The monetary model of exchange rate determination: the case of Greece (1974–1994)

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Abstract: The flexible-price monetary model is examined for the Greek drachma-US dollar exchange rate. The Johansen multivariate technique of cointegration is applied to an unrestricted form of the monetary model. Using quarterly data covering the period 1974–1994, strong evidence is found in favour of the existence of co-integration between the nominal exchange rate, relative money supply, relative income and relative interest rates. The monetary model is validated as a long-run equilibrium condition.

Keywords: Greece; monetary model; exchange rate determination.

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1 Introduction

The past three decades have seen enormous growth in the literature on exchange rate economics. Given the importance attached to the exchange rate in the success or failure of an open economy, it is not surprising that exchange rate economics is one of the most heavily and interesting research areas in the discipline. Exchange rate movements are perhaps the most important factors affecting sales and profit forecasts, capital budgeting plans and the value of international investments. In this respect, changes in exchange rates have a significant impact on the world's political and economic stability and the welfare of individual countries.

A European currency that has not received much attention in the finance literature before its withdrawal is the Greek drachma. Greece is a full member of the European Union (EU) and a member of the Economic Monetary Union (EMU) and its currency was added firstly to the European Currency Unit (ECU) on 17 September, 1984 and secondly it has been withdrawn after the entrance of Greece into the EMU in June 2000 by the introduction of Euro in the market on 1st January, 2001.

We have decided to run the econometric analysis for the time period 1974–1994 and not before or afterwards because 1974 saw the collapse of dictatorship in Greece, and in 1994 and afterwards the Central Bank of Greece decided to adopt a managed floating exchange rate policy as a last step before its entrance to the ERM and EMU, and the consequent withdrawal of the drachma in order to adopt the Euro (Table 1). More specifically, between end 1994 and early 1998, the Bank of Greece and the government decided to follow a policy of managing the exchange rate by not allowing the drachma to depreciate as fast as was required to compensate for the inflation differential *vis-a-vis* its trading partners. This was done deliberately in order to put a break on inflation and inflationary expectations which in the past were strongly fed by the recurrent devaluations of the drachma that, sometimes, more than fully accommodated the excessive wage and price rises. While there was some real appreciation of the drachma, the resulting disequilibrium was not as large as some argued, especially participants in the financial markets. Studies conducted by the Bank of Greece indicated that the real overvaluation of the currency was of the order of 10% after taking into account the Balassa-Samuelson effect and a number of other factors. This anti-inflationary exchange rate policy started having its impact on inflation after less than one year and was accompanied by other policy measures, with the result that domestic disequilibria were narrowing and the policy was gradually gaining credibility (Thomopoulos, 2004).

The monetary model of exchange rate determination will be viewed as a theory of long-run equilibrium, most appropriate for economies experiencing major monetary shocks. In testing the theory of long-run equilibrium, the concepts and tests of co-integration are suitable. Since the application involves a multivariate relation, it is necessary to employ maximum likelihood-based tests that allow for more than one co-integrating vector. We will test the monetary approach by implementing Johansen's (1988) maximum likelihood procedure. Unlike the co-integration technique of Engle and Granger (1987), this procedure allows for the existence of more than one co-integrating vector, a possibility that can occur in a multivariate framework. MacDonald and Taylor (1994a), argued that modelling and forecasting the exchange rate is a hazardous occupation. The object of this paper (based on the work done by MacDonald and Taylor, 1994a), is to demonstrate that at least one of the main exchange rate models, namely the monetary model, does not behave as badly as is widely thought if it is given better treatment. We will use a multivariate co-integration technique to test for the existence of a long-run relationship supporting the monetary equation.

Table 1 Exchange rate regimes for Greece

December 22, 1965–March 8, 1975	De facto band around the US Dollar: band width is ±2%. Officially pegged to the US Dollar
March 8, 1975–November 1977	De facto crawling peg to US Dollar: officially pegged to a basket of currencies
December 1977–June 1981	De facto crawling band around US dollar: band width is ±2%. Parallel market premia rises in this period and hits 31% in early 1981
July 1981–August 1984	Managed floating
September 1984–August 1989	De facto band crawling band around DM: band width is ±2%
September 1989–end 1994	De facto peg to DM
1995–March 1998	Managed floating exchange rate (strong drachma policy)
March 1998–end 2001	On March 15, 1998 the drachma entered the ERM I at a central rate of 357 per ECU. 1/1/1999 Greece entered the ERM II at a central rate of 353.109 per euro. January 2001, Greece entered into the EMU
January 1, 2002–	Currency union: Euro

Source: Reinhart and Rogoff (2002, p.69) and Bitzenis' modifications

Using quarterly data covering the period 1974–1994 of the exchange rates of the Greek drachma and US dollar, strong evidence is found in favour of the existence of co-integration between nominal exchange rates and a vector of explanatory variables. Furthermore, statistical testing of restrictions on the coefficients in the monetary model leads to the rejection of the restrictions.[1] The conclusion is that the monetary model can still be a valid representation of the long run behaviour of exchange rates and that the restrictions imposed on the model are, in general, not valid and may have been a factor contributing to the failure of the model in previous studies.

Diamandis and Kouretas (1996) used the Johansen-Juselius multivariate co-integration technique and an unrestricted monetary model and analysed the exchange rates between Greek drachma and four bilateral rates (against Deutschemark, franc, dollar, and sterling) and concluded that the null hypothesis for co-integration was accepted for the Deutschemark-Greek drachma exchange rate, and the Franc-Greek drachma exchange rate, and it was rejected for the Dollar-Greek drachma exchange rate and for the Sterling-Greek drachma exchange rate (p.361). The time period chosen was April 1975–February 1994 (they used monthly data for the period). We have used a similar but extended period January 1974–December 1994 for the Greek drachma-US dollar exchange rate only. Our results differ from the previous results as we found strong evidence in favour of the existence of co-integration between the nominal exchange rate, relative money supply, relative income and relative interest rates. The monetary model is validated as a long-run equilibrium condition.

The outline of the remainder of this paper is as follows: Section 2 the models of exchange rate determination are discussed. In Section 3, the exchange rate determination model's empirical validity is presented. The purpose of this section is to consider the econometric validity of some popular models of exchange rate determination. Finally, in Section 4, the flexible-price unrestricted monetary model for the case of Greece is reported. The data set used and some multivariate co-integration tests are also presented in this section. Section 5 concludes the paper.

2 Models of exchange rate determination

Early contributions to the post-war literature on exchange rate determination include Nurkse (1945) and Friedman (1953). Both of these contributions are, to a large extent, concerned with the role of speculation in foreign exchange markets. Nurkse warns against the dangers of 'bandwagon effects', which may generate market instability (Bilson, 1981; Frankel and Froot, 1987; Allen and Taylor, 1990).[2] Friedman's classic apologia for floating exchange rates is remarkable in its anticipation of much of the literature of the following two decades and is still cited as the seminal paper on stabilising speculation.

2.1 Mundell-Fleming model

The first model was an open Keynensian model which had been developed in its essentials by Meade (1951). Meade laid the foundations for simultaneous analysis of internal and external balance in an open economy. The central focus of Meade's analysis was on the conditions that had to be satisfied if a country was to succeed in achieving, simultaneously, internal balance and external balance. The major conclusion was that this requires the use of two policy instruments, with differentiated effects on income and the balance of payments. This model was further developed in a series of papers by Mundell (1961–1963) and Fleming (1962), and became to be known as the Mundell-Fleming model, which has become one of the more resilient models, partly because it is a formal extension of the equally resilient IS-LM system. Mundell and Fleming followed Meade's mathematical representation and thus abstracted from the stock-flow implications of interest rate differential changes. An important assumption of this model is that the domestic price level can vary with respect to the world price level, and that, therefore, the 'law of one price' does not apply. The Mundell-Fleming model provides a solution to the problem of securing simultaneously internal and external balance, solely by means of the appropriate mix of monetary and fiscal policies, and without recourse to changes in exchange rates, tariffs or any other balance of payments policies. The Mundell-Fleming model is based on a flow theory of capital movements; it provides a self-correcting mechanism for the balance of payments whereby deficits or surpluses generate changes in the money stock which, in turn, restore external equilibrium via the interest rate and real income changes. The insight of the model was that net excess demand for foreign exchange is just the overall balance of payments (current plus capital account). Under a free float, this must be equal to zero in equilibrium. Combining this equilibrium condition with standard equilibrium conditions for the goods market (the IS curve) and the money market (the LM curve) it is then possible to solve for the exchange rate and the other endogenous variables (normally real output and the interest rate) and to determine the comparative static effects of fiscal and monetary policy. The integration of asset markets and capital mobility into open-economy macroeconomics was a major innovation of the Mundell-Fleming model, but it contains a fundamental flaw: it is cast almost entirely in flow terms. The Mundell-Fleming model was largely rejected because of this flaw. In particular, it allows current account imbalances to be offset by flows across the capital account, without any requirement of eventual stock equilibrium in the holding of net foreign assets (MacDonald, 1988, Chapters 4 and 5).

