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Public Capital and Output Growth in Portugal: An Empirical Analysis

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Abstract

The paper investigates the growth effects of public capital in Portugal using annual data for the period 1965–95. Both a production function and a vector autoregressive model are estimated. Public capital is shown to be a significant long-term determinant of output growth. The size of the estimated production elasticity indicates, in line with studies for other countries, a substantial growth payoff from public investment. Disaggregating public capital shows that investment related to, among other things, roads, railways, and airports is more productive than public investment in other major categories.

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Contents

I. Introduction

Since Portugal joined the European Union (EU) in 1986, it has received on average 3.3 percent of GDP in transfers per annum from the EU. These transfers—primarily designed to promote infrastructure investment, human capital accumulation, and private investment—boosted the expansion of public investment (including capital transfers) from 4.8 percent of GDP in 1986 to 6.3 percent of GDP in 1998. As a result, gross public capital formation in Portugal (as a share of GDP) is currently the second highest in the EU area (see Figure 1). On average, the real value of the public capital stock grew by 5.1 percent during 1986–95, which is considerably above that of the United States (2.1 percent) but below that of Spain (see Table 1). However, the highest average change in the real value of the Portuguese capital stock was recorded during the 1974–85 period, just after the shift in the political regime,[2] indicating that even before joining the EU a substantial share of resources was devoted to public capital accumulation.

Various authors have tried to determine the productivity effects of public capital by estimating a Cobb-Douglas production function that includes public capital as an input.[3] Aschauer (1989, 1990) was one of the first to investigate this issue for the United States in an attempt to explain the productivity slowdown in the 1970s. He found that a 1 percent increase in the public capital stock increased private capital productivity by 0.39 percent, suggesting that public capital is an important determinant of production. Since then, many authors have employed this approach, but some have also pointed to its lack of attention to feedback effects because it assumes that the causality runs from public capital to output. Recently, a number of authors (e.g., Otto and Voss, 1996; Batina, 1998; and Sturm, Jacobs, and Groote, 1999) have employed a vector autoregressive (VAR) model with a view to capturing the dynamic interactions between output, public capital, and private capital. The VAR approach models every endogenous variable as a function of its own lagged value and the lagged values of the other endogenous variables and can therefore assess whether there is any feedback from private sector variables to the public capital stock.[4]

The objective of this paper is to study the effects of public capital on output growth in Portugal using annual data over the period 1965–95. To this end, two approaches are taken. First, a production function incorporating public capital, private capital, and employment is estimated using both the conventional technique of ordinary least squares (to ensure

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[2]In April 1974, a revolution replaced "The New State" with a democratic regime.

[3]See Table 2 for an overview of the studies to date.

[4]Other approaches include the estimation of cost functions, cross-sectional studies using country level data, and calibrated structural models. For example, see Gaspar and Pereira (1995) for an application of a computable general equilibrium model to assess the growth effects of EU-financed capital expenditures in Portugal.

comparability with other studies) and Johansen's (1988) cointegration procedure.[5] The results indicate that public capital is a significant long-term determinant of real output growth. In the second approach, an (unrestricted) VAR model is estimated to assess the causal dynamics between public capital and output. A positive, Granger (1969)-causal relationship is found, which runs from public capital to Portuguese production, providing support for the view that public capital has contributed to Portugal's economic growth.

The remainder of the paper is structured as follows. In Section II results are presented for the single equation approach. Section III addresses short-run dynamics by specifying a vector autoregressive model. Section IV summarizes the main findings.

II. The Production Function Approach

A. Conceptual Framework

Assume a Cobb-Douglas production function that incorporates the public capital stock, G, as an input:[6]

$$Y = AK^{\alpha}G^{\beta}L^{\gamma}, \quad \alpha, \beta, \gamma > 0, \tag{1}$$

where A denotes an index of economy-wide productivity, K is private capital, L denotes employment, and Y is output. In this setup, an increase in public capital raises output directly (i.e., $Y_G=\beta(Y/G)>0$, where a subscript denotes a partial derivative), but also indirectly through its positive effect on the marginal productivity of private capital and labor (i.e., $Y_{KG}>0$ and $Y_{LG}>0$).

By taking natural logs on both sides of the equation, and denoting lowercase variables as the natural log of the respective uppercase variable, the following equation results:

$$y = a + \alpha k + \beta g + \gamma l, \quad \alpha>0, \beta>0, \gamma>0. \tag{2}$$

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[5]A number of authors have employed Johansen's procedure to estimate production functions (e.g., Batina, 1998; Flores de Frutos, Gracia-Diez, and Pérez-Amaral, 1998; Ghali, 1998; and Mamatzakis, 1999), but most earlier studies employ ordinary least squares. See below for a further discussion of cointegration issues.

[6]Arrow and Kurz (1970) were one of the first to study theoretically the implications of incorporating public capital in a neoclassical growth model.

The coefficients α, β, and γ are the output elasticities of the factor inputs. Inclusion of public capital in the production function raises the issue of returns to scale. Imposing the restriction of constant returns to scale across all inputs (i.e., $\alpha+\beta+\gamma=1$), which is a common assumption in the literature, yields an expression featuring decreasing returns with respect to private inputs:

$$y-k = a + \beta(g-k) + \gamma(l-k), \quad \alpha+\gamma<1. \tag{3}$$

An alternative model assumes constant returns to scale in both private inputs, allowing for increasing returns to scale across all inputs:

$$y-k = a + \beta g + \gamma(l-k), \quad \alpha+\gamma=1. \tag{4}$$

This specification has been employed at times in the endogenous growth literature (see, for example, Barro, 1990).

