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Luca PIERONI — Giorgio d'AGOSTINO
Marco LORUSSO

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CAN WE DECLARE MILITARY KEYNESIANISM DEAD?

Luca Pieroni[a] Giorgio d'Agostino[b] Marco Lorusso[c]

[a, b] *Department of Economics, Finance and Statistics, University of Perugia,*

Via Pascoli 20, 06123 Perugia (Italy)

[c] *Department of Economics, University of Verona,*

Viale dell'Università 4, 37129 Verona (Italy)

****Corrisponding author:*** *Luca Pieroni; lpieroni@unipg.it*

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Abstract

This paper empirically tests the Keynesian hypothesis that government defence spending positively impacts on aggregate output, by using a long run equilibrium model for the US and the UK.

Our contribution, with respect to previous works, is twofold. First, our inferences are adjusted for structural breaks exhibited by the data concerning fiscal and monetary variables. Second, we take into account different dynamics between defence spending on aggregate output, showing that the results are sensitive to sub-sample choices. Though the estimated elasticities in both countries show a lack of significance in the more recent years of the sample, defence spending priorities addressed to international security may revitalize pro-cyclical effects in the UK, by an industrial policy of defence shared with the EU members.

KEY WORDS: Military spending; Output; Long run models

J.E.L.: E12, E52, E62, H56

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1. Introduction

One of the dominant approaches to macroeconomic research in the past several decades based policy predictions on the IS-LM model. In this framework, the debates between Keynesians and monetarists concerning the effectiveness of monetary and fiscal policy played a central role in the analysis of short-run fluctuations (Romer, 2000). One assumption largely criticized of this aggregate macroeconomic model involved in the monetary policy behaviour of the central bank concentrated on the aims targeting money supply. On the other hand, empirical policy researches have shown that the central banks mainly use the tool of the interest rate to determine the monetary policy (MP) with respect to characterize the money supply (Taylor, 2000).

Although such a framework is useful for understanding how the monetary policy affects the economy, through a closed relationship between inflation-real interest rates, it is ill-equipped to investigate how the fiscal policy shocks impact on aggregate output through the *composition* of government spending. Focus on the latter was motivated by our interest in understanding how society might best avoid the distortions created by the presence of misallocation of government spending. In this paper, we provide some empirical evidence of the effects of the composition of fiscal policy on the aggregate output when the categories of defence and civilian spending are explicitly distinguished within the government sector. Firstly, we assess the model by identifying fiscal policy shocks as motivating forces for the nonstationarity of output. Indeed, equilibrium of the IS-MP framework implies that, if the shock of government spending components, namely defence and civilian spending, are unobservable shocks I(1), these forcing variables will determine a long run equilibrium along with output and real interest rate. Secondly, we are interested in documenting and discussing the effects of a particular kind of government spending – the defence spending - on the long run output since the empirical evidence does not provide a clear picture if defence spending stimulates, through higher demand and innovations, the economy or retards economic performance by crowding out effects (Gold, 2005). Thus, this paper reviews the debate in line with the new-keynesian approach developed by Atesoglu (2002), by updating the sample of data in the US and by comparing the estimated results with that obtained for the UK economy.

Theoretically, the well-known hypothesis of the Keynesian approach, that treats the military budget as a source of aggregate demand for goods and services, suggests that positive government spending should induce economic stimulation by means of an income multiplier effect. In the extreme case, this government economic policy is known as *military Keynesianism*, when the fiscal policy devotes large amounts of spending to finance the defence sector (Mintz and Hicks, 1984). The channel through which military spending can affect the economy is based on boosting utilisation of capital stock and higher employment. Positive changes in capital stock utilisation may lead to increase profit rate which, in turn, drives to higher investment in short run (Smith and

Dunne, 2001; Dunne et al., 2004; Kollias et al., 2004). However, the “utilisation effect of capital stock” may have much less pronounced output effects in a longer span (Dakurah et al., 2001; Dritsakis, 2004). The defence economics literature identified with the opportunity cost of defence spending the effects of investment crowding out, which turns out to be a drag on economic take-off (Sandler and Hartley, 1995).

Focusing our attention on empirical analysis, it is known that the robustness of the aforementioned test of a long run model effect of defence spending on aggregate output could be better obtained by working with quarterly frequency data. While for the US, National Income and Product Accounts (NIPA) produces quarterly data for the categories of government defence and civilian spending, we have an issue with the classification of government spending in the UK, because quarterly observations are only available for the series of total public consumption, while disaggregated information on the composition of public spending is available on an annual basis. Thus, we provide to reconstruct quarterly data on the defence and non-defence public spending of the UK by disaggregating the relative annual series in line with the Chow and Lin’s procedure (1971).

In summary, this paper theoretically justifies and empirically tests two hypotheses: (i) the effects of defence spending on output depend on the long run equilibrium model that also includes the variables of monetary policy and government civilian spending; (ii) defence spending, as a component of public spending, positively and significantly impact on the long run output. In both countries, the empirical identification of a cointegrating vector shows a coherence of data with the predictions of the Keynesian model. By assessing the estimated parameters of the models, we find that the relationship between defence spending and output is strongly sample-dependent with a fall in the elasticity values in more recent years, though a Keynesian stimulus for the UK economy may arise by improving the efficiency of the defence industry within the auspicated EU defence policy.

The structure of this paper is as follows. We discuss conceptual issues in Section 2. Section 3 provides an overview of econometric specifications. Section 4 presents the data, shows tests for the identification of the model and discusses the estimation results to shed some light on the Keynesian effects of defence spending on output. Concluding remarks are offered in Section 5.

