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Testing the Intertemporal Approach to Current Account Determination: Evidence from Greece

by

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ABSTRACT

The paper investigates the empirical relevance of the intertemporal model of current account determination in the case of Greece during the 1950-1993 period. The core of the econometric analysis is based upon the recently popularised VAR testing framework of Campbell and Shiller (1987, 1991). The present value model is found to describe surprisingly well the dynamic behaviour of the Greek current account balance till 1987 yet fails to account for the full magnitude of capital flow fluctuations thereafter. This failure of the model is attributed to the detected presence of a significant policy regime shift in this year. The gradual lifting of capital controls and the deregulation of the Greek financial system, which have been taking place since the mid "80s, turn out to have increased the influence of speculative factors on the determination of short-run capital movements and possibly disrupted the long-run sustainability of current account deficits.

I. INTRODUCTION

Over the past fifteen years the bulk of theoretical work on the determination of the current account has mostly concentrated upon extending the rational expectations permanent income hypothesis model of private consumption to the open economy. The so called intertemporal model of current account determination developed out of this research programme treats the current account balance as the end product of forward looking saving and investment decisions and predicts that the current account should be equal to the expected future decline in an economy"s net output, that is GDP minus the sum of investment and government expenditure. Since its introduction by Sachs in 1981, the model has gone through various stages of elaboration and criticism gaining a prominent position as a policy analysis tool in contemporaneous open economy macroeconomics [see, Obstfeld and Rogoff (1996)].

The seminal developments in the theoretical front have inevitably spurred the interest of economists in evaluating the ability of the intertemporal model to account for the actual time-series behaviour of a country"s current account. The mainstream of the empirical research on the subject is expressed by the seminal work of Sheffrin and Woo (1990), Otto (1992), and Gosh (1995) on developed countries and Gosh and Ostry (1995) on developing ones. This line of applied work takes as a prerequisite the validity of the national intertemporal budget constraint (IBC henceforth), which is statistically expressed by means of the stationarity of the current account, and concentrates on the derivation and statistical evaluation of the *short-run* testable implications of the intertemporal model of the current account. The focal point of the approach is a reduced form expression obtained by solving the theory model which equates the current account balance to the stream of expected declines in net output. This reduced form expression embodies the joint hypothesis of the intertemporal allocation of consumption and perfect capital mobility that underlies the intertemporal model and formulates the hypothesis in question as a set of restrictions on the dynamics of a vector autoregression (VAR). The empirical evaluation of these short-run restrictions is performed in the context of the VAR based formal and informal test statistics proposed by Campbell (1987) and Campbell and Shiller (1987, 1991). The principal advantage of the Campbell-Shiller methodology lies in its ability to produce a *benchmark* current account series consistent with consumption-smoothing behaviour and perfect capital mobility against which the actual current account series can be compared. In this way not only the capability of the intertemporal model to track the cyclical fluctuations of the actual current account can be ascertained but also the variance of the benchmark current account series can help in interpreting any documented deviations from the hypothesis of perfect capital mobility that is nested within the theory model [see, Gosh (1995)].

The present paper embarks upon investigating empirically the validity of the testable implications of the intertemporal model of the current account in the case of the Greek economy. The principle scope for such an undertaking is to provide answers relating to the relevance of the intertemporal model in accounting for the evolution of current account balances and the degree of effective capital mobility in Greece for the period spanning from 1950 to 1993. This is of particular interest to policymakers because the abolition of capital and exchange rate controls and the deregulation of the domestic financial system, which have been gradually taking place since 1986 in view of the potential participation of the country to the EMU, are likely to increase episodes of capital flight complicating the formulation and implementation of macroeconomic policy in Greece [see, Christodoulakis and Karamouzis (1993), and Gibson and Tsakalotos (1993)]. The evaluation of the intertemporal model of the current account through the Campbell-Shiller formal and informal tests will provide evidence on the validity of the underlying hypotheses of consumption-smoothing behaviour and perfect capital mobility and should help us identify whether any detected deviations from the theory model over the sample period can be attributed to speculative factors determining the short-run flows of capital from and to Greece. Finally, particular care will be exercised during the implementation of the econometric analysis so that the possible impact of any significant policy regime changes that may have occurred during the sample period can be unveiled and be taken into consideration. To this end the recently proposed structural break and parameter stability cointegration tests of Gregory and Hansen (1996) and Hansen (1992) respectively are employed in determining the stationarity status of the current account measure used in this study prior to the implementation of the Campbell-Shiller framework. In the case of the Greek economy these tests happen to provide indirect evidence supporting recently expressed fears of the current account deficit being on a potentially unsustainable path.

The remainder of the paper is structured as follows: Section II presents the theoretical structure of the consumption-smoothing model as applied to the open economy. The derivation of the testable implications of the intertemporal approach to the current account as well as the key features of the techniques deployed toward the empirical evaluation of the model are discussed in Section III. Section IV deals with data issues, provides a brief discussion of institutional arrangements concerning capital and exchange rate controls in Greece, conducts the econometric analysis using annual national accounts data covering the period 1950-1993 and, finally, comments on the economic content of the findings. Section V concludes the paper.

II. THE THEORETICAL FRAMEWORK

The theoretical framework of this paper is founded upon a discrete time version of Sachs's 1982 intertemporal model of the current account. This model constitutes an extension of the rational expectations permanent income hypothesis model of private consumption to an open economy setting. On the assumption of perfect capital mobility, it predicts that the current account will absorb shocks to output, investment, and government expenditure, so that consumption can be maintained at a level consistent with the permanent income hypothesis.

In formal terms consider a small open economy represented by a single infinite lived agent who can borrow and lend freely in international capital markets at the (constant) world real interest rate r. The identity that links net foreign asset accumulation, that is the current account, to the economy's saving-investment balance and constitutes the economy's dynamic budget constraint is:

$$b_{t+1} = (1+r)b_t + (q_t - c_t - i_t - g_t) \tag{1}$$

where b_t is the economy's stock of net foreign claims at the end of period t, q_t is output or gross domestic product in period t, c_t stands for private consumption expenditure, i_t is public and private investment, and g_t denotes government consumption expenditure. All variables are expressed in real terms.

Taking expectations of Equation (1) conditional on the information available at time t and recursively eliminating the future values of the stock of real net foreign claims we are led to

$$b_t = -E_t \sum_{j=0}^{\infty} \left(\frac{1}{1+r}\right)^{j+1} \left(q_{t+j} - c_{t+j} - i_{t+j} - g_{t+j}\right) + \lim_{j\to\infty} E_t \left(\frac{1}{1+r}\right)^{j+1} b_{t+j+1} \tag{2}$$

For intertemporal budget balance the indefinite financing of interest payments on existing debt through unlimited new borrowing will not be allowed by foreign lenders, thus the no-Ponzi game condition should apply in order to exclude bubbles. This condition, which is also known as the transversality condition, states that the present value of the stock of foreign claims b_t should decline as the planning horizon gets infinitely large, i.e.

$$\lim_{j \to \infty} E_t \left(\frac{1}{1+r} \right)^{j+1} b_{t+j+1} = 0 \tag{3}$$

Imposition of the transversality condition (3) on (2) produces the national intertemporal budget constraint:

$$b_t = -E_t \sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^{j+1} d_{t+j} \tag{4}$$

where $d_t = q_t - c_t - i_t - g_t$ stands for the trade deficit. Equation (4) suggests that at any point in time the stock of net external claims equals the present discounted value of the stream of future trade deficits.

