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Infrastructures and productivity: an updated survey

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Abstract

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The relationship between infrastructures and productivity has been the subject of an ongoing debate during the last two decades. The available empirical evidence is inconclusive and its interpretation is complicated by econometric problems that have not been fully solved. This paper surveys the relevant literature, focusing on studies that estimate aggregate production functions or growth regressions, and extracts some tentative conclusions. On the whole, my reading of the evidence is that there are sufficient indications that public infrastructure investment contributes significantly to productivity growth, at least for countries where a saturation point has not been reached. The returns to such investment are probably quite high in early stages, when infrastructures are scarce and basic networks have not been completed, but fall sharply thereafter. Hence, appropriate infrastructure provision is probably a key input for development policy, even if it does not hold the key to rapid productivity growth in advanced countries where transportation and communications needs are already adequately served.

Key words: infrastructures, public capital, productivity

JEL Classification: H54, O47

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1.Introduction

An important part of public investment expenditure finances the construction and maintenance of infrastructures (such as transport, energy and water supply networks) that provide productive services to the private sector of the economy as well as direct consumption benefits. This paper surveys the literature that has tried to measure the impact of such expenditure on productivity growth through the estimation of aggregate (national or regional) production functions in which public capital or infrastructures enter as an input, or through cross-country growth regressions.

Although research in this area has been quite active for two decades, we are still far from having reached a consensus on the contribution of infrastructure investment to economic growth. A good many studies have addressed the issue using different econometric specifications and various regional and national samples with mixed results. Some of them conclude that the rate of return on infrastructure capital is very high and that public investment is an important determinant of the rate of growth of national or regional income. Other studies, however, obtain much more pessimistic results and conclude that the contribution of public capital to aggregate output is very small or non-existent -- or at least cannot be detected using the traditional framework of an aggregate production function. Since this discrepancy can, to a considerable extent, be traced back to the use of different empirical specifications in the existing studies, it will be necessary to start by briefly discussing some econometric issues. This will be done in Section 2, trying to keep the technicalities to a minimum. Sections 3-5 of the paper review some of the most interesting studies on the subject, starting with Aschauer's controversial (1989a) paper and the reactions to it. I will concentrate on a series of studies that have tried to quantify the direct impact of infrastructure investment on productivity through the estimation of aggregate production functions and convergence equations. I will not deal with the relationship between private and public investment or discuss the numerous studies that have approached the subject through the estimation of cost or profit functions. Section 6 closes the paper with a brief summary and some reflections on the implications of the existing empirical evidence[1].

1: See Gramlich (1994) and Draper and Herce (1994) for two broad reviews of the literature on infrastructure investment which cover many issues I do not address and provide useful references. Munnell (1992) provides a non-technical introduction to many of the questions discussed in this paper. Kessides (1993), Hurst (1994),

2. Empirical specifications and some econometric issues

Most econometric studies of the determinants of national or regional productivity start out from the assumption that there exists a stable relationship between aggregate output on the one hand and the stocks of productive factors (labor and different types of capital) and the level of technical efficiency on the other. This relationship is described by an aggregate production function that is typically assumed to be of the Cobb-Douglas type:

$$(1)\ Y_{it} = A_{it}\, Kit^{\alpha k} P_{it}^{\alpha p} L_{it}^{\alpha l}$$

In this expression Y_{it} denotes the aggregate output of region or country i at time t, L_{it} is the level of employment, K_{it} the stock of private physical capital, P_{it} the stock of public capital or productive infrastructures and A_{it} an index of technical efficiency or total factor productivity (TFP) which summarizes the current state of the technology and captures omitted factors such as geographical location, climate and endowments of natural resources. The coefficients α_i (with $i = k, p, l$) measure the elasticity of output with respect to the stocks of the different factors. An increase of 1% in the stock of public capital, for example, would increase output by α_p%, holding constant the stocks of the other factors and the level of technical efficiency.

Given time series data on output and factor stocks in one or more countries or regions, we can use standard statistical techniques to determine the values of the coefficients α_k, α_l and α_p that best explain the observed behavior of output in terms of the evolution of factor inputs. To this end, it is convenient to take logarithms of (1) in order to obtain a linear relationship between the variables of interest. Using lower-case letters to denote logarithms and adding a random disturbance to the equation, we have:

$$(2)\ y_{it} = \alpha_{it} + \alpha_k k_{it} + \alpha_p p_{it} + \alpha_l l_{it} + \alpha_{it}.$$

A second transformation of this expression is also useful. For each variable x, let $\Delta x_{it} = x_{it}+1 - x_{it}$ be its *first difference*. If we lead equation (2) by one period and subtract from it equation (2) itself, we obtain an expression that relates the rate of growth of output in each period (Δy_{it}) with the rates of accumulation of the different productive factors and the rate of technological progress ($\Delta\alpha_{it}$)[2]:

$$(3)\ \Delta y_{it} = \Delta\alpha_{it} + \alpha_k\, \Delta k_{it} + \alpha_p\, \Delta p_{it} + \alpha_l\, \Delta l_{it} + \Delta\varepsilon_{it}$$

Equation (2) is sometimes called a specification in levels of the production function, while equation (3) is a specification in (first) *differences*.

One of the problems that arise in the estimation of any of these specifications is that a_{it} (the level of technical efficiency) and $\Delta\alpha_{it}$ (the rate of technical progress) are not directly observable. To estimate the desired relation, therefore, it is necessary to make additional assumptions about the behavior of these terms. To illustrate some of the specifications commonly used in the literature, let us assume that the rate of technical progress (g) is constant over time and common to all the countries in the sample. In this case we have $\alpha_{it} = \alpha_{io} + gt$ and the specification in levels would be of the form

$$(2')\ y_{it} = \alpha_{io} + gt + \alpha_k k_{it} + \alpha_p p_{it} + \alpha_l l_{it} + \varepsilon_{it},$$

where the term a_{io} captures differences across countries in the level of technical efficiency (which are assumed to remain constant over time) and other unobservable factors that may affect output. The existence of such specific effects raises important estimation issues. Although some authors have argued that it may be reasonable to assume a common value of a_{io} because most technical knowledge is in principle accessible everywhere, casual observation suggests that levels of technological development differ widely across countries and possibly regions, and that other unobserved factors may be important. If this is so, failure to control for such differences (or for any other relevant variables) will bias the estimates of the remaining parameters whenever the regressors in the equation are correlated with the omitted variables. In other words, we can only legitimately subsume technological differences across countries in the error term if they are uncorrelated with factor stocks. This seems unlikely, however, as the level of total factor productivity is one of the key determinants of the rate of return on investment. The problem also arises, and may be even more important, in the case of public capital, as anything that increases income is likely to increase the demand for infrastructures. Hence, estimates of equation (2') that do not control for specific effects may be subject to a sort of endogeneity bias reflecting reverse causation from income to public capital. I will return to this issue in greater detail below.

2: Notice that the variables in equation (2) are in logarithms. The change in the logarithm of a variable from one period to the next is approximately equal to the percentage increase in the variable.

The standard solution for this problem is to turn to panel data techniques in order to control for unobserved national or regional effects. One possibility is to estimate a fixed effects model. This involves introducing dummy variables (or some equivalent device) in order to estimate a different regression constant for each country or region. A second possibility is to estimate a random effects model. In this case, the a_{io}'s are treated as different realizations of a stochastic variable, rather than as parameters to be estimated. The regression would now yield a single constant term (a_o) that would capture the average value of a_{io}, and differences in technical efficiency across countries would be captured by a perturbation term which would include a national component.[3]

Each of these forms to control for the possible existence of specific effects has its own advantages and disadvantages. The main shortcoming of the fixed effects model is that it ignores most of the information contained in the cross-section variation of the data. When we employ this specification we are essentially estimating the production function with the variables measured in deviations from their average values for each country (taken over the entire sample period). As a result, the parameters are identified by the variation over time of output and its determinants in the different countries. The random effects model, on the other hand, does make use of the cross-sectional variation in the data but it has the important disadvantage that if the specific effects are correlated with the regressors, as seems quite likely, the estimation will yield unreliable results (the coefficient estimates will be inconsistent). In principle, it is possible to determine which specification is better by making use of the so-called Hausman test, which is basically a test of the hypothesis that the specific effects and the regressors are uncorrelated.

The two types of techniques we have just summarized can also be applied to control for period-specific effects (introducing a dummy variable for each year of the sample or a temporal component in the error term). These period-specific effects may be used to capture the impact of technical progress or the effect of cyclical shocks if both processes follow a similar pattern in the entire sample. Both types of effects can also be used in combination with a specification in first differences. Their interpretation, however, will be different in each case. If we take first differences of equation (2'), for example, we see that the (time-independent) country specific effect disappears. Hence, the specification in first differences (3) implicitly allows for cross-country (but time-independent) differences in levels of technical efficiency. When we include national dummies in this equation, their coefficients will capture permanent differences across countries or regions in rates of technical progress.

To complete the specification, many studies add other variables to the estimated equations. A common practice is to include the rate of unemployment or some indicator of capacity utilization to control for cyclical shocks, or to use these variables to correct the capital stock series prior to estimation. Some studies disaggregate the private or public capital stocks into various components (equipment and structures, or roads and other infrastructures), or add other variables to the list of productive factors (e.g. some indicator of the stock of human capital).

As we will see, different specifications of the same production function often yield very different results. To determine which results are more reasonable we have at least two sets of criteria. The first one has to do with the statistical "goodness" of the estimation. In each case, we should check the goodness of fit of the model, measured by the R^2 of the regression, the precision of the coefficient estimates, measured by their standard error or the corresponding *t* statistic, and the results of the relevant specification tests. A second, common sense, criterion would be to require

a) a specification sufficiently flexible so as to capture differences across countries and periods which, in all likelihood, should be there and

b) sensible estimates of all the parameters of the model, and not only of the coefficient of public capital.

Economic theory provides some help in this respect. Under the assumptions of perfect competition and constant returns to scale, the coefficients of private capital and labor in the production function should be equal to the shares of these factors in national income. Hence, the coefficient of private capital (or the sum of the coefficients of private and public capital, given that the second factor is a free input and private capital is entitled to the surplus) should be between 0.30 and 0.40, and the coefficient of labor (or the sum of the coefficients of labor and human capital, since the return to the second factor is a component of wages), should be between 0.60 and 0.70. Although none of the assumptions which yield this prediction holds in a strict sense in real economies, estimates of input coefficients that are very far from these benchmark values should be regarded as suspect and may indicate some problem with the specification.

3: That is, the error term would now be of the form $\varepsilon_{it} = \eta_{it} + \delta_i$, where δ_i is a random variable with zero mean. This formulation implicitly assumes that the sample of countries or regions constitutes a random draw from a larger population, an assumption which is probably not very reasonable in the present context. For a more detailed discussion of both types of specific effects models, see for example Kmenta (1986, pp. 525 and ff.). Holtz-Eakin (1994) also discusses both types of models and their use in the public capital literature.

Convergence equations

Many empirical studies of the determinants of growth have proceeded by estimating a convergence equation of the type popularized by Barro (1991a)[4], that is, a regression of the form

$$(4)\ \Delta q_{it} = X_{it}\gamma - \beta q_{it}$$

where Δq_{it} is the growth rate of output per worker or per capita in country i over the period that starts at time t, q_{it} the logarithm of the initial level of the same variable and X_{it} a vector of other relevant variables. The initial level of income is included in the regression to capture a convergence effect that will favor poorer economies under the standard assumption of decreasing returns to scale in capital.

Equation (4) provides a flexible framework that has been used in the literature to investigate the impact on growth of a host of variables, including public investment and many other fiscal indicators. One important advantage of this approach is that, unlike production functions, convergence equations can be estimated without data on factor stocks, which are often hard to come by. In early studies in this literature, the empirical specification was frequently ad hoc and only loosely tied with the theory. In more recent years, however, researchers have increasingly focused on the estimation of "structural" convergence equations derived explicitly from formal models.

One of the most popular specifications in the literature is the one derived by Mankiw, Romer and Weil (MRW 1992) from an extended neoclassical model à la Solow (1956) with exogenous technical progress. Working with a log-linear approximation to the model around its steady state, MRW derive an equation linking the growth rate of output per worker over a given period (Δq_{it}) with the logarithm of the initial value of the same variable (q_{it}), the rate of growth of the labor force (n_{it}) and the shares in GDP of various types of investment measured over the same period. If we start from a constant-returns production function that includes both private and public capital as inputs (as in equation (1) above) and assume a common rate of depreciation for both types of capital, the equation takes the following form:[5]

$$(5)\ \Delta q_{it} = g + \beta(a_{io} + gt) + \beta \frac{\alpha_k}{1 - \alpha_k - \alpha_p} \ln \frac{s_{kit}}{\delta + g + n_{it}} + \beta \frac{\alpha_p}{1 - \alpha_k - \alpha_p} \ln \frac{s_{pit}}{\delta + g + n_{it}} - \beta q_{it}$$

where

$$(6)\ \beta = (1-\alpha)\,(\delta+g+n),$$

g is the rate of technical progress (which is assumed to be constant over time and across countries), δ the depreciation rate, α_i the coefficient of capital of type i in the aggregate production function, t the time elapsed since the beginning of the sample period, a_{io} the logarithm of the index of technical efficiency or TFP at time zero, s_i the share of investment in type i capital in GDP and n_i the rate of growth of the labor force.[6]

It is important to understand that the estimation of equation (5) does not imply that we are literally accepting the assumptions of the underlying Solow-type model (i.e. we do not need to assume that investment rates are necessarily exogenous for estimation purposes or constant over time). What we are doing is simply assigning to some of the parameters of the Solow model (in particular, to sk, sp and n) the observed average values of their empirical counterparts during a given period. During this period, the economy will behave approximately as if it were approaching the steady state of the Solow model that corresponds to the contemporaneous parameter values. In the next period, of course, we are likely to observe different values of the investment and population growth rates and therefore, a different steady state, but this poses no real difficulty. In essence, all we are doing is constructing an approximation to the production function that allows us to recover its parameters using data on investment flows rather than factor stocks. This is very convenient because such data are widely available and can be expected to be both more reliable and more comparable over time and across countries than most existing estimates of factor stocks.

5: Barro and Sala i Martin (1990, 1992) derive a similar expression from a variant of the optimal growth model of Cass (1965) and Koopmans (1965) with exogenous technical progress. The resulting equation is similar to (5) except that the investment rate (which is now endogenous) is replaced by the rate of time discount among the determinants of the steady state. The convergence coefficient, β, is now a more complicated function of the parameters of the model, but it still depends on the degree of decreasing returns to capital and on the rates of population growth, depreciation and technical progress.

6:In principle, β should vary across countries reflecting differences in rates of population growth (n_i). Most studies, however, impose a common value of this parameter for all countries. In this case, the n that enters equation (6) should be interpreted as the average value of this variable in the entire sample.

The empirical implementation of equations (4) or (5) does not, in principle, raise special problems. Given data on output, employment and private and public investment flows for a sample of countries or regions, we can use (5) to recover estimates of the rate of convergence and the parameters of the production function. The convergence equation can be estimated using either cross-section or pooled data. Most of the earlier convergence studies took the first route, averaging the variables over the entire sample period and working with a single observation for each country or region. The second possibility, which has become increasingly popular over time, involves averaging over shorter subperiods in order to obtain several observations per country.

In either case, one difficulty that immediately becomes apparent is that three of the variables on the right-hand side of the equation (g, δ and a_{io}) are not directly observable. In the first two cases, the problem is probably not very important. Although these coefficients can be estimated inside the equation (and this has been done occasionally), the usual procedure in the literature is to impose "reasonable" values of these parameters prior to estimation. The standard assumption is that $g = 0.02$ and $\delta = 0.03$, but researchers report that estimation results are not very sensitive to changes in these values. As we have already seen, the possibility that initial levels of technical efficiency (a_{io}) may differ across countries does raise a more difficult problem. Our earlier discussion about the estimation of models with specific effects applies also to convergence equations, as it is unlikely that initial TFP is uncorrelated with other regressors, and in particular with income per capita.

Reverse causation

As we have already seen, a standard concern when we are trying to estimate the impact of infrastructures on productivity is that reverse causation from income to public investment may generate an upward bias in the estimated coefficient of infrastructures in the production function. The source of the reverse causation bias is that the feedback effects of income on infrastructure can generate a correlation between the latter variable and the disturbance of the production function, thereby violating the conditions that are necessary for the consistency of ordinary least squares (OLS) estimators.

