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Are Public Investment and Education Expenditures Productive in the Argentine Case? A FMOLS and DOLS Analysis, 1960-2019

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Abstract

This paper examines whether public investment spending and public education expenditures (a proxy for human capital) in Argentina have a positive and significant effect on economic output and labor productivity for the 1960-2019 period. The paper estimates a simple model that incorporates the impact of public and private investment spending, education expenditures (including at the secondary level), and the labor force. It presents a modified empirical counterpart to the simple model and tests for unit roots and performs both a Johansen cointegration test and a Gregory and Hansen cointegration test with an endogenously determined regime shift. FMOLS and DOLS long-run estimates for the 1960-2019 period suggest that public and private investment spending, government education expenditures, and the labor force have a positive and significant effect on the level of economic output and labor productivity. For comparison purposes, both methodologies are used because these estimators are extremely consistent, particularly the DOLS estimator, even in the presence of both endogeneity and serial correlation of any order. The concluding section summarizes the major results and discusses potential avenues for future research on this important topic.

Keywords: Complementarity Hypothesis, DOLS estimator, Education expenditures, FMOLS estimator, Endogenous growth, Gregory-Hansen cointegration single-break test, Lee-Strazicich two-break unit root test, Johansen Cointegration Test, Private and Public Investment spending.

JEL codes: C22, H54, O54

1. Introduction

Argentina began the process of economic stabilization and reform with the country's adoption of a currency board system in 1991—the so-called "Convertibility Plan"-- under the administration of Carlos Saul Menem (Note 1). The Plan set the Argentinean peso to the dollar on a one-to-one basis, thus eliminating the ability of the government to finance budget deficits via direct money. A major accomplishment of the currency board was to reduce dramatically the rate of inflation from 2,314 percent in 1990 to 4.1 percent in 1994, and less that 1 percent in 1998! Unfortunately, the stabilization of the Argentine economy during the nineties was not sustainable in view of the impact of several external shocks and ill-conceived policies that paved the way for the economic and financial debacle associated with the collapse of the Convertibility Plan in 2001-2002. For example, the Asian and Russian crises in the late 1990s led to a significant flight of capital and a substantial rise in real interest rates with their adverse effects on investment and consumption spending. Next, the devaluation of the Brazilian currency (the real) in 1999 had a severe effect on the Argentine economy because at least a quarter of its exports were destined to that country (ECLAC, 2012). Finally, the economic situation was further aggravated by the fact that the dollar continued to appreciate in real terms relative to the Euro and the Yen, thus further undermining the competitiveness of the Argentine economy given its hard peg to the dollar and its policy of unrestricted capital mobility (see Baer et. al., 2002).

Table 1. Argentina: Investment as a Share of GDP (in percent), 1980-2019

Year	Private Investment	Public Investment
1980	19.2	6.1
1990	9.4	4.6
1992	14.9	1.8
1996	16.1	2.0
2000	15.4	1.0
2002	9.2	0.7
2004	10.5	1.3
2006	13.2	2.5
2008	15.1	3.3
2010	14.9	3.5
2011	14.1	2.8
2012	13.8	3.0
2013	13.7	3.0
2014	13.5	3.1
2015	13.6	3.1
2016	14.0	3.7
2017	15.6	2.6
2018	15.3	1.3
2019	13.1	1.0
Years	Simple Average	
1970-1979	13.6	9.1
1980-1989	15.0	4.9
1990-1999	15.7	1.6
2000-2010	12.7	2.1
2011-2019	12.6	2.4

Source: World Development Indicators. Washington, D.C., The World Bank, 2021; M.E.P., Argentina: Sustainable Output Growth After the Collapse. Buenos Aires, Ministerio De Economia Argentina; and ECLAC, 2022.

To make matters worse, IMF-sponsored stabilization policies adopted by the Argentine government in the wake of these shocks led to short- and long-term economic (negative) effects on consumption, investment, and economic growth (see Calva, 1997; Maia and Kweitel, 2003; Stiglitz, 2012; Taylor 1997; Baer et. al., 2002; Weisbrot and Sandoval, 2007; and Weisbrot, 2011). These policies entailed across-the-board cuts in public spending and tight restrictions on credit creation in order to meet stringent fiscal deficit targets, reduce the rate of inflation, and free resources to service the external debt (Note 2). Not surprisingly, they have resulted in a lackluster and erratic performance of private capital formation during the past two and a half decades. Table 1 below shows that Argentina's private investment as a proportion of GDP fell dramatically during the "lost decade" of the 1980s, reaching a low of 9.4 percent in 1990 which amounted to less than half its level in 1980. Following the adoption of the Convertibility Plan it rose to a high of 16.1 percent in 1996, from which it fell again to 9.2 percent in 2002 and a dismal 7.6 percent in 2003 (not shown in Table 1) as a result of the economic crisis and IMF-sponsored austerity measures (Note 3). In this connection, most economists believe that it is essential for Argentina (and other Latin economies) to improve and sustain its investment performance if it is going to lay the groundwork for rapid and sustained economic growth and employment opportunities for its rapidly expanding labor force (see Calva, 1997; Moguillansky, 1996; Baer et al., 2002; and Weisbrot and Sandoval, 2007).

A number of prominent researchers have cited the dramatic fall in public investment spending on economic and social infrastructure, brought about by the need to meet the stringent fiscal deficit targets of the stabilization program, as one possible factor in explaining the poor economic performance of Argentina and other Latin American countries. Table 1 shows that public investment spending in economic and social infrastructure as a proportion of GDP fell precipitously from 4.6 percent in 1990 to barely 1 percent in 1994, averaging only 2 percent during the 1995-99 period before falling again in the 2001-2003 period to less than 1 percent. Moreover, the average public investment spending on economic infrastructure for the 1990s and early 2000s is only a third of that of the 1980s and barely one fifth of the average level recorded during the 1970s. Table 1 also reveals that under the pro-investment and pro-growth policies of both Kirschner administrations (2003-2015), public investment as a percentage of GDP increased dramatically after 2006, recording, on a consistent basis, levels around 3 percent. It is also notable that during this period *private* investment spending as a proportion of GDP rose by one full percentage point during the 2011-2017 period as compared to 2003-2010.

