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IN THE SPANISH ECONOMY

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1. INTRODUCTION

Redistribution targets have provided the main rationale for the growth in public spending in recent years. From 1975 to 1991, the redistribution of income and wealth accounted for 60% of the increase in total non-interest government expenditure. However, public spending, despite its positive effects, is not a free good to society. Poverty relief, state pension funds and unemployment benefits contain two elements of cost. The first is associated with the diversion of resources from directly productive private uses. The second arises from government intervention, which tends to upset price signals and, consequently, efficiency in the allocation of resources. It is difficult to evaluate such costs, although there is some evidence of a negative correlation between public sector size and economic growth (for example, Grossman (1988); Grier and Tullock (1989)). As far as Spain is concerned, the empirical evidence gathered in Raymond (1992) shows that the expansion in current public expenditure explains nearly one quarter of the deceleration in GDP growth in the periods 1960-1974 and 1975-1991.

Several factors can account for this negative relationship and, among them, an inadequate government spending structure can be singled out. In Spain, the size of public spending assigned to social capital -especially infrastructure investment- has traditionally been small. The relationship between public infrastructure and productivity has been addressed in a series of recent papers (Aschauer (1989); Munnell (1990a) and (1990b); Ford and Poret (1991); Berndt and Hansson (1992); Holtz-Eakin (1992); Easterly and Rebelo (1993)). Although results differ according to the methodology and data employed (aggregate or sectoral time series, cross-sections for countries or regions and panel for regions) and the definition used for the public capital variable, most of the evidence gathered suggests a positive relationship between public capital and private productivity. The findings of Bajo and Sosvilla (1992) for the Spanish case point in the same direction, although no distinction is made between public capital as a whole and infrastructure.

This paper examines the effect of public capital accumulation on private-sector productivity in the Spanish economy, with special emphasis on the role of public infrastructure. The starting hypothesis is that

investment in infrastructure, particularly transportation and communications, can have a greater effect on private productivity than other types of investment (construction of government buildings, for example). An empirical approach to this question is possible through the estimation of production functions in which public infrastructure stock is an explanatory variable, using the infrastructure series recently constructed by Argimón and Martín (1993) for the Spanish economy.

The paper is structured as follows. Section 2 highlights certain stylised features that characterize the time variations in private-sector productivity in the period 1964-1990. Public policy aimed at improving infrastructure is among the possible factors that may explain these trends. Section 3 presents the theoretical framework. Section 4 begins with a brief description of the series used and contains econometric estimates of the parameters of an aggregate production function, in which infrastructure is an additional input. Finally, Section 5 summarises the main findings and tentatively evaluates their possible implications. The paper includes two appendices, one on the tests used and the other on causality.

2. PRIVATE-SECTOR PRODUCTIVITY: 1964-1990

The overall effect of the public sector on private sector production cannot be examined without considering the role played by changes in private inputs. One approach to this preliminary analysis is to measure the growth of total factor productivity (TFP), i.e. the growth in production that cannot be attributed to increases in capital and labour inputs. Growth in TFP will, thus, reflect both output growth derived from greater productive efficiency and changes in any other factor that affects the production function and which has not been explicitly taken into account.

A standard way of calculating the rate of change in TFP in the private sector is through the Solow residual, defined as:

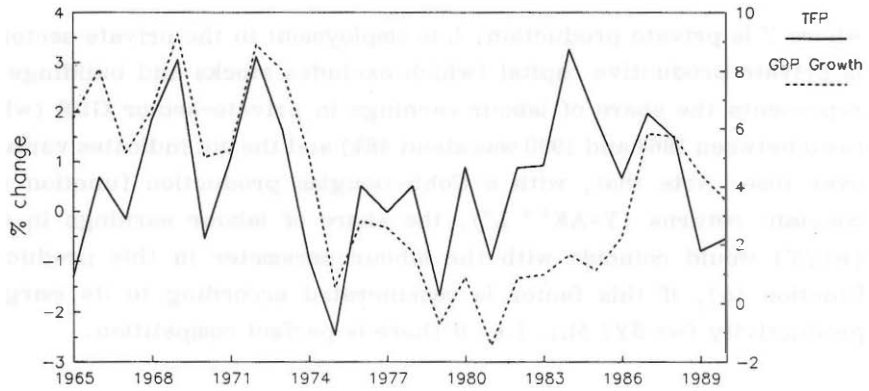
$$\dot{\text{TFP}} = \frac{\dot{Y}}{Y} - \alpha \frac{\dot{L}}{L} - (1-\alpha) \frac{\dot{K}}{K} \quad (1)$$

where Y is private production, L is employment in the private sector, K is private productive capital (which excludes stocks and buildings), α represents the share of labour earnings in private-sector GDP (whose mean between 1964 and 1990 was about 46%) and the dot indicates variation over time. Note that, with a Cobb-Douglas production function with constant returns ($Y=AK^{1-\alpha} L^\alpha$), the share of labour earnings in GDP (wL/Y) would coincide with the labour parameter in this production function (α), if this factor is remunerated according to its marginal productivity ($w= \delta Y/ \delta L$), i. e. if there is perfect competition.

Figure 1 shows the rate of change in private-sector TFP (left-hand scale) and compares it with the growth rate of private GDP (right-hand scale) obtained from annual National Accounts data. Figure 2 shows TFP in the private sector and in the economy as a whole. The differences are seen to be very small. The range of variation of TFP in the private sector is broader than in the overall economy, possibly indicating that, in many cases, the public sector has played a compensating role.

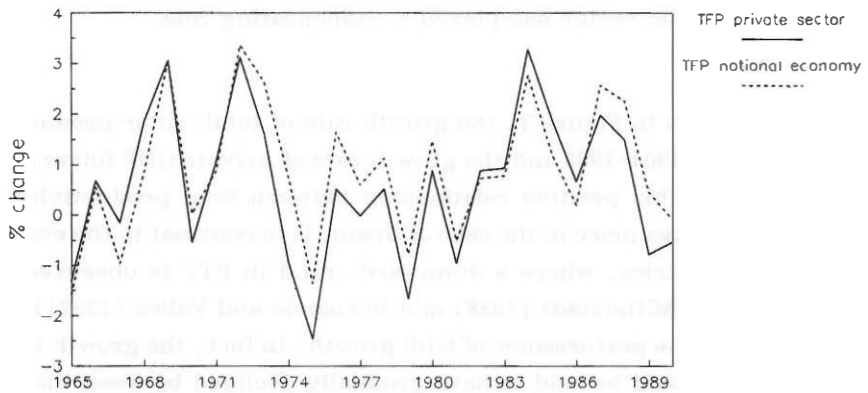
As shown in Figure 1, the growth rate of total factor productivity for the period 1964-1990 and the growth rate of private GDP follow a very similar path. This positive relationship between total productivity and GDP, which takes place in the case of Spain, is in contrast to the evidence for other countries, where a downward trend in PTF is observed (see Englander and Mittelstädt (1988) and Hernando and Vallés (1992)), with no relation to the performance of GDP growth. In fact, the growth in total productivity could be said to have gradually declined between the 1970s and the early 1980s. However, once the economy started to recover, TFP grew significantly, and it did not fall again until there were clear symptoms that the expansive cycle was over. This behaviour of productivity holds even if other methods for computing the coefficients in (1) are applied. For instance, if the parameter α is obtained from the econometric estimation of a Cobb-Douglas production function with

Figure 1
Total Factor Productivity
Private Sector



Left Axis, TFP Growth
 Right Axis, GDP Growth

Figure 2
Total Factor Productivity
Private Sector and National Economy



constant returns to scale, the time variation in TFP shows the same pattern, thus confirming its apparent procyclicality.

