Military Spending and the Growth-Maximizing Allocation of Public Capital: A Cross-Country Empirical Analysis

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Abstract

In this paper drawing from the theoretical framework developed by Shieh et al., (2002), we present an endogenous growth model to empirical analyze the growth maximizing allocation of public capital among military spending and investment in infrastructure. Using this general model of public capital formation, we derive the growth-maximizing values of the shares of public capital allocated to it's two different types, as well as the growth-maximizing tax rate (amount of total public capital as a share of GDP). Then we proceed with an empirical investigation of the theoretical implication of the model that both the effects of the shares of public capital and the tax rate on the long-run growth rate are non-linear, following an inverse U-shaped pattern. Using data of public investment in infrastructure and military capital formation, we investigate the long run relationship between economic growth and the allocation of public capital using panel cointegration analysis in a sample of 55 developed and developing countries. Our empirical results confirm the theoretical implications of the model for the majority of the countries in the sample. This finding is more consistent for the OECD countries although the same result can be drawn for a large part of the developing countries. .

Keywords: *public capital, military spending, economic growth, panel unit root tests, panel cointegration.*

JEL codes: E62, H56, O40, C23.

1. Introduction

Over the last three decades, after the publication of Aschauer´s (1989) empirical paper on the productivity of public capital and Barro´s (1990) theoretical paper on the effects of government spending on economic growth, the analysis of the macroeconomic effects of public investment has attracted research interest in growth economics. The theoretical research was mainly focused on analyzing how public spending and public capital may enhance productivity and promote economic growth. At the same time, the empirical literature was trying to confirm the existence of a positive empirical relationship between productive public spending and economic growth, by employing a variety of econometric techniques, data sets, and measures of public spending. In principle, public spending enhances economic growth through its external effect in the production function of private firms. This effect can be modeled by adding into the production function either the aggregate flow of public spending, following Barro (1990), or the aggregate stock of public capital, as in Turnovsky (1997). A new line of theoretical research recognizes the possibility that different types of public spending (*e.g.*, infrastructure, education, health, military expenses) may exert a different effect on economic growth. In this line of research, Devarajan, Swaroop and Zou (1996) develop an endogenous growth model using a production function which includes two types of government spending, Kalaitzidakis and Kalyvitis (2004) into infrastructure and maintenance expenditure, and Chen (2006) into productive and consumptive spending.

Largely overlooked in this strand of the literature has been the question of the economic growth impact of military spending. This oversight is surprising given the significant portion that military spending absorbs from public capital in both developed and developing countries. This question has received substantial attention in the early economic growth literature especially in developing countries initiated by Benoit's (1973; 1978) empirical study. Benoit (1973; 1978), in her pioneering work in this area using data from 44 developing countries, found that defense spending does indeed stimulate economic growth. Government provision of national defense which maintains property rights can be viewed as maintaining both internal and external security, increasing hence the incentives to accumulate private capital and attract foreign investments. On the other hand, the defense sector also provides a variety of public infrastructure like roads communication networks etc, while also enhances human capital accumulation through training and provision of educational services. These are at least some ways that military spending can foster economic growth domestically. However, military spending also diverts resources that could more productively be employed in sectors of the economy other than defense. In that sense, defense spending may degrade long run economic growth. Hence the net effect that military spending has on economic growth appears to be ambiguous. This results has been confirmed by many subsequent empirical studies devoted in addressing the validity of *Benoit hypothesis* utilizing a variety of empirical models and different data sets.[1]

The empirical research on the productive effects of public capital has followed two basic lines of research. The one line of research, following Benoit (1973; 1978) directly estimates aggregate production functions to identify the productive types of public capital. The most indicative examples in this strand of the literature are the studies by Baffes and Shah (1998) and Evans and Karras (1994). The empirical results often validate Benoit's initial assumption on the productivity of military spending. The other line of research estimates Barro-type growth regressions in which different types of public expenditure may serve as explanatory variables, as in Barro and Sala-i-Martin (1995, ch. 12). In most cases the different types of public expenditure enter linearly into the growth regression, as, for example, in Devarajan, Swaroop and Zou (1996). The results usually indicate a non significant or negative effect of public expenditure variables on economic growth. These results are interpreted as an over-provision of public capital.

Based on this conflicting empirical evidence, recently Shieh *et al.*, (2002) developed an intertemporal optimizing endogenous growth model to analyze how the government's resource allocation between the defense and the non-defense sectors affect economic growth and social welfare. They concluded that there is indeed an optimal share of defense expenditures that maximize the economic growth rate but this rate is smaller than the welfare maximizing share. In the same line of research and building upon Shieh *et al.*, (2002) theoretical development, the present paper derives the growth-maximizing values of the shares of public investment allocated to the two different types of public capital (*i.e.*, military spending and public infrastructure), as well as the growth-maximizing tax rate. Then the paper proceeds with an empirical investigation of the theoretical implication of the model assuming that both the effect of the share of public investment and that of the tax rate on the long-run growth rate are non-linear, following an inverse *U*-shaped pattern.

Using data of public capital and military capital formation from a sample of both developed and developing economies, we derive empirical estimations that confirm the theoretical implications of the model. The innovative feature of our empirical analysis is that we test a specific non-linear functional specification implied by the theoretical model. In addition tackling the problem of spurious regression arising from the non-stationarity of the data set, the empirical part of the paper makes use of panel unit-root tests and cointegration analysis to conclude that there is a fairly strong evidence in favor of the hypothesis that long run causality exists between the optimum allocation of government expenditures among military spending and public infrastructure.

The rest of the paper is organized as follows. Section 2 sets up a representative firm model and derives the long-run equilibrium conditions, analyzing the transitional dynamics of the economy. It also derives the growth-maximizing levels of the tax rate and the shares of public capital investment. Section 3 sets up the econometric procedures followed in the empirical part of the paper which is based on panel cointegration techniques. Section 4 presents an empirical investigation of the theoretical model in a sample of developed and developing economies, while Section 5 concludes the paper.

2. Theoretical Framework

In this section we present a simple decentralized version of the Shieh et al (2002) endogenous growth model, to derive the testable hypothesis of the model. We consider a closed economy populated by identical agents who consume and produce a single commodity, *Y*. There is no population growth. The labor force is equal to the population, with labor supplied inelastically. On the production side of the closed economy, the representative firm *j* produces its output, Y_j, using a conventional Cobb-Douglas technology:

$$Y_j = K_j^{\alpha}\left(hL_j\right)^{1-\alpha} \tag{1}$$

where $0<\alpha<1$, K_j denotes the stock of private capital, and L_j the labor used by firm *j*. The productivity of labor, *h*, is a function of the existing stock of public infrastructure, *Z*, and military capital, *M*, per worker so that:

$$h = \frac{Z^{\beta}M^{1-\beta}}{L} \tag{2}$$

where *L* is the total labor force, and $0<\beta<1$. The individual firm takes *h* as given. Private capital depreciates at the constant rate δ_k. Therefore, letting *I* denote gross private investment, the net private capital stock accumulates at the rate:

$$\dot{K} = I - \delta_k K \tag{3}$$

New output may be transformed to any type of capital, but in the case of private capital this process involves adjustment costs. The cost of investment faced by domestic firms is:

$$\Psi\left(I,K\right)=\left(1+\frac{\phi}{2}\frac{I}{K}\right)I \tag{4}$$

where $\phi>0$ is the adjustment cost parameter. As typically put forward by the relevant literature, the adjustment cost of private capital is proportional to the rate of investment per unit of installed capital.[2]