2. Theory: a simple macroeconomic model

In this section, an IS-MP model, that identifies the policy fiscal shocks by using the defence and civilian spending components of the public budget sector, will be formulated. To organize the discussion, a stripped-down baseline model is expounded as a version of the one described by Atesoglu (2002), to characterize a number of broad principles that underlie optimal policy management. We then consider fiscal policy implications by adding various real world complications to test how the prediction from theory is linked with policy-making in practice. Specifically, it will serve as a basis for the empirical work to assess the impact of the government defence spending on economic stimulation.

Because we are interested in characterizing fiscal policy rules in terms of composition of the government budget, the model we use evolves as in Romer (2000) and Taylor (2000), and is derived by assuming that the real interest rate is predetermined by the central bank[1]. The main change in the monetary policy rule is that it replaces the assumption to target the money supply with a simple interest rate rule, as supported by the central bank's behaviour in the developed countries (Taylor, 1993).

On the other hand, the importance of this assumption may depend on its applications. For example, it might be reasonable to ignore that the real interest rate may depend on aggregate output, when applied to the effects of government spending in the civilian and defence categories, if the aim is to examine their effects on aggregate output rather than to assess the new Keynesian model[2].

Below, we formally document the theoretical framework and discuss the assumption of the model. Let R_{jt} denote the measure of type-*j* interest rate chosen as a target indicator by the Central

[1] In the complete version of the new macroeconomic model, the real interest rate is explained by additional equations.

[2] It is worth noting that the straightforward assumption that the central bank is able to follow a real interest rate rule makes the model Keynesian.

Bank in period t to drive the monetary policy[3]. Then, the aggregate output, the amount of the final goods and services produced in the economy, is denoted as Y_t. Since the aggregate income is $Y_t = W(r_{jt})$, the mathematical formulation of the IS equilibrium equation requires that:

$$Y_t = -\zeta \ R_t + \mu_t \tag{1}$$

where μ_t is a stochastic term that includes shocks of fiscal policy and/or net export. The right-hand side of the IS equation describes the known inverse relationship between the (real) interest rate targeted by the central bank's choices and aggregate output. The stochastic term of Equation (1) plays a central role in the following analysis since we will concentrate our estimations on the effects of government defence spending. It is worth keeping in mind the intuitive meaning behind it. If this specific component of fiscal policy increase(s), the shock on the IS curve generates a positive shift on output and a new equilibrium in the output-real interest rate space is produced. Let M and G denote defence and civilian components of total government spending, respectively, we will identify these shocks as "fiscal policy shocks".

Let us now turn to the real interest rate. This variable is assumed to be only dependent on inflation such that the behaviour rule generates a monetary policy (MP). For the sake of simplicity, we assume that: $r_t = \bar{\pi}$, where $\bar{\pi}$ is the inflation rate assumed to be predetermined (known) by the central bank. A number of implications emerge from this baseline case on which monetary policy is firmly based. Focused on evaluating fiscal policy shocks, the main is that the central bank adjusts the nominal short rate one-for-one with perfect foresight of (expected future) inflation. That is, it should instantaneously adjust the nominal interest rate such that it does not alter the real interest rate (and aggregate demand).

To sum up, since the central bank's choice of the real interest rate is strictly predetermined by inflation rate, the real interest rate rule can be approximated by a horizontal line in the output-real

[3] See Atesoglu (2007) for a discussion on the choice rule of the interest rate target.

interest rate space. The IS curve, that includes the government components of expenditure, in turn, can be used to assess the impact on aggregate output.

Rather than work through the details of the derivation, which are available in Appendix 1, we directly introduce the key aggregate relationships by the reduced form of the model. For convenience, the theoretical framework abstracts from the way by which public expenditure was financed. This abstraction does not affect any qualitative conclusions as we will discuss. The model specification is formulated as follow:

$$Y_t = \beta_0^* + \beta_2^* M_t + \beta_3^* G_t + \beta_4^* R_t + \psi_t \qquad (2)$$

where ψ_t term of Equation (2) contains net export shocks as shown in Appendix 1. $\beta^* = (\beta_0^*, \beta_2^*, \beta_3^*, \beta_4^*)$ represents the vector of parameters to be estimated. Though the model is quite simple, it nonetheless contains the main ingredients of the descriptively richer frameworks that are used for policy analysis. Within the model, as in practice, the instruments of fiscal policy based on the composition of government spending, account for the short term fluctuations. We would like to remark that the presence of non-stationary(trending) variables of government expenditure may affect the long run relationship. In the next Section, a dynamic reduced form model will enable to test the presence of long run effects of the Keynesian stimulus on the economy.

3. Econometric framework

Given equation (2), we discuss its specification as a cointegrated system. It firstly considers the vector autoregressive (VAR) formulation and describes the corresponding vector error correction (VECM) representation. In Section 4, this model will then be applied to test the impact of defence spending on output in the US and UK.

Formally, we consider an extended VAR(p) specification for a $mx1$ vector of variables:

$$X_t = \mu_0 + \mu_1 T + \phi_h D_{th} + \sum_{i=1}^{p} A_i X_{t-i} + \varepsilon_t \quad t = 1, \dots\dots, T \qquad (3)$$

with $\mu_i = (0,1........)$ and $D_{th} = \begin{cases} 0, & t < h \\ 1, & t \geq h \end{cases}$

where μ_0 is a $mx1$ constant term, μ_1 is a $mx1$ coefficient vector related to the deterministic trend, D_{th} is a $dx1$ vector containing the likely presence of structural changes (shift dummies) and ϕ_h the corresponding mxd matrix of parameters[4]. A_i is a mxm matrix of unknown parameters, while ε_t is a Gaussian white noise process with covariance matrix Ω and p the lag order of the VAR. Equation (3) can be rewritten in a VECM form as:

$$\Delta X_t = \mu_0 + \Pi X^*_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \sum_{j=1}^{p-1} \gamma_j \Delta D_{t-j} + \varepsilon_t \qquad (4)$$

where $\Pi = (\sum_{i=1}^{p} A_i - I_p)$, $\Gamma_i = -\sum_{j=i+1}^{p} A_j$ and $\gamma_j = -\Gamma_j \phi$ with $j = 1.....p-1$. The matrix of parameters Π $(m\,x\,m+2)$ describes the long-run relationships of the VECM among the variables in vector $X^*_{t-1} = [X_{t-1}; D_t; T]'$. A necessary condition is that the polynomial characteristics associated with the VAR can determine the stability of the system. Γ_i refers to the short-run dynamics of the system ΔX_{t-1}, while ΔD_{t-j} characterises the persistence of a shock of the variables included in the cointegration space by means of the vector of shift dummy variables.