The intertemporal budget constraint faced by the representative agent can now be obtained through the use of the GDP identity, i.e.

$$c_{t+j} = q_{t+j} - i_{t+j} - g_{t+j} - d_{t+j} \tag{5}$$

Discounting each term in Equation (5) by $(1/1+r)^{j+1}$, summing over 0 to ∞, and substituting for the present discounted value of future trade deficits (or net exports) from Equation (4), the agent's IBC is expressed as

$$\sum_{0}^{\infty} \left(\frac{1}{1+r} \right)^{j+1} c_{t+j} = \sum_{0}^{\infty} \left(\frac{1}{1+r} \right)^{j+1} \left(q_{t+j} - i_{t+j} - g_{t+j} \right) + b_t \tag{6}$$

Given the intertemporal budget constraint, the description of the mechanism that gives rise to current account imbalances requires explaining the way the components of aggregate expenditure and output are determined. Let us assume that the preferences of the representative agent are given by

$$\sum_{j=0}^{\infty} \beta^{j} E_t[u(c_j)] \tag{7}$$

where β is the subjective discount rate, and $u(\cdot)$ represents the instantaneous utility function. The hypothesis that the country is small in world capital markets implies separability of consumption and investment decisions [see, Cooper and Sachs (1985) and Gosh (1995)]. This in turn suggests that planned consumption c_t^* can be chosen independently of investment and output. On the additional assumptions of autonomous government expenditure, uncovered interest rate parity, and a quadratic utility function, the optimal path for consumption can be derived by maximising (7) subject to (6). After some tedious algebra we arrive at the following expression for planned consumption c_t^*:

$$c_t^* = \frac{r}{\gamma}\left\{ b_t + (1+r)^{-1} E_t\left[\sum_{j=0}^{\infty} (1+r)^{-j} ncf_{t+j} \right] \right\} \tag{8}$$

where $ncf_t = q_t - i_t - g_t$ is the economy's net output or, in Gosh's terminology, the economy's net national cash flow and γ is a constant of proportionality reflecting the tilting dynamics of consumption [see, Sachs (1982) and Gosh (1995)].[1] Equation (8) is the permanent income hypothesis consumption function in an open economy. Permanent income or, in this case, permanent national cash flow is simply r times net national wealth, that is the term in braces in (8). Consequently, planned consumption is simply proportional to permanent national cash flow. The magnitude of γ determines whether the country tilts consumption toward the present or the future [Sachs (1982)]. Specifically, if $\gamma > 1$ the country consumes less than its permanent cash flow and, in effect, tilts consumption toward the future. Apparently the opposite holds true whenever $\gamma < 1$. Of course, when γ amounts to unity there is no consumption-titling component in (8) and consumption becomes equal to the economy's permanent cash flow.

Since the consumption-tilting parameter γ does not necessarily equal unity in empirical applications, Gosh (1995) and Gosh and Ostry (1995) argue that a reduced form expression for the current account which will be consistent with consumption-smoothing and will facilitate econometric evaluation of the present value model can be obtained by removing the consumption-tilting component from the current account

1. In the model under consideration the subjective and market discount rates are not necessarily identical so that the consumption-tilting parameter γ is equal to $\beta(1+r)r/\beta(1+r)^2 - 1$. In the special case of $\beta = 1/(1+r)$, γ will be unity suggesting no tilting dynamics for consumption.

series.[2] This is achieved by defining the optimal consumption-smoothing current account, that is the level of capital flows which is consistent with agents fully smoothing their consumption in the face of shocks to national cash flow, as

$$ca_t^* = ncfr_t - \gamma c_t \tag{9}$$

where $ncfr_t = y_t - i_t - g_t$ is net national cash flow inclusive of interest payments, and y_t stands for gross national product, that is q_t plus net factor payments rb_t. The intuition behind Equation (9) is that the consumption-tilting component is usually a non-stationary time series and subsequently its presence can complicate the empirical evaluation of the theory model through the implementation of the Campbell-Shiller methodology. The removal of this component amounts to detrending the current account variate and permits econometric analysis to focus upon the short-run fluctuations of the current account around its trend.

Substituting (9) into (8) and solving for ca_t^* yields

$$ca_t^* = -E_t\left[\sum_{j=1}^{\infty}(1+r)^{-j}\,\Delta ncf_{t+j}\right] \tag{10}$$

where $\Delta = 1 - L$ and L is the lag operator. This expression equates the consumption-smoothing current account to the discounted value of the stream of future declines in net national cash flow. This formulation is a direct extension of Campbell"s 1987 "saving for a rainy day" hypothesis since it suggests that the economy will increase (decrease) national saving by running a current account surplus (deficit) whenever it expects a *temporary* decrease (increase) in net national cash flow in the future. On the other hand, an anticipated *permanent* change in national cash flow, say due to an increase in output, will cause a one-for-one change in consumption leaving the current account unaltered.[3] Consequently, expression (10) summarises the theoretical content of the intertemporal model of current account determination and could serve as the basis for the model's empirical evaluation.

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2. Note that when the consumption-tilting parameter is set to one the current account identity results.

3. These results can be easily verified by simple comparative static analysis on the current account and consumption functions, that is Equations (8) and (10). For a comprehensive discussion of the implications of the intertemporal model of current account determination the reader is referred to the excellent survey article of Obstfeld and Rogoff (1996).

III. THE TESTABLE IMPLICATIONS OF THE MODEL

The ability of the intertemporal model to account for the short-run fluctuations in a country's current account can be investigated by assessing the extent to which the optimal (or benchmark) consumption-smoothing current account, as defined by Equation (10), coincides with the actual consumption-smoothing current account ca_t, where ca_t is defined in a manner analogous to ca_t^* as

$$ca_t = ncfr_t - \gamma c_t \tag{11}$$

For such an assessment to be performed, nonetheless, the testable implications of the present value relationship (10) should be clearly outlined.

Let us assume that the net national cash flow variable (ncf_t) is stationary in first differences then Equation (10) suggests that the optimal current account ca_t^* will be an I(0) variable in levels because it is defined as the infinite discounted sum of stationary variables. Moreover, under the null hypothesis that the present value model is true, that is $ca_t = ca_t^*$, the actual consumption-smoothing current account will also be a stationary variable due to (9). Hence, if private consumption and interest-inclusive national cash flow are stationary in first differences, the stationarity of ca_t implies that c_t and $ncfr_t$ will be cointegrated with cointegrating parameter γ. Thus, a cointegration regression between $ncfr_t$ and c_t should enable the necessary estimate of the consumption-tilting parameter γ to be obtained and the actual consumption-smoothing component of the current account to be identified. It should be noted though that the stationarity of the current account variable (prior to any decomposition into consumption-smoothing and consumption-tilting components) follows directly from the transversality condition (3) holding and, according to Trehan and Walsh (1991) and Wickens and Uctum (1993), constitutes a sufficient condition for a nation to satisfy the national IBC, that is Equation (4), and remain solvent in the *long-run* under unchanged policies. The purpose of detrending the current account through (11) and computing the consumption-smoothing component becomes now clearer. It merely intends to facilitate the evaluation of the short-run performance of the intertemporal model abstracting from the issue of *long-run* sustainability. If however the actual consumption smoothing component of the current account, which ought to be stationary by construction in view of the preceding argument, is eventually found to exhibit non-stationary behaviour (either deterministic or stochastic) then this finding should be taken as indirect evidence against the long-run sustainability of the country's international indebtedness.

The cointegration property of the data, if established, proves particularly helpful in setting up a formal test of the *short-run* theory predictions implied by the present value relationship (10). Following the Campbell (1987) and Campbell and Shiller (1987, 1991) paradigm, the representative agent is assumed to form expectations about the future values of Δncf_t by estimating a bivariate vector autoregression (VAR) of order p that involves ca_t and Δncf_t. The p-order VAR expressed for ease of exposition in companion form is:

$$\begin{bmatrix} \Delta ncf_t \\ \cdot \\ \cdot \\ \Delta ncf_{t-p+1} \\ ca_t \\ \cdot \\ \cdot \\ ca_{t-p+1} \end{bmatrix} = \begin{bmatrix} a_1 & \cdot & \cdot & \cdot & a_p & b_1 & \cdot & \cdot & \cdot & b_p \\ 1 & & & & & & & & & \\ & \cdot & & & & & & & & \\ & & 1 & & & & & & & \\ c_1 & \cdot & \cdot & \cdot & c_p & d_1 & \cdot & \cdot & \cdot & d_p \\ & & & & & 1 & & & & \\ & & & & & & & \cdot & & \\ & & & & & & & & 1 & \end{bmatrix} \begin{bmatrix} \Delta ncf_{t-1} \\ \\ \\ \Delta ncf_{t-p} \\ ca_{t-1} \\ \\ \\ ca_{t-p} \end{bmatrix} + \begin{bmatrix} u_{1t} \\ 0 \\ \cdot \\ 0 \\ u_{2t} \\ 0 \\ \cdot \\ 0 \end{bmatrix} \qquad (12)$$

where blank elements are zero. Letting $\mathbf{z}_t = \left[\Delta ncf_t \quad . \quad . \quad \Delta ncf_{t-p+1} \quad ca_t \quad . \quad . \quad ca_{t-p+1}\right]'$, the system in (12) can be written more compactly as $\mathbf{z}_t = \mathbf{A}\mathbf{z}_{t-1} + \mathbf{v}_t$. The matrix $\mathbf{A}$ is known as the companion matrix of the VAR. Taking into account that expectations are formed rationally in the underlying theoretical model, it follows that $E(\mathbf{z}_{t+i} / \mathbf{H}_t) = \mathbf{A}^i \mathbf{z}_t$ for all i, where H_t is the limited information set containing lagged values of z_t.

