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Abstract

Using panel data, this paper tests whether public and private capital have a positive and significant effect on aggregate output and labor productivity for Mexico during the 1960-2001 period. The richer information set made possible by the sectorial data enables this study to utilize the methodologically sound "group-mean" Fully Modified Ordinary Least Squares (FMOLS) procedure developed by Pedroni to generate consistent estimates of the relevant panel variables in the cointegrated production (labor productivity) function. The results suggest that, in the long run, changes in the stocks of public and private capital and the economically active population (EAP) have a positive and economically significant effect on output (and labor productivity). The period is also broken down into two sub-periods: 1960-81 (state-led industrialization) and 1982-2001 (neoliberal model). The estimate for the public capital variables clearly shows that it had a relatively more important economic effect during the earlier state-led period.

Keywords: Fully Modified Ordinary Least Squares (FMOLS), Panel Unit Roots, Panel Cointegration Test, Complementarity Hypothesis, Mexican Labor Productivity

JEL Codes: O10, O50, and O40

I. Introduction.

Mexico's relatively successful state-led, inward-oriented strategy of economic growth and development was effectively brought to an end with the onset and aftermath of the debt crisis in the early 1980s. In fact, it is easy to forget in this era of market-based and outward-oriented reforms that under state-led import substitution industrialization (ISI) Mexico enjoyed one of the highest and sustained rates of economic growth per capita in the developing world--during the so-called "miracle years" of the late fifties and sixties. Yet, in less than two decades, Mexico's cash-strapped governments have been forced to dismantle and privatize their massive state-owned sector, and following the country's accession to the GATT in 1986, its economy has been transformed from a heavily protected and highly regulated one to, arguably, one of the most open and unregulated economies of the region. This market-led, outward-oriented process was effectively locked in, both economically and institutionally, with the passage and phased implementation of the NAFTA beginning in 1994.

However, the country's transition from a closed to an open economy has been anything but easy. It has often been marred by a series of economic crises and financial setbacks, most dramatically in 1994-95 with the onset and aftermath of the so-called "peso crisis ," and most recently, with the severe economic slowdown the economy has experienced following the relatively mild U.S. recession in 2001. These economic crises and financial setbacks have led the Mexican government to the involuntary adoption and implementation of several IMF-sponsored adjustment programs which have resulted in sharp cuts in real government spending across the board, the severe contraction and increase in the cost of credit, and last but not least, repeated devaluations of the domestic currency in real terms [see Ibarra, 1996; Gonzalez, 2002; and

Lustig 2001].

Not surprisingly, the economic growth and investment performance of the Mexican economy under market-based reforms has fallen both far short of the initial (and inflated) expectations of its more ideological advocates as well as when compared to its performance under state-led ISI.[1] One possible factor in explaining Mexico's poor growth and investment performance is the sharp fall in public capital spending demanded by the stringent fiscal deficit targets of the various stabilization programs. For example, Figure 1 below shows that overall public investment spending as a proportion of GDP (RG) fell precipitously from 12.1 percent in 1981 to barely 2.2 percent in 2001; the dramatic fall in government investment is further revealed by the fact that average public investment spending as a proportion of GDP for the 1990s stood at just 3.6 percent, which is less than half as much as the level recorded during both the 1980s and 1970s.

From the standpoint of the composition of public spending, Figure 1 shows that public investment channeled to industry (RGI) fell sharply after 1990 (from 3.3 percent of GDP) to only 1.3 percent of GDP in 2001. On the other hand, state investment directed toward the service sector (RGS) rose to over 2 percent of GDP after 1990 (as high as 3.2 percent in 1991) and maintained that level until 1994, after which it fell to levels slightly below those recorded in the industrial sector (1 percent of GDP in 2001). Finally, the share of public investment channeled to the primary sector (RGP) has been historically low since 1970, and with the onset and aftermath of the debt crisis in 1982-83, it has declined to barely 0.12 percent of GDP in 2001.[2]

In recent years, a number of investigators have undertaken (univariate) empirical studies which suggest that the dramatic fall in public capital investment experienced by developing countries such as Mexico is of particular concern because investments in economic and social infrastructure 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 Aschauer, 1989; Albala-Bertrand and Mamatzakis, 2001; Barro, 1990; Cardoso, 1993; Green and Villanueva, 1991; Hermes and Lensink, 2004; Khan and Reinhart, 1990; Moguillansky, 1996 ; Nazmi and Ramirez, 1997; Ram, 1996; Ramirez, 2002]. However, the major problem with several of these studies is that the data set available to test the complementarity hypothesis is often in annual terms and for a limited time period. Thus, even when cointegration tests are performed and error correction models are generated , the reliability of the estimates is questionable because unit root tests have low power when the number of observations is less than fifty as is often the case with univariate (annual) time-series studies.[3]