Under general conditions, the VECM equation (4) is I(1) and cointegrated and can be written as[5]:

$$\Delta X_t = \mu_0 + \alpha\beta * X^*_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \sum_{j=1}^{p-1} \gamma_i \Delta D_{t-j} + \xi_t \qquad (5)$$

where $\beta^* = [\lambda', \eta, \theta]$, $\eta = -\lambda' \mu_1$ and $\theta = -\lambda' \phi$. In equation (5) α is a mxr matrix, $\beta*$ is a $(m+2)xr$ matrix and $\mathrm{r}\,(0 < r < m)$ is the cointegration rank of the system.

[4] In the literature no exact definitions of structural breaks or structural changes have been given, since breaks or changes are interpreted as changes of regression parameters (Maddala and Kim, 1998). In what follows it is sufficient to refer to structural changes or structural breaks as changes of the deterministic components of the time series, such that the terms *breaks* and *changes* as equivalent.

[5] The set of the necessary and sufficient conditions so that Equation (4) is I(1) and cointegrated are: i) the roots of the characteristic polynomial are outside the unit circle; ii) $\Pi_t = \alpha\beta'$ where α and β' are matrices of full rank r', $0<r<m$; iii) the matrix obtained by multiplying the orthogonal complement of the matrix and the parameter matrix of long run is non-singular (Pesaran et al., 2000).

VECM equation (5) is the extended model of this article. The residual $rx1$ vector $u_t = \beta^{*'} X_t^*$ in equation (5) is trend-stationary and, under suitable unitary identifying normalization, can be interpreted as being a vector of deviations of observable variables from the long run equilibrium relationships.

With respect to the theoretical discussion in Section 2, we have assumed that the cointegrating rank is given by $r = 1$. The long run equilibrium levels are predicted by equation (5) by identifying the block decomposition of $X_t = (X_{1t}^{'}, X_{2t})^{'}$, where $X_{1t} = (Y_t)$ and $X_{2t} = (M_t, G_t, R_t)$ is the $3x1$ vector containing real defence and civilian spending and the real interest rate. The deviation of estimations from observable output can therefore be obtained as:

$$u_{t-1} = \beta * X_{t-1}^* = \left[I_1, -\lambda_2^{'}, -\vartheta, -\eta \right] \begin{bmatrix} X_{1t-1} \\ X_{2t-1} \\ D_t \\ T \end{bmatrix} = X_{1t-1} - \lambda_2^{'} X_{2t-1} - \vartheta D_t - \eta T \qquad (6)$$

As we shall see in the next section, it is possible that some institutional decisions regarding monetary or fiscal policies can modify the structure of long run patterns of time series. From an econometric point of view, their exclusion may be a cause of possible misspecifications of the model and of the inconsistency in the estimation results. In contrast, modelling structural changes influences the cointegrating rank inference. This question refers to the decision problem of whether one may still use the standard cointegration tests to avoid possible power losses and size distortions caused by modelling structural shifts or whether it is recommended to use cointegration tests that take such breaks into account. In the latter case, the proposals by Johansen et al. (2000) and Saikkonen and Lutkepohl (2000a) can be regarded as generalizations of the procedures by Johansen (1992 and 1995) and Saikkonen and Lutkepohl (2000b), respectively.

In order to empirically test the best dynamic specification that rationalizes the data, nested models are obtained by setting $\eta = 0$, in which the presence of a linear deterministic trend is excluded from equation (6), or by setting $\theta = 0$ where a model without a shift dummy is specified, or by a long run specification that restricts both the hypothesis tests. From the conditions to derive equation (6), it follows that a cointegrated system is obtained by a reduced rank of the Π matrix. In a parsimonious long run dynamic model, inference on the number of cointegration relationships can be carried out by testing the hypothesis:

$$H(r) = rank(\Pi) \le r \text{ against the alternative } H(m) = rank(\Pi) \le m \qquad (7)$$

for $r = 0,1,....m-1$. By maximizing the log-likelihood of equation (6) under both the null and alternative hypotheses, we derive statistics of the likelihood ratio or trace that have non standard distribution. Thus, the empirical specifications that include the presence of a constant term or linear deterministic trend uses the tabulate quantiles of the trace statistics derived by Johansen (1995) while, when a break(s) is incorporated in the level of the time series, the rank test is carried out by Saikkonen and Lutkepohl (2000a).

4. Testing model implications

4.1. Data

The data used for testing the model of Equation (5) for the two countries were obtained from different sources. For the US, quarterly data of the government sector at current prices are classified in defence and civilian categories of government spending and were available by the National Income and Product Accounts (NIPA), while the other US macroeconomic series and deflator were taken from the *International Financial Statistic* (IFS), reports redacted yearly by the International Monetary Found.

On the other hand, the data used for the UK economy are mainly taken from the (IFS). An issue with government spending data for the UK is that quarterly observations are only available for the series of total government spending, while disaggregated information on the composition of government spending, that includes defence spending, is available from different sources. For the aims of our analysis, we collected the annual time series data for the UK on the current government defence spending from the Stockholm International Peace Research Institute (SIPRI) and then split it up into the quarterly data.

