

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1975-1/1997-4. The reason is that if we had estimated the VECM based on cointegration analysis covering the period 1975-1/1998-4, the forecast for 1998-1 would have been a function of the outcomes observed in 1998, as the lagged cointegrating vector entering the VECM would have been estimated using the whole sample, 1975-1/1998-4. Space limitations do not allow us to present the results of the cointegration analysis covering the period 1975-1/1997-4 here. We restrict ourselves in saying that they are identical to the ones presented in Table 3. The (reduced-form) cointegrating vector is given by $s_t = 4.9261 + 1.2903\,p - 2.6873\,p^*$. Finally, the estimated VECM from which Figures 2 and 3 emerge are conditioned on Δp^*_t (given that p^* again appears to be weakly exogenous to the system) and do not present any kind of mispecification.

[15] There seems to be little doubt that CIP condition holds as a long-run relationship (for a survey of the evidence, see Gibson, 1996 and the evidence therein). CIP states that $F_{t,T} = S_t \dfrac{1+r_{t,T}}{1+r^*_{t,T}}$, where $F_{t,T}$ is the forward rate prevailing at time t for delivery at time T, S_t is the spot exchange rate, and $r_{t,T}$ and $r^*_{t,T}$ the domestic and foreign effective rate of return between time t and T respectively. For the purpose of the test which we employ it is preferable to use the implicit rate (i.e. calculate the value of $F_{t,T}$ based on the observed variables of S_t, $r_{t,T}$ and $r^*_{t,T}$) rather than use the observed forward market for two reasons. First, because we neutralise the effect of factors which might lead to short-run deviations

regarding the future spot exchange rate, if it exceeds the upper limit of a target-zone, there exist indications suggesting that market participants expect (in a risk-adjusted sense) that the future spot rate will not be within the announced target. In short, it is an indication that the target zone is not perceived to be credible[16]. Figure 4 presents, on a monthly basis, a comparison between the implied one-year drachma/DM forward rate (calculated using end-of-period spot exchange rates and end-of-period annualised overnight interbank interest rates[17]) and the spot exchange rate for the period for which the Bank of Greece announced official exchange rate targets (1995 to 1998)[18]. The scale is normalised to unity in order to facilitate inference. Various interesting points emerge. First, the markets never perceived the official zone-targets as credible as the forward premium had always been higher than 5% above the spot rate. Second, the credibility of the drachma's peg to the DM improved considerably between January 1995 and May 1997 with the implied forward premium declining from 12% to 6%. Third, the credibility of the peg stops improving as early as June 1997 (almost one year before the devaluation), before the onset of the South-East Asia crisis. When the latter broke, a speculative attack against the drachma took place (in November

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from CIP, e.g. short-lived political uncertainty. Second, because in this way we correct the effects of data imperfections.

[16] Implicitly, the test assumes that foreign exchange markets are efficient. This implies that the forward rate equals the market's expectation regarding the future spot rate plus a risk premium ($F_{t,T} = S^e_t(T) + \rho_t$). If the forward rate $F_{t,T}$ exceeds the upper bound of a target zone S_{up} there exist two possibilities: either the expected future spot rate is higher than the upper limit of the zone ($S^e_t(T) > S_{up}$); or the perceived risk associated with the value of the future spot rate is too high. In both cases, the peg suffers from a credibility problem.

[17] Overnight interbank interest rates are widely viewed as the best indicators of the monetary conditions prevailing at a given point in time. For Greece, we use the annualised end-of-month interest rate observed in the overnight interbank money market. For Germany we use the annualised end-of-month call-money rate. The data source for Greek rates is the Economic Bulletin of the Bank of Greece (various editions). The data source for the German rate is Datastream (code GERMDRD).

[18] The exchange rate targets set by the Bank of Greece have as follows: For 1995: depreciation not higher than 3% against the ECU (Bank of Greece, Report of the Governor for 1994, p. 48). For 1996: stable exchange rate against the ECU (Bank of Greece, Report of the Governor for 1995, p. 49); For 1997: stable exchange rate against the currencies which constitute the ECU (Bank of Greece, Report of the Governor for 1996, p. 37); for 1998; exchange rate fluctuations which, on average, will not deviate from the central parities of the drachma against the currencies which constitute the ECU by ±2.5% (Bank of Greece, Report of the Governor for 1997, p. 54).

