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A MODEL OF EXCHANGE RATE POLICY: EVIDENCE FOR THE US DOLLAR-GREEK DRACHMA RATE 1975-1987

by

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CONTENTS

1. INTRODUCTION

The aim of this paper is to model movements of the exchange rate between the US dollar and Greek drachma over the period 1975.I-1987.IV during which a managed float regime was adopted. A sticky price monetary model which allows for one period wage contracts and an exchange rate reaction function of monetary authorities is set up, and then a reduced form error correction (EC) specification is derived that relates the change in the exchange rate to changes in the lagged exchange rate, the price differential, the real income, the interest rate, the real money supply and the real wage as well as the lagged deviation from the Purchasing Power Parity (PPP).

The motivation for the paper derives from two recent contributions to the literature: the first (Meese and Rogoff, 1983) provides evidence that structural models of exchange rate behaviour could not outperform a simple random walk model. The second contribution (Boothe and Glassman, 1987) which follows the lead of Meese and Singleton (1982) views the empirical problems of exchange rate theories in the context of spurious regressions (Granger and Newbold, 1974; Phillips, 1986) associated with the nonstationarity of economic time series. The EC model (Hendry et al., 1984) can be regarded as a way to alleviate the spurious regression problem without discarding valuable information by retaining the equilibrium constraints postulated from economic theory.

In Section 2 the theoretical model is presented. Section 3 presents the unit root and cointegration tests. In Section 4 the reduced-form EC model is estimated and the results are discussed. Concluding remarks are given in Section 5.

2.A. THE THEORETICAL MODEL

Consider first the long run PPP which relates the exchange rate to the difference between the domestic and foreign price levels:[1]

$$s(t)=\alpha+\beta(p(t)-fp(t)) \tag{1}$$

and then the deviations u(t) from PPP:

$$s(t)=\bar{s}(t)+u(t) \tag{2}$$

where s(t) ($\bar{s}$(t)) and p(t) (fp(t)) denote the log of the actual (equilibrium) exchange rate defined as domestic currency per foreign currency unit and the log of the Greek (US) consumer price level respectively.

The log of the long run real money demand (md(t)-p(t)) is specified as a function of the log of the real income y(t) and the interest rate r(t) on an alternative asset:

$$md(t)-p(t)=c+\gamma y(t)-\delta r(t)+v(t) \tag{3}$$

where γ, δ >0, and v(t)~IN(0,σ^2)

The log of the nominal money supply is assumed to follow a random walk with a drift:

$$m(t)=c+m(t-1)+\xi(t) \tag{4}$$

where ξ(t)~IN(0,σ^2).

Given (4), equation (3) determines the price level $\bar{p}$(t) that is consistent with money market equilibrium.

The third building block of the theory links the log of the nominal wage w(t) with the expected price level at time t (E(t-1)[p(t)]) through one period wage contracts which are set at the beginning of the period t with the aim of maintaining a constant real wage (Fisher, 1977):

$$w(t)=k+E(t-1)[p(t)] \tag{5}$$

where k denotes the target real wage. It is assumed that the expectations formation mechanism is rational:

$$p(t)=E(t-1)[p(t)\mid\Omega(t-1)]+\zeta(t) \tag{6}$$

with ζ(t)~(0,σ^2); Ω(t-1) denotes the information set.

Owing to the existence of wage contracts, it is assumed that prices adjust slowly to equilibrium value according to:

$$p(t)-p(t-1)=\theta_1(\bar{p}(t)-p(t-1)) \tag{7}$$

with $0\leq\theta_1\leq1$. The market clearing assumption corresponds to the value of $\theta_1=1$.

It is assumed that the monetary authorities pursue the following short run exchange rate rule:

$$s(t)-s(t-1)=\psi_0(\overline{s}(t)-s(t-1))+\psi_1(w(t)-w(t-1))-\psi_2(\overline{D}(t)-D(t-1)) \tag{8}$$

with $0\leq\psi_0\leq1$, $\psi_1>0$ or $\psi_1<0$ and $\psi_2>0$.

