CENTRE OF PLANNING AND ECONOMIC RESEARCH

No 48

A Cointegration Analysis of the Official and Parallel Foreign Exchange Markets for Dollars in Greece

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November 1995

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ACKNOWLEDGEMENTS

An earlier version of the paper was presented at the Centre of Planning and Economic Research staff seminar. We wish to thank seminar participants for their constructive criticism. It has also benefited from discussions and insightful comments by Richard Baillie, Ching-Fan Chung, Steve Matusz, Rowena Pecchenino, Robert Rasche, Peter Schmidt and Jeffrey Wooldridge. Finally we thank Robert Rasche for providing us with his program for the DOLS estimation procedure. Any remaining errors and omissions are our responsibility.

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October 1995

CONTENTS

ABSTRACT

This paper examines the monetary model of exchange rate determination from a long-run perspective in the presence of a "parallel" or "black" market for U.S. dollars in Greece using monthly data for the recent float, in four ways. First, unit root tests that maintain both stationarity and non-stationarity about either mean or trend are employed to determine the order of integration of our data. Second, using the Johansen's multivariate cointegration technique we found one significant cointegration vector. Johansen's FIML and Stock and Watson's (1993) DOLS approach were employed to estimate the cointegration coefficients. Third, formal stability tests as described by Hansen and Johansen (1993) were used, and it is shown that the dimension of the cointegration space may exhibit sample dependency, but the estimated coefficients are not unstable in recursive estimations. Finally, a new efficient and consistent test that maintains the null of cointegration developed by Shin (1994) was utilized, and once again the evidence in favour of cointegration was accepted.

1. INTRODUCTION

The emergence of parallel or "black" markets is a well known feature of many developing countries for several decades, with parallel exchange rates deviating, in some cases, considerably from official rates. Parallel markets are the result of goverment regulations or procedures that directly restrict access to the official market or otherwise result in non-equilibrium exchange rates. Unsatisfied supply and demand then "spills over" into the parallel market or is "diverted" from the official market in response to more favourable black market exchange rates. The size of this market varies from country to country and depends on the type of the exchage and trade restictions imposed along with the degree to which these restrictions are implemented by the goverment agencies. Thus, in countries of Latin America, Asia and Africa that face chronic balance-of-payments problems and there exists excess demand for foreign currency at the official rate which is not satisfied by the insufficient reserves that the central banks hold, the black market premia is very high (e.g. in Ghana the premium was as high as 700% in the 1980s while in Venezuela it was 85%) and the parallel markets for foreign exchange are well organized. In contrast, in countries where there is a moderate degree of trade and foreign exchange restrictions, the role of the parallel market is minimal.

Over the last two decades a substantial literature has been developed on the issue of modelling and the macroeconomic implications of parallel markets for foreign currency (Gupta, 1981, Agenor, 1992 and Agenor et al., 1993, provide a theoretical and empirical analysis of these markets while Kiguel and O'Connell, 1995, provide a summary of black market premia in a variety of countries). One approach of studying these types of markets is based on the influential paper by Dornbush et al. (1983) who developed a stock-flow model for the transactions in the foreign exchange market of Brazil. This model allows the simultaneous determination of the premium on foreign exchange and the rate of change of the stock of black markets as a result of the interaction of stock and flow conditions in the parallel market. Such models are consistent with real trade models used to analyze the effects of smuggling and the development of a parallel market for foreign exchange (Pitt, 1984). Recently Phylaktis (1991) and Phylaktis and Manalis (1995) have used this model to analyse the parallel market for U.S. dollars in Chile and Greece, respectively. A different approach is based on the monetary approach to the exchange rate. Blejer (1978a,b) provides an extension of the monetary model in order to account for the existence of a parallel exchange rate that is determined freely by market forces and responds to disequilibria in the domestic money market while the official rate is administratively determined by the goverment responding to a reaction function that may be derived from a condition of utility maximization. Using data for Brazil, Chile and Colombia he found support for this variant of the monetary model. Agenor (1991) extended this version of the monetary model by introducing illegal trade transactions, currency substitution features and forward looking expectations. He tested this model for twelve developing countries and he found evidence favourable to these extensions.

The recent development in the econometrics of nonstationarities and cointegration have provided a new framework within which we can analyse and model the behaviour of official and parallel foreign exchange markets by treating the relationship betwen the official and parallel exchange rates or the relationship implied by the monetary model as a long-run equilibrium relationship. Within this framework Blangiewicz and Charemza (1990), Booth and Mustafa (1991), Baghestani and Noer (1993) have employed the Engle-Granger (1987) two-step cointegration methodology and have provided evidence for the existence of a long-run relationship between the official and parallel exchange rate for several currencies and then the short run dynamics were modelled. Akgiray et al. (1989), Agenor and Taylor (1993) and Zarangas (1995) investigate the issue of causality between the official and parrallel exchange rate for several currencies and they found that causality runs from either exchange rate in most cases. Finally, Van den Berg and Jayanetti (1993) analysed the long-run validity of the monetary model where the parallel exchange rate is considered as a function of the monetary variables suggested by the standard monetary model and the official exchange rate. They applied the Johansen (1988, 1991) multivariate cointegration technique and they found at least one cointegrating vector for the currencies of India, Pakistan and Sri Lanka.

The objective of this paper is to examine the long-run validity of the monetary approach to the exchange rate in the presence of a "parallel" market for U.S. dollars in Greece. The "parallel" market for U.S. dollars has been operating in Greece since the end of World War II. Its size has been considerable with the premiun being on average 15%. Even when the Greek goverment decided to allow the drachma to float freely against major currencies in April 1975 since at the same time it imposed trade and foreign exchange restrictions so that the official exchange rate was not purely market determined but was still rather administratively determined the 'parallel' market for U.S. dollars was still operating that undermined these restrictions while smuggling of goods was also taking place. Greece's joining of the European Economic Community in 1981 had its beneficial effects on the liberalization of trade and capital flows since Greece had to follow the rules set by the European Economic Community for the creation of the Single Market and the Monetary Union. Recently, Diamandis and Kouretas (1995a,b) have tested the long-run validity of the monetary model for the case of several Greek drachma bilateral official exchange rates using the Johansen (1988, 1991) multivariate cointegration technique and found strong support in favour of the model, while MacDonald and Taylor (1993, 1994) have done so for several major bilateral exchange rates. However, in this paper we assume that the "parallel" drachma-U.S. dollar exchange rate is the one that is actually market determined and we wish to examine whether a stable linear long-run relationship between this rate, the official rate and the monetary variables suggested by the monetary model exists. To conduct our analysis we employ several recent developments in the econometrics of nonstationarities and cointegration. First, in order to determine the order of integration of our data, the new KPSS test (Kwiatkowski, Phillips, Schmidt and Shin, 1992), which has stationarity as its null hypothesis as opposed to the standard Dickey-Fuller test which is a unit root test, is applied. Second, we further analyse the stochastic properties of the data by testing for the presence of structural breaks in the time series by applying the Zivot and Andrews (1992) sequential test which considers the break as endogenous, as opposed to Perron's (1989) data-dependent approach. Third, Johansen's (1988, 1991) methodology is used as basis for testing the existence of cointegration. Estimates of the cointegration relationships are obtained using Johansen's FIML method and Stock and Watson (1993) dynamic OLS (DOLS) approach. Fourth, we employ three tests based on recursive estimation for the evaluation of parameter constancy proposed by Hansen and Johansen (1993). Finally, a new efficient and consistent test that maintains the null of cointegration developed by Shin (1994) was employed, to provide firmer conclusion about the long-run validity of the monetary model. The overall findings suggest that the modified monetary model is a valid framework for analyzing the long-run movements of the 'parallel' drachma-dollar exchange rate, and this result is robust in alternative estimation methods.