In the econometric specification, all equations to be estimated include a capital utilization rate, *cu*, to capture the effects of the business cycle on factor use. Because the capital utilization rate enters the equation in an additive fashion, it does not affect the optimal capital-labor ratio.[7] Many studies also include a constant and a time trend to capture Hicks-neutral technological progress. The current study will represent the technology variable, *a*, by a constant. This specification reflects the underlying hypothesis that economic growth in Portugal has been mainly driven by factor accumulation and not by increases in factor productivity. In addition, a dummy variable for the period 1975–85 will be introduced to capture the period in between the new regime after the 1974 revolution and EU accession in 1986.

B. Evidence for OECD Countries

Most time series studies employ a Cobb-Douglas production function to estimate the output effects of public capital. On average, these studies estimate a production elasticity of public capital (β) of 0.25 for various OECD countries when the production function is estimated in levels. Estimates of β vary considerably across countries but lie in the interval 0.20–0.30 at a 95 percent level of confidence (Table 3). If the model is estimated in first differences, estimates of β are on average higher and confidence intervals are wider. Panel studies—based

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[7]Some studies (e.g., Tatom, 1991) also include the relative price of energy in the equation to capture supply shocks but it is not immediately clear why any price variable should be included in a production function.

on regional data for a single country—find in general much lower estimated coefficients, which could be ascribed to "leakages" reflecting the fact that, at the regional level, not all beneficial spillover effects of public investment can be internalized (see Munnell, 1992).

C. The Data

The empirical analysis employs annual data for Portugal over the period 1965–95. Data on GDP, the number of employed persons, the private capital stock, and the public capital stock are obtained from the Historical Series for the Portuguese economy.[8] All series are expressed in constant prices. Estimates of the private and public capital stock are constructed by employing the perpetual inventory method (OECD, 1993). This approach computes the value of the capital stock by summing over past investments, appropriately adjusted for the rate of depreciation. It is assumed that asset lifespans are the following: residential buildings, 70 years; investments in machinery and equipment, 16 years; and public works (roads, railways, etc.), 35 years. The perpetual inventory method has been widely applied in the literature but is not free of criticism. Some authors (e.g., Sturm and de Haan, 1995) have shown that assumptions concerning the lifespans of capital goods matter for the size of the production elasticity estimate for public capital.

Table 4 shows the composition of the estimated Portuguese capital stock. The ratio of public to private capital was 18 percent for Portugal in 1995 (27 percent if only structures and equipment are included), compared with, for example, a ratio of 31 percent for the United States (58 percent if only equipment and structures are counted) with a PPP-based per capita income more than twice that of Portugal. Nearly 60 percent of public capital consists of core infrastructure such as roads, railways, airports, and the like. Investment in equipment and transport material amounts to only 9 percent of the public capital stock.

D. Empirical Results for Portugal

Ordinary least squares estimates

For purposes of comparison with the literature, equations (2)–(3) are first estimated in levels using ordinary least squares.[9] Because no time series data on capital utilization over a sufficiently long time span are available, the estimated output gap is used as a proxy. Potential output is obtained by employing a Hodrick-Prescott (1997) filter to the actual real output series. An alternative measure, the unemployment rate, u, is used as an indicator of demand

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[8]The time series were kindly provided by the Bank of Portugal.

[9]Issues of cointegration and the econometric validity of the ordinary least squares results will be discussed below.

pressure in factor markets.[10] The estimation results for equation (2) are presented in Table 5.[11] Fitting an unrestricted production function yields a production elasticity of public capital significantly different from zero (using standard inference criteria), amounting to 0.19, just at the lower bound of the confidence interval derived from the results of previous studies for other countries (Table 3). This implies that a 1 percent increase in the public capital stock raises GDP by 0.19 percent. The private capital elasticity amounts to 0.37, whereas the labor elasticity is on the order of 0.67. These values are closely in line with traditional assumptions on capital and labor shares in industrialized countries but higher than the Portuguese labor share of value added. Summing over the three input coefficients yields a value of 1.2, a little above unity. To test whether production can be characterized by constant returns to scale, a Wald test on the coefficients was conducted, which indicated that the restriction of constant returns to scale cannot be rejected. Accordingly, the focus below will be on the specification that assumes constant returns to scale.

Imposing the restriction of constant returns to scale—in line with Aschauer's (1989) specification, which serves as a useful benchmark (see equation (3))—yields slightly higher values for the output elasticity of public capital, β, ranging from 0.22–0.27 (Table 6). The estimated private capital elasticity amounts to 0.33–0.48, which seems to be a plausible range of values both in terms of capital income shares and in terms of the results obtained for other countries. Given that the public capital stock amounted to 51 percent of GDP in 1995, an estimate of the production elasticity of public capital in the range of 0.22–0.27 implies a marginal productivity of public capital of 43 to 52 percent a year. This is roughly four to five times the implicit nominal interest rate on public debt in that year. Gramlich (1994) finds even larger rates of returns (on the order of 100 percent a year or more) for the United States. These returns on public capital are very large and should be interpreted with caution.[12]

Estimating an equation that includes the unemployment rate (rather than capital utilization) as an indicator of demand pressure in factor markets yields a slightly larger β coefficient (specification II in Table 6). Specification III includes an interaction dummy variable for the period 1975–85 to capture possible productivity effects on public capital during the period when the role of the state was greatly expanded. The negative interaction coefficient is large

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[10]The unemployment rate is obtained from das Neves (1994) and the IMF's World Economic Outlook database.

[11]The results are obtained using PcGive Version 9 (Doornik and Hendry, 1997) and Eviews Version 3.1.

[12]Munnell (1992) concludes that the estimated returns on public capital for the United States are too large to be credible but stresses that these results should not be discarded altogether, since evidence from cost-benefit studies of individual projects and cross-sectional studies indicates that investment in public infrastructure may have a large payoff.

and suggests that public investment actually contributed to reducing output in these years.[13] In addition, the negative value of the "intercept" dummy (as observed in all the specifications) indicates that during this time period of accelerated public investment, total factor productivity growth was also lower.[14] Because of the presence of autocorrelation in the error terms, the equations are also estimated using the Cochrane-Orcutt (C-O) procedure, which changes the coefficient on public capital marginally.