The basic idea is that public investments in highways, bridges, sewerage systems, water supplies, and education and health services often generate substantial positive spillover benefits for the private sector by reducing the direct (and indirect) costs of producing, transporting, and delivering goods and services to consumers. (See Albala-Bertrand and Mamatzakis, 2001: Aschauer, 1989; Cardoso, 1993; Devarajan et al., 1996; Ramirez, 2010; Serven and Solimano, 1993; and Weisbrot and Merling, 2018) (Note 4). Moreover, cuts in public investment may undermine some or all of the long-term efficiency gains expected from the implementation of market-based, outward-oriented reforms such as privatization of state-owned firms and the liberalization of trade and finance (see Killick, 1995; and Stiglitz, 2012). After all, the newly privatized firms in liberalized (open) markets will need adequate and reliable economic infrastructure in order to produce, transport, and market their goods and services at home and abroad in a cost-effective manner (Note 5).

In view of this, this paper analyses the impact of public investment spending on infrastructure and education (a proxy for human capital) on the economic output and labor productivity of the Argentine economy. The choice of Argentina is warranted for a number of reasons. First, Argentina is a large and strategically important country in Latin America and the second largest economy in South America. This is a situation that promises to continue as a result of the country's participation in the important regional trade agreement named Mercosur and its status as a major regional exporter of maize and soybeans. Second, beginning with the Menem administration (1989-1999) and continuing under the ill-fated administrations of Fernando De La Rua and Duhalde (2000-2002), Argentina pursued a far-reaching neoliberal strategy of economic growth that was heralded by pundits and multilateral organizations as a prototype for the region—at least before the onset of the crisis. However, under both Kirschner administrations (2003-2015), but particularly that of Cristina Fernandez de Kirschner (2007-2015), the Argentine government reversed itself and pursued a more activist set of Keynesian-style policies designed to promote broad-based economic growth with an emphasis on the promotion of public investment and education (Note 6).

Beginning in 2016, under the administration of the right-center president Mauricio Macri (2015-2019), the pendulum swung back to the neoliberal policies of the past. [These policies are likely to intensify under the incoming administration of the recently elected far-right candidate, Javier G. Milei (2023-2027), a self-described "'anarcho-capitalist' libertarian"]. Under the Macri administration, the government adopted a three-year draconian IMF-sponsored program in June 2018 to combat a severe economic and financial crisis that erupted in 2018. The crisis was largely the result of an ill-conceived export tax cut early on in his administration and a large and rapid run-up in foreign borrowing. Moreover, there was an equally rapid and ill-timed removal of currency controls in the midst of rising interest rates in the U.S. which, in turn, fueled massive capital flight, and last but not least, a severe drought that ravaged the country's all-important agro-export sector. The crisis sent the peso reeling more than 70 percent during 2018-20 and, not surprisingly, private investment and industrial output plummeted, unemployment and poverty rose sharply, real wages nose-dived, and interest soared to almost 70 percent in 2019-20! Although beyond the scope of this empirical paper, beginning in 2020 the adverse economic and health effects of the Covid-19 pandemic plummeted the country into yet another devastating economic and social crisis which culminated in the defeat of the Peronist candidate Sergio Massa and the election of Milei in November 2023 (ECLAC, 2023; and Weisbrot and Merling, 2018) (Note 7).

The paper is organized as follows. Section II provides a conceptual framework for incorporating the public or human capital stock in a modified neoclassical production function. The model presented in this section is intended solely to motivate the ensuing discussion since the relevant parameters cannot be estimated directly due to data limitations present in the Argentine case. Next, the paper introduces a rough empirical counterpart to the model presented in the previous section, and discusses the nature and limitations of the data used in this study. Section IV presents unit root and cointegration tests, while Section V reports FMOLS and DOLS long-run estimates for the modified production and labor productivity functions. The concluding section summarizes the major results and discusses potential avenues for future research on this important topic.

2. Conceptual Framework.

On the supply side, the positive externalities generated by additions to the public (or human) capital stock can be formalized by incorporating them in an augmented Cobb-Douglas production function of the following form (see Barro and Sala-I-Martin, 1995; and De Mello 1997):

$$Y = A\,\varphi\,[L, K_p\,, E] = A\,L^{\alpha}\,K^{\beta}\,E^{(1-\alpha-\beta)} \qquad (1)$$

where Y is real output, K_p is the private capital stock, L is labor, and E denotes the externality generated by additions of the public capital stock or human capital stock (α and β are the shares of domestic labor and private capital respectively, and A captures the efficiency of production. Initially, it is assumed that α and β are less than one, such that there are diminishing returns to the labor and capital inputs.

The externality, E, can be represented by a Cobb -Douglas function of the type:

$$E = [L, K_p, K_g^{\gamma}]^{\theta} \qquad (2)$$

where γ and θ are, respectively, the marginal and the intertemporal elasticities of substitution between private and public (human) capital. Let γ >0, such that a larger stock of public (or human) capital generates a positive externality to the economy; i.e., knowledge or technological progress is an accidental by-product of capital investment by relatively small firms in the form of education or public investment spending. If θ > 0, intertemporal complementarity prevails and, if θ < 0, additions to stock of public (human) capital crowd out private capital over time [see Jones, 2011].

Combining equations (1) and (2), we obtain,

$$Y = A\,L^{\alpha+\theta(1-\alpha-\beta)}\,K_p^{\beta+\theta(1-\alpha-\beta)}\,K_g^{\gamma\theta(1-\alpha-\beta)} \qquad (3)$$

A standard growth accounting equation can be derived by taking logarithms and time derivatives of equation (3) to generate the following dynamic production function:

$$g_y = g_A + [\alpha + \theta(1\text{-}\alpha\text{-}\beta)]g_L + [\beta + \theta(1\text{-}\alpha\text{-}\beta)]g_{Kp} + [\gamma\theta(1\text{-}\alpha\text{-}\beta)]g_{Kg} \qquad (4)$$

2.1 Preferences

The demand side of the economy can be included into the model via the following intertemporal utility maximization framework:

$$\text{Max } u(t) = \int_{\circ}^{\infty} u(c(t))\, e^{-\rho t}\, L(t)\, dt \qquad (5)$$

$$\text{s. t. } k_p = A\,k_p^{\beta+\theta(1-\beta)}\,k_g^{\gamma\theta(1-\beta)} - c - \delta k_p, \text{ and } k_p(0) \geq 0.$$

where, for convenience, lower-case letters are defined in per capita terms and ρ is the discount rate, L(t) is the size of the family, c(t) is per capita consumption, and δ represents the rate of depreciation. For convenience, the initial population is normalized to 1 so that the analysis in aggregate and per capita terms is the same. The instantaneous utility function of the representative consumer is assumed to exhibit constant relative risk and can be written in the following general form:

$$u(c(t)) = (c(t)^{1-\sigma} - 1) / (1-\sigma) \tag{6}$$

σ denotes the relative risk aversion coefficient or the inverse of the elasticity of substitution between current and future consumption; i.e., σ is an index of the representative consumer's willingness to exchange current consumption for future consumption. Letting u(c) = lnc, for simplicity, and solving the standard optimal control problem in equation (5), we obtain the following equation:

$$\dot{c}/c = 1/\sigma\{A[\beta + \theta(1-\beta)]k_p^{\beta + \theta(1-\beta) -1}\, k_g^{\gamma\theta(1-\beta)} - \rho\} \tag{7}$$