To illustrate the nature of the problem, suppose we are trying to estimate a production function

$$(7)\ y_i = \theta p_i + \alpha k_i + \varepsilon_i$$

where y is log output, p and k the (logs of the) stocks of infrastructures and other capital and and ε_i a random disturbance. Suppose also that the supply of infrastructures is an increasing function of y given by

$$(8)\ p_i = by_i + cx_i + \nu_i$$

where x is some other relevant variable and ν_i a disturbance term. In this setup, a positive shock to income in equation (7) (a positive value of ε_i) will increase (y_i and hence) p_i through equation (8). As a result, the regressor p_i will be correlated with the disturbance of the production function and the estimate of its coefficient obtained by OLS will be biased.

To calculate the bias, let us ignore the additional regressors, k and x. (It can be shown that their introduction does not qualitatively alter the results I will present shortly). Solving (7) and (8) simultaneously, we can write y_i and p_i as functions of the disturbance terms, ε_i and ν_i. Substituting (8) into (7) (after setting $\alpha = c = 0$), we have

$$y_i = \theta(by_i + \nu_i)\ \varepsilon_i \rightarrow (1 - \theta b)y_i = \theta\nu_i + \varepsilon_i \rightarrow (9)\ y_i = \frac{\theta\nu_i + \varepsilon_i}{1 - \theta b}$$

and inserting (9) into (8)

$$(10)\ p_i = by_i + \nu_i = b\,\frac{\theta\nu_i + \varepsilon_i}{1 - \theta b} + \nu_i = \frac{\theta b\nu_i + b\varepsilon_i + (1 - \theta b)\nu_i}{1 - \theta b} = \frac{b\varepsilon_i + \nu_i}{1 - \theta b}$$

Let us now see what happens when we try to estimate equation (7) by OLS. Under our assumptions, the (expected value of the) OLS estimator of θ will be given by:

$$(11)\ \theta = \frac{Epy}{Ep^2}$$

Using (9) and (10) we have

$$(12Epy) = E\,\frac{b\varepsilon + \nu}{1 - \theta b}\,\frac{\theta\nu + \varepsilon}{1 - \theta b} = \frac{E(b\varepsilon + \nu)(\theta\nu + \varepsilon)}{(1 - \theta b)^2} = \frac{E(\theta b\varepsilon\nu + b\varepsilon^2 + \theta\nu^2 + \nu\varepsilon)}{(1 - \theta b)^2} = \frac{bE\varepsilon^2 + \theta E\nu^2}{(1 - \theta b)^2}$$

and

$$(13)\ Ep^2 = E\left(\frac{b\varepsilon + \nu}{1 - \theta b}\right)^2 = \frac{E(b\varepsilon + \nu)^2}{(1 - \theta b)^2} = \frac{E(b^2\varepsilon^2 + 2b\varepsilon\nu + \nu^2)}{(1 - \theta b)^2} = \frac{b^2E\varepsilon^2 + \theta E\nu^2}{(1 - \theta b)^2}$$

Hence, the OLS estimator of θ will be given by:

$$(14)\ \theta = \frac{Epy}{Ep^2} = \frac{bE\varepsilon^2 + \theta E\nu^2}{b^2E\varepsilon^2 + E\nu^2} = \theta\,\frac{E\nu^2}{b^2E\varepsilon^2 + E\nu^2} + b\,\frac{E\varepsilon^2}{b^2E\varepsilon^2 + E\nu^2} = \theta\,\frac{E\nu^2}{b^2E\varepsilon^2 + E\nu^2} + \left(\frac{1}{b}\right)\frac{b^2E\varepsilon^2}{b^2E\varepsilon^2 + E\nu^2}$$

which shows that its expected value will be a weighted average of the true value of the parameter, θ, and of the inverse of the slope of the feedback relationship from income to the stock of infrastructures, $1/b$. Rearranging terms, we can also write this expression in the form

$$(15)\ \theta = \theta + \left(\frac{1}{b} - \theta\right)\frac{b^2E\varepsilon^2}{b^2E\varepsilon^2 + E\nu^2}$$

where we see that the sign of the bias depends on the difference between θ and $1/b$ (and is not necessarily positive when $b > 0$ as is sometimes asserted).

Hence, the existence of a feedback relationship between income and the stock of infrastructures will bias our results if we try to estimate the production function by OLS. In practice, however, things are not necessarily quite as bad as the previous discussion may suggest because public capital stocks are likely to respond to income shocks only with a considerable lag due to the long gestation period of infrastructure projects. Thus, we should probably replace equation (8) by something like the following system

$$(16)\ P_{it} = (1 - \delta)P_{it} + IP_{it}$$

$$(17)\ IP_{it} = f\,(y_{it-1}, y_{it-2}\ldots)$$

where IP_{it} is infrastructure investment in period t. Notice that with this specification the problem disappears. Now, a positive shock to income in (8) will increase future investment but will not affect the current stock of public capital, implying that P_{it} can still be uncorrelated with the contemporaneous disturbance in equation (8).

It would be too hasty, however, to dismiss the problem entirely, for it may very well arise in many of the specifications used in the literature. For instance, the omission of fixed effects in the production function in levels is likely to cause trouble even in the model described by equations (8), (16) and (17). In this case, the composite error term in (8) would be of the form $(a_i + \varepsilon_{it})$ and its time invariant component (the fixed effect) would indeed affect P_{it} because it will have influenced investment in all previous periods. Hence, P_{it} is very likely to be correlated with $(a_i + \varepsilon_{it})$, which will again bias its coefficient.

Reverse causation can also be a problem when the production function is estimated in differences, as is often done, partly to remove the fixed effects bias. We now have

$$(18)\ \Delta y_{it} = g_i + \theta\Delta p_{it} + a\Delta k_{it} + \Delta\varepsilon_{it}$$

where we are allowing for the possibility that the rate of technical progress, g, may differ across countries. If equation (18) is well specified, its disturbance term $\Delta\varepsilon_{it}$ should only contain true random shocks to the growth rate that cannot be anticipated by governments and should not therefore feed back into Δ_{pit} through (16) and (17). But if this is not the case and the error term contains some systematic component of the growth rate that agents can anticipate (e.g. a fixed country effect in rates of technical progress), we may well find that Δ_{pit} is again correlated with the (enlarged) disturbance, particularly if the period over which we are computing growth rates is long enough for induced changes in investment rates to work their way through to the stock of public capital.

Data quality problems

Another common concern in the literature we will review is the poor quality of the data. The problem is likely to be particularly acute in the case of cross-country studies. One reason is that public investment includes rather different things in different countries because government involvement in educational and health care systems is more intense in some places than others. Other source of noise is the existence of multiple administrations and public enterprises, whose investment may not be captured by the existing statistics in a consistent way. Finally, existing data on stocks of infrastructures or public capital are generally scarce and may not be comparable across countries because of methodological differences in their construction. All these factors are likely to introduce a significant amount of measurement error in many existing data sets on public investment flows or public capital stocks. As is well known, this will in turn bias the relevant coefficient towards zero.[7]

7: For a discussion of the effects of measurement error (in a slightly different context) see de la Fuente and Doménech (2006a).

3. Time series studies with national data

Ratner (1983) provides one of the first existing estimates of an aggregate production function that includes public capital as an input.[8] After adjusting the stocks of private and public capital by the rate of capacity utilization, this author estimates a production function in levels with annual data for the private sector of the US economy and finds that the coefficient of public capital is positive and significant although fairly small (0.056). (See equation [1] of Table 1). Using a very similar specification with longer time series for the same country, Ram and Ramsey (R&R, 1989) and Aschauer (1989a) estimate a much larger coefficient for this variable (equations [2] and [3] in Table 1).[9] This last author also disaggregates the public capital stock and finds that the types of infrastructure with the largest impact on productivity are those having to do with transport and the supply of energy and water.

The work of Aschauer has received a considerable deal of attention and can be considered the point of departure of the large literature on the subject. Perhaps the main reason for its great impact is that it appeared at a time when both academics and policy-makers were searching without success for the causes of the alarming productivity slowdown that started in the mid 1970s in most industrial countries. In this context, Aschauer's results were quite appealing for they provided a plausible diagnosis of the problem and a simple policy prescription: increase public investment in infrastructure. (See Munell, 1990a).

Table 1

Time series estimates of the production function with aggregate US data

	[1]	*[2]*	*[3]*
α_k	[0.234]	[0.242]	[0.26]
α_p	0.056 (2.70)	0.240 (4.38)	0.39 (16.23)
α_l	0.710 (12.62)	0.518 (10.37)	0.35 (4.85)
R_2	0.969	0.94	0.976
variables in:	levels	levels	levels
other vars.:	t(+)	energy prices (-), t(+), cu	cu(+), t(+)
country:	USA	USA	USA
period:	1949-73	1948-85	1945-85
source:	Ratner (1983)	R&R (1989)	Aschauer (1989a)

Notes:
- t statistics inside parentheses below each coefficient.
- The coefficients that appear within brackets are estimated indirectly, using the assumptions on the degree of returns to scale of the production function made by each author. Ratner and R&R assume that the production function displays constant returns to scale in private and public capital and labor. Aschauer imposes the same assumption after testing the implied restriction on the coefficients.
- Other variables: cu is an index of capacity utilization and t a trend. The symbol "+" (-) inside parentheses indicates that the coefficient of the variable is positive (negative) and significant, (0) means that the coefficient is not significantly different from zero.

8:There is an earlier attempt by Mera (1973) using regional Japanese data, but it is hard to compare with the more recent literature.
9: Wylie (1996) essentially replicates Aschauer's analysis with Canadian data obtaining similar results. The coefficient of public capital is extremely large (0.517) and exceeds the one on private capital.

Aschauer's conclusions, however, were soon questioned on various grounds. Several authors have argued that his estimate of the elasticity of output with respect to public capital is too high to be credible and suggested that his results may be due to various econometric problems.[10] A first difficulty is the possibility of reverse causation. In this interpretation, public investment would be a superior good and the observed correlation between this variable and productivity growth would simply reflect the fact that governments tend to invest more in periods of rapid growth. Aschauer, however, is aware of this possibility and supplies some evidence that this is not the problem. First, he observes that his results do not change substantially when public capital is lagged. Moreover, if there were a reverse causation problem, we would expect to see a strong correlation between productivity growth and many components of public consumption and investment. A significant positive correlation, however, is only found for the component of public investment that is devoted to productive infrastructure. As we will see below, Fernald (1999) uses sectoral data to provide additional evidence that causation runs from infrastructure investment to productivity growth, and not the other way around. His key finding is that increases in the stock of roads seem to induce faster productivity growth in those industries that are more intensive users of such infrastructures -- a pattern we would not expect to find if causation ran from output growth to public investment.

A second and probably more serious objection is that Aschauer's analysis does not control for other possible determinants of productivity growth. As Holtz-Eakin (1994) observes, the American postwar data Aschauer used contain essentially a single observation: the simultaneous fall in public investment and productivity growth that starts in the mid 1970s. It is possible, however, that this pattern is only a coincidence. The fall in the US rate of public investment at the beginning of this decade seems to be due mostly to two factors: the completion of the interstate highway network, and the stabilization and subsequent decrease in the school-age population (Gramlich, 1994). On the other hand, the productivity slowdown could be the result of the rapid increase in energy prices following the oil shocks of the 70s and 80s, which may have rendered obsolete a significant fraction of the private capital stock,[11] or of some other factor. In any event, the coincidence of the two processes may be due to chance, and the apparent significance of public investment may simply reflect the fact that this variable serves as a proxy for some other omitted factor/s which would be the true cause of the problem. The results of Ford and Poret (1991) would seem to be consistent with this interpretation. These authors replicate Aschauer's analysis for eleven OECD countries and find a significant positive correlation between public investment and productivity growth only in half the cases.[12]

The third criticism of Aschauer's work takes us to relatively complex econometric issues. Some studies argue that Aschauer's results may be an example of the "spurious regressions" problem discussed by Granger and Newbold (1974).[13] These authors argue that in many cases the apparently good results of regressions in levels between non-stationary variables (i.e. between variables which display a trend) are not reliable and suggest the use of a specification in first differences in order to obtain consistent estimates.[14]

When this is done, the results are often, but not always, less favorable to the hypothesis that infrastructure investment has a substantial effect on productivity growth. This is illustrated by the difference between equations [1] and [2] of Table 2, taken from Serra and García-Fontes (SGF, 1994), who work with aggregate data for Spain. When the data are differenced, the estimated infrastructure coefficient is reduced by a third and loses its significance. This pattern, however, does not hold in the case of Portugal, where the coefficient of interest is essentially the same in both specifications (see equations [6] and [7] in Table 2).

10: Munnell (1992), for instance, finds it difficult to believe that the coefficient of public capital may be larger than that of private capital -- particularly because a large part of public investment finances "non-productive" activities which are not captured by existing national income statistics.

11: Tatom (1991b) (cited in Hurst (1994)) finds that controlling for the price of oil reduces the coefficient of public capital in Aschauer's specification to one third. R&R (1989), however, still obtain a very large coefficient for public capital after controlling for energy prices (equation [2] in Table 1).

12: A second fact that seems to point in the same direction is the dramatic difference between Aschauer's estimates and those of Ratner, which are obtained with a shorter sample. It must be noted, however, that the US infrastructures series used by both authors underwent a major revision prior to Aschauer's work. Hurst (1994) reports that Tatom (1991b) has reestimated Radner's model (using his specification and the original sample period) with the revised series and obtains a coefficient of 0.28 for public capital, which is closer to Aschauer's estimate than to Radner's original result.

13: See Aaron (1990), Hulten and Schwab (1991), Jorgenson (1991) and Tatom (1991a,b). These authors find that public capital loses its significance when a specification in first differences is used.

14: A variable is stationary if its distribution does not change over time. This excludes, for example, those variables which display a trend. Using Monte Carlo simulations, Granger and Newbold show that when the estimation is done in levels the standard inference procedures often lead to the acceptance of a spurious relationship between independently generated non-stationary variables. On the other hand, the estimation in first differences, which eliminates the trends, generally yields more reliable results. Nelson and Plosser (1982) find that a large number of macroeconomic series are non-stationary. This finding raises doubts about the validity of good deal of empirical work in macroeconomics.

Table 2

Time series estimates of the production function, Spain and Portugal

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*	*[7]*	*[8]*
α_k	0.40 (15.64)	0.45 (5.29)	[0.42]	[0.18]	0.613 (9.48)	0.374 (16.06)	[0.320]	0.443 (15.81)
α_p	0.27 (11.69)	0.18 (1.02)	0.19*	0.60 (7.74)	0.290 (4.52)	0.186 (3.15)	0.199 (2.03)	0.387 (7.43)
α_l	0.59 (10.70)	0.55 (2.91)	0.39*	0.22 (10.86)	0.290 (4.52)	0.667 (5.10)	0.481 (4.51)	0.104 (0.39)
R_2	0.99	0.85	0.99		0.998	0.998		
variables in:	levels	diffs.	levels	levels	levels	levels	diffs	levels
Estimation:	OLS	OLS	coint.	coint.	coint.	OLS	OLS	coint.
other vars.:	cu(+)		cu(+)			cu, rev	cu, rev	cu, rev
country:	Spain	Spain	Spain	Spain	Spain	Portugal	Portugal	Portugal
period:	1969-88	1969-88	1964-88	1964-89	1964-89	1965-95	1965-95	1965-95
source:	SGF (1993)	SGF (1993)	Bajo-Sosv. (1993)	Arg. et al (1993)	Mas et al (1993a)	Ligthart (2000)	Ligthart (2000)	Ligthart (2000)

Notes:
- *t* statistics inside parentheses below each coefficient.
- Bajo and Sosvilla report the Wald statistics proposed by Phillips and Hansen rather than *t* statistics. The symbol (*) indicates that the coefficient is significant at 1%.
- The coefficients that appear within brackets are estimated indirectly, using assumptions on the degree of returns to scale of the production function.
- *Other variables: cu* is an index of capacity utilization; *rev* is a dummy variable for the period 1975-85 (between the Portuguese revolution and the country's accession to the EU).
- As a proxy for capacity utilization, Ligthart uses an estimate of the output gap constructed using the Hodrick-Prescott filter. This author does not allow for technical progress (implicitly assumes constant TFP), except in equation [7], where he includes a constant in a specification in first differences.