Under the null of the present value model being true, projection of (10) onto the information set H_t and subsequent use of (12) produces the following set of restrictions on the companion matrix of the VAR:

$$\mathbf{g}' = -\sum_{i=1}^{\infty} (1+r)^{-i} \mathbf{h}' \mathbf{A}^i \qquad (13)$$

where g and h are column vectors involving *2p* elements all of which are zero except from the *p+1* element of g and the first element of h which are unity. Since the variables entering the VAR are stationary under the null, the infinite sum in (13) should converge so that

$$\mathbf{g}' = -\mathbf{h}'(1+r)^{-1}\mathbf{A}\left[\mathbf{I} - (1+r)^{-1}\mathbf{A}\right]^{-1} \tag{14}$$

Post-multiplying (14) by $\left[\mathbf{I} - (1 + r)^{-1}\mathbf{A}\right]$ yields

$$\mathbf{g}'\left[\mathbf{I} - (1+r)^{-1}\mathbf{A}\right] = -\mathbf{h}'(1+r)^{-1}\mathbf{A} \tag{15}$$

The latter equation summarises the full set of cross-equation restrictions that the present value relationship (10) imposes on the companion matrix under the maintained hypothesis. The validity of the *2p* linear restrictions in (15) can be easily evaluated by means of a Wald test.

Another less stringent test of the intertemporal model derives from the direct implication of the present value relationship (10) that ca_t should contribute in forecasting future changes in net national cash flow [see, Campbell and Shiller (1987) and Gosh (1995)]. Stated differently, the model predicts that the current account should Granger cause subsequent changes in Δncf_t. This weaker test of the theory could be undertaken by estimating the VAR in (12) and performing a standard Granger causality test on the national cash flow equation.

The VAR framework presented above, however, is not merely useful for conducting formal statistical tests of the present value relation. It can also be employed as an informal means of gauging the economic significance of any documented statistical deviations from the consumption-smoothing model. As pointed out by Gosh (1995) the variance of the optimal consumption-smoothing current account provides a benchmark against which the variance of the actual consumption-smoothing current account can be compared. Such a comparison is likely to shed light on the relevance of the perfect capital mobility hypothesis. The latter constitutes an integral part of the maintained hypothesis as the present value relation (10) has been arrived at by postulating no impediments to capital flows. Therefore, if actual capital movements are to be consistent with the perfect capital mobility hypothesis, the variances of the actual and optimal current account series should be *equal.* If, however, actual capital flows prove to be more (less) volatile than optimal flows then effective capital mobility will be greater (smaller) than perfect. In case of capital controls being in force, the detection of comparatively deficient capital movements in relation to the benchmark value predicted by the intertemporal model provides evidence on the effectiveness of such controls. On the contrary, evidence of excessive volatility of ca_t with respect to that of ca_t^* suggests that speculative forces could be lurking behind the evolution of actual capital flows.

A test based on the variances ratio of the actual and optimal consumption-smoothing current account components can only be carried out on the proviso of constructing the optimal current account series. Given the present analytical framework, the optimal current account variable is given straightforwardly as the VAR forecast of the present-value of future changes in national cash flow. To see this recall that $E_t \Delta ncf_t = \mathbf{h}'\mathbf{A}^i \mathbf{z}_t$, thus substitution of this expression into the right hand side of (10) yields ca_t^* as a function of the companion matrix of the VAR in (12), that is

$$ca_t^* = -\mathbf{h}'(1+r)^{-1}\mathbf{A}\left[\mathbf{I}-(1+r)^{-1}\mathbf{A}\right]^{-1}\mathbf{z}_t = \mathbf{B}\mathbf{z}_t \tag{16}$$

where $\mathbf{B} = -\mathbf{h}'(1+r)^{-1}\mathbf{A}\left[\mathbf{I}-(1+r)^{-1}\mathbf{A}\right]^{-1}$ is a row parameter vector containing $2p$ elements.[4]

As a final informal check on the ability of the optimal current account to track the short-run fluctuations in its actual counterpart, the sample correlation between the variables at issue should also be examined. If actual capital flow movements are to be largely consistent with the fundamental predictions of the present value model, a high correlation between the optimal and actual consumption-smoothing components is required.

4. If the present value model is true, that is $ca_t = ca_t^*$ then $\mathbf{B} = \mathbf{g}'$.

IV. THE EMPIRICAL ANALYSIS

A. The Data Set

The empirical relevance of the intertemporal model of current account determination is investigated for the case of Greece using annual data covering the period 1950-1993.

Over most of the period under consideration capital movements have been subject to extensive restrictions while the domestic financial system has been heavily regulated. The principle aim of the Deadalian legislation in force was the prevention of large capital outflows from Greece to the rest of the world by residents and non-residents alike. These capital controls ranged from limitations on the convertibility of both current and capital account flows to restrictions on the net foreign asset and liability positions of non-bank Greek residents as well as on specific portfolio investment decisions of foreign financial institutions in Greece [see, Christodoulakis and Karamouzis (1993)]. On the other hand, no effective obstacles existed to incoming capital. In fact, the attraction of capital into the country has been encouraged by a series of appropriate regulations, while the maximisation of capital inflows has been one of the goals of the country's macroeconomic policy during the period under study [see, Maroulis (1992)]. Following the adoption by the Greek government of the Single European Act, the progressive lifting of capital and exchange rate controls and the deregulation of the domestic financial system was initiated in 1986. Recently, the need to comply with the conditions set by the Maastricht Treaty for the participation of Greece to the EMU has proliferated the process of capital movement liberalisation. Nonetheless, the final impediments to short term current account flows were not abolished till May 1994.[5] Consequently, our empirical investigation is conducted over a sample period characterised by the presence of restrictions on the free flow of capital yet the persistence of pursuing these restrictions has been relaxed toward the end of the period examined.

Various issues of the widely available *IMF International Financial Statistics* (*IFS* henceforth) have been used to compile the data series for this paper. The correspondence between the variables in the text and the data in the *IFS* country tables is as follows: a) gross national product (y_t) is line *99a*, b) gross domestic product (q_t) is line *99b*, c) private consumption (c_t) and government consumption expenditure (g_t) are lines *96f* and *91f* respectively, d) total investment expenditure (i_t) is the sum of lines *93e* and *93i*. All variables are expressed in real terms through division by the GDP deflator, that is line *99bip* in the *IFS* country tables (Base year = 1990).

5. For a historical exposition on the progress of capital movement liberalisation in Greece, the reader is referred to Maroulis (1991), Karamouzis (1992) and recent annual reports of the Governor of the Bank of Greece.

B. Unit Root and Cointegration Tests

In accordance to the methodological guidelines of Campbell and Shiller (1987, 1991) outlined in the previous section, the validity of the present value relationship (10) and the empirical relevance of the perfect capital mobility hypothesis in Greece can be assessed through the setting up of a VAR in the actual consumption-smoothing component of the current account and the first difference of interest-exclusive national cash flow (Δncf_t). Since the variables entering the VAR must be stationary time series, the primary concern of the investigator lies with determining the order of integration of the variables required for the definition of ca_t, namely $ncfr_t$ and c_t, as well as the level of the interest-exclusive national cash flow variable.

