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The Effect of Fiscal Variables on Economic Growth in Asian Economies: a Dynamic Panel Data Analysis

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Abstract

This paper investigates the effect of relationship between fiscal variables and economic growth in Asian economies using a generalized method of moments (GMM) method as a dynamic panel data analysis over the 1985-2001 periods. These data contain a number of time invariant and time varying variables, where the time varying variables are averaged over four year It examines two different channels through which fiscal policy can affect long run economic growth in Asia. The first channel is when components and aggregate government expenditure affects the real per capita GDP and the second channel is when the components and aggregate of other fiscal variables affects the real per capita GDP. The dynamic panel data result especially GMM-SYS establishes a long run relationship between fiscal policy and economic growth. We find positive and statistically significant impact of health and education expenditure, aggregate of government expenditure and aggregate of other fiscal variables on real per capita GDP. Furthermore, we find that the defence expenditure, distortionary taxation and budget balance are significantly and negatively related to real per capita GDP.

Keywords: Economic growth, Fiscal variables, OLS, Within group, GMM-DIFF, GMM-SYS

1. Introduction

Fiscal policies have a benign role for economic growth in the Asian region, namely to provide a stable macro environment for investment. The changed environment of liquidity constraints on external borrowing and slowdown in output growth has led to new attention being directed towards the role and contribution of fiscal policies in reviving growth in the region (Gangopadhyay and Chatterji, 2005). In the debate on economic policy, fiscal policy is predominantly viewed as an instrument to mitigate short-run fluctuations of output and employment. By a varying government spending or taxation, fiscal policy aims at altering aggregate demand in order to move the economy closer to potential output.

Fiscal policy was neither a cause of the crisis nor a critical determinant of economic growth. Nevertheless, its role in both the pre-crisis and post-crisis period in Asian countries has been seen as crucial, primarily in terms of its contribution to economic growth. From the view of the perspective of contemporary debate in the pre-crisis decade, policy concerns focused on the perceived overheating of Asian economies rather than concerns with fiscal and external sustainability.

There are large differences among the Asian countries in their levels of living and other circumstances, as well as the policies that they have pursued. Larger government size is likely to be an obstacle to efficiency and economic growth because the taxes necessary to support government expenditures distort incentives to work and to invest, absorb funds that otherwise would have been used by the private sector in profitable investment opportunities, generally reduce efficient resource allocation, and hence reduce the level of output. In addition, government operations are often carried out inefficiently, and the regulatory process imposes excessive burdens and costs on the economic system. Thus, countries with greater government expenditure as a proportion of output should experience lower economic growth. These arguments, together with the debt crises experienced have led many countries to start a mass deregulation of market and privatization of public enterprises. Based on the above argument, and as we mentioned earlier, Keynesian economics predicts government expenditure should lead to economic growth.

When looking at the growth performance in the Asian countries in recent decades, two observations are noteworthy. First, growth has declined and become stagnant significantly since 1985. Secondly, government expenditure is not inhibits the full exploitation of the growth potential of Asian economies. There is a broad consensus that these developments in fiscal policies contribute to the relatively weak growth performance in the Asian countries.

Fiscal positions vary significantly across countries and sub regions. Significant fiscal deficit and accumulation of public debt are relatively new phenomena for most Asian economies. However, expenditure growth outpaced revenue growth in many Asian economies, leading to persistent budget deficits and high indebtedness. Weak fiscal positions have left little room for further fiscal expansion in most Asian economies when faced by economic slowdown. Moreover, measuring fiscal policy has always posed a difficult challenge.

The objective of this study is to examine the effect of fiscal variables on economic growth in Asian economies. Thus, this study aims at filling a gap in research devoted solely to investigating the effect and relationship between fiscal policy and economic growth using of newly developed methods of dynamic panel data by Arellano and Bond (1991) and Blundell and Bond (1998).

This paper is organized as follows. Section 2 contains a brief literature review. In section 3, the model is applied to the thirteen Asian economies. Section 4 presents empirical results and the last section concludes.

2. Review of related literature

The most recent empirical literature, mainly based on panel data regressions, show that economic growth is significantly affected by fiscal policies, although there remains some lack of agreement on the sign of the effects. Economic theory suggests that fiscal multipliers are more likely to be positive when economies are relatively closed, government debt is low and fiscal expansion focuses on spending, there is considerable slack in capacity of productive. There is also some evidence of negative fiscal multipliers which is no clear consensus on the precondition for such an outcome. Gerson (1998) surveyed the theoretical and empirical literature on the effect of fiscal policy variables (government expenditure program and taxes) on economic growth. He concluded that educational attainment and public health status had significant, positive effects on per capita output growth; economies that were open to international trade grew faster than those that were closed, therefore fiscal policies that encouraged openness should encourage growth. Caselli *et al.* (1996) found robust positive contribution of the government expenditure ratio (net of defence and educational expenditure) to growth. Kneller *et al.* (1999) found that public expenditure and taxation only affected growth if they were productive and distortionary, respectively; productive government expenditure was found to positively affect growth, whereas distortionary taxation was found to be harmful for growth. With this distinction they argued that both sides of the government budget should be considered in estimating the impact of fiscal policy on growth, as their financing offset the growth-enhancing effects of productive expenditure.

Studying the relationship between government expenditure and economic growth is becoming of crucial importance to divide government activities in several categories and methodologies. Zagler and Dürnecker (2003) surveyed the literature on fiscal policy and economic growth. They presented a unifying framework for the analysis of long run growth implications of government expenditures and revenues. They found that the level of education expenditure and the growth rate of public infrastructure investment both exhibited a positive impact on the growth rate of the economy. Tanzi and Zee (1997) examined systematically the various ways that the main fiscal instruments (tax policy, public expenditure policy, budget policy) influenced economic growth through their impact on the determinants of growth. Yasin (2003) studied the relationship between government expenditure and economic growth. His studies re-examined the effect of government spending on economic growth using panel data set from Sub-Saharan Africa. The results from both estimation techniques indicated that government spending, trade-openness, and private investment spending all had positive and significant effect on economic growth. Biswas and Ram (1986) used data from 55 countries over the period 1960-1977 and found that defence expenditure has no significant effect on output growth. Abu-Bader and Abu-Qarm

(2003) used multivariate cointegration and variance decomposition techniques to investigate the causal relationship between government expenditure and economic growth. Cross section growth regressions had been used to assess the relationship between defence expenditure and economic growth. They found that when considering overall government expenditure, there was bidirectional causality between government spending and economic growth with a negative long run relationship in the cases of Israel and Syria, and a unidirectional negative short-run causality from economic growth to government spending in the case of Egypt.

Landau (1986) examined the possibility that the impact of defence expenditure on output growth was nonlinear, with relatively low levels of defence expenditure enhancing output growth, but relatively low levels of defence expenditure inhibiting growth. He found that this was in fact the case, with a positive relationship between defence expenditure and output growth holding until defence expenditure reached about 4 percent of GDP and a negative relationship taking over at about 9 percent of GDP. For sub-samples restricted to Latin America and Africa, he found a significant, positive relationship between defence expenditure and the share of government education and health expenditure in GDP.

