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Is there a Greek-Turkish Arms Race?: Evidence from Cointegration and Causality Tests

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CENTRE OF PLANNING AND ECONOMIC RESEARCH

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ABSTRACT

Greece and Turkey are both members of NATO and are two of the principal players in the Balkan region. Their respective defence burdens, i.e. the share of military expenditure to GDP, are the highest in NATO. Their bilateral relations are marred by serious friction and conflict of interest and have on a number of occasions come close to an armed confrontation. Their strategic interaction and mutual weapons build-up has recently attracted the attention of researchers in the field testing the hypothesis of a Greek-Turkish arms race with conflicting results. This paper, using cointegration and causality tests, reports findings that provide empirical support to the notion of a systematic armaments competition between the two countries. More importantly though, this conclusion depends upon the identification of a regime shift estimated to occur in 1985.

I. INTRODUCTION

The Balkans have traditionally been an area of turmoil and instability. With the collapse of bipolarity and the Cold War security arrangements old as well as new local rivalries and disputes reappeared in the security agenda of the Balkan region. Greece and Turkey, two of the major players in the region, are considered by many to be the principal adversaries in the Balkans (Larrabee, 1992; Constas, 1991). Both have constantly ranked among the countries with the highest defence burden (military spending as a share of GDP) in NATO and in Europe. For example, in the period 1970-1994 Greece allocated an average of 5.8% of GDP to defence and Turkey 4.5% compared to a NATO average of 3.3% for the same period. In fact, in terms of the human and material resources that are yearly allocated to defence both countries are, in comparative terms, the most militarised countries in NATO (Kollias, 1995a; Chletsos and Kollias, 1995). The ongoing Greek-Turkish rivalry is well documented in the international relations literature (Wilson, 1979; Constas, 1991; Larrabee, 1992; Clogg, 1983, 1991) while empirical studies that have estimated demand functions for Greek and Turkish military spending have reported results indicating that the demand for defence expenditure in either country is strongly influenced by the allocations to defence by the other party (Kapopoulos and Lazaretou, 1993; Kollias, 1995a, 1996; Chletsos and Kollias, 1995).

The strategic interaction and the mutual weapons accumulation in the two countries has inevitably triggered the interest of applied researchers who have set out to address the question of whether the hypothesis of a Greek-Turkish arms race can find a modicum of empirical verification. Using in principal variants of the Richarson (1960) "reaction model" and applying Granger causality analysis on measures of the rival countries' military spending, the studies conducted so far report findings that can at best be deemed conflicting. The early work of Majeski and Jones (1981) and Majeski (1985) on the matter covering the 1949-1975 period, report significant reciprocal instantaneous interaction between the military expenditure series of the two countries pointing to the presence of an arms rivalry. Stavrinos (1992), extending the analysis through the use of cointegration techniques and a sample spanning from 1949 to 1988, finds evidence of unidirectional instantaneous causality running from the Greek to Turkish military expenditure but not the reverse. Finally, econometric results presented by Georgiou (1990) and Georgiou et al. (1996) for the 1958-1987 and 1960-1990 periods respectively do not appear to support the premise of a Greek Turkish arms race.[1]

The scope of this paper is to readdress the question of a systematic Greek-Turkish arms race in an effort to shed some light on the inconclusiveness surrounding the results documented in the empirical literature. The arms race hypothesis will be investigated in the context of a bivariate Vector Autoregression (VAR) in the military spending aggregates of the two countries through the application of Granger causality tests. In determining the direction of causality particular attention will be paid to the stationarity properties of the series involved

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[1]. But see Kollias (1994) for issues concerning the reliability of Georgiou's results.

and the presence of cointegration between them. In contrast to previous studies, however, specific allowance will be made for assessing the possible impact of structural breaks through the adoption of the recently proposed tests of Hansen (1992) and Gregory and Hansen (1996) which enable data based determination of policy regime shifts in a cointegrating relationship. The presence of significant policy regime shifts in the 1950-1995 period, for which our empirical tests are conducted, can not be ruled out from the outset since the period in question has seen important developments in the bilateral relations and strategic interaction of the two rival nations. It will be argued that failure to account for such regime changes may constitute one of the reasons lurking behind the mixed results reached in the literature.

The rest of the paper is organised as follows. The next section gives a brief historical background to the main issues of tension between the two countries. Section III lays out the specifics of the econometric techniques deployed and discusses the outcome of the empirical investigation on the hypothesis of a Greek-Turkish armaments competition. Finally, Section IV summarises and concludes the paper.