In light of the above, this paper estimates a pooled model that attempts to determine whether public capital in three major sectors of the Mexican economy has a positive and significant effect on Mexican output (and labor productivity) over the 1960-2001 period? The information contained in the time series data is thus enhanced by the cross-sectional (sectorial) data which makes it possible to *reliably* test whether increases in government investment spending enhance overall output and labor productivity in Mexico. The focus on Mexico is particularly relevant because it is one of the few countries in Latin America and the Caribbean that has reliable and disaggregated time-series data on public and private investment spending on a sectorial basis going as far back as the 1950s. More importantly, perhaps, it also allows policymakers to determine where the effects of public investment spending, if any, are most significant. And since Mexico is a country faced with severe constraints in generating public revenues, any additional information that improves the allocation of scarce resources should prove highly useful to the country's policymakers.

The paper is organized as follows. Section II provides an economic rationale for including the public capital stock as an argument in a modified neoclassical production function, and discusses the empirical methodology to be employed in subsequent sections. Section III pools data for the primary (agricultural and mining), industrial (manufacturing), and service (banking, telecommunications, and energy) sectors and estimates a "stacked" production function (and labor productivity function) over the 1961-2001 period. This section also applies recently developed panel unit root tests to the relevant variables to determine if they are stationary and a panel (and group) cointegration test developed by Pedroni [1999a] is used to determine whether there is a stable long-term relationship among the relevant panel regressors of the modified pooled production (labor productivity) function. In addition, it proceeds to estimate the pooled production (productivity) function via a "group-mean" panel fully modified Ordinary Least Squares (FMOLS) estimator developed by Pedroni [1999b; 2001] which not only generates consistent estimates of the parameters in relatively small samples, but also controls for potential endogeneity of the regressors and serial correlation. This study thus represents an important contribution to the extant literature on the complementarity hypothesis because it addresses the important question of spurious correlation among the variables in the pooled (stacked) model. The last section summarizes the chapter's major findings and offers some policy prescriptions.

II. The Model and Econometric Methodology.

Following the lead of Barth and Cordes [1980] and Aschauer [1989], it is possible to analyze the potential impact of public capital spending on output and the marginal productivity of private capital in Mexico by appealing to the modified neoclassical production function given in equation (1) below.

$$Y = A\, F(L, K_p, K_g) \tag{1}$$

$$F_1, F_2 > 0;\ F_{11}, F_{22} < 0;\ F_{12} > 0;$$

$$F_3 \lesseqgtr 0;\ F_{23} \lesseqgtr 0;\ F_{13} \lesseqgtr 0.$$

where A is an index of multi-factor productivity; Y is the level of real output; L denotes employment; K_p is the stock of private capital; and K_g refers to the public capital stock.

By treating the public capital stock as a separate input in the production function, a ceteris paribus increase in public investment gives rise to three conceptually distinct effects.[4] First, if the public capital stock is productive and complements the private capital stock, a *ceteris paribus* increase in the public capital stock will increase output directly in the same way that an increase in any other factor of production raises output ($F_3 > 0$). Secondly, it will indirectly increase private investment and output by raising the marginal productivity of the private capital stock ($F_{23} > 0$) relative to the real interest rate. Third, it will increase output via its positive impact on the marginal productivity of labor; i.e., by increasing the amount of both private and public capital per worker (F_{12} and $F_{13} > 0$).

Of course, in the case where public and private capital are direct substitutes ($F_{23} < 0$), an increase in public investment by state-owned enterprises in sectors that directly compete with the private sector generates a positive direct effect, but a negative indirect effect that could more than offset it; i.e., when the following condition arises: $[(F_3 + F_{13}) + (F_{23}) + (F_{12})] < 0$. Finally, in the case where private and public capital are independent ($F_{23} = 0$) , a *ceteris paribus* increase in public investment will generate a direct positive effect on output.[5]

Empirical Model

As in Albala-Bertrand and Mamatzakis [2001] and Ramirez [2002], this paper first estimates a pooled Cobb-Douglas production function that includes the stock of public capital as an input of production. In logarithmic form, the model estimated over the 1960-2001 period is given by

$$y_{it} = a_t + bl_{it} + ck_{pit} + dk_{git} \qquad (2)$$

where lower case letters denote logs and the variables are defined as in equation (1) above. a is a technology index, t is time, and the coefficients b, c, and d are elasticities. The data consists of 3 cross-sectional units (primary, industrial, and service sectors), denoted i = 1, . . .3, observations at each of 42 time periods, t= 1, . . . 42, for a total of 126 observations.