2.2 Flexible monetary model

The monetary approach to the exchange rate starts

> "from the definition of the exchange rate as the relative price of two monies and attempts to model the relative price in terms of the relative supply of and demand for those monies." (MacDonald and Taylor, 1989, Chapter 1)

The early, flexible-price monetary model relies on the twin assumption of continuous Purchasing Power Parity (PPP) and the existence of stable money demand functions for the domestic and foreign economies. The flexible-price monetary model is, implicitly, a market-clearing general equilibrium model in which continuous PPP among national price levels is assumed.

The basic flexible-price monetary model equation is:

$$S_t = (m^s - m^{s^*})_t - \phi\, y_t + \phi^* y_t^* + \lambda r_t - \lambda^* r_t^* .$$

In this model an increase in the domestic money supply, relative to the foreign money stock, will lead to a rise in exchange rate S_t, that is, a fall in the value of the domestic currency in terms of the foreign currency. An increase in domestic output, as opposed to the domestic money supply, appreciates the domestic currency. Furthermore, a rise in domestic interest rates depreciates the domestic currency. A relative rise in domestic real income, other things equal, creates an excess demand for the domestic money stock. In an attempt for agents to increase their real money balances, domestic residents reduce expenditure, and prices fall until money market equilibrium is achieved. Via PPP, falling domestic prices (with foreign prices constant) imply an appreciation of the domestic currency in terms of the foreign currency. Similarly, depreciation follows from an increase in the domestic interest rate as this reduces the domestic demand for money. (This result is in sharp contrast to the view where an increase in the home interest rate leads to an appreciation of the exchange rate) (MacDonald, 1988, pp.93–94). Because the domestic interest rate is endogenous in the flexible-price monetary model, however, it is not completely logical to consider increases in r_t which are independent of movements in r_t^*, or domestic or foreign monies or incomes. A further assumption of the model is that uncovered interest parity holds continuously that is, the domestic-foreign interest differential is just equal to the expected rate of depreciation of the domestic currency. The expected change in the exchange rate and the expected change in the interest differential, both of which reflect inflationary expectations, are interchangeable in this model.

2.3 Sticky-price monetary model

The very high volatility of real exchange rates during the 1970s float, obviously refuting the assumption of continuous PPP, led to the development of the sticky-price monetary approach to the exchange rate determination. Sticky-price monetary models, originally initiated by Dornbusch (1976), allow short-term overshooting of the nominal and real exchange rates above their long-run equilibrium levels. The main features of the sticky price model would be captured in a framework in which the domestic currency prices of domestic goods are sticky but domestic currency prices of foreign goods can move with the exchange rate. This model is a representation of long-run equilibrium towards which the economy tends to adjust, while in the short run it is possible that the exchange rate may overshoot its long-run equilibrium value. Because goods prices are sticky in the short run, a cut in the nominal money supply implies an initial fall in the real money supply and a consequent rise in interest rates in order to clear the money market. The rise in domestic interest rates then leads to a capital inflow and an appreciation of the nominal exchange rate, which, given sticky prices, also implies an appreciation of the real exchange rate. Foreign investors suffer a foreign exchange loss when the proceeds of their investment are reconverted into their local currency. Since the expected rate of depreciation must be non-zero for a non-zero interest differential, the exchange rate must overshoot its long-run equilibrium PPP level. Short-run equilibrium is achieved when the expected rate of depreciation is just equal to the interest differential. In the medium run, however, domestic prices begin to fall in response to the fall in money supply (MacDonald and Taylor, 1993). This alleviates pressure in the money market and domestic interest rates begin to decline. The exchange rate then depreciates slowly in order to converge at the long-run PPP level. This model explains the paradox that countries with relatively high interest rates tend to have currencies whose exchange rate is expected to depreciate. The initial rise in interest rates leads to a steep appreciation of the exchange rate, after which a slow depreciation is expected in order to satisfy uncovered interest parity.

2.4 Real interest differential monetary model

Frankel (1979) argued that a shortcoming of the Dornbusch sticky-price monetary model was that it did not allow a role for differences in secular rates of inflation. He argued that changes in the nominal interest rate reflect changes in the tightness of monetary policy. When the domestic interest rate rises relative to the foreign rate, it is because there has been a contraction in the domestic money supply relative to domestic money demand without a matching fall in prices. The higher interest rate at home than abroad attracts a capital inflow, which causes the domestic currency to appreciate instantly. Thus, we get a negative relationship between the exchange rate and the nominal interest differential. As a consequence of the flexible price assumption, changes in the nominal interest rate reflect changes in the expected inflation rate. When the domestic interest rate rises relative to the foreign interest rate, it is because the domestic currency is expected to lose value through inflation and depreciation. Demand for the domestic currency falls relative to the foreign currency, which causes it to depreciate instantly. This is a rise in the exchange rate, defined as the price of foreign currency. Thus, there is a positive relationship between the exchange rate and the nominal interest differential. Frankel (1979) suggests an equation of exchange rate determination in which the spot rate is expressed as a function of the relative money supply, relative income level, the nominal interest differential (with the sign hypothesised negative), and the expected long-run inflation differential (with the sign hypothesised positive).

2.5 Portfolio-balance model

At the end we have the portfolio balance model that assumes imperfect substitutability between domestic and foreign assets. In common with the flexible-price and sticky-price monetary models, the level of the exchange rate in the portfolio balance model is determined, at least in the short-run, by supply and demand in the markets for financial assets. The portfolio balance model is an inherently dynamic model of exchange rate adjustment, which includes in its terms of reference asset markets, the current account, the price level, and the rate of asset accumulation. The portfolio balance model, like the sticky-price model, allows one to distinguish between short-run equilibrium and the dynamic adjustment to long-run equilibrium. It also allows for full interaction between the exchange rate, the balance of payments, the level of wealth, and stock equilibrium (MacDonald, 1988). In the short-run, in the portfolio balance model the exchange rate is determined solely by the interaction of supply and demand in asset markets.

3 Empirical evidence for the models of exchange rate determination

The *Mundell-Fleming* model was largely rejected on a priory grounds as a serious contender for the explanation of exchange rate movements at the beginning of the recent float (MacDonald and Taylor, 1994a). Johnson (1958) had stressed the distinction between stock and flow equilibria in the open-economy context and this was to become a hallmark of the monetary approach to balance of payments analysis and subsequently, the monetary approach to the exchange rate. Indeed, since an exchange rate, by definition, is the price of one country's money in terms of that of another, it is perhaps natural to analyse the determinants of that price in terms of the outstanding stocks of and demand for the two monies.

The empirical evidence on the monetary exchange rate model can be divided into two periods. The first-period relates to studies of the interwar period and of the recent float until about 1978 and is largely supportive of the *monetary* model. The second-period, which covers the period of the recent float extending beyond the late 1970s, is not so supportive of the monetary model.

One of the first tests about the monetary model was conducted by Frenkel and Johnson (1976a) for the Deutschemark-US dollar exchange rate over the period 1920–1923. Frenkel and Johnson (1976b) reported results supportive of the *flexible-price* model during this period. An early attempt to estimate the monetary model was conducted by Bilson (1978) with the exchange rate between the Deutschemark and the Pound Sterling as the dependent variable over the period January 1972 through April 1976. Bilson incorporated dynamics into the model and used a Bayesian estimation procedure. His results were in broad accordance with the monetary approach. Bilson's unrestricted estimates revealed coefficients that were mostly insignificantly different from zero. Hodrick's (1978) tests for the flexible price model, for the US Dollar-Deutchemark and Pound Sterling- US dollar over the period June 1972 to June 1975 were also highly supportive.

Putnam and Woodbury (1979) estimated the model for the Sterling-Dollar exchange rate over the period 1972–1974, and reported that most of the estimated coefficients were significantly different from zero at the 5% significance level, and all were correctly signed according to the flexible-price model. However, the money supply term was significantly different from unity.