Two statistical procedures have been employed toward investigating the autoregressive representation of ncf_t, $ncfr_t$, and c_t for the presence of unit roots: the standard Augmented Dickey-Fuller (ADF) unit-root test [see, Said and Dickey (1984)] and the stationarity test of Kwiatkowski et al. (1992) or in terms of their initials KPSS thereafter. The former procedure relies on the estimation by OLS of the following auxiliary regression,

$$X_t = \zeta_0 + \zeta_1 t + \rho X_{t-1} + \sum_{i=1}^{k} \delta_i \Delta X_{t-i} + \xi_t \tag{17}$$

where X_t is the variable of interest and ξ_t is a sequence of normal independent random variables with mean 0 and variance σ^2, to evaluate through a t-ratio test (t_τ) the null hypothesis of a unit root, that is $\rho = 1$, against the alternative of stationarity around a deterministic trend. Although the t statistic (t_μ) from the simpler ADF regression which excludes the trend and tests for mean stationarity under the alternative is also computed, preference should be given to specification (17) even in the event of t being redundant. The reason is that the presence of the deterministic trend renders the unit root test invariant to the value of the drift term [see, Kiviet and Phillips (1992)]. The maximum lag length k required for correcting serial correlation in the associated ADF auxiliary regressions is selected on the basis of evidence provided by sequential t-ratio tests on the significance of the highest order lag in the estimated autoregression coupled with extensive residual diagnostic testing [see, Ng and Perron (1995)]. In contrast to the ADF statistics, the KPSS semi-parametric procedure tests for level (η_μ) or trend stationarity likely to enable more clear-cut conclusions to be drawn with regard to the order of integration of the series under investigation given the small size of our sample. Since these tests involve different maintained hypotheses, four are the possible outcomes of their joint implementation: a) if the ADF rejects and the KPSS does not, the series is said to be trend (or mean) stationary, b) if the ADF does not reject but the KPSS does, the (η_τ) against the alternative of a unit root.[6] The combined use of these testing principles is series is said to be following a random walk, c) if both statistics fail to reject, there is insufficient information in the data to discriminate between stationary and non-stationary outcomes, and d) if both tests reject inconclusiveness results. Finally, the first differences of the relevant variables are subjected to the same battery of tests in accordance to the suggestions of Dickey and Pantula (1987) so that the presence of higher order integrated processes can also be examined.

Table 1 reports the findings of the unit root and stationarity tests on the ncf_t, $ncfr_t$, and c_t aggregates and their differences. Brief inspection of the ADF statistics computed for the variables in levels reveals that the presence of at least one unit root cannot be ruled out at the 95% confidence level for all variables considered. This general conclusion is not contradicted by the corresponding η_μ and η_τ statistics which assume values that lead to outright rejections of the stationarity null. When the outcome of the application of the same statistical procedures on the first differences of ncf_t, $ncfr_t$, and c_t is examined, a second unit root in the autoregressive representation of the series at issue cannot be detected by either the ADF or the KPSS testing principle. In fact the former tests produce quite strong rejections of the unit root null while the latter fail to reject the maintained stationarity hypothesis at the usual 5% significance level. Thus all variables investigated for unit roots are found to be adequately described by I(1) processes as far as the specific sample period considered is concerned.

Since both private consumption and interest-inclusive national cash flow are found to exhibit I(1) behaviour, one can now proceed with the construction of the actual consumption-smoothing component of the current account. Recalling the discussion of Section III, this component can be derived by examining whether the aforementioned pair of variables forms a cointegrating vector. If it does then the consumption-smoothing component of the current account, which corresponds to the residuals of the cointegration regression between $ncfr_t$ and c_t, will be a zero mean stationary time series. Therefore, a cointegration regression of $ncfr_t$ on c_t should be performed and the stationarity of its residuals has to be subsequently evaluated. In contrast to the established practice of estimating such a regression by OLS, the Fully Modified (FM) OLS method of Phillips and Hansen (1990) has been used instead so that the inconsistency of the standard OLS estimates and the inapplicability of Gaussian inference on the estimated cointegrating vector can be alleviated [see, Ogaki (1993)].

6. The KPSS test statistic for level or trend stationarity is given by:

$$\eta = \frac{1}{s^2(k)T^2}\sum_{t=1}^{T} S_t^2$$

where $S_t = \sum_{i=1}^{t} u_i$, u_t are the residuals from the regression of X_t on a constant or a constant and a trend for level or trend stationarity respectively, $s^2(k)$ is the non-parametric estimate of the "long-run variance" of u_t, and k stands for the lag truncation parameter.

TABLE 1

Unit Root and Stationarity Tests

	Dickey-Fuller Tests		KPSS Tests	
Variable	t_μ	t_τ	η_μ	η_τ
c_t	1.211 [0]	-2.657 [0]	1.195^{**}	0.18^{*}
Δc_t	-5.645^{**} [0]	-5.823^{**} [0]	0.312	0.126
ncf_t	0.843 [1]	-2.311 [1]	1.199^{**}	0.194^{*}
Δncf_t	-4.889^{**} [2]	-5.068^{**} [2]	0.268	0.133
$ncfr_t$	0.48 [1]	-2.382 [0]	1.201^{**}	0.146^{*}
$\Delta ncfr_t$	-4.584^{**} [2]	-4.582^{**} [2]	0.186	0.131

Notes:

- ** and ** indicate significance at the 95% and 99% confidence levels respectively.*
- *t_μ and t_τ are the Dickey-Fuller test statistics for level and trend stationarity respectively. The asymptotic and finite sample critical values for these tests are obtained from Table 1 in MacKinnon (1991).*
- *η_μ, and η_τ stand for the KPSS test statistics for level and trend stationarity respectively. The l(4) formula of Schwert (1987) is used for the determination of the lag truncation parameter. The asymptotic and finite sample critical values for these tests appear in Table 1 in Sephton (1995).*
- *The numbers in square brackets correspond to the lagged differenced terms required to account for serial correlation in the Dickey-Fuller auxiliary regressions.*

The FM estimated regression between interest-inclusive national cash flow and private consumption expenditure reveals tilting dynamics for consumption in Greece. The estimate of the consumption-tilting parameter γ appearing in Table 2 Part A has a magnitude of 0.858. This value of γ is smaller than unity and complies with the Greek current account being in deficit over the entire sample period. As to the stationarity properties of ca_t, these have been investigated, in analogy to the approach adopted in testing for unit roots in the data, by two different testing principles, namely the Engle and Granger (1987) no cointegration standard ADF test and the recently proposed cointegration KPSS test of Shin (1994). The findings from implementing these testing principles on the residuals of the cointegration regression are slightly ambiguous (see, Table 2 Part A). The C_μ test, which stands for the KPSS statistic for mean stationarity, fails to reject the null of cointegration attaining a value of 0.151. This value is clearly below the 5% asymptotic critical value of 0.314 tabulated by Shin (1994). The conclusion of ca_t being described by a mean stationary process does not find equally forceful corroboration by the Engle-Granger ADF_μ no cointegration statistic whose value is insignificant at the same confidence level. Consequently, there seems to be insufficient sample information to enable a decision to be reached on the stationarity status of ca_t.

A probable justification for the documented failure of ADF_μ to reject the no cointegration null rests with the potentiality of ca_t undergoing a policy regime shift during the estimation period which biases ADF_μ toward no reject outcomes. The presumption of a structural break does not seem particularly farfetched since, as our earlier description of the institutional arrangements governing current account flows in Greece suggests, the gradual lifting of capital and exchange rate controls began in 1986. In order to investigate the possibility of the consumption-smoothing current account being affected by this institutional change, the variable in question is plotted in Figure 1. Examination of the overtime evolution of ca_t, i.e. the bold line in Figure 1, uncovers a parallel mean shift in the series around the mid eighties providing informal support to the claim of a policy regime change. Nonetheless, formal evidence on the significance of the impact of such a change on the stationarity status of ca_t as well as on the timing of the potentially seminal structural break requires implementation of appropriate testing procedures.