Next, assuming that the production process exhibits constant returns to scale for the private inputs but increasing returns over all inputs , private and public, we set b + c =1 while b + c + d >1, so that equation (2) can be estimated as a labor productivity function,

$$y_{it} - l_{it} = a_t + c\,(k_{pit} - l_{it}) + dk_{git} \qquad (3)$$

If, d, the elasticity of output (or labor productivity) with respect to public capital is positive and statistically significant, then the public capital variable is an important determinant of economic growth and labor productivity at the national and sectorial level. Finally, in the more problematic case of constant returns in all three inputs, we set b + c + d = 1, and (3) becomes,

$$y_{it} - l_{it} = a_t + c\,(k_{pit} - l_{it}) + d(k_{git} - l_{it}) \qquad (4)$$

or alternatively (4) can be estimated as a capital productivity function. This study estimated empirical versions of equations (2), (3) and (4) that also included time dummies and qualitative variables.

Data.

The data used in this study were obtained from official government sources such as Nacional Financiera, S.A., *La Economia Mexicana en Cifras* (various issues), the Banco de Mexico, *Informe Anual* (various issues), and INEGI, *Anuario Estadistico de los Estados Unidos Mexicanos* [1998; 2002]. Other relevant economic data have been obtained from OECD, *Economic Surveys: Mexico* [1992; 1995;1998; 2003] and the International Finance Corporation, *Trends in Private Investment in Developing Countries: Statistics for 1970-2000* [2001; 1998]. All of the data are in real terms and expressed in natural logs. For example, y_{it} is the natural log of (pooled) real GDP (in 1970 pesos) and l_{it} refers to the natural log of the(pooled) economically active population (thousands of individuals).[6] The capital stock data was generated using a standard perpetual inventory model of the following form,

$$K_t = K_{t-1} + I_t - \delta K_{t-1} \tag{5}$$

where K_{t-1} is the stock of capital at time t-1, I_t is the flow of gross investment during period t, and δ is the rate at which the capital stock depreciates in period t-1. In this study the initial capital stock was estimated by aggregating over seven years of gross investment (1953-1959), while an estimate of the rate of depreciation (5 percent) was obtained from Reynolds [1971] and Looney[1985].[7]

III. Panel Results.

Pooled time series data, much like uni-variate time series data, tend to exhibit a time trend and are therefore non-stationary; i.e., the variables in question have means, variances, and covariances that are not time invariant. Engle and Granger [1987] argue that the direct application of OLS or GLS to non-stationary data produces regressions that are misspecified or spurious in nature. These regressions tend to produce performance statistics that are inflated in nature, such as high R^2's and t-statistics, which often lead investigators to commit a high frequency of Type I errors [Granger and Newbold, 1974].

In recent years, a number of investigators, notably Levin, Lin and Chu (2002), Breitung (2000), Hadri (1999), and Im, Pesaran an Shin (2003) have developed panel-based unit root tests that are similar to tests carried out on a single series. Interestingly, these investigators have shown that panel unit root tests are more powerful (less likely to commit a Type II error) than unit root tests applied to individual series because the information in the time series is enhanced by that contained in the cross-section data. In addition, in contrast to individual unit root tests which have complicated limiting distributions, panel unit root tests lead to statistics with a normal distribution in the limit [see Baltagi, 2001].

With the exception of the IPS test, all of the aforementioned tests assume that there is a common (identical) unit root process across the relevant cross-sections (referred to in the literature as pooling the residuals along the within-dimension). The LLC and Breitung tests employ a null hypothesis of a unit root using the following basic Augmented Dickey Fuller (ADF) specification:

$$\Delta y_{it} = \alpha y_{it-1} + \Sigma\beta_{ij}\Delta y_{it-j} + X_{it}\delta + \nu_{it} \qquad (6)$$

where y_{it} refers to the pooled variable, X_{it}' represents exogenous variables in the model such as country fixed effects and individual time trends, and ν_{it} refers to the error terms which are assumed to be mutually independent disturbances. As indicated above, it is also assumed that $\alpha=\rho-1$ is identical across the three cross-sections, but the lag order for the difference terms across the three sectors is allowed to vary. By contrast, the less restrictive IPS test (and other widely used tests such as the ADF Fisher Chi-square) estimates a separate ADF regression for each of the three cross sections to allow for individual unit root processes; i.e., ρ_i may vary across cross-sections (referred to in the literature as pooling the residuals along the between-dimension).