II. THE GREEK-TURKISH CONFLICT: AN OVERVIEW

The Greek-Turkish disputes are not entirely new but, as Larrabee (1992) points out, they take on a greater significance and a new dimension given the overall deterioration of the security situation in the region that has followed the profound politico-strategic changes of recent years. Following the post-war division of Europe both countries became members of the same alliance, despite a long history of conflicting interests and often hostile relations which have led to armed confrontations. In principle, as members of NATO they share common interests and defence objectives. Their bilateral relations however, are marred by a number of serious disputes. This uneasy coexistence has often threatened with collapse NATO's southern flank. The disintegration of the post-war security arrangements in the region has intensified their antagonisms which cover a wide range of issues. They include strong disagreements over the continental self of the Aegean Sea as well as control of the airspace over it and the status of certain Greek islands. However, the island of Cyprus has for more than three decades been the single most important source of tension and friction between the two countries the Turkish invasion of which in 1974 acted as the catalyst in their relations (Clogg, 1991). Constas (1991) notes that the use of military force by Turkey for the advancement of national objectives and the creation of a fait accompli in the island has left an indelible mark on the external threat perceptions of each country. Since then Greece and Turkey have at least four times been close to war but apart from sabre rattling neither thus far crossed the Rubicon. In 1986 a frontier incident between border patrols resulted in the death of one Greek and two Turkish soldiers. In 1987 a Turkish attempt at oil exploration in disputed areas of the Aegean resulted in the most serious crisis since a similar incident in 1976. The engagement of Greek and Turkish jets in "dogfights" over disputed areas in the Aegean is a weekly feature of their bilateral interaction. In 1985 Greece officially declared a new defence doctrine[1] which perceived no Soviet/Warsaw Pact threat, identifying Turkey as the main direct threat to its interests (Platias, 1991). In recent years, Turkey has explicitly threatened with war if Greece extended its territorial waters to 12 miles. In late 1994 when the new international treaty that allows countries to extend their territorial waters to 12 miles came into force both navies were put on full alert. Greece is a signatory to the treaty but Turkey objects to its provisions despite the fact that it claims a 12-mile Turkish territorial waters zone in the Black Sea. More recently, in early 1996, an armed confrontation was

[1]. In fact, although not explicitly declared, Greece since the 1974 Turkish invasion of Cyprus has regarded Turkey as the single most important source of external threat to its national interests while the threat from Warsaw Pact forces was considered as indirect and possible only in the context of a wider East-West armed confrontation. This view was (and is) universally held by strategic analysts and defence policy planners across the whole political spectrum in Athens.

apparently once again narrowly avoided when Turkey claimed that the status of certain uninhabited Greek islands was not clearly defined by the relevant treaties.[1]

Like other local disputes the Greek-Turkish conflict has thus far remained in the shadow of the all enveloping East-West confrontation. In broad terms, as Constas (1991) notes, the shared perception among western allies was that this conflict is a "manageable" one, susceptible to collective (NATO) or hegemonic (USA) involvement. However, the recent momentous changes have probably eroded alliance discipline and loyalty and NATO may no longer be able to contain the Greek-Turkish conflict. With dim prospects of a substantial improvement in Greek-Turkish relations Larrabee (1992) suggests that as long as their differences remain unresolved, there is always a chance that any incident could touch off a conflict which will cause wider strategic implications in the whole region. It is not surprising then to note that both countries have been yearly allocating substantial resources to their military establishments (see, Table 1) and as Sezer (1991) notes, an armaments competition appears to have been a permanent feature of their interaction over the past decades. Can, nonetheless, the armaments race argument be substantiated empirically? This is the question that we will now turn to address.

TABLE 1
Greek and Turkish Military Expenditure as a Percentage of GDP

Year	1960	1966	1970	1974	1976	1980	1986	1990	1993
Greece	4.9	3.7	4.8	5.6	6.9	5.7	6.1	5.8	5.4
Turkey	4.7	4.3	4.3	3.9	6.2	4.3	4.8	4.9	4.8
NATO	4.4	4.2	3.9	3.7	3.7	3.5	3.4	3.1	2.9

Source: SIPRI Yearbooks.

[1]. The present status quo between the two countries was established by the treaties of Lausanne-1923, Montreux-1936 and Paris-1947. An overview of the main provisions of the treaties can be found in the *Yearbook 1988, Hellenic Foundation of Defence and Foreign Policy* (pp. 85-92).

III. THE EMPIRICAL ANALYSIS

Following the established practice, the question of whether Greece and Turkey indulge in a systematic arms race is investigated by attempting to determine the direction of causality between Greek and Turkish military expenditures. Military spending is used as a proxy for physical arms build-up according to the notion that increases in the arsenal of the two rivals should be echoed by upward changes in their military spending.[1] The data on the military spending aggregates are of annual frequency covering the period 1950-1995, originate from the SIPRI yearbooks (various issues) and are expressed in 1985 US dollars. The Greek and Turkish military expenditure variables are logarithmicly transformed prior to estimation and are denoted by $lmgr_t$ and $lmtur_t$ respectively.

A. The Granger Causality Test

A commonly deployed technique in establishing the presence of an arms race between two nations or groups of nations is the Granger causality test. In principle, the Granger causality test is formulated on the basis of a VAR in the variables of interest, namely:

$$lmgr_t = \pi_0 + \sum_{i=1}^{k} \pi_i lmgr_{t-i} + \sum_{i=1}^{k} \delta_i lmtur_{t-i} + \psi_{1t} \tag{1}$$

$$lmtur_t = \theta_0 + \sum_{i=1}^{k} \theta_i lmgr_{t-i} + \sum_{i=1}^{k} \zeta_i lmtur_{t-i} + \psi_{2t} \tag{2}$$

where k is the maximum lag length, provided that $lmgr_t$, and $lmtur_t$ are covariance stationary time series. If J_t is a universe of information up to and including period t then the Granger (1969) definition of causality states that $lmtur_t$ causes $lmgr_t$, given J_t if $lmgr_t$ has a smaller forecast error variance when the information contained in past values of $lmtur_t$ (i.e. $lmtur_s$, $s < t$) is used than if it is not used at all. Thus, if past $lmtur_t$ contributes significantly to forecasting current $lmgr_t$, that is the coefficients δ_i, $i=1, 2,...,k$, in (1) are jointly significant in terms of a standard Wald test, then $lmtur_t$ is said to Granger cause $lmgr_t$. Similarly the null hypothesis that $lmgr_t$ does not Granger cause $lmtur_t$ is rejected if the parameters θ_i, $i=1, 2,...,k$, in (2) are jointly significant. In the present context, the arms race hypothesis will hold true if each nation's current military expenditure is determined by the nation's own as well as its opponent's past military expenditure behaviour, that is if the military expenditure aggregates Granger cause each other.