Labour productivity, as reflected in Figure 3, has also followed a very similar course to that of private GDP, except for the final years of the sample. Since 1985, the growth gap between these two variables has widened, with the growth in productivity tending to stabilise at low levels. Consequently, a decline in labour productivity is evident during the period 1964-1990. The procyclicality observed in the first part of the sample is explained, in the literature on this topic, by the existence of technological shocks associated with the cycle, increasing returns or labour hoarding. Here again, the lack of a recovery in the growth rate of labour productivity in the final period stands out in comparison with the evidence of procyclicality in the productivity of labour found for other countries (see Hernando and Vallés (1992)).

Capital productivity, whose variation over time is shown in Figure 4, also changes its pattern of behaviour in the early 1980s. During the first few years of the sample, when the economy was steadily growing, the contribution of capital was small but, when GDP reduced its growth rate, the gap between the two variables narrowed. However, the recovery in the growth rate of GDP, which started in 1981, was accompanied by a noteworthy recovery in capital productivity growth.

To sum up, we have three stylised facts in relation to productivity: procyclicality in the growth of TFP throughout the period 1964-1990 (evidence which contrasts with the systematic decline observed in other countries), countercyclicality in labour productivity in the 1980s (which contrasts with the procyclicality of labour observed in other countries), and an increase in capital productivity at the end of this period, whereas previously the pattern had essentially been one of stability. Therefore, a series of stylised features in the performance of the different types of productivity can be argued to have consolidated in the 1980s.

A series of factors evident in past years might partly explain the observed change in the behaviour of the productivity of the two inputs. First, the public sector made a notable effort to renew and improve the

Figure 3
Labour Productivity
Private Sector

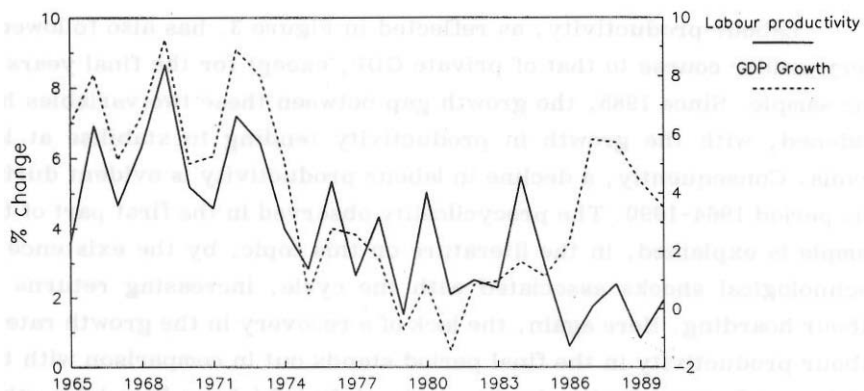
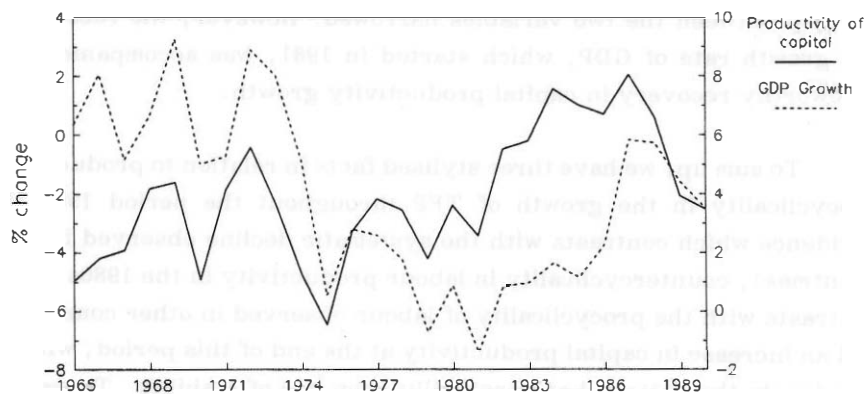


Figure 4
Productivity of Capital
Private Sector



infrastructure stock (see Figure 5). In fact (see Table 1, column 1), government investment has been rising in recent years, reaching around 5% of GDP. This improvement may have been the driving force behind the increase in capital productivity, at a time when the capital/labour ratio stood at its lowest values.

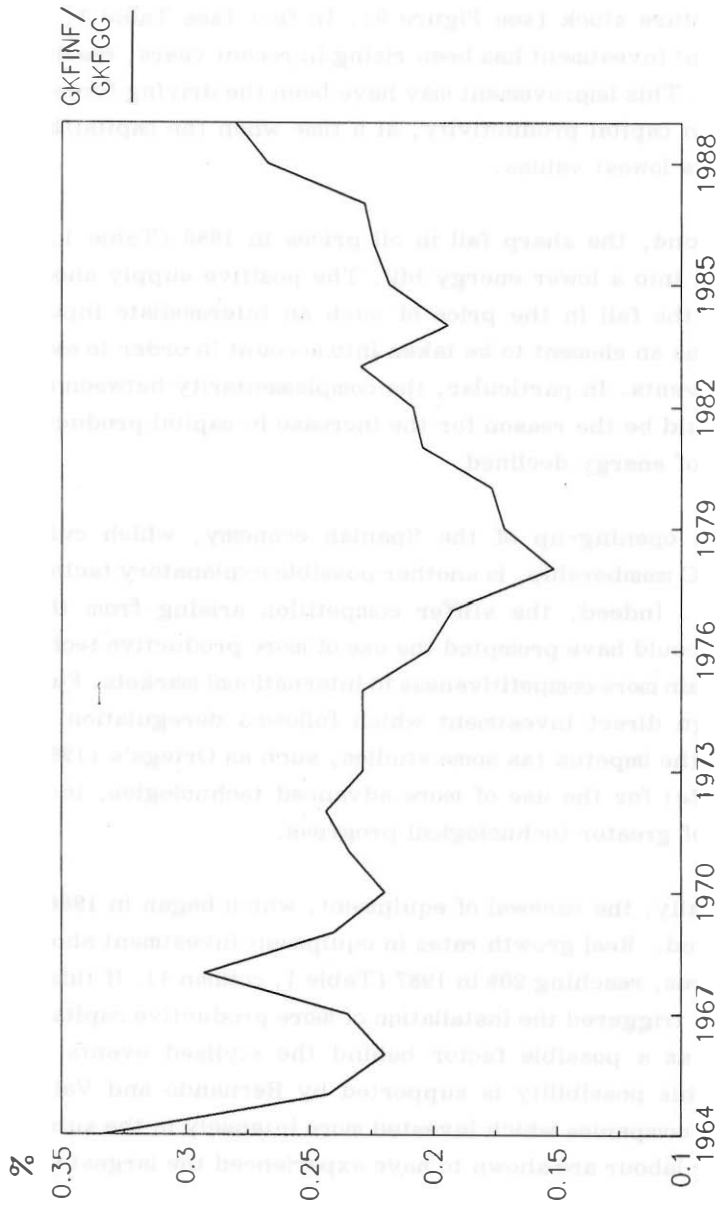
Second, the sharp fall in oil prices in 1986 (Table 1, column 2) translated into a lower energy bill. The positive supply shock brought about by the fall in the price of such an intermediate input could be regarded as an element to be taken into account in order to explain these stylised events. In particular, the complementarity between energy and capital could be the reason for the increase in capital productivity when the price of energy declined.

The opening-up of the Spanish economy, which culminated in Spanish EC membership, is another possible explanatory factor (Table 1, column 3). Indeed, the stiffer competition arising from this greater openness could have prompted the use of more productive technologies in order to gain more competitiveness in international markets. Furthermore, the foreign direct investment which followed deregulation could have provided the impetus (as some studies, such as Ortega's (1992), appear to conclude) for the use of more advanced technologies, incorporating elements of greater technological progress.