Infrastructure and military capital depreciates at rates δ_Z and δ_M, respectively. If we let G_Z and G_M denote gross public investment for infrastructure and military capital, respectively, then the net stock of each type of public capital accumulates as follows:

$$\dot{Z}=G_Z-\delta_Z Z \tag{5}$$

$$\dot{M}=G_M-\delta_M M \tag{6}$$

The government finances its total expenditure (investment in both types of public capital) through tax revenues collected via a tax rate τ imposed on total output produced by domestic firms. Hence, the government budget constraint is:

$$G_Z+G_M=\tau Y \tag{7}$$

If we define the share of total government expenditure that goes towards military capital formation as μ, then:

$$G_M=\mu\tau Y \quad \text{and} \quad G_Z=\left(1-\mu\right)\tau Y \tag{8}$$

We assume, for the moment, that both the tax rate τ and the share μ are fixed and constant over time. So the government can set both variables arbitrarily. Since, however, these policy instruments are going to affect the long-run growth rate of the economy, the tax rate and the component shares that maximize the growth rate will be derived later on.

The representative firm j in our economy solves the following infinite horizon profit maximization problem:

$$
\begin{aligned}
&max\int_0^{\infty} e^{-rt}\left[\left(1-\tau\right)Y_j - wL_j - \left(1+\frac{\phi}{2}\frac{I_j}{K_j}\right)I_j\right]dt \\
&s.t. \quad \dot{K}_j = I_j - \delta_k K_j
\end{aligned} \tag{9}
$$

where r is the real interest rate, w is the real wage rate, while the price of the commodity is normalized to one. The familiar optimality conditions with respect to I_j, and K_j are respectively:

$$1+\phi\frac{I_j}{K_j} = q \tag{10}$$

$$r = \frac{1}{q}\left[\dot{q} + \left(1-\tau\right)\alpha\left(\frac{K_j}{L_j}\right)^{\alpha-1} h^{1-\alpha} + \frac{\phi}{2}\left(\frac{I_j}{K_j}\right)^2\right] - \delta_k \tag{11}$$

where q is the shadow value of the private capital stock. Equation (10) equates the marginal cost of investment to the shadow value of capital, while equation (11) is the arbitrage condition that equates the interest rate to the rate of return of private capital, net of physical depreciation. The rate of return to private capital consists of three components: the change in its shadow value, the value of its marginal product, and its effect on the cost of investment.

Substituting (10) into (11), replacing for (2) and aggregating across firms, the optimality conditions with respect to I and K can be written as:

$$\frac{I}{K} = \frac{q-1}{\phi} \tag{12}$$

$$\dot{q} = (r+\delta_k)q - (1-\tau)\alpha\left(\frac{Z}{K}\right)^{\beta(1-\alpha)}\left(\frac{M}{K}\right)^{(1-\beta)(1-\alpha)} - \frac{(q-1)^2}{2\phi} \quad (13)$$

From equations (3), (5), (8), and (12), the growth rates of private and the two types of public capital considered herein are given by:

$$\frac{\dot{K}}{K} = \frac{q-1}{\phi} - \delta_k \quad (14)$$

$$\frac{\dot{Z}}{Z} = (1-\mu)\tau\frac{K}{Z}\left(\frac{Z}{K}\right)^{\beta(1-\alpha)}\left(\frac{M}{K}\right)^{(1-\beta)(1-\alpha)} - \delta_Z \quad (15)$$

$$\frac{\dot{M}}{M} = \mu\tau\frac{K}{M}\left(\frac{Z}{K}\right)^{\beta(1-\alpha)}\left(\frac{M}{K}\right)^{(1-\beta)(1-\alpha)} - \delta_M \quad (16)$$

Now, let us define z as the ratio of infrastructure to private capital stock, $z \equiv \frac{Z}{K}$, and m as the ratio of military to private capital stock, $m \equiv \frac{M}{K}$. Then, using relations (13), (14), (15) and (16) we get:

$$\frac{\dot{z}}{z} = -\delta_Z - \frac{q-1}{\phi} + \tau\frac{1-\mu}{z}z^{\beta(1-\alpha)}m^{(1-\beta)(1-\alpha)} + \delta_k \quad (17)$$

$$\frac{\dot{m}}{m} = -\delta_M - \frac{q-1}{\phi} + \tau\frac{\mu}{m}z^{\beta(1-\alpha)}m^{(1-\beta)(1-\alpha)} + \delta_k \quad (18)$$

$$\dot{q} = (r+\delta_k)q - (1-\tau)\alpha z^{\beta(1-\alpha)}m^{(1-\beta)(1-\alpha)} - \frac{(q-1)^2}{2\phi} \quad (19)$$

The stationary solution of the above system of differential equations must have at least one real solution, in order for output and the capital stocks to follow a balanced growth path. Under the simplifying assumption, $\delta_k = \delta_Z = \delta_M = \delta$ the equilibrium values of q, z and m are jointly determined by the following relations:

$$\dot{z} = 0 \Rightarrow q = 1 + \phi\tau \frac{1-\mu}{z} z^{\beta(1-\alpha)} m^{(1-\beta)(1-\alpha)} \tag{20}$$

$$\dot{m} = 0 \Rightarrow q = 1 + \phi\tau \frac{\mu}{m} z^{\beta(1-\alpha)} m^{(1-\beta)(1-\alpha)} \tag{21}$$

$$\dot{q} = 0 \Rightarrow (1-\tau)\alpha z^{\beta(1-\alpha)} m^{(1-\beta)(1-\alpha)} = (r+\delta)q - \frac{(q-1)^2}{2\phi} \tag{22}$$

Then, from equations (20) and (21) we get:

$$\frac{z}{m} = \frac{1-\mu}{\mu} \tag{23}$$

Substituting equation (23) into equations (21) and (22), we get:

$$q = 1 + \phi\tau (1-\mu)^{(1-\beta)(1-\alpha)} \mu^{1-(1-\beta)(1-\alpha)} z^{-\alpha} \tag{24}$$

$$(1-\tau)\alpha \left(\frac{1-\mu}{\mu}\right)^{(1-\beta)(1-\alpha)} z^{1-\alpha} = (r+\delta)q - \frac{(q-1)^2}{2\phi} \tag{25}$$

By totally differentiating the above equations with respect to τ and μ, and taking into account the fact, implied by equation (14), that the growth rate of the economy is an increasing function of the shadow price of private capital, we get:

$$\frac{dg_Y}{d\tau} \begin{cases} >0 & if \quad \tau < 1-\alpha \\ =0 & if \quad \tau = 1-\alpha \\ <0 & if \quad \tau > 1-\alpha \end{cases} \tag{26}$$

$$\frac{dg_Y}{d\mu} \begin{cases} >0 & if \quad \mu < 1-\beta \\ =0 & if \quad \mu = 1-\beta \\ <0 & if \quad \mu > 1-\beta \end{cases} \tag{27}$$

In accordance with the theoretical developments made by Shieh *et al.*, (2002), the above relations indicate that the effects of both τ and μ on the growth rate are non-linear, following an inverse *U*-shaped pattern. When these two policy variables are relatively small (large), an increase in their values raises (reduces) the growth rate of the economy.