In an attempt to resolve the encountered inconclusiveness, the sequential no cointegration test of Gregory and Hansen (1996) has been applied. This test has the same null hypothesis as the standard Engle-Granger procedure but its alternative is set up as cointegration around a one time structural break of unknown timing. This test has the advantage of making no arbitrary assumptions with regard to the occurrence of structural change allowing its data based determination. In the event of the cointegrating vector suffering a parallel shift in its mean, as the pre-test diagrammatic analysis of ca_t seems to indicate, the Gregory-Hansen procedure involves estimating the following cointegration regression,

$$ncfr_t = \varphi_0 + \varphi_1 DT(\lambda)_t + \gamma c_t + \psi_t \tag{18}$$

TABLE 2

Fully Modified OLS Estimates of the Consumption Tilting Parameter and Cointegration Tests

General Cointegration Regression: $ncfr_t = \varphi_0 + \varphi_1 DT(\lambda)_t + \gamma c_t + \psi_t$

A. Sample Period: 1950-1993

Regression applied	γ	ADF_μ	C_μ	L_c	*MeanF*	*SupF*	ADF^*
No Dummy $\varphi_1 = 0$	0.858 (0.02)	-3.35 [k=0]	0.151	0.615^*	5.782^*	8.704	-
With Dummy $\varphi_1 \neq 0$	0.902 (0.006)	-	-	-	-	-	-6.563^{**} [λ=0.86] [k=2]

B. Sample Period: 1950-1987

Regression applied	γ	ADF_μ	C_μ
No Dummy $\varphi_1 = 0$	0.899 (0.007)	-6.18^{**} [k=2]	0.078

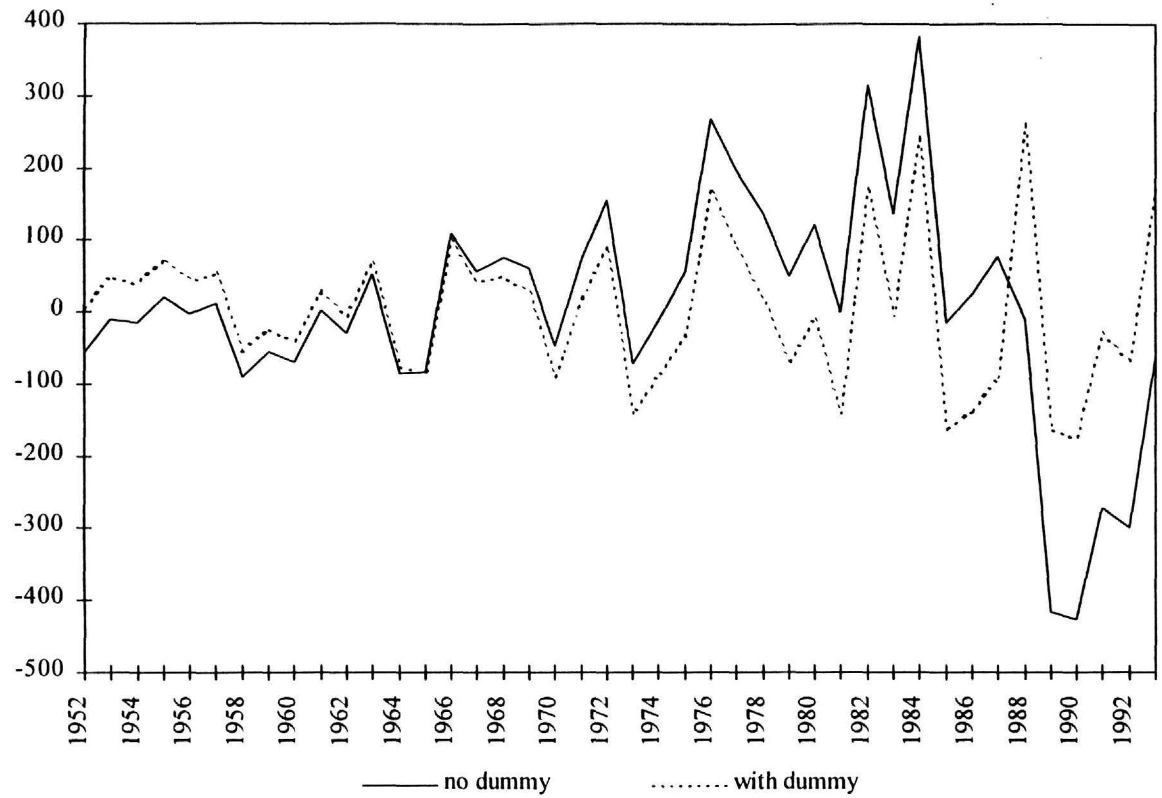
Notes:

- ** and ** indicate significance at the 95% and 99% confidence levels respectively.*
- *γ stands for the consumption tilting parameter.*
- *The cointegration regressions are estimated by the Fully Modified OLS technique proposed by Phillips and Hansen (1990).*
- *ADF_μ is the standard t ratio unit root statistic on the residuals from the no dummy cointegration regression. The critical values for the ADF_μ statistic are taken from the relevant response surface estimates of Table 1 in MacKinnon (1991). The figures in parentheses are standard errors while k corresponds to the number of lagged differenced terms in the ADF auxiliary regressions required for serial correlation correction.*
- *C_μ stands for the KPSS cointegration test when the cointegration regression involves only a constant. The l(4) formula of Schwert (1987) is used for the determination of the lag truncation parameter. The critical values for this test are obtained from Table 1 in Shin (1994).*
- *SupF, MeanF, and L_c are the Hansen (1992) tests on the parameter stability of the no dummy cointegration relationship. Asymptotic critical values for these test statistics appear in Tables 1, 2, and 3 of the Hansen (1992) article respectively.*

ADF^* is the smallest t-ratio no cointegration statistic obtained from the sequential estimation of the cointegration regression that involves the rolling regime shift dummy $DT(\lambda)_t$. λ stands for the breakpoint fraction while the asymptotic critical values for the test come from Table 1A in Gregory and Hansen (1996).

FIGURE 1

The Current Account Consumption-Smoothing Component: 1952-1993



Note:

The current account consumption-smoothing component series depicted above correspond to the residuals from the cointegration regressions appearing in Table 2 Part A.

where, $t=1,2,\ldots,T$, $\lambda = T_b / T$, T_b is the break, and $DT_t(\lambda) = 1$ if $t > T\lambda$ or 0 otherwise. The estimation of the cointegration regression is performed sequentially with the breakpoint $\lambda = T_b / T$ ranging from $2 + k / T$ to $(T-1)/T$ and k representing the number of lags required for serial correlation correction in the associated ADF auxiliary regression on the residuals from (18).[7] Each estimation run produces a value for the ADF t-ratio test. The lowest ADF t statistic obtained from this sequential procedure, denoted by ADF^* henceforth, is compared to the asymptotic critical value provided in Table 1A of Gregory and Hansen (1996). If ADF^* is found to be statistically significant, $ncfr_t$ and c_t are said to be forming a cointegrating relationship whose position has changed to a new "long-run" equilibrium at time λ^*.[8] To put it differently, ca_t is described by a stationary process undergoing a structural regime shift in its mean.

The sequence of the ADF t-ratio statistics produced through the Gregory-Hansen procedure using (18) is depicted in Figure 2, while the smallest ADF value encountered along with the associated estimate of the consumption tilting parameter γ are reported in Table 2 Part A. Examination of Figure 2 reveals that the no cointegration null is decisively rejected. The values assumed by the ADF t-statistic are clearly below the 5% asymptotic critical value line over a limited range of λ values with the lowest t-value being encountered in 1987, that is $\lambda^* = 0.86$. Specifically, the value of the ADF statistic that assigns the greatest weight to the policy regime shift alternative (i.e. ADF^*) amounts to -6.563, which is significant even at the 1% level, showing clearly that cointegration is achieved only when the policy regime change is explicitly incorporated into the analysis. This is also evident from the stationary behaviour exhibited by the consumption-smoothing current account derived from the ADF^* cointegration regression (see the dotted line in Figure 1). Moreover, Table 2 Part A shows that the consumption-tilting parameter γ is estimated with greater precision in comparison to the no dummy variable case remaining below unity as the historical evolution of the Greek current account series requires, while the timing of the structural break as determined by the Gregory-Hansen method effectively coincides with the Greek monetary authorities' announcement of progressive relaxation of capital and exchange rate controls.