Table 1 below reports (summary) panel unit root tests on the relevant variables given in equation (2) above. As can be readily seen, most of the tests (with the exception of the LLC test in one case) fail to reject the unit root null for all the variables in level form, but the tests do reject the null of a unit root in difference form. The table also reports the widely used Hadri-Z test statistic, which, as opposed to the aforementioned tests, uses a null of no unit root. Again, the results of this test are consistent with those of LLC, IPS, and Breitung because it rejects the null in favor of a unit root for the variables in level form. Thus, the evidence suggests that the variables in question do evolve as non-stationary processes and the application of OLS (or GLS) to equations (2), (3) and (4) above will result in biased and inconsistent estimates. It is therefore necessary to turn to panel cointegration techniques in order to determine whether a long-run equilibrium relationship exists among the non-stationary variables in level form.

Panel Cointegration Analysis.

To determine whether a cointegrating relationship exits, the recently developed methodology proposed by Pedroni [1999a] is employed. Basically, it employs four panel statistics and three group panel statistics to test the null hypothesis of no cointegration against the alternative hypothesis of cointegration..[8] In the case of panel statistics, the first-order autoregressive term is assumed to be the same across all the cross sections, while in the case of group panel statistics the parameter is allowed to vary over the cross sections. If the null is rejected in the panel case, then the variables of the production (labor productivity) function are cointegrated for *all* the sectors. On the other hand, if the null is rejected in the group panel case, then cointegration among the relevant variables exists for *at least* one of the sectors.[9]

The statistics are distributed, in the limit, as standard normal variables with a left hand rejection region, with the exception of the variance ratio statistic. Table 2 below presents the aforementioned panel (and group) statistics for equations (2) and (4) along with the respective variance ratios and rho statistics (non-parametric tests). For models (2) and (4) there is strong evidence of panel cointegration according to both the Augmented Dickey Fuller (ADF)-t and Phillips and Perron (non-parametric)-t statistics. (The evidence for model (3) suggests that the null hypothesis of no cointegration cannot be rejected at the 5 percent level of significance, and, given space constraints, is not reported in Table 2 but is available upon request.)

This study also performed an ADF Fisher unit root test proposed by Maddala and Wu [1999] to determine whether the residuals of each of the three cross-sections of equations (2), (3) and (4) exhibit a unit root (see Table 3 below). In this test, the null hypothesis of a unit root in the residuals (no cointegration) for all three cross sections is set against the alternative hypothesis of some cross sections without a unit root (cointegration). The p-values reported in Table 3 for each cross section suggest that a unit root can be rejected at least at the 5 percent level for models (2) and (4), but not for model (3) where the unit root null can only be rejected for the primary sector. Also, the ADF Fisher statistic and the Choi Z-stat. for the stacked residuals of models (2) and (4) indicate that the null hypothesis of non-stationarity is strongly rejected in the case of models (2) and (4) but not model (3).

The finding that the (stacked) residuals of models (2) and (4), including the residuals from the individual cross sections, do not contain a unit root suggests that there exists an equilibrium (stable) relationship that keeps the relevant variables in the pooled production (labor productivity) function in proportion to one another in the long run. This is a highly important finding because often in panel studies investigators unwittingly apply the GLS method to relationships that are non-stationary in nature, thereby generating spurious results.

Fully Modified OLS Analysis.

Having established that there is a linear combination that keeps the pooled variables in proportion to one another in the long run, we can proceed to generate individual long-run estimates for equations (2) and (4). In view of the fact that the OLS estimator is a biased and inconsistent estimator when applied to cointegrated panels, we utilize the "group-mean" panel fully modified OLS estimator (FMOLS) developed by Pedroni [1999b; 2001]. As indicated in Section I, the FMOLS estimator not only generates consistent estimates of the β parameters in relatively small samples, but it controls for the likely endogeneity of the regressors and serial correlation.[10] Formally, the FMOLS estimator for the i-th panel member is given by,

$$\beta_i^* = (X_i'X_i)^{-1} (X_i' y_i^* - T\delta) \qquad (7)$$

where y* is the transformed endogenous variable, δ is a parameter for autocorrelation adjustment, and T is the number of time periods.