It must also be noted that the estimation in first differences presents some shortcomings of its own. Munnell (1992) observes that this specification frequently yields implausible estimates of the parameters of private factors in the aggregate production function and argues that it would not be reasonable to expect a stable relationship between public investment and output growth over a one-year period. Secondly, some recent developments in econometrics suggest that the level estimates may be more reliable than it was at first thought. In particular, the estimation in levels of a relationship between non-stationary variables works well (is consistent) when the variables are cointegrated, i.e. when there exists a linear combination of them that is stationary. In fact, the OLS estimates obtained in this situation would be better than usual in large samples, although it is also true that the usual significance tests lose their validity because the distribution of these estimators will not be the standard one.[15]

The econometrics literature provides various cointegration tests and different procedures for estimation and inference with cointegrated variables.[16] Some of these techniques have been applied in the public capital literature. Table 2 summarizes the results of four studies that make use of cointegration techniques with Spanish and Portuguese data. Bajo and Sosvilla (1993) use the public capital stock series supplied in the MOISEES data base (Corrales and Taguas, 1991), while Argimón et al (1993) and Mas et al (1993) work with series which reflect a more restrictive concept of productive public capital or infrastructures. Although the estimation techniques differ and the exposition is not always clear, the results generally coincide. The existence of cointegration is accepted in most cases. Hence, the consistency of the estimates is guaranteed, and the hypothesis of spurious regressions can be rejected. The estimated coefficient of public capital is quite high and seems to be significant.[17] The application of these new techniques, however, does not conclusively establish Aschauer's thesis for, among other problems, the estimated coefficients vary considerable across studies and are not always reasonable when interpreted as output elasticities.[18]

15: Let xt and yt be two non-stationary series. In general, the variable et = yt - a - bxt is not stationary, but it may be so if there exists some value of b which makes the two trends "cancel each other." In that case, we say that xt and yt are cointegrated. Stock (1987) shows that the OLS estimators in levels are superconsistent when there is cointegration, i.e. that they converge to the true value of the parameter as the size of the sample increases and do so faster than under the standard assumptions.
16: Cointegration tests are basically a test of the stationarity of the residual of a regression in levels. For an introduction to cointegration and related issues and additional references see Griffiths, Carter and Judge (1993, chapter 21) or Maddala (1992, chapter 14).
17: Two of the studies report *t* statistics without much of an explanation. On the basis of our discussion, however, this statistic would not be the most appropriate one to determine the significance of the estimated coefficients.
18: There are a number of other studies that analyze the relationship between public capital and productivity in different countries using cointegration techniques. Among others, see Mamatzakis (1999) for Greece, Otto and Voss (1996) for Australia, Flores de Frutos et al (1998) for Spain, Ramírez (2000) and Albala-Bertrand and Mamatzakis (2001) for Chile, Ramírez (2002) for Mexico, Everaert and Heylen (2001) for Belgium and Kavanagh (1997) for Ireland. With the exception of the last one, all of these papers estimate significant and relatively large positive coefficients for public capital. Some of these authors also report the results of Granger causality tests that suggest that causality runs from public capital to output but not in the reverse direction.

Some sectoral evidence

Fernald (1999) analyzes the impact of road construction on productivity growth using industry-level US data. As I have already noted, his results support the view that the correlation between infrastructure and productivity does not reflect a reverse causation problem. If this were the case, we would expect to see no systematic variation across industries in the strength of the correlation. If road construction does increase productivity, on the other hand, the effect should be stronger in those sectors that are intensive users of road transport. Fernald finds that this is indeed the case and concludes that investment in highways is productive. As we will see, his findings are also broadly consistent with Aschauer's conclusions on the high rate of return on infrastructure investment, but with one important qualification: while the construction of the interstate highway system seems to have contributed substantially to productivity growth in the fifties and sixties, there are no reasons to expect that additional investment in roads will yield similar results. Hence, infrastructure investment may be subject to a saturation effect, with returns to investment dropping sharply once the basic networks required to articulate a territory are completed.

Fernald starts out from an industry-level production function of the form

$$(19)\ Y_{st} = A_{st}F^{s}(K_{st}, L_{st}, T_{st})$$

where Y_{st} is the aggregate output of sector s at time t, K is private non-vehicle capital and T measures internal transport services. Transport services are produced combining vehicles (V_{it}) and public road capital (P_t), which is a shared input for all sectors, through a Cobb-Douglas technology that is common to all industries:

$$(20)\ T_{st} = V_{st}^{\,a}\ P_t^{\,b}.$$

Substituting (20) into (19) and differentiating totally with respect to time, it is easy to obtain the following expression:

$$(21)\ \Delta y_{st} = \Delta a_{st} + \alpha_{sk}\,\Delta k_{st} + \alpha_{sl}\,\Delta l_{st} + \alpha_{sv}\,\Delta \nu_{st} + \alpha_{sp}\,\Delta_{pt}$$

where the Δ's denote growth rates and α_{si} is the elasticity of output with respect to the i-th input in sector s. Fernald assumes constant returns to scale in private inputs (K, L and V) and perfectly competitive markets. Under these conditions, the elasticity of output with respect to each factor (α_{si}) is equal to the input's share in output (θ_{si}). Using this result and data on factor shares, he computes a Solow residual or extended TFP growth measure which includes the contribution of road capital by

$$(22)\ \Delta b_{st} = \Delta y_{st} - \theta_{sk}\,\Delta k_{st} - \theta_{sl}\,\Delta l_{st} - \theta_{s\nu}\,\Delta \nu_{st}.$$

Combining equations (21) and (22), the growth rate of extended TFP can be written as a function of the technological shock and the growth rate of the stock of roads:

$$(23)\ \Delta b_{st} = \Delta a_{st} + \alpha_{sp}\,\Delta p_t.$$

This equation can be rewritten in a way that makes use of cross-industry variations in the share of vehicle capital in output by noting that, under the previous assumptions, the elasticity of sectoral output with respect to the stock of roads turns out to be proportional to this share. Notice that

$$(24)\ \alpha_{sp} \equiv \frac{F^{s}_{p}P}{F^{s}} = \frac{F^{s}_{p}P}{F^{s}_{\nu}V}\,\frac{F^{s}_{\nu}V}{F^{s}} = \frac{F^{s}_{T}bV^{a}_{s}P^{b}}{F^{s}_{T}aV^{a}_{s}P^{b}}\,\theta_{s\nu} = \phi\theta_{sv}$$

with $\phi = b/a$. Substituting this expression in (23),

$$(25)\ \Delta b_{st} = \Delta a_{st} + \phi\theta_{s\nu}\,\Delta p_t$$

we see that the impact of infrastructures on productivity growth will be larger in vehicle-intensive sectors. Notice that equation (25) can be used to recover estimates of the elasticity of sectoral outputs with respect to the stock of roads. Since the average vehicle share ($\theta_{s\nu}$) is 1.6%, the average output elasticity will be given by $0.016^{*}\phi$.

Fernald estimates equation (25) after transforming it in a way that seeks to avoid the endogeneity bias arising from the possible correlation between productivity growth (which depends in part on the unobservable technological shocks Δa_{st}) and the demand for public capital. He uses a SUR specification (with a different equation for each sector), imposing the equality of ϕ across equations and working with different industry disaggregations. He also allows the value of ϕ to change over time by introducing a dummy variable for the second part of the sample period (1974-89), when the construction of the interstate highway system had been largely completed.

The results of the estimation are shown in Table 3. When ϕ is constrained to adopt a single value for the entire sample period (equations [1] and [4]), the coefficient is positive, significant, and quite large, implying an average elasticity of output with respect to road capital of around 0.35. When the coefficient is allowed to vary over time, the earlier result is preserved for the first half of the sample period, but the coefficient is not significantly different from zero in the period after 1973. In the preferred specification, which controls for congestion in a rather ad-hoc way (equation [5]), the point estimate of the elasticity of output with respect to road capital drops from 0.27 in the first half of the sample period to 0.085 in the second, and the latter coefficient is not significantly different from zero.

Table 3
Fernald (1999)

	[1]	*[2]*	*[3]*	*[4]*	*[5]*
ϕ	22.1 [3.3]		22.9 [3.6]		
ϕ1953-73		17.4 [4.0]		19.3 [4.3]	17.1 [3.1]
$\Delta\phi$1974-89		-25.3 [11.2]		-14.7 [10.0]	
ϕ1974-89					5.3 [4.5]
data for:	9 ind group	9 ind group	1-digit SIC	1-digit SIC	9 ind group
congestion effect:	no	no	no	no	yes for >1973

Notes:
- Standard errors in brackets below each coefficient.
- Aggregation: nine industry groupings or ten 1-digit SIC industries. Annual data for 1953 to 1989.
- Equation [5] controls for a congestion effect after 1973 by dividing the road stock by an estimate of total miles driven. Evidence of congestion is found only for the second part of the sample period.

Using these results, Fernald estimates that highway construction contributed about 1.4 percentage points to aggregate annual productivity growth before 1973 and only about 0.4 points after this year. Hence, infrastructure investment may indeed account for a substantial share of the observed slowdown in US productivity. It does not follow, however, that an increase in public investment will trigger substantial productivity gains in the future, as the rate of return on road construction seems to have fallen dramatically with the completion of the basic interstate network.

4. Cross-country evidence

A large number of studies have explored the relationship between infrastructure and productivity using pooled data sets that combine observations for several years and different countries or regions. This section reviews the available cross-country evidence and the following one focuses on studies that use regional data sets for a given country.

Using pooled data allows the researcher to exploit the cross-section variation of the data in addition to its time-series dimension and should mitigate some, but not all, of the econometric problems that arise in time-series studies for a single country. In particular, panel results are less open to the spurious regressions problem arising from common trends in the data. On the other hand, relying on the cross-section variation in the data does not eliminate and may even aggravate the reverse causation problem that may affect studies of the productivity effects of public capital. The direction of the bias, however, is not clear ex-ante and will depend on the way public investment is financed. At the country level, it seems likely that rich nations will demand more public capital than poor ones, a fact that is likely to generate an upward bias in the estimated coefficient of public capital. A similar situation will tend to arise within a given country if regional governments are financed by "domestic" tax revenues, but the bias may be reversed if infrastructure investment is financed centrally and is used as a mechanism for regional redistribution. As we have seen, in order to avoid the potential endogeneity bias, it is important to control for fixed country or regional effects, as their omission will generate a correlation between the error term and infrastructure variables that will render OLS estimates of production function coefficients inconsistent.

4.1. Production functions

Working with panel data for the G7 and a specification in first differences, Aschauer (1989b) reports results that are quite similar to those of his previously cited study. (See equation [1] in Table 4). His estimates, however, are somewhat problematic. Since data on capital stocks are not available for all the countries in the sample, Aschauer uses the share of private and public investment in GDP as proxies for the corresponding stocks. In the absence of further corrections (as dictated for instance by a full structural convergence equation specification), this procedure would only be valid if the capital-output ratio remained constant both over time and across countries, which seems unlikely. Secondly, Aschauer estimates a regression equation with a single constant term, which amounts to imposing a constant rate of technical progress for all countries and periods. Since public investment is the variable with the clearest downward trend of all regressors, its coefficient will tend to pick up the decline in the rate of productivity growth in this specification, yielding a positive coefficient for public capital which, once more, may reflect a spurious correlation between this variable and the growth rate of output.[19]

Evans and Karras (E&K, 1994a) undertake a similar exercise with panel data for a sample of seven countries but investigate the sensitivity of the results to different specifications. As in the previous study, a production function specification in first differences without country or period effects (equation [2] in Table 4) yields a positive and significant coefficient of public capital, although much smaller than the one reported by Aschauer. This coefficient, however, loses its significance (and in fact becomes negative) when we add period or country effects (see equations [3] and [4]). On the other hand, the coefficients of some of the private factors also adopt implausible values with these specifications.

19: Aschauer (1994) also provides some panel estimates. His results are, in general terms, quite positive, even when he includes fixed period and country effects, but are not directly comparable with those of the studies discussed in this section. In this paper Aschauer estimates, among other specifications, a convergence equation in which the dependent variable is the rate of growth of output per employed worker and the regressors include the initial value of output per worker and the growth rate of the stock of public capital, but not the growth rate of the private capital stock or the private investment rate. Hence, the estimated coefficients would presumably incorporate an indirect effect that would capture the impact of public investment on growth through the induced increase in private investment.

Table 4

Production function estimates with panel data at the national level

	[1]	[2]	[3]	[4]	[5]	[6]
α_k	0.24 [0.05]	0.348 [0.110]	0.251 [0.140]	0.230 [0.146]	0.26 n.r.	0.22 (2.1)
α_p	0.41 [0.13]	0.182 [0.085]	-0.103 [0.097]	-0.108 [0.100]	0.223 (4.83)	0.13 (2.0)
α_l	0.65 [0.08]	0.637 [0.103]	0.387 [0.099]	0.394 [0.103]	0.57 n.r.	0.66 (5.6)
R_2	0.61	n.r.	n.r.	n.r.	n.r.	
variables in:	differences	differences	differences	differences	differences	levels
country effects:	no	no	no	random	no	fixed
time effects:	no	no	fixed	fixed	no	trend
sample:	G7	7 DCs	7 DCs	7 DCs	22 OECD	28 LDCs
period:	1966-85	1963-88	1963-88	1963-88	1960-2001	1981-91
other variables:	CU(0), DOIL(-)					OPEN (+), USA (+), HC (+)
source:	Aschauer (1989b)	Evans & Karras (1994a)	Evans & Karras (1994a)	Evans & Karras (1994a)	Kamps (2006)	Dessus & Herrera (2000)

Notes:
- Standard errors (within brackets) or t statistics (in parentheses) below each coefficient.
- Other variables: CU = capacity utilization; DOIL = dummy variable for the years of the oil crises (1974 and 1979).
- The sample used by Evans and Karras includes Belgium, Canada, Finland, Germany, Greece, the UK and the US.
- Kamps only reports t statistics for public capital but does indicate that the rest of the coefficients are significant at 5%.
- n.r. = not reported.
- Dessus and Herrera: Equation [6] is estimated by three-stage least squares as part of a simultaneous equation model. OPEN is a measure of openness, USA is an indicator of TFP in the US and HC the stock of human capital, measured by average years of schooling.

Kamps (2006) constructs estimates of net stocks of public capital for 22 OECD countries using investment data from the OECD Analytical Database and a perpetual inventory procedure with geometric depreciation. He estimates a Cobb-Douglas production function in first differences using both his own series and data from earlier national and OECD estimates that relied on different estimation procedures in different countries. Kamps reports that results for individual countries generally improve when his data are used. When data from OECD (1997) are used (which are available for a subsample of 12 countries), the estimated coefficient of public capital is negative in four cases and not statistically significant in six more, leaving only two countries with positive and significant infrastructure coefficients. When the exercise is repeated with Kamps' data, the counterintuitive negative signs disappear and the estimated coefficients are significant in five out of twelve cases (and in 12 out of 22 cases for the entire sample). The author also indicates, however, that country-level equations often produce implausible estimates of labor coefficients, possibly due to severe multicollinearity problems. To avoid such problems, he pools the data obtaining a preferred specification (shown in equation [5] in Table 4) that yields plausible and significant coefficients for private inputs and an estimate of 0.22 for the elasticity of output with respect to public capital. This estimate is close to the one obtained by Evans and Karras (1994a) using a similar specification. Unfortunately, however, Kamps does not investigate the robustness of his results to the introduction of fixed period effects.

Seeking to avoid endogeneity bias, Dessus and Herrera (2000) estimate a production function with public capital as part of a simultaneous equation model that includes two equations describing the accumulation of private and public capital that are derived from a flexible accelerator model. The production function includes human capital (measured by average years of schooling) as an input and controls for an openness indicator and for an index of TFP in the US as proxies for technical progress. Public capital is interpreted in a broad sense as the accumulated investment of all public administrations and government-owned enterprises. The system is estimated by three-stage least-squares and includes fixed country effects and country-specific trends. The estimated output elasticities are shown in equation [6] of Table 4. The coefficient of public capital is positive and significant but, as may be reasonably expected, it is also significantly smaller than the coefficient of private capital. The authors report that their results imply very similar rates of return for public and private capital (14.2% and 14.9% respectively). Hence, they conclude that while public capital does seem to be rather productive, there are no signs of a shortage of this input.

Estefahani and Ramírez (2003) and Yeaple and Golub (2007) also develop and estimate simultaneous equation models that endogenize investment in infrastructures. Yeaple and Golub use data for a sample of 10 industrial sectors in 18 developing and developed countries to analyze the effect of three types of infrastructures (roads, telecommunications and power supply) on productivity and on the pattern of sectoral specialization. They find that roads have a positive effect on productivity in practically all sectors, while telecommunications and power infrastructures have significant effects only in a few cases. Infrastructure stocks also seem to affect the sectoral composition of manufacturing output. Estefahani and Ramírez (2003) use data on aggregate output and stocks of telecommunications and power supply infrastructures in a sample of 75 countries. They find positive, significant and rather high output elasticities for both types of infrastructures (between 0.078 and 0.095 for telephones and between 0.128 and 0.156 for power generation facilities).

Physical infrastructure indicators

Canning (1998) constructs a data set on physical infrastructure indicators that covers the period 1950-95 for a large sample of countries. He collects, collates and homogenizes data from a variety of sources on kilometers of (total and paved) roads and railway lines, number of telephones and telephone main lines and electricity generating capacity.