Dornbusch (1979) reported results broadly supportive of the flexible-price model for the mark-dollar exchange rate over the period March 1973 to May 1978, in a specification incorporating the long-term interest rate differential. Although Dornbusch introduced this differential as an econometric expedient, an interpretation may be placed on this term that is consistent with Frankel's *real interest differential* equation. Thus, Frankel (1979) introduced the real interest differential model for the Mark-Dollar exchange rate over the period July 1974–February 1978. He argued that a shortcoming of the Dornbusch (1976) formulation of the sticky-price monetary model was that it did not allow a role for differences in secular rates of inflation. His model was an attempt to allow for this defect, and the upshot was an exchange rate equation that included the real interest rate differential as an explanatory variable. Frankel argued that since the coefficients of the interest rate and expected inflation terms were both significant, both the flexible and sticky-price models were rejected in favour of the real interest differential model. After this, estimates of the real interest differential model reported by Dornbusch (1980), Haynes and Stones (1981), Frankel (1984) and Backus (1984) cast serious doubt on its ability to track the exchange rate in-sample: few coefficients were correctly signed; the equations had poor explanatory power as measured by the coefficient of determination; and residual autocorrelation was a problem. In particular, estimates of monetary exchange rate equations for the Deutche Mark-US Dollar for the post-1978 period often report coefficients that suggest that a relative increase in the domestic money supply leads to a rise in the foreign currency value of the domestic currency (exchange rate appreciation). Frankel (1982) called this phenomenon – the price of the mark rising as its supply is increased – "the mystery of the multiplying marks". He attempted to explain the mystery of the multiplying marks by introducing wealth into money demand equations. By including home and foreign wealth in his empirical equation, Frankel came up with a monetary approach equation that fit the data well and in which all variables, apart from the income terms, were correctly signed and most were statistically significant. Dornbusch (1980) estimated the monetary model for the Deutchemark-US dollar exchange rate and found that, even with the coefficient of relative money stocks constrained to unity, the estimates did not support the model. He concluded that current-account developments and portfolio shifts arising from limited substitutability among securities were important additional determinants of exchange rates.

Hacche and Townend (1981) estimated a monetary model for the effective Pound-Sterling exchange rate. The only coefficient that was significantly different from zero was the one on official intervention, which does not enter the model unless the assumption of Unbiased Interest Parity (UIP) is relaxed. Rasulo and Wilford (1980) and Haynes and Stone (1981) have suggested that the root of the problem may be traced to the constraints imposed on relative monies, incomes and interest rates. The imposition of such constraints may be justified on the grounds that if multi-collinearity is present, constraining the variables will increase the efficiency of the coefficient estimates. However, Haynes and Stone showed that the constraints used in the monetary approach equations were particularly dangerous because they could lead to biased estimates and sign reversals.

Frankel (1983) presented estimates for the Mark-Dollar rate and found coefficients that were mostly either not significant or of the incorrect sign; he characterised the results as a 'disaster'. Frankel (1984, p.242) concluded that "the presence of wrong signs … and the predominance of low significance levels render the results discouraging for the monetary equation". Backus (1984) estimated a number of models for the exchange rate between the US and Canadian dollars. He found evidence that the monetary models were excessively restricted; Durbin-Watson coefficients were extremely low, and tests against a more general model rejected the restrictions.

An attempt to rehabilitate the monetary model led to the development of a second generation of monetary models, the *sticky-price monetary* model. Dornbusch (1976) introduced the overshooting model which was further developed by Buiter and Miller (1981). They allowed for a non-zero rate of core inflation and considered the impact of natural resource discoveries on output and the exchange rate. Driskill and Sheffrin (1981) presented an estimate of an equation representative of the Dornbusch (1976) overshooting model for the Swiss Franc-US Dollar rate for the period 1973–1977 and reported results largely favourable to the sticky-price model. Wallace (1979) reported results supportive of the model for the float of the Canadian Dollar against the US Dollar during the 1950s. Backus (1984) finds little support using US-Canadian data for the period 1971–1980. Unlike Wallace, he found few statistically significant coefficients. Hacche and Townend (1981) provided a more dynamic version of the sticky-price model. But the estimated equation was unsatisfactory: many coefficients were insignificant and wrongly signed, and the equation did not exhibit sensible long-run properties. Papell (1988) argued that the price and exchange rate dynamics underlying the Dornbusch sticky-price model cannot be captured by single-equation estimation methods. To capture such dynamics, he argued, it is necessary to use a systems method of estimation that incorporates the cross-equation constraints derived from the structural equations and the assumption of rational expectations. His procedure allows domestic income and interest rates to be modelled endogenously, but not the money supply. Papell used the exchange rates of Germany, Japan, the UK and the USA for the period 1973–1984, and found that most of the estimated structural coefficients had the expected sign, were of reasonable magnitude, and were statistically significant. Barr (1989) and Smith and Wickens (1989) empirically implemented a version of the sticky price model formulated by Buiter and Miller (1981). All reported favourable in-sample estimates of the model.

Testing the monetary model beyond 1978 produced poor results in terms of the signs and significance of the coefficients. In the period since 1987, several studies have been conducted to test the long-run properties of the monetary model using co-integration, specifically the two step procedure suggested by Engle and Granger (1987). These studies – which include Baillie and Selover (1987), Frankel and Meese (1987) and Kearney and MacDonald (1990) – failed to find co-integration between the exchange rate and either relative money supplies and relatives prices, or the vector of standard explanatory variables in the monetary model. Moreover, it has been generally found that exchange rates, relative money supplies and relatives prices are $I(1)$; relative levels of real income are $I(0)$ with trend; and that short-term interest rate differentials are $I(0)$. These findings are even more harmful, since the failure of the monetary model was earlier attributed to its nature as a long-run model.

Driskill and Sheffrin (1981) argued that the poor performance of the monetary model could be traced to a failure to account for the simultaneity bias introduced by having the expected change in the exchange rate on the right-hand side of the monetary equations. One potential method of circumventing such simultaneity is offered by the rational expectations solution of the monetary model, which effectively yields an equation purged of the interest differential – forward exchange rate effect. Kearney and MacDonald (1990) estimated the monetary model for the Australian dollar-US dollar and could not reject the restrictions implied by the rational expectations hypothesis.

As noted by Boughton (1988a), a further explanation for the failure of the monetary approach equations may be traced to the relative instability of the underlying money demand functions and the simplistic functional forms that are normally implicitly assumed for money demand. MacDonald and Taylor (1991), using multivariate co-integration techniques, tested the validity of the monetary model. They explained the failure of some other researchers to find co-integration between the exchange rate and other variables to the inadequacy of the Granger (1986)[3] two-step method. The alternative approach suggested by them is to test for co-integration using a more appropriate multivariate technique, the Johansen technique.[4] They found evidence in favour of the existence of co-integration for three exchange rates (the Sterling, Mark and Yen against the Dollar). They concluded that "... in contrast to the findings of other researchers ... an unrestricted monetary model does provide a valid explanation of the long-run nominal exchange rate ...". MacDonald and Taylor (1992) argued that the flexible monetary model "explains the paradox that countries with relatively high interest rates tend to have currencies whose exchange rate is expected to depreciate ...". They re-examined the flexible monetary model in unrestricted form for the Sterling/Dollar exchange rate using monthly data and found evidence in favour of the existence of up to three significant co-integrating vectors between the exchange rate, relative money supply, industrial production and long-term interest rates. They also showed that some coefficient restrictions are rejected when imposed on the full set of the co-integrating vectors. Recently, MacDonald and Taylor (1993), applied multivariate co-integration analysis and dynamic modelling techniques to a number of exchange rates and found some evidence to support the monetary model as a long-run equilibrium toward which the exchange rate converges, while allowing for complicated short-run dynamics.

MacDonald and Taylor (1994a) re-examined the monetary model for the sterling-dollar exchange rate, using a multivariate co-integration technique and they found that the unrestricted monetary model is a valid framework for analysing the long-run exchange rate. They demonstrated that there were up to three statistically significant co-integrating vectors between the exchange rate and domestic and foreign money supplies, industrial outputs and long-term interest rates. They also indicated with their tests for a unit root, that all series are $I(1)$ processes.[5] Using the Johansen technique, they rejected the hypothesis that there are no co-integrating vectors: there would appear to be up to three such relationships. Their finding of at least one co-integrating vector indicates that the monetary model would seem to have some long-run validity.