Thus, higher frequency of data were obtained by disaggregating the relative annual series by the methods of Chow-Lin (1971), Fernandez (1981), Litterman (1983) and Santos Silva and Cardoso (2001), including alternative auxiliary time-series. The performance of each method was evaluated by means of the Akaike Information Criterion (AIC) and the Root Mean Squared Percentage Error (RMSPE), as in Aristei and Pieroni (2008). Based on the quarterly series of aggregate government spending as high-frequency indicator, the Chow-Lin (1971) method was preferred, though it is worth noting that the differences in the performance of the disaggregation methods are relatively small. As a result, quarterly time series of civilian spending of the UK is obtained from the difference between aggregate and defence spending.

For both countries, we transformed the original variables into real terms of logarithms by the index price for the GDP at the constant value of 2000, except the real interest rate[6]. On the other hand, the real interest rate was set up as the difference between the nominal three month treasury bill rate and the annual rate of growth in the consumption price index (*cpi*). It is worth noting that the data available on this indicator constrained the beginning of a more extended sample: for the US, the sample for the empirical tests spans from 1957:1 to 2005:4, while for the UK covers the period from 1957:1 to 1998:4[7]. Figure 1 and 2 describe the patterns of the four macroeconomic variables (in logarithm) that are included in Equation (5), for the UK and US, respectively. As it is possible to note in the first Figure, the GDP (top left) and non-military series (bottom left) in the UK sketch smooth growing patterns, with an unsurprisingly volatility stressed only in connection with unanticipated shocks.

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[6] Gold (2005) criticizes the results of Atesoglu (2004) derived by a chained price index for GDP to deflator military spending series. The core of his criticism is that since inflation in the defence sector has tended to out pace overall inflation, this may understate defence inflation and overstate the growth in defence spending. However, the use of deflator of the GDP in the US to obtain real values of the GDP is close to the results that are possible to obtain with a chained price index (Landefeld et al., 2003).

[7] It is worth noting that though recent works use data after 1998 for the UK (et al., 2004), the Stockholm International Peace Research Institute (SIPRI), suggests us avoiding their use after this point data since the time-series are not statistically updated.

Figure 1 − *Quarterly macroeconomic variables in logarithm for the United Kingdom*

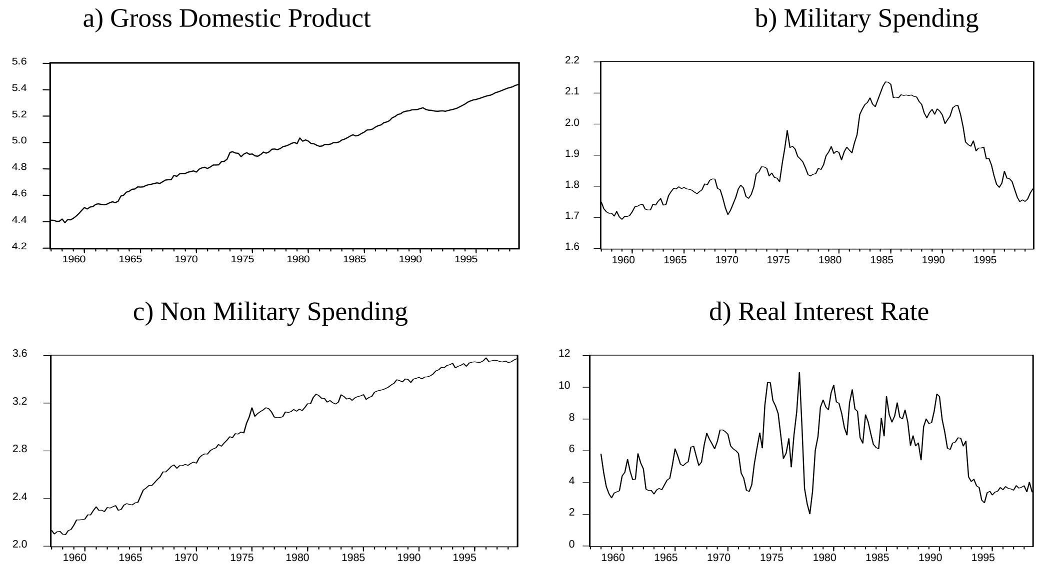
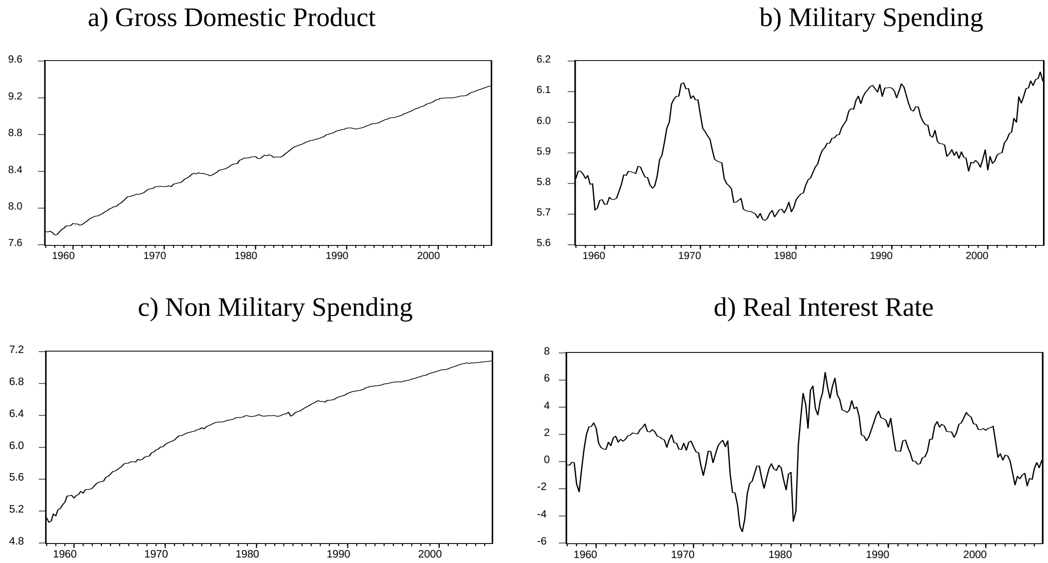