Canning (1999) and Canning and Bennathan (C&B, 2000) use these data and additional variables (taken generally from the Summers and Heston data set and from Barro and Lee, 1993) to investigate the contribution of infrastructure investment to productivity growth using panel cointegration techniques developed by Kao and Chiang (1997). Both studies estimate an error-correction specification that takes the form of a production function in first differences augmented with lag and lead terms to allow for short-run dynamics around the long-run relationship between inputs and outputs described by the production function. The coefficients of the production function (which includes fixed period effects and implicitly allows for fixed country effects) are assumed to be the same for all countries, while the short-run coefficients (i.e. those of the lead and lag terms) are allowed to vary across territories. The authors argue that this procedure should help mitigate any potential reverse causation problems. They contend that while it is plausible to assume that the technical relationship described by the production function is stable across countries, the feedback relationship from income to factor demands will most likely vary across countries and will therefore be picked up by the short-run coefficients. Hence, while there may be two cointegration relations between output and factor stocks for each individual country, only the forward one, from inputs to outputs, should hold uniformly when the data are pooled.

Table 5

Canning and Bennathan (2000)

	[1]	*[2]*	*[3]*
α_k	0.392 (11.90)	0.371 (8.58)	0.365 (6.41)
α_h	0.059 (1.54)	0.035 (0.64)	0.112 (1.57)
roads	0.048 (2.30)	0.03 (0.12)	0.117 (3.73)
electricity	0.057 (3.13)	0.012 (0.50)	0.134 (4.05)
sample:	62 ctries.	low income, 31 countries	higher income, 31 countries

Notes:
- *t* statistics in parentheses below each coefficient.
- *roads* is kilometers of paved roads per worker.
- *Other variables*: Physical capital stocks are constructed by accumulating investment flows using Summers and Heston data and assuming a depreciation rate of *7%*. Human capital is proxied by average years of schooling, taken from Barro and Lee (1993).
- Constant returns to scale are assumed. The coefficient of labor is not estimated directly but recovered using this assumption (as 1 minus the sum of the rest of the elasticities).

Table 5 shows some of the results obtained in the second of these studies using a Cobb-Douglas specification and yearly data.[20] To interpret the table, it should be noted that infrastructure stocks enter the production function twice, directly and though the stock of physical capital, in which they are included. Hence, an insignificant coefficient of the infrastructure variables will indicate that this factor is just as productive as other types of physical capital, and a positive and significant coefficient would suggest that an increase in infrastructure will increase output holding constant the total stock of capital. All infrastructure variables are expressed in per worker terms. They are physical indicators, of kilometers of paved roads and installed generating capacity (kilowatts).

20: One of the main differences between Canning (1999) and Canning and Bennathan (C&B, 2000) is that the first paper includes among the regressors an indicator of the number of telephone lines. This variable is very significant and its inclusion tends to drive out the rest of the infrastructure indicators. C&B argue that the phone variable should be dropped because its significance is likely to be driven by demand effects. They argue that its estimated coefficient is entirely implausible as it would imply rates of return to investment in telephone lines of the order of 10,000%.

C&B first estimate the model using the entire sample and then repeat the exercise after splitting the sample by income levels. Their results suggest that infrastructure investment does contribute to productivity but that "extra returns" over and above those available from other types of investment can only be found in relatively rich countries.[21]

To complete their analysis in a way that allows them to calculate rates of return on investment, the authors also compile data on the construction cost of different infrastructures. Their calculations suggest a complex pattern with very significant differences across countries in both construction costs and rates of return. For a majority of countries, the expected rates of return on infrastructure investment are comparable with or even below those on other types of physical capital. There are some countries, however, where the returns to infrastructure appear to be quite high in relative terms either because there seem to be shortages of infrastructures relative to other inputs or because construction costs are rather low. These countries are mostly middle-income developing countries.

4.2. Cross-country growth equations

This section reviews some results on the growth effects of public investment that have been obtained as part of broader cross-country analyses of the determinants of economic growth. The empirical model estimated in most of these papers is a non-structural growth equation relating the growth rate of output to a set of explanatory variables which typically includes initial income per capita, some measure of investment in physical and human capital, the rate of population growth and possibly other regressors chosen from a long list of potential determinants of output growth (including such things as measures of political or social instability, openness to international trade and many other macroeconomic indicators). Different fiscal variables are then added to this benchmark specification, either one at a time or in different combinations, in order to establish their partial correlation with output growth.

Some of the studies I will review make some effort to identify the channels through which fiscal policy measures affect growth by changing the set of conditioning variables. A common approach involves comparing the coefficient of the variable of interest in two different growth regressions, one that controls for the investment rate (and/or other immediate sources of growth), and a second one which does not. If the variable of interest has a smaller coefficient and/or loses significance when the investment rate is introduced as a regressor, it is inferred that this variable works, at least in part, by stimulating investment. Otherwise, its impact on growth is attributed to its direct effect on productivity. These regressions are sometimes complemented by simple accumulation equations, with investment as the dependent variable, which are used to check whether the variable is indeed correlated with the rate of capital accumulation. A second and closely related procedure attempts to separate the "direct" impact of a given expenditure variable on growth or investment from the "indirect" (disincentive and crowding-out) effects that arise from its (tax or deficit) financing by comparing the coefficients of the expenditure indicator in two specifications, one which controls for public revenues and one which does not.

One of the earliest and most thorough studies of the impact of fiscal policy on growth is the one undertaken by Landau in a series of three papers (1983, 1985 and 1986). To investigate the effects of different types of government expenditures, this author estimates a series of growth equations that include initial income per capita *(qo)* and various investment ratios among the regressors using both cross-section and panel data for a large number of countries. His results suggest that the expansion of the government sector may have contributed to the growth slowdown in the developed countries and retarded the development of LDCs and that the contribution of public investment to productivity growth is small at best and may even be negative.

21: Canning and Bennathan (2000) also estimate a translog production function (but only separately for each infrastructure indicator at a time plus the standard inputs). The results are qualitatively similar to those reported in Table 5 but this more flexible functional form allows them to investigate the pattern of complementarity or substitutability across inputs (which is very severely restricted by the Cobb-Douglas assumption). The results suggest that both roads and electricity generating capacity are complementary with physical and human capital, and that diminishing returns are strong in both types of infrastructure alone (ie. they estimate positive cross product terms and negative square terms).

Table 6 summarizes some of the more relevant results of the latter two studies.[22] Equations [1]-[3] are estimated using panel data for a sample of 65 LDCs. The first equation controls for government revenues (TAX), the budget deficit (DEFICIT) and the private investment rate (SKPR) and includes the public investment rate (SKG) along with other types of government expenditures. In equations [2] and [3] the revenue and deficit variables are omitted, and in the latter private investment is also excluded from the set of conditioning variables. Controlling for private investment, public revenues and the budget deficit, government investment seems to make a small positive (but insignificant) contribution to growth (equation [1]). When the revenue variables are dropped, the coefficient of public investment drops to zero (equation [2]). When private investment is omitted, the coefficient of public investment becomes negative (equation [3]). From this pattern of coefficients Landau concludes that the direct effects of public investment are positive but small, and that its net effect on growth is insignificant or even negative once we take into account its adverse impact on private investment and the distortions generated by its financing. The results are even more discouraging for a sample of 16 developed countries (equation [4]). The coefficient of public investment is now significantly negative in an equation that controls for private investment (and other types of public expenditure).

Table 6
Main results of Landau (1985 and 1986)

	[1]	*[2]*	*[3]*	*[4]*
SKG	0.098 (1.24)	0.004 (0.008)	-0.021 (0.47)	-0.197 (2.98)
TAX	-0.079 (1.12)			
DEFICIT	-0.066 (1.48)			
SKPR	0.08 (1.90)	0.059 (1.37)		0.074 (1.88)
SH	0.031 (4.80)	0.0032 (4.87)	0.034 (5.03)	0.015 (4.71)
q_o	-0.307 (4.26)	-0.311 (4.80)	-0.281 (4.59)	-0.0014 (8.60)
N	151	151	151	375
R^2	0.741	0.714	0.710	0.65
Data	panel	panel	panel	panel
	4-yr	4-yr	4-yr	annual
Period:	1960-80	1960-80	1960-80	1952-76
Sample	65 LDCs	65 LDCs	65 LDCs	16 OECD
Source:	(1986)	(1986)	(1986)	(1985)

Notes:
- *t* statistics in parentheses below each coefficient. *N* is the number of observations or the degrees of freedom of the equation. The dependent variable is the average growth rate of real income per capita.
- *SH* is investment in human capital, proxied by a weighted average of the primary, secondary and university enrollment rates.
- Other control variables (coefficients not reported in the table): Equations [1]-[3]: log time trend, change in the terms of trade, contraction and recovery year dummies, dummies for Australia, Canada, Norway and Switzerland, public consumption net of education and defense, transfer payments, and spending on education, all measured as real shares in GDP.
Equation [4]: total population, growth rate of population, foreign aid received, distance to closest port for landlocked countries and average rainfall, shares of public consumption and transfers to households in GDP.
- Fiscal variables and investment indicators are typically moving averages of three lagged years ending in the first year of the current subperiod.

A second and highly influential set of studies is due to Barro (1991a,b). Using cross-section data for a sample of over seventy countries during the period 1960-85, this author estimates a series of growth and investment equations following essentially the same procedure as Landau. Table 7 summarizes Barro's key results on public investment, which are somewhat more optimistic than Landau's. The dependent variable varies across equations and is either the growth rate of income per capita *(GQ)* or the total investment rate *(SK)*, which includes both private and public investment measured as a fraction of GDP. Government investment *(SKG)* seems to have a positive effect on growth (equations [1] and [3]) and on private investment (equation [5])[23], although the relevant coefficients are not always significant. Controlling for total investment, however, the coefficient of *SKG* is practically zero (equations [2] and [4]), a result which suggests that public investment is approximately as productive as private investment. Barro is rather cautious in drawing conclusions, however, and argues that the relevant variable would be the stock of public capital, rather than the public investment ratio. Using a rough estimate of its stock as a regressor, the coefficient of public capital becomes insignificant in the growth equation and negative and significant in the investment equation (not shown in the table).

23: The dependent variable in equation [5] it the total investment ratio, which includes public and private investment. Since the coefficient of the public investment variable is significantly larger than one, the estimate implies that private investment increases with public investment.

Table 7

Barro (1991a, b)

	[1]	*[2]*	*[3]*	*[4]*	*[5]*
dep var =	GQ	GQ	GQ	GQ	SK
SKG	0.128 (1.24)	-0.015 (0.13)	0.255 (2.80)	-0.026 (0.26)	2.21 [0.38]
SK		0.073 (1.87)		0.106 (3.93)	
YPC	-0.0075 (7.50)	-0.0068 (6.80)	-0.0107 (2.49)	-0.0183 (4.16)	0.009 (0.50)
SH	0.0312 (4.22)	0.024 (2.79)		0.011 0.012	
GPOP				-0.59 (2.11)	
R2	0.62	0.65	0.56	0.69	0.66
N	76	76	72	72	72
source:	(1991a)	(1991a)	(1991b)	(1991b)	(1991b)

Notes:
- Cross-section data for 1960-85. *t* statistics (in parenthesis) or standard errors [in brackets] below each coefficient
- *SH* = secondary enrollment ratio, *GPOP* = growth rate of population.
- *Additional control variables:* All equations control for the real share of government consumption (net of defense and education expenditures) in GDP. In addition,
Equations [1] and [2]: primary school enrollment in 1960, number of revolutions or coups per year, assassinations per million inhabitants per year, price deflator for investment spending, deviation of the investment deflator from the sample mean.
Equations [3]-[5]: initial income per capita squared, military expenditure as a fraction of GDP, Gastil's index of political rights, dummy variables for socialist and mixed economies, dummy for countries which experienced war or revolution during the period, and dummies for Africa and Latin America.

Another interesting paper on fiscal policy and growth is the one by Easterly and Rebelo (1993). These authors follow a somewhat different approach than either Barro or Landau. Instead of seeking precise estimates of the growth effects of various fiscal instruments, they concentrate in trying to identify "robust" empirical regularities. In fact, one of their main conclusions is that reliable "structural" estimates may be very difficult to obtain due to the presence of severe multicollinearity in the data and the existence of endogeneity problems.

Table 8

Impact on growth and private investment of public investment. Easterly and Rebelo (1993)

	[1G]	*[1I]*	*[2G]*	*[2I]*
dependent variable =	GQ	SKPR	GQ	SKPR
[1] total consolidated public investment	0.04 (1.02)	-0.194 (2.08)	-0.004 (0.89)	-0.241 (2.57)
Public investment by level of government:				
[2] general government	0.453 (4.13)	1.008 (3.89)	0.388 (3.18)	0.771 (2.88)
[3] public enterprises	-0.001 (0.01)	-0.623 (3.40)	-0.13 (1.15)	-0.63 (3.04)
Sectoral public investment:				
[4] agriculture	-0.231 (1.13)	-0.943 (2.64)	-0.304 (1.36)	-0.74 (2.24)
[5] education	1.49 (2.26)	1.987 (1.29)	1.18 (1.60)	1.96 (1.40)
[6] health	0.011 (0.02)	0.027 (0.02)	-0.37 (0.49)	2.29 (1.95)
[7] housing and urban infrastructure	1.49 (2.82)	2.108 (1.65)	0.91 (1.48)	1.01 (0.85)
[8] transport and communications	0.661 (2.48)	0.01 (0.00)	0.626 (2.48)	-0.17 (0.43)
[9] industry and mining	0.218 (1.39)	-0.351 (1.35)	0.082 (0.53)	-0.359 (1.14)

Notes:
- *t* statistics in parentheses below each coefficient
- Conditioning variables: Equations [1] = per capita real GDP in 1960, primary and secondary enrollment rates in 1960, assassinations per million, revolutions and coups and war casualties per capita) and share of government consumption in GDp. Equations [2] = same as in [1] plus M2/GDP and trade share in GDP.
- Panel data with decade averages, 1960-90. The number of countries for which data are available varies across decades.

After exploring the impact of various fiscal indicators, Easterly and Rebelo undertake an analysis of the growth effects of public investment using a new and fairly detailed data set on public capital expenditures by sector and level of government compiled by them using in-house World Bank reports. They start out from two benchmark growth and investment equations to which they subsequently add different public investment indicators one at a time. The benchmark equations are regressions of the growth of income per capita (*GQ*) and the private investment rate (*SKPR*) on initial per capita GDP, the primary and secondary enrollment ratios, the share of government consumption in GDP and three measures of political and social instability (assassinations per capita, revolutions and coups, and war casualties per capita). Columns [1G] and [1I] of Table 8 report the coefficients of different public investment variables when these are added one at a time to the benchmark growth and investment equations. As a quick check on the statistical robustness of these results, the estimation is repeated in columns [2G] and [2I] in Table 8 after adding two new control variables to the equations (the ratio of M2 to GDP and the share of foreign trade in GDP).

The results show a rather mixed pattern and suggest that disaggregation may be important for understanding the impact of government investment on growth. Total consolidated public investment (including investment by public enterprises) seems to have virtually no effect on growth and tends to crowd out private investment (equation [1]). These adverse results, however, seem to be due to investment by public enterprises (equation [3]), as capital expenditures undertaken directly by the public administrations (equation [2]) are positive and significantly correlated with growth and private investment. When investment is disaggregated by sectors (equations [4]-[9]) several of its components have large positive effects. Although some of these positive coefficients are sensitive to the introduction of new conditioning variables in columns [2G] and [2I] (in particular, those of education and urban infrastructure, shown in rows [5] and [7]), investment in transport and communications has a robust positive effect on productivity growth (though not on private investment). The coefficients of this last variable and of general government investment, moreover, remain significant when these variables are instrumented to correct potential endogeneity problems.[24]

Hulten (1996), Rieber (1999) and Milbourne et al (MOV, 2003) estimate structural convergence equations derived from an extension of Mankiw, Romer and Weil's (MRW, 1992) generalized Solow model that includes public capital as a separate input. Rieber (1999) uses a variant of the MRW model (without human capital) in which the convergence equation is written in terms of the stock of public capital (rather than the public investment ratio) and divides the stock of public capital by the stock of private capital in order to capture congestion effects. He estimates the model using annual data for a panel of 12 OECD countries covering the period 1970-93. To avoid the small sample bias that can affect the estimation of the convergence coefficient in dynamic panels,[25] he uses a specification proposed by Holtz-Eakin (1993c) that involves pooling increasingly longer differences of the data starting from a fixed initial date. He initially allows for fixed country effects but these turn out not to be significant. Hulten (1996) and MOV (2003) use cross-section data to estimate the standard version of the MRW model (written in terms of investment rates and including human capital). To deal with the possible simultaneity bias, MOV reestimate the model using IV techniques. They consider as potentially endogenous all their regressors except for initial income per capita and use as instruments the observed values at the beginning of the sample period of a series of other variables (that include the size and growth rate of the population, different investment indicators, the average schooling of the male and female population and an index of ethno-linguistic fractionalization). Finally, Hulten (1996) extends the basic model to control for the effectiveness with which infrastructure is operated. He assumes that this factor's contribution to productivity can be a function both of its quantity and of its effectiveness. As a proxy for the second factor, he constructs an effectiveness index that combines data on the probability that telephone calls are successful, power losses as a fraction of total electricity generation, the percentage of paved roads that is in good condition and the share of locomotives that use diesel technology.