Hassan *et al.* (2003) stated that there were essentially four arguments showing military expenditure retarding economic growth. First, higher defence expenditures could crowd out both public and private investment that might be more growth-oriented and need-based than those of defence spending. This crowding out of essential investment might have an adverse impact on the long-run economic growth. Second, defense expenditure can cause balance of payment problems if hard-earned foreign exchanges were used to purchase arms and defence hardware. Third, defence might inhibit growth by diverting resources from the export sector, which was often considered an engine of growth. Finally, the defence sector limited growth through inefficient bureaucracy and excess burdens created by taxes necessary to finance military spending. Since defence spending could cause both positive and negative effects, its final impact on growth would depend on the strength of the opposing forces.

Devarajan *et al.* (1996) investigated the relationship between the compositions of public expenditure and economic growth. Using a simple, analytical model, they derived conditions under which a change in the mix of public spending could lead to a higher steady-state growth rate for the economy. Based on the model, their empirical results suggested that expenditures that were normally considered productive could become unproductive if there was an excessive amount of them. Glomm and Ravikumar (1994) considered the relationship between government expenditure on infrastructure or education and economic growth, and the implication of their models' yield depended on how the expenditures were being conceived and how they looked at the effects of taxes that had to be raised to finance the expenditure. Therefore the general implications that seem to follow from these models are that one expects partially a positive correlation of growth with productive expenditure (e.g infrastructure and education) and partially a negative correlation with government consumption and distortionary taxes.

Abdullah *et al.* (2008) used the Pedroni Cointegration method to establish a long run relationship between fiscal policy and economic growth. They found a positive and statistically significant impact of health and education expenditure, aggregate of government expenditure and aggregate of fiscal policy on real per capita GDP. They also found that the defence expenditure, distortionary taxation and budget balance are significantly and negatively related to real per capita GDP. Barro and Sala-i-Martin (1995) found that government expenditure in education, health, and other services could contribute indirectly towards raising the marginal productivity of private sectors via their contribution on human capital accumulation. Chen and Gupta (2006) examine the government expenditure in health and education and other structural factors that may have an effect on economic growth. They apply the GMM estimation technique which is the set explanatory variables included in the growth regression specification are based on the endogenous growth theory and can all be considered to be important determinants of economic growth. The results show that the coefficient on government expenditure in health and education is negative but is small in absolute value.

3. Empirical model

As follow Hoeffler (2002) and Abdullah *et al* (2008), in the Solow model (1956) in which growth in output per worker depends on initial output per worker [y(0)], the initial level of technology [A(0)], the rate of technological progress (g), the savings rate (S), the growth rate of the labour force (n), the depreciation rate (δ), and the share of capital in output (α). Thus, the model predicts that a high saving rate will affect growth in output per worker positively, whereas high labour force growth (corrected by the rate of technological progress and the rate of depreciation) will have a negative effect on growth in output per worker. The basic Solow model is

$$\ln y_t - \ln y_0 = \ln y_0 + \ln A_0 + g_t + (\alpha/1-\alpha) \ln S - (\alpha/1-\alpha)\ln(n+g+\delta) \quad (1)$$

where y_t denotes the logarithm of output per worker in period t.

In the augmented version of the Solow model investment in human capital is an additional determinant of growth in output per worker

$$\ln y_t - \ln y_0 = \ln y_0 + \ln A_0 + g_t + (\alpha/1\text{-}\alpha\text{-}\beta)\ln S_k + (\beta/1\text{-}\alpha\text{-}\beta)\ln S_h - (\alpha/1\text{-}\alpha)\ln(n+g+\delta) \quad (2)$$

where S_k and S_h denote the proportion of output invested in physical and human capital, respectively.

Equations (1) and (2) have for example been used as the framework for empirical analysis by Mankiw *et al*. (1992), Islam (1995) and Caselli (1996). In this section, a simple model is set out that provides an organizing framework for thinking about the ways in which the elements and components of fiscal variables affect growth. Therefore, we adopt the framework introduced by Mankiw *et al.* (1992), Demetriades and Law (2006), Ghura and Hadjimichael (1996), Hoeffler (2002), and Knight *et al.* (1993). This study provides a growth model from the conventional growth accounting framework and the production function below takes the standard neoclassical form with a minor modification which includes human capital in the Cobb-Douglas production function:

$$Y_t = K_t^{\alpha} H_t^{\beta}(A_tL_t)^{1-\alpha-\beta},\ 0 < \alpha < 1. \quad (3)$$

where Y is real output at time t, K and L are the stocks of physical capital and labour, respectively, at time t, H is the stock of human capital, A is a similar measure for physical capital, and α and β the share of capital and human capital on output. A is a labour-augmenting factor reflecting the level of technological development and efficiency in the economy and the subscript t indicates time. This equation states merely that at any moment, the total output of the economy depends on the quantity and quality of physical capital employed, the quantity of labour employed, and the average level of skills of the labour force. Output can only increase if K, L, A, or H also increases, and perpetual increases in output per worker can only occur if the stock of capital per worker or the average quality of labour or of capital also increases perpetually.

We assume that $\alpha + \beta < 1$ which implies that there are decreasing returns to all capital raw labour and labour-augmenting technologies are assumed to grow according to the following functions:

$$L_t = L_0e^{nt} \quad (4)$$

$$A_t = A_0^{gt+P\theta} \quad (5)$$

where n is the exogenous rate of growth of the labour force, g is the exogenous rate of technological progress, P is variables of vector of government expenditure and fiscal policy that can affect the level of technology and efficiency in the economy, and θ is a vector of coefficient related to these variables.

Demetriades and Law (2006) state that variable A depends on exogenous technological improvements and the level of other variables. Variable A in this study is differing from A used by Mankiw *et al*. (1992). This modification is more likely to be particularly relevant to the empirical cases of the link between government expenditure, fiscal policy and economic growth. The technological improvements are encouraged by development in public investment spending and fiscal policy which tend to contribute to economic growth (Ramirez and Nazmi, 2003).

In the steady state, output per worker grows at the constant rate g, which is the exogenous component of the growth ret of the efficiency variable A (Demetriades and Law, 2006). Hence, this output can be obtained directly from the definition of output per effective worker as follows:

$$Y_t/A_tL_t = (k_t)^{\alpha}(h_t)^{\beta} \quad (6)$$

$$Y_t/L_t = A_t(k_t)^{\alpha}(h_t)^{\beta}$$

Let $y_t^* = (Y_t/L_t)^*$

Taking logs both of Equation (6) and log income per worker at a given time; time 0 for simplicity is

$$\ln(Y/L)^* = \ln A + \alpha\ln k^* + \beta\ln h^* \quad (7)$$

Where $A_t = A_0e^{gt-P\theta}$

The main equation of this model used for estimation purposes in equation (8) below:

$$\ln(Y/L)^* + \ln A_0 + g_t + (\alpha/1\text{-}\alpha\text{-}\beta)\ln S_k + (\beta/1\text{-}\alpha\text{-}\beta)\ln S_h - (\alpha+\beta/1\text{-}\alpha\text{-}\beta)\ln(n+g+\delta) \quad (8)$$

Equations (8) indicate steady state output per worker or labour productivity where a vector of fiscal variables proxies exist, while S_k is the savings in physical capital, S_h is the savings in human capital, and δ is the rate of depreciation.