Table 7 presents results for various components of public capital: core construction projects, *gc*, public buildings, *gb*, and transport equipment and machinery, *ge*. The equation including all three components shows that none of the coefficients on public capital are significant and two of them (core infrastructure and buildings) have the wrong sign. In addition, the coefficient on employment becomes implausibly large. That the coefficient on each component is insignificant while, as shown above and in Table 7, the coefficient on aggregate public capital is significant, suggests that there maybe important interrelationships between the different components of public capital. Statistically, the insignificant coefficients on public capital also reflect multicollinearity between the various components. Nevertheless, to obtain some idea about differences in productivity of various types of public capital, the equations are also estimated for each component separately. The coefficient on core infrastructure is now significant and close to the estimate found for the aggregate capital stock. The growth effect of transport material and equipment is also statistically significant, but substantially smaller than core infrastructure (i.e., β=0.10 compared with β=0.18 in the benchmark case). Government buildings and equipment do not appear to play a significant role in explaining capital productivity, which is confirmed by the statistically insignificant coefficient. These results should be interpreted with care, given the poor results that were obtained when all three components of public capital were included together.

Integration and cointegration issues

So far, the issue of stationarity of the variables has not been addressed. If variables are nonstationary, the usual test statistics have nonstandard distributions, implying that the use of standard inference tests may give rise to seriously misleading inferences. A related problem is the possibility of finding spurious relationships between variables. Only when variables are cointegrated—expressing the presence of a long-run equilibrium relationship between a group of nonstationary economic time series—can the equations be estimated in levels. The first step is to determine the order of integration of the variables. To this end, tests for unit roots are performed on the levels, first differences, and second differences of the variables. The results of the augmented Dickey-Fuller (1981) tests are presented in Table 8. The evidence suggests that the variables g, k, l, and y are all integrated of order one (i.e., they are $I(1)$ variables) and

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[13]The sum of the two coefficients for public capital is negative.

[14]The dummy may also capture somewhat the negative output effect of higher oil prices after the oil shocks of 1973 and 1979.

thus nonstationary in levels, but stationary in first differences. The public capital stock appears to be $I(2)$ but its coefficient is very far from unity (i.e., -0.35), indicating that it could be an $I(1)$ variable. Most of the variables expressed in terms of the private capital stock are stationary in levels, except the labor-capital ratio, which seems to be integrated of order two.

The second step is to examine whether the series of the variables used in the estimation of the production function are cointegrated. To this end the Engle-Granger (1987) procedure can be employed, which consists of a unit root test on the residuals of the estimated equation.[15] The Engle-Granger statistics (see the bottom row in Tables 5 and 6) indicate that the null hypothesis of no cointegration cannot be rejected at the 95 percent level of confidence.[16] From the literature it is well known that the Engle-Granger procedure has low power in finite samples[17] and may therefore be unable to detect cointegration when it is present in the data (see Kremers, Ericsson, and Dolado, 1992).

As an alternative to the single equation Engle-Granger test, the Johansen (1988) procedure—based on estimating a VAR—can be employed to test for cointegration. The following equation is estimated:

$$\Delta x_t = \alpha^* \beta' x_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} + \epsilon_t, \tag{5}$$

where x_t is a vector of variables, α^* is an adjustment coefficient, β' is a long-run elasticity (or set of eigenvectors), and ϵ_t is an error term. The maximum likelihood test statistics (i.e., the maximum eigenvalue statistic, λ_{max}, and the trace statistic, λ_{trace}, both adjusted for the degrees of freedom) for the system of equations (5) are reported in Table 10. In contrast to the Engle-Granger results, both test statistics strongly reject the null hypothesis of no cointegration in favor of at least one cointegrating vector.[18] There is little evidence of more than one cointegrating relationship.

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[15]Owing to its simplicity, the Engle-Granger method has been widely applied in the literature. Sturm and de Haan (1995) employ it to argue that Aschauer's model should be estimated in first differences because their test results could not identify cointegration.

[16]The standard critical values cannot be used for present purposes because they were applicable to the actual values of the variable being tested, whereas here, only estimated values of the relevant process are available. MacKinnon (1991) has derived relevant critical values for finite samples from Monte Carlo simulations, which are used in the present case.

[17]In addition, the Engle-Granger procedure imposes an invalid common factor restriction on the dynamics by performing the test on a single equation.

[18]The capital utilization rate, 1975–85 dummy, and interaction term are not included.

In sum, the evidence on cointegration is mixed. The Engle-Granger results point to a lack of cointegration, which would invalidate the earlier ordinary least squares results. However, the Johansen test (generally recognized to be superior to Engle-Granger tests) indicates cointegration. In light of these results, the ordinary least squares estimates should be interpreted with caution, and carefully compared with those obtained by other techniques.[19]

Estimates based on Johansen's approach

One advantage of the Johansen cointegration approach is that it is based on a dynamic multiple equation system (i.e., a VAR).[20] As such, it captures the feedback that might be present between the variables at hand. For example, if public capital crowds out private capital this would be captured, unlike in the ordinary least squares estimates. The text table (see below) reports the standardized cointegrating vector (i.e., the β' eigenvector), showing that all coefficients have their anticipated signs; it features a public capital elasticity of 0.39, whereas the private capital elasticity amounts to 0.44 if no restrictions are imposed. However, the employment elasticity is very imprecisely estimated and only 0.10, which seems to be implausibly small. The sum of the coefficients is close to unity (i.e., 0.93) and the homogeneity restriction cannot be rejected (i.e., $\chi^2(1)=0.01$ [$p=0.98$]), confirming the results of the previous ordinary least squares estimation. The coefficient on public capital in the equation restricted to constant returns to scale (see the second row of the text table) generally does not differ much from the one in the unrestricted equation. The coefficient on employment is higher and significant now, but still low in view of the labor share. When compared with the ordinary least squares results in Table 5 (see the second row, with the results presented without the dummy variable, intercept, and capital utilization rate), the coefficient on public capital does differ substantially. It is interesting to note that the coefficient on public capital in the restricted maximum likelihood equation is only slightly smaller than the coefficient of 0.39 that was obtained by Aschauer (1989, 1990) for the United States.