Equation (7) can be interpreted as follows in the presence of intertemporal complementarity between public and private capital (i.e., θ > 0): the economy grows at a positive rate whenever the marginal product of capital, net of depreciation, can be kept above the rate of time preference (discount) [the dot over c denotes the rate of change of per capita consumption]. The marginal productivity of private capital, in turn, is augmented and kept above the discount rate by additions to the stock of public or human capital. Finally, the larger the intertemporal elasticity of substitution of current consumption for future consumption, as captured by the inverse of the relative risk coefficient, σ, the higher the rate of growth of the economy. Put differently, the sacrifice of current consumption is less costly to the representative consumer when present and future consumption are good substitutes (see Jones, 2011).

A more intuitive understanding of equation (7) can be obtained by assuming that private capital exhibits constant returns to scale (i.e., β + θ(1-β) = 1, which implies that θ =1). Equation (7) now can written as follows:

$$\dot{c}/c = 1/\sigma\,(Ak_g^{\gamma(1-\beta)} - \rho) \tag{8}$$

Equation (8) states that the long-run growth rate of the economy depends on the degree of relative risk aversion, the rate of time preference, the productivity of private capital, and the degree of complementarity, γ, between stocks of private and public capital, but not on the size of the private capital stock *per se*. In other words, increases in the stock of public capital lead to temporary increases in the growth rate of output in the presence of diminishing returns to private capital. Equation (8) can also be modified to accommodate endogenous growth models by assuming that the elasticity of output growth with respect to capital is unity, i.e., γ(1-β) = 1 (the so-called "AK" class of models). In the absence of diminishing returns to private capital, permanent increases in the stock of public capital generate permanent increases in the output growth rate of the economy (see Barro and Sala-I-Martin, 1995; and Jones, 2011).

3. Empirical Model

Due to the poor quality or paucity of data for key variables over a long period of time, it is often not possible to generate direct estimates of the variables in equations (3) and (4). Instead, researchers have used proxies for variables such as the labor force and/or the stocks of private and public capital such as population data rather than labor force data, or substituted investment data (as a proportion of GDP) for capital stock data (see Aschauer, 1989; Cardoso, 1993; Lin, 1994; and Ramirez, 2010). However, models that use these proxies have to impose unduly restrictive assumptions (e.g., such as a fixed capital-output ratio) or unrealistic assumptions (a constant labor force participation rate) that can generate both misspecified relationships and significant measurement errors.

We are fortunate, in the case of Argentina, to have labor force data going as far back as 1960, but we do not have consistent estimates of the public and private capital stock series, or for that matter, reliable estimates of the rate of depreciation from which such a series could be generated. Researchers in the field of economic development have tried to address this problem by estimating proxies such as the ratio of public and private investment spending to gross domestic product. Following their lead, this study estimates a rough empirical counterpart of the production function in equation (4) for the 1960-2019 period. The most general formulation of the modified production function is given below,

$$y = \alpha + \beta_1 l + \beta_2(i_p) + \beta_3(i_g) + \beta_4(e_g) + \beta_5 D_1 + \beta_6 D_2 + \varepsilon \qquad (9)$$

lower case letters denote natural logarithms; y is real GDP (1993 pesos); l, as indicated above, refers to the labor force (thousands occupied); i_p denotes the ratio of private investment to GDP, while i_g represents public investment spending on economic infrastructure as a proportion of GDP, viz., roads, bridges, and ports (Note 8) --it therefore excludes investment expenditures by state-owned enterprises which are more likely to crowd out private investment spending and output; e_g is real government expenditures on education as a proportion of GDP, and it is a proxy for human capital—a data series for secondary education expenditures (se_g) as a proportion of GDP was also available for the period under review and it was utilized to determine whether the impact of education expenditures would vary by the type of expenditure. This variable is expected to affect output (and labor productivity) in a positive fashion. D_1 is a dummy variable that takes a value of one for the crisis years, and 0 otherwise, while D_2 equals 1 for the impact of the currency board, and 0 otherwise. Finally, ε is a normally distributed error term.

3.1 Data

The data used in this study were obtained from official government sources such as the Direccion Nacional de Politicas Macroeconomica, Ministerio de Economia y Produccion (Ministry of Economy and Production, various issues) and the Instituto Nacional De Estadistica y Censos de la Republica Argentina (National Institute of Statistics and Census of Argentina). Other relevant economic data have been obtained from the World Development Indicators for Argentina, World Bank (2021); ECLAC, Statistical Yearbook for Latin America and the Caribbean, 2022, and the International Finance Corporation (Everhart and Sumlinski, 2001).

The dependent variable is also estimated as a labor productivity variable by taking the natural log of the ratio of real GDP to the labor force. Defining the dependent variable in this manner reverses the expected sign of the labor variable because of diminishing returns to the labor input. The sign of β_1 is anticipated to be positive in the GDP formulation while, as indicated above, it is expected to be negative in the labor productivity specification. β_2 is expected to be positive, while the sign of β_3 can be positive or negative depending on whether increases in public investment complement or substitute for private capital formation. The sign of β_4 is also expected to be positive because government expenditures on education (a proxy for human capital) may directly or indirectly crowd in private investment expenditures and thus affect output (and labor productivity) in a positive fashion. B_5 is anticipated to be negative for obvious reasons, while β_6 is expected to be positive.

4. Unit Roots, Structural Breaks, and Cointegration Analysis

Initially, conventional unit root tests (without a structural break) were undertaken for the variables in question given that it is well-known that macro time series data tend to exhibit a deterministic and/or stochastic trend that renders them non-stationary; i.e., the variables have means, variances, and covariances that are not time invariant (see Granger and Newbold 1974). This study tested the variables in question for a unit root (non-stationarity) by using an Augmented Dickey-Fuller test (ADF) with a lag length automatically determined by the Schwarz Information Criterion (SIC).