The plan of this paper is as follows: Section 2 presents the Greek foreign exchange market and the monetary model of exchange rate determination. Data sources and time series properties of the individual variables are described in Section 3. The econometric methodology for modelling and testing cointegration is presented in Section 4. Section 5 reports the empirical results, with the interpretations and concluding remarks given in Section 6.

2. THE GREEK FOREIGN EXCHANGE MARKET AND THE MONETARY MODEL

"Parallel" or "black" market activities in U.S. dollars in Greece dates back to the end of World War II, when as a consequence of the enormous government budget deficit and the loans that had to be made for the financing of the German occupation troops, the Greek economy suffered from hyperinflation for a period of three years (1945-1948) in addition to the existing unstable political and social conditions. Thus, people lost their faith in the national currency and most of the transactions were made in U.S. dollars or in gold sovereign. This situation continued even after the implementation of the major reconstruction plan in the 1950s. Trade and foreign exchange restrictions that were in force until the mid-1980s sustained sizeable market for U.S. dollars in Greece and according to the estimates by Pavlopoulos (1987) the volume of transactions in 1984 in this market was approximately 400 million U.S. dollars.

The Greek foreign exchange market moved to a managed float system in April 1975, when the link to the U.S. dollar was abandoned and a variable trade-weighted system was adopted, where the U.S. dollar had the greater weight. With the Greek entrance in the European Economic Community in 1981, the Bank of Greece adjusted the trade-weighted system to place a greater weight on the Deutschemark and on the other European currencies and smaller on the dollar. The movement to the managed float system was coupled by tight controls on all kinds of capital flows (Manalis, 1993). This has led to a small size of the Greek foreign exchange market in terms of transactions and a limited number of market agents. Unstable political conditions and the first oil price shock gave a new momentun to the activities in the parallel market for U.S. dollars during the second half of the 1970s and the first half of the 1980s. Figure 1 shows the evolution of the parallel and official drachma-dollar exchange rate from 1975 to 1989, while Figure 2 shows the evolution of the parallel market premium for the same period. The premium is positive apart from short periods in the second half of the 1980s where it turned to a discount. This negative premium is explained by the fact that for some periods after 1985 the Bank of Greece has forced the commercial banks not to accept foreign currency without proper identification of the seller. In that case the seller was willing to undersell his foreign currency in the black market.

During the 1980s several significant economic events took place in the Greek economy that eventually led to the shrinkage of the parallel markets for dollars. First, following Greece's joining of the European Economic Community in January 1981 the import duties for goods and services were reduced and by January 1988 there were no tariffs for imported goods from the countries of the European Economic Community which are its major trade partners. Second, the drachma was added in the ECU basket in

Figure 1: Official and Parallel Greek Drachma-US Dollar exchange rates

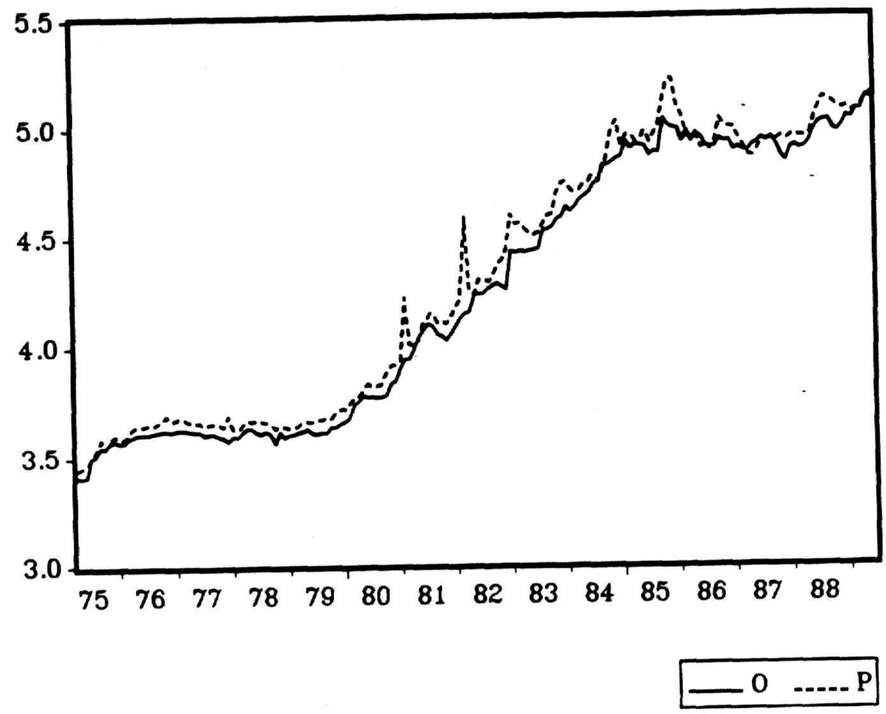
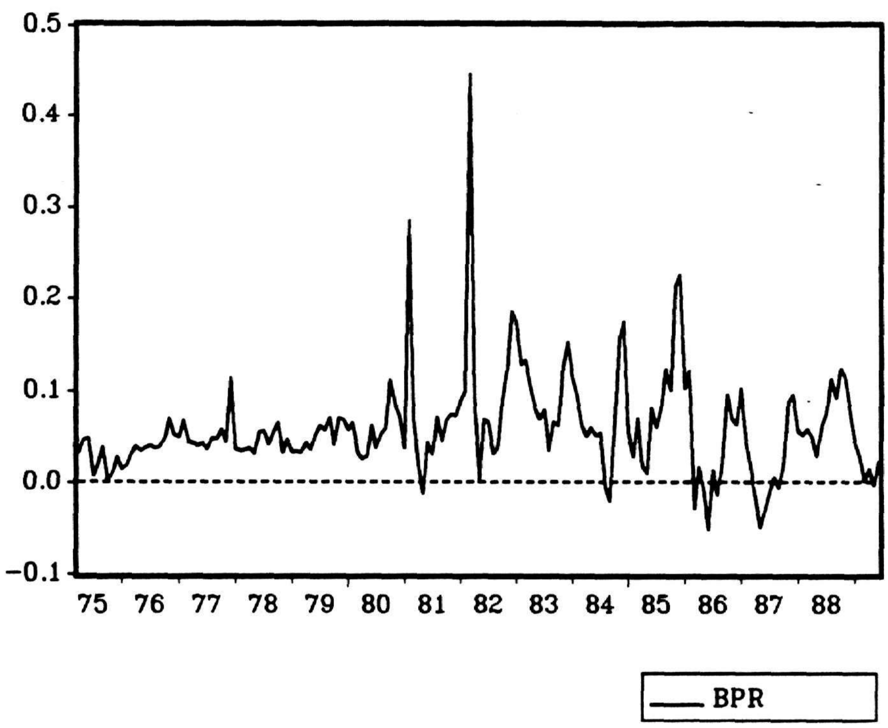