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[19]In the absence of cointegration, but with nonstationary variables, the literature recommends taking first differences of the variables to obtain stationary time series. Table 9 includes the estimation results in first differences and shows that the coefficient on public capital appears to be similar to that from the ordinary least squares equation, as long as no interaction dummy for the 1975–85 period is included. A notable drawback of first differencing is that it discards information on the long-term relationship between the variables.

[20]The system of equations is estimated by the maximum likelihood technique.

Standardized Eigenvectors Under Johansen's Approach 1/ 2/

	y	g	k	l
β' unrestricted	1.000 (n.a.)	-0.387 (-7.43)	-0.443 (-15.81)	-0.104 (0.39)
β' restricted 3/	1.000 (n.a.)	-0.370 (-12.74)	-0.441 (-44.10)	-0.188 (-12.33)

1/ t-statistics are in parentheses.
2/ Note that the coefficients of g, k, and l need to be multiplied by minus one to derive the respective elasticities.
3/ A constant returns to scale restriction is imposed.

III. The Unrestricted Vector Autoregression Approach

This section employs an unrestricted vector autoregression model to analyze the dynamic interaction between public capital, private capital, employment, and output. Granger causality analysis, impulse response functions, and variance decompositions are employed to quantify the dynamic relationships.

A. Method

The VAR approach sidesteps the need to specify a structural model by modeling every endogenous variable as a function of its own lagged values and the lagged values of the other variables in the system. In the literature, VARs have been criticized for being atheoretical because no a priori theoretical relationship between the variables is assumed. However, the VAR can be used to provide empirical evidence on the dynamic responses of macroeconomic variables to impulses in the public capital stock in order to discriminate between alternative theoretical models of public capital.

In its most general form, a VAR with p lags can at time t be written as follows:

$$\vec{z}_t = \Psi_1 \vec{z}_{t-1} + \Psi_2 \vec{z}_{t-2} + \ldots + \Psi_p \vec{z}_{t-p} + \Omega \vec{x}_t + \epsilon_t \tag{6}$$

where $\vec{z}_t$ is a k vector of endogenous variables, $\vec{x}_t$ is a j vector of exogenous variables, $\Psi_i, \ldots, \Psi_p$ and Ω are matrices of coefficients, and ϵ_t is an error term. In the present case, the VAR is estimated in levels and consists of four endogenous variables, $\vec{z}_t = \left[g_t \;\; k_t \;\; l_t \;\; y_t\right]'$ and a constant term.[21] The system features two lags that were chosen using a likelihood ratio test (see Table 11). Although the four variables are nonstationary and the Johansen (1988)-based results suggested one cointegrating relationship, the VAR is estimated in an unrestricted (levels) format (see Sims, Stock, and Watson, 1990, for a discussion of this issue) by ordinary least squares. As previously noted, estimating the VAR in levels is allowed, as long as the relevant variables are cointegrated.[22]

B. Granger-Causality Tests

Granger-causality tests can be used to study the short- and medium-run linkages between public capital and other macroeconomic variables. It can address a number of issues that have been raised in the literature. Some authors (e.g., Aschauer, 1990; and Munnell, 1992) have argued that the direction of causation may run from high levels of output to larger public investment rather than the other way around. Two hypotheses have been put forward to explain this reverse causation. First, public expenditure may be a luxury good that rises more than proportionally with national income. Second, public capital may move pro-cyclically; during recessions less tax revenue is collected, implying that governments may need to cut public investment to meet their fiscal targets. Other authors have focused on the linkage between public and private investment. Aschauer (1985) found that public investment crowds out private investment in the United States. On the other hand, a large number of theoretical studies assume that public investment initiates private investment through its positive effect on the marginal productivity of private capital.

The results of bivariate Granger-causality tests indicate that the public capital stock Granger-causes output, but output does not Granger-cause public capital (Table 12).[23] This is consistent with the hypothesis that variations in public capital play a part in economic fluctuations. There is bidirectional causality between output and private capital. However, a larger F-value (and thus significance) is attached to output positively Granger-causing private capital. Note that public capital does not Granger-cause private capital, indicating that direct "crowding out" is not present, but this does not preclude private capital being indirectly reduced through other variables (see below). The Granger causality between employment and

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[21]The VAR includes a constant, but the 1975–85 dummy and the capital utilization rate are not included.

[22]This strategy involves no costs in terms of the consistency of the estimators but some costs are incurred in terms of reduced efficiency of estimation.

[23]The results are derived by running a bivariate regression of the relevant pair of variables in the group of variables. Each equation contains lagged values of the left-hand-side variable plus lagged values of the other variable under consideration. In essence, this is equivalent to running a two-variable VAR. The Granger analysis tests whether the lags of the latter are significantly different from zero.

output is found to run from output to employment, implying that employment responds with a lag to fluctuations in output.