Before reporting the unit root tests, it is important to acknowledge that when dealing with historical time series data for developing countries such as Argentina or Chile, investigators are often constrained by the relatively small number of time series observations (usually in annual terms). This is the case in this study where the total sample size, at 60 annual observations, is slightly above the threshold level of 50 recommended by Granger and Newbold, which may compromise the power of the unit root (and cointegration) tests—not to mention distort the size or significance of the tests as well (see Charemza and Deadman, 1997). However, a growing literature contends that the power of unit root (and cointegration) tests depends on the length or *time span* of the data rather than the number of observations per se. That is, for a given sample size n, the power of the test is greater when the time span is longer. Thus, unit root or cointegration tests based on 60 observations over 60 years have considerable more power than those based on 100 observations over 100 days (see Bahnam-Oskooee 1996; and Shiller and Perron, 1985) (Note 9).

Table 2A presents the results of running an ADF test (one lag) for the variables in both level and differenced form under the assumption of a stochastic trend only, i.e., the test is run with a constant term and no time trend (Note 10). It can be readily seen that all the variables in level form are nonstationary; i.e., they appear to follow a random walk with (positive) drift (see Nelson and Plosser, 1982). In the case of first differences, however, the null hypothesis of non-stationarity is rejected for all variables at least at the 5 percent level (Note 11). Thus, the evidence presented suggests that the variables in question follow primarily a stochastic trend as opposed to a deterministic one, although there is the possibility that for given sub-periods they follow a mixed process (Note 12).

Table 2A. Unit Root Tests for Stationarity, Sample Period 1960-2019

Variables	Levels	First Difference	5% Critical Value[1]	1% Critical Value
ln(Y)	-0.48	-5.81**	-2.92	-3.56
ln(Y/L)	-2.21	-5.53**	-2.92	-3.56
lnL	0.72	-6.46**	-2.92	-3.56
lnIp	-0.43	-6.47**	-2.92	-3.56
lnIg	-1.55	-7.10**	-2.92	-3.56
lneg	-1.07	-6.53**	-2.92	-3.56
lnseg	-0.33	-5.64**	-2.92	-3.56

[1]MacKinnon critical values for rejection of hypothesis of a unit root. *Significant at the 5 percent level; **Significant at the 1 percent level.

Table 2B. KPSS (LM) No Unit Root Tests for Stationarity with constant, Sample Period 1960-2019

Variables	Levels	First Difference	5% Critical Value
lnY	0.869*	0.083	0.463
ln(Y/L)	0.666*	0.132	0.463
lnL	0.892*	0.205	0.463
lnIp	0.854*	0.152	0.463
lnIg	0.654*	0.094	0.463
Lneg	0.699*	0.066	0.463
Lnseg	0.503*	0.185	0.463

Asymptotic critical values for rejection (LM-Stat.).*denotes significance at 5 percent level.

In view of the relatively low power of the ADF unit root tests when the data generating process is stationary but with a root close to the unit root, Table 2B presents the results of running a confirmatory KPSS stationarity test (Kwaitkowski et al., 1992). This test has a no unit root (stationary) null hypothesis, thus reversing the null and alternative hypotheses under the Dickey Fuller test. The reported results in both level and differenced form under the assumption of a constant are, by and large, consistent with those reported in Table 2A. For example, the null hypothesis of no unit root can be rejected for all the variables in level form at the 5 percent level of significance; i.e., they appear to follow a random walk with (positive) drift. In the case of first differences, however, the null hypothesis of stationarity cannot be rejected for all variables at least at the 5 percent level. Thus, the evidence presented suggests that the variables in question follow primarily a stochastic trend as opposed to a deterministic one, although there is a possibility that for given sub-periods they follow a mixed process.

4.1 Single-Break Unit Root Test

The conventional results reported in Table 2 may be misleading because the power of the ADF test may be significantly reduced when the stationary alternative is true and a structural break is ignored (see Perron, 1988); that is, the investigator may erroneously conclude that there is a unit root in the relevant series. In order to test for an unknown one-time break in the data, Zivot and Andrews (1992) developed a data dependent algorithm that regards each data point as a potential break-date and runs a regression for every possible break-date sequentially. The test involves running three regressions (models): model A which allows for a one-time change in the intercept of the series; model B which permits a one-time change in the slope of the trend function; and model C which combines a one-time structural break in the intercept and trend (Waheed et. al., 2006). Following the lead of Perron, most investigators report estimates for either models A and C, but in a relatively recent study Sen (2003) has shown that the loss in test power (1-β) is considerable when the correct model is C and researchers erroneously assume that the break-point occurs according to model A. On the other hand, the loss of power is minimal if the break date is correctly characterized by model A but investigators erroneously use model C. In view of this, Table 3 reports the Zivot-Andrews (ZA) one-break unit root test results for model C in level form along with the endogenously determined one-time break date for each time series.

Table 3. Zivot-Andrews One-break Unit Root Test, Sample Period 1960-2019

Variables	Levels	Break Year	5% Critical Value	1% Critical Value
Ln(Y)	-3.27	1981	-5.08	-5.57
In(Y/L)	-3.81	1985	-5.08	-5.57
lnL	-4.32	1995	-5.08	-5.57
lnIp	-4.49	1981	-5.08	-5.57
lnIg	-4.62	1991	-5.08	-5.57
lneg	-4.02	1992	-5.08	-5.57
lnseg	-2.58	2010	-5.08	-5.57

Estimations undertaken with Eviews 11.0.

The estimates reported in Table 3 for the series in level form are consistent with those in Table 2. For all of the series in question, Table 3 shows that the null hypothesis with a structural break in both the intercept and the trend cannot be rejected at the 5 percent level of significance (Note 13). In addition, the Z-A test identifies endogenously the single most significant structural break in every time series. In view of space constraints, Figure 1 below shows the endogenously determined break-date for the labor productivity, In(Y/L), series.