Much less empirical work has been carried out on the *portfolio balance* approach to the exchange rate than on the monetary class of models, presumably because of the problems which researchers have encountered in mapping theoretical portfolio balance models into real-world financial data. Compared to the monetary approach to the exchange rate, less empirical work has been conducted on the portfolio balance model, perhaps due to the limited availability of good disaggregated data on non-monetary assets. Branson et al. (1977) were supportive of the Portfolio Balance model. They estimated a log-linear version of the model for the Deutche mark-US Dollar exchange rate over the period August 1971–December 1976. However, they dropped the terms relating to domestic and foreign bond holdings. Later, they re-estimated their model using two-stage, least-squares and reported more satisfactory estimates; however, residual autocorrelation remained a problem. A log-linear exchange rate equation was estimated for the longer period August 1971–December 1978, for the Mark-Dollar, but the results did not differ significantly from the earlier ones; again persistent auto-correlation was a problem. Branson et al. (1979) estimated the model for five currencies against the mark for a variety of different sample periods over the 1970s. Although their results seemed supportive of the portfolio balance model, in terms of statistically significant and correctly signed coefficients, a note of caution must again be sounded, since the residuals in their ordinary-least-squares equations were all highly auto-correlated. Bisignano and Hoover (1982), argued that the portfolio balance approach should be implemented using only bilateral data for foreign assets, and to be consistent, domestic and foreign bond holdings should be included in the reduced form of the model.[6] Incorporating such modifications in their estimates of the portfolio balance equation[7] for the Canadian Dollar-US Dollar over the period March 1973 to December 1978, they reported moderately successful econometric results; in particular, they showed that it is wrong to neglect domestic and foreign non-monetary asset stocks in exchange rate reduced forms.

Dooley and Isard (1982) were the first to attempt to construct data on domestic and foreign bond holding without assuming that the current account deficit is financed entirely in one of the two currencies under consideration. They estimated their model for the Dollar-Mark exchange rate over the period May 1973 through June 1977. They pointed out that their model was better than the forward rate as a predictor of the change in the exchange rate.

Boughton (1988b) introduce term-structure effects into an empirical portfolio balance model and estimated jointly a ‘semireduced form’ consisting of a real exchange rate portfolio balance equation that includes long-run and short-term interest rates, an equation for the short-term rate, and a forecasting equation for the long and short-term interest rate spread. He used the Dollar-Yen and Dollar-Mark exchange rates for the period 1973–1985. His estimation results were broadly satisfactory in terms of the sign and statistical significance of the estimated coefficients. MacDonald and Taylor (1994b), using data on the Dollar-Franc 1976–1990 and a variant of the Campbell-Shiller technique for testing present-value models, demonstrated that the static monetary equation has some long-run validity. By assuming monetary exchange-rate fundamentals, the speculative-bubbles hypothesis was rejected and the forward-looking rational expectations restrictions were also rejected. Kanas (1997) examined whether the monetary exchange rate model represents a long run relationship among nominal exchange rates, money supplies, interest rates and real incomes of five countries that participate in the ERM. The results also strongly support the hypothesis of co-integration for all the ten ERM country-pairs considered. Furthermore, multiple co-integrating vectors are found for all cases. These results can be interpreted as evidence that the monetary model represents a stable long-run relationship for all ERM countries considered.

Makrydakis (1998) examined the monetary model of exchange rate determination as a long-run equilibrium of the Korean Won-US Dollar rate using monthly data from 1980 to 1995 and concluded that the unrestricted version of the monetary model provides a valid framework for analysing long-run movements in the exchange rate. Miyakoshi (2000) re-examined the flexible-price monetary approach to the exchange rate between the Korean Won and the three currencies: the US Dollar, the German Mark and the Japanese Yen. He concluded that at least one co-integrating vector exists; some popular monetary restrictions on this model are valid for the Korean Won-German Mark rate and the Korean Won-Japanese Yen rate. Especially, all variables in the model are correctly signed and mostly statistically significant for the Korean Won-German Mark rate. Dutt and Ghosh (2000) examine the validity of the monetary approach under both fixed and flexible exchange rates for the Japanese Yen-US Dollar exchange rate and they find support for the results by MacDonald and Taylor (1991, 1993).

Francis et al. (2001) using Canadian-US Dollar data provide evidence for strong support for the long-run monetary model of exchange rates. Groen (2002) investigated the validity of the monetary exchange rate model for both US Dollar and Pound Sterling exchange rates of Germany, Japan, Switzerland and either the UK or the USA with purely time series-based co-integration techniques as well as panel data-based co-integration techniques. Empirically the validity of the monetary exchange rate model implies both a co-integration rank restriction and parameter restrictions on the corresponding co-integrating vector. Both types of restrictions were accepted for the two four-country samples within the VEC approach. In contract, neither the Johansen approach nor the panel Engle-Granger approach is able to accept both types of restrictions. Civcir (2003) applied the monetary models to the Turkish Lira/US Dollar Exchange rate and concluded that the sticky price versions of the monetary model support the hypothesis of co-integration and the fully dynamic out-of-sample forecast from the equilibrium-correcting monetary models significantly outperforms forecasts from random-walk models and differenced vector autoregressive models.

4 The flexible monetary model: the Greek case

In this paper, we follow the approach adopted by MacDonald and Taylor (1994a) for the flexible-price monetary model. The objective of this paper is to employ modern econometric techniques to test the long-run properties of an unrestricted version of the monetary model.

The quarterly data, relating to the Greek Drachma-US Dollar exchange rate and Greek and US macroeconomic variables, are all taken from IFS, OECD periodicals and data tapes and run from January 1974–December 1994.

The basic monetary model[8] can be written in unrestricted stochastic form as follows:

$$s = \beta_1 m + \beta_2 m^* + \beta_3 y + \beta_4 y^* + \beta_5 r + \beta_6 r^* + \varepsilon$$

where

$$\beta_1, \beta_4, \beta_5 > 0, \;\; \beta_2, \beta_3, \beta_6 < 0.$$

and

s: spot exchange rate Drachma/Dollar (direct quote),[9] EXRAT
m: narrow money supply M1 for Greece, GRM1
m^*: narrow money supply M1 for USA, USAM1
y: industrial production sa for Greece, INDGR
y^*: industrial production sa for USA, INDUSA
r: Greek long run interest rate, GRLIR
r^*: US long run interest rate, USLIR
ε: disturbance term.

All variables, apart from the interest rate terms, are expressed in natural logarithms (LN).

MacDonald and Taylor (1992) argued that using the OLS testing procedure, the coefficients of the monetary model all have wrong signs. Using the same procedure, we found the same results.

Ordinary least squares estimation

Dependent variable is LNEXRAT			
84 observations used for estimation from 74Q1 to 94Q4			
Regressor	*Coefficient*	*Standard error*	*T-Ratio* (*Prob.*)
CON	–8.6694	1.8820	–4.6064 (0.000)
LNGRM1	–0.20076	0.14152	–1.4186 (0.160)
LNUSAM1	2.6141	0.33461	7.8126 (0.000)
LNINDGR	0.95126	0.38816	2.4507 (0.017)
LNINDUSA	–1.3915	0.40366	–3.4471 (0.001)
GRLIR	–0.037944	0.011107	–3.4164 (0.001)
USLIR	0.057882	0.012832	4.5109 (0.000)
R-squared	0.96977	F-statistic $F(6, 77)$	411.6625 (0.000)
R-bar-squared	0.96741	SE of regression	0.13000
Residual sum of squares	1.3012	Mean of dependent variable	4.4717
SD of dependent variable	0.72012	Maximum of log-likelihood	55.8441
DW-statistic	0.59262		

Using the Johansen procedure, taking the co-integrating vector which corresponds to the second largest eigenvalue, we have the correct signs:

$$S_t = 0.12391\ \text{GRM1} - 0.496\ \text{USAM1} - 0.0609\ \text{INDGR} + 0.0042\ \text{INDUSA}$$
$$+ 0.185\ \text{GRLIR} - 0.326\ \text{USLIR}.$$

Some commonly imposed monetary restrictions: $H_1 : \beta_1 = -\beta_2$, $H_2 : \beta_1 = 1$, $H_3 : \beta_2 = -1$, $H_4 : \beta_3 + \beta_4 = 0$, $H_5 : \beta_5 + \beta_6 = 0$.

Specifically, the exchange rate used is expressed as Drachma per Dollar (direct quote), the chosen monetary aggregate is the narrow money supply $M1$, the income measure is the industrial production, the long-term interest rates are the government bond yield rates for the USA and the treasury bill for Greece. For both countries the industrial production series are seasonally adjusted. Because of collinearity between short-term and long-term interest rates, we included only long-term interest rates in the co-integration analysis. We have chosen the period 1974–1994 with quarterly data because we agree with the belief by MacDonald and Taylor (1994a, p.288) that the total length of the sample period, rather the frequency of observation is the important factor when examining the long-run properties of time series.