24: In fact, instrumental variables estimation yields extremely large positive coefficients for these two variables, a finding which would tend to support Aschauer's (1989a) conclusion that the rate of return on public investment is exceedingly high.
25: See Nickell (1981).

Table 9

Different estimations of MRW's model with public capital

	[1]	[2]	[3]	[4]	[5]
ln s_k		0.33 [0.08]	0.52 [0.13]	0.344 (3.60)	0.180 (2.05)
ln s_p		0.27 [0.11]	0.22 [0.36]	0.355 (2.81)	0.107 (0.89)
ln effectiveness					0.794 (4.24)
α_k	0.615 (11.10)	0.177	0.237	0.240	0.200
α_p	0.096 (3.53)	0.145	0.100	0.248	0.118
R^2	0.82	0.51			
Data:	panel*	cross-section	cross-section	cross-section	cross-section
Estimation:	OLS	OLS	IV	OLS	OLS
country effects:	not signif.				
time effects:	no				
sample:	12 OECD	72 ctries	72 ctries	46 LDCs	46 LDCs
period:	1970-93	1960-85	1960-85	1970-90	1970-90
other variables:	ln q_0 (-), n	ln q_0 (-), *ln* s_h (+),n(0),	ln q_0 (-), n(0), *ln* s_h (+)	ln q_0 (-), n(-), *ln* s_h (+)	ln q_0 (-), n(-), *ln* s_h (+)
source:	Rieber	MOV	MOV	Hulten	Hulten

Notes:
- MOV do not report the coefficients of the production function, only those of the log terms in MRW's growth equation. I have recovered the output elasticities implied by these coefficients.
- Other variables: all equations control for the log of initial income per capita (ln q_0) and for population growth (n) using a specification similar to equation (5). MOV and Hulten also control for enrollment rates as a proxy for investment in education (s_h).
(*) As noted in the text, the panel used by Rieber is somewhat peculiar. Each observation corresponds to an increasingly longer period which always starts in 1970.
- *t* statistics in parentheses and standard errors in brackets.

Table 9 shows the main results of these three studies. In all cases, the coefficient of public capital estimated by OLS with the standard specification is positive and significant (see equations [1], [2] and [4]), ranging 0.10 to 0.25. On the other hand, MOV's IV estimation yields inconclusive results. Using IV reduces the point estimate of the coefficient of public capital, although not by much, but considerably increases the standard error of the estimator, causing this variable to lose its significance. On the other hand, the results of the Hausman test the authors use to determine whether IV estimation is required or not are inconclusive. Hulten's results, finally, suggest that having a large stock of infrastructure capital may not be enough. His effectiveness index is highly significant and drives out the public capital variable. When the first variable is included in equation [5], the size of the coefficient of public capital drops by more than two thirds and loses its significance. The result, however, is difficult to interpret because it seems unlikely that the performance measures Hulten uses to construct his efficiency index are independent of the stock of infrastructure. At any rate, his findings provide a useful reminder that, as Pritchett (1996) argues, what matters is the flow of services you get out of your stock of infrastructure, and not how much you spent building it.

Devarajan et al (1996) examine the growth implications of the composition of government expenditure using an extension of Barro's (1990) model with two types of public spending. Since the production technology displays constant returns in private capital and government spending (while labor does not enter the production function), the reduced form of the model is of the *AK* type, and the equilibrium path involves a constant rate of growth with no transitional dynamics. For a constant level of public expenditure (measured as a fraction of GDP) the equilibrium growth rate is an inverted U-shaped function of the shares of the two components of public expenditure in the government budget. In the Cobb-Douglas case, the growth-maximizing expenditure shares are proportional to the output elasticities of the different expenditure items. As the authors emphasize, in this setup the coefficient of a given expenditure share in a growth equation will depend not only on its output elasticity but also on the level of expenditure: productive spending may actually decrease the growth rate if given an excessive weight in the budget because it will crowd out other expenditures that are even more productive at the margin. Hence, regression results may in principle (and conditional on the correct specification of a model that relies on rather implausible technological assumptions) be used to determine how spending patterns differ from an "optimum" that would maximize the growth rate for the given volume of total expenditure.

In their empirical analysis, Devarajan et al (1996) use annual data for a panel of 43 LDCs to estimate several specifications of an equation that relates the (5-year forward moving average) growth rate of income per capita to the composition of (central) government spending and a set of control variables comprised by the share of (central) government spending in GDP, the black market premium on foreign exchange, a set of continent dummies and an "external shock" variable that is constructed as a weighted average of import and export prices and the world interest rate. They do not control, however, for private investment rates, human capital, population growth or (in coherence with their model) for initial income, although they do indicate that their results are not sensitive to the inclusion of this last variable. They also report some results for a sample of 21 developed countries. Two alternative decompositions of public spending are considered: an "economic" one into current and capital expenditures, and a functional one that provides data for four expenditure categories (defense, health, education and transport and communications).

Table 10

Devarajan et al (1996)

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*	*[7]*	*[8]*
SHSKG	0.11 (1.80)		-0.045 (1.72)	-0.059 (3.41)	0.072 (4.57)			
SHSKGP^2	-0.003 (2.62)							
SHCURR		0.24 (2.39)						
SHCURR^2		-0.001 (1.95)						
SHT&C						-0.145 (3.16)	-0.037 (1.14)	0.089 (3.50)
sample	43 LDCs	43 LDCs	43 LDCs	43 LDCs	21 DCs	43 LDCs	43 LDCs	21 DCs
fixed effects	no	no	no	yes	no	no	yes	no

Notes:
- *SHCURR* = share of current expenditures in total central government expenditures; *SHSKG* = public investment as a share of total central government expenditure, *SHT&C* = share of transport and communications in total central government spending.
- *Other explanatory variables:* All equations for the sample of LDCs control for the variables indicated in the text. In the DC sample, the only one of these variables that is included is the weight of government spending in GDP. Equations [6]-[8] include as additional regressors the shares of defense, health and education in total government spending. The coefficients of these variables are not significantly different from zero in most specifications, although some significant coefficients are obtained for some further disaggregations of the expenditure shares.

The results, summarized in Table 10, are consistent with the predictions of the theoretical model and suggest rather different conclusions for developed countries and for LDCs. While the first group of countries would apparently benefit from an increase in the weight of spending in capital (*SHSKG*) or in transport and communications infrastructure (*SHT&C*), LDC governments seem to be overspending on these items. Evidence of non-linearities in the shares of capital (*SHSKG*) and current (*SHCURR*) expenditure is detected in equations [1] and [2], where squared terms are included as regressors. Rather surprisingly, most countries in the LDC sample seem to be in the downward-sloping branch of the inverted U curve in terms of capital expenditures and on the upward sloping branch in terms of current spending. The coefficient of the first variable is negative (equation [3]) and that of the second positive (not reported in the Table) when only linear terms are included in the equation. The negative coefficient on the share of capital expenditure survives (and is increased by) the inclusion of fixed effects in the equation (equation [4]), but is reversed for a sample of developed countries (equation [5]). In equations [6] and [7] the expenditure variables are the shares of the different functional expenditure categories. The coefficient of spending on transport and communications is negative and significant in equation [6] but loses its significance in a fixed effects specification (equation [7]) and becomes positive and significant in the sample of developed countries (equation [8]).

Table 11

Miscellaneous results on government and growth

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*
dep var =	*GY*	*GY*	*GQ*	*GQ*	*GQ*	*GQ*
GCURR	0.0088 (0.16)					
SKG	0.2737 (2.04)	-0.108 (1.02)				
SKPR	0.3279 (3.10)	0.158 (3.27)				
ST&C			-0.0001 (0.40)	-0.00044 (4.54)		
INFR.					0.012 (2.08)	0.0089 (5.62)
q_o					-0.00065 (1.53)	
GPOP	????	0.573 (1.94)	-0.0002 (0.98)	-0.0007 (3.73)	-0.00012 (0.59)	
POL			-0.00011 (3.71)	-0.00011 (4.24)		-0.00011 (4.85)
R2	0.20	0.737	0.40	0.54	0.42	0.51
N	38	24	57	57	57	57
sample	38 LDCs	24 LDCs	57 count.	57 count.	57 count.	57 count.
period	1980-85	1970-80	1970-85	1980-92	1970-85	1980-92
source:	Diamond (1985)	K&R (1990)	S-R (1998)	S-R (1998)	S-R (1998)	S-R (1998)

Notes:
- All equations estimated with cross-section data.
- *Sources: K&R* = Khan and Reinhart (1990), S-R = Sánchez-Robles (1998).
- *Dependent variable: GQ* = growth of output per capita; *GY* = growth of total output.
- *Independent variables:*
GCURR = public current expenditures as a share of GDP;
ST&D = government expenditure on transport, communications, fuel and energy as a share of GDP.
POL = proxy for political instability
GPOP = growth rate of the population (or the labor force in eqs. [1] and [2])
INFR = physical infrastructure index, constructed as a weighted average of several standardized measures of physical infrastructure units. The weights are derived from an auxiliary regression of the growth rate of income per capita on the principal components of the physical infrastructure indicators and a set of conditioning variables that differs slightly from the one used in the equations reported in the table. For the 1970-85 sample, the infrastructure indicators are railways (Kms.?) per head, roads (Kms.?) per square Km. and energy production capacity per capita; for the 1980-92 sample, the railroad variable is replaced by the number of telephones per capita.
- *Other conditioning variables not shown in the table*
Equation [1]: growth rate of the labor force.
Equation [2]: growth rate of exports
Equations [3]-[6]: dummy variable for NICs

A number of other studies have analyzed the growth effects of various types of fiscal policies using cross-country growth regressions with mixed and even contradictory results. Table 11 summarizes some of the relevant findings. Although the specifications are not fully comparable,[26] Diamond's (1985) results (equation [1]) seem to contradict those of Devarajan et al (1996): in his sample of LDCs, growth increases with capital expenditures and decreases with current spending. Khan and Reinhart (1990) find a negative but insignificant correlation between public investment and growth in a sample of 24 LDCs (equation [2]). Sanchez-Robles (1998) provides some evidence that physical infrastructure indicators perform better in growth regressions than measures of monetary expenditure. Using cross-section data to estimate a growth regression for various samples, she finds that while the share of GDP devoted by government to investment in transport, communications and energy (*ST&C*) is not significant (equations [3] and [4]), an indicator that summarizes the initial endowment of public capital measured in physical units (*INFR*, see the notes to the table for its construction) does enter the equation with a positive and significant coefficient. It must be noted, however, that the conditioning variables change somewhat across specifications, making it difficult to interpret her findings.

26: Diamond works with expenditure shares in GDP, rather than in total government spending, and controls for different things that Devarajan et al.

5. Regional evidence

Let us now turn to the evidence that is available at the regional level, starting with the US and continuing with Spain and other countries. As we will see, a large number of these studies conclude that infrastructure indicators tend to lose their significance in fixed effects or differenced specifications. Most of these studies, however, work with a sample of US states that starts in 1970. Results for the regions of Spain and other Mediterranean countries, on the other hand, tend to be much more optimistic, raising the possibility that a saturation effect may be at work.

5.1. Evidence for the US states

Munnell (1990b) constructs estimates of the private and public capital stocks for the states of the US[27] and estimates a production function with panel data using a specification in levels without specific effects. Her results coincide with those of Aschauer (1989a) in that the effect of the stock of infrastructures on productivity is positive and significant. (See equation [1] of Table 12). On the other hand, Munnell's estimate of the coefficient of public capital is much lower than Aschauer's (around 0.17 rather than 0.39). One possible explanation for this difference has to do with the existence of interregional spillovers which would be captured by national but not by regional estimates. In a second exercise (equation [2]) Munnell disaggregates the public capital stock into various types of infrastructures. Confirming Aschauer's results once again, she finds that the infrastructures having the greatest impact on productivity are roads and the water supply network. Another early study at the regional level which also uses US state data is the one by García-Milà and McGuire (1992a). These authors focus on the role of roads and educational expenditure, including both factors in a production function in levels with fixed period effects (equation [3] in Table 12). The coefficient of the stock of roads is positive and significant and its size is not very different from Munnell's estimate.[28]

Table 12

Panel data results for the US States. Munnell (1990b) and García-Milà & McGuire (1992a)

	[1]	*(t)*	*[2]*	*(t)*	*[3]*	*(t)*
constant (a_o)	5.70	(39.3)	5.72	(42)	-4.598	(12.11)
trend (g)	0.002	(2.70)				
α_k *(total stock)*	0.30	(28.9)	0.31	(28.1)		
α_{ks} *(structures)*					0.027	(1.55)
α_{km} *(machinery)*					0.449	(15.99)
α_l	0.59	(42.6)	0.55	(35.4)	0.465	(15.31)
α_p *(total stock)*	0.17	(9.4)				
α_{pr} *(roads)*			0.06	(3.8)	0.044	(8.89)
α_{pw} *(water)*			0.12	(9.6)		
α_{po} *(other)*			0.01	(0.7)		
α_h *(education)*					0.087	(9.41)
R^2	0.99		0.993		0.995	
variables in:	levels		levels		levels	
fixed effects:	no		no		by period	
period:	1970-86		1970-86		1970-83	
other variables:	U(-)		U(-)		E(+), Pop(+), Ind(-)	
source:	Munnell (1990b)		Munnell (1990b)		G-M&M (1992a)	

Notes:
- Other variables: *U* = unemployment rate; *E* = current expenditure on education by state governments; *Ind* = share of manufacturing in total output (average over the entire sample period for each state); *Pop* = total population.
- García-Milà and McGuire divide the stock of road capital by the land area of the state. Their human capital indicator is the median years of schooling of the population aged 25 or over in each state (average of the values corresponding to 1970 and 1980).

28: An even earlier study is the paper by da Silva, Costa, Ellson and Martin (1987). These authors estimate a translog production function with cross section data for the US states in a single year. One interesting finding of this study, which uses a more flexible functional form than the standard one in the literature, is that the elasticity of output with respect to infrastructure falls with the ratio of public to private capital, or with the per capita stock of public capital, thus providing some evidence of a saturation effect.

Although the nature of the sample and the more reasonable size of the coefficients obtained in the two studies just cited suggest that their results may be more reliable than those of Aschauer, some recent papers have questioned the validity of this evidence on the basis of econometric problems similar to the ones we have discussed in the previous section. García-Milà and McGuire (1992b)[29] and García-Milá, McGuire and Porter (1993) estimate a Cobb-Douglas production function using panel data for the 48 continental US states during the period 1970-83. Table 13 summarizes the results obtained with different specifications of the model. The estimation in levels and without specific effects (equation [1]) essentially reproduces Aschauer's (1989) results.[30] The coefficient of public capital, however, falls drastically when we add state-specific effects (equations [2] and [3]) and becomes negative when the model is estimated with the data in first differences (equations [4] to [6]). At the same time, however, the estimated coefficient of labor increases to unreasonable values, approaching one in the last specifications.

Table 13

Panel data results for the US States. G. Milà et al (1992 and 1993)

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*
α_k	0.255 (8.57)	0.461 (6.67)	0.239 (5.73)	0.273 (2.80)	0.295 (3.00)	0.351 (3.39)
α_p	0.394 (14.93)	0.035 (1.19)	0.053 (1.80)	-0.082 (1.41)	-0.100 (1.59)	-0.121 (1.55)
α_l	0.383 (10.96)	0.704 (19.63)	0.743 (22.27)	0.898 (17.75)	0.923 (17.70)	0.986 (16.53)
R^2	0.985	0.947	0.956	0.468	0.448	0.415
variables in:	levels	levels	levels	differences	differences	differences
period effects:	fixed	fixed	fixed	fixed	fixed	fixed
regional effects:	no	fixed	random	no	random	fixed

Notes:
- Source: García-Milà and McGuire (1992) and García-Milà, McGuire and Porter (1993).
- Annual data for the 48 continental US states, 1970-83.
- *t* statistics in parentheses below each coefficient.