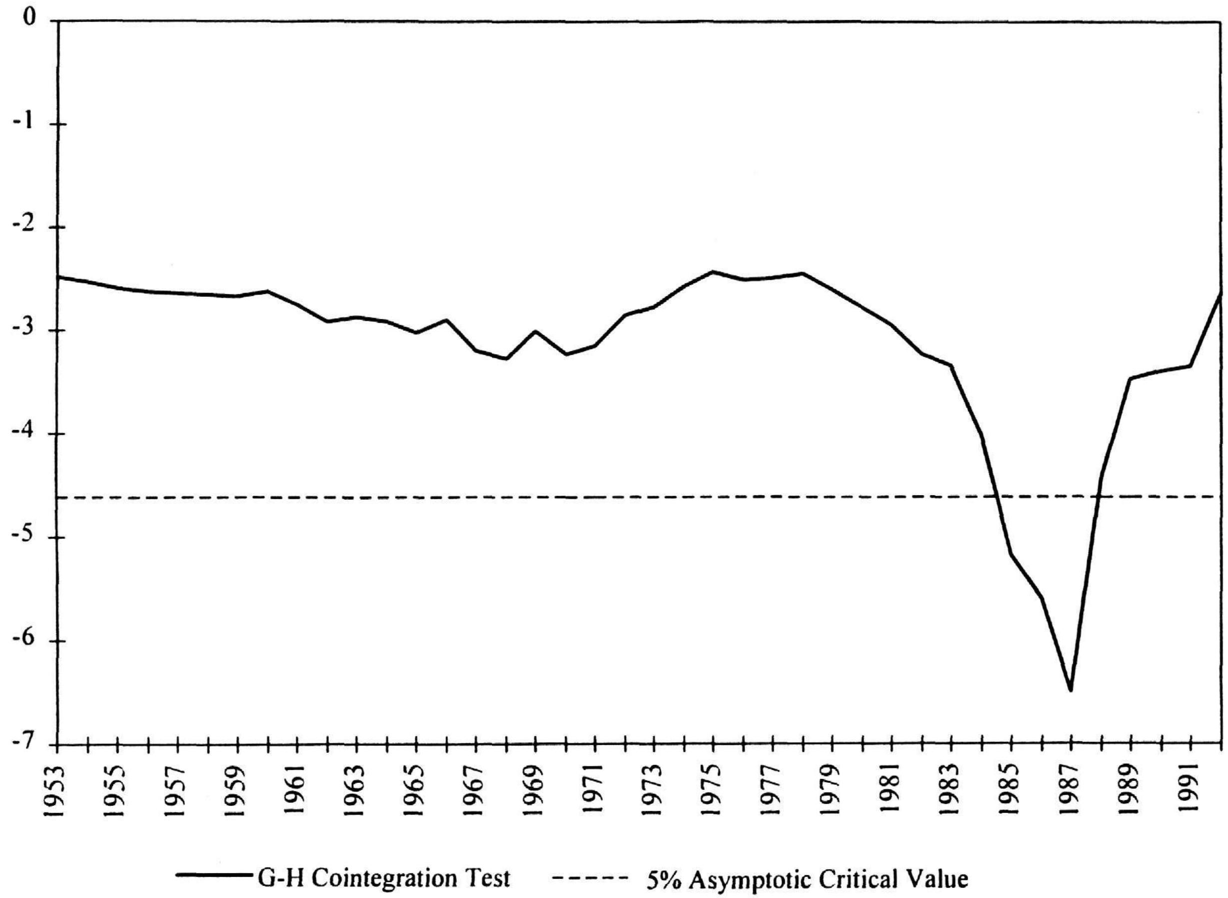
Although the Gregory-Hansen test permits the possibility of regime change to be entertained in cointegrated models, it does not answer explicitly the question of whether there has been a regime shift because it contains the standard model of cointegration with no regime change as a special case under the alternative. Gregory and Hansen (1996) recommend the complementary use of the parameter instability tests suggested by Hansen (1992) as a means of reducing the danger of erroneous inferences been drawn. Hansen (1992) proposes three tests, namely *SupF*, *MeanF*, and L_c, designed to evaluate the coefficient stability of the FM estimated standard, i.e. no dummy, cointegration

7. The number of lag differenced terms (k) required for the computation of the ADF_μ and ADF^* cointegration statistics is determined in accordance to the same criterion used in the case of the corresponding unit root tests.

8. Note that λ^* is the estimated sample break fraction corresponding to ADF^*.

FIGURE 2

Gregory-Hansen Sequential Cointegration ADF Test*



regression without imposing any *a priori* assumption of a specific break date. The null hypothesis of the cointegration relationship being stable is common to all three test statistics. Under the alternative though, the *SupF* test is targeted on uncovering any abrupt shift in the cointegrated vector, while the *MeanF*, and L_c statistics are designed to detect any gradually escalating changes in the cointegration regression coefficients.[9]

The findings from the application of the Hansen tests on the standard cointegration regression appear in Table 2 Part A. Comparing the values obtained for the *SupF*, *MeanF*, and L_c statistics with the corresponding critical values tabulated by Hansen in Tables 1, 2, and 3 of his 1992 article respectively, it is easily inferred that the cointegration relationship between $ncfr_t$ and c_t has suffered a gradual shift during the estimation period. This is so because the *MeanF*, and L_c tests are found to reject the stability null at the usual 5% significance level while the *SupF* provides no evidence of a sudden parameter change at the same significance level. Additional evidence on the presence of the 1987 regime shift come from the estimation of the no dummy cointegration regression over the 1950-1987 period. The outcome of this exercise shown in Table 2 Part B indicates that once the sample is truncated to exclude the final six observations, cointegration between $ncfr_t$ and c_t can clearly be established by both the ADF_μ and C_μ statistics.

Collectively taken the results on the time series properties of the consumption-smoothing component of the current account point to the conclusion that ca_t becomes stationary only when the detected presence of the 1987 policy regime shift is explicitly accounted for in the cointegration regression between the interest-inclusive national cash flow and private consumption expenditure. Recalling that ca_t ought to be stationary by construction due to the removal of the consumption-tilting component, the detection of the policy induced shift in the mean of the consumption-smoothing current account series signifies indirectly violation of the transversality condition (3) by the data. In other words the relaxation of capital and exchange rate controls and the deregulation of the financial system in Greece has probably set the current account on a long-run unsustainable path. Consequently, the setting up of the VAR for evaluating the intertemporal model of the current account requires ca_t to be constructed on the basis of the ADF^* Gregory-Hansen cointegration regression so that the impact of the policy regime change can be taken explicitly into consideration.

C. Formal and Informal Tests of the Model

Having established stationarity for the variables of interest, the vector autoregression in the first difference of interest-exclusive national cash flow and the consumption-smoothing current account required for the formal and informal appraisal of the present value model can now be estimated. In accordance to standard practice, the variables entering the VAR are expressed as deviations from their means and the

9. For a comprehensive description of the computational and distributional details of the *SupF*, *MeanF*, and L_c statistics the interested reader is referred to Hansen (1992).

policy regime shift dummy $DT(\lambda^*)_t$, so that only the dynamic restrictions of the theory are tested [see, Campbell (1987), Campbell and Shiller (1987, 1991) and Gosh (1995)]. The lag length of the VAR system is determined through the use of the Schwartz (1978) criterion. Furthermore, the standard errors of the VAR coefficients are corrected for heteroscedasticity using the technique of White (1980) because the variables in the VAR are not logarithmicly transformed and, in effect, heteroscedasticity emerges as a serious possibility. Table 3 Part A reports the estimated coefficients and the associated White heteroscedasticity consistent standard errors from the second order VAR eventually selected together with the computed values of the formal and informal tests of the present value model obtained for the 1950-1993 period.

Following the analysis of Section III, the first testable implication of the model under investigation relates to the ability of ca_t to forecast subsequent changes in national cash flow. If the representative agent has more information about the overtime evolution of national cash flow than that contained in this variable's past history then the consumption-smoothing component of the current account must Granger-cause Δncf_t. The results of the VAR estimation in Table 3 Part A turn out to provide strong empirical support to the Granger causality presumption and, consequently, to the underlying theory. Notice that the parameter estimates of ca_{t-1} and ca_{t-2} carry negative signs as the theoretical model would have predicted, while the coefficient estimates on the same set of variables are individually significant in terms of conventional asymptotics. More importantly, though, the standard Granger-causality $\chi^2(2)$ test on the joint significance of the coefficients in question is found equal to 23.481 leading to a profound rejection of the no Granger-causality null.

Further informal evidence supporting the intertemporal optimising theory model stems from Figure 3 Part A which presents the time-series plots of ca_t and its optimal counterpart ca_t^*. The latter quantity is computed from Equation (14) using the coefficient estimates from the unrestricted VAR and postulating *a priori* an annual real world interest rate of 4% for discounting purposes.[10] It is reminded that if the intertemporal model were true the time series plots of ca_t and ca_t^* should differ only by sampling error. Hence, statistically significant deviations of the actual current account from its VAR forecast suggest departures from the maintained hypothesis of capital flows being consistent with consumption-smoothing behaviour and perfect capital mobility. In spite of the simplicity characterising the structure of the theoretical model, the visual comparison of the two series in Figure 3 Part A reveals a surprisingly high capability of ca_t^* to track closely the year to year fluctuations of the Greek (consumption-smoothing) current account over almost the entire sample period. Nonetheless, ca_t seems to exhibit a slightly greater volatility to that considered optimal in

10. All empirical computations involving discounting have been carried out using the values of 2, 4, 6, and 8 per cent for r without significant quantitative differences in the results being detected. The eventual selection of the 4% discount rate is founded upon its equality to the average interest rate on three month US treasury bills which can be thought of as a sensible proxy for the constant world interest rate r in the model. Naturally, results unreported herein for reasons of brevity are available upon request from the author.