Table 4 below presents estimates of the cointegration vectors and t-ratios for models (2) and (4) considered in this study. The basic model was also estimated with common time dummies to deal with potential cross-sectional dependency arising from common shocks, such as the 1982-83 debt crisis and the 1994-95 "peso crisis" (reported in eq.1a for both specifications). The estimates for model (2) suggest that all the variables have a positive and highly significant effect on aggregate output in the long run. For example, the elasticity of output with respect to the private capital variable suggests that a *ceteris paribus* increase of 10 percent in private capital raises output by 3.3 percent in the long run, while a similar increase in the public variable increases output by 2.0 percent. The inclusion of time dummies in the production function does not alter the general results, but it does reduce the magnitude of the coefficient for the labor variable and increases that of the public variable.[11] The basic production function was also estimated with dummy variables D1 and D2 and they are reported in Table 4 as eqs. 2 and 3. D1 equals 1 for the crises years 1976, 1982-83, 1995 and 2001 and 0 otherwise, while D2 equals 1 for the petroleum led expansion of 1978-81 and 0 otherwise. As can be seen from Table 4, D1 has an (expected) negative and statistically significant effect, while D2 is positive and statistically significant. The inclusion (or exclusion) of the dummies does not alter significantly the coefficients for the quantitative variables in the production function.

Turning to the estimates in eq. (1) for the labor productivity function reported in Table 4, they suggest that private capital per worker has a relatively more important and statistically significant effect on labor productivity. For example, the estimates suggest that a *ceteris paribus* 10 percent increase in the private capital stock per worker raises labor productivity by 7.3 percent in the long run, while a similar increase in public capital per worker raises labor productivity by 2 percent . However, when common time dummies are included, the magnitude, if not significance, of the private variable decreases, while both the magnitude and significance of the public variable increases notably.[12] The estimates reported in eqs. (2) and (3) indicate that the dummy variables have the anticipated signs and are highly significant. Again, their inclusion or exclusion does not appear to alter the estimates for the quantitative variables.

Before concluding, Table 5 reports estimates of the cointegration vectors for the production (and labor productivity) functions when the period is partitioned into the following sub-periods: 1960-81 and 1982-2001. The rationale for doing this, as shown in Figure 1, stems from the sharp drop in the public investment ratio as a result of several IMF-sponsored stabilization programs implemented after the onset and aftermath of the debt crisis. In view of this, it is not unreasonable to expect, *a priori,* that the economic significance of the public capital variable should be less in the later sub-period relative to the earlier one. As can be readily seen from the estimates, the magnitude of the coefficient for the public capital variable is significantly smaller in both specifications during the 1982-2001 period and, in eqs. (2) and (4), it is only statistically significant at the 10 percent level. By contrast, during the earlier state-led investment period (1960-81), the public capital variable is both economically and statistically significant in all reported equations. The magnitude of the estimate for the private capital variable also decreases during the later period, but this variable remains highly significant throughout both sub-periods. It should be noted that time dummies were used in the estimation of both sub-periods to allow for cross-sectional dependency (see eqs. 2 and 4 in particular).

IV. Conclusion.

The paper estimated a pooled production (productivity) function for Mexico during the 1960-2001 period which suggests that private capital, public capital, and the economically active population have a positive and significant effect on output and labor productivity. In contrast to previous studies that have examined the complementarity hypothesis, the pooling of data across three sectors-- viz., primary, industrial, and services-- enabled this study to expand the information set and thus generate a more reliable test of this hypothesis. Second, this study tested the panel variables for unit roots and showed that they exhibited a unit root (i.e., they evolved as non-stationary processes). The latter is a highly significant finding because most investigators have applied OLS (or GLS) to non-stationary (panel) variables, thereby generating spurious results. Finally, in contrast to most extant (panel) studies, this paper utilized the methodologically sound FMOLS procedure developed by Pedroni to generate consistent estimates of the relevant panel variables in the cointegrated production (labor productivity) function.

From a policy standpoint, no strong policy recommendations can be made given the partial equilibrium framework of the analysis undertaken in this paper. However, the reported estimates seem to suggest that revenue-constrained governments of Latin America, such as the Mexican one, can improve their economic performance by changing the composition of spending towards economic and social infrastructure and away from collective consumption goods that compete directly with those provided by the private sector. Social overhead investments in roads, bridges, ports, education, and health are likely to increase the marginal productivity of the private inputs directly (as well as indirectly via relative price effects), thereby increasing private investment, output, and labor productivity. If anything, the sharply differing estimates for the public capital variable during the two sub-periods suggests that politically expedient (across-the-board) cuts in public investment spending should be avoided because they may well undermine the long-term efficiency gains anticipated from the recently adopted open economy model of economic development.

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TABLE 1 Pool Unit Root Tests: Individual Effects Estimation.