Finally, the renewal of equipment, which began in 1986, has been unparalleled. Real growth rates in equipment investment showed double digit figures, reaching 20% in 1987 (Table 1, column 4). If this renewal of equipment triggered the installation of more productive capital, it may be regarded as a possible factor behind the stylised events of the last period. This possibility is supported by Hernando and Vallés (1992), where the companies which invested more intensely in the substitution of capital for labour are shown to have experienced the largest increases in TFP.

The only stylised feature which does not fit with any of the aforementioned trends is the countercyclical pattern of the growth rates of labour productivity. However, besides its coincidence with the

Figure 5
Ratio Infrastructure Investment/Total Investment



GKFINF: Government investment in Infrastructure
 (Central, Regional and Local)
 GKFGG: General Government Total Investment

TABLE 1
Trends in total productivity in the period 1985-1991:
Explanatory factors

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Public investment (% of GDP)	Relative energy prices ^a	Trade openness (% of GDP) ^b	Equipment investment (real rate of growth) ^c	Unemployment rate (%)	Proportion of temporary employment (%)	First-time job seekers (% total unemployment) ^e
1964/1973	2.8	0.66	22	11	1.5	-	-
1974/1984	2.4	1.76	30	-1	9.9	-	-
1985	3.5	2.21	36.1	7.9	21.6	-	40
1986	3.5	1.00	37.6	12.5	21.2	-	41.1
1987	3.3	0.89	40.1	20.5	20.5	18.15	37.9
1988	3.7	0.71	41.9	16.3	19.5	23.29	35.4
1989	4.4	0.76	44.1	12	17.3	26.38	31.7
1990	4.8	0.74	45.1	1.4	16.3	30.3	28.7
1991	5.1	0.66	48.0	-2.5	16.3	32.3	25.2

Sources: Instituto Nacional de Estadística, Banco de España and authors' estimation.

(a) Relative price of energy imports. Source: Data base of MOISES model (base 1986).

(b) Real exports plus real imports, as a percentage of real GDP. Source: Data base of MOISES model (base 1986).

(c) Sources: Instituto Nacional de Estadística and Banco de España. For the period 1964 to 1984, the average growth of real private productive investment, taken from the data base of the MOISES model (base 1986), is shown.

(d) Fixed-term contracts in relation to total contracts. Source: Encuesta Población Activa (EPA)

(e) Source: Encuesta Población Activa (EPA)

intuitive hypothesis which states that labour productivity should be countercyclical (as a result of diminishing marginal productivity), this isolated event could be the consequence of job destruction in former years, implying that the skills of those entering the labour market in the expansive stage were relatively poorer. In fact, it was first-time job seekers who fuelled the recovery in employment (see columns 5 and 7 of Table 1) which, furthermore, was especially strong in temporary contracts (Table 1, column 6). Therefore, there are factors, basically of an institutional nature, which can explain why labour productivity showed an atypical pattern in Spain.

To sum up, we have pointed out some stylised events in the final part of the sample, such as the acceleration in TFP and in capital productivity, and several factors which can account for these events. This paper will concentrate on the role played by the public capital stock and, in particular, by communications and transportation infrastructure. To this end, we will use as our analytical framework an aggregate production function which includes public capital as an additional input. The following section formally presents this framework.

3. ANALYTICAL FRAMEWORK

The starting point is an aggregate production function such as:

$$Y_t = A_t f(L_t, K_{pt}, K_{gt}) \quad (2)$$

where Y_t is aggregate private production of goods and services, L_t is aggregate employment in the private sector, K_{gt} represents public infrastructure stock, K_{pt} private productive capital stock and A_t is a measure of technological progress.

The production function is assumed to be of the Cobb-Douglas type:

$$Y_t = A_t L_t^\alpha K_{pt}^\beta K_{gt}^\gamma \quad (3)$$

where:

$$A_t = A_0 \cdot e^{gt} \quad (4)$$

with g representing the growth rate of exogenous technological progress. Taking into account (4), equation (3), in its logarithmic version, would be

$$y_t = a_0 + gt + \alpha l_t + \beta k_{pt} + \gamma k_{gt} \quad (5)$$

where small case letters denote the variables in logarithms.

Equation (5) is the relationship whose parameters are sought. Reparametrisation yields:

$$(y - k_p)_t = a_0 + gt + \alpha(1 - k_p)_t + \gamma(k_g - k_p)_t + (\alpha + \beta + \gamma - 1)k_{pt} \quad (6)$$

Therefore, if we estimate:

$$(y - k_p)_t = b_0 + b_1 t + b_2(1 - k_p)_t + b_3(k_g - k_p)_t + b_4 k_{pt} \quad (7)$$

the lack of statistical significance of the coefficient b_4 will indicate the existence of constant returns in all private inputs, including public capital. Similarly, to test the existence of constant returns in private inputs, the equation to be estimated is:

$$(y - k_p)_t = c_0 + c_1 t + c_2(1 - k_p)_t + c_3 k_{gt} + c_4 k_{pt} \quad (8)$$

where the result that the coefficient c_4 is not statistically different from zero would indicate the existence of constant returns in private inputs (since $c_4 = \alpha + \beta - 1$). Note that this implies the existence of increasing returns to scale in all inputs, both private and public.

Generally speaking, the existence of increasing returns to scale raises some problems. However, here, it does not, since the level of

public capital is set by an agent -the government- which finances the services entailed in this capital through compulsory revenue (and not by charging the cost to potential users). Hence, the private sector does not determine the level of public input which, on the contrary, is exogenous. In this sense, the equality of private and public capital with respect to their marginal productivity net of depreciation does not occur automatically. This would happen if there were a planner who would optimise social welfare and set the level of public capital so as to bring about this equality.

An alternative approach would be to start from the definition of total factor productivity set forth in expression (1), attempting to explain its variations on the basis of a regression such as:

$$\Delta \text{TFP} = a_0 + a_1 \Delta k_{gt} + a_1 \cdot \text{other variables} \quad (9)$$

where Δ is the first-difference operator, and k_g is the variable of public capital. The inclusion of public capital would attempt to reflect the effect of this variable on private productivity, although the coefficient obtained would no longer have the structural interpretation of the coefficient of public capital in the private production function.

Several theories can suggest which additional variables could be used in (9). First, insofar as the performance of total factor productivity is affected by disturbances in demand, the use of capacity utilisation (in first differences) would allow the effect of these shocks to be monitored. Second, certain theories state that there is some interaction between the accumulation of capital and technological progress. Thus, if technological progress is incorporated into new capital, we should find that the accumulation of capital has a positive effect on productivity (i.e. that in using $\Delta(k-1)$ or Δk as variables that reflect the renewal of capital goods in expression (9), these variables become statistically significant). Another theory that takes into account this interactive effect is the vintage model, which shows that, if the new capital is more productive than the capital already installed, any acceleration in the accumulation of capital should be related to increases in productivity (i.e. in (9) the

variables $\Delta^2(k-1)$ or $\Delta^2(k)$, which measure the acceleration of capital accumulation, should be statistically significant: see Wolff (1991)).

The following section presents the main empirical results obtained from both approaches.

4. PRIVATE PRODUCTIVITY AND PUBLIC CAPITAL

The data base used in the estimation of private production functions is MOISEES, at constant 1980 prices, (see Molinas et al. (1990)) except for the data for private-sector employment, which is the series presented in García Perea and Gómez (1993) that is constructed taking into account National Accounts criteria. However, the lack of a series of transportation and communications infrastructure stock -which is only one part of public capital- necessitated the construction of the series.