According to Lane (1991) there are a lot of reasons for the apparent failure of the monetary model. His approach was to detect the reasons for the failure of the monetary model, with the alternative of constructing a theoretical model that is consistent with available empirical evidence, including that on the stationarity of the variables and their order of integration. His conclusion was that "it is perhaps less surprising that the monetary model has failed empirically than that it ever appeared to succeed at all". On the other hand, MacDonald and Taylor (1991; 1992) explain the failure of the monetary model in the way in which various researchers tried to find co-integration between exchange rates and other variables using an inappropriate testing method (*Engle and Granger two-step procedure*). This method has been criticised on the following grounds:

First, although Engle and Granger (1987) proposed seven tests of co-integration, the most widely used tests are the DF and ADF test for the stationarity of the empirical residuals derived from the co-integrating regression. But, Phillips and Perron (1988) argued that methods like the Dickey-Fuller (DF), and Augmented Dickey-Fuller (ADF), do not test adequately for the existence of unit root, and they have low power. Moreover, the distribution of the test statistics is not invariant with respect to the presence of a non-zero mean or a time trend. Since there is no clear-cut and consistent evidence on the robustness of alternative unit roots, we have to depend on the results of the DF and ADF tests for the purpose of determining the order of integration of the variables. However, we do not need to depend on the same method to test for co-integration by testing the stationarity with the Engle and Granger method. Secondly, the method makes the implicit assumption that the co-integrating vector is unique. Finally, the method produces results that are not invariant with respect to the direction of normalisation that is the choice of the dependent variable.

The Engle-Granger method has, to a large extent, been replaced by the multivariate technique developed by Johansen (1988; 1989) and Johansen and Juselious (1990). The *Johansen technique* has the following merits: first, it fully captures the underlying time series properties of the data. Secondly, it provides estimates of all of the co-integrating vectors that may exist among a vector of variables, and offers test statistics for the number of co-integrating vectors which has an exact limiting distribution. Thirdly, it allows direct hypothesis testing on the coefficients of the co-integrating vectors, a facility for testing the restrictions imposed on the monetary model. Finally, this test may, therefore, be viewed as more discerning in its ability to reject a false null hypothesis.

4.1 Testing the flexible monetary model

Before we start to analyse the data doing the necessary tests for units roots and co-integration we should provide some basic definitions of time series analysis that might be helpful for the testing of the model.

Gujarati (1995) argued that

> "any time series data can be thought of as being generated by a stochastic or random process; and a concrete set of data can be regarded as a realisation of the stochastic process."[10]

In *Stochastic Models* we assume that the time series has been generated by a stochastic process. If we suppose a time series $Y(t)$ where $t = 1, ..., n$ we can say that it is a stochastic model if the series $Y1, Y2, ..., Yn$ is drawn randomly from a probability distribution.

To apply standard inference procedures in a dynamic time series model we need the various variables to be stationary, since the majority of econometric theory is built upon the assumption of *stationarity*.

A stochastic process is said to be stationary if its mean and variance are constant over time and the value of covariance between two time periods depends only on the distance or lag between the two time periods and not on the actual time at which the covariance is computed. If the characteristics of the process change over the time then the process is *non-stationary*. A special case of a non-stationary process, in which we can difference[11] one or more time the series and produce a stationary process, is called the homogeneous non-stationary process. The number of times that the original series must be differenced (integrated) before a stationary series is produced is called order of integration. If a time series is differenced once and the differenced series is stationary, we say that the original series is *integrated of order 1*, denoted by $I(1)$. Similarly, if the original series has to be differenced twice (the first difference of the first difference) before it becomes stationary, the original series is *integrated of order 2*, or $I(2)$. In general, a time series that needs to be integrated d times before it becomes stationary, is an $I(d)$ series.

Engle and Granger (1987) discuss the main differences between processes that are $I(0)$ and $I(1)$. They point out that an $I(0)$ series:

- has finite variance which does not depend on time
- has only a limited memory of its past behaviour
- tends to fluctuate around the mean
- has autocorrelation that decline rapidly as the lag increases.

For the case of an $I(1)$ series, the main features are:

- the variance depends upon time and goes to infinity as time goes to infinity
- the process has an infinitely long memory
- it wanders widely
- the autocorrelation tend to one in magnitude for all time separations.

A graphical way of having a first evidence[10] of stationarity is based on the so-called autocorrelation *function* (ACF) and the correlogram. If we plot the auto-correlation function of a variable, and see that it starts at a very high value and declines very gradually, then it is likely to have a non-stationary process. The ACF at lag κ, denoted by $\rho_k = \gamma_k / \gamma_0 =$ covariance at lag k / variance (if $\kappa = 0$, $\rho_0 = 1$).

Since both covariance and variance are measured in the same units of measurement, ρ_k is a pure number. It lies between –1 and +1, as any correlation coefficient does. If we plot ρ_k against k, the graph we obtain is known as the *population correlogram*. Since the sample autocorrelation function at lag κ is $\hat{\rho}_k = \overline{\gamma}_k / \overline{\gamma}_0$, which is simply the ratio of the sample covariance to sample variance, a plot of $\hat{\rho}_k$ against κ is known as the *sample correlogram*.

A particular important case is the distinction between Trend Stationary Process and Difference Stationary Processes. A *Trend Stationary Process* has the form

$$y_t = \beta_0 + \beta_1 T + \varepsilon_t$$

in which the stationarity of the series is defined around a deterministic trend: de-trending the series we obtain a stationary process.

A *Difference Stationary Process* has the following non-stationary form:

$$y_t = \alpha_0 + y_{t-1} + \varepsilon_t.$$

These two processes generate very similar series that are difficult to distinguish at first inspection. For example, the series which we have to analyse could be described alternatively by one of the two processes. One of the meanings of the Unit Root test is to distinguish between these two specifications. Particularly, we test the length of the memory of the series: if the series present unit root, then every innovation (every random shock in the series) produces a permanent effect in the series. On the other hand, if the series is a Trend Stationary process, any innovation has only a temporary effect, and the series tends to return at the long run trend.

4.1.1 Testing for unit roots

Unit root tests (Harris, 1995; Gujarati, 1995)[12]

An alternative test of stationarity is known as the *unit root test*. The easiest way to introduce this test is to consider the following model:

$$Y_t = Y_{t-1} + u_t$$

where u_t is the stochastic error term that is known as a *white noise error term*.

In order to test for the presence of unit roots, and hence the order of integration of individual series, a number of statistical tests may be used. We are going to perform the tests proposed by Dickey and Fuller (1981).

Dickey Fuller and Augmented Dickey Fuller tests

The most widely used tests are those developed by Dickey and Fuller. The basic Dickey-Fuller (DF) statistic to test the order of integration of the time series y_t is based on the regression:

$$y_t = \alpha_0 + \rho y_{t-1} + u_t \quad [13]$$

that is usually convenient to re-parameterise as follows:

$$\Delta y_t = \alpha_0 + (\rho - 1) y_{t-1} + u_t .$$

The model can be extended as

$$\Delta y_t = \alpha_0 + \alpha_1 T + (\rho - 1) y_{t-1} + u_t$$

in order to allow for the possible presence of a deterministic trend in the process for y_t.

The *t*-statistics on $(\rho - 1)$ is then used to test the null hypothesis that this coefficient is equal to zero (i.e., that $\rho - 1$ and there is a unit root in the process).

The critical value for *t*-statistics is computed by Dickey and Fuller for any different model specifications. If the computed absolute value of *t*-statistics exceed the DF absolute critical value, then we reject the hypothesis that the series presents a Unit Root. Furthermore, Dickey and Fuller suggest a number of ϕ-type statistics to test multiple restrictions on the different parameters of the previous regression. The characteristics of this ϕ-test are given in Table 2 for which the critical values are showed in Dickey and Fuller (1981).

Table 2 Models and *F*-type statistics

Null model	*Alternative model*	*Test*
$y_t = y_{t-1} + u_t$	$y_t = \alpha_0 + \rho y_{t-1} + u_t$	Φ_1
$y_t = y_{t-1} + u_t$	$y_t = \alpha_0 + \alpha_1 T + \rho y_{t-1} + u_t$	Φ_2
$y_t = \alpha_0 + y_{t-1} + u_t$	$y_t = \alpha_0 + \alpha_1 T + \rho y_{t-1} + u_t$	Φ_3

The Dickey-Fuller tests are based on the assumption that the variable follows a simple first-order autoregressive process, and that the disturbance term is independently and identically distributed. When a problem of serial correlation appears, the Dickey-Fuller equation should be modified to take account of a more complex structure of the time series analysed.

For this reason the Augmented Dickey-Fuller test is developed introducing lagged difference terms of the dependent variable in the regression, in the form:

$$\Delta y_t = \alpha_0 + \alpha_1 t + (\rho - 1) y_{t-1} + \sum_{i=1}^{n} \gamma_i \Delta y_{t-i} + u_t$$

where *n* is chosen so as to ensure that the residuals are white noise: $\Delta y_{t-i} = (y_{t-i} - y_{t-i-1})$.