TABLE 3

VAR Based Tests of the Present Value Model of the Current Account

A. Estimation Period: 1952-1993

2nd order VAR	Δncf_{t-1}	Δncf_{t-2}	ca_{t-1}	ca_{t-2}	DW
Δncf_t	-0.242	0.219	-0.68	-0.372	1.96
	(0.162)	(0.129)	(0.166)	(0.195)	
ca_t	0.037	0.496	-0.164	-0.396	1.83
	(0.141)	(0.112)	(0.144)	(0.169)	
	Granger causality test		$\sigma(ca_t)/\sigma(ca_t^*)$	$corr(ca_t, ca_t^*)$	Wald test
	$\chi^2(2) = 23.481$		1.535	0.893	$\chi^2(4) = 16.909$
	[0.000007]		**<0.285>**	**<0.072>**	[0.002]

B. Estimation Period: 1951-1987

1st order VAR	Δncf_{t-1}	ca_{t-1}			DW
Δncf_t	-0.082	-0.91			1.87
	(0.143)	(0.21)			
ca_t	-0.048	-0.006			1.93
	(0.132)	(0.193)			
	Granger causality test		$\sigma(ca_t)/\sigma(ca_t^*)$	$corr(ca_t, ca_t^*)$	Wald test
	$\chi^2(1) = 18.725$		1.165	0.984	$\chi^2(2) = 2.865$
	[0.00002]		**<0.23>**	**<0.023>**	[0.2387]

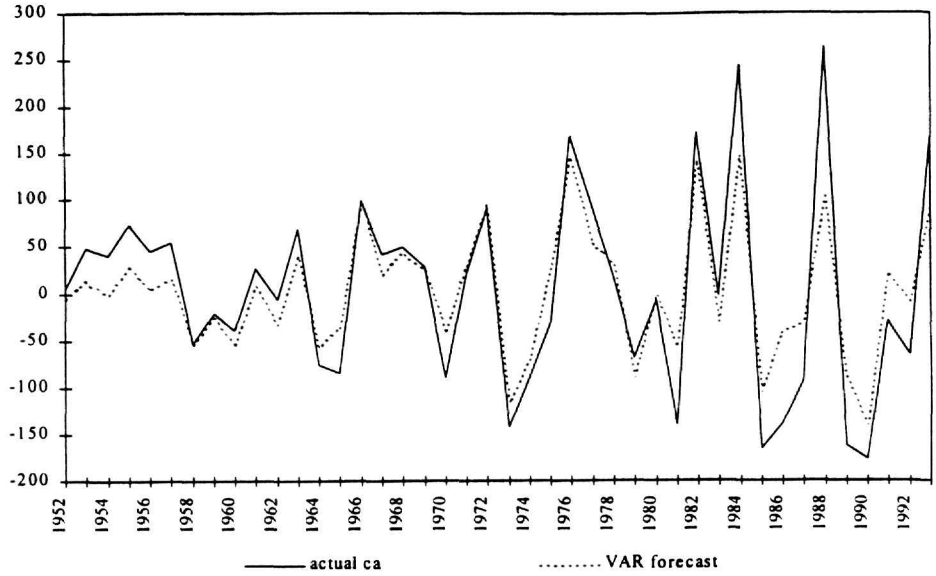
Notes:

- *The actual consumption-smoothing current account component (ca_t) used in Part A is obtained from the policy regime shift cointegration regression of Table 2 Part A while that used in Part B comes from the cointegration regression of Table 2 Part B.*
- *The VAR in Part A does not involve a constant term because the variables entering it are expressed as deviations from their means and the step regime shift dummy variable $DT(.86)_t$. The variables entering the VAR in Part B are expressed as deviations from their means.*
- *The optimal or predicted current account consumption smoothing component (ca_t^*) is obtained as a forecast from the unrestricted VAR estimation using Equation (14).*
- *DW is the Durbin-Watson statistic.*
- *The numbers in parentheses are White (1980) heteroscedastic consistent standard errors, while the numbers in square brackets are p-values.*
- *corr denotes the correlation between the actual and forecasted consumption-smoothing current account components, and σ denotes standard deviations.*
- *The Granger causality test evaluates the joint significance of the lagged ca_t terms in the Δncf_t equation.*
- *The numbers in bold are bootstrap standard errors from a non-parametric Monte Carlo simulation of 10,000 runs which retains the original VAR coefficients and draws randomly from the original estimated residuals.*
- *The $\chi^2(2p)$ Wald test evaluates the 2p linear cross-equation restrictions imposed by the intertemporal model upon the companion matrix of the VAR. The White heteroscedastic consistent variance-covariance matrix required for the construction of the Wald χ^2 test comes from the bootstrap simulation.*

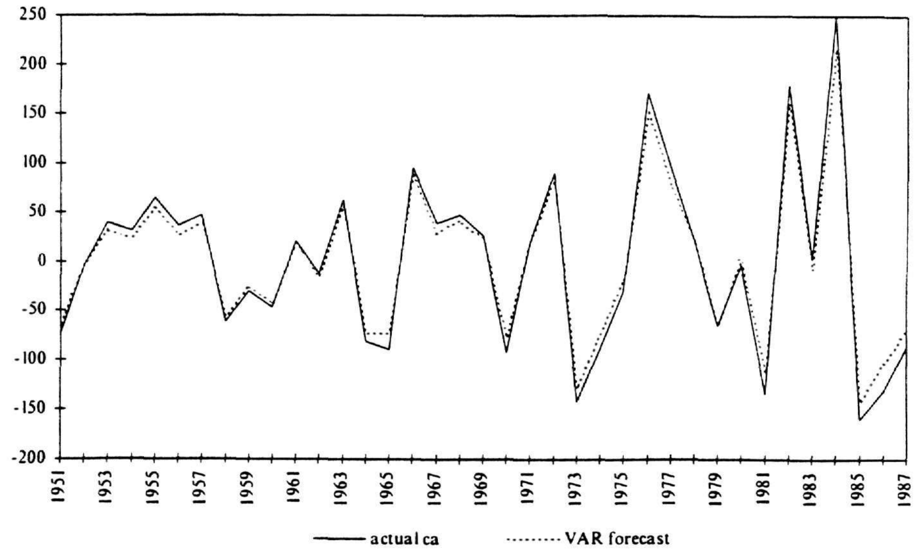
FIGURE 3

Consumption-Smoothing Current Account Versus its VAR Forecast

A. Estimation Period: 1952-1993



B. Estimation Period: 1951-1987



Note:

The actual ca_t series presented in Part A corresponds to the residuals from the cointegration regression that involves the step dummy variable $DT(0.86)_t$. The actual ca_t series in Part B is the residuals series from the cointegration regression appearing in Table 2 Part B.

terms of the model, especially toward the end of the estimation period, signifying that speculative forces may have become important in driving short-term capital flows in Greece. In order to determine whether this observation implies statistically significant departures from the consumption-smoothing model the variance ratio and correlation coefficient tests should be examined.