A detailed description of the computation of the infrastructure figures can be found in Argimón and Martín (1993). However, a summary of such a description is worthwhile. The main statistical source for the construction of these stock series was the General Government Accounts (CAP), available from 1958 to 1989, with a much more detailed information for the more recent period than for the sixties and seventies. Similarly, the volume of information is much larger at the Central Government level than for regional and local governments. For the latter, several hypotheses had to be applied for the breakdown of public spending into transportation and communications infrastructure and other investments (given their negligible quantitative importance, Social Security and autonomous administrative agencies were not taken into account). The stock series is constructed from the real investment series in transportation and communications infrastructure, assuming a 5% rate of depreciation (the same rate used in Corrales and Taguas (1989) for the construction of the public capital series and in the MOISEES data base). Altogether, four series were calculated: one for General Government, based on national accounts (GG(NA)); another for Central Government, also in terms of national accounts (CG(NA)); another one for Central Government, but in terms of public accounts (CG(PA)), (among other

differences, this includes defence spending, most of which is recorded under public consumption in national accounts); and a final one for Central Government, which, in addition to investment, includes capital transfers (CG (PA with KT)). These series not only differ in level but also over time (see Argimón and Martín (1993)).

If the stock series on infrastructure and communications of General Government is compared with the public capital series used in the MOISEES model (see Baiges, Molinas and Sebastián (1989)), infrastructure stock is seen to represent around 23% of total public capital. If only total capital stock of Central Government and Regional and Local Governments is taken into account, transportation and communications infrastructure accounts for 38% of the total. The quantitative difference between both series and their different qualitative components would require the testing of the hypothesis that public investment in this type of capital (transportation and communications infrastructure) has a noteworthy positive effect on productivity in the private sector.

To this end, we shall estimate the Cobb-Douglas production functions which, assuming constant returns to scale, can be expressed in logarithms as:

$$(y - k_p)_t = a + b(1 - k_p)_t + c(k_g - k_p)_t \quad (10)$$

where small case letters denote variables in logarithms, y_t is private GDP (calculated as real GDP at factor cost less public labour earnings and minus General Government consumption of fixed capital), k_{pt} is private capital stock, l_t is employment in the private sector and k_{gt} is stock of public capital or of public-sector infrastructure.

Before interpreting the regression results, which impose constant returns to scale for all inputs, it should be noted that estimations were made without applying this constraint. The low value (near zero) of the coefficients for all Central Government series, as well as the low value of their t-ratios, in the regressions which test the existence of constant returns in private inputs and in all inputs (equations (7) and (8)),

support the conclusion that there are constant returns in all inputs. However, for the General Government series, there are signs of increasing returns. Nonetheless, when this test for the General Government series is performed in an error correction model, estimated by non-linear least squares (NLS), the existence of constant returns cannot be rejected. As far as these regressions are concerned, it should be borne in mind that, even though the distribution of the coefficients is not the standard one, we do know that computed standard errors have, in general, a lower value than the true standard errors. A t-ratio below 2 may, therefore, be an indication of the minor role played by the variable whose statistical significance is being questioned.

Table 2 shows the results obtained from the estimation of equation (10) for the different infrastructure series constructed. Two estimations are presented, one obtained by OLS and another by non-linear least squares (NLS). The specification of the latter is represented by an error correction mechanism (ECM) model, whose general expression is given by:

$$\begin{aligned} \Delta(y-k_p)_t = a' + b'\Delta(1-k_p)_t + c'\Delta(k_g - k_p)_t \\ + e\left[(y-k_p)_{t-1} - b''(1-k_p)_{t-1} - c''(k_g - k_p)_{t-1}\right] + d\Delta(y-k_p)_{t-1} \end{aligned} \quad (11)$$

Appendix I gives the details on the variables used in each of the estimated equations. Given the non-stationarity of the data used, cointegration statistics are presented. In the estimation by OLS, this statistic is the ADF, and in the NLS estimation it is the t-ratio of the ECM coefficient (see Appendix I).

As Table 2 shows, in all estimations (both for the different series and using OLS and NLS) the labour coefficient ranges between 0.2 and 0.3. With constant returns to scale and perfect competition, the labour parameter matches the share of labour in the product. In fact, if each factor is remunerated according to its marginal productivity, b , the coefficient estimated for $(1-k_p)$ should be such that $b = (\partial Y / \partial L)(L/Y)$. However, the estimated coefficient is much lower than the share of labour earnings in GDP (whose mean is 46% of GDP). Nonetheless, if we consider that the product is distributed solely between the two competitive inputs

TABLE 2
Private production function with different public infrastructure stock
(1964-1989)

OLS				
Infrastructure variable used				
	GG (NA)	CG (NA)	CG (PA)	CG (PA, with KT)
	(1)	(2)	(3)	(4)
$1-k_p$	0.35 (45.76)	0.22 (10.86)	0.29 (28.14)	0.30 (22.47)
$k_g - k_p$	0.21 (6.52)	0.60 (7.74)	0.59 (8.72)	0.49 (6.07)
ADF	-2.38	-3.79*	-4.14**	-2.6
DW	0.71	1.28	1.47	0.78
SE	0.016	0.014	0.013	0.017
NLS				
Public infrastructure variable used				
	GG (NA)	CG (NA)	CG (PA)	CG (PA, with KT)
	(5)	(6)	(7)	(8)
$1-k_p$	0.36 (10.46)	0.20 (7.52)	0.31 (20.26)	0.27 (12.05)
$k_g - k_p$	0.21 (2.83)	0.71 (8.60)	0.59 (8.34)	0.67 (7.30)
ECM	-0.5 (-2.79)	-0.77 (-6.00)**	-0.79 (-5.56)**	-0.68 (-5.68)**
DW	1.97	2.22	2.30	2.09
SE	0.012	0.08	0.08	0.08

Results of the estimation of equation (10) by OLS and equation (11) by NLS. The values of the coefficients under NLS are those obtained within the error correction mechanism and represent long-run relationships.

Note: The values in brackets are t-ratios. The variables are in logarithms. DW is the Durbin-Watson autocorrelation test; ADF is the statistical value of t in the test of stationarity of the residuals of the estimated equation, which is described in Appendix I; ECM is the estimated value of e in equation (11), whose significance is interpreted as a cointegration test, as explained in Appendix I, which also lists the variables included in the short term of each equation estimated; SE is the standard error of the regression.

** 5% level of significance.

* 10% level of significance.

(labour and private capital), since public capital is not remunerated in relation to its marginal productivity, then labour's share in output - calculated as the ratio between the labour coefficient (c in equation (10)) and the sum of c and the private capital coefficient ($1-c-b$) - rises to 44%, according to the figures in column (1). The main difference between the results of the NLS and OLS estimations is that the public capital coefficient obtained under NLS is generally higher than that obtained through OLS. With respect to the fit obtained with the different series constructed, the General Government stock series has the most unsatisfactory results (in relation to the cointegration tests).

The most remarkable result in Table 2 is that, for all series, the productivity of public infrastructure capital is higher than that of the private sector. In addition, the Central Government infrastructure series constructed according to the public accounting criterion, and including capital transfers, presents a worse fit than the series which does not include capital transfers.

As far as comparison of the national and the public accounting series of Central Government infrastructure is concerned, the estimated coefficients, the fit of the equation and the ADF and ECM statistics are very similar to each other (columns (2), (3), (6) and (7)). It should be recalled that the public accounting series includes defence spending, whereas national accounts does not. The only remarkable difference in the estimation is that the labour coefficient is, in the case of the public accounting series, somewhat higher (0.29 compared with 0.22 in the OLS estimation). In both cases, it is very low since under the hypothesis of perfect competition and constant returns, this coefficient should match the participation of wages in private GDP. As earlier mentioned, one possible explanation for this low value may lie in the nature of public capital, whose level is determined by the public sector, and cannot be controlled by the private one, nor is it remunerated according to its marginal productivity. In this sense, if the private labour and capital coefficients of columns 1 to 3 in Table 2 are again compared (i.e. if the labour coefficient is compared with the sum of the coefficients for private labour and capital), labour would have a relative participation of 44% when the General Government series is used (compared with 56% for private

capital), 55% with the Central Government-national accounts series and 74% with the Central Government-public accounts series. In any event, this low value of the labour coefficient is in line with the results obtained when the public capital series is used (see Bajo and Sosvilla (1992)). It should be pointed out that the output elasticity of public capital is lower for the Central Government-national accounts series, which does not include military spending, than for the Central Government-public accounts series.