1997). On that occasion, the Bank of Greece managed to beat the speculators by increasing significantly the overnight interbank money market rate (reflected in the peak reported in Figure 4) and proceeding to foreign exchange interventions. However, as international monetary conditions started returning to normality (early 1998), we observe that despite continuous reassurances at the highest level that there exists no imminent change in the parity of the drachma, the drachma/DM forward premium takes values well above those of early 1997. All in all, Figure 4 is consistent with a scenario according to which markets lost confidence to the drachmas' peg to the DM in mid-1997, a fact which became more evident after the South-East Asia crisis. The realisation of this fact prompted the Bank of Greece to devalue the drachma at an unexpected point in time (three months after successfully repelling a speculative attack) and, from that point of view, take the markets by surprise.

5. SUMMARY AND CONCLUDING REMARKS

This paper has examined whether the collapse of an exchange rate peg is predictable on the basis of the movements of the underlying macroeconomic fundamentals and the observed behaviour of currency markets. Our verdict is a conditional yes. Our results suggest that the *ultimate* collapse of the peg is predictable if the latter is clearly inconsistent with valid long-run macroeconomic relationships. Furthermore, concerning the case study examined, our findings indicate that currency markets behaved in a way suggesting that they anticipated a collapse of the peg well before it took place. However, the *exact timing* of the devaluation took the markets by surprise. There emerge three main messages of wider interest from these results. First, fundamentals matter in long-run exchange rate determination. Second, markets acknowledge this fact and act accordingly. And third, even under conditions of devaluation expectations, Central Banks can still surprise the markets, provided they choose the correct timing.

Regarding the particular case of Greece, our findings suggest that the strong-drachma policy was helpful in terms of achieving inflation convergence but not inflation equalization with Germany. As economic theory suggests (see, among others, De Grauwe, 1996 and Persson and Tabellini, 1996), and as it actually happened in countries where the authorities maintained an overvalued exchange rate for too long (e.g. in Italy and Spain in 1992-93 and Mexico in 1994), eventually, Greece had to abandon the peg[19]. From that point of view, the Greek devaluation of 1998 involves a certain *deja-vu* element. The difference is that in the case of the countries mentioned above, the markets forced the devaluation. In the case of Greece, the central bank had learned from the others' experiences and moved before the speculators did.

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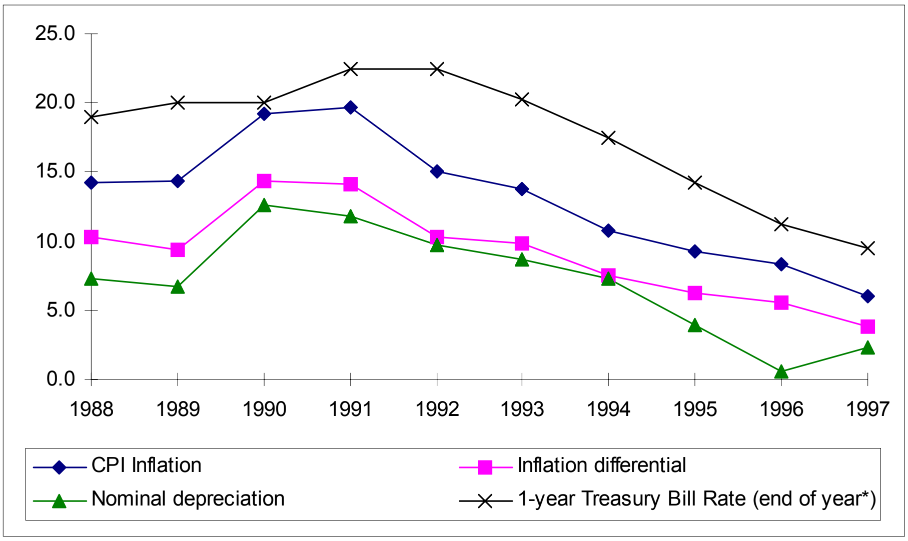
Figure 1



Annual CPI inflation, inflation differential against the EU and nominal depreciation against the ECU

Source: European Economy No 64, Statistical Appendix and Bank of Greece, Economic Bulletin, various editions