Figure 2: Parallel market premium



September 1984 with the obligation to join the Exchange Rate Mechanism at a later stage; this implies tighter fiscal and monetary policies which will affect the official and consequently the parallel exchange rate. Third, two discrete devaluations of the drachma were implemented in January 1983 and in October 1985 and each one was equal to 15%. The case of a negative premium after the second devaluation may also be explained by the likelihood that the parallel market agents were expecting a higher percentage of devaluation of the drachma than the realized one and that led to selling dollars at a discount. Finally, in January 1986 the liberalization process for capital flows began which was completed in May 1994 when all capital controls on short-term capital were lifted.

In view of the above we argue that the parallel market exchange rate depends on the underlying supply and demand for foreign exchange, which, according to the monetary model, are influenced by the domestic and foreign money supply, outputs and interest rates. However, the parallel exchange rate is also dependent on the level of the official exchange rate as well as on the diverse set of policies and institutions that regulate the official exchange market. If this latter set of policies and institutions is stable, we can then investigate whether there is a linear long-run equilibrium relationship between the parallel and official exchange rate (the spillover effect) and the monetary variables (the underlying shifts in supply and demand).

Under these assumptions a typical equation for testing the monetary approach using the parallel market exchange rate is

$$e_{pt} = \beta_0 + \beta_1 e_{ot} + \beta_2 m_t + \beta_3 m_t^* + \beta_4 y_t + \beta_5 y_t^* + \beta_6 r_t + \beta_7 r_t^* + u \qquad (1)$$

where e_p is the black market exchange rate, e_o is the official spot exchange rate, m is the domestic money supply, y is the domestic real output and r is the domestic interest rate, while an asterisk denotes the corresponding foreign variable, u_t is a white noise disturbance term, and all the variables except for the interest rates, are expressed in natural logarithms. The monetary model implies that $\beta_2>0$, $\beta_3<0$, $\beta_4<0$, $\beta_5>0$, $\beta_6>0$ and $\beta_7<0$. Finally for our case we also assume that $\beta_1>0$.

3. DATA AND PRELIMINARY EMPIRICAL FINDINGS

For this study parallel (black) market Greek drachma-U.S. dollar exchange rates were taken from the monthly series in various issues of the *World Currency Yearbook* (1989, 1986-1987) (International Currency Analysis, Inc., Brooklyn, New York), and its predeccesor, Pick's Currency Yearbook (1975-1976, 1968) (Pick Publishing Corporation, New York). The data for the official drachma-U.S. dollar exchange rate and the monetary aggregates, the output measures and the U.S. short-term interest rate were obtained from the CD-ROM edition of the IMF's International Financial Statistics, while the Greek short-term interest rate was obtained from the Monthly Bulletin of the Bank of Greece. Specifically, the official and parallel exchange rates are Greek drachmas per U.S. dollar, the monetary aggregate used is the M1, the income measure is the industrial production index, the U.S. short-term interest rate is the three-month treasury bill rate and the Greek interest rate is the 3-6 month deposit rate. Finally, the analysis is conducted for the period April 1975 (when the Greek drachma was allowed to float freely against the major currencies) through June 1989 (the last date data on the parallel exchange rate are available).

As a precondition for testing for cointegration between the parallel exchange rate, the official exchange rate and the respective macroeconomic variables suggested by the monetary model, the individual series are tested for nonstationary behaviour. For this purpose two alternative tests are employed. First, we apply the standard augmented Dickey-Fuller (ADF) test (Dickey and Fuller, 1981), that tests for the null of a unit root in the series against the alternative of stationarity. To allow for seasonality in the data, all regressions include eleven monthly seasonal dummies.[1] Recently, DeJong et al. (1992) have argued that the Dickey-Fuller class of tests have low power against trend stationary alternatives. For this purpose, we also applied the Kwiatkowski, Phillips, Schmidt and Shin (KPSS, 1992) test for the null of stationarity against the alternative of a unit root. Briefly, the KPSS test uses the components model

$$y_t = \alpha + \delta t + X_t + v_t, \qquad X_t = X_{t-1} + u_t \tag{2}$$

where y_t is the sum of the deterministic trend t, a random walk X_t, and a stationary error v_t. Also u_t are i.i.d. $(0, \sigma_u^2)$. Since v_t is assumed to be stationary for the null hypothesis that

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[1]. The use of seasonals does not change the limiting distribution of the ADF test statistics. See Dickey, Bell and Miller (1986), Appendix B.

y_t is trend stationary we simply require that $\sigma_u^2 = 0$. When σ_u^2 equals zero, X_t is a constant and is added to the intercept of equation (2). We can also set $\delta = 0$, in which case we test the null hypothesis that y_t is stationary around a level rather than around a trend. The KPSS test makes no allowance for inclusion of seasonal dummy variables and as a result, we did not include seasonals in the regressions.[1]

The combined results of the ADF and KPSS tests are given in Table 1 and we can unambiguously conclude that the levels of all the series contain a unit root. By contrast, the first differences of the series are found to be I(0) processes. Therefore, we treat all the series as being I(1) processes.

Perron (1989) argues that most macroeconomic time series are trend stationary if one allows for a one-time change in the intercept or in the slope (or both) of the trend function. Perron's methodology has been recently critisized on the grounds that it considers the break point as to be uncorrelated with the data, hence, as being an exogenous event. Recently, Zivot and Andrews (1992) treat the selection of the break point as the outcome of an estimation procedure and transform Perron's (1989), conditional (on structural change at a known point in time) unit root test into an unconditional unit root test (i.e. the break point is considered to be endogenous). Following Zivot and Andrews (1992), we test the null hypothesis of an integrated process with drift against the alternative hypothesis of trend stationarity with a one-time break in the intercept and the slope of the trend function at an unknown point of time (T_B), using the following augmented regression equation:

$$y_t = \hat{\mu} + \hat{\theta} DU_t(\hat{\lambda}) + \hat{\beta} t + \hat{\gamma} DT_t(\hat{\lambda}) + \hat{\alpha} y_{t-1} + \sum_{i=1}^{k} \hat{c}_i \Delta y_{t-i} + \hat{e}_t \tag{3}$$

where $DU_t = 1$ and $DT_t = t$ if $t > T_B$ and 0 otherwise.

Table 2 presents the results of the Zivot-Andrews sequential ADF test and we conclude that we are unable to reject the null hypothesis of a unit root for all the series even in the case that we allow for the presence of one-time structural break in each series.