C. Impulse Response Analysis

To study the dynamic properties of the VAR, impulse-response functions are employed. These functions trace out the effect of a one standard deviation shock to the orthogonalized residuals of equation i (where i=y, k, l, and g) on current and future values of the endogenous variables in the system. Because of the dynamic structure—in which each equation consists of its own lagged values and the lagged values of all the other endogenous variables—an innovation in one variable is transmitted to all other variables. The ordering of the variables is g, k, l, and y, reflecting the underlying presumption that output is the most endogenous variable in the system. Public capital is assumed to be least sensitive to contemporaneous innovations in the other variables, reflecting the fact that public capital is predominantly the outcome of exogenous government decisions. Various other orderings of the variables were employed and yielded qualitatively similar results. Figures 2a and 2b present the impulse-response functions for the four variables, where the first column of figures displays the results of innovations in the public capital stock. The solid lines in the figures trace the response of a variable over a 10-year time period and the dashed lines represent the 95 percent confidence intervals.[24] All variables are in logarithms, implying that the vertical axis represents percentage changes in a variable (i.e., a 0.01 movement corresponds to a 1 percent change). Note that the confidence intervals are relatively large, indicating that a considerable amount of uncertainty is present so that the results should be interpreted with care.[25]

GDP responds the strongest to innovations in the private sector variables, that is, the private capital stock, employment, and output itself. As can be seen from the lower left-hand panel, the public capital stock has a positive effect on GDP growth, and it adds more to growth in the medium run than during the first year. This could be interpreted as evidence that it takes some time for public capital to become fully productive. However, in light of the large error band, especially during the years immediately following a shock to public investment, this should be interpreted cautiously. Initially, public capital does not respond much to innovations in output, private capital and employment. Over time, private capital and output do contribute to public capital accumulation, as may be expected, given that private and public capital are usually complementary inputs. The medium-term effects of innovations in output on public capital are small, but not significantly smaller than the effect of public capital on output. This might give some support to the reverse causation hypothesis: during an expansionary phase of the business cycle, with larger tax revenues and a less tight fiscal situation, the government becomes more willing to finance public investment projects.

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[24]These were computed using the option of analytic asymptotic standard errors in Eviews 3.1.

[25]Eichenbaum (1992) argues that this uncertainty is typical of unrestricted VAR models.

On impact, innovations in public capital depress private capital formation, and it takes approximately 10 years before the negative effect on private capital has died out. In the long run, the effect on private capital is zero. This suggests that in the short run some crowding out is present, but this was not confirmed in the Granger-causality analysis, implying that crowding out occurs through other variables. Employment and private capital are complementary factors of production in the short run, which is consistent with, for example, a Cobb-Douglas production structure. Both private factors of production contribute positively to GDP growth in the short and medium run.

D. Variance Decompositions

Variance decomposition is another method used to analyze the dynamics of the system of variables. It provides information on the quantitative importance of random shocks to the variables in the system. In Table 13 the rows give the variance of the *k*-steps ahead forecast error explained by contemporaneous shocks in one of the three other variables, whereby the four rows for each variable add up to 100 percent.[26] It is evident that in the short and medium run, a significant share of the variation in output is due to innovations in private capital (43 percent in year 5) and employment (34 percent in year 5). Innovations in public capital contribute only about 10 percent to the variance in output in the medium term. If the low productivity of public capital during the period 1975–85 is filtered out by including the intercept dummy, public capital contributes 22 percent to the variance in output in year 5. Public capital appears to be largely exogenous in the short and medium run; its forecast error is mainly due to its own innovations, indicating that it does not respond much to private economic activity. This is in line with the assumption that contemporaneous shocks to public investment and thus the capital stock stem mainly from government decisions that are independent of other variables considered here.

IV. Concluding Remarks

This paper has analyzed the short- and long-run output effects of public capital using data for the Portuguese economy over the period 1965–95. Public capital is shown to be a significant long-run determinant of output growth. This supports the earlier work of Gaspar and Pereira (1995), showing that EU-supported public investment has a positive effect on economic activity. The size of the estimated production elasticity suggests that a 1 percent increase in the public capital stock increases output by some 0.20–0.35 percent. If a conservative view is taken and the lower bound is adhered to, this would imply a marginal productivity of public capital over 40 percent in 1995—four times the implicit nominal rate of interest on public debt in that year. These high numbers are roughly in line with results found by studies for the United States and various other countries. Disaggregating public capital shows that