Figure 2 − *Quarterly macroeconomic variables in logarithm for the United States*



For example, it is plausible to argue that the sharp increase in the civilian spending in 1974 and 1975 were developed in response to an increase in the price of oil in October 1973 and, in turn, to the downturn business cycle. A completely different pattern characterizes the real defence spending time series (top right of Figure 1) with an inverse hump-shaped pattern. It is worth noting that defence spending has declined dramatically, both in real terms and as a share of the GDP, since the mid-1980s. The reasons strictly depend on changes of British government strategies. In particular, the 1985/86 peak in the pattern of defence spending marked the end of the UK’s NATO commitment, based on the increase of the 3% per annum of British government spending. Disarmament following the end of the Cold War resulted in reductions in UK defence spending of more than 20% in real terms between 1990 and 2000. However, this decline appears to have halted, and partly reversed, since the end of the 1990s (Emmerson et al., 2004). For these reasons, much of the explanatory power to describe the long run relationship between military spending and output might basically reside in the pattern of defence spending determined until the mid 80s. We analyzed the robustness of this hypothesis by estimating the defence spending and output elasticity for the UK in the sub-sample (1957:1-1985:4) and by comparing them with those of the full sample. In Figure 2 (top right) that concerns real defence spending in the US. What is immediately evident is that, while the long run pattern of defence spending has remained stable enough (or slightly increasing) in the last fifty years, the profile of the graph appears to be event-driven with large cyclical spikes corresponding to wars (or threat of wars) (Gerace, 2002; Gold, 2005). The levels of real government defence spending show that a sharp first peak in the data reflects the Vietnam War, a second one is in correspondence with the worsening tensions of the Cold War during Reagan's Presidency and the first Iraq attack while, after ten-years in which the defence spending dropped, there was an upswing in response to terrorist attacks[8].

In addition, the empirical studies have shown that the patterns of the US military spending and, in general, of the public sector during the 60s are unambiguously more volatiles and it might be responsible for a strong Keynesian stimulus with respect to the successive periods. This is, for example, the Gold's thesis (1997) that sustains that beginning from the 70s, a narrower and more stable range of the US defence spending (with respect to the long run trend of the output) generated an ineffective impact on the economy, the latter affected by the decline in output volatility over the past decades. In this regard, below we shall test the significance of defence spending on output for the sub-sample 1970:1-2005:4, assuming the presence of internal substitutability effects of the US government expenditure components (i.e. defence and civilian spending) in favour of the civilian sector. For this sub-period, long run output responses are, therefore, expected to be greater (and

[8] The events of war or the threats of war adopted in our sample, that lead to large military buildups, are close to the political events described by Ramey and Shapiro (1998) and used in Burnside et al. (2004) to identify changes in fiscal policy in the neoclassical context.

statistically significant) for trended-increased civilian spending with respect to the dynamics of the military sector. It is worth remarking that the central theme of Keynesian economics, associated with the effectiveness of fiscal policy as a stabilization tool, is maintained and fluctuations are, therefore, associated with variations in the efficiency with which productive resources are used. On the other hand, the statistical hypothesis of a non-stationary data-generating process for defence spending, as well as for the other variables of the model specified in equation (2), leads to the need to test the possibility of effects of the long run on output.

From a Keynesian point of view, a substantial decline in output volatility may also be attributed to better monetary policies. Martin and Rowthorn (2005) have documented that a rise or fall in the volatility of economies coincided with changes in inflation volatility, suggesting that this may have also been a contributing factor. Thus, starting from the assumption of the model in Section 2, we concentrate on the measure of real interest rates and its impact on monetary policies regarding output. The real interest rate patterns for the two countries (bottom right of Figure 1 and 2) reveal the presence of some volatility in the time series during the 70s. While it is known that exogenous shocks, caused by the oil crisis in 1973, invested countries over the world and, in turn, the sharp decrease in the real interest rate due to high levels of inflation, simple inspection of the US (Figure 2) shows the likely presence of a structural break, related to the fourth quarter of 1979, as a change in the manner of conducting monetary policy. In fact, Federal Reserve switched from pegging the Federal Funds' interest rate to a policy of reserve targeting, resulting in more variability in interest rates.

Finally, the drop in the real interest rate for the US economy, generated by the 2001 terrorist attacks, is in line with an exceptional active response of the FED to an unexpected negative impulse of the business cycle. As strongly suggested by Saikkonen and Lutkepohl (2000a), both of these break points will be used to assess the robustness of the long run relationship and the estimated parameters of the theoretical model. This is what we shall do in next sub-section.

4.2. Cointegrating Tests, Estimated Cointegrating Vectors and Policy Implications

Given $X_t = (X_{1t}^{'}, X_{2t})^{'}$, defined as before, the unrestricted VAR (Equation (3)) was estimated over the countries' samples. The number of lags (*p*) were not fixed a priori but derived by the information criteria. The parsimonious choice of lags, namely $p = 6$ and $p = 5$ for the complete sample of the US and UK, respectively, reveals that the disturbances of the unrestricted VAR model can be approximated as the realisations of a white noise multivariate process[9].

After fixing the lags of the VAR (and hence of the corresponding VECM parameterization), the analysis is carried out by selecting the cointegration rank of the system. Consistently with the Equation (6), a linear trend was restricted to belong to the cointegrated space for the US because it seems clear, on the basis of Figure 2, that at least three of four variables contain a deterministic trend. Moreover, as suggested by the descriptive analysis, we included a shift-dummy for a first break point related to the US monetary policy change in October 1979. This institutional change determined more variability in the level of the real interest rate, leading to the issue of rank instability in the cointegrating matrix. Finally, a second shift-dummy was included in the specification model to account for the 9/11 terrorist attack, as repeatedly used in studies that assessed the (economic) effects of this unexpected event (Blomberg et al., 2004; Virgo, 2001).