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[1]. Evidence from the Monte Carlo simulation suggest that a smaller number of lags causes large size distortions while a larger number of lags tends to reduce the power of the test.

TABLE 1
Unit Root and Stationarity Tests

Variable	Level τ_μ	τ_τ	η_μ	η_τ	First Difference τ_μ	η_μ
e_o	0.096[5]	-1.667[5]	1.371[a]	0.172[a]	-5.22[5][a]	0.157
	(0.057)	(0.343)			(0.405)	
e_p	-0.461[1]	-1.914[1]	1.368[a]	0.166[a]	-11.07[1][a]	0.138
	(0.240)	(0.065)			(0.102)	
m	-0.474[4]	-3.173[4]	1.420[a]	0.223[a]	-9.40[4][a]	0.038
	(0.144)	(0.123)			(0.253)	
y	-2.652[5]	-3.595[5]	1.229[a]	0.258[a]	-8.29[5][a]	0.091
	(0.255)	(0.244)			(0.346)	
r	-0.297[0]	-1.938[0]	1.152[a]	0.245[a]	-12.96[0][a]	0.098
	(0.998)	(0.987)			(0.996)	
m*	-0.447[3]	-2.108[3]	1.405[a]	0.214[a]	-8.59[3][a]	0.103
	(0.300)	(0.276)			(0.409)	
y*	-1.745[3]	-2.787[3]	1.267[a]	0.128[a]	-10.77[3][a]	0.135
	(0.110)	(0.098)			(0208)	
r*	-1.587[12]	-1.620[12]	1.262[a]	0.263[a]	-5.37[12][a]	0.123
	(0.053)	(0.045)			(0.289)	

Notes: The symbols e_o and e_p denote, respectively, the spot official exchange rate and the parallel (black) exchange rate, while m, y, r and m*, y*, r* are, respectively, the Greek and U.S. money supply M1, industrial output index and the short-term interest rate. τ_μ and τ_τ denote the standard augmented Dickey-Fuller tests for the null of non-stationarity, when a constant and a constant and a time trend is included in the equation, respectively. Numbers in brackets after the ADF statistics indicate the lag length used in the autoregression to ensure residual whiteness while numbers in parenthesis indicate marginal significance level of the Lagrange Multiplier with 12 degrees of freedom. All regressions include eleven seasonal dummy variables. The critical values at the five percent significance level for the two tests and for T = 171, T = 170, T = 168, T = 167, T = 166 and T = 159 are respectively -2.8785, -2.8786, -2.8788, -2.8789, -2.8790 and -2.8798 and -3.4370, -3.4371, -3.4374, -3.4376, -3.4377 and -3.4389, (MacKinnon, 1991).
η_μ and η_τ is the KPSS test for the null of stationarity, when a constant and a constant and a time trend is included in the equation. Both tests were calculated with a lag truncation parameter of 12 lags. The five percent significant level for the two tests and for T = 171, is equal to 0.4616 and 0.151, respectively (Sephton, 1995, Table 2).
([a]) indicates significance at the five percent critical level.

Given these results for the univariate case we now proceed to investigate whether a stable *linear* long-run relationship between the parallel drachma-U.S. dollar exchange rate and the respective official exchange rate and the macroeconomic variables can be established.

TABLE 2

Unit Root Tests under a Structural Break

Variable	$t(\alpha)$	T_B
e_o	-2.427	1983.01
e_p	-2.099	1982.01
m	-5.013	1983.03
y	-4.763	1978.09
r	-4.345	1978.09
m_{us}	-3.122	1986.06
y_{us}	-3.299	1981.11
r_{us}	-3.700	1982.05

Notes: Notation as in Table 1. $t(\alpha)$ is the Zivot-Andrews (1992) minimum t-statistic allowing for shifts in the mean and the slope over all the T-2 regressions for testing $\alpha = 1$. The asymptotic critical value at the 5% sifnificance level is -5.08 (Zivot and Andrews, 1992, Table 4).

The estimated equation is:

$$y_t = \hat{\mu} + \hat{\theta} DU_t(\hat{\lambda}) + \hat{\beta} t + \hat{\gamma} DT_t(\hat{\lambda}) + \hat{\alpha} y_{t-1} + \sum_{i=1}^{k} \hat{c}_i \Delta y_{t-i} + \hat{e}_t$$

where $DU_t = 1$ and $DT_t = t$ if $t > T_B$ and 0 otherwise.

4. ECONOMETRIC METHODOLOGY

Our cointegration analysis is based on Johansen's (1988, 1991) multivariate cointegration methodology and on Shin's (1994) residual based test for the null of cointegration, that is considered to be a multivariate extention of the KPSS stationarity test. Furthemore, the estimation of the cointegration vectors is done using the Johansen's FIML method and Stock and Watson (1993) DOLS approach.

According to Johansen (1988, 1991) any p-dimensional vector autoregression can be written in the following "error correction" representation.

$$\Delta X_t = \sum_{i=1}^{k} \Gamma_i \Delta X_{t-i} + \Pi X_{t-k} + \phi D + \mu + \varepsilon_t \tag{4}$$

where X_t is a p-dimensional vector of I(1) processes, μ is a drift, ε_t is an i.i.d. p-dimensional vector with zero mean and variance-covariance matrix Λ, D is a vector of nonstochastic variables, such as centered seasonal dummies which sum to zero over a full year by construction and they are necessary to account for short-run effects which could otherwise violate the Gaussian assumption and/or intervention dummies. T is the sample size.[1]

The matrix Π provides the long-run information in the data. When $0 < \mathrm{rank}(\Pi) = r < p$, Π can be decomposed into:

$$\Pi = \alpha \beta' \tag{5}$$

where both α and β are $p \times r$ matrices, and the rows of β' form the r distinct cointegrating vectors, while α may be interpreted as the "error correction" parameters.

Johansen's (1988, 1991) procedure provides maximum likelihood estimates of α, β and Π as well as two likelihood ratio test statistics to determine the rank of the cointegration space. With the trace statistic the null hypothesis is that there are at most r cointegrating

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[1]. Gonzalo (1994) shows that the performance of the maximum likelihood estimator of the cointegrating vectors is little affected by non-normal.

vectors, while with the maximum eigenvalue statistic, we test for the presence of r versus r+1 cointegrating vectors.

Subsequently, Johansen and Juselius (1990, 1992) developed tests for linear restrictions on the individual elements of α and β such as those imposed on the monetary model.