Equation (8) denotes that deviations from PPP, wage inflation and deviations of the external debt D(t) from the target $\bar{D}(t)$ determine the stance of the exchange rate policy. If the sign of the parameter ψ_1 is positive (negative) means that the monetary authorities accommodate (stabilize) movements in the wage inflation by depreciating (appreciating) the exchange rate.

Finally, it is assumed that the external debt at the end of the period t-1 is defined as:

$$D(t-1)=\lambda\sum_{i=-\infty}^{t-1}u(t-1) \tag{9}$$

with $\lambda<0$, and

$$\Delta D(t-1)=\lambda u(t-1) \tag{10}$$

where Δ denotes the first difference operator. Formulation (9) shows that the external debt negatively depends on all past accumulated equilibrium errors. In other words, equation (10) denotes that the external debt falls, if the exchange rate stands above the target rate.

2.B. DERIVING A REDUCED-FORM EQUATION

First we solve for p(t) which clears the money market and then combine the resultant expression with equation (7) to get:

$$p(t)=\theta_1(\theta_0-c)+(1-\theta_1)p(t-1)+\theta_1m(t-1)-\gamma\theta_1y(t)+\delta\theta_1r(t)+\theta_1(\xi(t)-v(t)) \tag{11}$$

Equations (5), (6) and (11) put together give:

$$w(t)=(\theta_1\theta_0+k-\theta_1c)+(1-\theta_1)p(t-1)+\theta_1m(t-1)-\gamma\theta_1y(t)+\delta\theta_1r(t)+\theta_1(\xi(t)-v(t))-\zeta(t) \tag{12}$$

Finally, we substitute equations (1), (9) and (12) in equation (8) and take the first difference of the resultant form we have:

$$\Delta s(t)=\varphi_0\Delta s(t-1)+\varphi_1\Delta(p(t)-fp(t))+\varphi_2\Delta y(t)+\varphi_3\Delta r(t)+\varphi_4\Delta(m(t-1)-p(t-1))+\varphi_5\Delta(w(t-1)-p(t-1))+\varphi_6u(t-1)+\omega(t) \tag{13}$$

where $\phi_0=1-\psi_0$, $\phi_1=\beta\psi_0$, $\phi_2=-\psi_1\theta_1\gamma$, $\phi_3=\psi_1\theta_1\delta$, $\phi_4=\psi_1\theta_1$, $\phi_5=-\psi_1$, $\phi_6=\psi_2\lambda$, $\omega=\Delta\psi_1[\theta_1(\xi-v)-\zeta]$.

Observing (13), it can be seen that an EC specification has been derived from the structural model. If the rates of change are set equal to zero, then the long run solution u=0 is theoretically consistent with PPP. Since there is no unique way of estimating the structural parameter ψ_2 but only the reduced-form coefficient $\phi_6=\lambda\psi_2$, a normalization on $\lambda=-1$ could be undertaken.

3. TESTS OF UNIT ROOTS AND COINTEGRATION

Before estimating equation (13), the univariate time-series properties of all the variables concerned are established.

Under the null hypothesis of a unit root, the time-series are not stationary and the classical t-test and F-test are inappropriate because the limiting distribution of the asymptotic variance of the parameter estimates is not finitely defined (Fuller, 1985). Furthermore, Phillips (1986) has shown that these tests actually diverge (grow) as the sample size increases. To test whether a nondeterministic series contains a unit root, i.e., it is integrated of order one (I(1)), the test developed by Dickey and Fuller (1981) can be used. The technique is to estimate the regression

$$\Delta x(t)=-ax(t-1)+b+e(t) \tag{14}$$

with $e(t)\sim IN(0,\sigma_e^2)$, and then test the hypothesis H_o:a=0 by comparing the calculated t ratio of $\hat{a}$ with the reported percentiles in Fuller (1976, p.373). This is known as the Dickey-Fuller (DF) test. If however, e(t) is not approximately white noise, Dickey and Fuller (1981) have suggested to enlarge model (14) by adding in a lag polynomial in $\Delta x(t)$ as a means of removing serial correlation. The reported percentiles in Fuller (1976) can also be used for the t ratio of $\hat{a}$ in the augmented regression (ADF test). Recently Phillips (1987) and Phillips and Perron (1988) proposed nonparametric tests which allow for weakly dependent and heterogenous disturbances.