García-Milà, McGuire and Porter (1993) argue that the preferred specification is the one in first differences with fixed state effects (equation [6]). This choice, which leads to a pessimistic conclusion regarding the impact of public investment on productivity,[31] is based on the systematic application of various specification tests. Thus, the choice of a model in first differences is justified on the basis of a cointegration test for panel data, and that of a fixed effects model is based on a Hausman test. Finally, they run various tests for endogeneity and measurement error, rejecting the hypothesis that the poor results are due to these problems.

Table 14

Panel data results for the US States. G. Milà et al (1996)

	[1]	*[2]*	*[3]*	*[4]*	*[5]*	*[6]*
α_k	0.327 (10.33)	0.515 (7.36)	0.191 (4.61)	0.289 (2.90)	0.303 (3.02)	0.348 (3.30)
α_l	0.319 (9.61)	0.704 (20.28)	0.756 (23.85)	0.898 (17.64)	0.919 (17.53)	0.985 (16.34)
α_{pr}	0.370 (18.01)	0.127 (4.25)	0.120 (4.51)	-0.007 (0.13)	-0.024 (0.39)	-0.058 (0.77)
α_{pw}	0.069 (3.35)	0.064 (4.07)	0.043 (2.71)	-0.002 (0.07)	-0.012 (0.47)	-0.029 (1.07)
α_{po}	-0.010 (0.49)	-0.071 (3.50)	-0.048 (2.40)	-0.056 (1.63)	-0.049 (1.37)	-0.022 (0.55)
R^2	0.987	0.755	0.915	0.469	0.450	0.414
variables in:	levels	levels	levels	differences	differences	differences
period effects:	fixed	fixed	fixed	fixed	fixed	fixed
regional effects:	no	fixed	random	no	random	fixed

Notes:
- Annual data for the 48 continental US states, 1970-83.
- *t* statistics in parentheses below each coefficient.

29: Although the title of this paper coincides with that of García-Milá and McGuire (1992a), it is a different and more recent work.
30: García-Milá and Mcguire (1992b) also experiment with a decomposition of the public capital stock into three categories: roads, sewage and water supply, and the rest. The impact on productivity of the first type of investment seems to be considerable, that of the second is still positive but smaller, and the effect of the third type is zero. The first two variables are significant in all the specifications in levels, but none of them maintains its significance in the equations in first differences.
31: By contrast, the conclusion of the first two authors in the first of the cited papers (García-Milà and Mcguire, 1992b) is moderately optimistic. In this article, the first differences specification is rejected on informal grounds.

García-Milà, McGuire and Porter (1996) repeat essentially the same exercise after disaggregating the public capital stock into three components: roads and highways (with coefficient α_{pr} in Table 14), water and sewers (α_{pw}) and other structures (α_{po}). The results, summarized in Table 14, are qualitatively similar to those reported above, although the coefficients of two of the components of the public capital stock (roads and water works) are positive and significant in specifications in levels with fixed or random effects, while the coefficient of the third component is negative and significant. All these variables, however, lose their significance in the specifications in differences, which turn out to be preferred according to various econometric tests. Balmaseda (1996a) replicates G. Milà et al (1996) using both the Munell (1990b) and Holtz-Eakin (1993) data sets for public capital at the state level and reaches similar conclusions. He also estimates some additional specifications (assuming perfect mobility of private inputs, allowing the production function coefficients to vary across states and disaggregating the public capital stock into its components) with similar results in all cases: in the preferred specification, which generally involves fixed state effects with the data in first differences, the public capital indicators never display a positive and significant coefficient.

Table 15

Panel data results for the US states. Evans and Karras (1994b) and Baltagi and Pinnoi (1995)

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
α_k	0.386 [0.016]	0.18 [0.043]	0.003 [0.028]	0.31 (30.1)	0.29 (11.62)	0.31 (17.60)	0.018 (0.80)	0.02 (0.08)
α_p	0.096 [0.021]	-0.048 [0.031]	-0.029 [0.046]	0.16 (9.04)	-0.03 (0.90)	0.004 (0.19)	0.12 (2.78)	0.02 (0.29)
α_l	0.541 [0.022]	0.717 [0.039]	0.885 [0.035]	0.59 (43.2)	0.77 (25.53)	0.73 (29.38)	0.96 (27.93)	1.01 (24.83)
vars. in:	levels	levels	difs	levels	levels	levels	difs	difs/IV
period effects:	no	fixed	fixed	no	no	no	no	no
regional effects:	no	fixed	no	no	fixed	random	fixed	fixed
other variables:	u(-), GS(+)	u(-), GS(n)	u(-), GS(+)	u(-)	u(-)	u(-)	u(-)	u(-)
source:	E&K (1994b)	E&K (1994b)	E&K (1994b)	B&P	B&P	B&P	B&P	B&P

Notes:
- Annual data for the 48 continental US states, 1970-86.
- *t* statistics in parentheses or standard errors in brackets below each coefficient.
- Other variables: *u* = unemployment rate; *GS* = "public services," measured by the current expenditure of state governments on education, roads, health services, police, fire control, sewer and garbage disposal.
- The estimated equations do not seem to include a trend or fixed time effects. In the specifications in levels this would amount to the implicit assumption of no technical progress.

The same conclusion emerges from the work of Evans and Karras (1994b) and Baltagi and Pinnoi (1995) who obtain similar results using a panel of state data to estimate the production function (see Table 15). As usual, the public capital variable displays a positive and significant coefficient when the equation is estimated in levels without specific effects (equations [1] and [4] in Table 15), but loses its significance when specific effects are introduced or the equation is estimated in first differences (equations [2], [3], [5] and [6]). In the second study, public capital remains significant when the equation is estimated in first differences (equation [7]), but the authors detect evidence of measurement error in this specification.[32] Once they correct for this problem using instrumental variables techniques, the coefficient of public capital is again insignificant (equation [8]), although the coefficients of the private inputs are rather implausible in this specification (and in Evans and Karras' preferred equations).

32: Following Griliches and Hauman (1986) Baltagi and Pinnoi interpret changes in the coefficient of public capital across specifications in differences of varying length as an indication that this variable is measured with error. To correct this problem, in equation [8] they instrument the public capital variable using its lagged levels and first differences.

Table 16

Panel data results for the US States. Holtz-Eakin (1994) and Holtz-Eakin and Schwartz (1995a)

	[1]	[2]	[3]	[4]	[5]	[6]
α_k	0.359 [0.0112]	0.301 [0.0302]	0.361 [0.0233]	0.504 [0.142]	0.316 [0.0103]	0.0746 [0.027]
α_p	0.203 [0.019]	-0.0517 [0.0267]	0.0077 [0.0235]	-0.115 [0.126]	-0.0378 [0.0103]	0.0456 [0.027]
α_l	0.497 [0.0144]	0.691 [0.0262]	0.659 [0.0225]	0.643 [0.137]		
variables in:	levels	levels	levels	difs*	difs**	difs**
period effects:	fixed	fixed	fixed	no	trend (+)	trend (+)
regional effects:	no	fixed	random	no	no	fixed
period:	1969-86	1969-86	1969-86	1969-86	1971-86	1971-86
source:	H-E (1994)	H-E (1994)	H-E (1994)	H-E (1994)	H-E&S (1995a)	H-E&S (1995a)

Notes:
- Standard errors within brackets below each coefficient. Annual data for the 48 continental states of the US-.
- (*) Equation [4] is estimated using data on "long differences" (difference between the values at the beginning and the end of the sample period), rather than in first (annual) differences.
- A Hausman test indicates that the fixed effects model (eq. [2] is preferable to the random effects specification (eq. [3]).
- (**) Equations [5] and [6] are estimated using a SUR specification, with each equation corresponding to a "long difference" of different length, obtained by fixing the initial year and taking longer and longer subperiods.
- In equations [5] and [6] the public capital variable is "infrastructure capital", and includes streets and highways, sanitation and sewage, electric, gas and water utilities. The equations are also estimated using the entire stock of public capital with similar results.

In a number of papers, Holtz-Eakin (1994) and Holtz-Eakin and Schwartz (1995a, 1995b) also arrive at negative conclusions regarding the impact of public investment on productivity on the basis of an analysis of US state data. Holtz-Eakin (1994) provides additional evidence that the introduction of specific effects leads to the loss of significance of public capital variables. Although this author rejects the use of first differences, arguing that the short-term correlation between the relevant variables could contain too much cyclical noise to allow us to estimate a sensible production function, his results show that the inclusion of specific state effects is sufficient for the public capital stock to lose its significance, as illustrated in Table 16 (equations [1]-[4]). In his opinion, the existence of such effects is more than likely, and the fragility of the results to their inclusion should be enough to raise reasonable doubts as to the magnitude of the contribution of public investment to productivity growth. Holtz-Eakin and Schartz (1995a) estimate a structural convergence equation with panel data for the US states, using a variant of the specification proposed by Mankiw, Romer and Weil (1992). They find a negative coefficient for public capital (equation [5]) and a small and insignificant positive coefficient in a specification with fixed state effects (equation [6]), which is accompanied by a rather implausible coefficient for private capital.

Holtz-Eakin (1994) and Holtz-Eakin and Schwartz (1995b) investigate the spillover effects of public capital across state boundaries. Holtz-Eakin (1994) reestimates some of the specifications reported in Table 16 with the state data aggregated into 8 regions and reports that the estimated coefficients do not change significantly, thus contradicting Munnell's conjecture that public capital coefficients should increase with the level of aggregation because estimates obtained from disaggregated data do not capture all the relevant external effects. (See equations [1]-[3] in Table 17). Holtz-Eakin and Schwartz (1995b) estimate a production function in which the effective stock of public capital of each state depends both on its own stock and on those of neighboring states. They assume, in particular, that the variable that enters directly into the (Cobb-Douglas) production function is the effective stock of public capital of each region (kpi^e), defined by[33]

$$(26)\ k_{pi}^e = k_{pi} + \delta\Sigma_n w_{in} k_{pn}^e$$

where k_{pi} is the (log of) the actual stock of public capital in region *i* and k_{pn}^e the effective stock of its *n*-th neighbor. The summation that appears in this equation is taken over all the neighbors of region *i*, w_{in} is the weight assigned to the *n*-th neighbor, and the parameter δ measures the rate at which the spillover effects of public capital decay across states. Letting k_p^e and k_p denote the vectors of effective and actual public capital stocks for the different states, equation (26) can be written in the form

$$(27)\ k_p^e = k_p + \delta W k_p^e$$

where *W* is a matrix of neighbor weights. This expression can be solved for the effective capital stocks as a function of the observable actual stocks.

33: The discussion suggests that the variables that appear in equation (26) should be measured in logarithms, but does not make it explicit.

$$(28)\ k_p^e = (I - \delta W)^{-1} k_p.$$

Equation (28) implies that each state's effective stock of highway capital is a weighted sum of its own physical stock and those of all other states, with weights decreasing with distance (measured by the number of border crossings) and possibly other factors.

Holtz-Eakin and Schwartz consider three alternative specifications of the weighting matrix. The first one ("total", with w_{in} = 1 for all *n*) involves using the sum of the (logs of?) the effective capital stocks of all neighbors. In the second one ("average"), this variable is replaced by a simple average of each state's neighbor's effective stocks ($w_{in} = 1/N_i$ where N_i is the number of neighbors of state *i*). In the third one, the weight of each neighbor of *i* is inversely proportional to its share in the total land area of the region formed by state i and all its neighbors. This last specification implies that the externalities generated by a given amount of public capital are larger if it is concentrated in a smaller region.

The results for these three specifications are shown in equations [4]-[6] in Table 17. As in the specification without externalities, the coefficient of public capital in the production function is negative, and the point estimate of the externality parameter (δ) is negative in two cases and insignificantly different from zero in the other. From these results the authors conclude that there is no evidence of interstate spillovers from public capital.

Table 17

Spillover effects across states? Holtz-Eakin (1994) and Holtz-Eakin and Schwartz (1995b)

	[1]	[2]	[3]	[4]	[5]	[6]
α_k	0.253 [0.021]	0.272 [0.0566]	0.312 [0.044]	0.415 [0.0304]	0.414 [0.030]	0.414 [0.030]
α_p	0.201 [0.048]	-0.120 [0.042]	-0.0497 [0.0372]	-0.0222 [0.0249]	-0.0191 [0.0238]	-0.0206 [0.0301]
α_l	0.563 [0.045]	0.722 [0.051]	0.698 [0.0451]	0.624 [0.0256]	0.624 [0.0233]	0.624 [0.0263]
δ				-0.148 [0.0352]	0.0489 [0.122]	-0.215 [0.152]
variables in:	levels	levels	levels	long diffs	long diffs	long diffs
period effects:	fixed	fixed	fixed	yes	yes	yes
regional effects:	no	fixed	random	fixed	fixed	fixed
data for	8 regions	8 regions	8 regions	48 states	48 states	48 states
neighbours				total	average	wtd avge
source:	H-E (1994)	H-E (1994)	H-E (1994)	H-E&S (1995b)	H-E&S (1995b)	H-E&S (1995b)

Notes:
- Standard errors within brackets below each coefficient. Annual data for 1969-86.
- In equations [4]-[6] the public capital variable is the stock of roads and highways.

Shioji (1999) explores the growth effects of public capital with data for the states of the US and the Japanese prefectures. Using panel data, he estimates a non-structural convergence equation of the form

$$(29)\ \Delta q_{it} = \alpha_i - \beta q_{it} + \gamma p_{it}$$

where q_{it} is gross state product per capita in state *i* at time *t* for the US and GDP per employed worker for Japan, Δq_{it} the average growth rate of this variable over the five-year subperiod starting at *t*, α_i is a state or prefecture fixed effect, and p_{it} the log of the stock of public capital per worker at the beginning of the subperiod, which is added to the equation as a determinant of the steady-state level of income per capita. The equation is estimated with the variables measured in deviations from their contemporaneous sample averages, and the stock of public capital is disaggregated into several components in some of the specifications. Given the chosen specification, the coefficient of public capital, γ, cannot be interpreted directly as the output elasticity of this factor and (since private investment rates are omitted from the equation), it will also capture any indirect effects of infrastructure that work through the induced accumulation of private factors. Shioji argues that this specification may be preferable to a production function or a structural convergence equation because, by using the initial stock of public capital as the explanatory variable, it is less likely to be subject to endogeneity bias and because it allows for the likely lag between infrastructure investment and output growth.

Equation (29) is estimated using a variety of econometric specifications with the results summarized in Table 18 for the US and in Table 19 for Japan. In the first specification, which is a pooled regression without fixed effects (equation [1] in Tables 18 and 19), the coefficient of the aggregate stock of public capital is positive and significant in the US and zero in Japan. Shijoi argues that this result is likely to reflect an endogeneity bias induced by the omission of the fixed effects. Since local and state governments are largely self-financed in the US, states that are rich for whatever reason are likely to have larger stocks of public capital, thus accounting for the positive correlation uncovered by the estimation. In the case of Japan, however, the result is reversed because public investment is largely centralized and is the key instrument of an active regional policy that seeks to mitigate income disparities across prefectures.

Table 18
Shioji (1999), results for the USA

	[1]	[2]	[3]	[4]	[5]	[6]
β	0.052 [0.009]	0.295 [0.050]	0.266 [0.043]	0.150 [0.024]	0.324 [0.056]	0.314 [0.053]
P	0.536 [0.108]	-0.104 [0.071]	-0.07 [0.075]	0.352 [0.122]		
Pedu					-0.262 [0.061]	-0.351 [0.062]
P infrast.					0.114 [0.068]	0.157 [0.069]
estimation method	pooled	LSDV	GMM(a)	GMM(b).	LSDV	GMM(a)

Notes:
- Standard errors in bracket below each coefficient.
- Data for 1963-93 at five-year subintervals for the 48 continental states of the US.
- Income variable = gross state product per capita.
- *P* is the aggregate stock of public capital, *Pedu*. the stock of educational capital (mostly school buildings), and *Pinfrast*. the stock of infrastructures, which includes streets and highways, sewage and utilities.

Table 19
Shioji (1999), results for Japan

	[1]	[2]	[3]	[4]	[5]	[6]
β	0.055 [0.007]	0.104 [0.013]	0.063 [0.010]	0.076 [0.011]	0.126 [0.015]	0.109 [0.014]
P	-0.000 [0.069]	0.335 [0.079]	0.590 [0.126]	0.236 [0.102]		
Pedu					0.007 [0.069]	0.266 [0.081]
P infrast.					0.176 [0.056]	0.241 [0.062]
Pcons					0.019 [0.034]	0.034 [0.038]
Pagric					0.127 [0.035]	0.227 [0.039]
estimation method	pooled	LSDV	GMM(a)	GMM(b).	LSDV	GMM(a)

Notes:
- Standard errors in bracket below each coefficient.
- Data for 1955-95 at five-year subintervals for 46 Japanese prefectures.
- Income variable = GDP per employed worker.
- *P* is the aggregate stock of public capital, *Pedu*. the stock of educational capital (mostly school buildings), *Pinfrast* includes public housing, sewage and garbage disposal, water, city parks, roads, ports, airports and industrial water supply. *Pcons* includes investment in land conservation and *Pagric.* refers to agricultural and fisheries infrastructures.