Johansen (1992a) provides an alternative representation of system (3 and 4) that assumes weak exogeneity and a particular a priori normalization to obtain

$$\Delta x_t^1 = \upsilon_t^1 \tag{6}$$

$$x_t^2 = \vartheta_0 + \vartheta x_t^1 + \upsilon_t^2 \tag{7}$$

where $x'_t = [x_t^{1'} \mid x_t^{2'}]$ with $x_t^1\,[(p - r)_{xl}]$ and $x_t^2[r_{xl}]$. As Phillips (1991) show, the error processes υ_t^1 and υ_t^2 are stationary in this formulation.In a system characterized by r cointegrating vectors, an arbritrary normalization as described in equations (6) and (7) may be obtained. Recently, Stock and Watson (1993) have shown that we can obtain an estimator which is equivalent to a maximum likelihood estimator by estimating the normalized cointegrating vectors, ϑ, by OLS through equation (7) that has been augmented with both leads and lags of Δx_t^1. The inclusion of leads and lags terms accounts for possible correlation between the error processes υ_t^1 and υ_t^2. Furthermore, it is usually the case that the errors of the augmented specification follow a moving average process and therefore an adjustment of the variance-covariance matrix of the estimated parameters is required to take account for the nonspherical error process in the augmented structure of (7). In their work Stock and Watson (1993) suggest a number of alternative approaches for the required adjustment. For our case we employ the dynamic OLS (DOLS) approach that utilizes a covariance matrix estimated by an autoregressive spectral estimator with two lags.

The major difference between the Johansen and Stock-Watson estimators stems from the fact that the former is derived from a system equations methodology while the latter is obtained from a single equation technique, and thus a normalization *prior* to estimation is required (the coefficient of the exchange rate is set to one). Therefore, we expect that the two estimates will usually differ. Moreover, while the Johansen technique simultaneously determines the rank of the cointegration space and provides estimates of the cointegrating vectors, the Stock and Watson approach is based on a specific normalization that requires *prior* knowledge of the dimension of the cointegration space. In this study we base the normalization necessary for applying the Stock and Watson technique on the dimension of the rank of the cointegration space determined by the Johansen likelihood ratio statistics.

Recently, Sephton and Larsen (1991) have shown that the Johansen's test may be characterized by sample dependency. In addition Stock and Watson (1993) also argue that the Johansen estimates of the cointegrating vectors may display considerable instabilities in recursive estimates (DOLS estimates may suffer, too). Hansen and Johansen (1993) have suggested methods for the evaluation of parameter contancy in cointegrated VAR models, formally using estimates obtained from the Johansen FIML technique. Three tests have been constructed under the two VAR representations. In the "Z-representation" all the parameters of model (3) are re-estimated during the recursions while under the "R-representation" the short-run parameters Γ_i, i = 1...k, are fixed to their full sample values and only the long-run parameters in (4) are re-estimated.[1] The first test is called the *Rank* test and we examine the null hypothesis of sample dependency of the cointegration rank of the system. A second test considers the null hypothesis of the constancy of the cointegration space for a given cointegration rank. Likelihood ratio test statistics are constructed for each sample size in order to test whether the cointegration space estimated from the full sample falls within the space spanned from the estimated vectors of each individual space. The test is a χ^2

[1]. The motivation for the "R-representation" is that by fixing the estimates of the short-run parameters we reduce the variance of the parameters in which cointegration analysis is primarily interested, i.e. the long-run (Hansen and Johansen, 1993).

distributed with (p-r)r degrees of freedom, where p stands for the number of endogenous variables and r for the cointegration rank. Finally, the third test examines the constancy of the individual elements of the cointegration vectors, through the unique relationship that exists between the eigenvalues and the cointegration vectors.

Johansen's methodology is truly a test for testing the null hypothesis of *no cointegration* rather than that of *cointegration*, and additionally both likelihood ratio tests have low power against the alternative of cointegration. Therefore, since our primary interest is the hypothesis of cointegration, it is often argued that cointegration would be a more natural choice of the null hypothesis. Recently, Shin (1994) provides a consistent residual-based test of the null of cointegration, and this test is a multivariate extension of the KPSS stationarity test that we discussed in Section 3. Shin (1994) simply adds an m-vector Z_t of I(1) variables in equation (2), so we get the following cointegrating relationship

$$y_t = \alpha + \delta t + Z_t^* \beta + X_t \tag{8}$$

As with the univariate case we can distinguish three cases; the cointegrating regression without intercept and trend, with intercept only and with intercept and trend. The null hypothesis is again $\sigma_u^2 = 0$. Two points must be made with respect to this test. First, the residuals must be obtained from an efficient estimation method such as Johansen's method is. Second, when we have a multivariate context a meaningful comparison with Johansen's results can only be drawn when at most one *linear* cointegrating relationship has been found. In the presence of more than one *linear* accepted cointegrating vector the multivariate KPSS test is not suitable.

5. COINTEGRATION RESULTS

In testing the long-run validity of the monetary model for the Greek drachma - U.S. dollar exchange rate, when the parallel along with the official foreign exchange market is considered, our (8×1) monthly data vector is given by

$$X_t = [e_p, e_o, m_t, m_t^*, y_t, y_t^*, r_t, r_t^*] \qquad t = 1975.04........1989.06$$

In addition we needed a vector of dummy variables given by

$$D_t = [S_1, S_2, S_3, S_4, S_5, S_6, S_7, S_8, S_9, S_{10}, S_{11}, D_{83}, D_{84}, D_{85}],$$

where $S_1,...,S_{11}$ are centered seasonal dummies while D83 and D85 are two shift dummy variables to account for the effects of the two devaluations and they take the value of 1 on 1983.01 and 1985.10, respectively, and the value of zero otherwise. Finally, D84 is a shift dummy variable with value 1 for t = 1984.10,...,1989.06, 0 otherwise, and is needed to account for the inclusion of the drachma in the ECU basket.

Before proceeding with the application of Johansen's methodology one must choose the lag structure of the VAR equation (3). We used as a benchmark a model with thirteen lags under which serial correlation is present in none of the equations of the system. This model was then tested against the alternative ones with progressively fewer lags. Two selection criteria were applied: The Sims (1980) likelihood ratio for the choice of the lag structure, and the Ljung-Box Q statistic for the presence of serial correlation in the residuals of each equation.The marginal significance level of both statistics was found to be above the 5 percent level for a model with 9 lags. Models with lower lag structure were found to suffer from non-whiteness in their residuals.[1]

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[1]. Eleven seasonal dummies were included in the estimation of each equation to take account of any seasonal effects in the series.

In Table 3 we report the trace and maximum eigenvalue statistics obtained using Johansen's multivariate maximum likelihood technique for estimating cointegrating relationships.[1],[2] The trace and the maximum eigenvalue tests reject the null hypothesis that there are zero cointegrating vectors or seven common trends. Both tests suggest that one cointegration vector exists for the drachma-dollar case. This finding indicates that the monetary model is a valid long-run framework for analyzing movements of the parallel exchange rate as a function of the official exchange rate (the spillover effect) and the monetary variables (the underlying shifts in the supply and demand). Table 3 presents two additional significant diagnostics for our procedure. First, when analyzing the VAR model, it is sometimes the case that only a subset of the variables in the Z_t-vector are needed in the cointegration space. Thus, to check whether the parallel exchange rate enters into the cointegrating relationship, a likelihood ratio test of a zero restriction on the cointegrating vector was performed. The hypothesis that the cointegrating relationship excludes the parallel exchange rate variable was rejected at the five percent significance level. The same outcome was obtained when exclusion restrictions are imposed for the official exchange rate and the macroeconomic variables. Second, although applying the univariate Dickey-Fuller class of unit root test statistics and the KPSS test is a standard procedure that we also followed in Section 3 it is also understood that if the data are being determined in a multivariate framework, a univariate model is, at best, a bad approximation to the multivariate counterpart, at worst, completely misspecified leading to arbitrary conclusions. Therefore, we have also applied the stationarity test proposed by Johansen and Juselius (1992) in which we test for the null hypothesis of stationarity for each variable conditional on one cointegrating vector for our case. This is also a likelihood ratio test that is distributed as a χ^2 statistic with (p-r) degrees of freedom. The results are reported in Table 3 and stationarity is rejected for all the series of the analysis.