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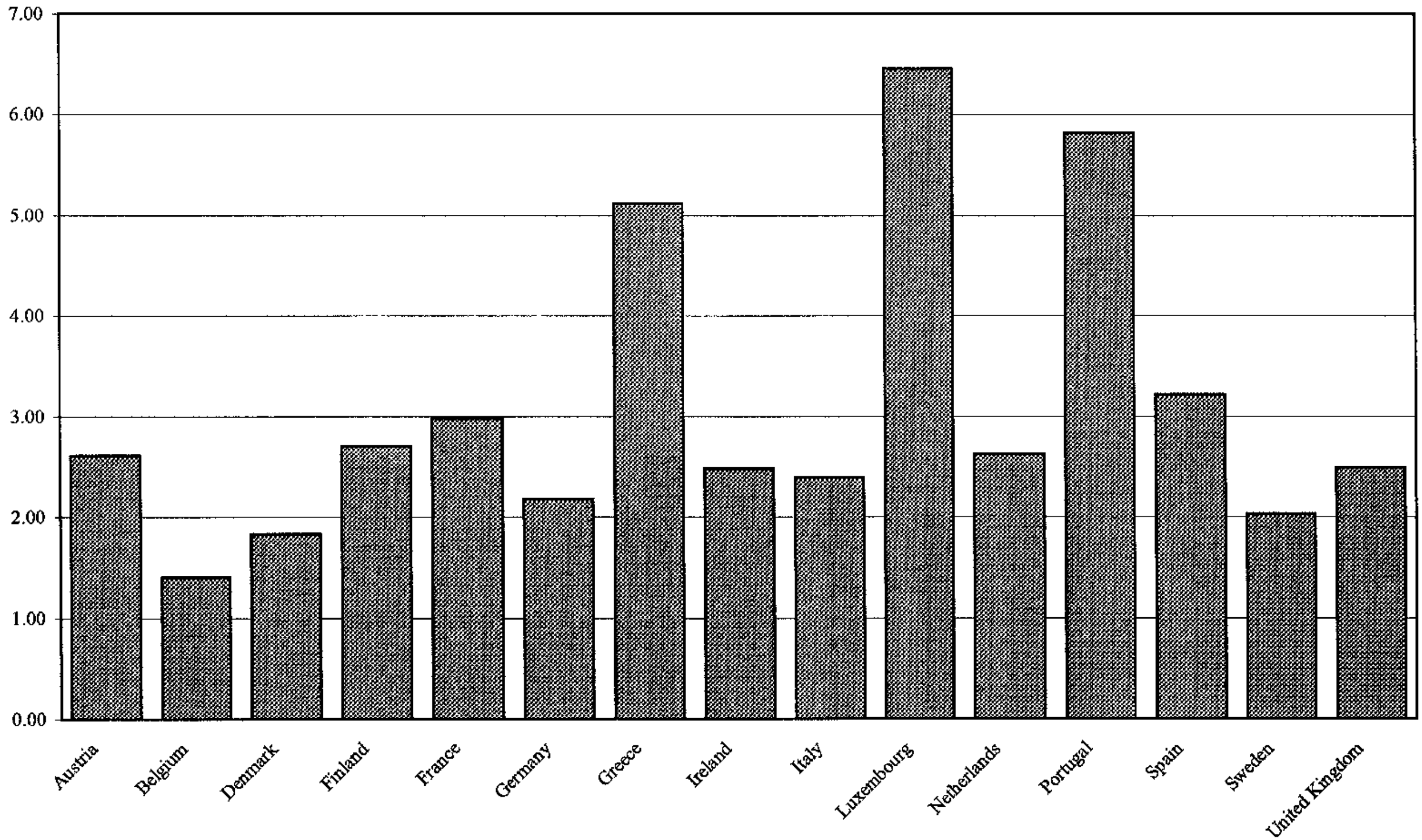
[26]The same variable order as in the case of the impulse-response analysis is employed.

investment related to, among other things, roads, railways, and airports is more productive than public investment in other major categories.

Public capital is found to Granger-cause output, supporting the hypothesis that public investment contributes positively to output fluctuations. Variance decompositions suggest, however, that public capital does not explain a quantitatively important amount of the variation in output, although the result reflects in part the low productivity of public capital during the 1975–85 period. On the issue of crowding out of private by public investment, the results were mixed. The impulse-response analysis shows that public capital may crowd out private capital. But Granger-causality tests could not validate this, implying that crowding out occurs through other variables. In the short and medium run, public capital does not respond much to changes in private sector variables. Hard evidence in support of the reverse causation argument—which alleges that public investment responds positively to upswings in the business cycle—could not be found.

In light of limitations of the econometric methods employed in this study (and studies for other countries), it is important to be cautious in deriving policy conclusions from the empirical findings presented here. Keeping this caveat in mind, the results consistently indicate, in line with other studies, a substantial growth payoff from public investment. Additional research to gauge the precise size of the positive effect of public capital on Portuguese growth is warranted.

Figure 1. Average Gross Public Capital Formation as share of GDP for Various Countries, 1994-98 1/



Source: *World Economic Outlook* database.

1/ Includes capital transfers for Portugal.