Table 1: Unit root tests

<u>*Equation*</u>	$\Delta y_t = \psi^* y_{t-1} + \sum_{i=1}^{p} \psi^*_i \Delta_{t-1} + \mu + \gamma t + u_t$ [1]		$\Delta y_t = \psi^* y_{t-1} + \sum_{i=1}^{p} \psi^*_i \Delta_{t-1} + \mu_t + u_t$ [2]	
	s_t	ADF(0) = 0.114	Δs_t	ADF(2) = -2.644*
	p_t	ADF(5) = - 0.334	Δp_t	ADF(0) = -3.452*
	p^*_t	ADF(5) = -1.067	Δp^*_t	ADF(3)=-1.409
			Δp^{*+}_t	ADF(6) = -5.741**

Order of ADF lag structure in parenthesis; [1]: Critical values are –3.46 at the5% level; -4.062 at the 1% level ; [2]: Critical values are –1.94 at the 5% level; -2.589 at the 1% level; * rejects the null hypothesis at the 5% level , ** rejects the null hypothesis at the 1% level ; [+] ADF equation estimated including two dummy variables (see text) ; The lag order of the ADF equations was selected to ensure that u_t presents no autocorrelation

Table 2: Determination of VAR specification

PANEL A: DETERMINATION OF VAR ORDER

VAR estimated: $\mathbf{z_t} = \mathbf{A_1}\,\mathbf{z_{t-1}} + \ldots + \mathbf{A_k}\,\mathbf{z_{t-k}} + \mathbf{\Psi}\,\mathbf{D_t}$

$\mathbf{z'} = [constant\ \ s_t\ p_t\ p^*_t\ ]$, $\mathbf{D'_t} = [D_{1t}\ \ D_{2t}\ \ D_{3t}]$

	k=5	*k= 4*	*k=3*	*k=2*	*k=1*
Information selection criteria					
Schwarz	-29.187	-29.378	-29.657	-30.022	-29.617
Hannan-Quinn	-30.125	-30.168	-30.299	-30.516	-29.962
Akaike	-31.012	-30.758	-31.591	-31.509	-30.658
System mispecification tests					
Vector $AR_{autocorrelation}$	0.7352	1.063	1.205	1.228	2.8848**
Vector Normality χ^2	23.899**	27.466**	26.685**	24.307**	18.982**
Vector $\chi^2_{\text{heteroscedasticity}}$	0.8337	0.88255	0.8824	1.0374	1.8423**

PANEL B: DETERMINATION OF SYSTEM SPECIFICATION

VAR estimated: $\mathbf{z_t} = \mathbf{A_1}\,\mathbf{z_{t-1}} + \mathbf{A_2}\,\mathbf{z_{t-2}} + \mathbf{\Psi}\,\mathbf{D_t}$

$\mathbf{z'} = [\ constant\ \ s_t\ p_t\ p^*_t\ ]$, $\mathbf{D'_t} = [D_{1t}\ \ D_{2t}\ \ D_{3t}\ \ D_{1983\text{-}1}\ \ D_{1998\text{-}1}\ D_{1980\text{-}2}\ D_{1985\text{-}3}\ D_{1985\text{-}4}]$

Individual equations mispecification tests	
S_t: AR $_{\text{autocorrelation}}$ – F-test	0.5926
P_t: AR $_{\text{autocorrelation}}$ – F-test	2.1384
P^*_t: AR $_{\text{autocorrelation}}$ – F-test	1.2678
S_t: Normality χ^2	1.3697
P_t: Normality χ^2	4.9501
P^*_t: Normality χ^2	1.2196
S_t: ARCH F-test	0.7955
P_t: ARCH F-test	1.5646
P^*_t: ARCH F-test	0.4460
S_t: $\chi^2_{\text{heteroscedasticity}}$	1.0465
P_t: $\chi^2_{\text{heteroscedasticity}}$	1.0882
P^*_t: $\chi^2_{\text{heteroscedasticity}}$	1.2321
S_t: $\chi_i\chi_{j\ \text{auteroscedasticity}}$	0.7455
P_t: $\chi_i\chi_{j\ \text{auteroscedasticity}}$	0.5839
P^*_t: $\chi_i\chi_{j\ \text{auteroscedasticity}}$	1.0224

System mispecification tests	
Vector $AR_{autocorrelation}$	1.1235
Vector Normality χ^2	7.9276
Vector $\chi^2_{\text{heteroscedasticity}}$	0.8338
Vector $\chi_i{}^*\chi_{j\ \text{heteeroscedasticity}}$	0.6364