The first column of Table 1 reports the results from the US cointegration test over the entire sample. Let r denote the number of cointegrating vectors. As shown in the methodological section, the trace test is a sequential test that moves until the null hypothesis (7) cannot be rejected. For the entire data sample, the hypothesis of $H_0: \ r = 0$ is rejected at the 90% significant level. As for the presence of one cointegrating vector, the null hypothesis cannot be rejected at the usual significance level. Thus, in line with the previous empirically results of Atesoglu (2002) for the US economy, the data support the evidence of one cointegrating vector between endogenous variables, namely aggregate output and real interest rate, with I(1) fiscal policy shocks identified by the defence and

[9] The data used in this empirical section, as well as the estimations did not report to save space, are available for replication upon request.

civilian spending. This (Keynesian) evidence enables using this model to infer the relevant effects of the disaggregate measures of fiscal policy shocks.

The estimated parameters of the US cointegrating vector are given in Table 1 (bottom of column 1). As stated above, we use a maximum likelihood estimator to obtain the estimated elements of the cointegrating vector, while normalization has no impact on the information concerning structural parameters of the model reported in Appendix 1. In what follows, both for the US and, successively, the UK, the parameters associated with the aggregate output variable will be normalized to unity. Cointegration estimated parameters have signs consistent with those in equation (2) and their inference reveal a statistically significant relationship among the real variables of output, defence and civilian spending and interest rate. Moreover, the data provide fairly strong support that the US monetary policy change (October 1979) and the 9/11 terrorist attacks affect the long run equilibrium as well as it is statistically relevant to include the time trend.

Specifically, as in Atesoglu (2002), the significant effects of the US government defence spending on aggregate output seem to confirm the predictions of the theoretical model. However, taking the longest view first, our elasticity estimations are much weaker than Atesoglu (2002). The estimated value is reduced to 0.1 [with respect to Atesoglu's estimation, 0.57]. Inspection of the full sample in Figure 1 confirms a slight increase in the defence spending patterns, mainly sustained by the large military spending of the aforementioned political events. The intuition to test is that the presence of I(1) shocks of the US defence spending may be responsible of the slight positive relationship with aggregate output but, linked with Gold statement, this quantitative relationship may be sensitive to the sample-period. Thus, we re-estimate the long run model for the sub-sample from 1970:1 to 2005:4. The results presented in the second column of Table 1, without the presence of a time trend[10], support the hypothesis test, highlighting a much lower and insignificant defence spending impact on the economy, while it is registered a sharp increase in the impact of civilian spending (from 0.42 in the full sample to 1.10 in the sub-sample).

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[10] According to the hypothesis test, we cannot reject the hypothesis that the trend parameter η is not different from zero by a chi-square test.

Table 1 - *Cointegration Tests and Estimated Cointegrating Vectors*

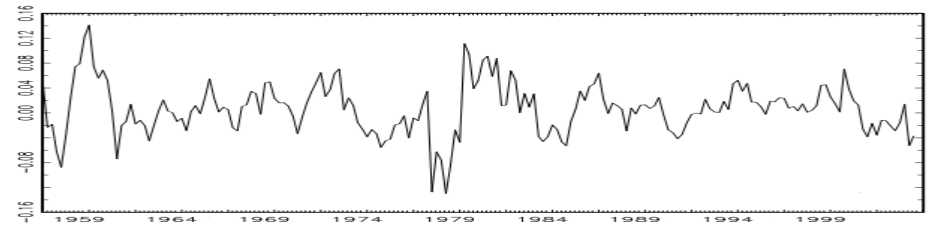
US (1957:1-2005:4)		US (1970:1-2005:4)		UK (1957:1-1998:4)		UK (1957:1-1985:4)	
Cointegration test statistics							
$H_0: r=0$	87.52	$H_0: r=0$	64.02	$H_0: r=0$	66.80	$H_0: r=0$	67.86
$H_0: r=1$	38.27*	$H_0: r=1$	28.21 *	$H_0: r=1$	19.91*	$H_0: r=1$	24.56*
$H_0: r=2$	21.31	$H_0: r=2$	12.25	$H_0: r=2$	9.35	$H_0: r=2$	10.26
$H_0: r=3$	8.60	$H_0: r=3$	0.68	$H_0: r=3$	1.77	$H_0: r=3$	4.68
Parameter Estimates							
β_1^*	-1	β_1^*	-1	β_1^*	-1	β_1^*	-1
β_2^*	0.105 *(0.066)*	β_2^*	0.053 *(0.410)*	β_2^*	0.085 *(0.818)*	β_2^*	0.375 *(0.077)*
β_3^*	0.426 *(0.000)*	β_3^*	1.106 *(0.000)*	β_3^*	0.904 *(0.000)*	β_3^*	0.528 *(0.000)*
β_4^*	-0.015 *(0.001)*	β_4^*	-0.010 *(0.053)*	β_4^*	-0.026 *(0.006)*	β_4^*	-0.008 *(0.024)*
ϑ_1		ϑ_1		ϑ_1	-0.729 *(0.000)*	ϑ_1	-0.188 *(0.000)*
ϑ_2	0.150 (0.000)	ϑ_2	0.121 (0.001)	ϑ_2		ϑ_2	
ϑ_3	-0.063 (0.053)	ϑ_3	-0.019 (0.565)	ϑ_3		ϑ_3	
μ	4.636 (0.000)	μ	1.062 (0.003)	μ	1.809 (0.001)	μ	2.540 (0.000)
η	0.005 (0.000)	η		η		η	

Note: The trace test statistic for cointegration and the maximum likelihood estimator for the cointegrating vector are obtained from Johansen (1995). Test statistics are adjusted for the presence of a structural break in the time series (Johansen et al., 2001). An asterisk (*) in the upper part of the Table indicates that the null hypothesis over the rank of cointegration cannot rejected at 90% significant level, while round brackets in the estimated parameters show the p-values.