Figure 2a. Portugal: Impulse Response Functions

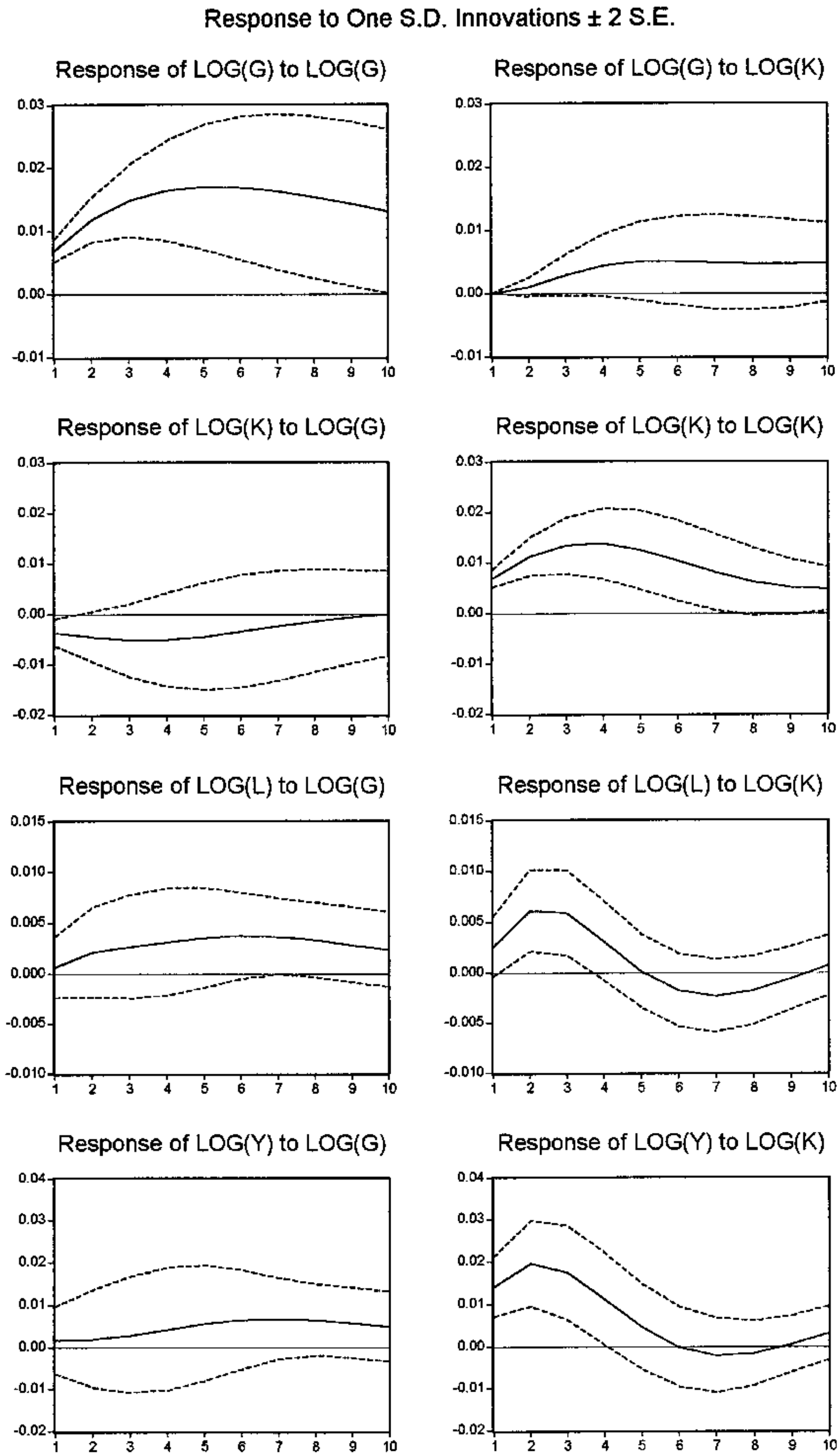


Figure 2b. Portugal: Impulse Response Functions

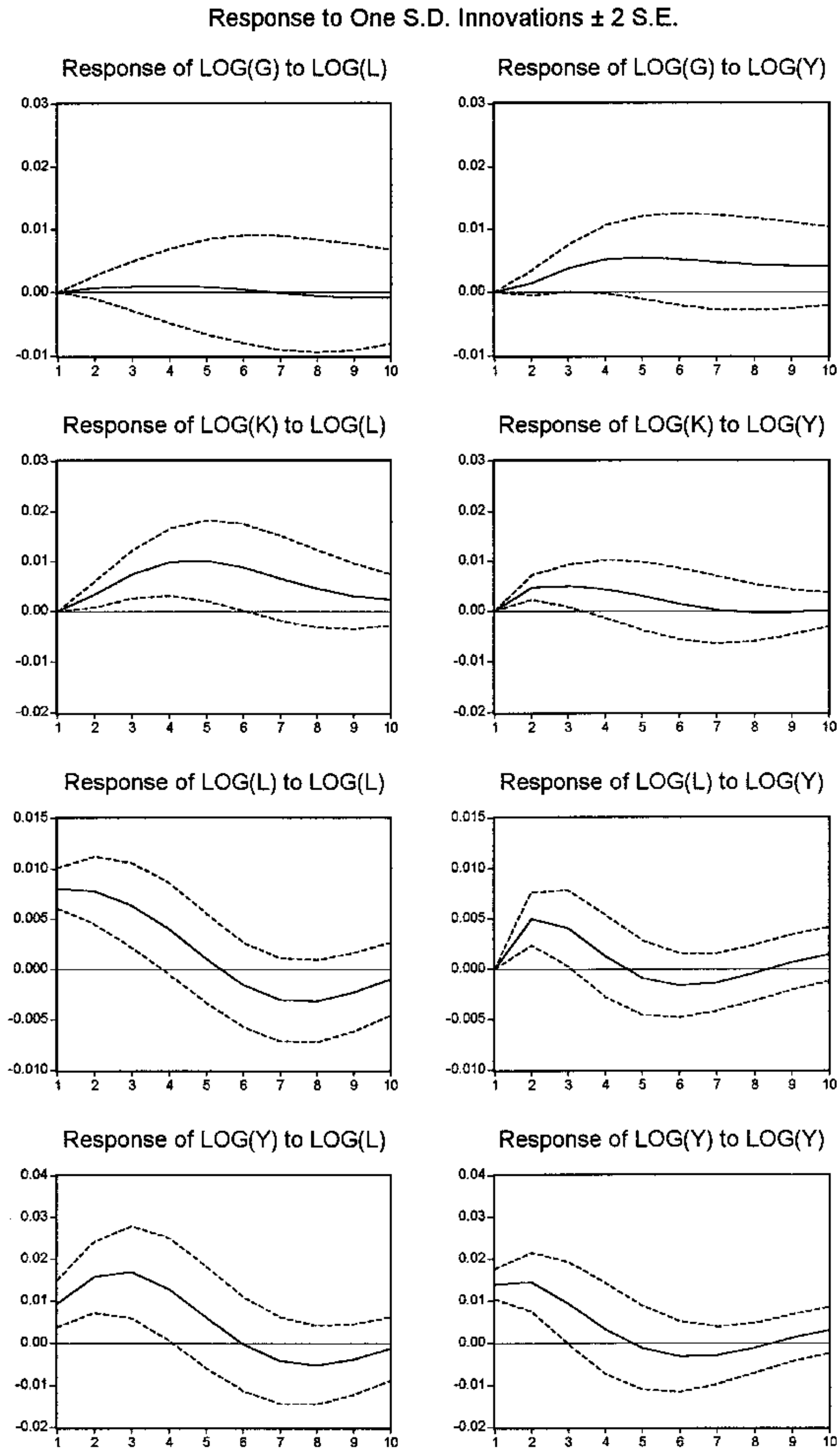


Table 1. Growth Rates of Private and Public Capital Stocks in Portugal, Spain, and the United States for Selected Time Periods

(Average annual growth rate of the capital stock)

	1966–95	1966–73	1974–85	1986–95 1/
Private sector				
Portugal	7.0	12.0	5.3	4.4
Spain	5.8	10.2	3.2	4.6
United States	3.0	4.0	3.0	2.2
Public sector				
Portugal	5.5	5.3	5.6	5.1
Spain	5.3	8.1	4.1	7.2
United States	2.0	2.8	1.8	2.1

Source: Data for Portugal are from accumulated investment flow data provided by the Bank of Portugal. Data for Spain are taken from Flores de Frutos and others (1998). United States data are from the U.S. Department of Commerce, Bureau of Economic Analysis (1998).