One of the contributions of this paper is the use of an infrastructure series as the relevant public capital variable. Thus, we can compare the results obtained from the series of public infrastructure stock with the ones which only consider public capital as a whole, without distinguishing between its different components. We will estimate the production function presented in equations (10) and (11), but including the rest of public capital as an additional regressor⁽¹⁾. This regression enables the test of the preliminary hypothesis that, first, the distinction between public investment in infrastructure and in the rest of public capital is relevant; and, second, that public infrastructure is more productive than the rest of public sector capital. As may be seen in Table 3, in all the regressions the coefficient of the rest of public capital ($r_g - k_g$) is very close to zero, with the exception of the General Government series. The empirical evidence for this series is not conclusive because whereas the OLS estimation would evidence that this coefficient is very close to zero, the NLS estimation would support the conclusion that the coefficient for the rest of public capital is the same as that for public infrastructure. Once again, it is the General Government series which has the worst results in relation to the cointegration tests.

To sum up, the empirical evidence reflected in Table 3 shows that not only the output elasticity to public infrastructure is higher than the output elasticity to the rest of public capital, but also that the rest of

⁽¹⁾ The public capital series is constructed according to the same criteria and for the same agents used in the construction of the infrastructure series.

TABLE 3

Private production function with infrastructure stock and all other
public stock
(1964-1989)

OLS				
Infrastructure variable used				
	GG (NA)	CG (NA)	CG (PA)	CG (PA, with KT)
	(1)	(2)	(3)	(4)
$1-k_p$	0.38 (19.33)	0.21 (9.77)	0.27 (15.59)	0.45 (4.92)
$k_g - k_p$	0.17 (3.84)	0.62 (7.78)	0.65 (8.49)	0.26 (1.55)
$r_g - k_p$	0.04 (1.47)	-0.01 (0.98)	-0.03 (1.50)	0.12 (1.66)
ADF	2.67	3.71	4.57**	2.61
DW	0.86	1.31	1.65	0.76
SE	0.016	0.015	0.013	0.017
NLS				
	GG (NA)	CG (NA)	CG (PA)	CG (PA, with KT)
	(5)	(6)	(7)	(8)
$1-k_p$	0.41 (10.92)	0.20 (6.30)	0.32 (13.55)	0.36 (4.36)
$k_g - k_p$	0.11 (1.42)	0.71 (7.70)	0.56 (6.33)	0.51 (3.22)
$r_g - k_p$	0.10 (2.19)	0.002 (0.02)	0.02 (0.67)	0.069 (1.16)
ECM	-0.67 (3.58)*	-0.77 (5.71)**	-0.78 (5.38)**	-0.71 (5.81)**
DW	2.03	2.20	2.35	2.12
SE	0.011	0.008	0.009	0.008

r_g : All other public capital

** 5% level of significance.

* 10% level of significance.

See note to Table 2.

public capital does not appear to affect productivity in the private sector, either positively or negatively. This result could be interpreted as indicating that this part of public capital (which includes, among other elements, government buildings) can be regarded more as an element of public consumption than as productive capital.

Another finding in Table 2 which, in principle, is not satisfactory is the bad fit (in terms of cointegration tests) of the estimation of the production function when the General Government series is used, compared with the good fit of this function when the series constructed for the Central Government are used. Looking for the possible causes of this result, we conducted stability tests for the equation, and also analysed, separately, the series on General Government, Central Government (national accounts) and Regional and Local Governments (taken as the difference between the two), with the results presented in Table 4. As can be seen, they evidence the existence, in the General Government series, of a structural break in 1985, when the transfer of powers to regional governments was intensified and consolidated. In particular, the output elasticity to the Regional and Local Government series seems to indicate the irrelevance of these infrastructure systems to private sector productivity.

Table 5 presents the results obtained when the production function is estimated separating Central Government infrastructure from that of Regional and Local Governments. It shows that both estimations - the one including the joint infrastructure stocks and the estimation including only Regional and Local Government infrastructure - evidence that territorial government infrastructure apparently does not affect productivity in the private sector.

To sum up, the overall evidence reflected in Tables 4 and 5 confirms that the bad fit of the General Government series stems from the different output elasticity to Central Government and to Regional and Local Government infrastructure, the latter not seeming significantly to affect private-sector productivity. The rationale for this apparent irrelevance of territorial infrastructure with respect to productivity in the private sector may be twofold. First of all, it could be the

TABLE 4

Structural change in the production function

	General Government		Central Government		Regional and Local Government	
	(1)	(2)	(1)	(2)	(1)	(2)
1985	0.67 (4.12)	3.25** (3.11)	0.71 (7.10)	1.38 (3.11)	0.07 (0.63)	0.32 (3.11)
1989	0.21 (2.83)	-	0.71 (8.60)	-	0.02 (0.67)	-

The first column of the table shows the last year of the sample used for the estimation of the function:

$$\Delta(y-k_p)_t = \alpha_0 + \alpha_1 \Delta(1-k_p)_t + \alpha_2 [(y-k_p)_{t-1} - \alpha_3(1-k_p)_{t-1} - \alpha_4(k_g-k_p)_{t-1}] + \alpha_5 \Delta(y-k_p)_{t-1}$$

Column (1) presents the estimated value of α_4 , and, in brackets, the t statistic. Column (2) shows the value of the F test of structural change, and, in parenthesis, the critical value at the 5% level of significance.

** 5% level of significance.

TABLE 5

Role of infrastructure stock of Regional and Local Governments

	(1)	(2)
	NLS	
(1-k _p)	0.46 (9.84)	0.21 (7.12)
(k _{tt} -k _p)	0.02 (0.67)	0.007 (0.88)
(k _e -k _p)	-	0.67 (7.23)
ECM	-0.30 (-2.08)	-0.82 (-5.77)**
R ²	0.697	0.868
SE	0.012	0.008

Results of the estimation by non-linear least squares of the equation:

$$\Delta(y-k_p)_t = \alpha_0 + \alpha_1 \Delta(1-k_p)_t + \alpha_2 [(y-k_p)_{t-1} - \alpha_3(1-k_p)_{t-1} - \alpha_4(k_{tt}-k_p)_{t-1} - \alpha_5(k_e-k_p)_{t-1}] + \alpha_6 \Delta(y-k_p)_{t-1}$$

where α_5 has been assumed to be zero in column (1).

The values of the coefficients are those obtained within the error correction mechanism. ECM is the value of the coefficient α_2 .

k_{tt}: logarithm of infrastructure stock of regional and Local Governments.

k_e: logarithm of Central Government infrastructure stock.

consequence of measurement errors in the General Government variable (in its Regional and Local Government component). In fact, the statistical sources available for regional and local public finances (incomplete and less detailed), the exclusion of Social Security spending in infrastructure (with little relevance in this case) and of autonomous government agencies (with a low level of investment but unquestionably higher than in the case of Social Security), can lead to measurement errors in the overall General Government variable which would affect the result obtained.

Second, this result may also be warranted by the transfer of powers to regional governments in the 80s. Indeed, the instability observed since 1985 in the General Government series is a reflection of the finding that the output elasticity to the infrastructure investment of Regional Governments is lower than that of Central Government. Moreover, both the different nature of the goods in which Regional Governments invest and the possible lower efficiency in the use of resources, due to their relative management inexperience, may lead to the different estimated output elasticities. However, the analysis does not provide for the separation of these possible explanations which, as a result, are tentative and, therefore, must be interpreted with caution.