In computing the so called informal test statistics, particular attention has to be paid to the fact that the optimal consumption-smoothing component of the current account is a forecast and, strictly speaking, is measured with error [see, Campbell and Shiller (1987, 1991) and Hardouvelis (1994)]. In the absence of corrective action, such errors of measurement could introduce significant small sample bias into statistics formed on the basis of ca_t^*. Hodrick (1992) suggested the resort to Monte Carlo simulation techniques as a way of minimising the probability of erroneous conclusions being drawn from variance ratio tests in the context of present value models. Thus, the standard errors printed in bold underneath the values of $\sigma(ca_t)/\sigma(ca_t^*)$ and $corr(ca_t, ca_t^*)$ in Table 3 are the end product of 10,000 Monte Carlo bootstrap simulation runs. Each simulation run involves forming bootstrap errors by drawing randomly from the original VAR resid uals. Assuming initial values for the variables entering the VAR, the bootstrap errors are embedded in the originally estimated VAR to produce artificial series for the (demeaned and policy dummy adjusted) Δncf_t and ca_t aggregates. Subsequently, artificial estimates of $\sigma(ca_t)/\sigma(ca_t^*)$ and $corr(ca_t, ca_t^*)$ are compiled by reestimating the second-order VAR using the artificial time-series derived for the VAR variables. The repetition of this experiment 10,000 times yields a sequence of artificial estimates on each of the informal test statistics which can be utilised to ascertain the precision with which the statistics in question are estimated. Specifically, the bootstrap errors appearing in Table 3 reflect the standard deviations of $\sigma(ca_t)/\sigma(ca_t^*)$ and $corr(ca_t, ca_t^*)$ across the 10,000 simulation runs.

The findings of the informal tests are not as clear-cut as the Granger causality test conducted earlier unravelling potentially significant departures from the intertemporal model of current account determination. The 1.535 value reported in Table 3 Part A for the variance ratio $\sigma(ca_t)/\sigma(ca_t^*)$ carries a bootstrap standard error of 0.285 which in turn suggests that the variance ratio is just 1.87 standard deviations away from unity. Thus, the maintained hypothesis of consumption-smoothing behaviour and perfect capital mobility is rejected only at the 10% level of significance. Although the rejection of the maintained hypothesis is not particularly strong, the larger than unity variance ratio does not only indicate that capital in Greece has been on average sufficiently mobile in smoothing consumption in the face of shocks to national cash flow but also does not preclude the eventuality of speculative factors being at play in determining short-run capital movements. This piece of evidence complements and extends empirical findings on the degree of capital mobility reached by other researchers in the past. In particular, Gibson and Tsakalotos in their 1993 study on capital flight over the 1975-1987 period report the estimated volume of short-term capital inflows/outflows in Greece to be highly variable and amounting to a relatively significant proportion of the current account deficit, yet not as large as that observed for the other European countries examined in their study. Furthermore, Zarangas (1995) investigating the perfect capital mobility hypothesis in the context of a portfolio balance model over the period of fixed exchange rates (1960-1975) documents a progressive increase in the volatility of current account fluctuations since 1968 which is not nonetheless large enough to substantiate the notion of perfect capital mobility for the period studied. The gradual proliferation in the variability of short-term current account fluctuations is also clearly visible in the graphical representation of actual ca_t in Figure 3 Part A. Notice that capital movements demonstrate an ever increasing amplitude over the 1950-1993 period with the volatility of ca_t increasing more rapidly since the introduction of capital market liberalisation measures.

With regard to the correlation coefficient between the actual and optimal current account series, this is found to be relatively high standing at 0.893, as one would have anticipated judging from the visual inspection of the two series in Figure 3 Part A. In contrast though to the variance ratio, the correlation coefficient is estimated with greater precision so that the perfect correlation between the benchmark and actual current account series implied by the null hypothesis is not statistically unattainable. This is so because the $corr(ca_t, ca_t^*)$ statistic stands just one and a half standard deviations away from unity.

Finally, the formal and most stringent test of the complete set of parametric restrictions imposed on the companion matrix of the VAR by the intertemporal model is considered. Recall that if the consumption-smoothing model constitutes a valid representation of actual current account movements then Equation (15) should hold. In the context of the second order VAR adopted herein this implies that the theory restrictions on the coefficients of matrix **A** are $c_i = a_i$, $i = 1,2$, $d_2 = b_2$, and $d_1 - b_1 = 1 + r$ or, equivalently, the four element row vector **B** in Equation (16) should be identical to $[0 \quad 0 \quad 1 \quad 0]$, that is the coefficient on the currently dated ca_t should be unity and all other coefficients should be nought. The validity of the restrictions embodied in Equation (15) can be easily assessed by means of a $\chi^2(4)$ Wald test. Following Hardouvelis (1994), the Wald test statistic on the overall fit of the model is computed using the 4x4 variance-covariance matrix of the restrictions in (15) across the 10,000 bootstrap simulation runs. The $\chi^2(4)$ statistic presented in Table 3 Part A is found equal to 16.91 and is statistically significant even at the 99% confidence level suggesting an outright rejection of the intertemporal model"s restrictions on the VAR.

Taken together, the findings of the formal and informal testing procedures, with the exception of the Wald test, do not imply a resolute dismissal of the intertemporal model and the perfect capital mobility hypothesis. In fact, the consumption-smoothing model proves particularly capable in explaining the peaks and troughs of the Greek current account series over the sample period. However its ability to account for the magnitude of these cyclical fluctuations is questionable particularly over the 1987-1993 period which is associated with the introduction of the different policy regime (see, Figure 3 Part A). In order to determine whether the institutional changes introduced since 1986 are lurking behind the model's failure to satisfy the Wald and marginally the variance ratio tests, the Campbell-Shiller procedure has been reimplemented over the 1950-1987 period. The consumption-smoothing component of the current account is constructed in this case using the residuals of the no dummy cointegration regression appearing in Table 2 Part B. The estimated coefficients of the first order VAR chosen by the Schwartz criterion along with the estimated formal and informal tests statistics can be seen in Table 3 Part B. The formal and informal Campbell-Shiller tests are now *overwhelmingly* supportive of the joint hypothesis of consumption-smoothing behaviour and perfect capital mobility with the theory model being capable to account for the direction and turning points of capital movements as well as the amplitude of the ca_t"s over time fluctuations (see, Figure 3 Part B). Hence, the suspicion of the documented 1987 policy regime change being largely responsible for the mixed results produced using the full sample seems to be well founded. In addition, the simple comparison of the values assumed by the informal tests over the two estimation periods reinforces the conclusion reached earlier concerning the role of the policy regime change in increasing the incidence of short-run speculative capital flows over the 1987-1993 period. Notice that the variance ratio test is found equal to 1.165 and insignificantly different from unity in the 1950-1987 subsample increasing to 1.535 and becoming significantly different from unity in the full sample, while the ability of the model to track the actual current account series, as measured by the correlation coefficient statistic, drops from 0.984 to 0.893 between the two sample periods.

V. CONCLUSIONS

The aim of this study has been to evaluate the empirical relevance of the intertemporal model of current account determination and its nested hypothesis of perfect capital mobility in the case of Greece over the period spanning from 1950 to 1993. The theoretical framework is provided by a simple discrete time version of the present value model of the current account whose focal point is a reduced form expression linking the current account to the stream of expected future declines in the economy's net output. The testable implications of this relationship are appraised through the adoption of the recently popularised VAR based formal and informal testing methodology for present value models developed by Campbell and Shiller (1987, 1991). In contrast to the existing literature on the subject, the possible impact of policy regime changes undermining the quality of inferences drawn is given meticulous attention prior to the implementation of the Campbell-Shiller approach through the estimation of the recently developed structural break and parameter instability cointegration tests of Gregory and Hansen (1996) and Hansen (1992) respectively for the consumption-smoothing current account series.

The prediction of the intertemporal model that the country will find it optimal to increase its indebtedness, that is to run current account deficits, if it expects its national cash flow to grow on average over time finds dramatic corroboration in the case of the Greek economy over the 1950-1987 period but it is weakly refuted when the entire sample period is considered. The reason for this differentiation in the ability of the model to describe the evolution of the Greek current account seems to lie with the detection in 1987 of a deterministic shift in the mean of the consumption-smoothing current account aggregate marking the presence of a significant policy regime change in the economy. This change in regime is chiefly associated with the EU induced relaxation of capital and exchange rate controls and the deregulation of the domestic financial market which have been taking place in Greece since 1986. The findings of the VAR based formal and informal tests reveal that the institutional changes at issue appear to have increased the role of speculative forces in the determination of short-run capital movements over the 1987-1993 period in the sense that capital mobility turns out to be greater than perfect, while the deterministic shift in the mean of the consumption-smoothing account is indicative of potentially unsustainable current account deficits in the long-run. Since the short time span elapsed from the gradual introduction of the new regime qualifies as premature any argument of the agents" in Greece having had the required time to adjust fully to the new institutional environment, the outcome of the empirical analysis serves as a clear warning to policymakers over the cautiousness with which macroeconomic policy should be pursued in Greece if currency and current account crises are to be averted in the near future.

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