3. Data and Econometric Model

For the quantitative assessment of the effect of the allocation of public investment on GDP growth we utilized a balanced data set of 17 OECD and 38 Non-OECD countries[3] covering the period from 1980 to 1995. The data on GDP at constant 1995 international prices, private and public investments and military expenditures are obtained from the *Global Development Network Growth Database* developed by the *World Bank*. The sample size over both countries and time was restricted by data availability on military spending by individual countries. To investigate the relationship between economic growth and the allocation of public capital between military expenditures and public infrastructure we proceed to the econometric estimation of a growth equation with both policy parameters appearing quadratically exhibiting thus an inverse *U*-shaped pattern. In addition, we include two other control variables in our growth regression equation: (*a*) the share of private investment to GDP, which has been shown to be a robust explanatory variable of GDP growth, and (*b*) the real per capita lagged GDP to capture any convergence process in the sampled countries. In particular we use the following model:

$$\hat{g}_{it} = \beta_{0i} + \beta_{yi} y_{it-1} + \beta_{zi} z_{it} + \beta_{ki} k_{it} + \beta_{k^2 i} k_{it}^2 + \beta_{mi} m_{it} + \beta_{m^2 i} m_{it}^2 \qquad (28)$$

where $\hat{g}_{it}$ is the growth rate of real per capita GDP for country $i = 1,\ldots,N$ at year $t = 1,\ldots,T$, y_{it-1} is the one-period lagged real per capita GDP, $z_{it} = I_{it}/y_{it}$ is the share of private investment to GDP, $k_{it} = G_{it}/y_{it}$ is the share of total public spending to GDP and serves as a proxy to the tax rate τ, $m_{it} = M_{it}/G_{it}$ is the share of military expenditures to total public spending, and β*'s* are the parameters to be estimated. The regression parameters, β*'s*, are all country-specific allowing thus for intra-country heterogeneity on the effect of the explanatory variables on the growth rate of real per capita GDP.

Relation (28) is considered as long-run equilibrium relationship of the optimum allocation of public capital and GDP growth. Provided that all variables in (28) are integrated of order one, *i.e.*, $I(1)$, valid economic inferences on this optimum allocation of public expenditures can be drawn *iff* this relation is indeed a cointegrated relation.[4] Otherwise spurious inference would be obtained. Given the short span of the data series, the use of conventional cointegration testing like *Dickey-Fuller* tests or *Johansen's* maximum likelihood cointegration methodology may result to invalid conclusions.[5] Instead the use of recently suggested panel based unit-root and cointegration tests result in a more valid statistical inference as they take into account in the most efficient way all the available statistical information contained in the data set. Before proceeding to the identification of the long-run relationship between economic growth and allocation of public expenditures among infrastructure and military spending, we need to verify that all variables in (28) are integrated of order one. To test this hypothesis, we employ the panel unit root tests of Im *et al.*, (2003) and Maddala and Wu (1999). These tests have an advantage over earlier generation tests such as Breitung and Meyer (1994), Quah (1994) and Levin *et al.*, (2002) in that they allow for greater flexibility under the alternative hypothesis.[6]

First, Im *et al.*, (2003, IPS herafter) using the maximum likelihood framework suggest a procedure based on averaging individual unit-root test statistics for panels. Specifically, their *t-bar* test statistic is based on the average of augmented Dickey-Fuller (ADF) statistics computed for each cross-section unit in the panel. Their approach allow for residual serial correlation and heterogeneity of the dynamics and error variances across groups. In summary to compute the statistic, one first need to estimate the individually ADF regressions for each of the i countries in the sample and then to construct the N corresponding ADF t-statistics. These individual statistics are averaged to obtain the IPS *t-bar* statistic. However, since the distribution for the individual ADF t-statistics are not centered around zero under the unit root null hypothesis, it is necessary to be standardized to ensure that the distribution of the resulted statistic does not diverge as the number of individual countries in the panel increases. Hence, the IPS standardized *t-bar* statistic is specified as follows:

$$W_{t-bar}\left(p,\rho\right)=\frac{\sqrt{N}\left(t-bar_{NT}-N^{-1}\sum_{i=1}^{N}E\left[t_{iT}\left(p_i,\rho_i=0\middle|\beta_i=0\right)\right]\right)}{\sqrt{N^{-1}\sum_{i=1}^{N}E\left[t_{iT}\left(p_i,\rho_i=0\middle|\beta_i=0\right)\right]^2}}\rightarrow N\left(0,1\right) \quad (29)$$

where $t-bar_{NT}=N^{-1}\sum_{i=1}^{N}t_i\left(\rho_i\right)$ is the average of individual ADF *t*-statistics t_i, ρ_i is the lag order of the individual $ADF\left(p_i\right)$ regressions with p_i being the order of AR term for the i^{th} individual in the sample. When the lag order is zero for all individuals, then the values of the moments $E\left[t_{iT}\right]$ and $E\left[t_{iT}\right]^2$ are obtained by Monte-Carlo simulations for different values of T and are tabulated by Im *et al.*, (2003).[7]

On the other hand, the Maddala and Wu (1999, MW hereafter) suggest, instead of averaging the individual ADF *t*-statistics, to utilize the pooled values of the associated individual marginal significance levels. For the N countries in the panel, we can sum these values to obtain the *Pearson-lambda* statistic, also commonly referred to as the *Fisher* statistic, which becomes:

$$P=-2\sum_{i=1}^{N}ln\,p_i \qquad (30)$$

with p_i being the *p*-values obtained from the individual ADF *t*-statistics. Since the marginal significance levels, for the individual ADF tests are uniformly distributed within the $(0,1)$ interval, this implies that the above test for each individual is distributed as a *chi*-squared with two degrees of freedom. Again, under the assumption that the individual statistics are independent, implies that the resulted Fisher statistic in (30) is distributed as a *chi*-squared with degrees of freedom twice the number of countries in the sample (*i.e.*, 2*N*). Breitung (1999) found that the IPS *t-bar* test statistic is sensitive to the specification of individual deterministic trends loosing it's explanatory power. On the other hand, the MW test statistic is not dependent on the lag lengths used in estimating the individual ADF tests, while on the other hand it takes into account the role of individual countries in contributing to the overall results for the panel.[8]

Establishing the existence of a unit-root, the next step is to statistically examine the long-run relationship among GDP growth, the share of public capital allocated in military spending, the GDP share of total public spending, and the control variables, *i.e.*, the share of private investment and the one-period lagged real per capita GDP. Given again the short span of our data set, the power of the traditional Johansen's ML procedure is severely distorted. To overcome this problem we utilize the panel cointegration test due to Levin *et al.*, (2002) which is based on the estimated residuals of the following long-run relation:

$$y_{it} = \delta y_{it-1} + \sum_{j=1}^{k} \gamma_j z_{it} + \varepsilon_{it} \tag{31}$$

where z_{it} are the $j = 1,\ldots,k$ deterministic control variables, and ε_{it} is usual *iid* error term. Then the *t*-statistic on δ is computed from the following relation:

$$t_\delta = \frac{\left(\hat{\delta} - 1\right)\sqrt{\sum_{i=1}^{N}\sum_{t=1}^{T} \tilde{y}_{it-1}^2}}{s_\varepsilon} \tag{32}$$

where, $\hat{\delta}$ is the OLS estimate of δ, $\tilde{y}_{it} = y_{it} - \sum_{s=1}^{T} f\left(t,s\right) y_{is}$, $\tilde{\varepsilon}_{it} = \varepsilon_{it} - \sum_{s=1}^{T} f\left(t,s\right)\varepsilon_{is}$, $f\left(t,s\right) = z_t'\left(\sum_t z_t z_t'\right) z_s$ and $s_\varepsilon^2 = \left(NT\right)^{-1} \sum_{i=1}^{N}\sum_{t=1}^{T} \tilde{\varepsilon}_{it}^2$. To deal with the heterogeneity among countries in the sample that the above test-statistic does not take into account, we use *Fisher's* test to aggregate the individual *p*-values of Johansen's maximum likelihood cointegration test statistics as suggested by Maddala and Kim (1998, p. 137). The test is computed in a similar manner with that in relation (30) with degrees of freedom equal again with twice the number of countries in the sample.