Finally, the results obtained under the TFP approach, which implies the estimation of equation (9) by OLS, are found in Table 6. As can be seen, we have gathered evidence that the capital utilisation ratio (lcu , which captures the existence of demand shocks) and public infrastructure stock (with the series for General Government and national and public accounts Central Government) play an important role in explaining variations in total factor productivity. In contrast, the theory of interaction between capital accumulation and technological progress was not empirically confirmed in either of its two versions. Thus, the variable $\Delta(k-1)$ was neither significant nor did it have the proper sign. The same can be said of the vintage effect. A test of the interaction between capital and technological progress with the variables Δk and $\Delta^2 k$ was attempted, with negative results in both cases, since they did not prove to be statistically significant.

TABLE 6

TFP approach

Exogenous V.	GG (NA)	CG (NA)	CG (PA)
Δk_{gt-1}	0.12 (3.19)	0.13 (2.89)	0.11 (2.90)
Δlcu_t	0.30 (3.23)	0.31 (3.25)	0.31 (3.25)
R^2	0.39	0.36	0.36
SE	0.012	0.012	0.012
DW	1.84	1.69	1.71

Results of the estimation of the equation:

$$\Delta TFP_t = a_1 \Delta k_{gt-1} + a_2 \Delta lcu_t$$

t statistics in brackets; R^2 , determination coefficient; SE, standard error of the regression; DW, Durbin-Watson statistic.

The performance of the three series is very similar, as evidenced in the last three columns of Table 6. In fact, the coefficient for infrastructure, when one period lag is included, ranges between 0.13 and 0.11, and the fit of the equation, measured by the standard error, is the same in all three cases (0.012). This makes it difficult to decide in favour of any of the three specifications. It should be recalled that the evaluation of these results is not the same as when a production function is estimated, since in the productivity approach there is no structural interpretation of the estimated coefficients. Furthermore, the analysis of total productivity, as reflected in Table 6, implies estimating in differences, whereas the production function approach leads to an estimation in levels.

Therefore, the results of these regressions support the positive interaction between public infrastructure capital and productivity in the private sector. However, given the lack of evidence of a relationship between capital accumulation and technological progress (which contradicts the empirical findings available for other countries: see Wolff (1991)), these results should be viewed as preliminary.

In short, the construction of a series of public infrastructure stock allowed for the test of its effect on private-sector productivity. A substantial positive effect is observed, higher than in the rest of public capital (which, in fact, does not seem to affect productivity in the private sector). Furthermore, the output elasticity to investment in infrastructure is higher than the elasticity to private productive capital. However, the output elasticity to public capital in infrastructure is rather high, although in line with the results of similar studies in other countries (see Munnell (1990a) and Table 7). Nonetheless, it should be recalled that the stochastic characteristics of the time series used (their non-stationarity) invalidate the use of standard errors to analyze the distribution of the estimated parameters. Therefore, the output elasticity to infrastructure may not be significantly different from that of private capital.

The two Central Government series based on national and public accounts produce the best fit of the four infrastructure series considered. Even though the better result of the Central Government series can be explained by the existence of measurement errors in the General Government variable, the increasing share of Regional and Local Governments in public investment, including spending on infrastructure, implies that the relatively worse behaviour of this General Government series is one of the most unsatisfactory aspects of the empirical evidence gathered. Although there are several possible explanations for the problems of stability arising in the use of this series, these results should be regarded as provisional and require a more detailed study of the potential problems hidden behind such instability.

Finally, the TFP approach provides further evidence of the relationship between public infrastructure capital and private productivity. Unlike the production function approach, there are no apparent differences between the series, either in the estimated coefficients or in the fit of the equations, as regards their explanatory power for the variations in total factor productivity.

A frequent criticism of this literature is that the observed empirical relationship between infrastructure and private investment may be the

result of a reversed causality scheme, whereby a higher level of development entails higher infrastructure spending. Appendix 2 contains the causality tests performed, which support the result that the direction of causality is the expected one: an increase in public transportation and communications infrastructure increases (causes) productivity in the private sector.

5. CONCLUSIONS

Private-sector productivity in Spain exhibits features distinguishing it from observations in other countries. In particular, in the boom period which began in 1985, there was an acceleration in total factor productivity (TFP), i.e. in that part of the product which is not attributable to mere increases in the quantities of the labour factor or private productive capital.

This paper has attempted to ascertain which part of the recent variation in this residual factor can be explained by trends in public transportation and communications infrastructure. The observed acceleration of TFP has taken place at the same time as a stronger public investment drive centered, to a large extent, on the renewal of transportation and communications infrastructure.

The estimation of a Cobb-Douglas production function for the private sector, which includes infrastructure stock as an additional input, results in not being able to reject the hypothesis that public infrastructure played an important role in the acceleration of TFP. In line with the available international evidence, the output elasticity of public infrastructure is much higher than the output elasticity of private productive capital, a comparison which holds even when possible omitted variables are included or when the regression is run by subsamples. Furthermore, empirical evidence shows that the output elasticity of public infrastructure is much higher than the output elasticity with respect to overall public capital (which includes, in addition to infrastructure, other investment such as government buildings, etc.). Finally, we find that the

direction of causality is the expected one, i.e. the observed relationship does not result from a higher level of development being necessarily accompanied by a higher demand for public spending in infrastructure.

As to the comparison of these results with the existent empirical literature on the subject, the elasticity of output to infrastructure obtained in this paper can be observed to lie in the upper band of the estimations found in other countries (see Table 7). However, it is similar to the elasticity estimated in studies that use time series at a national level. The lower value of the output elasticity obtained in other works that use more disaggregated data (states, regions, towns) may be due to the loss of the externality effect which public capital at a regional level has on private productivity in other regions. Moreover, the definition of infrastructure employed in those studies is broader than the one used in this paper, thus possibly affecting the estimation of the elasticity of output to public infrastructure. Whereas many studies are based on public capital as a whole, or take all public capital into consideration except defence spending, this paper sets public investment in

TABLE 7
Estimations of the elasticity of output to public capital

Author	Level of Aggregation	Specification	Elasticity
Aschauer (1989)	National	Cobb-Douglas, levels, logs	.39/.56
Holtz-Eakin (1988)	National	Cobb-Douglas, levels, logs	.39
Munnell (1990a)	National	Cobb-Douglas, levels, logs	.34
Costa et al. (1987)	States	Translog, levels	.20
Eisner (1991)	States	Cobb-Douglas, levels, logs	.17
Mera (1973)	Regions of Japan	Cobb-Douglas, levels, logs	.20
Munnell (1990b)	States	Cobb-Douglas, levels, logs	.15
Holtz-Eakin (1992)	States	Cobb-Douglas, levels, logs	.20/0
Eberts (1986), (1990)	Towns	Translog, levels	.03
Argimón et al. (1993)	National	Cobb-Douglas, levels, logs	.60

Source: Munnell (1992).