Fixed effects are introduced for both samples in equations [2]-[4] using three alternative econometric techniques. The simplest one (LSDV) involves the use of dummy variables to capture regional fixed effects, while the other two, labeled GMM(a) and GMM(b), are generalized method of moments techniques developed respectively by Arellano and Bond (1991) and by Blundell and Bond (1998) and differ essentially in that the first one identifies the parameters by exploiting the time variation in the data, while the second one also makes use of its cross-section variation. Because this feature makes GMM(b) estimates more sensitive to the likely endogeneity bias discussed above, Shioji argues that LSDV and GMM(a) estimates are likely to be the most reliable ones. Under these specifications (equations [2] and [3] in both tables) the pooled regression results are reversed: the coefficient of the public capital variable is now positive and significant in Japan and negative and insignificant in the US.

Seeking to explain this apparent puzzle, Shioji disaggregates the public capital stock into various components, which differ across countries reflecting data availability. His results (equations [5] and [6]) now suggest a positive and significant role for core infrastructures *(Pinfrast)* in both countries and for agricultural capital *(Pagric)* in Japan. The coefficient of educational capital *(Pedu)*, however, remains quite different across countries. Shioji interprets his results as a confirmation of a positive link between investment in core infrastructures and productivity growth. In addition, the exercise serves to illustrate the sensitivity of the results to the choice of estimation technique.

In summary, the conclusions of the most recent panel studies with US data are markedly more pessimistic than those of early estimates of the impact of public investment on productivity. As we have seen, a number of these studies conclude that the apparent significance of infrastructure capital in production function estimates may be due to the use of inappropriate specifications which could suffer from a spurious regressions problem or which do not allow us to control adequately for existing cross-regional differences. It must be said, however, that some of the specifications selected on the basis of different statistical tests do not yield reasonable results. For example, the specification preferred by García-Milà, McGuire and Porter (1993) (in first differences with fixed state effects) implies that the rates of technical progress of some states are consistently higher than those of the rest and, therefore, that income differentials across states will grow without bound. Labor's extremely high coefficient in this regression, moreover, suggests that the estimates are dominated by short-run cyclical effects which can distort the long-term technical relationship between inputs and outputs we would like to capture. Similar problems arise also in the case of Evans and Karras (1994b) but not in Holtz-Eakin (1994). At any rate, the recent evidence we have reviewed suggests, on the whole, that the optimistic results of Munnell and Aschauer may not be valid and point to the need for flexible and careful econometric specifications.

5.2. Regional evidence for Spain

Political decentralization following the transition to democracy has greatly increased the demand for regional data in Spain. Starting in the early 90s, a number of important studies have significantly increased the availability of data on key regional economic aggregates. One of the most relevant efforts involves the construction by a team of researchers from the Instituto Valenciano de Investigaciones Económicas (IVIE) of series of stocks of private and public physical capital that go back to mid 60s. These data have been used in several studies of the determinants of productivity in the Spanish regions. Unlike their US counterparts, most of these studies yield positive results about the impact of infrastructure investment on output growth using econometric techniques that control for possible regional specificities.[34]

The earliest estimates that exploit (a first version of) these data are due to Mas et al (1993a) and (1994). These authors estimate a production function in levels with a regional panel covering the period 1980-89. The public capital variable is the stock of productive infrastructure (including the transport network, water supply works and urban structures) estimated by IVIE (1993) and the private capital series is taken from Calabuig et al (1993). In the first paper, the preferred specification is the one in levels with random effects (selected on the basis of a Hausman test), while in the second one the authors choose a fixed effects specification (equations [1] and [2] in Table 20). In both cases the coefficient of public capital is positive and significant, although its size is much larger in the second study. To investigate the possible existence of cross-regional spillover effects, Mas et al (1994) also estimate an equation in which the infrastructure indicator for each region is the sum of its own stock of public capital and those of the neighboring regions (equation [3] in Table 20).[35] The (slight) increase in the size of the estimated coefficient is interpreted, rather optimistically, as evidence of the existence of such effects.[36]

34: Some authors (see for instance García-Milà and Marimón (1995)) have expressed concern about the possibility that the positive correlation between public investment and output growth detected in the studies reviewed in this section may simply reflect short-term demand effects arising from the construction of infrastructures. Although it is certainly possible that demand effects may generate such a correlation, studies based in the estimation of production functions should not be too sensitive to this problem. Notice that demand effects would have to work through an increase in employment and the stock of capital in construction and related sectors, which are already included among the inputs that appear on the right-hand side of the equation. If this were the only link between infrastructure investment and output growth, the stock of infrastructures per se should not be correlated with productivity once we control for private inputs.
35: Álvarez et al (2006) question the logic of Mas et al's "pseudo-test" for cross-regional externalities and fail to find evidence of the existence of such effects using more standard specifications. Other studies, however, report the opposite finding (see the references in Alvarez et al, 2006).
36: The change in the public capital coefficient is somewhat larger when the authors impose constant returns to scale in the private factors (K and L) rather than constant returns in all factors (neither of these restrictions is rejected by the data). In this case, public capital's coefficient increases from 0.243 (when the regressor is each region's own public capital stock) to 0.306 when the stock of neighboring regions is added to the domestic one. The change in the value of the coefficient, however, is smaller than the standard error of the estimate.

Table 20

Panel data results for the Spanish regions (Mas et al)

	[1]	[2]	[3]	[4]	[5]	[6]	[7]
θ_k	0.6308 (11.86)	0.435 (5.04)	0.401 (3.73)	0.4191 (14.52)	0.4508 (12.94)	0.4285 (15.57)	0.4059 (14.13)
θ_p				0.0697 (2.31)			
θ_{pinf} *(infra.)*	0.091 (3.10)	0.191 (2.70)			0.0831 (3.05)	0.0771 (2.89)	
θ_{pinfa} *(incl. ady.)*			0.214 (2.05)				0.1411 (3.48)
θ_{ps} *(social)*					-0.0247 (1.04)		
$\Sigma_i\, \theta_i$ *- 1*				0.0586 (1.63)	0.0735 (2.03)	0.0674 (1.88)	0.1069 (2.83)
θ_l	[0.28]	[0.374]	[0.385]	[0.5698]	[0.5643]	[0.5618]	[0.5599]
g		0.013 (4.21)	0.013 (3.63)	0.0173 (8.84)	0.0174 (8.98)	0.0168 (9.06)	0.0152 (7.53)
R^2	0.469	0.835	0.83	0.844	0.985	0.9845	0.9848
variables in:	levels	levels	levels	levels	levels	levels	levels
regional effects:	random	fixed	fixed	fixed	fixed	fixed	fixed
period:	1980-89	1980-89	1980-89	1964-91	1964-91	1964-91	1964-91
source:	(1993a)	(1994)	(1994)	(1996)	(1996)	(1996)	(1996)

Notes:
- *t* statistics in parentheses below each coefficient.
- *g* is the rate of technical progress.
- In equations [1]-[3], the coefficient of labor (θ_l) is not estimated directly, but recovered from the assumption of constant returns to scale in *K, P* and *L*.
- For columns [4]-[7], the estimated equation is of the form

$$y_{it} = a_{io} + gt + \theta_k(k_{it} - l_{it}) + \theta_p\,(p_{it} - l_{it}) + (\theta_k + \theta_p + \theta_l - 1)\; l_{it} + \varepsilon_{it}.$$

The coefficient of labor in this specification ($\eta = \Sigma_i\, \theta_i - 1$) will therefore be different from zero if the production function does not display constant returns to scale. The estimate of θ_l shown in the table is reconstructed using the estimated value of η and the other parameters.

In a later piece (Mas et al, 1996), the same authors reexamine the problem using longer series. The output and employment data are now taken from Fundación BBV (1995) and refer to private sector output, excluding the energy sector.[37] Mas et al estimate different variants of a fixed effects model disaggregating the public capital stock in various alternative ways. In equation [4] in Table 12, the regressor is the total stock of public capital, including accumulated investment in educational and health infrastructures ("social capital"). In equation [5], this variable is disaggregated into two components: an indicator of productive infrastructure, similar to the one used in previous studies by the same authors, and a measure of the stock of social capital. The coefficient of the first of these variables (θ_{pinf}) is around 0.08, while that of the second one (θ_{ps}) is not significantly different from zero (it should be kept in mind, however, that the output measure used in this study does not include public services). After excluding this second component of the stock of infrastructure in equation [6], the estimation is repeated in equation [7] using as a regressor the sum of the infrastructure stocks of each region and all the adjacent ones. The increase in the relevant coefficient (θ_{pinfa}), which is considerably more significant than in the previously cited study, is interpreted by the authors as an indication of the existence of cross-regional spillovers. Finally, Mas et al also find evidence of a gradual decrease in the elasticity of output with respect to the stock of infrastructure. To capture a possible saturation effect, the authors reestimate their production function with different subsamples of increasing length. The first estimation, which corresponds to the period 1964-73, yields a coefficient of public capital of 0.14, which falls to 0.077 when the entire sample period is used.

Serra and García-Fontes (SGF, 1994) construct estimates of the regional stocks of public capital using the data reported by Frutos (1991), the Regional Accounts of the National Statistical Institute and other sources, and estimate a production function both in levels and in first differences. Their results, although not as positive as those of Mas et al (possibly due to the lower quality of their data), display a rather different pattern than the US studies reviewed above. As can be seen in Table 21, the size of the coefficient of public capital is rather modest in the Spanish case when the estimation is in levels (equations [1]-[3]) but increases (instead of becoming negative) in a specification in first differences (equations [4]-[6]), although it also loses significance. On the other hand, the inclusion of specific effects in the equation in levels tends to improve the results in the Spanish case, and to worsen them with the US data.

37: The authors, however, do not discuss the construction of this aggregate which cannot, in principle, be reconstructed for the entire sample period using their sources.

Table 21

Panel data results for the Spanish regions (SGF, 1994)

	[1]	[2]	[3]	[4]	[5]	[6]
α_k	0.63 (19.73)	0.57 (17.34)	0.49 (10.88)	0.19 (1.88)	0.21 (2.15)	0.34 (2.99)
α_p	0.04 (1.74)	0.02 (0.75)	0.06 (2.53)	0.27 (1.53)	0.38 (1.93)	0.25 (1.23)
α_l	0.35 (7.68)	0.41 (10.36)	0.46 (8.21)	0.21 (2.95)	0.15 (2.06)	0.14 (1.95)
R^2	0.99	0.99	0.99	0.14	0.47	0.49
variables in:	levels	levels	levels	difs.	difs.	difs.
period effects:	no	fixed	fixed	no	fixed	fixed
regional effects:	no	no	fixed*	no	no	fixed*

- *(*) Note:* Given the shortness of the sample period, the authors have chosen to use a set of suprarregional dummies (rather than a dummy for each region) to distinguish between three groups of regions according to their average income levels during the sample period.

González-Páramo and Argimón (G-P&A, 1997) estimate a fixed effects model with data from the same sources as Mas et al (1996), disaggregating the public capital stock into a productive and a social component. Their results, summarized in columns [1]-[3] ot Table 21, are qualitatively similar to those of the previous study, although there are some changes in the results that may arise from differences in the data or in the specification. The finding that social capital now enters the equation with a positive and significant coefficient may be due to the fact that G-P&A work with total regional value added, without excluding the output of the public sector. On the other hand, the specification used by these authors does not seem to include either a trend or period fixed effects. This may explain why the estimated coefficients of capital and labor appear to be less plausible than those reported by Mas et al (1996). G-P&A also estimate a random effects model (equation [3]) including the regional land area (sup) among the regressors as a possible determinant of transport costs, together with the infrastructure stock. As expected, the estimated coefficient of this variable is negative and significant.

Dabán and Lamo (D&L, 1999) estimate various random effects specifications in levels using instrumental variables techniques to correct for possible endogeneity problems. Their output and capital stocks series are taken from the BDMORES database of the Ministry of Finance and cover the period 1980-93 at an annual frequency. The dependent variable is the output of the productive private sector, excluding both non-market services and residential rentals (both real and imputed). Among the explanatory variables, they include the stock of infrastructures, a human capital indicator (the number of workers with some university schooling) and three variables that may affect transport costs for a given stock of infrastructures: regional land area (*sup*), the standard deviation of altitude within the region (*alt*), and the dispersion of population (*pd*), measured by the number of population centers. Their specification controls for capacity utilization (*cu*) at the national level and for the regional rate of unemployment (*u*) to capture possible cyclical shocks.

Table 22

Panel results for the Spanish regions (various authors)

	[1]	[2]	[3]	[4]	[5]	[6]
θ_k	0.69 (22.68)	0.62 (16.13)	0.65 (17.41)	0.251 (3.70)	0.388 (3.78)	0.265 (2.65)
θ_p	0.19 (8.52)					
θ_{pinf} *(infra.)*		0.09 (3.89)	0.09 (3.62)	0.108 (4.42)	0.099 (2.90)	0.030 (1.00)
θ_{ps} *(social)*		0.13 (5.28)	0.13 (5.29)			
θ_h				0.249 (4.99)	0.016 (0.29)	-0.084(2.80)
$\Sigma_i \theta_i - 1$				0.145 (10.47)	0.137 (8.18)	
θ_l	[0.12]	[0.16]	[0.23]	[0.537]	[0.634]	[0.789]
R^2	0.978	0.978	0.975	0.726	0.516	
variables in:	levels	levels	levels	levels	q-difs*	
regional effects:	fixed	fixed	random	random	random	fixed
period effects:	no	no	no	no	no	fixed
period:	1964-91	1964-91	1964-91	1980-93	1980-93	1969-91
other variables:			sup (-)	sup(-), pd (-), alt (+), U(-), cu(?)	sup(-), pd (-), alt (+), U(-), cu(?)	initial income per capita (-)
source:	G-P&A	G-P&A	G-P&A	D&L	D&L	Gorost.

Notes:
- The coefficient θ_l shown in the Table is recovered using the assumptions made by each author about returns to scale and the relevant parameter estimates.
- *G-P&A* impose constant returns in all inputs (including land area in equation [3]). The coefficient of labor shown here is given by $\theta_l = 1 - \Sigma_{i \neq l} \theta_i$. The infrastructure stock used in equations [2] and [3] includes private toll highways.
- *D&L*. They employ GMM techniques, using as instruments the lagged values of the regressors and the public capital stocks of neighboring regions. The equation does not include a trend because this term is not found to be significantly different from zero in a preliminary estimate. The authors do not report the estimated coefficient of capacity utilization. As in Mas et al (1995), the coefficient $\Sigma_i \theta_i - 1$ captures the degree of returns to scale. See the notes to Table 6.
(*) Equation [5] is estimated in quasi-differences, using a Cochrane-Orcutt type procedure to correct for an autocorrelation problem detected in the residuals of equation [4].
- Gorostiaga instruments ln $(s_{pit}/(\delta+g+n_{it}))$ with ln $(\delta+g+n_{it-1})$ where $\delta = 0.05$ is the assumed depreciation rate, $g = 0.02$ the rate of technical progress and n the growth rate of the working-age population.

The results, summarized in columns [4] and [5] of Table 22, coincide in general terms with those of previous studies, at least in terms of the infrastructure coefficient, which lies around 0.10. Unlike other authors, however, Dabán and Lamo find evidence of increasing returns to scale and control for a human capital indicator as well as for orography and the pattern of population settlement. The educational variable seems to have a very large effect on productivity in the first specification (equation [4]), but loses its significance when a correction for autocorrelation is introduced in equation [5]. As for the second group of variables, both land area and the dispersion of the population are significant and enter the equation with the expected sign, but the orography variable displays a counterintuitive sign (other things equal, regions with more uneven terrain seem to be more productive).

Gorostiaga (1999) obtains rather more pessimistic results than the previous studies through the estimation of a convergence equation à la Mankiw, Romer and Weil (1992) derived from an extended Solow model that includes human capital and infrastructures as well as physical capital as inputs in the production function. This author estimates a fixed effects specification similar to the one proposed by Islam (1995) using panel data techniques and instrumenting the rate of public investment. She finds a negative coefficient for the stock of human capital and an elasticity of output with respect to the stock of infrastructures that is not significantly different from zero (equation [6] in Table 22). Although some aspects of the estimation are rather problematic,[38] the study is useful as a warning that, also in the Spanish case, the available results on the relationship between infrastructures and productivity are quite sensitive to the choice of specification.