Once the long-run relationship between the GDP growth and the share of public capital allocated in military spending and public infrastructure is established, the next step involves the econometric estimation of relation (28). Among the existing candidates, Pedroni's (1996; 2000) between dimension fully modified OLS estimator (FMOLS hereafter) is the best alternative as it takes into account the dynamic heterogeneity in cointegrated panels allowing the transitional dynamics to be different among different countries in the sample. In summary, the FMOLS estimator addresses consistently the problem of non-stationary regressors as well as that of possible simultaneity bias. An important advantage of between dimension FMOLS estimator compared to within dimension FMOLS is that is permits for greater flexibility in the presence of heterogeneity of the cointegrating vectors. Specifically, the test statistics constructed to statistically examine the cointegrating slope parameters of the model are not constrained to be the same under the alternative hypothesis. In addition, the between dimension point estimates are the mean values of the cointergating vectors when these are indeed heterogeneous (Pedroni, 2001, p. 728).

In general, Pedroni's (2000; 2001) between dimension FMOLS estimator involves the econometric estimation of the following cointegrated system of panel data assuming only one explanatory variable:

$$y_{it} = \alpha_i + \beta_i x_{it} + v_{it} \quad \text{and} \quad x_{it} = x_{it-1} + \tilde{v}_{it} \tag{33}$$

where $\omega_{it} = \left[v_{it}\tilde{v}'_{it}\right]$ is stationary and $\Phi_i \equiv \lim_{T\to\infty} E\left[T^{-1}\left(\sum_{t=1}^{T}\omega_{it}\right)\left(\sum_{t=1}^{T}\omega'_{it}\right)\right]$ is the long-run covariance matrix estimated using Newey-West estimator. It can be decomposed as: $\Phi_i = \Phi_i^0 + \Gamma_i + \Gamma'_i$ where Φ_i^0 is the contemporaneous covariance matrix, and Γ_i is a weighted sum of estimated autocovariances. In fact, Pedroni (2000; 2001) in deriving his FMOLS estimator follows Philips and Hansen (1990) semi-parametric correction to the OLS that eliminates the second-order bias caused by endogeneity in the regressors allowing at the same time for heterogeneity in the short run dynamics and the fixed effects. Specifically, his estimator is given by:

$$\hat{\beta}^* = N^{-1}\sum_{i=1}^{N}\left[\sum_{t=1}^{T}\left(x_{it} - \overline{x}_i\right)^2\right]^{-1}\left[\sum_{t=1}^{T}\left(x_{it} - \overline{x}_i\right)y_{it}^* - T\hat{\xi}_i\right] \tag{34}$$

where $y_{it}^* = \left(y_{it} - \overline{y}_i\right) - \hat{\Phi}_{22i}^{-1}\hat{\Phi}_{21i}\Delta x_{it}$, and $\hat{\xi}_i = \hat{\Gamma}_{21i} + \hat{\Phi}_{21i}^0 - \hat{\Phi}_{22i}^{-1}\hat{\Phi}_{21i}\left(\hat{\Gamma}_{22i} + \hat{\Phi}_{22i}^0\right)$. The associated *t*-statistic for the between dimension FMOLS can be computed from:

$$t_{\hat{\beta}^*} = N^{-0.5}\sum_{i=1}^{N} t_{\hat{\beta}_{FMi}^*} \quad \text{and} \quad t_{\hat{\beta}_{FMi}^*} = \left(\hat{\beta}_{FMi}^* - \beta_0\right)\left[\hat{\Phi}_{11i}^{-1}\sum_{t=1}^{T}\left(x_{it} - \overline{x}_i\right)^2\right]^{0.5} \tag{35}$$

where $\hat{\beta}^* = N^{-1}\sum_{i=1}^{N}\hat{\beta}_{FMi}^*$ with $\hat{\beta}_{FMi}^*$ being the within dimension FMOLS estimator. The between dimension estimator above can be used also to perform formal statistical testing for the hypothesized cointegrating regression in addition to those discussed previously. Specifically, Pedroni (1999) suggests three different test statistics that are based on the between dimension FMOLS estimator and are analogous to those reported for single time series. These cointegration tests allow for considerable heterogeneity among countries in the sample, including heterogeneity in both the long-run cointegrating vectors as well as heterogeneity in the dynamics associated with short-run deviations from these cointegrating vectors. Specifically, Pedroni (1999) develops a group ρ-statistic which is analogous to Philipls-Perron *rho*-statistic based on the estimated autoregressive parameter, and both a non-parametric and parametric t-statistic analogous to the Phillips-Perron and ADF t-statistics well established in single time series literature. All these statistics are standardized by the means and variances so that they are distributed as $N(0,1)$ under the null hypothesis.[9]

4. Empirical Results

The empirical analysis presented in this section was carried out for the whole sample of developed and developing countries as well as for the two sub-groups of countries, *i.e.*, OECD and Non-OECD. We have also considered various country sub-groupings according to the geographical location (*i.e.*, Latin America, Africa, Europe, Asia) to investigate the robustness of our results but no clear pattern emerged. First Table 1 summarizes the results of both IPS t-bar and MW *Fischer* statistics for all countries in the sample as well as for the two sub-groups.[10] The lag truncations for the individual unit root regressions were allowed to vary by individual country in all three samples. The same procedure was followed also for the panel based IPS and MW tests. The standard step down procedure applied in conventional time series analysis was used to choose the lag length. This involves starting with a sufficiently large number of lags and then eliminating each time the highest order lag until one of them is statistically significant. For the initial starting value, we used the $1/5$ of the sample length. Thus, with T=15, we started with an initial lag value of 3, and then allowed the data dependent procedure to choose the actual number of fitted lags. The actual number of fitted lags varied between 0 and 3 for each country in the sample.

The statistical testing results reported in Table 1 show that all series involved in each one of the three samples considered has a unit root at the 5 per cent confidence interval. This is confirmed from both IPS *t-bar* and MW *Fisher* statistics.

Given the short time span of our data, it is more likely that our data are coming from the early stages of the business cycle in the countries in the sample. However, when the first differences of that data series is used the hypothesis of a unit root is rejected in all cases and for all variables.

Next panel cointegration results for all sub groups of countries as well as for the whole sample are reported in Table 2. In principle, the hypothesis of no cointegration is rejected for all countries and the hypothesis of one cointegrating vector is accepted. Specifically, the Levin *et al.*, (2002) test with only fixed effects as well as with both fixed and time effects included, supports the hypothesis of a cointegrating relation between GDP growth and the explanatory variables included in relation (28). On the other hand, the Maddala and Kim (1998) *Fisher* test support the presence of one cointegrating vector in all sub-groups of countries and the whole sample. Finally, Pedroni's (1999) test statistics in the context of between dimensions FMOLS estimator accept the hypothesis of a cointegrating relationship between GDP growth and all variables included in (28). This is true only for the parametric ADF *t*-statistic whereas, for the sample of Non-OECD countries both the *rho* and non-parametric Phillips-Perron *t*-statistic reject the hypothesis of cointegration.