Before proceeding to estimate the model, it is necessary to write equation (8) in terms of per capita output. Note again that for Mankiw *et al.* (1992):

$$\ln A_0 = \alpha + \varepsilon \quad (9)$$

On the other hand, for Islam (1995) and Caselli *et al* (1996):

$$\ln A_t = \ln A_0 + g_t \quad (10)$$

Out model differs from Caselli *et al* (1996) and Islam (1995) where we assumes that S_h and gt do not vary over time but S_k and n can be assumed to vary over time. This means that $\ln A_0$, gt and S_h can be considered as a constant term A_0. Therefore, the steady-state output per worker or labour productivity (y^*) grows according to the following equation:

$$\ln(Y/L)^* = A_0 + \theta \ln P + (\alpha/1-\alpha-\beta) \ln S_k + - (\alpha+\beta/1-\alpha-\beta)\ln(n+g+\delta) \quad (11)$$

The above equation introduces a set of variables (P) which is assumed as exogenous that could affect economic growth in the long run. With the introduction of endogenous growth theory, P is no longer assumed as exogenous. The endogenous treatment of P allows us to suggest a possible set of explanatory variables. This model differs from neoclassical production functions in two important categories of variables namely technology related variables and policy related variables. The key assumption about productivity growth here is that a typical developing county purchases technology knowledge abroad from various suppliers. What technology will be purchased depends on the price of foreign technology as well as trade and exchanged rate policies that impact the final cost of the imported technology (Ramirez and Nazmi, 2003). In our model, we concentrate on policy related variables and we introduce fiscal variables as a proxy for policy related variables.

As with studies of the impact of health, education and defence expenditure on economic growth, some dispersion of results is a natural outcome of differences in data sets and specifications. Therefore we proposed empirical Model 1 is as follows for the effect of components of government expenditure on economic growth:

$$\ln Y_{it} = \beta_0 + \beta_1 \ln Y_{it-1} + \beta_2 \ln he_{it} + \beta_3 \ln ee_{it} + \beta_4 \ln de_{it} + \beta_5 \ln OFV_{it} + \beta_6 \ln S_{kit} - \beta_7 \ln(n+g+\delta)_{it} + \varepsilon_{it} \quad (12)$$

where Y_{it} is real GDP per capita, Y_{it-1} is the initial level of per capita GDP, he_{it} is a government expenditure on health to GDP, ee_{it} represents government expenditure on education to GDP, de_{it} is a government expenditure on defence to GDP, OFV_{it} is an aggregate of other fiscal variables as a share of GDP (obtained by summing up public sector wages and salaries, expenditure on other goods and services, transfers and subsidies, interest payment on government debt, capital expenditure (minus government expenditure on health, education and defence), tax revenues, nontax revenue, and grant), S_{kit} is the savings in physical capital, $(n+g+\delta)$:n is the rate of labour growth, g is the rate of technology growth or technological progress and δ is the rate of depreciation. The addition of g and δ is assumed to be constant across countries and over time, following Islam (1995), Mankiw *et al*. (1992) and Caselli *et al* (1996), technological progress and the depreciation rate were assumed to be constant across countries and that they sum up to 0.05. The natural logarithm of the sum of population growth and 0.05 was calculated for $\ln(n+g+\delta)$, i is a cross-section data for countries referred to, and t is a time series data, ε_{it} is an error term. The constant is denoted β_0 while $\beta_1 - \beta_7$ are the coefficients showing how much a one unit increase in each individual variable will affect the growth rate in economic growth.

We also proposed empirical Model 2 for the effect of ditortionary taxation and budget balance on economic growth as follows:

$$\ln Y_{it} = \beta_0 + \beta_1 \ln Y_{it-1} + \beta_2 \ln dt_{it} + \beta_3 \ln bb_{it} + \beta_4 \ln GE_{it} + \beta_5 \ln S_{kit} - \beta_6 \ln(n+g+\delta)_{it} + \varepsilon_{it} \quad (13)$$

where Y_{it} is real GDP per capita, Y_{it-1} is the initial level of per capita GDP, dt_{it} is a distortionary taxation as a share of GDP (obtained by taxes on income and profit + social contribution + taxes on payroll and + taxes on property), bb_{it} represents budget balance as a share of GDP [obtained by (tax revenue + nontax revenue + grants) – (current expenditure + capital expenditure (minus government expenditure on health, education and defence)], GE_{it} is an aggregate of independent government expenditure variables as a share of GDP (obtained by summing up the government expenditure on health, education, and defence), S_{kit} and $(n+g+\delta)_{it}$ are as defined earlier in Equation (12), i is a cross-section data for countries referred to, and t is a time series data, ε_{it} is an error term. The constant is denoted as β_0 while $\beta_1 - \beta_6$ are the coefficients showing how much a one unit increase in each individual variable will affect the growth rate in economic growth.

3.1 Dynamic panel data estimation procedure

Follow Hoeffler (2002), in a panel data model we can then explicitly account for permanent unobserved country specific effects, η_i. This provides a panel data model of form

$$y_{it} = \alpha + \beta y_{it-1} + \gamma x_{it} + \eta_{it} + \nu_{it} \tag{14}$$

where y_{it} denotes the growth rate of real per capita GDP, y_{it-1} is the initial level of real per capita GDP, t denotes points in time $t = 2, \ldots T$. For example y_{it} may reflect the average growth rate over a series of four year period, with y_{it-1} being the level of real per capita GDP at the beginning of these periods, and x_{it} being measured either at the beginning of each period, or as an average over each of the four year periods.

Since the relevant four year growth rate in Equation (14) is logarithmic difference in GDP per capita, we have the following dynamic panel data model

$$y_{it} - y_{it-1} = \alpha + \beta y_{it-1} + \gamma x_i + \eta_{it} + \nu_{it} \tag{15}$$

or equivalently

$$y_{it} = \alpha + \beta^* yit_{-1} + \gamma x_{it} + \eta_i + \nu_{it} \tag{16}$$

where $\beta^* = (\beta + 1)$. It is important to note that the typical panel in the study of economic growth then has a small number of time series period, i.e. T is small. Asymptotic results discussed below focus on the case where the number of countries N becomes large for fixed T.

3.1.1 OLS levels estimation and within groups estimation

In this section also we adopts the autoregressive model from Bonds (2002), compromising the estimation of the following equation:

$$Y_{it} = \alpha y_{it} + (\eta_i + \nu_{it});\ |\alpha| < 1;\ I = 1,2,\ldots, N;\ t = 2.3\ldots\ldots\ldots,T \tag{17}$$

where y_{it} is an observation on some series for individual i in period t, y_{it-1} is the observation on the same series for the same individual in the previous period, η_i is an unobserved individual-specific time-invariant effect which allows for heterogeneity in the means of the y_{it} series across individuals, and v_{it} is a disturbance term. A key assumption we maintain throughout is that the disturbances v_{it} are independent across individuals.

As Hsio (1986) shows, omitting unobserved time invariant country effects in a dynamic panel model will cause OLS levels estimate to be biased and inconsistent. The lagged dependent variable, y_{it-1}, is positively correlated with the permanent effects, η_i. As a result the OLS levels estimate of coefficient β^* in the typical regression is likely to be biased upward (Hoeffler, 2002).