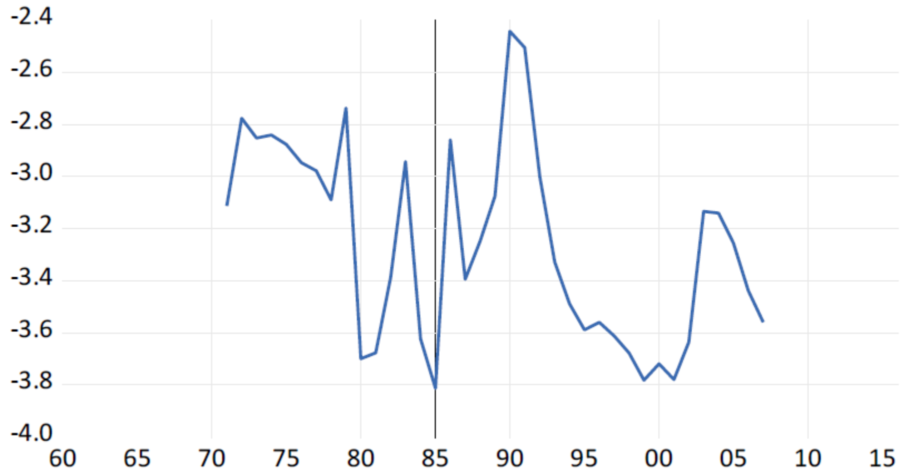

Figure 1. Zivot-Andrew Breakpoints

4.2 Two-Break Unit Root Test

So far, only tests for the presence of a single endogenously determined structural break have been undertaken. Lee and Strazicich (2003) have developed a two-break minimum Lagrange Multiplier (LM) unit root test that shows that assuming erroneously that there is one structural break in the data when, in fact, there are two leads to a further loss of power. Moreover, the LM unit root test developed by Lee and Strazicich (LS) enables the investigator to properly test for structural breaks under *both* the null and alternative hypotheses, thus eliminating size distortions that lead to the over-rejection of the null hypothesis of a unit root. This study therefore performed two-break unit root tests for all the variables in level form utilizing the LSUNIT (lags =1, model=crash, breaks=2) command in Rats 9.2. It was determined that the null hypothesis of a unit root under two endogenously determined structural (intercept) breaks could not be rejected at either the 1 or 5 percent level of significance (with the possible exception of the education variable which is marginally significant at the 5 percent level, see Table 4 below). The (LS) test was also performed with two endogenously determined structural intercept *and* trend breaks (lags =1, model = break, breaks=2) and the results are consistent with those reported for the intercept breaks (the education variable is no longer significant at the 5 percent level). (These results are available upon request.) Thus, the more powerful (LS) unit root test strongly suggests that all the included variables are I(1), by and large, consistent with the reported estimates for the (ZA) procedure. To save space, the (LS) Two-break (Bs) unit root results for the log of real GDP (Y), the log of productivity (Y/L), the log of the ratio of private capital formation to GDP (i_p), the log of the ratio of public investment to GDP, and the log of the ratio of educational expenditures to GDP (e_g) are reported in Table 4 below.

Table 4. Lee-Strazicich Two-Break Unit Root Test, 1960-2019

Variable	Coefficients	T-ratios	1% cv	5% cv
SY(-1)	-0.179	-2.66	-4.073	-3.563
Constant	0.032	5.12	---	---
B1: 1977	-0.055	-1.22	---	---
B2: 1984	-0.086	-1.92	---	---
SY/L (-1)	-0.139	-2.37	-4.073	-3.563
Constant	0.019	2.96	---	---
B1: 1977	-0.046	-1.17	---	---
B2: 1984	-0.073	-1.91	---	---
Sip (-1)	-0.343	-3.24	-4.073	-3.563
Constant	0.012	0.05	---	---
B1: 1968	0.359	2.14	---	---
B2: 2002	-0.384	-2.07	---	---
Sig (-1)	-0.251	-2.70	-4.073	-3.563
Constant	-0.058	-1.71	---	---
B1: 1990	-0.887	-3.99	---	---
B2: 2002	-0.455	-1.95	---	---
Seg(-1)	-0.523	-3.22	-4.073	-3.563
Constant	0.015	0.92	---	---
B1: 1984	-0.732	-4.10	---	---
B2: 1991	0.520	2.99	---	---

Notes: The coefficients on the $S_Y(-1)$,$S_{Y/L}(-1)$,$Si_p(-1)$,$Si_g(-1)$, and Se_g are lagged de-trended variables test for the presence of a unit root. B1 and B2 are the endogenously determined breaks in the intercept for the sample period. The estimations were undertaken with Rats 9.2.

4.3 Cointegration Analysis

Given that the variables are integrated of order one, I(1), it is necessary to determine whether there is at least one linear combination of these variables that is I(0). Does there exist a stable and non-spurious (cointegrated) relationship among the regressors in each of the relevant specifications? This was done by first using the cointegration method proposed by Johansen and Juselius (1990). The Johansen method was chosen over the one originally proposed by Engle and Granger (1987) because it is capable of determining the number of cointegrating vectors for any given number of non-stationary series (of the same order); it also has a well-defined limiting distribution (see Harris, 1995).

Table 5. Johansen Cointegration Rank Test (Trace), 1960-2019

<table>
<tr><td colspan="6">A. Series: lnY, lnL, lnIg and lnIp.
Test assumption: No Linear deterministic trend in the data.</td></tr>
<tr><td>0.404</td><td colspan="2">57.359</td><td>54.08</td><td colspan="2">None</td></tr>
<tr><td>0.305</td><td colspan="2">29.378</td><td>35.19</td><td colspan="2">At most 1</td></tr>
<tr><td>0.123</td><td colspan="2">9.706</td><td>20.26</td><td colspan="2">At most 2</td></tr>
<tr><td>0.047</td><td colspan="2">2.614</td><td>9.17</td><td colspan="2">At most 3</td></tr>
<tr><td colspan="6">B. Series: ln(Y/L), lnL, lnIg , and lnIp.
Test assumption: no linear deterministic trend in the data.</td></tr>
<tr><td>Eigenvalue</td><td colspan="2">Likelihood Ratio</td><td>5% Critical Value</td><td colspan="2">No. of CE(s)</td></tr>
<tr><td>0.404</td><td colspan="2">57.359</td><td>54.08</td><td colspan="2">None</td></tr>
<tr><td>0.305</td><td colspan="2">29.378</td><td>35.19</td><td colspan="2">At most 1</td></tr>
<tr><td>0.123</td><td colspan="2">9.706</td><td>20.26</td><td colspan="2">At most 2</td></tr>
<tr><td>0.047</td><td colspan="2">2.614</td><td>9.17</td><td colspan="2">At most 3</td></tr>
<tr><td colspan="6">Normalized Cointegrating Vector; coefficients normalized on ln(Y/L) in parenthesis.</td></tr>
<tr><td>Vector</td><td>ln(Y/L)</td><td>lnL</td><td>lnIg</td><td>lnIp</td><td>Constant</td></tr>
<tr><td>1.</td><td>1.000</td><td>0.147</td><td>-0.055</td><td>-0.271</td><td>-1.247</td></tr>
<tr><td colspan="2"></td><td>(0.078)</td><td>(0.019)</td><td>(0.045)</td><td>(0.440)</td></tr>
</table>