In Table 4, the estimates of the normalised cointegrating vectors, that are obtained from two alternative techniques are given. Johansen's estimates which are based upon eigevectors obtained from the solution of the eigenvalue problem resulted from the reduced rank regression approach and Stock and Watson's (1993) DOLS estimates which are obtained by applying OLS on an equation with the parallel exchange rate as dependent

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[1]. To determine whether the constant is restricted in the cointegrating vector or enters unrestrictedly in the VAR model we applied the testing procedure suggested by Johansen (1992b).

[2]. The estimation of the Johansen's results as well as the Hansen-Johansen stability tests were performed with the software package CATS 1.0 in RATS 4.2.

TABLE 3
Johansen - Juselius Likelihood Cointegration Tests

(5% Critical Values)

r	Trace	λmax	Trace	λmax
r<7	182.47[a]	62.75[a]	156.00	51.42
r<6	119.72	33.57	124.24	45.28
r<5	86.15	30.03	94.15	39.37
r<4	56.12	22.26	68.52	33.46
r<3	33.86	16.10	47.21	27.07
r<2	17.76	12.07	29.68	20.97
r<1	5.69	4.51	15.41	14.07
r<0	1.18	1.18	3.76	3.76

Notes: r denotes the number of eigenvectors. Trace and λmax denote, respectively, the trace and maximum eigevalue likelihood ratio statistics. The 5% critical values are taken from Osterwald-Lenum (1992, Table 1*). A model with an unrestricted constant in the VAR equation is estimated.
([a]) denotes statistical significance at the five percent critival level.

Tests of Exclusion Restrictions and Stationarity

Variable	Exclusion Restrictions	Stationarity Test
e_o	8.02[a]	82.39[a]
e_p	11.40[a]	82.03[a]
m	24.03[a]	81.65[a]
y	24.46[a]	81.45[a]
r	10.05[a]	78.04[a]
m*	11.01[a]	81.53[a]
y*	12.50[a]	76.36[a]
r*	15.90[a]	49.88[a]

Notes: For the test of exclusion restrictions figures are χ^2 statistics with one degree of freedom, and the critical value at the 5% significance level is 3.84. For the multivariate stationarity test figures are χ^2 with seven degrees of freedom, and the critical value at the 5% significance level is 14.07.
([a]) denotes statistical significance at the five percent critical level.

TABLE 4
Estimated Coefficients and Hypothesis Testing Eigenvector

	e_p	e_o	m	m*	y	y*	r	r*
ML	1.000	-0.858	-0.644	1.108	0.062	-0.218	-0.013	0.015
DOLS	1.000	-1.059	-0.041	0.084	0.799	-0.619	-0.013	0.008
		(0.040)	(0.018)	(0.029)	(0.206)	(0.188)	(0.008)	(0.004)

DOLS Recursive Estimation

Period	e_p	e_o	m	m*	y	y*	i	i*
1975:7-	1.000	-1.231	-0.124	0.064	0.457	-0.198	-0.006	0.003
1984:1		(0.071)	(0.034)	(0.015)	(0.223)	(0.035)	(0.002)	(0.001)
1975:7-	1.000	-1.105	-0.079	0.154	0.610	-0.611	-0.014	0.006
1985:1		(0.063)	(0.021)	(0.039)	(0.228)	(0.154)	(0.004)	(0.003)
1975:7-	1.000	-1.103	-0.190	0.386	0.874	-0.741	-0.015	0.004
1986:1		(0.065)	(0.056)	(0.103)	(0.199)	(0.206)	(0.006)	(0.002)
1975:7-	1.000	-1.135	-0.014	0.180	0.921	-0.619	-0.013	0.004
1987:1		(0.053)	(0.003)	(0.076)	(0.223)	(0.177)	(0.007)	(0.001)
1975:7-	1.000	-1.070	-0.066	0.150	0.750	-0.637	-0.013	0.001
1988:1		(0.041)	(0.013)	(0.067)	(0.205)	(0.180)	(0.007)	(0.0005)
1975:7-	1.000	-1.050	-0.096	0.158	0.728	-0.602	-0.014	0.007
1989:1		(0.037)	(0.033)	(0.075)	(0.198)	(0.175)	(0.007)	(0.003)

Hypothesis Testing

$$e_t = \beta_0 + \beta_1 e_p + \beta_2 m_t + \beta_3 m^* + \beta_4 y_t + \beta_5 y^* + \beta_6 r_t + \beta_7 r^* + u_t$$

H_1 (β_1 = 1)	H_2 ($\beta_2 = -\beta_3 = 1$)	H_3 ($\beta_4 = -\beta_5$)	H4 ($\beta_6 = -\beta_7$)	H_5 $H_2 \cap H_3$	H_6 $H_2 \cap H_4$	H_7 $H_3 \cap H_4$	H_8 $H_2 \cap H_3 \cap H_4$
13.26[a]	25.23	1.93	0.08	25.67	25.83	5.87	25.84
(0.00)	(0.00)[a]	(0.16)	(0.78)	(0.00)[a]	(0.00)[a]	(0.54)	(0.00)[a]

Notes: notation as in Table 1. The eigenvectors have been normalized with the estimated coefficient on the nominal official exchange rate. Numbers in parenthesis under the DOLS estimates provide standard errors. They are calculated using a Newey-West procedure and Bartlett lag window set at q = 2. H_1 to H_8 denote the likelihood ratio statistic for the null hypothesis provided in the respective column, constructed as in Johansen and Juselius (1990) and is distributed as central χ^2 under the null with $r \times s$ degrees of freedom where r denotes the number of cointegrating vectors and s is the number of restrictions. The numbers in brackets are marginal significance levels.
([a]) denotes statistical significance at the five percent critical level.

variable and a constant and contemporaneous and two leads and lagged differences of all the explanatory variables. Both estimates are shown to have reasonable size and the signs predicted from the monetary model (1). In addition, for the DOLS estimates we have calculated standard errors which reflect the adjustment for persistence in errors and are calculated using the Newey-West correction suggested by Stock and Watson. A lag of 2 is used to calculate Bartlett weigths used in this procedure. A comparison of the two estimates from the two techniques reveal similarities which are in variant with some of the simulations performed by Stock and Watson.