* Rejects the null hypothesis at the 5% level , ** Rejects the null hypothesis at the 1% level

Table 3: Cointegration analysis

PANEL A: DETERMINATION OF COINTEGRATION RANK

H_0	H_1	λ_{max} statistic	95% CV	λ_{rrace} statistic	95% CV
r = 0	r = 1	70.02**	22.0	81.51**	34.9
r ≤ 1	r = 2	10.38	15.7	11.49	20.0
r ≤ 2	r = 3	1.11	9.2	1.11	9.2

Standardized beta′ eigenvectors

	s_t	P_t	p^*_t	*constant*
	1.0000	-1.3090	2.7509	-5.0008
	-0.9469	1.0000	-1.9275	3.9060
	-0.1765	-0.0031	1.0000	-1.6151

Standardized alpha coefficients

S_t	0.093713	0.015330	0.042556
P_t	0.096472	0.0010448	-0.011680
P^*_t	0.0065151	0.013266	-0.0047130

PANEL B

(a) Testing cointegration restrictions in beta′ vector $\beta_0 s_t + \beta_1 p_t + \beta_2 p^*_t$ *+ constant* (cointegrating rank =1)

Restriction	*LR-test* $\chi^2(1)$
$H_0: \beta_0 = 0$ (s_t does not enter the cointegration space)	11.62**
$H_0: \beta_1 = 0$ (p_t does not enter the cointegration space)	15.551**
$H_0: \beta_2 = 0$ (p^*_t does not enter the cointegration space)	19.394**
H_0: *constant* = 0 (constant does not enter the cointegration space)	17.755**

(b) Testing cointegration restrictions in alpha matrix (cointegrating rank =1)

	LR-test $\chi^2(1)$
$H_0: \alpha_s = 0$ (s_t weakly exogenous to the system)	19.745**
$H_0: \alpha_p = 0$ (p_t weakly exogenous to the system)	58.588**
$H_0: \alpha_{p*} = 0$ (p^*_t weakly exogenous to the system)	2.761

PANEL C: RESTRICTED COINTEGRATION ANALYSIS (cointegrating rank =1 and $\alpha_{p*} = 0$)

Beta′ matrix	s_t	p_t	p^*_t	*constant*
	0.15332	-0.19896	0.41925	-0.76566

Standardized beta′ eigenvectors

$$s_t = 4.9937 + 1.2976\, p_t - 2.7344\, p^*_t$$

$$p_t = -3.843 + 0.7706\, p_t + 2.1072\, p^*_t$$

Alpha matrix

(std. errors in parenthesis)		
	s_t	0.68893 (0.14409)
	p_t	0.61819 (0.07153)
	p^*_t	0.0000 (0.0000)

* Rejects the null hypothesis at the 5% level, ** Rejects the null hypothesis at the 1% level