Alternatively, the Within Groups estimator differences all variables from their respective time means. It is assumed that all right hand side variables are strictly exogenous, which is violated at least by the lagged variable. As such, it introduces a significant correlation between non-exogenous variables and the time-demeaned error term (Bond, 2002), which decreases as the number of periods tends towards infinity (Baltagi, 1995). For this method model in Equation (17) is transformed by subtracting out the time series means of each variable for each country. The within groups estimator eliminates this source of inconsistency by transforming the equation to eliminate η_i. Specifically the mean values of y_{it}, y_{it-1}, η_i and v_{it} across T–1 observation for each individual i are obtained, and the original observations are expressed as deviations from these individual means. Within groups estimator is also inconsistent and standard results for omitted variables bias indicate that, at least in large samples, the within groups estimate $\beta^*$is likely to be biased downward (Bond, 2002). Consequently, the estimate β^* obtained from OLS levels can be regarded as an approximate upper bound on this coefficient, and the estimate obtained from within groups estimation can be regarded as an approximate lower bound (Hoeffler, 2002).

3.1.2 GMM estimators for dynamic panel data model

The GMM estimator proposed by Arellano and Bond (1991), known as two-step estimation, is constructed in two phases. Firstly, first differences regressions and/or instruments from the dynamic panel data model are calculated to control for unobserved effects; then, second, using lagged observations values of right-hand side explanatory variables in levels as their instruments. With lagged dependent variable and other endogenous regressors, the lagged levels are dated t-2 and earlier. If there are predetermined regressors, all their lagged levels are used as instruments.

Consider the panel data model specification

$$y_{it} - y_{i,t-1} = (\beta^* - 1) y_{i,t-1} + \gamma x_{it} + u_{it} \tag{18}$$

$$u_{it} = \eta_i + \nu_{it} \tag{19}$$

where $y_{i,t}$ is the logarithm of dependent variables, $x_{i,t}$ is the set of other endogenous variable, u_{it} is the time-specific effects, η_{it} is the country-specific effects, and ν_{it} is the error term. For i = 1, …, N, t = 2, …T. The single regressor

x_{it} is correlated with η_i and predetermined with respect to ε_{it}, meaning that $\mathrm{E}(x_{it}\ \nu_{it+s}) = 0, s = 0,...,T-t$, but $\mathrm{E}(x_{it}\ \nu_{it+s}) \neq 0, r = 1,...,t-1$. A commonly used estimator is the GMM estimator in the model in first difference (Arellano and Bond, 1991),

$$\Delta y_{it} = \beta^* \Delta y_{i,t-1} + \gamma \Delta x_{it} + \Delta u_{it}; \quad t = 2,....,T \tag{20}$$

After accounting for the time-specific effects and grouping all explanatory variables in a vector *x*, Equations (18), (19), and (20) can be rewritten as:

$$\Delta y_{it} = \beta^* \Delta y_{i,t-1} + \gamma \Delta x_{it} + \Delta \eta_{it} + \Delta \nu_{it} \tag{21}$$

The estimation of cross-country effects is based on a regression on time-averaged data. In order to sweep out unobserved individual country specific effects η_i that are a source of inconsistency in the estimates and specified and in order to obtain a consistent estimate of β^* as $N \to \infty$ for fixed T we take first difference of equation (16):

$$y_{it} - y_{i,t-1} = \beta^*(y_{i,t-1} - y_{i,t-2}) + \gamma(x_{it} - x_{i,t-1}) + (\nu_{it} - \nu_{i,t-1}) \tag{22}$$

Since the differenced lagged dependent variable and the differenced error term are correlated OLS estimation of (22) will not produce a consistent estimate of β^*, even if the regressor, $x_{i,t}$ is strictly exogenous.

The GMM estimator which was suggested by Arellano and Bond (1991) is known to be rather inefficient when instruments are weak because it makes use of the information contained in differences only. Blundell and Bond (1998) suggest making use of additional level information besides the differences. The combination of moment restrictions for differences and levels results in an estimator which was called GMM-system estimator.

Following Arrelano and Bond (1991) and Hoeffler (2002) we can use values of the predetermined x_{it} lagged one period or more as valid instruments in the first differenced growth equation. It is also straightforward to treat for example investment as an endogenous variable. This means that we are allowing for correlation between current investment and current shocks to GDP, as well as feedback from past shocks to GDP, i.e.

$$E(x_{it}\nu_{is}) \neq 0 \qquad \text{for } s \leq t$$

and

$$E(x_{it}\nu_{is}) = 0 \qquad \text{for } s > t \quad \text{only}$$

In this case, valid instruments in the differenced equation are values of the endogenous x_{it} lagged two periods or more.

Blundell and Bond (1998) show that estimators relying on lagged levels as instruments for current differences are likely to perform poorly when the series are close to a random walk. In this case the available instruments are only weakly correlated with the endogenous variables, and the GMM estimator is likely to suffer from serious finite sample bias, as well as imprecision (Hoeffler, 2002). Instead they suggest estimating a system combining two sets of equation. One set of equations is the differenced equation (22) as follow:

$$y_{it} - y_{i,t-1} = \beta^*(y_{i,t-1} - y_{i,t-2}) + \gamma(x_{it} - x_{i,t-1}) + (\nu_{it} - \nu_{i,t-1})$$

For which we use suitably lagged levels of y_{it} and x_{it} as instruments, as discussed for the first differenced GMM estimation. The other set of equation in the system are the levels equation

$$y_{it} = \alpha + \beta^* y_{i,t-1} + \gamma x_{it} + \eta_i + \nu_{it}$$

Provided the x_{it} regressor satisfies

$$E(\Delta x_{it}\eta_i) = 0 \tag{23}$$

and the initial conditions satisfy the restriction

$$E(\Delta y_{i2}\eta_i) = 0 \tag{24}$$

Arrelano and Bover (1995) proposed estimators of $\Delta y_{i,t-1}$ and Δx_{it} as instruments in the level equation. Assumed that equation (23) allows the level of x_{it} to be correlated with the unobserved country specific effects, η_i, but requires the changes in x_{it} to be uncorrelated with η_i. Given equation (23), assumption equation (24) will be satisfied provided the process equation (16) has been generating the y_{it} series for a sufficiently long time (Hoeffler, 2002).

The consistency of the GMM estimators depends on whether lagged values of the explanatory variables are valid instruments in the growth regression. We address this issue by considering two specification tests suggested by Arellano and Bond (1991) and Arellano and Bover (1995). The first is a Sargan test of over-identifying restrictions, which tests the overall validity of the instruments. Failure to reject the null hypothesis gives support to the model. The second test examines the null hypothesis that the error term is not serially correlated. As with the Sargan test, the model specification is supported when the null hypothesis is not rejected. In the system specification we test whether the differenced error term (that is, the residual of the regression in differences) is second-order serially correlated. Second-order serial correlation of the differenced residual would indicate that the original error term is serially correlated and follows a moving average process at least of order one. This would reject the appropriateness of the proposed instruments (and would call for higher-order lags to be used as instruments).

The test statistic AR(1) (m_1) and AR(2) (m_2) test for presence of serial correlation in the first differenced residuals of first and second order, respectively; they are asymptotically normally $N(0,1)$ distributed under the null of no serial correlation (Arrelano and Bond, 1991). First order autocorrelation (AR (1)) is expected to be negative significant but according to the second order autocorrelation (AR (2)) test there is no significant which the crucial point with respect to the validity of the instruments.