Two series which are integrated of the same order μ are said to be cointegrated of order μ,η-denoted as

CI(μ,η)-if there exists a linear combination of the two which is I($\mu-\eta$), $\eta>0$ (Engle and Granger 1987). If two variables are CI(1,1), it implies that they do not drift away from each other over the long run.

In order to test for cointegration between s(t) and p(t)-fp(t) the testing procedure suggested by Engle and Granger (1987) is followed. If it is established that both series are I(1), the regression

$$s(t)=\alpha+\beta(p(t)-fp(t))+u(t) \tag{15}$$

is estimated, and then the hypothesis H_o:u(t)~I(1) against the alternative H_1:u(t)~I(0) is tested by using the DF test. An alternative but low power test is the cointegrating regression DW statistic (CIRDW). Engle and Yoo (1987) have derived the critical values of the empirical distributions of DF and CIRDW for different numbers of variables in the cointegrating equation.

Table 1 reports the results of the unit root tests for the levels and the first differences of the series concerned.[2] The hypothesis of nonstationarity cannot be rejected at the 0.05 significance level for any of the levels of the variables with the exception of the real wage. By contrast, the hypothesis of a unit root is rejected for the first differences of all the series. These results lend support to Nelson and Plosser's (1982) and Meese and Singleton's (1982) conjecture that the process generating the log of exchange rates and other macroeconomic variables is well approximated by random walks. The fundamental implication of these findings with respect to the time-series modelling of the exchange rate is that first difference transformation of s is required to induce stationarity. Otherwise, by ignoring the problem, one can obtain spurious outcomes (Granger and Newbold, 1974; Phillips, 1986). For example, when conducting statistical inference in theoretical rational expectations models which rely on a stationarity assumption to derive time invariant representations of the exchange rates in terms of conditional expectations of the explanatory variables (Bilson, 1978), one may obtain spurious results when the variables are not stationary unless they are cointegrated (Engle and Granger, 1987).

The DF and CIRDW tests reported in table 2 show that the nominal exchange rate and the price differential are not cointegrated. Thus there seems to be no evidence of a long run equilibrium relationship between s(t) and p(t)-fp(t). This means that shocks which affect the discrepancy between domestic and foreign prices are not reflected in the nominal exchange rate in a similar way, implying that the dollar real exchange rate of the drachma fluctuates widely with no tendency to return to a predetermined path.

The value of CIRDW statistic implies that the first order residual autocorrelation is of the order of 0.91. Given the low power of CIRDW and DF tests to reject the null hypothesis of a unit root against alternatives close to the unit circle at such levels (Engle and Granger, 1987), an alternative route to modelling and testing the

long run relationship in the context of a dynamic model is used.