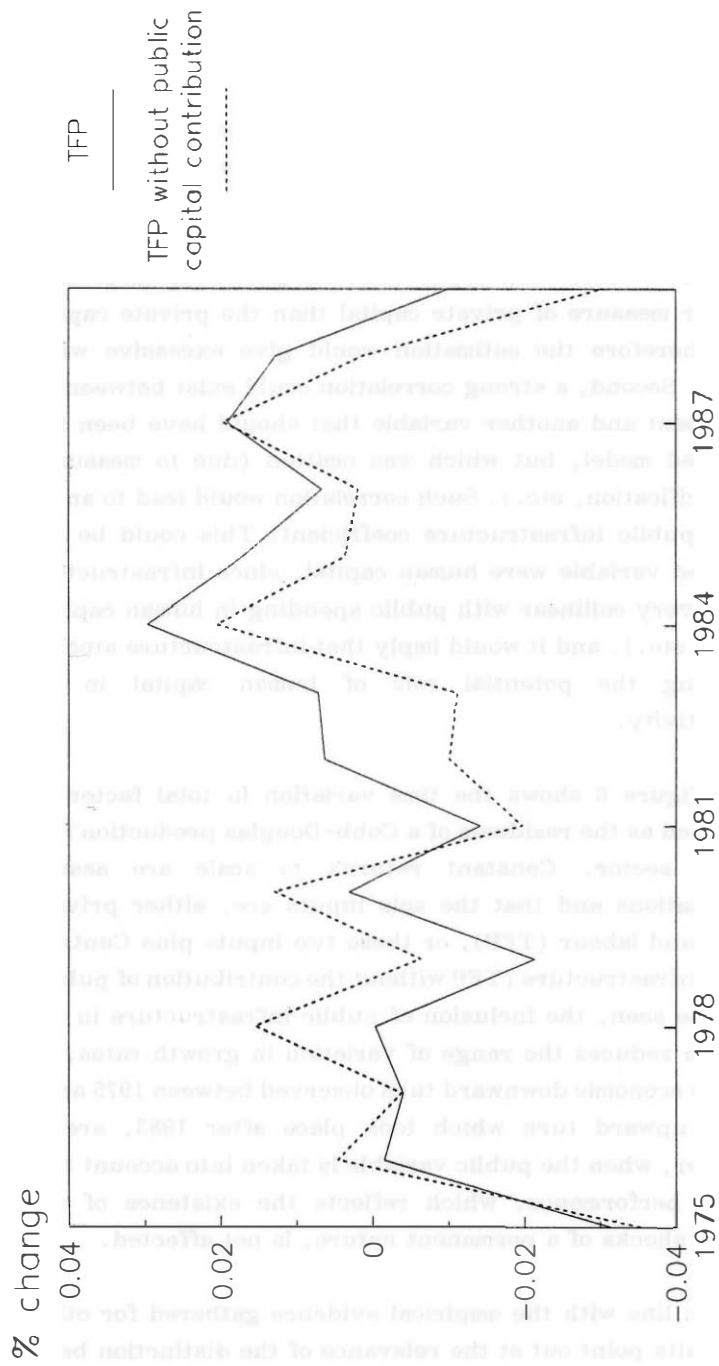
transportation and communications infrastructure apart from the rest of public capital.

Several elements may explain a possible upward bias in the estimation of the output elasticity. First, the true rate of depreciation of private capital may be closer to the rate assumed for public capital (5% in our case) than to the one used to construct private capital stock (10% in our case). Under such hypothesis, the public capital stock used could be a better measure of private capital than the private capital stock itself and, therefore the estimation would give excessive weight to public capital. Second, a strong correlation could exist between infrastructure investment and another variable that should have been included in the estimated model, but which was omitted (due to measuring problems, misspecification, etc.). Such correlation would lead to an overestimation of the public infrastructure coefficient. This could be the case if the excluded variable were human capital, since infrastructure investment can be very collinear with public spending in human capital (education, health, etc.), and it would imply that infrastructure stock would also be reflecting the potential role of human capital in private-sector productivity.

Figure 6 shows the time variation in total factor productivities calculated as the residuals of a Cobb-Douglas production function for the private sector. Constant returns to scale are assumed in both specifications and that the sole inputs are, either private productive capital and labour (TFP), or these two inputs plus Central Government public infrastructure (TFP without the contribution of public capital). As it can be seen, the inclusion of public infrastructure in the production function reduces the range of variation in growth rates, in such a way that the economic downward turn observed between 1975 and 1982, as well as the upward turn which took place after 1983, are less intense. Moreover, when the public variable is taken into account the annual path of TFP performance, which reflects the existence of supply and/or demand shocks of a permanent nature, is not affected.

In line with the empirical evidence gathered for other countries, the results point out at the relevance of the distinction between current

Figure 6
Private Productivity and Public Capital



public expenditure and capital expenditure as far as the study of the impact of budgetary policy is concerned. Moreover, the distinction drawn between public investment as a whole and investment in transportation and communications infrastructure highlights that not only the breakdown of public spending by current expenditure and capital investment is important, but that the breakdown of public investment by infrastructure and other types of investment is also relevant.

The positive effects of public infrastructure investment on private-sector productivity suggest some tentative policy conclusions. First, the process of budget consolidation undertaken in order to meet the requirements for joining the Economic and Monetary Union should not imply a reduction of public spending in infrastructure, as has occurred in recent years. Even more important, given the positive impact on productivity of this type of expenditure, any policy aimed at improving the competitiveness of the economy will have to ensure that budgetary consolidation is compatible with public investment efforts in the area of infrastructure⁽²⁾.

⁽²⁾ Moreover, the empirical evidence gathered in an ongoing study allows us to conclude, at least provisionally, that the crowding-out effect of public investment on private investment is negligible if not irrelevant, in the case of Spain. In other studies with a different conceptual framework, however, there are indications that this crowding-out effect is, in the long run, absolute (see Argimón and Roldán 1991)).

**APPENDIX I:
BRIEF DESCRIPTION OF THE TESTS USED**

- ADF test

Given a time series x_t , the ADF test is a test on the statistical significance of β in the regression:

$$\Delta x_t = \beta x_{t-1} + \sum_{j=1}^n \gamma_j \Delta x_{t-j}$$

where n is such that the regression residuals are white noise.

The null hypothesis that x_t is a first-order integrated series ($x \sim I(1)$) cannot be rejected, if we cannot reject that $\beta=0$. Conversely, if $\beta \neq 0$, then x_t is said to be stationary ($x_t \sim I(0)$). If the null hypothesis cannot be rejected, it must then be tested whether β' is different from zero in the regression:

$$\Delta^2 x_t = \beta' \Delta x_{t-1} + \sum_{j=1}^{n'} \gamma'_j \Delta^2 x_{t-j}$$

If $\beta' \neq 0$, x_t is said to be a first-order integrated series. If the null hypothesis that the series is $I(2)$ ($\beta'=0$ cannot be rejected), then, likewise, it must be tested whether x_t is an integrated series of order three, as opposed to order two.

When the ADF test is used as a cointegration test it is performed on the residuals of the estimated long run relationship. McKinnon (1990) gives critical values of the t statistic that is needed for our tests on β , since, under the null hypothesis, the t statistic does not have the standard distribution. Critical values vary with sample size and with the presence or absence of a constant or a trend. When this test is used as a cointegration test, the critical values also vary with the number of variables included in the regression which gives rise to the residual x_t , as well as with the inclusion or exclusion of a time trend in the cointegration relationship.

The critical values for a sample of 26 data (1964 to 1989) are:

	Size of the test	
	5%	10%
One variable		
without constant	-1.95	-1.62
constant	-2.98	-2.63
constant and trend	-3.59	-3.23
Three cointegration variables		
constant	-4.08	-3.69
constant and trend	-4.60	-4.19
Four cointegration variables		
constant	-4.54	-4.14
constant and trend	-5.02	-4.59

Source: McKinnon (1990).

A summary of the results of the ADF test for the different series used may be found in Table A.1.

- ECM test

Because of its simplicity, the cointegration test based on the ADF statistic raises the problem of imposing common factor restrictions, which have not been tested. Such limitation can reduce the power of this test in relation to others. In this sense, the statistical significance of the estimated coefficient of the error correction term in an Error Correction

TABLE A.1
Tests of unit roots

	I(0)	I(1)	I(2)
$(y - k_p)_t$		X	
$(1 - k_p)_t$		X^a	
$(k_g - k_p)_t$			
GG (NA)			X
CG (NA)		X	
CG (PA)		X	
CG (PA, with KT)		X	
GG (MOISEES)		X^a	

All variables are in logarithms.
(a) At the 10% level of significance.