3.2 Data

The data set consists of a panel of observations for thirteen Asian countries namely China, Hong Kong, China, Korea, Japan, Indonesia, Malaysia, Philippines, Singapore, Thailand, Bangladesh, India, Pakistan, and Sri Lanka for the period 1982-2001.These data contain a number of time invariant and time varying variables, where the time varying variables are averaged over four year. All data are collected from the World Development Indicator (World Bank CD-ROM 2005), Asian Development Bank (ADB, 2004), and The Government Finance Statistics in various years (GFS).

4. Empirical results

In this section, results will be presented for the dynamic panel data estimators of Model 1 and Model 2 in Table 1 and Table 2 which are a version of the augmented Solow model, where the logarithm of the government expenditure in health, education and defence, distortionary taxation, budget balance, aggregate of government expenditure, aggregate of other fiscal varibles, the savings in physical capital, and population growth Results will be outlined for GMM estimation in difference (DIF) and system (SYS) version and also the OLS level and within group's estimation techniques.

4.1 OLS and within groups results

Table 1 and Table 2 follow the suggestion of Arrelano and Bond (1991) and Blundell and Bond (1998, 1999), all regressions includes time dummies which we found to be jointly significant in every regression. The tables show the parameter estimates, (in parenthesis) the standard errors of the parameter estimates (robust to arbitrary forms of cross-sectional and time-series heteroskedasticity) and a selection of diagnostic statistics.

In Table 1 and Table 2, the left hand side variable is the change in the logarithm of explanatory variables. We begin our analysis with an OLS regression in column 2 through Table 1 and Table 2, the estimate of the coefficient for the initial per capita GDP ($\ln rgdp_{it-1}$) is negative [-1.244 (Model 1) and -1.754 (Model 2] and statistically significant at 1 percent level for all models. The negative coefficient on initial GDP as in most published growth regressions is interpreted as conditional convergence as suggested by the Solow model (Hoeffler, 2002). In Table 1, the estimate of the coefficient of the government expenditure in health ($\ln he$) is positive (3.672) and statistically significant at the 1 percent level, while the estimate of the coefficient of the government expenditure in education ($\ln ee$) is also positive (5.429) and statistically significant at the 1 percent level. Therefore, results in Table 1 shows that education and health expenditures have significant effect on real per capita GDP growth. Thus, we found that educational and health expenditures increase growth.

The estimate of the coefficient of the government expenditure in defence ($\ln de$) is negative [-4.160 (Model 1)] and statistically significant at the 1 percent level. Therefore, we found that defence expenditure has an adverse effect on economic growth that is, increase in defence expenditures will decrease economic growth.

Table 2 shows that there is a negative coefficient (-0.078) and statistically significance at the 5 percent level for the distortionary taxation ($\ln dt$) in Model 2. Thus distortionary taxation has adverse effect on economic growth which is a negative relationship. The estimate of the coefficient of the budget balance ($\ln bb$) is also negative (-1.549) and statistically significant at the 5 percent level. Thus budget balance has a significant and adverse effect on real per capita GDP growth.

The estimate of the coefficient for the aggregate of other fiscal variables is positive (0.184) and statistically significant at the 5 percent level in Model 1. Thus aggregate of other fiscal variables has a positive and significant effect on real per capita GDP growth. The estimate of the coefficient for the aggregate of government expenditure ($\ln GE$) is positive (1.950) and statistically significant at the 5 percent level in Model 2. Thus aggregate of government expenditure has a positive and significant effect on real per capita GDP growth.

By assuming the savings in physical capital (investment) ($\ln S_k$) is potentially endogenous variable and current population growth rate is potentially exogenous variable, these estimates already allow for the possibility of serially uncorrelated measurement error in either of these explanatory variables. In Table 1 and Table 2, we found that the estimate of the coefficient for the savings in physical capital (investment) (lnS_k) is positive for all models [0.369 (Model 1) and 0.945 (Model 2)] and statistically significant at the 5 percent level in column 2 (Model 1) and statistically significant at the 1 percent level for Model 2. Therefore, this indicates that savings in physical capital (investment) (lnS_k) has a significantly positive effect on the steady state level of per capita GDP growth. Column 2 in Table 1 and Table 2 also shows that the coefficient on population growth (ln(n+g+δ)) is negative [-1.113 (Model 1) and -1.018 (Model 2)] and statistically significant at the 1 percent level in both models. This is in line with the neoclassical growth model that predicts that as population increases, the steady state level of per capita GDP will decline through lowering of the capital labour ratio. Thus, our findings here is also in line with Solow model which is, the negative coefficient on initial GDP as in most published growth regressions is interpreted as conditional convergence while investment is positive and population growth is negative.

Overall, based on the Table 1 and Table 2 and the argument above, in accordance with expectations in the presence of firm-specific fixed effects, OLS panel estimates seem to deliver an upward-biased estimate of the lagged dependent variable coefficient.

Our analysis with a within groups estimator was used; the results are shown in the third column of Table 1 and Table 2. We found that the estimate of the coefficient for the initial per capita GDP (ln$rgdp_{it-1}$) is negative [-1.563 (Model 1) and -2.220 (Model 2)] and statistically significant at 1 percent level for both models. Same as in OLS estimate, the negative coefficient on initial GDP as in most published growth regressions is interpreted as conditional convergence as suggested by the Solow model. Therefore, comparing the estimated coefficient of the OLS levels regression and the within groups estimation, we found that the OLS levels provides a higher estimate for the coefficient on the initial real per capita GDP than the within groups estimation for all models in all tables.

In Model 1, the estimate of the coefficient of the government expenditure in health (ln*he*) is positive (4.121) and statistically significant at the 1 percent level, while the estimate of the coefficient of the government expenditure in education (ln*ee*) is also positive (6.037) and statistically significant at the 1 percent level. Thus, government expenditure on health and education has a positive and significant effect on real per capita GDP growth in within group's estimation. The estimate of the coefficient of the government expenditure on defence (ln*de*) is negative (-5.805) and statistically significantly at the 1 percent level. Thus defence expenditure has a negative and significant effect on real per capita GDP. Table 2 shows that there is a negative coefficient (-0.090) and statistically significance at the 5 percent level for the distortionary taxation (ln*dt*) in Model 2. The estimate of the coefficient of the budget balance (ln*bb*) is negative (-0.065) and statistically significant at the 5 percent level in column 3. Thus, distortionary taxation and budget balance has a positive and significant effect on real per capita GDP growth in within groups estimation.

Model 1, the estimate of the coefficient for the aggregate of other fiscal variables (ln*OFV*) is positive (0.418) and statistically significant at the 5 percent level in Table 1. Thus, the aggregate of other fiscal variables has a positive and significant effect on real per capita GDP growth. In Table 2, the estimate of the coefficient for the aggregate of government expenditure (ln*GE*) is positive (2.461) and statistically significant at the 5 percent level. Thus, aggregate of government expenditure has positive and significant effects on real per capita GDP.

The estimate of the coefficient for the savings in physical capital (investment) (lnS_k) is positive [0.471 (Model 1and 0.694 (Model 2)] and statistically significant at the 5 percent level in both models. The coefficient for the population growth (ln(n+g+δ)) is negative [-1.120 (Model 1) and -0.956 (Model 2)] and statistically significant at the 1 percent level in both models. Thus, institutions and savings in physical capital have a positive and significant effect on real per capita GDP growth and population growth is negative and has a significant effect on real per capita GDP growth.