All these tests employed so far are based on the analysis of possible cointegrated relations between GDP growth and the variables included in the empirical analysis. However, none of these tests takes into account the possibility of a cointegrating relationship that may exist between countries for any variable included in (28). If this is the case then the between dimension FMOLS estimator may yield biased results as the between transformation that is applied to the data prior to the estimation may destroy the cointegrating relationship that statistical testing confirmed previously. This may be true if for instance military spending of the sampled countries are relative to each other so that each m_{it} is cointegrated to one another. The ADF t-statistics of the demeaned variables confirms that between transformation does not render any of the series stationary.[11]

Between dimension FMOLS estimates of the cointegrating relationship for the whole sample of countries as well as for both OECD and Non-OECD countries are reported in Tables 3, 4 and 5. Regarding the whole sample estimates first, as it shown from Table 3, indeed the effect of both total public spending and military spending is exhibiting an inverse *U*-shaped pattern as both second-order parameters are negative and statistically significant at least at the 5 per cent level. Specifically, the first-order parameters were found to be 0.198 and 0.236, whereas the second-order ones -0.103 and -0.102 for total public spending and military spending, respectively. Concerning the rest two explanatory variables included in relation (28), private investment and lagged GDP, provide statistically significant estimates. Private investment influences the growth rate of real per capita GDP positively, 0.238, whereas the GDP of the previous period affect negatively the economic growth rate in countries in the sample at a lower but statistically significant rate, -0.058.

Regarding the estimates for the two sub-samples, first the between dimension FMOLS estimates for OECD countries are presented in Table 4. The results presented for the pooled sample (last row in Table 4) indicate that the share of private investment still affects positively the growth rate of GDP with a higher magnitude though, 0.298. The individual country estimates however, indicate a significant variation. The higher impact is observed for USA followed by Ireland and UK, whereas in Mexico, Greece and Poland the significance of private investment on economic growth is lessened. Finally, in Denmark, France and Turkey the relevant parameter estimates turned to statistically insignificant values. On the other hand, the one period lagged GDP is still affecting negatively the growth rate of per capita GDP. The corresponding parameter estimate for the pooled data is -0.046 but again with a greater variation across OECD countries. In USA, Norway, Mexico and Australia the corresponding parameters are not statistically significant different than zero implying that there is a structural dependent permanent boost in their growth rate. Contrary, in Ireland, Luxemburg and Poland, the GDP growth rate is highly dependent on their growth pace exhibiting the highest parameter estimates, whereas Belgium, Canada, Denmark and France exhibit the lowest impact with the corresponding parameter estimates being statistical significant around -0.020.

Again for the pooled data, the share of public investment is monotonically decreasing as the squared term turned to a statistically significant negative estimate, -0.081. Both first and second-order parameter estimates exhibit lower values compared with those obtained from the econometric estimation of the whole sample. With regard to intercountry differences, in Belgium, Canada, Finland and Ireland, the GDP share of total public spending appears to be at the optimum level as both the associated first- and the second-order parameters turned to statistically not significant values. For the remaining 13 OECD countries in the sample the presumed inverse *U*-shaped pattern emerges in all cases at a different magnitude though. Concerning military spending, in Canada, Japan and Luxemburg seems that it does not affect their respective economic growth. In all three cases the corresponding parameter estimates are not significant at the 5 per cent level. Nevertheless for the rest 14 countries the impact of military spending is monotonically decreasing at a different rate. Using these parameter estimates we have calculated the respective marginal effects of total public spending and military spending on the growth rate of GDP and the results are presented in Table 6.

First, concerning the marginal effect of the GDP share of total public spending, the estimates presented in Table 6 exhibit both negative and positive values which, however, are very close to their optimal level. Similarly, the point estimate for the whole panel of the OECD countries exhibit a value close to zero, 0.0063. This implies that public expenditures are rather close to their optimal level, besides the variation observed in individual country estimates that lie both above and below zero. Specifically, in France, Greece, Luxemburg and Norway, the marginal effect of the GDP share of total public spending is negative implying that relatively a large portion of public budget, compared to the size of these economies, is allocated to public investment. On the other hand, the highest positive values are observed in USA (0.0477), UK (0.0436), Turkey (0.0318) and Poland (0.0301) implying that public investment is below it's optimal value. Finally, in Denmark and Korea the relevant point estimates are closest to zero and thus, on their optimal level.

The same non-uniform pattern emerges also for military spending. However, the individual point estimates exhibit a greater variation than those for spending on public infrastructure. The mean estimate for the whole panel in 0.1580 indicating that the allocation of public capital in military spending is far below it's optimal level. In France, Greece, Turkey, UK and USA military spending is beyond it's optimal allocation with the highest absolute values coming from France and USA with point estimates of -0.1261 and -0.1173, respectively. For the rest of the countries the marginal effect of military spending is positive. The highest values are in Mexico (0.2502) and Norway (0.2189) indicating that these two countries are below the optimal allocation. Contrary, Australia and Belgium exhibit a very close to zero point estimate, implying that the allocation of public investment between infrastructure and military capital accumulation is growth-maximizing.

Turning now to the sample of non-OECD countries the corresponding parameter estimates are presented in Tables 5a and 5b. For the whole panel, the point estimates indicate that the lagged GDP is negatively affecting the current economic growth, whereas the relative size of private investment positively at a higher rate than the corresponding estimate for the sample of OECD countries, indicating that there is a lack of private investment in developing countries (see the last row in Table 5b). Concerning the effect of the GDP share of total public spending and the share of military spending again the non-linear relationship that emerges from the theoretical model is confirmed. The quadratic terms of both policy variables exhibit a negative statistical significant sign. Nevertheless the variation among developing countries is more intense than the sample of OECD countries.

Concerning the intertemporal dependence of economic growth on the level of economic activity, this is consistently negative in all countries in the sample. The lowest impact is observed in Brazil (-0.091), Guatemala (-0.088), Bangladesh (-0.087), South Africa (-0.079) and Paraguay (-0.078). On the other hand, the highest value is coming from Venezuela (-0.015), Egypt (-0.021), Colombia (-0.022), India (-0.031), Costa Rica (-0.031) and Peru (-0.031). The relative size of private investment domestically is higher in China (0.339), Argentina (0.301), India (0.289), Pakistan (0.276) and Philippines (0.269). Contrary in Iran, Madagascar, Ecuador, Namibia, Papua and El Salvador the corresponding parameter estimates are the lowest in the sample. Finally, in Bangladesh, Cote d'Ivoire, Dominican Republic, Guinea, Malawi, Malaysia, Trinidad-Tobago and Tunisia the respective parameter estimates turned to non statistical significant values.

With regard to the GDP share of total public expenditures, in Argentina, El Salvador, Guinea, Iran, Madagascar, Morocco, Pakistan and Uruguay both the first- and second-order parameters exhibit statistically non-significant values. The latter implies that in these countries the GDP share of public spending is rather at its optimal level. For the rest of the countries the inverse *U*-shaped pattern does not emerge only for five of those, namely, Colombia, Cote d'Ivoire, Dominican Republic, Ecuador and Trinidad-Tobago. In all these countries the quadratic term is statistically non-significant at the 5 per cent level. For the remaining twenty-five developing countries included in the sample, the non-linear relationship between the GDP share of total public spending and GDP growth emerges as both the own and the quadratic terms are statistically significant. Using these point estimates we have calculated the respective marginal effects presented in Table 7.