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[1]. A number of important points have been raised concerning the usefulness of military expenditures as a defence measure [see for example Anderton (1989) pp. 352]. However, the discussion of such issues is well beyond the scope of this paper.

In the event of the variables involved being non-stationary, the form of non-stationarity should be clarified prior to the implementation of the Granger causality test so that the danger of drawing misleading inferences can be averted. Specifically, if the series at issue are integrated of the same order, say I(1), and share a common trend, that is if they are cointegrated in the sense of Engle and Granger (1987), then Granger (1988) has shown that there should be Granger causality in at least one direction. In this case determining the causal ordering between $lmgr_t$ and $lmtur_t$ in a Granger sense requires the reformulation of the VAR in the familiar error correction specification (ECM), that is,

$$\Delta lmgr_t = \pi_0^* + \sum_{i=1}^{k} \pi_i^* \Delta lmgr_{t-i} + \sum_{i=1}^{k} \delta_i^* \Delta lmtur_{t-i} + \rho_1 \hat{z}_{t-1} + \omega_{1t} \tag{3}$$

$$\Delta lmtur_t = \theta_0^{'} + \sum_{i=1}^{k} \theta_i^* \Delta lmgr_{t-i} + \sum_{i=1}^{k} \zeta_i^* \Delta lmtur_{t-i} + \rho_2 \hat{z}_{t-1} + \omega_{2t} \tag{4}$$

where Δ is the first difference operator and $\hat{z}_t$ are the estimated residuals from the cointegration regression between $lmgr_t$ and $lmtur_t$, i.e:

$$lmtur_t = a + blmgr_t + z_t \tag{5}$$

The error correction term in (3)-(4) provides an additional channel of causation between $lmgr_t$ and $lmtur_t$ and consequently its statistical significance should be assessed in addition to evaluating the joint significance of the δ_i^*'s in (3) and that of the θ_i^*'s in (4), for a verdict on the direction of causality and the existence of an arms competition to be reached (see, Granger, 1988). If, on the other hand, the series do not happen to form a cointegrating vector then the application of the Granger causality test commands that $lmgr_t$ and $lmtur_t$ be rendered covariance stationary before entering the VAR. With the non-stationarities removed the computation of the Granger causality test proceeds in the usual manner.

B. Testing for Unit Roots

In the light of the preceding discussion our primary concern lies with determining the presence of non-stationary behaviour, either deterministic or stochastic, in the autoregressive representation of the time-series under investigation. Consequently, $lmgr_t$ and $lmtur_t$ are initially subjected to unit root testing by means of three statistical procedures, namely the recently proposed Phillips and Ploberger (1994) posterior information criterion (PIC), the standard Augmented Dickey-Fuller (ADF) unit-root test (see, Said and Dickey, 1984), and the *Z(t)* unit root test of Phillips and Perron (1988).

The PIC procedure determines the lag order (k) and trend degree (p) in an autoregressive model of the form,

$$\Delta X_t = \rho X_{t-1} + \sum_{j=0}^{p} \beta_j t^j + \sum_{i=1}^{k} c_i \Delta X_{t-i} + \varepsilon_t \tag{6}$$

where X_t is the variable of interest and ε_t is a sequence of normal independent random variables with mean 0 and variance σ^2, and returns the odds in favour of a unit root. Specifically if the PIC value exceeds unity then the unit root model, i.e. $\rho = 0$, should be chosen. The main advantage of the PIC criterion over the widely used ADF test is the ability of the former to avoid overspecification or underspecification of Eq. 6 under the null hypothesis by utilising data information for determining the lag order and the trend degree parameters.[1] Nevertheless as a means of verifying the outcome of the PIC procedure, both a standard ADF unit root test, denoted by $t_{adf}(k,p)$ henceforth, as well as a Phillips-Perron $Z(t)$ test statistic are computed. The former is obtained by employing the values of k and p returned by the PIC procedure. The latter is estimated successively for three Dickey-Fuller type auxiliary regressions that involve no-intercept, an intercept, and an intercept and trend as deterministic regressors respectively. The covariance parameters required for the computation of the relevant $Z(t)$ statistics are estimated by means of a VAR(1) prewhitened quadratic spectral kernel estimator with the Andrews (1991) automatic plug-in bandwidth (see, Andrews and Monahan, 1992). The resultant statistics are denoted by $Z(t)$, $Z_\mu(t)$, and $Z_\tau(t)$ respectively. Finally, in accordance to the suggestions of Dickey and Pantula (1987), the first differences of the variables are also subjected to the same battery of tests so that the presence of higher order integrated processes can be examined too.

The outcome of the PIC, ADF and Z(t) tests for the variables under consideration is presented in Table 2. For the levels of both military expenditure series the results favour a unit root because the PIC assumes values which clearly exceed unity. Specifically the PIC equals 19.735 for $lmgr_t$ and 1.395 for $lmtur_t$ with $k = 1$ and $p = -1$, where a minus one value for p signifies the absence of a constant term in Eq. 6. The presence of at least one unit root suggested by the PIC is given further grounds of support by the $t_{adf}(1,-1)$ and the Phillips-Perron $Z(t)$, $Z_\mu(t)$, and $Z_\tau(t)$ unit root statistics. None of these statistics turns out to be statistically significant in terms of the corresponding p-values computed from the response surface regressions of MacKinnon (1994) thereby failing to reject the unit root null hypothesis (see, Table 2). Moreover the same set of tests when applied to the first difference of $lmgr_t$ and $lmtur_t$ discloses no evidence of a second unit root. The PIC is practically nought for the first difference of both series with $k = 0$ and $p = -1$ for $\Delta lmgr_t$ and $k = 1$ and $p = -1$ for $lmtur_t$, while the ADF and Z(t) statistics emerge highly negative in size leading to a sound rejection

[1]. Since a detailed description of the computational details of the PIC procedure goes beyond the scope of this paper, the interested reader is referred to Phillips and Ploberger (1994).

of the unit root null even at the 99% confidence level. Hence one can confidently infer that the military spending aggregates of both countries behave as I(1) stochastic processes. Having established that the variables in question are integrated processes of the same order, one may now proceed to examine whether they also move together in the long-run, that is whether they share a common trend.