Mechanism (ECM) representation of the model is a more powerful test than the ADF statistics.

The ECM cointegration test is based on the statistical significance of β in the regression:

$$\Delta y_t = \alpha_0 + \alpha_1 \sum_1^k \Delta x_{1t} + \beta [y_{t-1} - c_1 \sum_1^k x_{1t-1}]$$

where y_t is the variable to be explained, x_i are the explanatory k variables, c_i are the coefficients of the cointegration relationship, and the constant (the mean) of the long-run relationship is recovered as the mean value of the expression in brackets. A more formal development, as well as the critical values for this test, may be found in Banerjee, Dolado and Mestre (1993), where this same relationship is estimated by ordinary least squares (OLS) through the specification of an unrestricted dynamic model.

The ECM representation can also be interpreted as a specification of the dynamic relationship between the different variables. The long-run

relationship is represented by the expression in brackets and the coefficients α_1 give us the short-run relationships.

The regression results for the different definitions of the variables and the different specifications may be found in part in the tables, under the heading of non-linear least squares (NLS). The estimated values for c_1 appear in this part of the different tables. The line ECM shows the value for β and its t-ratio. In all cases, the test required the inclusion of the lag in the variable to be explained, as an additional regressor. Likewise, for the results in Tables 2, 3 and 4, the public infrastructure variable was not included in the short run, since it was not statistically different from zero.

For a sample size of 25, the critical values are:

	Size of the test	
	5%	10%
Three cointegration variables		
constant	-3.64	-3.24
constant and trend	-4.18	-3.72
Four cointegration variables		
constant	-3.91	-3.46
constant and trend	-4.18	-3.72

Source: Banerjee, Dolado and Mestre (1993).

APPENDIX II: CAUSALITY

If two series are cointegrated then causality must exist, at least in one of the two possible directions (Granger (1988)). However, beforehand, an exogeneity test must be performed: it will allow us to determine whether the variables in the model, and whose causality is not being tested, are exogenous.

1. Checking exogeneity:

The exogeneity check is formulated as a Hausman specification test. Given that the model includes an error correction mechanism (ECM), a test must be performed to determine whether this mechanism helps to explain the behaviour of the variables (Charemza and Deadman (1992)).

In this case, the test entails including the previously estimated error correction mechanism among the instruments which are used to construct the Vector Autoregressive for each explanatory variable z_{pt} . The estimation is based on the regression:

$$\Delta z_{pt} = \sum_{i=1}^j \sum_{p=1}^q \alpha_{pi} \Delta z_{pt-i} + \sum_{i=1}^j \beta_i \Delta x_{t-i} + \sum_{i=1}^j \lambda_i \Delta y_{t-i} + \sum_{i=1}^j \gamma_i \text{ecm}_{t-i} \quad (1)$$

where x_t and y_t are the variables whose causality relationship is being tested and ecm_t is the error correction mechanism. If the null hypothesis for the ecm coefficient in the equation of each of the explanatory variables cannot be rejected, then the residuals of equation (1) for these variables are calculated, with the ECM omitted. The statistical significance of these residuals in the model equation is tested. It is formulated as an F test with $k_2 - k_1$, $T - k_2$ degrees of freedom, where T is the number of observations and k_1 and k_2 are the number of parameters estimated in the restricted model and in the unrestricted version (which includes the residuals), respectively. In the event that the null hypothesis that the variables are exogenous cannot be rejected, they would appear contemporaneously in the ECM equation and, consequently, in the causality test.

2. Causality test

Once exogeneity has been determined, a causality direction test based on Granger (1969) is built. The causality equation, in the event that we wish to check whether x causes y , is specified as:

$$\Delta y_t = a_0 + \sum_{i=1}^j a_i \Delta x_{t-i} + \sum_{i=1}^j \sum_{p=1}^q b_{pi} \Delta z_{p,t-i+1} + \sum_{i=1}^j c_i \Delta y_{t-i} + \sum_{i=1}^j d_i ecm_{t-i} \quad (2)$$

where $b_{pi} \neq 0$, if and only if z_p is exogenous, and ecm is the previously estimated error correction mechanism.

It is a test on the overall significance of the coefficients a_i and d_i and is formulated as an F with $k_2 - k_1$, $T - k_2$ degrees of freedom, where T is, again, the number of observations, and k_1 and k_2 are the number of parameters estimated under the restricted ($a_i = d_i = 0$) and unrestricted models, respectively. Likewise, the t statistics can be used as a criterion.

3. Results

Table A.2 shows the results of the tests performed in order to determine the exogeneity of $(1-k)$, carried out for each of the infrastructure variables. Only one lag of each of the variables was included. The error correction mechanism used as a regression was obtained from the non-linear least squares estimation of the entire dynamics, as reflected in the results presented in Table 3, i.e. the variable ecm , which is included in the exogeneity and causality tests, is obtained by estimating an equation of the type:

$$\begin{aligned} \Delta(y-k_p)_t &= a_0 + a_1 \Delta(1-k_p)_t + a_2 \Delta(y-k_p)_{t-1} + \\ &+ a_3 \left[(y-k_p)_{t-1} - a_4(1-k_p)_{t-1} - a_5(k_g-k_p)_{t-1} \right] \end{aligned}$$

and setting

$$ecm_t = (y-k_p)_t - \hat{a}_4(1-k_p)_t - \hat{a}_5(k_g-k_p)_t$$

As Table A.2 shows, it seems that in no case can the null hypothesis be rejected. Therefore, in the causality test, the explanatory variable I-k should appear contemporaneously.

The causality test resulted in the figures presented in Table A.3, which shows the values of the relevant coefficients and their respective t-ratios. In particular, under the columns labelled (1), the causality direction runs from infrastructure stock to productivity, and, under the columns labelled (2), the direction runs from productivity to infrastructure. The conclusions that can be reached from these results are that infrastructure affects productivity, whereas there is no evidence that variations in productivity generate changes in infrastructure.

TABLE A.2
Exogeneity tests

Infrastructure series	$F = \frac{S_1 - S_2 / k_2 - k_1}{S_2 / T - k_2}$
General Government (GG (NA))	0.71 (4.38)
Central Government (NA)	2.63 (4.38)
Central Government (PA)	4.22 (4.38)
Central Government (PA) + transfers	3.29 (4.38)

In brackets is the value of $F_{1,19}$ at the 5% level of significance.

TABLE A.3
Test of causality direction

	GG (NA)		CG (NA)		CG (PA)		CG (PA) + KT	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
α_3	0.23 (1.44)	-0.002 (0.01)	-0.04 (0.25)	-0.31 (1.30)	0.12 (0.83)	-0.40 (1.44)	-0.06 (0.32)	-0.26 (1.17)
α_5	-0.36 (2.29)	-0.22 (0.83)	-0.81 (4.13)	0.79 (2.04)	-0.68 (3.65)	0.83 (2.10)	-0.74 (3.87)	0.46 (1.29)
F	8.26**	1.63	25.62**	2.49	25.88**	2.40	24.45**	1.73

Results of the estimation of the equation

$$\Delta y_t = \alpha_0 + \alpha_1 \Delta(1-k)_t + \alpha_2 \Delta(1-k)_{t-1} + \alpha_3 x_{t-1} + \alpha_4 \Delta y_{t-1} + \alpha_5 \text{ecm}_{t-1}$$

where $y_t = (y - k_p)_t$, $x_t = (k_g - k_p)_t$ in column (1), and the opposite in column (2).

The values of the coefficients α_3 and α_5 , and their respective t-ratios are shown, as well as the F test of the overall significance of α_3 and α_5 . The value of this test at the 5% level of significance is 3.55.

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