Variable (levels)	LLC	Breitung	IPS	Hadri
	Method			
y	-2.28*	1.10	-0.92	7.18*
l	-0.95	1.15	0.53	7 .22*
k_p	1.00	-0.22	-1.26	3.58*
k_g	0.16	-0.64	0.91	5.06*
y - l	-1.79*	-0.28	-0.89	4.50*
k_p- l	0.37	-0.56	0.79	2.89*
k_g-l	0.47	-0.44	0.76	2.37*

Variable (differences)	LLC	Breitung	IPS	Hadri
	Method			
Δy	-12.43*	-9.35*	-10.68*	1.45
Δl	-12.05*	-10.21*	-10.19*	-0.63
Δk_p	-4.54*	-6.82*	-6.22*	1.60
Δk_g	-10.93*	-10.27*	-9.54*	-0.46
Δy-Δl	-11.24*	-10.95*	-10.86*	0.26
Δk_p-Δl	-13.25*	-8.18*	-11.41*	0.19
Δk_g-Δl	-13.13*	-8.84*	-11.34*	0.28

Note: LLC= Levin, Lin, Chu (2002), IPS= Im, Pesaran, Shin (2003). The statistics are asymptotically distributed as standard normal with a left hand side rejection area, except on the Hadri test, which is right sided. A * indicates the rejection of the null hypothesis of nonstationarity (LLC, Breitung, IPS) or stationarity (Hadri) at least at the 5 percent level of significance. Total number of observations (NT) ranged between 117 and 126. Estimations undertaken with EViews5.0.

Table 2. Panel Cointegration Tests for Production and Labor Productivity Functions.

A. Production Function	**Panel Statistics**	**Group Statistics**
Variance ratio	**1.598**	**-----**
Rho statistic	**-1.639**	**-1.114**
PP statistic	**-2.156***	**-2.196***
ADF statistic	**-2.185***	**-2.119***

B. Labor Productivity Function	**Panel Statistics**	**Group Statistics**
Variance ratio	**1.650***	**-----**
Rho statistic	**-1.609**	**-0.964**
PP statistic	**-2.164***	**-2.138***
ADF statistic	**-1.800***	**-1.739***

Notes: All reported values are asymptotically distributed as standard normal. The variance ratio test is right-sided, while the other Pedroni tests are left-sided. A * indicates the rejection of the null of unit root or no cointegration at the 0.05 level of significance. NT=126. Estimations undertaken with Rats 6.01.

TABLE 3. ADF Fisher Unit Root Test on Residuals: Individual Effects Estimation.
I. Model 2.

Method	Statistic	Prob
ADF Fisher Chi-square	40.835	0.000
ADF Choi Z-stat	-5.278	0.000

Intermediate ADF Test Results on Residuals

Cross Section	Prob.	Lag	Max Lag	Obs.
1	0.0006	1	3	38
2	0.0044	1	3	38
3	0.0005	1	3	38

II. Model 3.

Method	Statistic	Prob
ADF Fisher Chi-square	12.094	0.059
ADF Choi Z-stat	-1.588	0.056

Intermediate ADF Test Results on Residuals

Cross Section	Prob.	Lag	Max Lag	Obs.
1	0.0304	0	3	39
2	0.5992	0	3	39
3	0.1298	0	3	39

III. Model 4.

Method	Statistic	Prob
ADF Fisher Chi-square	45.213	0.000
ADF Choi Z-stat	-5.647	0.000

Intermediate ADF Test Results on Residuals

Cross Section	Prob.	Lag	Max Lag	Obs.
1	0.0021	1	3	38
2	0.0004	1	3	38

3	0.0002	1	3	38

Note: Cross section (1), (2), and (3) refer, respectively, to the primary, industrial, and service sectors. Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. Automatic selection of lags based on Schwartz Information Criterion: 0-3. Estimations undertaken with EViews5.0.

Table 4. Panel Group FMOLS Results, 1960-2001.

A. Production Function

	FMOLS Regressions			
Variables	(1)	(1a)	(2)	(3)
l	0.68 (9.87)*	0.53 (5.54)*	0.64 (9.30)*	0.70 (9.82)*
k_p	0.31 (4.70)*	0.34 (5.17)*	0.33 (4.79)*	0.27 (3.84)*
k_g	0.20 (2.85)*	0.26 (2.21)*	0.20 (2.97)*	0.22 (3.56)*
D_1	----	----	-0.07 (-3.35)*	-0.06 (-3.52)*
D_2	----	----	----	0.05 (2.98)*