TABLE 2
The Unit Root Tests

	Phillips-Ploberger Criterion				Phillips-Perron Z(t) test		
Variable	*PIC*	*k*	*p*	$t_{adf}(k,p)$	*Z(t)*	$Z_{\mu}(t)$	$Z_{\tau}(t)$
$lmgr_t$	19.735	1	-1	2.1527	2.537	-1.266	-1.025
$\Delta lmgr_t$	0.0000	0	-1	-6.036**	-6.037**	-6.807**	-6.883**
$lmtur_t$	1.395	1	-1	2.1529	2.466	-0.715	-3.401
$\Delta lmtur_t$	0.0007	1	-1	-4.078**	-4.406**	-5.179**	-5.178**

Notes:

◊ * and ** indicate significance at the 95% and 99% confidence levels respectively.

◊ PIC is the Phillips and Ploberger (1994) posterior information criterion odds in favour of a unit root. The unit root hypothesis is favoured if the PIC value exceeds unity.

◊ k is the selected lag order and p is the selected trend degree in the associated autoregressive model estimated by the Phillips and Ploberger (1994) method. The values of k and p are data determined. Note that a minus one value for p signifies the absense of an intercept in the underlying AR specification.

◊ $t_{adf}(k,p)$ is the standard Said and Dickey (1984) ADF test statistic corresponding to the PIC selected k and p values.

◊ Z(t), $Z_{\mu}(t)$, and $Z_{\tau}(t)$ are the Phillips and Perron (1988) unit root test statistics when the auxiliary regression contains no deterministic components, an intercept, and an intercept and a trend respectively. The computation of these statistics is based on the application of a quadratic spectral kernel with VAR(1) prewhittenning (see, Andrews and Monahan, 1992). The truncation lag parameter is determined by applying the Andrews (1991) automatic plug-in bandwidth estimator.

◊ The statistical significance for the ADF and Z(t) test statistics is determined through the use of the response surface regressions appearing in MacKinnon (1994).

C. Testing for Cointegration

Following Engle and Granger (1987), the presence of cointegration between $lmgr_t$ and $lmtur_t$ is investigated by subjecting the residuals $\hat{z}_t$ from the cointegration regression (5) to unit root testing. In contrast to the usual practice, the cointegration regression has not been estimated by OLS. Instead the Fully Modified (FM) OLS method of Phillips and Hansen (1990) is used so that the inefficiency of the standard OLS estimates and the inapplicability of Gaussian inference on the estimated cointegrating vector can be alleviated (see, Ogaki, 1993).[1]

The ADF testing principle is utilised as the means of evaluating the null hypothesis of no-cointegration. Naturally, the ADF auxiliary regression involving the FM residuals from (5) does not simultaneously incorporate any intercept as the latter has already been accounted for in the cointegration regression itself. The resultant ADF t statistic is denoted by *CADF* while the decision on the number of lagged differenced terms in the ADF auxiliary regression has been made on the basis of evidence provided by a 10% level sequential t-ratio test on the significance of the highest order lagged differenced term (see, Ng and Perron, 1995). The FM estimated cointegration regression and the associated *CADF* statistic are reported in Table 3. The value taken by *CADF* is -1.633 and carries a p-value of 0.707 which clearly suggests that the null hypothesis of no-cointegration cannot be rejected.[2] The lack of a cointegrating relationship between the military expenditure aggregates of Greece and Turkey detected by the *CADF* test statistic points to the initial conclusion that the Granger causality analysis should be performed using a VAR in the first differences of these variables. However, the failure of the *CADF* test to reject the unit root null may in fact be due to the presence of a structural break in the data biasing the testing procedure toward favouring the null hypothesis of no cointegration (see, Perron, 1989). This presumption gains considerable momentum if one looks at the time plots of $lmgr_t$ and $lmtur_t$ in Figure 1.

It can be easily seen that although these variables are tracing each other fairly closely over the sample period, they begin to diverge from each other around the mid eighties providing , in effect, a visual justification for the detected lack of cointegration between them.

[1]. The most commonly used technique in estimating cointegrating relationships is that of Johansen (1991). However, its implementation in small samples, as the one used in this study, is inadvisable since the Johansen estimator is subject to significant small sample bias (see, Phillips, 1994). Furthermore, the utilisation of a single equation estimation method instead of the Johansen system estimation technique seems more appropriate in our case because with two variables there can only be one cointegrating vector. Finally the FM estimation technique being semiparametric in nature requires consistent estimation of covariance parameters. The latter have been computed by employing the quadratic spectral kernel estimator proposed by Andrews and Monahan (1992) as in the case of the Phillips-Perron Z(t) statistics.

[2]. The significance of the *CADF* test is derived from the response surface regressions found in Tables 3 and 4 in MacKinnon (1994).