4.2 First differenced gmm and system gmm results

Hoeffler (2002) argues that for special cases of spherical disturbances, the one-step and two-step GMM estimators are asymptotically equivalent for the first-differenced estimator (Arellano and Bond, 1991). In this case the two-step estimator is more efficient. Therefore, Column 4 presents the results using the Arellano and Bond (1991) first differenced GMM estimator in Table 1 and Table 2 for Model 1 and Model 2. We assume that the initial GDP is predetermined and investment is endogenous. We also assume that current population growth is exogenous, in the sense of being uncorrelated with shocks to GDP per capita. This allows the use of both current and lagged levels of population growth as instruments in the first differenced equation. The OLS level will give an estimate of β that is biased upwards in the presence of individual-specific effects (Hsiao, 1986) and within groups will give an estimate of β^* that is seriously biased downwards in short panel (Nickell, 1981). Thus, a consistent estimate of β^* can be expected to lie in between the OLS levels and within groups estimates. If we observe that the first differenced GMM estimate is close to or below the Within groups estimate, it seems likely that the GMM estimate is also biased downwards in our application, perhaps due to weak instruments.

In Column 4 Table 1 and Table 2, we found that all coefficients are positive and statistically significant at the 1 percent and 5 percent level, except the coefficient of population growth ($\ln(n+g+\delta)$), government expenditure on defence, distortionary taxation, and budget balance are negative but still statistically significant at 1 percent and 5 percent level. From these tables we found that the first differenced GMM estimate of the coefficient on the initial real per capita GDP in all tables are negative [-1.472 (Model 1) and -1.927 (Model 2)] and statistically significant at the 1 percent level in both models. These results lies close to the within groups estimate. Therefore, we can conclude that our case for the first differenced GMM estimate seems to be downward biased, because it is very close to the within groups estimate which is expected to be seriously biased downwards in a short panel with six or fewer time periods (Nickell, 1981).

In Column 4, the estimate of the coefficient of the government expenditure in health (ln*he*) is positive [4.583 (Model 1)] and statistically significant at the 1 percent level, while the estimate of the coefficient of the government expenditure in education (ln*ee*) is positive [5.220 (Model1)] and statistically significant at the 1 percent level in Table 1. Therefore, the education and health expenditures have significant effects on real per capita GDP growth and this result is consistent with the results from the Fully Modified OLS which has been discussed earlier.

Table 2 shows that there is a negative coefficient (-0.055) and statistically significant at the 5 percent level for the distortionary taxation (ln*dt*). The estimate of the coefficient of the budget balance (ln*bb*) is negative (-0.113) and statistically significant at the 5 percent level. Thus, distortionary taxation and budget balance have a significant and adverse effect on real per capita GDP growth which is a negative relationship.

The estimate of the coefficient for the aggregate of other fiscal variables (ln*OFV*) is positive [0.124 (Model 1)] and statistically significant at the 1 percent level in Table 1. The aggregate of government expenditure (ln*GE*) is positive [2.154 (Model 2)] and statistically significant at the 1 percent level in Table 2.

In Table 1 and Table 2, the estimate of the coefficient for the savings in physical capital (investment) ($\ln S_k$) is positive [0.470 (Model 1) and 0.930 (Model 2)] and statistically significant at the 1 percent level in Model 1 and statistically significant at the 5 percent level in Model 2. The coefficient on population growth ($\ln(n+g+\delta)$) is negative [-0.797 (Model 1) and -0.735 (Model 2)] and statistically significant at the 1 percent level for both models. Again we found that the results for the savings in physical capital (investment) and population growth in first differenced GMM are in line with the Solow model which is the negative coefficient on initial GDP as in most published growth regressions is interpreted as conditional convergence while investment is positive and population growth is negative.

As mentioned previously, differenced GMM results potentially suffer from a bias in the direction of within group's results, due to weak instruments, related with persistent time series. In such case, GMM-SYS results would be preferable. The presented results do not seem to point towards a major problem, since the GMM-DIF estimations are situated quite central amongst the two extremes of OLS and within groups. This leads to the conclusion that the system GMM results shown in Table 1 and Table 2 are to be the preferred parameter estimates.

The system GMM estimator thus combines the standard set of explanatory variables in first differences with suitably lagged levels as instruments, with an additional set of explanatory variables in levels with suitably lagged first differences as instrument. As an empirical matter, the validity of these additional instruments can be tested using standard Sargan tests of over-identifying restriction (Arrelano and Bond, 1991). For system GMM the two-step estimator is always more efficient than the one-step estimator. However, Monte Carlo studies show that the efficiency gain is small and that the two-step estimator converges only slowly to its asymptotic distribution. In finite samples, the asymptotic standard errors associated with the two-step GMM estimator can be seriously biased downward (Blundell and Bond, 1998). Therefore, we follow Hoeffler (2002) who prefer to report the one-step estimates. The fifth columns of Table 1 and Table 2 report the results from using system GMM, again treating investment as endogenous and population growth as exogenous in the sense described before. The estimate of the coefficient on the initial real per capita GDP is not obviously biased which lies well above the within groups estimate and well below the OLS levels estimate, and the estimates of the coefficients are more precise than the ones obtained from first differenced GMM. We agreed with Hoeffler (2002) where the additional instruments using the system GMM estimator appear to be both valid and highly informative in this context.

In Column 5 (Table 1), the estimate of the coefficient of the government expenditure in health (ln*he*) is positive (4.524) and statistically significant at 1 percent level, while the estimate of the coefficient of the government expenditure in education (ln*ee*) is also positive (5.863) and statistically significant at the 1 percent level. Therefore, we found that educational and health expenditures increase growth and this result is also consistent with the results from OLS, within groups and first differenced GMM as discussed earlier. These findings may also be supported by the evidence that many other authors also found that a significant positive effect on growth from education and health expenditures (Barro and Sala-i-Martin 1995; Hansson and Henrekson 1994; Chen and Gupta 2006; Dreger and Reimers 2005). The estimate of the coefficient of the government expenditure on defence (ln*de*) is negative (-1.703) and statistically significantly at the 1 percent level in Model 1. Thus, defence expenditure has a negative and significant effect on real per capita GDP.

Table 2 shows that there is a positive coefficient (-0.018) and statistically significant at the 1 percent lever for the distortionary taxation (ln*dt*) in Model 2. The estimate of the coefficient of the budget balance (ln*bb*) is positive (-1.359) and statistically significant at the 1 percent level both models. Thus, distortionary taxation and budget balance have a significant and adverse effect on real per capita GDP growth.

The estimate of the coefficient for the aggregate of other fiscal variables (ln*OFV*) is positive [0.133 (Model 1)] and statistically significant at the 1 percent level in Table 1. The aggregate of government expenditure (ln*GE*) is positive [3.010 (Model 2)] and statistically significant at the 1 percent level in Table 2. This result also, which matches as overall the finding in Levine and Renelt (1992), Devarajan *et al*. (1996) and Gemmel *et al*. (1999) show that the composition of government expenditure is a highly significant factor in economic growth. Again we found that the aggregate of other fiscal variables and aggregate of government expenditure have a positive and significant effect on real per capita GDP growth in system GMM.