We have decided to proceed with checking for unit roots with an approach that essentially follows the indications illustrated by Dickey and Fuller in their seminal paper (Dickey and Fuller, 1981). Our methodology is:

- estimate the Dickey-Fuller equation with drift and trend
- check the stability of the residuals
- if the residual are not white noise, pass to the Augmented DF equation adding difference lags until the residuals are white noise[14]
- test for unit-root and for restrictions on the parameters (τ-type and φ-type tests).

(a) Exchange rates

The first step of the analysis is to estimate the full Dickey and Fuller (1981) model in the form:

$$\text{DEXRAT} = \text{CON} + \beta_1 T + (\rho - 1)\text{PREXRAT} + u_t.$$

The analysis of the residual from this regression shows a certain grade of autocorrelation, therefore we pass to analyse the ADF equation. We need to add two difference lags before we find stability in the residuals. Then the equation estimated is:

$$\text{DEXRAT} = \text{CON} + \beta_1 T + (\rho - 1)\text{EXRAT} + \sum_{i=1}^{2} \gamma_i \text{DEXRAT}_{t-i} + u_t$$

and the results obtained are:

Regressor	*Coefficient*	*Standard error*	*T-ratio*
CON	–0.38119	1.8182	–0.20965
T	0.34647	0.14595	2.3739
EXRAT(–1)	–0.10575	0.051945	*–2.0359*
DEXRAT(–1)	0.037536	0.10875	0.34515
DEXRAT(–2)	–0.25514	0.11014	–2.3166
R-squared	0.15114		
R-bar-squared	0.10646		
DW-statistic	1.9255		

First of all we test the unit root hypothesis using the t-statistics on EXRAT(–1). If the test rejects the hypothesis of unit root (H_0: non stationary, H_1: stationary), it is not necessary to proceed to test joint hypotheses on the parameters. In this case, the test can not reject the hypothesis of unit root at 95% significance. (Here, we have $|t_{\text{statistics}}| < |t_{\text{critical}}|$, therefore we fail to reject the null hypothesis, i.e., we have unit roots, considering that $|t_{\text{statistics}}| = 2.0359$ and $|t_{\text{critical}}| = 3.45$).

As has already been stressed above, autocorrelation functions are used graphically to produce a first evidence of having unit roots.

From the autocorrelation functions of EXRAT and DEXRAT we can conclude for the ACF of EXRAT, starting from 1, that it declines very gradually. On the other hand, the ACF of DEXRAT declines very suddenly and moves around the mean. Thus, EXRAT may be a non-stationary process and DEXRAT may be a stationary process. In other words, we have evidence that the variable EXRAT is an $I(1)$ process.

Then we consider different specifications of the ADF equation, testing jointly unit root, drift and trend with the f-type test previously illustrated. (H_0: no drift, no trend, unit roots).

We first test the more general restriction:

$$H_0 : (\alpha_0, \alpha_1, \rho) = (0, 0, 1)$$

$$\Phi_2 = \frac{RSS_R - RSS_U / 3}{RSS_U / 81} = 20.2735.$$

The value of the test exceeds the critical value 4.88. The hypothesis is rejected at 95% of significance.

Then we pass to test the hypothesis that the series presents a unit root and drift but no deterministic trend

$$H_0 : (\alpha_0, \alpha_1, \rho) = (\ \alpha_0, 0, 1\ )$$

$$\Phi_3 = \frac{RSS_R - RSS_U / 2}{RSS_U / 81} = 6.8717$$

that leads to reject the hypothesis (critical value 95% = 6.49).

Then we proceed to the φ_1 test:

Wald test of restrictions imposed on parameters

Based on OLS regression of DEXRAT on:				
CON	T	EXRAT(–1)	DEXRAT(–1)	DEXRAT(–2)
81 observations used for estimation from 74Q4 to 94Q4				
Coefficients A1–A5 are assigned to the above regressors, respectively				
List of imposed restriction(s) on parameter(s):				
a1 = 0				
a3 = 0				
Wald statistic	CHI-SQ(2) = *4.3835* (0.112)			

We fail to reject the hypothesis H_0: there is unit root, no drift (critical value 95% = 6.71). After applying these three tests we confirm that we have unit roots, including in the regression, only trend. We used two lags of the dependent variable EXRAT in order to relax the problem of autocorrelation. We run the OLS regression and we confirm that still have unit roots:

Ordinary Least Squares estimation

Dependent variable is DEXRAT			
81 observations used for estimation from 74Q4 to 94Q4			
Regressor	*Coefficient*	*Standard error*	*T-Ratio* (*Prob*)
T	0.33073	0.12438	2.6590 (0.010)
EXRAT(–1)	–0.10216	0.048737	*–2.0962* (0.039)
DEXRAT(–1)	0.035186	0.10750	0.32731 (0.744)
DEXRAT(–2)	–0.25715	0.10903	–2.3585 (0.021)
R-squared	0.15065	F-statistic $F(3, 77)$	4.5524 (0.005)
R-bar-squared	0.11755	SE of regression	7.1973
Residual sum of squares	3988.6	Mean of dependent variable	2.5902
SD of dependent variable	7.6617	Maximum of log-likelihood	–272.7526
DW-statistic	1.9262		
Diagnostic tests			
Test statistics	*LM version*	*F version*	
A: serial correlation	CHI-SQ(4) = 7.4043 (0.116)	F(4, 73) = 1.8361 (0.131)	

(b) Greek Money Supply (M1)

Next, we have the variable GRM1. We found the need for nine lags for the elimination of autocorrelation and heteroscedasticity. Through the tests, we understand that we must have drift and trend. We fail to reject the null hypothesis (H_0: non stationary), considering that $|t_{\text{statistics}}| = 2.6783$ and $|t_{\text{critical}}| = 3.45$.

(c) US Money Supply (M1)

And for this variable we fail to reject the null hypothesis, considering that $|t_{\text{statistics}}| = 2.7942$ and $|t_{\text{critical}}| = 3.45$.

We first test the more general restriction:

$$H_0 : (\alpha_0, \alpha_1, \rho) = (0, 0, 1)$$

$$\Phi_2 = \frac{RSS_R - RSS_U / 3}{RSS_U / 79} = 11.6010.$$

The value of the test exceeds the critical value 4.88. The hypothesis is rejected at 95% of significance.

Then we pass to test the H_0: there is unit root and drift but no deterministic trend.

$$H_0 : (\alpha_0, \alpha_1, \rho) = (\alpha_0 , 0, 1)$$

$$\Phi_3 = \frac{RSS_R - RSS_U / 2}{RSS_U / 79} = 8.3306$$

that leads to reject the hypothesis (critical value 95% = 6.49).

Finally we have the ϕ_1 test:

Wald test of restrictions imposed on parameters

Based on OLS regression of DUSAM1 on:				
CON	T	USAM1(–1)	DUSAM1(–1)	DUSAM1(–2)
DUSAM1(–3)		DUSAM1(–4)		
79 observations used for estimation from 75Q2 to 94Q4				
Coefficients A1–A7 are assigned to the above regressors, respectively				
List of imposed restriction(s) on parameter(s):				
a1 = 0				
a3 = 0				
Wald statistic		CHI-SQ(2) = *8.6997* (0.013)		

Therefore, we conclude that both trend and drift exist. In order to diminish the problem of autocorrelation we add four lags to the dependent variable USAM1.

(d) Greek industrial production (seasonally adjusted)

Following the same steps we have found these results:

Ordinary Least Squares estimation

Dependent variable is DINDGR			
83 observations used for estimation from 74Q2 to 94Q4			
Regressor	*Coefficient*	*Standard error*	*T-ratio (Prob.)*
CON	6.9080	3.5118	1.9671 (0.053)
T	0.012675	0.018913	0.67016 (0.505)
INDGR(–1)	–0.076995	0.044909	*–1.7145* (0.090)
R-squared	0.058392	*F*-statistic *F*(2, 80)	2.4805 (0.090)
R-bar-squared	0.034852	SE of regression	2.2375
Residual sum of squares	400.5097	Mean of dependent variable	0.32289
SD of dependent variable	2.2775	Maximum of log-likelihood	–183.0886
DW-statistic	2.3266		
Diagnostic tests			
Test statistics	*LM version*	*F version*	
A: serial correlation	CHI-SQ(4) = 7.5889 (0.108)	*F*(4, 76) = 1.9120 (0.117)	

The test cannot reject the hypothesis of unit root at 95% significance. (Here, we have $|t_{\text{statistics}}| < |t_{\text{critical}}|$, therefore we fail to reject the null hypothesis, i.e., we have unit roots, considering that $|t_{\text{statistics}}| = 1.7145$ and $|t_{\text{critical}}| = 3.45$.)