4. ESTIMATION, THEORY CONSISTENCY AND DATA COHERENCE

The reduced-form EC model (13) is estimated over the period 1975.I-1987.IV using the instrumental variable (IV) estimator to tackle the endogeneity problem associated with the correlation between $\Delta(p(t)-fp(t))$ and the error term $\omega(t)$. The results are reported in table 3. The value of the structural parameter ψ_0 is equal to one which implying that, other things being constant, the monetary authorities set the exchange rate equal to the target rate which is not however consistent with PPP. The value of the structural parameter $\psi_1 (= -0.9973)$ is significantly different from zero with a negative sign implying that the authorities appreciate the domestic currency when there is wage inflation. In other words, Greek policymakers faced with double-digit inflation rates over the sample pursued a short run antiinflationary exchange rate policy that appreciates the exchange rate in the presence of wage inflation as an attempt to mitigate the depreciating pressures on the domestic currency and thus to ease the burden of adjustment on the Greek producers. Furthermore, the value of the parameter ψ_2 is insignificantly different from zero, given the postulated value of $\lambda=-1$. The rate of change in the domestic real income has a significant positive impact on the rate of change in the exchange rate. In other words, an increase in the domestic real income raises the demand for real balances; for a given path of the nominal money supply, money market equilibrium requires a price level which is lower than the price which would have prevailed otherwise. The lower domestic price level is however associated with a lower wage which induces an offsetting exchange rate depreciation. It is worth noting that the real income elasticity of the money demand is equal to 1.98. The rate of change in the price level differential significantly affects the rate of change in the exchange rate. The value of the structural parameter β is not significantly different from one. The explanation of the negative association between $\Delta s(t)$ and $\Delta(m(t-1)-p(t-1))$ is that the rate of change of the real money supply measures the effect of current wage inflation on the exchange rate. On the other hand, the positive association between $\Delta s(t)$ and $\Delta(w(t-1)-p(t-1))$ means that the rate of change in the real wage measures the effect of unanticipated inflation on the exchange rate. The interest rate is statistically insignificant because of the low variation due to interest rate controls. The equilibrium error $u(t-1)$ insignificantly affects $\Delta s(t)$ implying that the PPP has not determined the stance of the long run exchange rate policy. This result confirms that obtained from the cointegration tests and also supports the evidence of the analysis of PPP for the Greek drachma (Karfakis and Moschos, 1989). The value of the parameter θ_1 is equal to 0.653 which is not significantly different from one.[3] This result seems to be consistent with the short run nature of

the wage contracts. The low value of R^2 implies that a significant part of the exchange rate policy remains unpredictable.

The diagnostic tests reject the presence of serial correlation, functional misspecification and heteroscedasticity at the 0.05 significance level. Furthermore, the normality test suggests that the error process presents a random sample from a normal distribution.

5. CONCLUDING REMARKS

This paper has used a structural model to derive and estimate a reduced-form EC specification for the US dollar-Greek drachma exchange rate over the period 1975.I-1987.IV. The evidence has showed that the short run exchange rate policy in Greece stabilizes movements in the wage inflation. In other words, the monetary authorities have pursued a short run antiinflationary exchange rate policy that appreciates the exchange rate in the presence of increasing wage costs as an attempt to mitigate the depreciating pressures on the domestic currency and thus to ease the adjustment required on the Greek producers. Furthermore, the nonestablishment of a systematic long run relationship between the exchange rate and the price differential implies that the exchange rate target of the monetary authorities has not been consistent with PPP.

NOTES

1. Owing to the existence of capital controls over the sample period, it is assumed that any form of the interest parity condition is to be violated.

2. The exchange rate (central bank midpoint rate, S), the Greek consumer price index (P, 1980=100), the US consumer price index (FP, 1980=100) and the nominal wage (W, 1980=100) are obtained from the IMF's International Financial Statistics, various issues. The money supply (currency in circulation plus private sight deposits with commercial banks and specialized credit institutions, M), and the three-six month deposit rate (r) are obtained from the Bank of Greece, the Monthly Statistical Bulletin, various issues. The gross domestic product at factor costs (Y, 1970 prices) is taken from the National Accounts of Greece, February 1989.

3. To obtain the value of the asymptotic standard error of θ_1 (=0.728), the formula given in Phillips and Wickens (1978, pp.110-111) used.