As these estimates reveal, in Bulgaria, China and Romania the GDP share of total public expenditures is above its optimal level as the marginal effects turned to a negative value very close, however, to zero. In all other countries the corresponding marginal effects are positive indicating under provision of public investment. The highest value is observed in Costa Rica (0.3694) followed by Peru (0.3444), Mauritania (0.3428), Bolivia (0.3159) and Malaysia (0.3073). In all these countries public investment is far below its optimal level. On the other hand, the lowest values are observed in Namibia (0.0570), Malawi (0.0818), Egypt (0.1008), Nicaragua (0.1095), Brazil (0.1278) and Kenya (0.1341).

Finally, concerning military spending in developing countries, the parameter estimates presented in Tables 5a and 5b indicate these are not affecting the economic growth in Bulgaria, Cote d'Ivoire, Dominican Republic, El Salvador, Guinea, Madagascar, Morocco, Nicaragua, Pakistan, Papua and Uruguay. The corresponding parameter estimates in these countries are all statistically non-significant than zero. On the other hand, in Bangladesh, Costa Rica, Guatemala, Malawi, Paraguay, Thailand, Trinidad-Tobago and Tunisia the non-linear relationship between military spending and GDP growth is not established as the quadratic term turned to statistically non-significant value. For the remaining nineteen Non-OECD countries, the inverse *U*-shaped pattern emerges. The corresponding marginal effects computed using the parameter estimates appearing in Tables 5a and 5b are presented also in Table 7.

The mean effect of military spending for the whole panel is almost the same with that of public expenditures, 0.2674. However, individual country estimates exhibit both negative and positive values. Specifically, in China, Chile, Colombia, India and Iran the marginal effect estimates are all negative indicating that these countries allocate more than the optimal expenditures in military spending with respect to the relative size of their public investment. However, still these values are not far from the optimal level. On the other hand, in Namibia, Mauritania, Malaysia, South Africa and Ecuador, military spending are well below their optimal level. Specifically, Namibia exhibits a point estimate of 0.2990 the highest among all countries in the sample. Contrary, Egypt, Argentina and India have achieved a better allocation of their public capital with respect to military spending as their marginal effects are very close to zero.

5. Concluding Remarks

The main goal of this paper was to study the growth implications of public capital formation empirically. We collected data on a wide range of countries to test the implications of our theoretical model about the nonlinear effects of the tax rate and the allocation of public investment on the growth rate. Our empirical investigation was limited, due to data availability, to only two types of public capital: military and infrastructure capital. The empirical results strongly support the implication of our theoretical model by indicating the existence of an inversed U shape pattern between the share of military spending and the growth rate, as well as between the share of total public investment in GDP and the growth rate.

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Table 1. Panel Unit-Root Tests.

Variable	IPS *t-bar* test statistic		MW test statistic	
	Levels	1st Diff.	Levels	1st Diff.
Whole Sample (N=55):				
GDP growth	-1.0712	-6.6072*	114.23	135.43*
Lagged GDP	-0.7153	-2.3765*	102.17	149.32*
Share of Private Investment	0.3827	-3.3784*	93.21	141.47*
Share of Total Public Spending	-0.2281	-4.7781*	97.82	155.36*
Share of Military Spending	-0.9573	-5.4429*	99.34	153.42*
OECD Countries (N=17):				
GDP growth	-0.2332	-2.3772*	41.54	51.36*
Lagged GDP	-0.0883	-2.2009*	37.98	48.92*
Share of Private Investment	1.1288	-2.4069*	36.64	55.64*
Share of Total Public Spending	-0.7159	-2.3437*	39.03	49.08*
Share of Military Spending	-1.2001	-3.3132*	36.55	53.21*
Non-OECD Countries (N=38):				
GDP growth	-1.1812	-2.3280*	90.87	108.91*
Lagged GDP	-0.7943	-1.9206*	89.76	110.23*
Share of Private Investment	0.4649	-2.4686*	92.34	107.56*
Share of Total Public Spending	-1.5138	-4.8177*	88.03	111.35*
Share of Military Spending	-0.6617	-5.1129*	90.48	109.55*

Note: * indicate rejection of the unit root hypothesis at the 5 per cent level. The tabulated critical values for the MW tests at the 5 per cent significance level are 124.34 for the whole sample, 43.77 for the sample of OECD countries and 101.87 for the sample of Non-OECD countries.

Table 2. Panel Cointegration Tests (the dependent variable is the GDP growth rate).

Levin *et al.*, (2002)		Fisher's test			Pedroni (1999) BD tests		
FE	*FTE*	$r = 0$	$r \leq 1$	$r \leq 2$	*rho*	*pp*	*adf*
Whole Sample (*N*=55):							
-11.76*	-18.00	139.6*	111.2	93.8	-1.89*	-2.04*	-2.75*
OECD Countries (*N*=17):							
-9.73*	-15.63	54.8*	40.2	35.7	-2.09*	-2.34*	-3.14*
Non-OECD Countries (*N*=38):							
-13.72*	-20.35	103.1*	95.6	81.1	-1.28	-1.48	-1.76*

Note: * indicate rejection of the null hypothesis of no cointegration at 5% significance level. FE denotes Levin *et al.*, (2002) test with only fixed effects while FTE denote the existence of both fixed and time effects. The tabulated critical values for the Fisher test at the 5 per cent significance level are 124.34 for the whole sample, 43.77 for the sample of OECD countries and 101.87 for the sample of Non-OECD countries. The critical value for Pedroni's (2000) between dimension test at the 5 per cent significance level is -1.65.

Table 3. Between Dimension FMOLS for the Whole Sample.

Variable	Estimate	t-statistic
$GDP_{(-1)}$	-0.058	(2.653)*
Share of Private Investment	0.238	(3.293)*
Share of Total Public Spending	0.198	(2.981)*
(Share of Total Public Spending)^2	-0.103	(2.334)*
Share of Military Spending	0.236	(2.134)**
(Share of Military Spending)^2	-0.102	(1.942)**

Note: * (**) indicate statistical significance at the 1 (5) per cent level. Only the panel estimates are reported in this table. The individual country-specific parameter estimates are available upon request.

Table 4. Between Dimension FMOLS Estimates for the Sample of OECD Countries.