In Table 1 and Table 2, as expected, the estimate of the coefficient for the savings in physical capital (investment) ($\ln S_k$) is positive [0.270 (Model 1) and 0.632 (Model 2)] and statistically significant at the 1 percent level in both models. The coefficient on population growth ($\ln (n+g+\delta)$) is negative [-0.844 (Model 1) and -0.871 (Model 2)] and statistically significant at the 1 percent level for both models. Our results matched many studies which controlled capital accumulation by including the rate of investment or savings. Levine and Renelt (1992) argue that the positive correlation between growth and the share of investment in GDP is one of the few robust findings from the cross-country growth regression literature. On the other hand, there is most growth studies accounted for population growth. The effect of population growth on growth in GDP per capita tended to be negative in some studies, but this finding is rather fragile (Levine and Renelt, 1992). Therefore, we conclude that system GMM savings in physical capital have a positive and significant effect on real per capita GDP growth and population growth is negative and has significant effect on real per capita GDP growth.

4.3 Diagnostic results

The Diagnostics part of the Table 1 and Table 2 show three diagnostic tests of the appropriateness of the instruments used. The first test is Wald (joint) tests to test the significance of all regressors. The Wald test for the joint significance of regressors (excluding time dummies) is statistically significant at the 1 percent level in all four estimations for all models. The Wald test indicates that the initial real per capita GDP are jointly significant.

The second test is a Sargan test of identifying restrictions under the null hypothesis of the validity of the instruments (Arellano and Bond, 1991; 1998). The validity of the instrument set is checked using a Sargan test. This test is asymptotically distributed as chi-squared under the null. The instruments used in the first differenced GMM or in the system GMM are not rejected by the Sargan test of over-identifying. In Table 1 and Table 2, we found that the Sargan test of the validity of instruments used is not statistically significant at the 10 percent level in first differenced GMM and system GMM for all models. With respect to the Sargan test of over-identifying restrictions, the high *p-value* suggests that we cannot reject the null hypothesis that the set of instruments is appropriate. Therefore, the Sargan test supports the validity of the GMM estimator and GMM system and do not indicate a serious problem with the validity of the instrumental variables. This is consistent with the presence of measurement errors (Blundell and Bond, 1998; 1999), as well as instruments used in the GMM-SYS estimation and earlier do pass the test.

The third tests are the tests of first and second-order serial correlation in the first-differenced residuals, reported as the asymptotically standard normal distribution values m_1(AR(1)) and m_2(AR(2)). As required, the test for first-order autocorrelation AR (1) rejects the null; the *p-values* of the Arellano and Bond statistics in Table 1 and Table 2 at the 1 percent significance level and 5 percent significance level. While the test for second-order autocorrelation AR(2) fails to reject the null hypothesis of no autocorrelation and the statistics reported are *p-values* giving the probability of correctly rejecting the null hypothesis of no autocorrelation. Moreover, as expected, we do not find the presence of statistically significant second-order serial correlation. Therefore, both AR (1) and AR (2) test support the validity of the first differenced GMM and the system GMM estimator of Table 1 and Table 2. It was mentioned previously that the consistency of GMM-SYS estimates would be related to the absence of serial correlation in the error terms. The presented estimations clearly seem to be consistent; the absence of serial correlation shows in the differenced residuals by significant negative first order serial correlation and no second order serial correlation.

We conclude that the Wald tests of the joint significance of the variables as well as the tests for autocorrelation and the Sargan test confirm that the GMM estimator estimated for Model 1 and Model 2 in Table 2 and Table 2 are appropriate. First-differencing introduces AR (1) serial correlation when the time-varying component of the error term in levels is serially uncorrelated (Arellano and Bond, 1991; 1998). Therefore, GMM estimator is consistent only when second-order correlation is not significant although first-order correlation need not be zero. Again, the first and second order serial correlations tests are all satisfied.

5. Conclusion

We estimate the growth equation using the generalized method of moments (GMM) method as proposed by Arellano and Bond (1991) and GMM-system estimator by Blundell and Bond (1998) as a dynamic panel data analysis. These dynamic panel estimator controls possible endogeneity of the regressors and the possible bias in specifications with nearly integrated regressors, as is the case for GDP per capita and the explanatory variables. We found that there is a positive and statistically significant effect of government expenditure on health and education, aggregate of government expenditure and aggregate of other fiscal variables on GDP per capita. Turning to the remaining explanatory variables, we found that defence expenditure, budget balance, and distortionary taxation are significantly and negatively related to GDP per capita. Our results on initial real per capita GDP, savings in physical capital (investment) and population growth rate are in line with the neoclassical growth model that predicts that as population increases, the steady state level of per capita GDP will decline through lowering of the capital labour ratio. Thus, our findings here is also in line with the Solow model which is the negative coefficient on initial GDP as in most published growth regressions is interpreted as conditional convergence while investment is positive and population growth is negative.

Overall, we concluded that fiscal policy is one of the most important instruments of government economic policy. To establish the relationship between fiscal policy and economic growth, the mechanisms through which fiscal policy affects the above mentioned factors of economic growth have been investigated in the long run. The significance arises for two reasons. First, they can have opposite impacts on the economy. Second, there is an outside lag inherent in fiscal policy. The long run impacts of fiscal policy is not only an interesting intellectual and theoretical exercise but it also has important implications for policy making.

References

Abdullah, H., Habibullah, M.S. and Baharunshah, A.Z. (2008), Fiscal Policy. Institutions and Economic Growth in Asian Economies: Evidence from the Pedroni's Cointegration Approach. *International Journal of Business and Managemen*t, Vol. 3, No. 4, pp. 107-126.

Abu-Bader, S. and Abu-Qarm, A.S. (2003). Government Expenditure, Military Spending and Economic Growth: Causality Evidence from Egypt, Israel, and Syria. *Journal of Policy Modeling*, 25, 567-583.

Arellano, M. and Bond, S. (1991). Some Tests of Specific Action for Panel Data: Monte Carlo Evidence and an Application to Employment Equations. *Review of Economic Studies*, 58, 277-297.

Arellano, M. and Bover, O. (1995). Another Look at the Instrumental Variable Estimation of Error-Component Models. *Journal of Econometrics*, 68, 29-52.

Baltagi, B.H. (1995). *Econometric analysis of panel data,* John Wiley and Sons.

Barro, R.J. and Sala-I-Martin, X. (1995). *Economic Growth,* McGraw-Hill: New York

Biswas, B. and Ram, R. (1986). Military Expenditures and Economic Growth in Developing Countries: An Augmented Model and Further Evidence, *Economic Development and Cultural Change*, 34, 361-372.

Blundell, R. W. Bond, S. R. (1999). GMM estimation with persistent panel data: an application to production functions, Cemmap Working Paper, CW99/4. London: Institute for Fiscal Studies.

Blundell, R.W. and Bond, S.R (1998). Initial conditions and moment restrictions in dnamic panel data models. *Journal of Econometrics*, 87, 115-143.

Bonds, S. (2002). Dynamic panel data models: a guide to micro data methods and practice. *Portuguese Economic Journal*, 1: 141-162.

Caselli, F., Esquivel, G. and Lefort, F. (1996). Reopening the Convergence Debate: A new Look at Cross-Country Growth Empirics. *Journal of Economic Growth*, 1, 363-389.

Chen, P. and Gupta, R. (2006). *An investigation of openness and economic growth using panel estimation,* University of Pretoria, Working Paper, 22.