Considering the ϕ tests we have these results:

$$H_0 : (\alpha_0, \alpha_1, \rho) = (0, 0, 1)$$

$$\Phi_2 = \frac{RSS_R - RSS_U / 3}{RSS_U / 83} = 6.6895.$$

The value of the test exceeds the critical value 4.88. The hypothesis is rejected at 95% of significance.

Doing the second test, $H_0 : (\alpha_0, \alpha_1, \rho) = (\alpha_0, 0, 1)$ we found that

$$\Phi_3 = \frac{RSS_R - RSS_U / 2}{RSS_U / 83} = 4.961$$

the null hypothesis that there is unit root, drift, but no trend cannot be rejected (critical value 95% = 6.49). We do not need to add difference lags to find stability in the residuals, since there was no indication of autocorrelation.

We run the regression again including only drift in the regression:

Ordinary Least Squares estimation

Dependent variable is DINDGR			
83 observations used for estimation from 74Q2 to 94Q4			
Regressor	*Coefficient*	*Standard error*	*T-Ratio (Prob.)*
CON	5.1108	2.2597	2.2617 (0.026)
INDGR(–1)	–0.051703	0.024258	*–2.1314* (0.036)
R-squared	0.053106	*F*-statistic *F*(1, 81)	4.5428 (0.036)
R-bar-squared	0.041416	SE of regression	2.2299
Residual sum of squares	402.7581	Mean of dependent variable	0.32289
SD of dependent variable	2.2775	Maximum of log-likelihood	–183.3210
DW-statistic	2.3738		
Diagnostic tests			
Test statistics	*LM version*	*F version*	
A: serial correlation	CHI-SQ(4) = 7.3325 (0.119)	*F*(4, 77) = 1.8654 (0.125)	

We have unit roots, considering that $|t_{\text{statistics}}| = 2.1314$ and $|t_{\text{critical}}| = 2.89$.

(e) US industrial production (seasonally adjusted)

Following the same procedure we found that we need to add one lag to the INDUSA. Through the Wald tests we concluded that there is drift, trend and unit root. Considering the ϕ tests we have these results:

$$H_0 : (\alpha_0, \alpha_1, \rho) = (0,0,1)$$

$$\Phi_2 = \frac{RSS_R - RSS_U / 3}{RSS_U / 82} = 17.3951.$$

The value of the test exceeds the critical value 4.88. The hypothesis is rejected at 95% of significance.

Doing the second test,

$H_0 : (\alpha_0, \alpha_1, \rho) = (\alpha_0, 0, 1)$ we found that

$$\Phi_3 = \frac{RSS_R - RSS_U / 2}{RSS_U / 82} = 13.0649$$

that leads us to reject the hypothesis (critical value 95% = 6.49).

Doing the Φ_1 test we also reject the hypothesis:

Wald test of restrictions imposed on parameters

Based on OLS regression of DINDUSA on:	
CON T INDUSA(–1) DINDUSA(–1)	
82 observations used for estimation from 74Q3 to 94Q4	
Coefficients A1–A4 are assigned to the above regressors, respectively	
List of imposed restriction(s) on parameter(s):	
a1 = 0	
a3 = 0	
Wald statistic	CHI-SQ(2) = *12.5586* (0.002)

Doing the OLS we found unit roots.

Ordinary Least Squares estimation

Dependent variable is DINDUSA			
82 observations used for estimation from 74Q3 to 94Q4			
Regressor	*Coefficient*	*Standard error*	*T-Ratio* (*Prob.*)
CON	10.4007	2.9350	3.5437 (0.001)
T	0.088395	0.024459	3.6140 (0.001)
INDUSA(–1)	–0.14351	0.040722	*–2.5242* (0.001)
DINDUSA(–1)	0.55444	0.094417	5.8722 (0.000)
A: serial correlation	CHI-SQ(4) = 6.4370 (0.169)	*F*(4, 74) = 1.5760 (0.190)	

(f) Greek long interest rate (treasury bill rate)

We fail to reject the null hypothesis, considering that $|t_{\text{statistics}}| = 2.160$ and $|t_{\text{critical}}| = 2.89$. We have unit root, using four lags for the variable DGRLIR in order to eliminate the problems of autocorrelation. We reject the ϕ_2 test, but we fail to reject the ϕ_3 test, therefore we introduce only drift in the regression.

Ordinary Least Squares estimation

Dependent variable is DGRLIR			
79 observations used for estimation from 75Q2 to 94Q4			
Regressor	*Coefficient*	*Standard error*	*T-Ratio* (*Prob.*)
CON	0.76546	0.34330	2.2297 (0.029)
GRLIR(–1)	–0.045963	0.021280	*–2.1600* (0.034)
DGRLIR(–1)	–0.041877	0.10561	–0.39653 (0.693)
DGRLIR(–2)	–0.082998	0.10736	–0.77310 (0.442)
DGRLIR(–3)	0.0050205	0.10780	0.046570 (0.963)
DGRLIR(–4)	0.45234	0.11060	4.0899 (0.000)

R-squared	0.24953	*F*-statistic *F*(5, 73)	4.8546 (0.001)
R-bar-squared	0.19813	SE of regression	0.72895
Residual sum of squares	38.7898	Mean of dependent variable	0.12025
SD of dependent variable	0.81404	Maximum of log-likelihood	–84.0002
DW-statistic	1.7937		
Diagnostic tests			
Test statistics	*LM version*	*F version*	
A: serial correlation	CHI-SQ(4) = 3.7185 (0.445)	*F*(4, 69) = 0.85207 (0.497)	

(g) US long interest rate (Bond Government Yield)

From the ϕ_2 test, we found that there is unit root without trend and drift in the simple form of Dickey-Fuller. We have that $|t_{\text{statistics}}| < |t_{\text{critical}}|$, therefore we fail to reject the null hypothesis, i.e., we have unit roots, considering that $|t_{\text{statistics}}| = 1.4699$ and $|t_{\text{critical}}| = 1.95$).

Ordinary Least Squares estimation

Dependent variable is DUSLIR			
83 observations used for estimation from 74Q2 to 94Q4			
Regressor	*Coefficient*	*Standard error*	*T-ratio* (*Prob.*)
USLIR(–1)	–0.0010761	0.0073210	*–0.14699* (0.884)
R-squared	–0.4200E–3	*F*-statistic	NONE
R-bar-squared	–0.4200E–3	SE of regression	0.60747
Residual sum of squares	30.2595	Mean of dependent variable	0.015783
SD of dependent variable	0.60734	Maximum of log-likelihood	–75.8972
DW-statistic	1.3742		
Diagnostic tests			
Test statistics	*LM version*	*F version*	
A: serial correlation	CHI-SQ(4) = 8.0825 (0.089)	*F*(4, 78) = 2.1038 (0.088)	

Having found that all variables are *I*(1), we can now proceed to test for co-integration.[15]

4.2 Testing for co-integration

The Concept of Co-integration

We have seen in the previous section that a series y_t is said to be integrated of order d or $I(d)$, if it requires to be differentiated d times to yield a stationary series. If we have a vector x_t containing n variables, all of which are $I(d)$, the series contained in x_t are said to be co-integrated, if there exist a linear combination:

$$z_t = \alpha x_t$$

such that z_t is $I(d-b)$, where α is the co-integrating vector.

Co-integration could be viewed as the expression of a dynamic equilibrium relation between variables, which, although singularly non-stationary, maintain a stationary relation in the long period. z_t could be interpreted as a measure of the extent to which the system of the co-integrated variables are out of long-run equilibrium, and can be called the 'equilibrium error'.