B. Labor Productivity Function

	FMOLS Regressions			
Variables	(1)	(1a)	(2)	(3)
k_p-l	0.73 (8.46)*	0.60 (7.87)*	0.72 (9.21)*	0.77 (9.71)*
k_g-l	0.20 (1.70)*	0.34 (3.72)*	0.20 (1.96)*	0.20 (1.65)*
D_1	----	----	-0.07 (-2.01)*	-0.04 (-2.14)*
D_2	----	----	----	0.05 (3.08)*

Notes: Estimates refer to (fixed-effects) long-run elasticities of output and labor productivity with respect to the relevant regressors. T-ratios are in parenthesis. Equation (1a) in both specifications includes common time dummies to account for (potential) cross-sectional dependency. A * denotes statistical significance at least at the 5 percent level. NT = 126. Estimations undertaken with Rats 6.01.

Table 5. Panel Group FMOLS Results for Sub-periods, 1960-81 and 1982-2001.

A.	**Production Function, 1960-81.**		**Labor Productivity Function, 1960-81**	
	(1)	(2)	(3)	(4)
l	0.65 (5.67)*	0.83 (6.15)*	—	—
k_p	0.28 (6.06)*	0.23 (2.10)*	0.24 (3.00)*	0.20 (1.89)*
k_g	0.36 (9.85)*	0.35 (9.21)*	0.36 (9.39)*	0.34 (9.21)*

B.	**Production Function, 1982-2001**		**Labor Productivity Function, 1982-2001**	
Variables	(1)	(2)	(3)	(4)
l	1.39 (9.91)*	1.37 (8.98)	—	---
k_p	0.12 (4.71)*	0.14 (8.62)*	0.11 (4.22)*	0.13 (5.27)*
k_g	0.06 (1.73)*	0.04 (1.42)	0.04 (2.04)*	0.05 (1.60)

Notes: Estimates refer to (fixed-effects) long-run elasticities of output with respect to the relevant regressors. T-ratios are in parenthesis. Equations (2) and (4) in both specifications also include common time dummies to account for (potential) cross-sectional dependency. A * denotes statistical significance at least at the 5 percent level. NT = 66 for 1960-81 period and NT=60 for 1982-2001 period. Estimations undertaken with Rats 6.01.

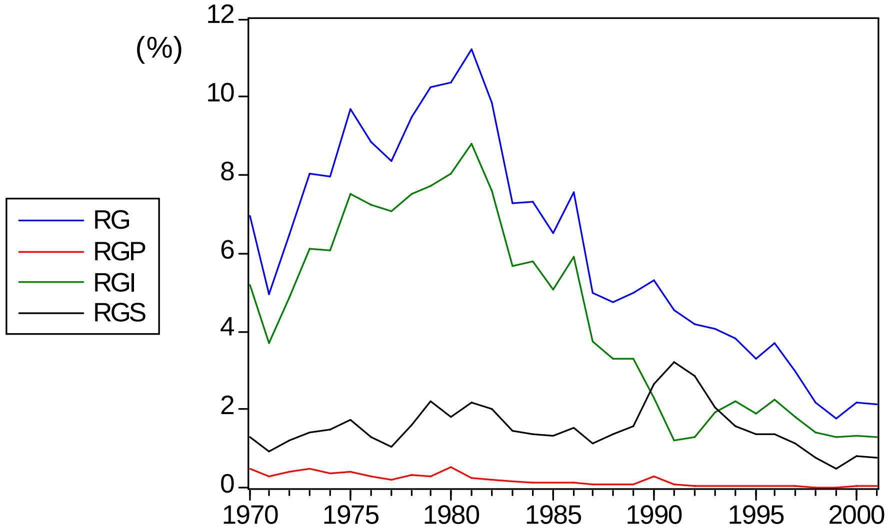

Figure 1. Public Investment as a Share of GDP by Sector.

Notes

1. For example, private investment as a proportion of GDP stood at 16.3 percent in 1993, and after the "peso crisis" of 1994-95, fell sharply to12.4 percent in 1995. It then rose to 18.3 percent in 1997 following the country's recovery (fueled by rapid growth in the U.S.), only to fall again to 15.9 percent in 2001 as a result of the downturn in the U.S. economy. Based on author's calculations and investment data in Everhart and Sumlinski [2001], Table C1., p. 58.

2.For further details see International Finance Corporation (IFC), Trends in Private Investment in Developing Countries, Statistics for 1970-2000. Washington, D.C.: The World Bank, 2001; and Instituto Nacional de Estadistica Geografia e Informatica (INEGI), Anuario Estadistico de los Estados Unidos Mexicanos. Aguascalientes, Ags.: INEGI, 1999.