TABLE 3

The Cointegration Tests

General cointegration regression: $lmtur_t = \alpha + \beta\varphi(\lambda)_t + \gamma lmgr_t + \varepsilon_t$

Regression applied	α	β	γ	$CADF$	L_C	$MeanF$	$SupF$	$CADF_{inf}$
No dummy $\beta = 0$	-0.018 (1.856)	-	1.039 (0.267)	-1.633 [1]	0.043	7.502^{**}	27.2^{**}	-
With dummy $\beta \neq 0$	1.522 (0.493)	0.612 (0.131)	0.793 (0.073)	-	-	-	-	-4.648^{*} [2] $\lambda=0.78$

Notes:

◊ * and ** indicate significance at the 95% and 99% confidence levels respectively.

◊ The cointegration regressions are estimated by the Fully Modified OLS technique proposed by Phillips and Hansen (1990). The covariance parameters required for the bias correction are estimated through the Andrews and Monahan (1992) VAR(1) prewhitened quadratic spectral kernel estimator with automatic plug-in bandwidth.

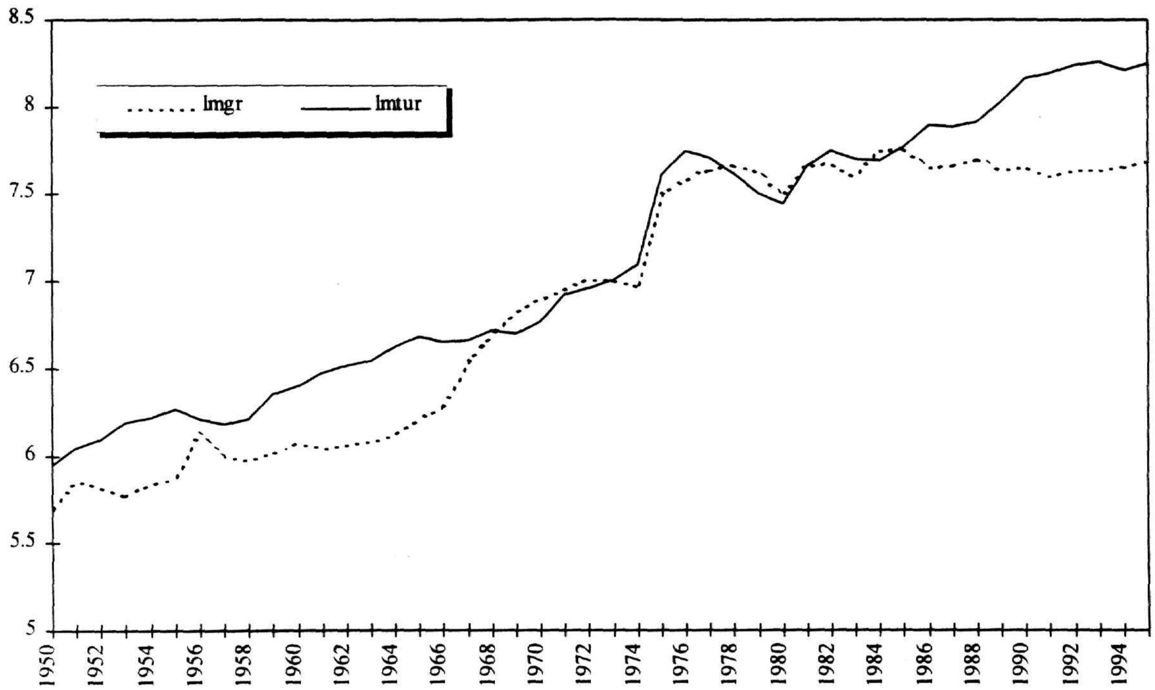
◊ CADF is the standard t ratio unit root statistic on the residuals from the no dummy cointegration regression. The statistical significance of the CADF statistic is ascertained by the relevant response surface regressions found in MacKinnon (1994). The figures in parentheses are standard errors while those in square brackets correspond to the number of lagged differenced terms in the ADF auxiliary regressions required for serial correlation correction.

◊ SupF, MeanF, and L_c are the Hansen (1992) parameter stability tests of the no dummy cointegration relationship. Asymptotic critical values for these test statistics appear in Tables 1, 2, and 3 in Hansen (1992) respectively.

◊ $CADF_{inf}$ is the smallest t-ratio no cointegration ADF statistic obtained from the sequential estimation of the cointegration regression that involves the rolling regime shift dummy $\varphi(\lambda)_t$. 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

Greek and Turkish Military Expenditure 1950-1995 (in logs)



In fact the level of Greek defence spending appears to be falling short of its Turkish counterpart from 1985 onwards. A possible explanation should rest with the significant deterioration Greek public finances experienced over the past decade[1] which seems to have seriously affected the ability of the country to arm itself at the same rate as in the past. On the other hand, in the case of Turkey, the sustained upward trend in military spending may be partly due to the intensification of its internal security problems regarding the Kurdish movement for autonomy. For the past decade or so the Turkish security forces have been fighting a costly and bloody war with Kurdish rebels in the southeastern provinces of the country (Gunluk-Senesen, 1995). Nonetheless, formal evidence on the impact of such a change on the stationarity status of the residuals from (5) as well as on the precise timing of the potentially important structural break demands application of suitably designed testing procedures.

In order to look into the possibility of a structural break impinging upon the unbiasedness of the *CADF* procedure, the sequential no cointegration test of Gregory and Hansen (1996) has been implemented. This test shares the same null hypothesis with the

[1]. For a comprehensive discussion on Greek fiscal problems and economic performance see Alogoskoufis (1995).