The Johansen technique is fast becoming an essential tool for applied economists wishing to estimate time series models.[16]

The following is a brief exposition of the Johansen technique. Starting with a multivariate vector autoregressive representation of n variables:

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + ... + \Pi_k X_{t-k} + \varepsilon_t$$

where $t = 1, 2, ..., T$ and X_t is an $N \times 1$ vector[17] of $I(1)$ variables. $\Pi_1, \Pi_2, ..., \Pi_k$ are $N \times N$ matrices of unknown parameters. We can re-parameterise the equation as:

$$\Delta X_t = \vartheta_1 \Delta X_{t-1} + \vartheta_2 \Delta X_{t-2} + ... + \vartheta_{k-1} \Delta X_{t-k+1} - \Pi X_{t-k} + \varepsilon_t$$

where $\vartheta_i = -I + \Pi_1 + \Pi_2 + ... + \Pi_i$ and $\Pi = I - \Pi_1 - \Pi_2 - ... - \Pi_k$, Π is known as the co-integrating matrix with a rank r, such that $\Pi X_t = 0$ represents long-run equilibrium. Now if we define two $N \times r$ matrices, α and β such that $\Pi = \alpha\beta'$. It can be shown that $\beta_t' X_t \sim I(0)$ where β_i' (the ith row of the β') is one of the r distinct, linearly independent co-integrating vectors. The procedure then boils down to testing for the value of r, the number of significant co-integrating vectors on the basis of the number of significant eigenvalues of Π. For this purpose, two test statistics are used: the maximum eigenvalue test (Max) and the trace test (Trace). The Johansen methodology allows direct hypothesis tests on the coefficients entering the co-integrating vectors (Cuthbertson et al., 1992).[18]

The results of applying the Johansen technique are supportive of the long-run properties of the monetary model. On the basis of the trace and maximum eigenvalue statistics obtained using Johansen's multivariate maximum likelihood technique for estimating co-integrating relationships,[19] we may reject the hypothesis that there are no co-integrating vectors. The maximum eigenvalue test and the trace test reject the null hypothesis $r = 0$, $r \leq 1$ and $r \leq 2$.

We demonstrated that there were up to three statistically significant co-integrating vectors between the exchange rate and domestic and foreign money supplies, industrial outputs and long-term interest rates. We have followed the procedure of regarding a co-integrating vector as significant only if it is indicated to be so by both the Max and the Trace tests.

Having established the existence of co-integration, we are now in a position to proceed in testing the restrictions on the coefficients of the co-integrating vectors along the lines proposed by Johansen and Juselius (1990). The most common, and perhaps most important, restriction to test the monetary model to determine whether there is proportionality between relative monies and the exchange rate. Additionally, a number of researchers have imposed equal and opposite coefficients on relative income and interest rates terms.

According to the Wald test outcome we can reject all the proposed restrictions.

H_1: Wald test of restrictions imposed on parameters

Based on OLS regression of LNEXRAT on:				
CON	LNGRM1	LNUSAM1	LNINDGR	LNINDUSA
GRLIR	USLIR			
84 observations used for estimation from 74Q1 to 94Q4				
Coefficients A1–A7 are assigned to the above regressors, respectively				
List of imposed restriction(s) on parameter(s):				
a1 = –a2				
Wald statistic	CHI-SQ(1) = 20.1082 (0.000)			

H_2: Wald test of restrictions imposed on parameters

Based on OLS regression of LNEXRAT on:				
CON	LNGRM1	LNUSAM1	LNINDGR	LNINDUSA
GRLIR	USLIR			
84 observations used for estimation from 74Q1 to 94Q4				
Coefficients A1–A7 are assigned to the above regressors, respectively				
List of imposed restriction(s) on parameter(s):				
a1 = 1				
Wald statistic	CHI-SQ(1) = 26.3962 (0.000)			

H_3: Wald test of restrictions imposed on parameters

Based on OLS regression of LNEXRAT on:				
CON	LNGRM1	LNUSAM1	LNINDGR	LNINDUSA
GRLIR	USLIR			
84 observations used for estimation from 74Q1 to 94Q4				
Coefficients A1–A7 are assigned to the above regressors, respectively				
List of imposed restriction(s) on parameter(s):				
a2 = –1				
Wald statistic	CHI-SQ(1) = 31.8949 (0.000)			

H_4: Wald test of restrictions imposed on parameters

Based on OLS regression of LNEXRAT on:				
CON	LNGRM1	LNUSAM1	LNINDGR	LNINDUSA
GRLIR	USLIR			
84 observations used for estimation from 74Q1 to 94Q4				
Coefficients A1–A7 are assigned to the above regressors, respectively				
List of imposed restriction(s) on parameter(s):				
a3 = –a4				
Wald statistic	CHI-SQ(1) = 41.6567 (0.000)			

H_5: Wald test of restrictions imposed on parameters

Based on OLS regression of LNEXRAT on:				
CON	LNGRM1	LNUSAM1	LNINDGR	LNINDUSA
GRLIR	USLIR			
84 observations used for estimation from 74Q1 to 94Q4				
Coefficients A1–A7 are assigned to the above regressors, respectively				
List of imposed restriction(s) on parameter(s):				
a5 = –a6				
Wald statistic	CHI-SQ(1) = 12.4215[0.000]			

We found that all of the restrictions are rejected. Thus, although we have found some support for the monetary model, in that the variables are co-integrated with the exchange rate, our analysis also suggests that the relationship may not be quite as simple as the basic flexible-price monetary approach suggests. The Johansen procedure suffers from a fundamental difficulty that becomes apparent when one tries to identify separate long-run relationships. This is because the procedure only allows us to test the restrictions across all co-integrating vectors simultaneously. Given this difficulty, MacDonald and Taylor (1992) did not rule out the possibility that there might exist a co-integrating vector that satisfies some or all of the restrictions implied by the monetary model.

5 Conclusion

In this paper an attempt has been made to re-examine the unrestricted flexible-price monetary model of exchange rates for the Greek Drachma/US Dollar. Judged by the results of the co-integration tests, the monetary model does not seem to be useless as a representation of the long-run behaviour of exchange rates. Two conclusions are reached: the monetary model can still be a valid representation of the long-run behaviour of exchange rates; the restrictions imposed on the model are in general not valid and may have been a factor contributing to the failure of the model in previous studies. However, it seems that both economic and econometric theory have still a long way to go to tackle the problem of explaining exchange rate movements. This is the reason why exchange rate determination will remain one of the most interesting research areas in economics.

Our results differ from Diamandis and Kouretas (1996, p.361) results since we found strong evidence in favour of the existence of co-integration between the nominal exchange rate, relative money supply, relative income and relative interest rates, while they concluded that the null hypothesis for co-integration was rejected for the Dollar-Greek Drachma exchange rate. Similar results as ours, but using different country case studies have been found by Kanas (1997), Makrydakis (1998), Miyakoshi (2000), Dutt and Ghosh (2000), Francis et al. (2001), Groen (2002) and Civcir (2003).

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Notes

[1]MacDonald and Taylor (1994a) found that all of the imposed monetary restrictions are rejected. They found some support for the monetary model and they suggest that the relationship may not be quite as simple as the basic flexible-price monetary approach (FLMA). If the FLMA is correct, then $\beta_1\beta_2$ should equal, respectively, $\beta_1 = +1$, $\beta_2 = -1$. $\beta_3\beta_4$ should, respectively, be negative and positive with numerical values equal to income elasticities from domestic and foreign money demand functions. $\beta_5\beta_6$ should, respectively, be positive and negative with numerical values similar to those from interest rate semi-elasticities in money demand functions.

[2]See also Bilson (1981), Frankel and Froot (1987) and Allen and Taylor (1990).

[3]See Granger (1986).

[4]See Johansen (1988).

[5]For two, of the series (US industrial production and US long-term interest rate) there is some evidence of stationarity around the trend. MacDonald and Taylor argued that this evidence is slight and they considered these series to be $I(1)$ processes.

[6]The reduced form exchange rate equation is $S_t = g(M_t, M_t^*, B_t, B_t^*, fB_t, fB_t^*)$. See Branson *et al.* (1977).

[7]The portfolio balance model can be write down as follows:

$$W = M + B + SF$$

$$M = M(r, r^*)W$$

$$B = B(r, r^*)W$$

$$SF = F(r, r^*)W$$

$$M_r < 0, M_r^* < 0, B_r > 0, B_r^* < 0, F_r < 0, F_r^* > 0.$$

[8]See MacDonald and Taylor (1994a).

[9]S is exchange rate expressed as domestic currency per one unit of foreign currency.

[10]See Gujarati (1995).

[11]In the regression $\Delta y_t = (\rho - 1)y_{t-1} + u_t$, Δ is the first-difference operator.

[12]For a complete exposition, see Harris (1995) and Gujarati (1995) for an accessible introduction to the literature on non-stationarity and co-integration.

[13]If we run this regression and actually find that $\rho = 1$, then we say that the stochastic variable Y_t has a unit root. In time series econometrics, a time series that has a unit root is known as a random walk time series. For more details see Gujarati (1995, pp.718, Chapter 21).

[14]However, the more the lags added into the equation, the smaller is the power of the test.

[15]All the variables have been found to be non-stationary in their levels and stationary in first differences, which means that they are $I(1)$.

[16]See Harris (1995).

[17]We assume that this vector has a k-th order Vector Autoregressive (VAR) representation with Gaussian errors ε_t.

[18]See Cuthbertson *et al*. (1992).

[19]We use the 95% critical values of the maximum eigenvalue and the trace statistics.

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