3. 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. To the extent that increases in public investment on economic and social infrastructure reduce the relative price of energy, transportation, and human capital to firms in the private sector, it will, *ceteris paribus*, reduce their prime costs, raise profit margins, and spur further investment and output.

4. It can also be argued that 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 Chile and Mexico) are in the process of awarding concessions to private firms to produce and provide quasi-public goods and services. However, as Prager [1992] correctly observes, relatively little or no attention has been given to the monitoring or supervision costs of outsourcing public works projects. If these costs are substantial, particularly in the medium run, the bias in favor of privatizing these types of expenditures is removed.

5. Critics of the complementarity hypothesis contend that, in addition to the direct (negative) effects of public investment spending, there are indirect ones that arise from the financing of these expenditures with government bonds, the printing of currency, higher current and future taxes, and increased foreign borrowing in structurally weak banking and financial sectors. Thus, even when there are alleged positive effects from public investment spending, they may be completely offset by the combined crowding-out effects arising from both the financing and promotion of these types of expenditures [see Devarajan and Zou, 1994 and Green and Villanueva, 1991]. Given that this study is not based on a general equilibrium framework of analysis where such issues are addressed, no strong conclusions can be derived from the reported estimates.

6.Several studies have used population data as a proxy for the labor force, but this imposes the unrealistic assumption of a constant labor force participation rate, thus generating relationships that are mis-specified and subject to significant measurement error.

7. The initial stocks of private and public capital were constructed on the basis of the methodology first developed by Harberger [1972] (reported in Hoffman, 2000, pp. 276-77). After allowing for depreciation, the growth rates of both the private and public stocks were assumed to be equal to the average annual growth rate of real gross domestic product over the 1950-59 period (6.2%). The private capital stock figures generated by this methodology are similar to those reported by Hoffman (*op.cit.*) and Looney [1985]. To ensure the robustness of the results, other estimates of the rate of depreciation were used (2.5 and 7 %), as well as different estimates of the initial capital stock (summing over five years), but the results were not altered significantly.

8. In his seminal paper, Pedroni [1999a] shows that the " within-dimension statistics are constructed by summing both the numerator and the denominator terms [of the panel cointegration statistics] over the N dimension [cross-sections] separately, whereas the between-dimension statistics [referred to as group cointegration statistics] are constructed by first dividing the numerator by the denominator prior to summing over the N dimension" (p. 6). Pedroni further notes that because the between-dimension [group panel statistics] do not presume a common first-order autoregressive parameter [for all the cross sections], it allows the investigator "...to model an additional source of potential heterogeneity across individual members of the panel" (*ibid*.). In all, Pedroni constructs seven cointegration statistics (four panel statistics and three group statistics) and they are reported in Table I, p.7.

9. Pedroni [1999a] argues that one of the attractive features of his newly developed tests is that "...they allow the cointegrating vector to differ across members [of the panel] under the alternative hypothesis" (p.4). He goes on to warn that " incorrectly imposing *homogeneity* [emphasis added] of the cointegrating vectors in the regression would imply that the null of no cointegration may not be rejected despite the fact the variables are actually cointegrated"(*ibid*).

10. Pedroni [199b] shows via small sample Monte Carlo simulations that the bias (and sampling variance) of the group mean FMOLS estimator (based on the "between" dimension of the panel) is very small "...even in extreme cases when both the N and T dimensions are as small as N=10 and T=10 and become minuscule as the T dimension grows larger" (p. 23). And, in general, provided that T exceeds N (which is clearly the case in this study), Pedroni shows that the small sample properties of both the estimator and the associated t-statistic are extremely well-behaved "...even in panels with very heterogeneous serial correlation dynamics, fixed effects and endogenous regressors"(p. 24). For a concrete application of this recently developed methodology to a test of the Purchasing Power Parity Hypothesis, see Pedroni [2001], pp. 727-731; see also an interesting study by Dreger and Wismar [2004], pp. 1-17, who use this methodology to test whether health care expenditures are a luxury good in OECD countries.

11. Pedroni [2001] observes that, despite their widespread use in panel estimation, common time dummies cannot account for other forms of dependency such as those that arise from

“....dynamic feedback effects that exist between variables of different [cross sections]...”(p. 730). He proposes using a GLS type approach to account for these (non-contemporaneous) effects that is beyond the scope of this paper. For further detail, see Pedroni [1999b, pp. 28-29].

12. The inclusion of both common time dummies and the D_1 and D_2 dummy variables in eqs. (2) and (3) does not make sense because it generates zero for all entries and a non-invertible matrix.