Figure 2

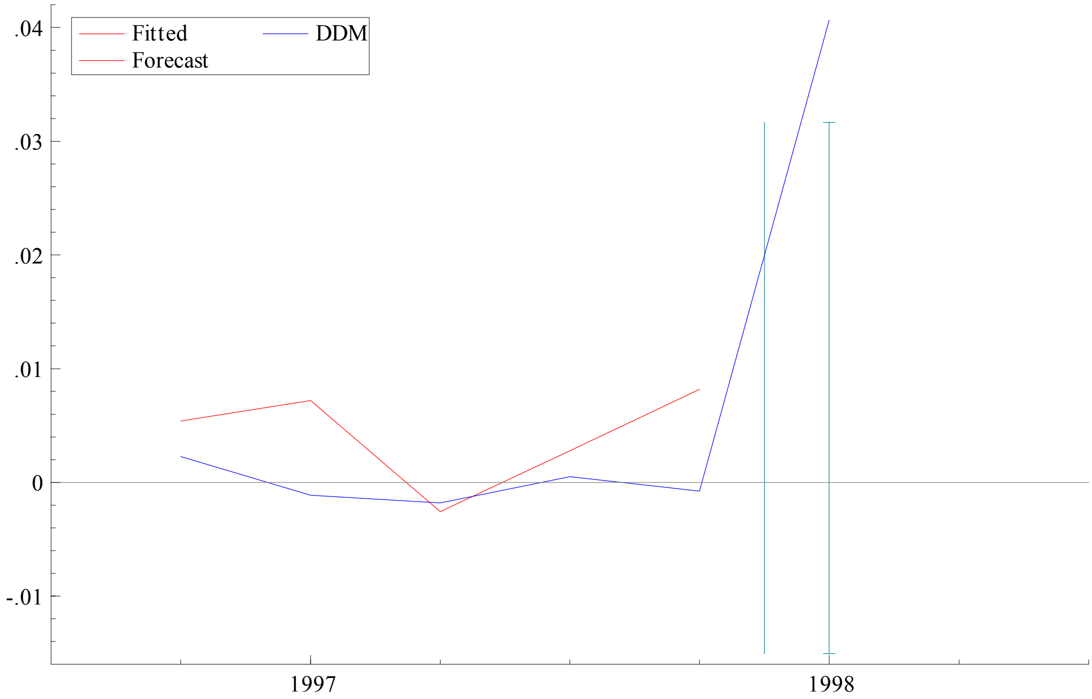


Figure 3

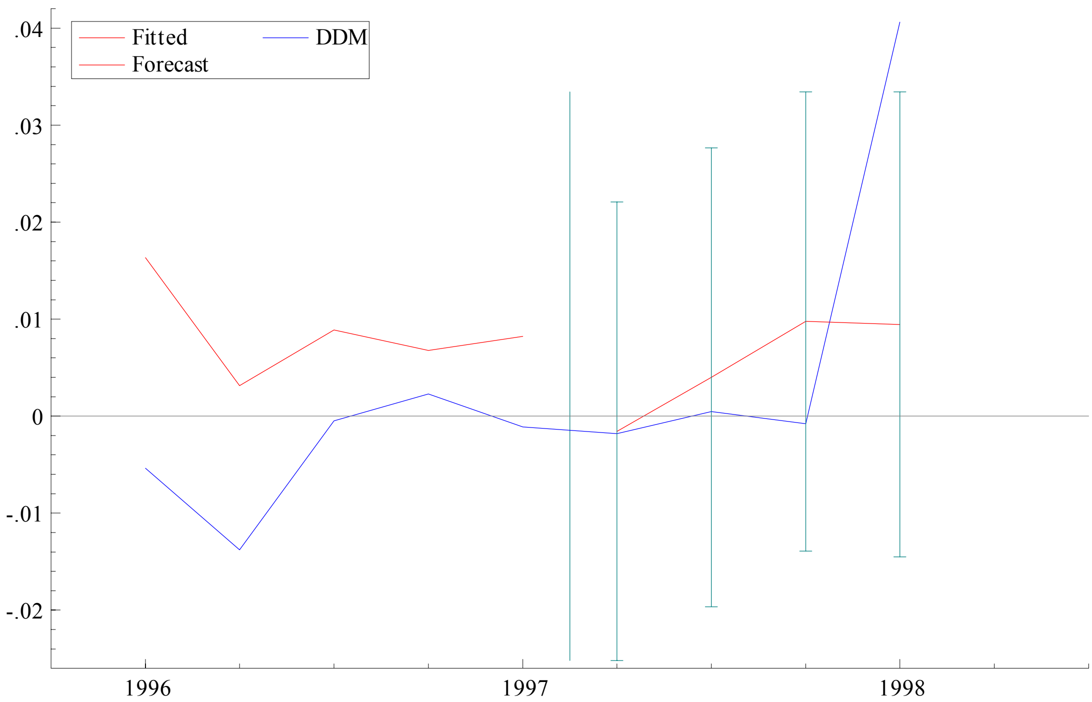
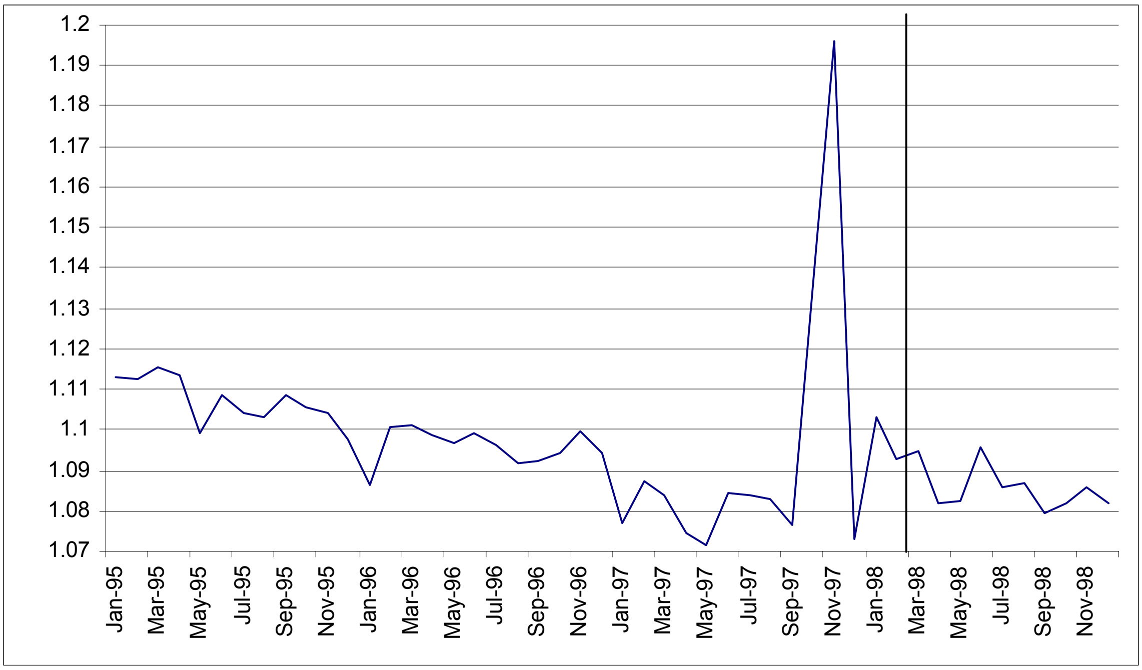