Note: Standard errors are in parenthesis. Estimation undertaken with Eviews 11.0

Table 5 below shows that the Johansen test for both the output and labor productivity equations show that the null hypothesis of no cointegrating vector can be rejected at least at the five percent level; i.e., there exists a unique linear combination of the I(1) variables that links them in a stable and long-run relationship (Note 14). The signs of the cointegrating equation are reversed because of the normalization process and they suggest that, in the long run, the private and government investment variables have a positive and highly significant effect on Argentine labor productivity. The relatively high private capital (investment) elasticity reported in Table 4 is consistent with the extant empirical literature for developing (and developed) countries, and may be explained by educational externalities in the form of better managerial know-how and the transfer of superior technology that “inflate” the private investment elasticity estimate by a positive factor θ (DeMello, Jr., 1997). For example, a *ceteris paribus* 10 percent increase in the ratio of private investment to GDP raises output per worker by an estimated 5.6 percent in the long run. Admittedly, the relatively high coefficient for the private investment variable may be due to measurement error, omitted variables such as human capital, and/or simultaneity bias.

4.4 Cointegration Analysis with Structural Breaks

Before turning to FMOLS and DOLS estimators, it should be noted that the cointegrating test performed in this study does not allow for structural breaks in the sample period, whether level (intercept) shifts or regime (intercept and slope) shifts. However, Gregory and Hansen (1996) have shown that ignoring these breaks reduces the power of conventional cointegration tests similar to conventional unit root tests and, if anything, should lead to a failure to reject the null hypothesis of no cointegrating vector, which is clearly not the case in the present study. This study therefore undertook a confirmatory G-H cointegration test with level *and* regime shifts and the results are consistent with the Johansen test. The G-H test with endogenously determined level (intercept) shift generated a minimum ADF* stat. = -6.027 [break point=2005] which is smaller than the tabulated 5 % critical value [-6.05 (1%); -5.57(5%); -5.33(10%)] reported by Gregory and Hansen. Thus, the null hypothesis of no cointegration with endogenously determined level break is rejected at the 5 percent level of significance. A regime test (intercept and slope) was also performed and the minimum ADF* stat. = -6.312 [break point= 2004] was essentially the same as the tabulated 5% critical value [-6.890 (1%); -6.320 (5%); -6.16 (10%)], thus leading to rejection of the null hypothesis once again. It should be noted that the break dates are found by estimating the cointegrating relationship for all possible break dates in the sample period. The Rats program selects the break date where the modified [trimmed] ADF* = inf ADF (τ) test statistic is at its minimum. The optimal lags for the G-H test were determined via the AIC and general–to-specific criteria.

5. FMOLS and DOLS Estimators

Having shown that the individual variables in question are I(1), and that there is at least one unique linear combination of these non-stationary variables (in level form) that is stationary, eq. (1) can be estimated via long-run cointegration regression models. These include the fully modified ordinary least squares (FMOLS) proposed by Phillips and Hansen (1990) and the dynamic ordinary least squares (DOLS) model proposed by Stock and Watson (1993). In this study we utilize both methodologies because these estimators are extremely consistent even in the presence of both endogeneity and serial correlation of any order. The FMOLS estimator employs a non-parametric correction to eliminate the problems caused by the long-run correlation between the cointegrating equation and stochastic regressor innovations, and thus generates asymptotically unbiased and fully efficient estimates. DOLS, on the other hand, uses leads and lagged differences of the explanatory variables to control for the endogenous feedback (Saikkonen, 1991). That is, the DOLS estimator uses a parametric adjustment to the errors in order to obtain an unbiased estimator of the long-run parameters. While DOLS and FMOLS solve the problem of endogeneity and eliminate small sample finite bias, the application of the FMOLS approach essentially requires that all variables must have the same order of integration and that the regressors must not appear as co-integrated; DOLS, on the other hand, generates estimates of the long-run parameters irrespective of the order of integration and the existence or absence of cointegration. In this connection, Kao and Chiang (2000) show that, on the basis of Monte Carlo simulations, DOLS outperforms FMOLS estimators in terms of finite sample biases. In view of this, this paper utilizes the DOLS approach to verify the estimates obtained via the FMOLS estimator. The paper also undertakes stability tests on the parameters in question via a test developed by Hansen (1992).

The long-run estimates for the various models estimated (with and without dummy variables to ensure robustness) are presented in Tables 6 and 7. Turning first to the FMOLS results for economic output in Table 6, it can be seen from Eqs. 1 (without the dummies) and 2 (with dummies) that most of the variables in question have their anticipated signs and are significant at the 5 percent level. For example, in Eq. 1, a one percent increase in the labor force (InL) increases the level of the real GDP (InY) in the long run by almost 1 percent, while a one percent increase in the ratio of private investment to GDP (InI_p) generates a long-run rise of 0.18 percent in the level of real GDP, ceteris paribus. Insofar as the impact of education expenditures (Ine_g) are concerned, the estimates for the equations with and without the dummy variables suggest that they have a positive and significant effect on economic output in the long run (comparable to that of private investment). The estimates suggest that the impact of education expenditures ($lnse_g$) is even greater if they are channelled to secondary education (completion of a high school degree) as shown in eq. 3. Similarly, a 1 percent increase in the ratio of public investment to GDP (InI_g) results in a 0.09 percent rise in real GDP in the DOLS specification (see Eq.4). Turning to the dummy variables, the estimates suggest that they have the anticipated signs and are significant in Eqs. 2 and 3. For example, the sign on D_1 suggests that economic and/or political crises have had a negative and significant effect on real GDP, while the sign for D_2 indicates that the Convertibility Plan had a positive and somewhat significant effect on real GDP.