Country	$GDP_{(-1)}$		Private Inv.		Public Spend.		$(\text{Public Spend.})^2$		Military Spend.		$(\text{Military Spend.})^2$	
	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat
Australia	-0.009	(1.082)	0.342	(2.643)**	0.034	(2.127)**	-0.032	(1.892)**	0.143	(2.624)*	-0.075	(2.352)**
Belgium	-0.022	(1.982)**	0.176	(1.873)**	0.019	(1.123)	-0.033	(0.876)	0.154	(1.998)**	-0.088	(1.813)**
Canada	-0.023	(2.623)*	0.377	(3.873)*	0.007	(0.872)	-0.008	(0.951)	0.131	(1.212)	-0.096	(1.109)
Denmark	-0.022	(1.751)**	0.093	(0.731)	0.072	(2.376)**	-0.140	(2.351)**	0.321	(2.724)*	-0.125	(2.672)*
Finland	-0.032	(2.879)*	0.197	(2.066)**	0.004	(0.153)	-0.009	(1.092)	0.193	(2.377)**	-0.085	(2.323)**
France	-0.021	(1.873)**	0.211	(1.163)	0.021	(2.314)**	-0.063	(2.253)**	0.093	(1.877)**	-0.187	(2.614)*
Greece	-0.044	(3.098)*	0.123	(1.837)**	0.033	(3.098)*	-0.151	(2.897)*	0.125	(1.798)**	-0.192	(2.091)**
Ireland	-0.076	(3.653)*	0.465	(3.652)*	0.007	(0.653)	-0.009	(0.809)	0.209	(2.841)*	-0.055	(2.104)**
Japan	-0.035	(4.076)*	0.287	(2.153)**	0.042	(2.125)**	-0.047	(1.899)**	0.062	(0.745)	-0.032	(0.653)
Korea	-0.028	(2.351)**	0.212	(3.231)*	0.015	(1.983)**	-0.033	(1.820)**	0.203	(2.809)*	-0.143	(2.612)*
Luxemburg	-0.064	(2.752)*	0.123	(1.034)	0.046	(2.231)**	-0.129	(2.213)**	0.091	(1.212)	-0.021	(0.834)
Mexico	-0.017	(1.531)	0.098	(1.829)**	0.032	(2.724)*	-0.078	(2.576)*	0.432	(3.651)*	-0.198	(3.124)*
Norway	-0.012	(0.763)	0.164	(2.580)*	0.036	(1.893)**	-0.116	(2.341)**	0.321	(2.746)*	-0.113	(2.578)*
Poland	-0.056	(3.423)*	0.103	(2.203)**	0.066	(2.673)*	-0.117	(3.133)*	0.194	(2.585)*	-0.064	(2.314)**
Turkey	-0.032	(2.063)**	0.132	(1.429)	0.054	(2.109)**	-0.122	(2.783)*	0.078	(2.087)**	-0.135	(1.893)**
UK	-0.041	(4.109)*	0.421	(3.651) *	0.061	(3.764)*	-0.048	(2.093)**	0.089	(1.883)**	-0.176	(2.798)*
USA	-0.007	(1.112)	0.541	(2.963) *	0.053	(3.093)*	-0.021	(1.793)**	0.072	(1.799)**	-0.184	(2.913)*
Panel	-0.046	(2.095)**	0.298	(3.542) *	0.033	(1.894)**	-0.081	(2.341)**	0.158	(2.764)*	-0.109	(2.412)

Note: * (**) indicate statistical significance at the 1 (5) per cent level

Table 5a. Between Dimension FMOLS Estimates for the Sample of Non-OECD Countries.

Country	$GDP_{(-1)}$		Private Inv.		Public Spend.		$(\text{Public Spend.})^2$		Military Spend.		$(\text{Military Spend.})^2$	
	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat
Argentina	-0.072	(3.143)**	0.301	(2.913)**	0.142	(1.435)	-0.056	(0.757)	0.104	(1.785)**	-0.214	(2.368)**
Bangladesh	-0.087	(3.672)*	0.076	(0.672)	0.223	(2.341)**	-0.067	(0.975)	0.215	(2.058)**	0.076	(1.021)
Bolivia	-0.034	(2.098)**	0.254	(3.287)*	0.341	(2.784)*	-0.105	(1.851)**	0.236	(2.341)**	-0.174	(1.985)**
Brazil	-0.091	(4.762)*	0.231	(3.164)*	0.298	(2.590)*	-0.368	(3.247)*	0.198	(1.974)**	-0.325	(3.258)*
Bulgaria	-0.063	(2.965)*	0.176	(2.872)*	0.142	(2.243)**	-0.524	(3.574)*	0.085	(1.102)	-0.107	(1.485)
Chile	-0.032	(2.134)**	0.209	(2.590)*	0.304	(3.222)*	-0.465	(2.412)**	0.041	(1.747)**	-0.164	(1.685)**
China	-0.041	(2.451)*	0.339	(2.098)**	0.085	(1.982)**	-0.662	(4.639)*	0.187	(2.132)**	-0.324	(2.365)**
Colombia	-0.022	(1.097)	0.183	(1.980)**	0.365	(2.748)*	-0.084	(0.936)	0.121	(2.025)**	-0.341	(2.478)*
Costa Rica	-0.031	(1.341)	0.213	(2.542)*	0.412	(3.124)*	-0.158	(1.968)**	0.365	(3.141)*	-0.089	(1.239)
Cote d'Ivoire	-0.036	(1.513)	0.097	(1.364)	0.213	(1.898)**	-0.056	(1.057)	0.128	(1.611)	-0.033	(0.857)
Dominican R.	-0.065	(2.562)*	0.076	(1.092)	0.189	(1.754)**	-0.105	(1.321)	0.087	(0.932)	-0.017	(0.635)
Ecuador	-0.071	(2.773)*	0.145	(2.671)*	0.431	(2.863)*	-0.032	(0.627)	0.111	(1.854)**	-0.147	(1.978)**
Egypt	-0.021	(0.791)	0.175	(2.154)**	0.142	(2.173)*	-0.214	(2.236)**	0.117	(1.814)**	-0.365	(2.698)*
El Salvador	-0.076	(3.098)*	0.160	(1.963)**	0.134	(1.482)	-0.039	(1.027)	0.086	(1.041)	-0.034	(0.748)
Guatemala	-0.088	(2.960)*	0.219	(2.672)*	0.312	(2.767)*	-0.354	(2.685)*	0.305	(3.321)*	-0.117	(1.236)
Guinea	-0.045	(2.341)**	0.118	(1.451)	0.152	(1.376)	-0.047	(1.385)	0.046	(0.754)	-0.014	(0.658)
India	-0.031	(1.451)	0.289	(2.352)**	0.331	(2.877)*	-0.426	(3.287)*	0.124	(2.285)**	-0.321	(2.323)**
Iran	-0.069	(2.673)*	0.102	(2.009)**	0.114	(1.092)	-0.012	(0.365)	0.163	(2.169)**	-0.458	(3.852)*
Kenya	-0.033	(2.125)**	0.176	(2.608)**	0.212	(2.064)**	-0.325	(2.933)*	0.136	(1.694)**	-0.214	(2.478)*
Madagascar	-0.076	(2.798)*	0.112	(1.907)**	0.109	(0.982)	-0.053	(1.058)	0.041	(0.789)	-0.009	(0.369)

Note: * (**) indicate statistical significance at the 1 (5) per cent level

Table 5b. Between Dimension FMOLS Estimates for the Sample of Non-OECD Countries (continued).