Demetriades, P. and Law, S.H. (2006). Finance, Institutions and Economic Development. *International Journal of Finanace and Economics*, 11, 1-16.

Devarajan, S., Swaroop, V. and Zou, H. (1996). The Composition of Public Expenditure and Economic Growth. *Journal of Monetary Economics,* 37(2), 313-344.

Dreger, C. and Reimers, H.E. (2005). Health care expenditures in OECD countries: A panel unit root and cointegration analysis. *The Institute for the Study of Labor (IZA), Discussion Paper*, No. 1469.

Gangopadhyay, S. and Chatterji, A. (2005). *Economic Globalization in Asia,* Ashgate Publishing.

Gerson, P. (1998). The Impact of Fiscal Policy Variables on Output Growth” International Monetary Fund. *Working Paper* , 98/1.

Ghura, D. and Hadjimichael, M.T. (1996). Growth in Sub-Saharan African. *IMF Staff Paper*, 43, 605-635.

Glomm, G. and Ravikumar, B. (1994). Public Investment in Infrastructure in a Simple Growth Model. *Journal of Economic Dynamics and Control,* 18, 1173-1188.

Hansson, P. and Hendrekson, M. (1994). A New Framework for Testing the Effect of Government Spending on Growth and Productivity. *Public Choice*, 31, 381-401.

Hassan, M.K., Waheeduzzaman, M. and Rahman, A. (2003). Defense Expenditure and Economic Growth in the SAARC Countries. *The Journal of Social, Political, and Economic Studies*, 28(3), 275-293.

Hoeffler A.K. (2002). The augmented Solow Model and the African growth debate. *Oxford Bulletin of Economics and Statistics,* 64(2), 0305-9049.

Hsiao, C. (1986). *Analysis of panel data,* Cambridge; Cambridge University Press.

International Monetary Fund, (various years), a manual on Government Finance Statistics (GFS).

Islam, N. (1995). Growth Empirics: A Panel Data Approach. *Quarterly Journal of Economics*, 110, 1127-1170.

Kneller, R., Bleaney M. F. and Kneller, N. (1999). Fiscal Policy and Growth: Evidence from OECD Countries. *Journal of Public Economics,* 74:171-190.

Knight, M., Loayza, N. and Villanueva, D. (1993). Testing the Neoclassical Theory of Economic Growth, A Panel Data Approach. *International Monetary Fund Staff Paper*, 40(3).

Landau, D. (1986). Government and Economic Growth in Less Developed Countries: An empirical study, *Economic Development and Culture Change,* 35, 35-75.

Mankiw, G., Romer, D. and Weil, D. (1992). A Contribution to the Empirics of Economic Growth. *Quarterly Journal of Economics*, 107, 407-437.

Nickell, S. (1981). Biases in Dynamic Models with Fixed Effects. *Econometrica*, 49, 1417-1426.

Ramirez, M.D. and Nazmi, N. (2003). Public Investment and Economic Growth in Latin America: An Empirical Test. *Review of Development Economics*, 7, 115-126

Solow, R. M. (1956). A Contribution of the Theory of Economic Growth. *Quarterly Journal of Economics,* 70: 65-9.

Tanzi, V. and Zee, H. (1997). Fiscal Policy and Long-Run Growth. *International Monetary Fund Staff Papers,* 44, 65-94.

World Bank. (2005). World Development Indicators 2004 and 2005 on CD-ROM, World Bank, Washington D.C.

Yasin, M. (2003). Public Spending and Economic Growth: Empirical Investigation of Sub-Saharan Africa. *Southwestern Economic Review*, 59-68

Zagler, M. and Dürnecker, G. (2003). Fiscal Policy and Economic Growth. *Journal of Economic Survey*, 17(3), 397-418.

Table 1. Estimation of the Model 1; Dependent variable $\Delta \ln rgdpc_{it}$

	OLS	Within Groups	DIF-GMM	SYS-GMM
(ln$rgdp_{it-1}$)	-1.244*(0.271)	-1.563*(0.185)	-1.472*(0.112)	-0.904* (0.019)
(lnhe)	3.672* (4.388)	4.121* (3.974)	4.583* (4.343)	4.524* (8.942)
(lnee)	5.429* (3.668)	6.037* (3.374)	5.220* (1.961)	5.863* (5.807)
(lnde)	-4.160* (2.872)	-5.805* (3.669)	-3.095* (1.089)	-1.703* (4.347)
(lnOFV)	0.184**(0.169)	0.418** (0.203)	0.124* (0.055)	0.133* (0.113)
(lnS_k)	0.369**(0.195)	0.471** (0.167)	0.470* (0.162)	0.270* (0.117)
(ln(n+g+δ))	-1.113* (0.288)	-1.120* (-5.67)	-0.797* (0.311)	-0.844* (0.231)
No. of obs.	52	52	52	52
Wald test:	269.3* (0.000)	977.6* (0.000)	3616.* (0.000)	789.8* (0.000)
Sargan test:	-	-	27.7 (0.428)	167.3 (0.981)
AR(1) test:	-1.733**(0.043)	-2.016**(0.044)	-2.892* (0.004)	-1.779* (0.000)
AR(2) test:	-0.095(0.924)	-0.773 (0.439)	-1.862 (0.163)	-1.325 (0.185)

Note: Heteroskedasticity consistent standard error is reported in parentheses; (*, **, ***) denotes the level of significance levels at 1%, 5%, and 10%, respectively. Time dummies are included and a constant is included. – The figures reported for the tests of first and second order correlation under the System-GMM column AR(1) and AR(2) as well as for the Wald test and Sargen test and are the p-values of the null hypothesis.

Table 2. Estimation of the Model 2; Dependent variable $\Delta \ln rgdpc_{it}$

	OLS	Within Groups	DIF-GMM	SYS-GMM
(ln$rgdp_{it-1}$)	-1.754*(0.291)	-2.220*(0.345)	-1.927*(1.560)	-0.906*(0.026)
(lndt)	-0.078**(0.054)	-0.090**(0.077)	-0.055**(0.088)	-0.018*(0.047)
(lnbb)	-1.549**(0.029)	-0.065**(0.026)	-0.113**(0.063)	-1.359*(0.493)
(lnGE)	1.950**(1.238)	2.461**(2.458)	2.154*(0.720)	3.010*(1.931)
(lnS_k)	0.945*(0.207)	0.694**(0.312)	0.930**(0.801)	0.632*(0.135)
(ln(n+g+δ))	-1.018*(0.261)	-0.956*(0.284)	-0.735*(2.353)	-0.871*(0.327)
No. of obs.	52	52	52	52
Wald test:	329.8*(0.000)	551.5*(0.000)	8179.*(0.000)	229.8*(0.000)
Sargan test:	-	-	26.4(0.607)	154.1(0.151)
AR(1) test:	-1.793**(0.043)	-1.813**(0.040)	-2.774*(0.006)	-1.513*(0.005)
AR(2) test:	-0.267(0.789)	-0.858(0.391)	-1.379(0.168)	-1.188(0.235)

Note: Heteroskedasticity consistent standard error is reported in parentheses; (*, **, ***) denotes the level of significance levels at 1%, 5%, and 10%, respectively. Time dummies are included and a constant is included. – The figures reported for the tests of first and second order correlation under the System-GMM column AR(1) and AR(2) as well as for the Wald test and Sargen test and are the p-values of the null hypothesis.