Table 6. Argentina: FMOLS and DOLS estimates, 1960-2019 (Dependent variable =lnY_t)

	FMOLS			**DOLS**		
Variables	Eq.(1)	Eq. (2)	Eq. (3)	Eq.(4)	Eq.(5)	Eq. (6)
Constant	3.71 (15.1)**	3.68 (15.6)**	1.65 (5.51)**	3.93 (11.0)**	3.83 (17.2)**	1.07 (3.85)**
lnL	0.98 (33.9)**	0.99 (35.7)**	1.22 (39.1)**	0.96 (22.6)**	0.98 (27.2)**	1.27 (45.3)**
lnIp	0.17 (5.90)**	0.18 (6.09)**	0.20 (4.44)**	0.18 (3.25)**	0.18 (4.62)**	0.15 (2.30)**
lnIg	0.08 (9.02)**	0.08 (9.05)**	0.05 (3.34)**	0.09 (6.94)**	0.09 (9.40)**	0.06 (3.03)**
Ineg	0.15 (9.54)**	0.15 (9.72)**	---	0.17 (7.35)**	0.17 (9.65)**	---
Inseg	---	---	0.63 (3.11)**	---	---	1.07 (2.67)**
D1	---	-0.02 (-1.90)**	-0.04 (-2.54)**	---	-0.03 (-3.11)**	-0.03 (-1.83)**

D2	---	0.04 (1.98)**	0.03 (1.50)*	---	0.03 (2.11)**	0.05 (2.37)**
Adj.R^2	0.98	0.98	0.97	0.99	0.99	0.97
S.E.	0.04	0.04	0.05	0.03	0.03	0.05
L.R. var.	0.001	0.001	0.002	0.001	0.000	0.002

t-ratios in parenthesis. *Significant at the 10% level; **significant at the 5% level. L.R. var. = Long-run variance.

The FMOLS estimates reported in Table 6 are corroborated by the more robust DOLS estimates for the equations with and without dummies (Eqs.4-6). For example, a one percent increase in education expenditures as a percentage of GDP (Ine_g) increases the level of real GDP in the long run by 0.17 percent in Eq. 4. Again, the effect is even greater if those expenditures are directed towards secondary education as shown in Eq. 6. Moreover, a quick perusal of the standard error and long-run variance for the reported DOLS equations suggests that they are a better fit to the data. Hansen stability tests for all equations (available upon request) suggest that the null hypothesis of parameter stability cannot be rejected for any of the reported equations (p-values > 0.2). The residuals for all of the reported equations were also checked for normality via the Jarque-Bera test, and the null hypothesis that the residuals are normally distributed could not be rejected for any of the reported equations at the 5 percent level (available upon request).

Table 7. Argentina: FMOLS and DOLS estimates, 1960-2019 [Dependent variable = ln(Y/L)t]

	FMOLS			**DOLS**		
Variables	Eq. (1)	Eq. (2)	Eq. (3)	Eq (4)	Eq (5)	Eq. (6)
Constant	3.71 (15.1)**	3.68 (15.6)**	1.65 (5.51)**	3.93 (16.6)**	3.84 (17.2)**	1.07 (3.85)**
InL	- 0.02 (-0.44)	- 0.07 (-0.30)	-0.22 (-6.97)**	-0.03 (-1.20)	-0.02 (-0.79)	-0.27 (-9.74)**
InIp	0.17 (5.90)**	0.18 (6.09)**	0.20 (4.44)**	0.18 (4.51)**	0.18 (4.62)**	0.15 (2.30)**
InIg	0.08 (9.02)**	0.08 (9.05)**	0.05 (3.34)**	0.09 (9.84)**	0.09 (10.3)**	0.06 (3.04)**
Ineg	0.15 (9.54)**	0.15 (9.72)**	---	0.17 (12.3)**	0.17 (11.7)**	---
lnseg	---	---	0.63 (3.10)**	---	---	1.07 (2.68)**
D1	---	-0.02 (-1.89)**	-0.04 (-2.54)**	---	-0.03 (-3.11)**	-0.03 (-1.83)**
D2	---	0.04 (1.98)**	0.03 (1.50)*	---	0.03 (2.11)**	0.06 (2.37)**
Adj.R2	0.79	0.80	0.58	0.85	0.86	0.74
S.E.	0.04	0.04	0.05	0.03	0.02	0.05
L.R. var.	0.002	0.001	0.003	0.001	0.000	0.002

t-ratios in parenthesis. *Significant at the 10% level; **significant at the 5% level. L.R. var. = Long-run variance.

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Table 7 below reports FMOLS and DOLS results for the labor productivity [In(Y/L)] function. As can be seen from the estimates without dummies in Eqs. 1 and 4, the included regressors have their anticipated signs and, except for the labor force variable, are significant at the one percent level. For example, in Eq. 1 a one percent increase in the ratio of private investment to GDP generates an increase of 0.17-0.18 percent in labor productivity in the long run, while a similar increase in the ratio of public investment to GDP results in an increase of 0.08-09 percent (for both the FMOLS and DOLS specifications). Again, education expenditures have a relatively large (0.15-0.17) and significant long-run effect on labor productivity in both the FMOLS and DOLS specifications (comparable to the impact of private investment). This is particularly the case when these expenditures are channeled towards secondary education where the effect ranges from 0.63-1.07 percent (see Eqs.3 and 6). The dummy variables in Eqs. 2 and 5 also have their anticipated signs and are significant in both methodologies. The adjusted R^2 s for the labor productivity functions are somewhat lower than those for economic output, but the standard errors and long-run variances are comparable (if not lower) than those reported in Table 6. Finally, as in the case for economic output, stability and normality tests for these labor productivity equations did not lead to a rejection of the null hypothesis at conventional levels of significance.

6. Conclusion

This paper presented a simple model that explicitly includes the impact of the public (or human) capital stock on both the supply and demand sides of a hypothetical economy. The discussion showed that if significant complementarities are present (i.e., if a positive externality is present), then diminishing returns to the private inputs can be prevented or postponed indefinitely. The conceptual model laid the groundwork for the empirical analysis of the Argentine economy during the 1960-2019 period in Sections 3 and 4. Several key findings were obtained.

First, Zivot-Andrews and Lee-Stratizicich unit root tests in the presence of single and two-time breaks indicate, respectively, that the null hypothesis of non-stationarity cannot be rejected for the relevant series in level form but can be rejected in first differences. This represents a significant contribution to the extant literature which does not address the low power of conventional unit root tests in the presence of structural breaks. Second, the Johansen and Gregory-Hansen cointegration tests revealed that the null hypothesis of no cointegration can be rejected at the five percent level, thus suggesting that the I(1) variables have a unique and stable relationship that keeps them in proportion to one another in the long run even in the presence of an endogenously determined structural regime (intercept and trend) break. This is a highly important finding because previous empirical studies have applied the OLS method directly to nonstationary variables in level form, thus generating spurious or miss-specified regressions. Third, this paper is the *first* to report DOLS and FMOLS estimates which suggest that private and public investment spending as a percent of GDP, as well as education spending (as a percent of GDP), have a positive and significant long-run effect on both the level of real GDP and labor productivity. Hansen stability tests for all equations suggest that the parameters are stable over the period in question. The qualitative variables designed to pick up the effect of economic crises and the impact of

Convertibility Plan have their expected signs and they are statistically significant.