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 incidence 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 seems to suggest, the Gregory-Hansen procedure involves estimating the following cointegration regression,

$$lmtur_t = \alpha + \beta\varphi(\lambda)_t + \gamma lmgr_t + \varepsilon_t \tag{7}$$

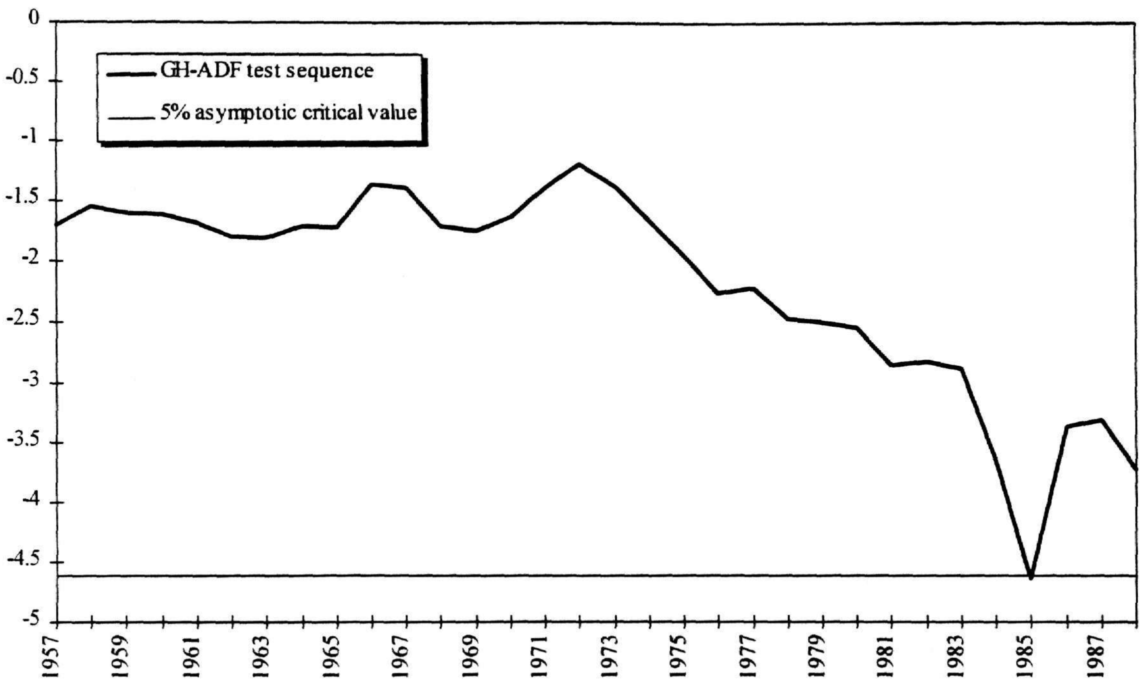
where, $t=1,2,...,T$, $\lambda = T_B / T$, T_B is the break date, and $\varphi(\lambda)_t = 1$ if $t > T\lambda$ or 0 otherwise. The estimation of the cointegration regression is performed sequentially with the breakpoint $\lambda = T_B / T$ assuming values in $\Im$=[0.15, 0.85] and k representing the number of lags required for serial correlation correction in the associated ADF auxiliary regression on the residuals from (7).[1] Each estimation run produces a value for the ADF t-ratio test. The lowest ADF t statistic obtained from this sequential procedure, denoted by $CADF_{inf}$ henceforth, is compared to the asymptotic critical value provided in Table 1A of Gregory and Hansen (1996). If $CADF_{inf}$ is found to be statistically significant, $lmgr_t$ and $lmtur_t$ are said to be forming a cointegrating relationship whose position has changed to a new "long-run" equilibrium at time λ^*.[2] To put it differently, $\hat{z}_t$ 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 (7) is depicted in Figure 2, and the smallest ADF value encountered along with the associated estimated cointegration regression are reported in Table 3. Examination of Figure 2 reveals that the no cointegration null is soundly rejected. The ADF t-statistic takes its lowest value in 1985, that is $\lambda^* = 0.78$, and lies just below the 5% asymptotic critical value line. Specifically, $CADF_{inf}$, that is the value of the ADF statistic that assigns the greatest weight to the policy regime shift alternative, amounts to -4.648 suggesting that cointegration is achieved only when the 1985 policy regime change is explicitly incorporated into the analysis. More importantly though the timing of the structural break as determined by the Gregory-Hansen method complies with the observation made earlier, through the pre-test diagrammatic examination of the data, that the series under investigation commence to diverge from each other on this particular year.

1. The trimming of the values assumed by λ is deemed necessary due to the test suffering power loss when structural breaks occur outside the $\Im$=[0.15, 0.85] region (see, Andrews, 1993). Moreover, note that the number of lag differenced terms (k) required for the computation of the $CADF_{inf}$ cointegration statistic is determined in accordance to the same criterion used in the case of the standard $CADF$ test.