Though this result may solve the puzzle of the effect of defence spending generated by the gap of data inspection and empirical results (Gold, 2005), the increase in civilian spending complementarities on aggregate output needs an explanation for the central role it assumes in the economy and for its relevant fiscal policy implications. It is empirically documented that the decline in the pattern of US defence spending was substituted by an increasing civilian investment in new technology. This different government allocation shifted the military sector's central role to one of creating spin-off and complementary relationships of demand in the economy towards the civilian sector and showed the switch of the defence to an economically mature sector[11].

Figure 3 - ***Vector Error Correction model for the United States***

a) Full sample (1957:1-2005:4)



b) Sub-sample (1970:1-2005:4)

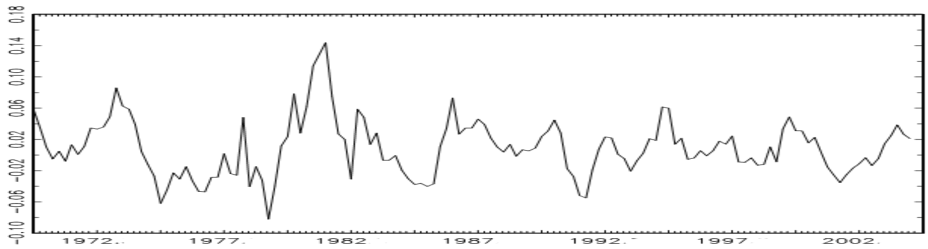
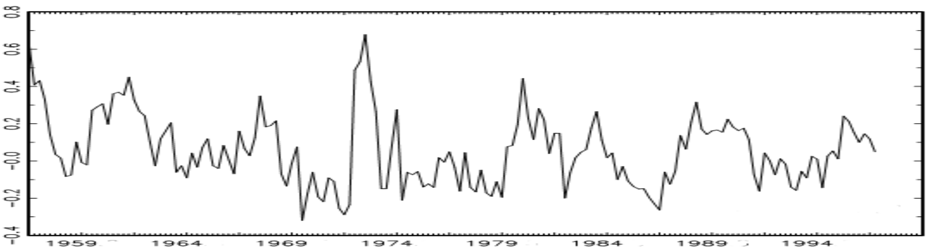
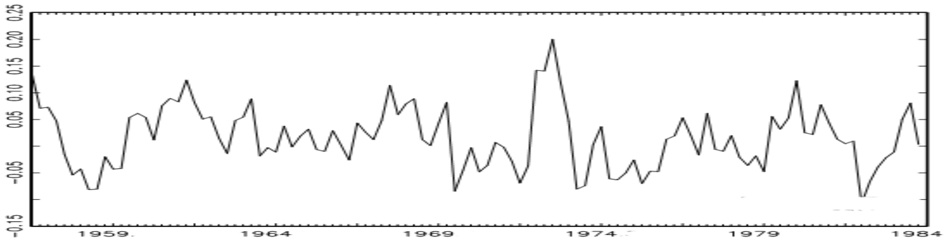


Figure 4 - ***Vector Error Correction model for the United Kingdom***

a) Full sample (1957:1-1998:4)



b) Sub-sample (1957:1-1985:4)



[11] The evidence suggests that the dynamics described made the policy makers aware of appropriateness of policies even if the spending induced by the war's lobby and defence industry are relevant components of the US economy. As an example, it is known that the reaction after the 9/11 terrorist attack to account for the predictable downturn business cycle cut the interest rate and increased the level of government defence spending. The latter policy, financed by a federal government debt was, however, perceived as a temporary event addressed to guarantee national and international security. Contrary to the expectations of government spending substitutability, was documented the constant growth of non-defence category around its equilibrium pattern justifying the leading role in a new-Keynesian perspective.

In conclusion, the long run policy mechanism at work seems, therefore, to structurally sustain greater returns from civilian investments even with political events (wars or threat of wars) that increased US military spending.

However, this may only be a part of the story. More convincingly, the fiscal and monetary policy responses, designed as tools for aggregate demand management, are jointly responsible for the fall in discretional defence spending changes in US output volatility. A source for such empirical result may be identified in the role of monetary policy as one determinant of economic and political stability. The increased credibility of inflation targets and the role of the central banks in the last two decades have been considered as being responsible for the fall in economic volatility (Martin and Rowthorn, 2005). We find confirmation of this assumption in Table 1, where the significance of the real interest rate for both empirical specifications is a strong support for the model specification and shows its relevance for policy-makers as a countercyclical tool.

One purpose of this article is to provide some comparative evidence of the impact of defence spending on output between the US and UK. We anticipate that there is no evidence of changes in the number of cointegrating vectors (1 for each specification) in the VAR. As a confirmation, we reported the estimated vectors of the error correction models both for the complete and sub-samples of the two countries (Figures 3 and 4). In all cases the estimated residuals range around the long run equilibrium patterns.

Viewed from the perspective of the estimated parameters reported in the last two column of Table 1, the long run specification gives a good description of the UK defence spending-output relationship only in one historical period preceding the end of the UK’s NATO commitment, namely in the period 1957:1-1985:4, characterized by a sharp rise of budget in military spending.