Our finding of one cointegrating vector allows us to proceed to test several commonly imposed monetary restrictions implied by equation (1). Table 4 presents the appropriate likelihood ratio statistics described in Johansen and Juselius (1990, 1992) for testing these restrictions. The overall conclusion is that the hypothesis of proportionality between the two exchange rates is rejected. Similarly, the proportionality hypothesis of the parallel exchange rate and the two monetary aggregates as well as most of the other restrictions are rejected apart from the hypothesis of equal and opposite coefficients on relative income and interest rates.

Since the period under consideration involves interesting subperiods of important financial changes in the Greek economy, we conduct the stability tests which are based on the recursive analysis suggested by Hansen-Johansen (1993), through which we can analyse the stability of our cointegrating results. For the recursive analysis, we chose as starting point January 1984, thus we were able to examine possible structural effects that may had on the long-run relationship implied by the monetary model, the inclusion of the drachma in the ECU basket in September 1984, the devaluation of the drachma in October 1985 and the process of the abandoment of the capital controls in Greece which started in January and completed in May 1994 (Christodoulakis and Karamouzis, 1993). Figure 3 shows that we were unable to reject the null hypothesis that the dimension of the cointegration space is sample dependent at the 5 percent significance level, since the evidence in favour of one cointegration vector is established only after the liberalization process on the capital movements have begun. Figure 4 shows the results of the application of the second test and we unambiguously conclude that we were unable to reject the null hypothesis of constancy of the cointegration space for a given cointegration rank. This is a χ^2 statistic distributed with (p-r)r degrees of freedom its values have been scaled by the 95% quantile in the χ^2-distribution such that unity corresponds to a test with a 5% significance level. The final test is presented in Figure 5. It provides strong support of the constancy of the estimated cointegration vector since, through time, it is observed that the path of the respective eigenvalue always lies within the bounds of its 95 percent confidence interval. The basic idea of this test is that if the cointegrating vector has undergone a structural change this will be reflected in the estimated eigenvalues. Again, the test statistic has been scaled, such as to produce values below one under the null hypothesis of parameter stability. As a means of comparison with the last two tests of the Hansen-Johansen procedure in Table 4 we present a recursive analysis for the DOLS estimates for the same period and we show that the DOLS estimates are shown to be quite stable in recursive estimation. These results further reinforce our conclusion that the monetary model of exchange rate determination is a valid framework to analyse movements of the parallel drachma-U.S.dollar exchange rate from a long-run perspective.

Figure 3. Trace tests

1 is the 5% significance level

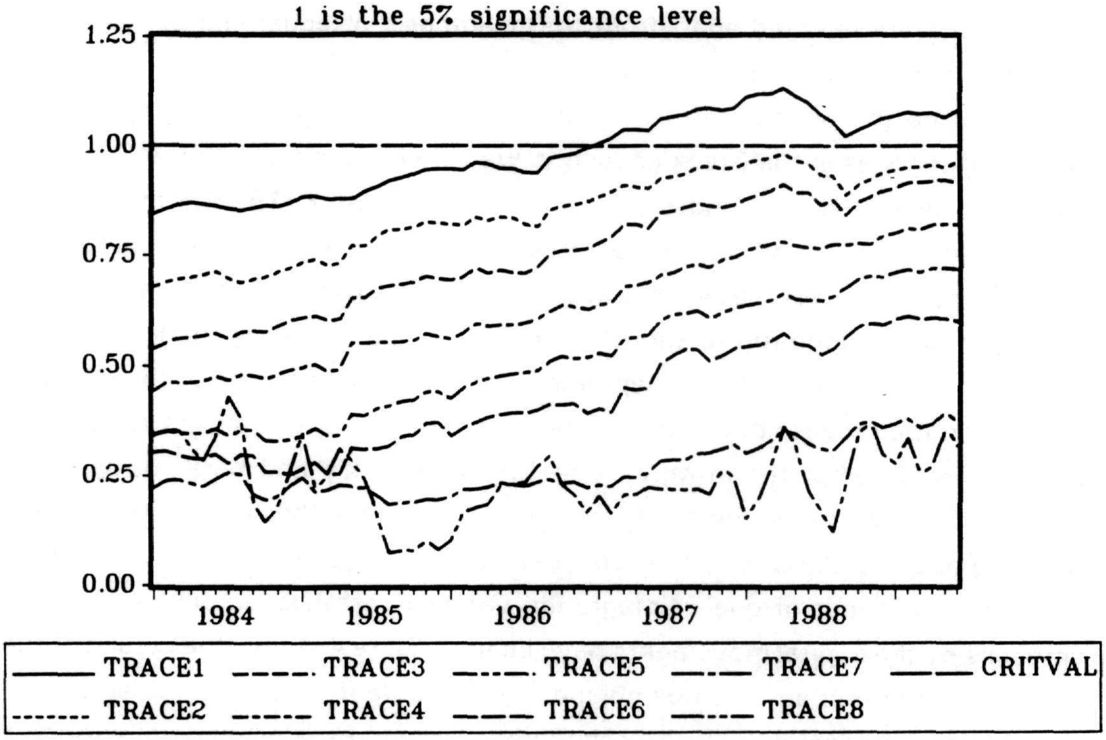

Figure 4. Test of known beta eq. to beta(t)