2. λ^* stands for the estimated sample break fraction corresponding to $CADF_{inf}$

FIGURE 2
Gregory-Hansen No-cointegration ADF t-Test Sequence



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. This is so 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 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 geared toward detecting any gradually escalating changes in the cointegration regression coefficients. In principal the Hansen (1992) procedure produces a sequence of fixed breakpoint Chow F-tests with the breakpoint taking values over the $\Im$=[0.15, 0.85] range. The *SupF* is the highest value in the F-test sequence while *MeanF* stands for the mean value of the sequence.[1]

The findings from the application of the Hansen tests on the no dummy cointegration regression appear also in Table 3, and the F-test sequence produced is plotted against the breakpoint range in Figure 3. 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, the L_c test provides no evidence of parameter instability while the *MeanF* and *SupF* are found to reject the stability null even at the 1% significance level. On the basis of this evidence is laborious to distinguish whether there has been a gradual as opposed to a swift change in the cointegrating vector, although the latter scenario seems the most probable one in view of the sheer size of the *SupF* statistic; it amounts to 27.2. The sudden shift scenario gains in credibility once the plot of the F-test sequence is examined. A brief inspection of Figure 3 reveals that the F-test sequence attains three distinct peak values in 1975, 1981 and 1985 with the first two marginally exceeding the 5% SupF critical value line and the last, corresponding to the *SupF* statistic, being far well above it. Taking into consideration that the Gregory and Hansen (1996) $CADF_{inf}$ statistic also assumed its lowest value in 1985, one is led to infer that there has been a seminal policy regime change in 1985 and unless allowance is made for its presence the military expenditures of the two countries cannot form a cointegrating vector. Since cointegration around the 1985 regime policy shift between $lmgr_t$ and $lmtur_t$ has been established, it is now possible to proceed with the formulation of the relevant error correction specification on the basis of which the hypothesis of an arms race between Greece and Turkey can be investigated.

D. Testing for the Presence of an Arms Race

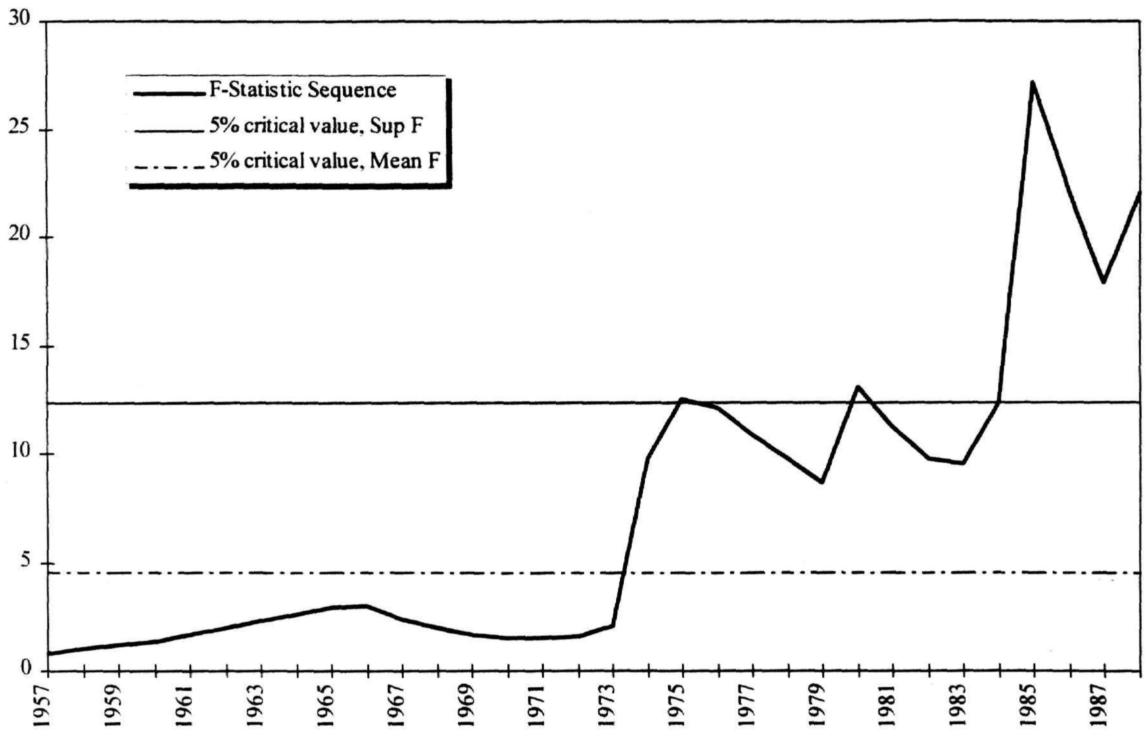
In accordance to the methodological guidelines set out earlier, the ECM specification (3)-(4) is estimated by OLS applied to each equation. The error correction term is formed from the residuals of Eq. (7) which involves the regime shift dummy $\varphi(\lambda^*)_t$ for 1985. The optimum lag length *k* for specifying the ECM has been determined by the joint use of a 10% level sequential F-type test on the joint significance of the highest order lagged terms in the ECM coupled with extensive diagnostic testing of the residuals. Given the annual frequency of the data set, an initial lag length of 2 was selected. However, irrespective of the initial maximum lag length chosen, the model failed consistently the Jarque-Berra normality test pointing to the presence of outliers in the residuals. Subsequent estimation of the model by recursive least squares and computation of a sequence of Chow tests unveiled the need to incorporate two impulse dummies for 1975 and 1981 in order to achieve a congruent specification.[2] The dummy variables can be thought of as accounting for the lagged response

[1]. A comprehensive description of the computational and distributional details of the *SupF*, *MeanF*, and L_c statistics is provided in Hansen (1992).

[2]. These results are not reported for reasons of brevity. Nonetheless they are available upon request from the authors.

FIGURE 3

Hansen Cointegration Regression Parameter Stability F-Test Sequence



of the expenditure variables to sample specific events, namely the 1974 invasion of Cyprus by Turkish troops, the 1980 military coup in Turkey and the impact that the ascension to power in 1981 of the socialist party PASOK had on Greek defence and foreign policy since on the whole, the socialist administrations have adopted a more aggressive stance and rhetoric towards Turkey. Having established congruency, a lag length of one has been found to be the optimal. The resultant ECM model is then used as the platform for assessing the null hypothesis of no Granger causality, that is absence of an arms race between the two rival countries. This hypothesis is evaluated by two standard Wald χ^2 tests performed on each equation of the system. These tests assess the statistical significance of the error correction term and the opponents' lagged differenced military spending regressors in each equation respectively. The results of the OLS estimation along with the computed χ^2 tests are summarised in Table 4.