Country	$GDP_{(-1)}$		Private Inv.		Public Spend.		(Public Spend.)^2		Military Spend.		(Military Spend.)^2	
	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat	Estimate	*t*-stat
Malawi	-0.056	(2.533)*	0.121	(1.563)	0.223	(2.231)**	-0.352	(2.965)*	0.168	(1.874)**	-0.014	(0.298)
Malaysia	-0.059	(2.498)*	0.153	(1.609)	0.423	(3.129)*	-0.487	(3.875)*	0.325	(2.635)*	-0.174	(1.705)**
Mauritania	-0.061	(2.781)*	0.176	(1.892)**	0.531	(3.672)*	-0.667	(4.148)*	0.410	(3.985)*	-0.324	(2.652)*
Morocco	-0.043	(2.124)**	0.231	(2.761)*	0.112	(1.176)	-0.022	(0.842)	0.082	(0.285)	-0.104	(1.229)
Namibia	-0.037	(1.563)	0.124	(2.341)**	0.312	(2.898)*	-0.445	(2.421)**	0.321	(2.638)*	-0.121	(1.675)**
Nicaragua	-0.069	(2.873)*	0.251	(2.981)*	0.213	(2.421)**	-0.321	(2.796)*	0.039	(0.365)	-0.089	(1.109)
Pakistan	-0.034	(1.718)**	0.276	(3.082)*	0.112	(0.917)	-0.077	(1.328)	0.074	(0.795)	-0.017	(0.751)
Papua	-0.066	(2.901)*	0.121	(1.913)**	0.309	(2.231)**	-0.412	(3.596)*	0.125	(1.174)	-0.008	(0.228)
Paraguay	-0.078	(3.092)*	0.208	(2.123)**	0.341	(2.314)**	-0.502	(3.154)*	0.312	(2.985)*	-0.104	(1.365)
Peru	-0.031	(1.321)	0.217	(2.542)*	0.410	(3.316)*	-0.368	(3.754)*	0.236	(2.457)*	-0.177	(2.074)**
Philippines	-0.042	(1.982)**	0.269	(2.823)*	0.275	(2.773)*	-0.326	(3.165)*	0.147	(1.977)**	-0.214	(2.298)**
Romania	-0.061	(2.314)**	0.234	(3.212)*	0.074	(1.715)**	-0.463	(3.253)*	0.214	(2.074)**	-0.365	(2.852)*
South Africa	-0.079	(3.245)*	0.179	(2.124)**	0.412	(3.129)*	-0.524	(4.875)*	0.325	(2.695)*	-0.415	(3.625)*
Thailand	-0.038	(1.453)	0.198	(2.322)**	0.312	(2.902)*	-0.487	(4.632)*	0.128	(2.014)**	-0.027	(0.852)
Trinidad	-0.044	(2.130)**	0.089	(1.331)	0.289	(2.417)**	-0.217	(1.258)	0.225	(2.852)*	-0.063	(0.985)
Tunisia	-0.060	(2.983)*	0.161	(1.412)	0.308	(3.325)*	-0.372	(2.985)*	0.207	(3.285)*	-0.038	(0.795)
Uruguay	-0.071	(3.415)*	0.192	(2.613)*	0.123	(1.143)	-0.055	(1.089)	0.085	(1.188)	-0.014	(0.698)
Venezuela	-0.015	(0.679)	0.214	(2.761)*	0.074	(1.742)**	-0.645	(3.822)*	0.241	(2.365)**	-0.342	(2.985)*
Panel	-0.061	(2.761)*	0.334	(2.871)*	0.337	(2.981)*	-0.285	(2.857)*	0.298	(2.385)**	-0.098	(1.984)**

Note: * (**) indicate statistical significance at the 1 (5) per cent level

Table 6. Marginal Effects of Military Spending and Total Public Spending on GDP Growth for the Sample of OECD Countries.

Country	Total Public Spending	Military Spending
Australia	0.0234	0.0382
Belgium	0	0.0321
Canada	0	0
Denmark	0.0034	0.1570
Finland	0	0.0876
France	-0.0067	-0.1261
Greece	-0.0063	-0.0886
Ireland	0	0.1498
Japan	0.0292	0
Korea	0.0093	0.0625
Luxemburg	-0.0001	0
Mexico	0.0172	0.2502
Norway	-0.0104	0.2189
Poland	0.0301	0.1352
Turkey	0.0318	-0.0525
UK	0.0436	-0.0874
USA	0.0477	-0.1173
Panel	0.0063	0.1580

Table 7. Marginal Effects of Military Spending and Total Public Spending on GDP Growth for the Sample of Non-OECD Countries.

Country	Total Public Spending	Military Spending	Country	Total Public Spending	Military Spending
Argentina	0	0.0286	Madagascar	0	0
Bangladesh	0.2230	0.2150	Malawi	0.0818	0.1680
Bolivia	0.3159	0.1804	Malaysia	0.3073	0.2705
Brazil	0.1278	0.0575	Mauritania	0.3428	0.2768
Bulgaria	-0.0273	0	Morocco	0	0
Chile	0.2216	-0.0319	Namibia	0.0570	0.2990
China	-0.0279	-0.0878	Nicaragua	0.1095	0
Colombia	0.3650	-0.0606	Pakistan	0	0
Costa Rica	0.3694	0.3650	Papua	0.1436	0
Cote d'Ivoire	0.2130	0	Paraguay	0.2763	0.3120
Dominican Rep.	0.1890	0	Peru	0.3444	0.1750
Ecuador	0.4310	0.2045	Philippines	0.2094	0.0872
Egypt	0.1008	0.0158	Romania	-0.0374	0.1314
El Salvador	0	0	South Africa	0.2263	0.2281
Guatemala	0.2778	0.3050	Thailand	0.2215	0.1280
Guinea-Bissau	0	0	Trinidad	0.2890	0.2250
India	0.2509	-0.0081	Tunisia	0.1932	0.2070
Iran	0	-0.0504	Uruguay	0	0
Kenya	0.1341	0.0854	Venezuela	-0.0138	0.0732
Panel	0.2696	0.2674			

Endnotes

[1] A more comprehensive review of the relevant studies can be found in Adams *et al.*, (1991), Sandler and Hartley (1995) and Ram (1995).

[2] The inclusion of adjustment costs is useful because it allows the presence of transitional dynamics when comparative statics are utilised to analyse the dynamic behaviour of the model. Notice that we will not include adjustment costs for public capital formation, as this would only complicate the solution of the model without adding further insights.

[3] The OECD countries included in the sample were: Australia, Belgium, Canada, Denmark, Finland, France, Greece, Ireland, Japan, Korea, Luxemburg, Mexico, Norway, Poland, Turkey, UK and USA. Accordingly the sample of non-OECD includes: Argentina, Bangladesh, Bolivia, Brazil, Bulgaria, Chile, China, Colombia, Costa Rica, Cote d'Iviore, Dominican Rep., Ecuador, Egypt, El Salvador, Guatemala, Guinea-Bissau, India, Iran, Kenya, Madagascar, Malawi, Malaysia, Mauritania, Morocco, Namibia, Nicaragua, Pakistan, Papua, Paraguay, Peru, Philippines, Romania, South Africa, Thailand, Trinidad-Tobago, Tunisia, Uruguay and Venezuela.

[4] The inconsistency of pooled estimators in both static and dynamic heterogeneous panel models has been demonstrated by Pesaran and Smith (1995) and Pesaran *et al.*, (1996).

[5] Pierse and Shell (1995) argued that the explanatory power of individual unit-root tests is limited when the time span of the data is short.

[6] Panel based unit root tests have been initially suggested by Quah (1992; 1994). However the tests suggested by Quah do not accommodate heterogeneity across groups which is most likely the case with aggregate country data. On the other hand, the unit root test developed by Levin *et al.*, (2002) and Breitung and Meyer (1994) allows for heterogeneity only in the constant term in the ADF regression equation.

[7] Im *et al.*, (2003) showed that the standardized *t-bar* statistic converges to the standard normal distribution as both N and T tend to infinity.

[8] In particular, the MW Pearson-lambda statistic provides insights whether the results for the panel are generally being driven by the strength of one or two outlier countries or whether it is a general tendency of all countries in the panel.

[9] See Pedroni (1999) for more details and the exact formulas for calculating the test statistics.

[10] In order to conserve space the individual ADF tests for each country in the sample are not reported but are available upon request.

[11] The corresponding test statistics are not reported here but are available upon request.