Although suggestive, the positive long-run estimates reported in this study for public expenditures are by no means conclusive and need to be supported by micro-based case studies and/or sectorial (regional) studies. As more disaggregated data becomes available on a regional or sectorial basis, it will be possible to conduct panel cointegration studies to determine whether public investment and education spending have greater positive (or negative) effects on economic output and labor productivity in certain regions or sectors of the Argentine economy.

From a policy standpoint, the findings in this paper are important because they suggest that cash-strapped governments of Latin America, particularly the newly elected Argentine one, should be mindful of slashing government expenditures across-the-board without focusing on the *composition* of that spending. For a given level or rate of government spending, policymakers can maximize the growth potential of their economies by directing scarce resources to investments in economic and social infrastructure, particularly education at the primary and secondary levels. These investments, through a positive externality effect, are likely to increase the marginal productivity of the private inputs directly (as well as indirectly), thereby increasing private investment, output, and labor productivity.

Competing interests

Sample: The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Informed consent

Obtained.

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The Publication Ethics Committee of the Macrothink Institute.

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The data that support the findings of this study are available on request from the corresponding author. The data are not publicly available due to privacy or ethical restrictions.

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No additional data are available.

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Notes

Note 1. Argentina's privatization, liberalization, and deregulation program is discussed and analyzed in Baer et. al. (2002) and Weisbrot et. al. (2002).

Note 2. Weisbrot (2002) observes that in 2002 the IMF demanded that the Argentine government enact spending cuts of 10 percent across-the-board, in addition to a 30 percent reduction in outlays for goods and services and a 13 percent cut in salary and pensions for government employees (p. 13).

Note 3. It should be mentioned that during the 1990s only a handful of countries, notably Chile and Costa Rica, have managed to record investment ratios comparable to those before the onset of the debt crisis in the early 1980s.

Note 4. This paper only addresses the direct output effects of increasing public investment spending. It ignores the impact of public investment spending on the relative prices that private firms face for key inputs and services.

Note 5. Of course, the public sector need not provide these public goods directly; the goods can be contracted out to the private sector in accordance with government regulations and guidelines. In fact, many governments in Latin America (including Argentina) have awarded concessions to private firms to produce and provide quasi-public goods and services. However, if the monitoring or supervision cost of outsourcing public works projects is high, then the bias in favour of privatizing these types of expenditures is removed (see Ramirez, 2010).

Note 6. Nestor Kirschner, a member of the Justicialist (Peronist) Party, served as President of Argentina during the 2003-2007 period, and his wife, Cristina Fernandez de Kirschner, succeeded him as President of Argentina in 2007. Under both Kirschner administrations, the economy grew at average annual rates exceeding 8 percent and levels of poverty and unemployment experienced a dramatic fall from their crisis levels in 2001-02; there also was a significant increase in government spending on housing, health and economic infrastructure, as well as a generous extension of social security coverage and a substantial rise in real wages (see Weisbrot, 2011, pp. 8-12).

Note 7. Estimates by ECLAC (2023) indicate that the growth rate in real GDP fell by 2.6 percent in 2018 and 2.0 percent in 2019, after expanding by 2.8 percent in 2017—it is projected to fall by an additional 9.9 percent in 2020 or 10.8 percent in real per capita terms, partly as a result of the adverse economic effects of the Covid-19 pandemic. The data also indicate that the rate of inflation surged to an estimated 52.9 percent in 2019 from a rate of 25 percent in 2017, and that the unemployment rate rose by 3.5 percentage points to an estimated 11.5 percent in 2020 compared to 2017. Economic activity is expected to continue to fall in 2020 and beyond because of the combined effects of Covid-19 and the US56.3bn stand-by agreement with the IMF calls for substantial real cuts in in public expenditures on subsidies (energy and transportation), public sector wages, and public investment on economic infrastructure. For further details, see ECLAC (2022), Statistical Annex, pp. 231-266; and Weisbrot and Merling (2018, pp. 1-19).

Note 8. It should be noted that the reported government investment data contains a portion that is devoted to health and education expenditures, and should be treated separately as public (human) capital investment. At the risk of some overlap, I have included in this study a separate proxy for human capital, viz., government expenditures on education as a proportion of GDP. There is also data for enrollment ratios at the primary, secondary and tertiary education levels, but the data was significantly more limited and sporadic in nature and had to be interpolated for some years. See World Development indicators, published by World Bank (2021).

Note 9. For example, Hakkio and Rush (1991) contend that in nearly non-stationary time series "the frequency of observation plays a very minor role" in cointegration [and unit root] analysis because "cointegraton is a long-run property, and thus we often need long spans of data to properly test it" (p. 579). Similarly, Bahmani-Oskooee (1996) observes that in cointegration (and unit root) analysis using annual data over 30 years "is as good as using quarterly data over the same period" (p. 481). To some degree, this addresses the strong analytical and policy inferences drawn from a relatively small sample size.

Note 10. A stochastic trend is one where the random component of the series itself, say variable xt, contributes directly to the long run pattern of the series, either upward or downward. However, in the case of a deterministic trend the deviations from the non-stationary mean over time are quickly corrected. It is also possible for the variable in question to display both a stochastic and deterministic trend process over time. For further details see Charemza and Deadman, (1997, pp. 84-92).

Note 11. The order of the lag length was determined by applying both the Akaike Information Criterion (AIC) and the Schwarz Bayesian Information Criterion (SIC). For all the variables in this study, the ADF tests with one lag showed the lowest value for both the AIC and SIC criteria.

Note 12. This study also performed an ADF test (one lag) on the variables in logarithmic form with a deterministic trend. The results indicate that the null hypothesis of non-stationarity cannot be rejected for any of the variables in level form with a deterministic trend, suggesting that the variables in question do not exhibit a deterministic time trend throughout the period under review.

Note 13. The Z-A one-break point unit root test was also performed for the relevant time series in differenced form under the assumption of model C and the null hypothesis was rejected at the 5 percent level or lower in all cases.

Note 14. The dummy variables were treated as exogenous variables in the cointegration test. The variables in question are also cointegrated with the inclusion of the education variable. The trace statistic in this case is equal to 88.56 and the null hypothesis of no cointegration is rejected at the 5 percent level (critical value= 76.972). The Max-eigen value test also indicates one cointegrating vector [39.977 > 34.80 (5% critical value)].