Figure 4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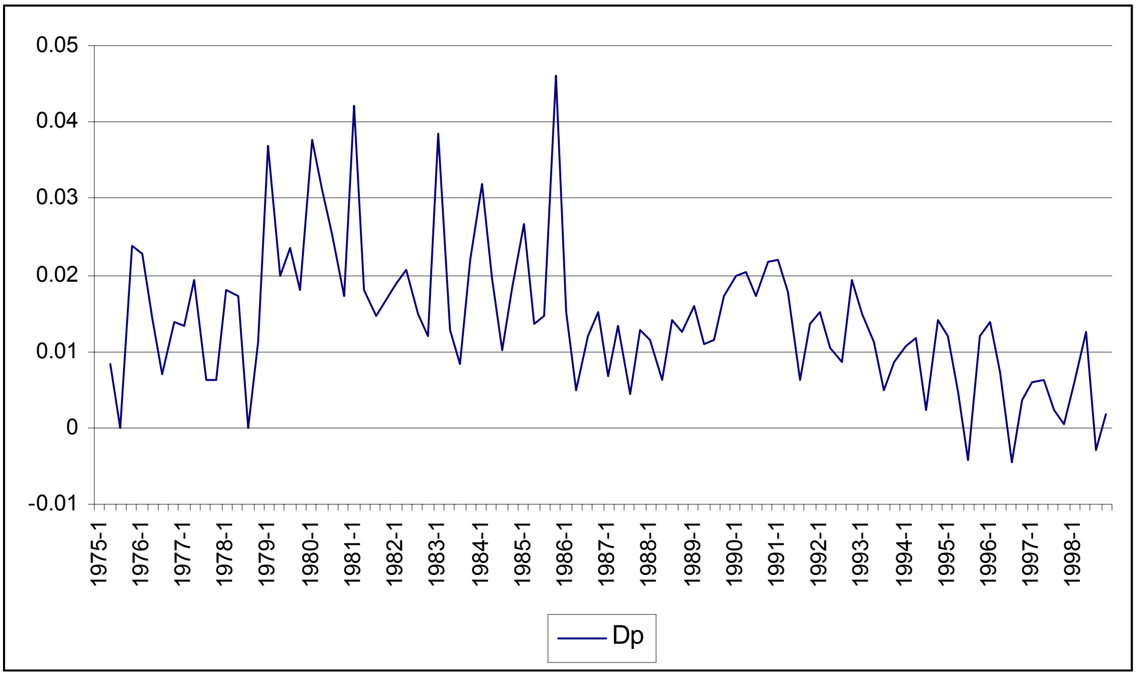
APPENDIX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