38: In particular, the choice of proxy for the rate of investment in human capital and the instrument chosen for the rate of public investment are questionable. The first of these variables includes only direct public expenditure on education and, unlike most of the studies on the subject, does not make use of available information on enrollment or achievement rates. The chosen variable, moreover, is available only after 1980, forcing the author to extrapolate it backward over the rest of the sample period. As for the second issue, Gorostiaga essentially uses the lagged rate of population growth as an instrument for the public investment rate. This variable is unlikely to be a good instrument because public investment in Spain has been driven by redistributive concerns over much of the sample period and there is no clear correlation across regions between income levels and population growth rates.

Delgado and Alvarez (2000) construct a synthetic infrastructure indicator using a principal components technique to combine various measures of infrastructure endowments in physical units. Their series cover the period 1985-95 and are constructed using information on transport networks and facilities (kilometers of roads, highways and railways and measures of the capacity of ports and airports), telecommunications (number of telephone lines) and energy supply networks (kilometers of oil and gas pipelines and electricity supply lines). These variables are normalized by the land area of the region or by its population (ports and airports). The infrastructure indicator is used to estimate a regional production function where the stock of physical capital is measured, as usual, in monetary units. The authors estimate a fixed effects model and a model in first differences, using at times IV techniques with lagged regressors as instruments. One problem with their specification is that it does not seem to allow for technical progress: there is no trend in levels, no constant in differences and no fixed period effects.

Table 23 summarizes the main results. The coefficient of the infrastructures variable is positive, significant and quite large in the specification in levels but becomes considerably smaller and becomes only weakly significant when the model is estimated in first differences. On the other hand, infrastructures maintain their significance in both levels and first differences when instrumental variables are used to deal with potential endogeneity and measurement error problems.

Table 23

Delgado and Álvarez (2000)

	[1]	[2]	[3]	[4]
α_k	0.29 (6.48)	0.36 (7.84)	0.22 (2.83)	0.19 (3.65)
α_p	0.20 (4.47)	0.10 (1.63)	0.25 (3.15)	0.25 (3.50)
α_l	0.40 (6.96)	0.46 (5.72)	0.47 (6.33)	0.47 (4.85)
data in	levels	diffs.	levels	diffs.
reg effects	fixed		fixed	
estimation	OLS	OLS	IV	IV
period	1985-95	1985-95	1985-95	1985-95

Álvarez et al (2003) estimate a regional production function using two alternative infrastructure series: Delgado and Alvarez's (2000) data set based on physical indicators, and a new version of the IVIE series of net capital stocks measured in monetary units (FBBVA, 1998). Table 24 summarizes some of the main results. Equations [1]-[3] are estimated with the FBBVA data and equations [4]-[6] use Delgado and Alvarez's data on physical infrastructure indicators. Constant returns to scale in all three inputs are imposed in equations [3] and [6] but not in the rest. All equations include regional fixed effects.

The results are, on the whole, positive. The coefficient of infrastructures is positive, significant and large in all equations but one. The exception is equation [2] where fixed period effects are introduced to control for technical progress and constant returns to scale are not imposed. The authors indicate that there is a high correlation (over 0.90) between the stocks of infrastructures and private capital on one hand and time on the other, and that this makes it difficult to untangle their effects without additional restrictions. Imposing constant returns does help, in the sense that it helps recover what seem to be more sensible results, but the authors also note that the data seem to reject this restriction.

Table 24

Álvarez et al (2003)

	[1]	[2]	[3]	[4]	[5]	[6]
α_k	0.49 (11,3)	0.11 (2.5)	0.26 (5.0)	0.59 (13.5)	0.11 (2.9)	0.21 (4.1)
α_p	0.22 (8.3)	0.01 (0.2)	0.11 (4.3)	0.20 (4.7)	0.11 (3.1)	0.21 (7.7)
α_l	0.29 (10.6)	0.43 (7.0)	[0.63]	0.26 (5.2)	0.40 (7.3)	[0.58]
period effects:	no	yes	yes	no	yes	yes
CRTS	no	no	yes	no	no	yes
period:	1980-95	1980-95	1980-95	1980-95	1980-95	1980-95
infr. data:	FBBVA	FBBVA	FBBVA	D&A	D&A	D&A

- Note: all equations include regional fixed effects. D&A = Delgado and Álvarez (2000).

De la Fuente (2002) and de la Fuente and Doménech (D&D, 2006b) estimate a model that combines a Cobb-Douglas regional production function in differences (taken over two-year periods) with a technical progress function that allows for technological catch-up across regions using two consecutive versions of the IVIE data set (Fundación BBV (1998) and Mas et al (2002)). The specification includes regional fixed effects and either period fixed effects or a trend and trend squared. The results are shown in Table 25. As in most previous studies, the infrastructures variable is significant and displays a positive coefficient. The size of this coefficient, however, is significantly smaller than in some of the previous studies reported above, particularly in the case of D&D (2006b)—a result that implies more plausible rates of return to infrastructure than those of other studies discussed in this section.

Table 25

de la Fuente (2002) and de la Fuente and Doménech (2006b)

	[1]	[2]
α_k	0.297 (5.73)	0.171 (3.27)
α_p	0.106 (2.14)	0.0567 (3.25)
α_l	[0.597]	[0.772]
α_h	0.286 (7.30)	0.835 (2.04)
vars in	diffs	diffs
period effects:	no	yes
reg. fixed effects	yes	yes
other vars	tech gap, t, t2	tech gap
period	1964-93	1965-95
source	D 2002	D&D 2006

5.3. Regional evidence for other countries

Table 26 summarizes the results of a series of studies that have estimated Cobb-Douglas production functions using regional data for Greece and Italy. Bonaglia, La Ferrara and Marcellino (BFM, 2000) and Picci (2000) both use the series of regional private and public capital stocks constructed by Bonaglia and Picci (2000). The results I report are those obtained with a broad measure of public capital (including educational, administrative and health facilities as well as core productive infrastructures) which appears to be similar to the one used in R&S (2002) for Greece. By contrast, Bronzini and Piselli (B&P, 2009) rely on the series of stocks of public capital constructed by Montanaro (2003), using as a regressor a measure of the stock of core infrastructures that includes roads and water and electricity supply networks. Most papers use standard OLS specifications in levels with fixed regional effects or in first differences. The exception is Bronzini and Piselli (2009), who estimate the model in levels using panel cointegration techniques. These authors, in particular, use the fully-modified OLS estimator developed by Pedroni (2000), which corrects for endogeneity of the regressors and for serial correlation of the error terms.

Table 26

Regional results for Greece and Italy

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]
α_k	0.088 (2.00)	0.166 (2.26)	0.140 [0.021]	0.171 (17.2)	0.072 (2.77)	[0.35]	0.427 [0.031]	0.146 [0.019]
α_p	0.202 (2.53)	0.451 (2.46)	0.005 [0.029]	0.358 (16.3)	0.184 (2.69)	0.109 [0.004]	0.192 [0.007]	0.190 [0.011]
α_l	0.811 (16.39)	0.531 (6.21)	[0.855]	0.836 (15.7)	0.462 (8.12)	[0.65]	[0.573]	0.557 [0.033]
vars in	levels	diffs	levels	levels	diffs	levels	levels	levels
period effects:	trend		yes	yes	yes	yes	yes	yes
reg. effects	fixed	fixed	fixed	fixed	no	fixed	fixed	fixed
other vars	CU	CU				R&D(+), HC(+)	R&D(+), HC(+)	R&D(+), HC(+)
period	1982-92	1982-92	1970-94	1970-95	1970-95	1985-01	1985-01	1985-01
country	Greece	Greece	Italy	Italy	Italy	Italy	Italy	Italy
Source:	R&S	R&S	BFM	Picci	Picci	B&P	B&P	B&P

Notes:
- *Source:* R&S = Rovolis and Spence (2002); Picci = Picci (2000); B&P = Bronzini and Piselli (2009); BFM = Bonaglia, La Ferrara and Marcellino (2000).
- *Other variables: HC* = human capital, measured by average years of schooling. R&D = stock of R&D capital.
- B&P: Equations [6] and [7] assume constant returns to scale in private capital and labor. In equation [6], TFP is computed using the coefficients shown in the Table for these two factors (which reflect observed factor shares) and this variable is used as the dependent variable. In equation [7] the regressor is average labor productivity and the coefficient of labor that is shown in the table is recovered using the constant returns assumption. In equation [8] the dependent variable is total output and no assumption is imposed regarding returns to scale.

Generally speaking, the results are similar to those obtained with Spanish data. The coefficient of public capital is positive, significant and relatively large in most cases, even though all specifications control for regional effects. The one exception is BFM (equation [3]), where the coefficient of public capital is not significantly different from zero. The discrepancy between this paper and Picci's equation [4] is rather surprising given that both studies seem to use very similar data and a common econometric specification. At any rate, BFM (2000) also report that public capital has a large and significant positive coefficient in the center and south macro regions when the estimation is repeated independently for different subsamples, and that the same is true for the sample as a whole for the core component of productive infrastructures.

6. Conclusion

The relationship between infrastructures and productivity has been the subject of a debate that is still ongoing in the literature. The available empirical evidence is inconclusive and its interpretation is complicated by econometric problems that have not been fully solved. Early work on the subject, notably by Aschauer (1989a and b) and Munell (1990b), concluded that the elasticity of national or regional output with respect to public capital is large and very significant, and that the rate of return on public investment is exceedingly high. A number of more recent studies, however, have questioned these results on the basis of various econometric problems. Some of these studies find that the significance of public capital disappears when a specification in first differences is used or fixed effects are introduced to control for unobserved national or regional specificities, and conclude that the accumulation of public capital does not appreciably contribute to productivity growth. Other recent papers, by contrast, confirm the significance of infrastructure indicators using cointegration or panel data techniques and, in many cases, point to rates or return that are much more modest than those estimated in early work, but also much more credible.

Table 27 summarizes most of the evidence I have reviewed in Sections 4 and 5 of this paper. For each of four samples (comprised, respectively, by the states of the US, the regions of Spain, various groups of countries and regions of different nations) and three broad groups of specifications (convergence equations and production functions estimated with the data in levels, with and without specific (fixed or random) effects, and production functions estimated with the data in differences) the table shows the total number of equations reported in the text tables in previous sections of the paper and the fraction of this total in which the estimated coefficient of the relevant public capital or infrastructures variable is positive and significant (+), negative and significant (-) and insignificantly different from zero (0), using a *t* value of 1.8 as the cut-off point for significance.[39] The one exception has to do with the reporting of several papers by Canning and coauthors where different physical infrastructure indicators are added to an equation where the total stock of physical capital (including infrastructures) is one of the regressors. As the authors point out, the coefficients of the infrastructure variables in this specification will tell us whether or not infrastructures are more or less productive than the rest of the stock of capital, not whether it does make a positive contribution to productivity. Hence, I include these results in the (+) group whenever the infrastructure coefficients are either significantly positive or not significantly different from zero (as the total stock of capital always has a positive and significant coefficient).

39: In the case of Easterly and Rebelo, I consider only those equations where the public investment variable is one of the following: total public investment, general government investment, and investment in transport and communications. In each case, I take into account both the basic equation and the extended one that includes additional regressors.

Table 27

Summary of reported results

1. Data in levels without specific effects

	no. of eqs.	(+)	(0)	(-)	avge α_p	avge. t
US states	10	100%	0%	0%	0.204	9.43
Spanish regions	3	0%	100%	0%	0.03	1.16
cross country	29	45%	45%	10%		1.11
regions of other ctries.						

2. Data in levels with specific effects

	no. of eqs.	(+)	(0)	(-)	avge α_p	avge. t
US states	16	31%	56%	13%	0.004	0.51
Spanish regions	21	95%	5%	0%	0.129	3.99
cross country	4	50%	25%	25%		0.25
regions of other ctries.	11	91%	9%	0%	0.176	9.19

3. Data in differences

	no. of eqs.	(+)	(0)	(-)	avge α_p	avge. t
US states	15	7%	87%	7%	-0.030	-0.58
Spanish regions	8	63%	38%	0%	0.190	2.17
cross country	8	75%	25%	0%	0.121	0.79
regions of other ctries.	2	100%	0%	0%	0.318	2.58

Whenever possible, I have also added the average coefficient of infrastructures or public capital (α_p) and the average *t* ratio for each group of estimates. Average coefficient values should be interpreted with care because different studies use different infrastructure or public capital aggregates. It should also be kept in mind that the two averages reported in the table are generally computed over different sets of regressions. Average *t* ratios are calculated taking into account the signs of these statistics (ie. those of the coefficient estimates) and make use of most of the estimates given in the relevant tables. The exceptions (i.e. the observations that are used to calculate the % of significant estimates but not the average *t* ratio) are two sets of estimates that produce *t* ratios that cannot be directly compared with those of most specifications: the Canning et al results mentioned above, and a few equations where the stocks of infrastructures of neighboring regions are added together with the region's own stock to construct a single infrastructure regressor that mixes spillovers and direct, own-region effects. Finally, there are some equations where infrastructure stocks are disaggregated into several components. In this case, I have taken the largest coefficient and *t* ratio associated with a core infrastructure variable (typically roads), ignoring health or education variables.

The average value of α_p has been calculated using only those equations that produce estimates of this production function parameter using data on either the total stock of public capital, some broad infrastructure aggregate or a core infrastructure indicator. Hence, I have excluded (in addition to the observations not used to calculate average *t*'s) most convergence equations and some equations where infrastructures are disaggregated into components or include the stock of neighboring regions. When I do not report average *t* or average α_p it is because over half of the total observations in the relevant group of estimates are unsuitable to calculate these averages for the reasons I have just discussed.

As can be seen in Table 27, the cross-country evidence is inconclusive, but also very hard to assess due to differences across studies in sample composition and econometric specification and to the likely lack of homogeneity of the cross-country fiscal data. Regional studies should be free of most of these problems, but they also display a mixed pattern of results that at best allows the drawing of rather tentative conclusions. Early positive results with regional samples by Munell and other authors are probably unreliable because they fail to control for specific effects that are quite likely to be there and can generate substantial biases. The elimination of specific effects through differencing often leads to the loss of significance of public capital variables. As I have already noted, however, such specifications can be questioned on the grounds that first-differenced data may contain too much short-term noise to allow for the estimation of a production function. The fact that estimates of the coefficients of private inputs obtained with these data are typically quite unreasonable seems to point in this direction, notwithstanding the results of formal tests that are invoked in some studies to select differenced specifications.

This leaves us with estimations in levels with specific regional effects as the most likely source of reliable results. Focusing on those studies that have followed this approach, the one thing that stands out in Table 27 is the difference in results across the two main regional samples we have considered. Public capital variables are almost always significant in panel data specifications for the Spanish regions, and often insignificant in similar exercises conducted with US data. As Fernald (1999) notes, however, the existing data for the US states starts in 1970, i.e. at approximately the time when the interstate highway system was completed. This author conjectures that this fact may explain the negative results of most US studies. The evidence from Spain and other Mediterranean countries is consistent with this hypothesis, as the per capita stock of infrastructure are much smaller in these countries than in the US, and so are the findings by Mas et al (1996) and Fernald (1999) that the coefficient of infrastructure tends to fall as the sample period is extended.

On the whole, my reading of the evidence is that there are sufficient indications that public infrastructure investment contributes significantly to productivity growth, at least for countries where a saturation point has not been reached. The returns to such investment are probably quite high in early stages, when infrastructures are scarce and basic networks have not been completed, but fall sharply thereafter. Hence, appropriate infrastructure provision is probably a key input for development policy, even if it does not hold the key to rapid productivity growth in advanced countries where transportation and communications needs are already adequately served.

On the other hand, the great disparity of results we find in the literature is worrisome, to say the least, and points to the need for further work in this area. It seems reasonable to interpret some of the results I have reported as an indication of the persistence of various data and econometric problems, as well as of the possible need for a more flexible framework than the one generally used in the literature. First of all, there is in all likelihood an important data problem that may bias estimates of infrastructure coefficients toward zero. Many of the public capital series used in the studies we have reviewed (particularly at the cross-country level) are rather tentative estimates constructed starting from incomplete primary sources and covering relatively short periods. Secondly, it is quite likely that we have not yet fully solved other relevant econometric problems, starting with the possible endogeneity of the regressors. A third factor to consider is that the monetary cost of a given infrastructure may not be a good measure of the productive services it supplies. The usual specifications of the production function assume implicitly that a dollar spent in roads has the same impact in all regions (holding the stocks of other inputs constant). It seems clear that this is not the case, as both construction costs and the productive impact of different types of infrastructures may differ across regions reflecting factors for which most studies do not control, such as orography and climate. As we have seen, there are also some indications in the literature that controlling for this factors and using physical infrastructure indicators helps improve the results somewhat.

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