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Table 1. Dickey-Fuller unit root tests

s	(p-fp)	y	r	(m-p)	(w-p)
			Level		
-0.51(0)	4.20(0)	-2.90(2)	-1.10(1)	-0.49(1)	-4.20(0)
[7.76]	[7.40]	[7.90]	[7.93]	[8.82]	[5.98]
			First Difference		
-3.75(1)	-4.04(0)	-7.33(1)	-4.83(0)	-4.70(0)	-4.45(0)
[6.40]	[4.04]	[6.26]	[7.45]	[3.89]	[8.36]

Note: Figures in parentheses denote the number of lagged dependent variables in the regression. The selection between zero and nonzero lags was based on the Lagrange multiplier (LM) test for fourth-order serial correlation of the residuals. Figures in squared brackets refer to the values of the LM(4) statistic. The rest of the entries are the values of the DF test, the critical value of which at the 0.05 level is -2.93 for T=50 (Fuller 1976, p.373).

Table 2. Cointegration tests

$$s(t) = \underset{[0.02]}{3.8769} + \underset{[0.042]}{1.2893}(p(t)-fp(t)) + \hat{u}(t)$$

CIRDW = 0.1714 $\qquad R^2 = 0.9517$

$$\Delta\hat{u}(t) = \underset{[0.06]}{-0.0211}\hat{u}(t-1)$$

$$DF = \underset{\{8.83\}}{-0.3358}$$

Note: Number in squared brackets are estimated standard errors. Figure in the brace refer to the value of the LM(4) test for autocorrelation of the residuals of the auxiliary equation. The critical values of the CIRDW and DF tests at the 0.05 level are 0.78 and -3.67 respectively for T=50 (Engle and Yoo 1987).
See also notes to table 1.

Table 3. Results of the EC model of the US dollar-Greek Drachma exchange rate

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$$\underset{[0.1647]}{\Delta s(t)=0.045\Delta s(t-1)}+\underset{[0.9716]^*}{2.513\Delta(p(t)-fp(t))}+\underset{[0.6121]^*}{1.257\Delta y(t)}+\underset{[1.8366]}{2.54\Delta r(t)}+$$

$$\underset{[0.32]^*}{0.997\Delta(w(t-1)-p(t-1))}-\underset{[0.2242]^*}{0.651\Delta(m(t-1)-p(t-1))}-\underset{[0.0761]}{0.061\hat{u}(t-1)}$$

$R^2=0.2635$ SEE=0.0518

<u>DIAGNOSTIC TESTS</u>

1. <u>Correlation</u>
   DW=1.9240
   Sargan's Test: $Q_1(1)=0.6961$ (>0.30)
   Sargan's Test: $Q_2(12)=6.4137$ (>0.80)
   LM Test: $Q_3(4)=1.0753$ (>0.30)

2. <u>Functional Form</u>
   Ramsey's Reset Test: $Q_4(1)=1.0713$ (>0.30)

3. <u>Normality</u>
   Jarque-Bera Test: $Q_5(2)=0.0222$ (=0.90)

4. <u>Heteroscedasticity</u>
   LM Test: $Q_6(1)=0.3675$ (>0.50)

Note: Sample: 1975.I-1987.IV. Observations=50. Degrees of freedom=40. Numbers in squared brackets are estimated standard errors. An asterisk (*) indicates significance at the 0.05 level. Instruments used: $\Delta s(t-1)$, $\Delta p(t-1)$, $\Delta fp(t-1)$, $\Delta y(t)$, $\Delta r(t)$, $\Delta(w(t-1)-p(t-1))$, $\Delta(m(t-1)-p(t-1))$, and $u(t-1)$. Three seasonal dummies are also included in the regression which are not reported. SEE=standard error of the equation. The Ramsey's Reset test of functional specification uses the square of the fitted values. The Jarque-Bera normality statistic is based on a test of skewness and kurtosis of residuals. The test of heteroscedasticity is based on the regression of squared residuals on squared fitted values. The $Q_i(n)$ statistics are asymptotically distributed as $X^2(n)$ under the null hypothesis. Numbers in parentheses below the calculated X^2 statistics are marginal significance levels which denote the probability of finding a X^2 entry value smaller than or equal to the calculated X^2 statistic under the null hypothesis.

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