TABLE 4

The OLS Estimated ECM Specification and the Granger Tests for Temporal Causality

Regressant	Estimated coefficients				Granger causality χ^2 Wald tests	
	constant	π_1^*	δ_1^*	ρ_1	δ_1^*=0	ρ_1=0
$\Delta lmgr_t$	0.021 (0.015)	0.088 (0.133)	0.025 (0.152)	0.145 (0.086)	0.026 [0.87]	2.824 [0.09]
	Diagnostics: R^2=0.517, S.E.=0.084, AR: F(2,36)=0.387, H: F(6,31)=0.669, NORM: χ^2(2)=5.207, ARCH: F(1,36)=0.898					
	constant	θ_1^*	ζ_1^*	ρ_2	θ_1^*=0	ρ_2=0
$\Delta lmtur_t$	0.02 (0.01)	-0.034 (0.092)	0.318 (0.105)	-0.129 (0.059)	0.135 [0.713]	4.747 [0.029]
	Diagnostics: R^2=0.69, S.E.=0.058, AR: F(2,36)=0.824, H: F(6,31)=0.309, NORM: χ^2(2)=0.939, ARCH: F(1,36)=0.012					

Notes:

◊ *S.E. is the standard error of the equation, AR stands for the F version of the LM test for serial correlation, ARCH tests for conditional heteroskedasticity, H tests for unconditional heteroskedasticity due to the squares of the regressors, and NORM is the Jarque-Berra test for normality.*

◊ *The figures in parentheses are standard errors while those in square brackets are p-values levels.*

◊ *The model involves two impulse dummy variables assuming the value of 1 for 1975 and 1981 respectively. These dummies account for the lagged response of the military expenditure variables to the 1974 invasion of Cyprus and the 1980 military coup in Turkey.*

Recall that for the hypothesis of an arms race to be substantiated, the Granger causality tests should detect bi-directional causality between $lmgr_t$ and $lmtur_t$. The empirical findings reached tend to support this notion for the 1950-1995 period. Initially note that both the $\Delta lmtur_{t-1}$ term in equation (3) as well as the $\Delta lmgr_{t-1}$ term in equation (4) are found to be statistically insignificant by the computed Wald tests. From Table 4 it can be easily verified that the χ^2 tests assessing the null hypotheses $H_0^1\!:\delta_1^* = 0$ and $H_0^2\!:\theta_1^* = 0$ have a *p*-value of 0.87 and 0.713 respectively indicating that the change in defence spending of each country does not respond to the lagged change in the spending of its adversary. Nevertheless this implied lack of causal feedback between the variables of interest is reversed when the significance of the error correction terms in each equation, that is $H_0^3: \rho_1 = 0$ and $H_0^4: \rho_2 = 0$, is examined. The p-values of the associated Wald tests turn out to be 0.09 and 0.029 respectively (see, Table 4). These results are indicative of the presence of bi-directional instantaneous causality between the Greek and Turkish military expenditure variates or equivalently, of the existence of an armaments competition between the two countries at the 10% significance level. Collectively taken these findings provide a novel insight into the Greek-Turkish arms race debate. First a reciprocal and instantaneous response of each country to the opponents' current defence spending has emerged from the empirical analysis as one would have anticipated given the volatility and tension characterising the bilateral relations of the two countries. More importantly though our analysis has identified that a change in this antagonistic pattern of behaviour has been taking place since the mid 80s. In actual fact failure to account for the 1985 defence policy regime change would have led to the derivation of erroneous inferences with regard to the arms race competition question. This is so because the lack of cointegration between the defence spending variates in the absence of the policy shift dummy would have caused us to mispecify the VAR used for the computation of the Granger causality tests. The latter tests would then have missed to bring forth the pre-1985 armaments competition between Greece and Turkey.[1]

[1]. We have also estimated a VAR only in the first differences of the defence spending series. This exercise produces Granger causality tests which find no evidence of temporal causality between the series in question supporting the argument made in the text. Again in order to economise in space these results are not reported, yet can be obtained upon request.

IV. CONCLUSIONS

The end of the Cold War and of the East-West arms race has raised the prospects of potential gains from reductions in military expenditures. However, tension, conflict and interstate rivalry are still present in many regions. Greece and Turkey, two of the major players in the post-bipolar strategic environment of the Balkans, have a long and well documented history of tense bilateral relations that have often brought the two countries close to armed conflict. This study, using cointegration and causality tests, examined whether the hypothesis of a Greek-Turkish armaments race can be empirically corroborated. In contrast to recent empirical studies (Georgiou, 1990; Stavrinos, 1992; Georgiou et al., 1996) the obtained results cannot refute the hypothesis of an arms race between the two members of NATO at the 90% confidence level. The empirical verification of the Greek-Turkish arms race hypothesis, however, is found to depend upon the identification of a policy regime change estimated to occur in 1985. We believe that this change in defence policy is related to the straining fiscal finances confronting the Greek economy over the past decade which in turn translate to the realised inability of Greek defence spending to trail its Turkish counterpart and its stagnation around its 1985 level. Had this structural change remained unaccounted for, the existence of an arms rivalry between the two countries in the pre-1985 era would have gone undetected.

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