In fact, inspection of the results in the full sample (column 3) reveal that, among the variables of interest, the estimated parameter of defence spending is close to zero with implausible large standard errors, such as its impact on output seems to be irrelevant. Namely, much of the long run impact is observed in other estimated parameters. The significant large and positive values of elasticities in the civilian spending and the negative impact of the real interest rate are in line with a Keynesian theoretical perspective discussed for the US case. However, the sub-sample results in column 4 of Table 1 do provide striking evidence of substantial changes in the parameter estimates for (β_2), with a significant impact of defence spending on aggregate output. These estimates are particularly plausible and can be used to infer the relevance of government changes on fiscal policy. As it implies that defence spending strongly impacts on the economy before the end of the Cold War while, overall during the 90s, data inspections show that the UK government's choices made it seem no longer willing to sustain high defence spending when the (international) threat of security had become low.

On the other hand, as it is frequently argued in mainstream economics literature and shown in this paper, the "positive analysis" of the impact of defence spending on economic performance depends on the sample period that, like a picture, summarizes the structural economic changes and political events of the country(ies). It is known that the response to the 9/11 terrorist attack militarily involved the UK because of its international role and its strong political agreements with the US in the defence sector. Thus, the costs of preventive terrorist actions and the UK's military obligations in the Iraq war sharply rose its defence budget in 2002. Contrary to the politicians' statements, the sizeable military presence in Iraq also generated a high level of public defence spending in the following years (Emmerson et al., 2004). On the other hand, the recent increase in defence spending of the UK, though focused on the aim of international political stability, stimulated discussions concerning defence policy sharing, i.e. EU defence policy, and the central role that a renewed defence industry might have on stimulating the economy's demand-size. As described by Kollias et al. (2004), the auspicated EU defence policy, that in principle should increase the defence spending of each EU country, might be politically justified by improving the efficiency of defence industries and military R&D. Aggregate demand and multiplier effects may, therefore, be generated by increased capacity utilisation and technological innovation and be the economic foundation to sustain an efficient a trans-Atlantic defence market (Hartley, 2003).

Therefore, in this context of *new military Keynesianism* of the first years of the 21st century in the UK, an extension of our sample period may predict an under-estimation of the long run complementarity relationships between defence spending and output and, therefore, its economic mechanism would be the reasons to survive.

5. Concluding remarks

This paper aims to empirically test whether government defence spending, as a component of public spending, significantly affects the long run aggregate output pattern. We use a Keynesian theoretical framework that explicitly account for its potential role in explaining fiscal policy fluctuations. On the other hand, since the components of fiscal policy shocks are identified as the motivating force for the nonstationarity of aggregate output, a stable long run relationship among the macroeconomic variables is a necessary condition to accomplish their impact.

The econometric results are carried out for the US and UK over the periods 1957-2005 and 1957-1998, respectively, while a sensitivity analysis is included by estimating the theoretical model for sub-samples and by including shift dummies to account for institutional or policy changes.

By discussing the empirical results, we found that aggregate data provides consistent evidence that defence spending, as well as civilian spending are cointegrated with output and real interest rate, in line with the theoretical suggestions both for the US and UK economies. On the other hand, answering the question whether defence spending provides economic stimulation is more complex. Although we obtain a positive and significant impact of government defence spending on output, that supports the hypothesis of a military Keynesianism, underpinnings the dimension and the pattern of elasticities for sub-samples, the hypothesis at work becomes questionable. For the US, the estimated elasticity of government defence spending on output is really low. Even if the dimension of impact might be surprising for some, these estimates are highly in line with the descriptive evidence of the time series. More than a part of the long run pattern between government defence spending and output, the significance of this elasticity appears linked with the persistence in event-driven government spending. On the other hand, the positive response of the output to shocks of military spending in the UK estimations seems to depend on international agreement after the Second World War. This paper has shown that government choices in the mid 80s to give up the UK's NATO commitment have determined a decrease in military spending. Switching government priorities in favour of supplying civilian goods and services rather than financing federal defence spending may be responsible for significant fall in output elasticities.

Given these dynamics, a straightforward prediction of a revised and declining role of the defence sector for the two economies can be made. However, under the threat of international terrorism, new army policy initiatives (and the consequent rise in the defence spending) were announced between the end of 2001 and the middle of 2002, so that government priorities regarding international security may revitalize the pro-cyclical effects of the military sector on aggregate output. Because of the robustness of our findings, any sample extensions are left for future work.

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APPENDIX 1 – *The IS-MP model*

In line with theoretical suggestions of Romer (2000) and Taylor (2000), the empirical specification for testing the impact of defence spending on aggregate real output in Equation (1) is build up by a new macroeconomic Keynesian model under the hypothesis that the real interest rate, R, is given. Below, we sketch the straightforward structural cross model in which the variables are expressed in real terms.

$Y_t = C_t + I_t + X_t + M_t + G_t$ $\quad$ Y_t = Aggregate output; C_t = Consumption; I_t = Investment; X_t = Real Net Export; M_t = Defence Spending; G_t = Civilian Spending;

Consumption Function:

$C_t = d + e(Y_t - T_t)$ $\quad$ T_t = Real Taxes;

Tax Function:

$T_t = n + gY_t$

Investment Function:

$I_t = h - iR_t$ $\quad$ R_t = Real interest rates;

Net Export Function:

$X_t = l - mY_t - nR_t$

The parameters of the reduced form of Equation (1) in the body of the text, $\beta_0^*, \beta_2^*, \beta_3^*$ and β_4^*, can, therefore, be determined by substituting the aforementioned functions of the economic aggregates into the output-income equation. Formally we obtain an extended relationship as given:

$$Y_t = d + e(Y_t - (n + gY_t)) + (h - iR_t) + (l - mY_t - nR_t) + M_t + G_t$$

Finally, solving the equation for Y_t, we obtain:

$$\beta_0^* = \frac{d - en + h + l}{1 - e(1+g) + m}; \quad \beta_2^* = \beta_3^* = \frac{1}{1 - e(1+g) + m}; \quad \beta_4^* = \frac{-(i+n)}{1 - e(1+g) + m};$$

that represent the parameters to estimate in Equation (2).

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