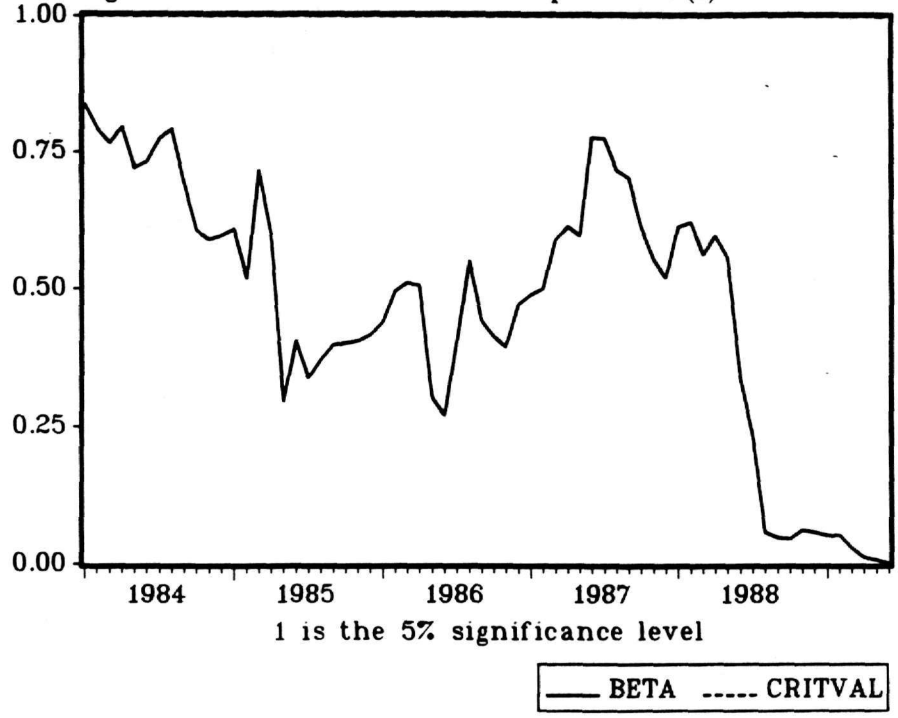

Figure 5. Test for Lambda1

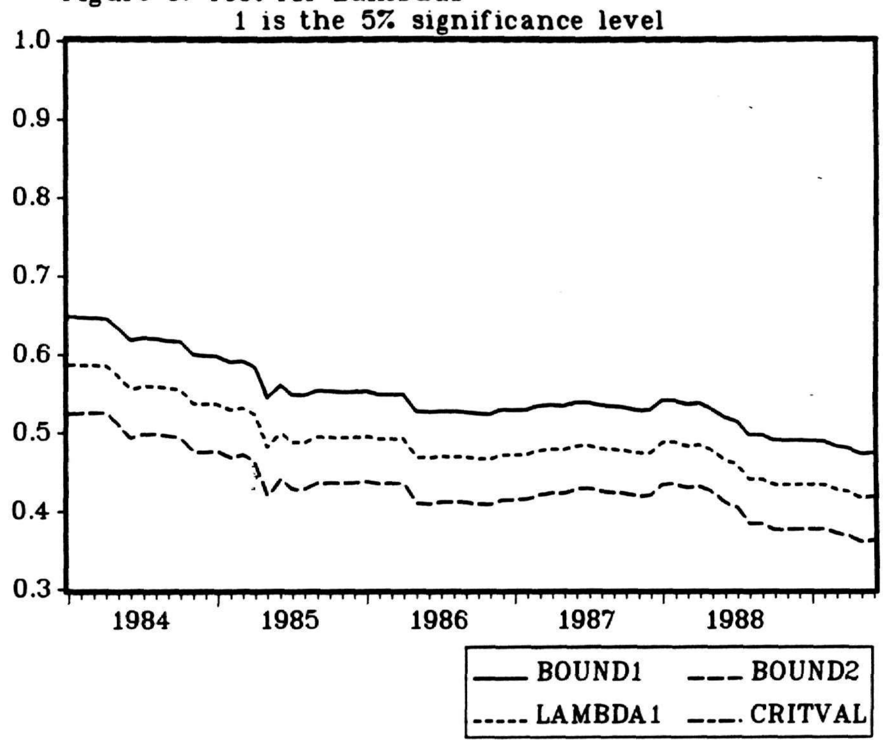

Table 5 gives the results from the application of the multivariate KPSS test for testing the null hypothesis of cointegration applied on the residuals of the accepted cointegration vector from the Johansen analysis. The picture emerging from these results reconfirm the existence of a *linear* long-run relationship between the parallel exchange rate and the official exchange rate and the Greek and U.S. money supplies, outputs and short-term interest rates. Thus, we conclude that the long-run validity of the monetary model in explaining movements of the exchange rates is robust.

TABLE 5

Residual Test for the Null Hypothesis of Cointegration
KPSS Cointegration Test

η_μ	η_τ
0.043	0.038

Notes: η_μ and η_τ is the KPSS test for the null hypothesis of cointegration against the alternative of no cointegration applied on the residuals of the Johansen's ML estimates, when a constant and a constant and a time trend, respectively, enter the estimated equation. In order to construct a consistent estimator of the long-run variance of the cointegrating regression residuals a lag trucation parameter with 12 lags was chosen. The five and one percent critical values are provided by Shin (1994, Table 1 for up to 5 regressors). Since in our case we have eight regressors we notice that the critical values are declining monotonically in a rate that guarantees thee outcome of our results.

Number of Regressors

	1	2	3	4	5	
η_μ	0.314	0.221	0.159	0.121	0.097	5%
	0.533	0.380	0.271	0.208	0.150	1%
η_τ	0.121	0.101	0.085	0.073	0.061	5%
	0.184	0.150	0.126	0.109	0.087	1%

6. INTERPRETATIONS AND CONCLUDING REMARKS

In this paper we have examined the monetary model of the exchange rate modified to incoroprate the existence of a substantial parallel market for U.S. dollars in Greece. The data used cover the period from April 1975 to June 1989. Although in April 1975 the Greek drachma was allowed to float freely against major currencies at the same time tight capital and trade controls have been implemented. Coupled with political uncertainty these restrictions led to the emerge of a parallel market for U.S. dollars of a considerable size during the period under examination. After a period where the premium has been continuously increasing this market eventually diminished as a result of Greece's entrance in the European Economic Community and the consequent abandoment of trade restrictions and capital controls (the capital movement liberalization process began in January 1986 and was completed in May 1994).

Our analysis employs recent developments in the econometrics of nonstationarities and cointegration. Several important findings were reported. First, the Johansen multivariate cointegration technique was applied and one statistically significant cointegrating vector between the parallel drachma-U.S. dollar exchange rate, the official exchange rate, and the variables suggested by the monetary model. Second, the coefficients were estimated using Johansen's FIML technique and the Stock and Watson DOLS procedure. They had the correct sign and reasonable size and they exhibit no significant difference across the two methods of estimation.Third using the Hansen and Johansen (1993) testing methodology we show that the dimension of the cointegration space may exhibit sample dependency, but the estimated coefficients did not exhibit instabilities in recursive estimation. Similar results were obtained for the DOLS estimates. Finally, given that the Johansen's test is truly a test for the null of no cointegration and that both likelihood ratio test have low power against the alternative of cointegration we applied the recently developed efficient and consistent residual-based multivariate KPSS test due to Shin (1994) that has *cointegration* as its null hypothesis, and we were unable to reject it. This implies that our finding of one cointegrating vector is robust, leading to the conclusion that the monetary model is a valid framework to analyse long-run movements of the parallel drachma-U.S. dollar exchange rate.

These findings have several important policy implications. First, we should reject the view that in the presence of a parallel market for U.S. dollars, the official rate is irrelevant, merely a goverment bookkeeping convention, and unrelated to the parallel exchange rate which might be assumed to be the market-clearing rate. Second, given the above relationship along with the effects that monetary variables have on the parallel exchange rate, the Greek authorities should take into consideration the existence of the black market for U.S. dollars in Greece for that period, so that the foreign exchange and trade restrictions to be as effective as possible. This is necessary since it was observed that in several cases and especially during the two devaluations of the drachma black market effects had dampened the direct effects of restrictions on the official reserves. Overall, the existence of the parallel market for U.S. dollars in Greece weakened the effectiveness of capital controls and trade restrictions. Finally, the experience of the Greek economy over the last twenty years shows that indeed the adoption of a managed float exchange rate is necessary but not sufficient for the determination of the exchange rate by the market mechanism. Equally important is the lack of capital controls and trade restrictions, otherwise a parallel market for foreign exchange emerges. The Greece's joining of the European Economic Community in 1981 and the consequent harmonization of its exchange rate policy along with the fiscal and monetary policy to the targets set by the European Union led to the liberalization of the trade and capital movements (the short-term capital controls were lifted in May 1994) and the subsequent diminishing of the black markets for U.S. dollars in the end of 1989, while at the same we observed an increase in the efficiency of the foreign exchange market.

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