1/ Data for Spain are only up to 1992.

Table 2. Overview of Empirical Studies: The Production Function Approach

Author	Country	β	Specification	Data
Ratner (1983)	U.S.	0.06	CD, LL	TS, 1949–73
Aschauer (1989)	U.S.	0.39	CD, LL	TS, 1949–85
Ram and Ramsey (1989)	U.S.	0.24	CD, LL	TS, 1948–85
Munnell (1990)	U.S.	0.31	CD, LL	TS, 1949–87
		0.37 1/	CD, LL	
Aaron (1990)	U.S.	0.41	CD, LL	TS, 1952–85
		0.27	CD, DL	
Tatom (1991)	U.S.	0.13	CD, LL	TS, 1948–89
		0.04 2/	CD, DL	
Ford and Poret (1991) 3/	U.S.	0.30	CD, LL	TS, 1949–87
		0.25	CD, DL	
	Germany	0.53	CD, DL	TS, 1961–87
	Canada	0.63	CD, DL	TS, 1963–88
	Belgium	0.52	CD, DL	TS, 1967–88
	Finland	0.54	CD, DL	TS, 1967–88
	Australia	0.34	CD, DL	TS, 1967–87
Hulten and Schwab (1991)	U.S.	0.21	CD, LL	TS, 1949–85
		0.03 2/	CD, DL	
Berndt and Hansen (1991)	Sweden	n.a. 4/	CD, LL	TS, 1960–88
Finn (1993)	U.S.	0.16	CD, LL	TS, 1950–89
Bajo-Rubio and Sosvilla-Rivero (1993)	Spain	0.19 5/	CD, LL	TS, 1964–88
Eisner (1994)	U.S.	0.27	CD, LL	TS, 1961–91
Ferreira (1994)	U.S.	0.08 6/	CD, LL	TS, Q, 1975–86
Sturm and de Haan (1995)	Netherlands	0.41	CD, LL	TS, 1949–85
		0.26	CD, LL	

Table 2. Overview of Empirical Studies: The Production Function Approach (Concluded)

Author	Country	β	Specification	Data
Dalamagas (1995)	Greece	0.53 7/ 6/	TL	TS, 1950–92
Ai and Cassou (1995)	U.S.	0.15	CD, DL	TS, 1947–89
Otto and Voss (1996)	Australia	0.17	CD, LL	TS, Q, 1959III–92II
Wylie (1996)	Canada	0.11–0.52	CD, LL	TS, 1946–91
Crowder and Himarios (1997)	U.S.	0.17–0.38	CD, LL	TS, 1947–89
Flores de Frutos et al. (1998)	Spain	0.21 5/	CD, LL	TS, 1964–92
Ramirez (1998)	Mexico	0.12 6/	CD, DL	TS, 1950–90
Mamatzakis (1999)	Greece	0.25	CD, LL	TS, 1959–93
Costa et al. (1987)	U.S.	0.19–0.26	TL	CS, 48 states, 1972
Merriman (1990)	U.S.	0.20	TL	CS, 48 states, 1972
	Japan	0.43–0.58	TL	P, 9 regions, 1954–63
Munnell and Cook (1990)	U.S.	0.15	CD, LL	P, 48 states, 1970–86
Aschauer (1990)	U.S.	0.11	CD, LL	P, 50 states, 1965–83
Eisner (1991)	U.S.	0.17 2/	CD, LL	P, 48 states, 1970–86
Garcia-Milà and McGuire (1992)	U.S.	0.04–0.05	CD+TL, LL	P, 48 states, 1969–83
Munnell (1993)	U.S.	0.14–0.17	CD, LL	P, 48 states, 1970–86
Evans and Karras (1994)	U.S.	n.a. 2/	CD, TL, LL, DL	P, 48 states, 1970–86
Holtz-Eakin (1992)	U.S.	n.a. 2/	CD, LL	P, 48 states, 1969–86
Pinnoi (1994)	U.S.	0.08	TL	P, 48 states, 1970–86
Baltagi and Pinnoi (1995)	U.S.	n.a. 2/	CD, LL	P, 48 states, 1970–86
Mas et al. (1996)	Spain	0.08	CD, LL	P, 17 regions, 980–89
Garcia-Milà et al. (1996)	U.S.	n.a. 2/	CD, DL	P, 48 states, 1970–83

Key: CD=Cobb-Douglas, LL=estimated in log levels, DL=estimated in first differences of logs, TL=translog in levels, TS=time series, CS=cross-section, P=panel data, and Q=quarterly data.

1/ No constraints on the production function imposed.
2/ Coefficient is insignificant at the 5 percent level.
3/ Study of 11 OECD countries. Only the coefficients of the listed countries were significant.
4/ Finds implausible values of the coefficients.
5/ Cointegrating relationship identified.
6/ Public investment rather than the public capital stock.
7/ Only when fiscal deficit is included in the equation, otherwise the coefficient is insignificant.

Table 3. Summary Descriptive Statistics for Studies Estimating the Production Elasticity of Public Capital

	Time Series: National Data				Cross-Section/Panel		All Studies
	U.S.		OECD 1/ 2/		U.S.	OECD 1/	
	Levels	First difference	Levels	First difference	Levels	First difference	Levels
Average	0.25	0.22	0.25	0.39	0.14	0.17	0.22
Number of observations	11	3	17	9	7	9	26
Standard deviation	0.11	0.06	0.10	0.17	0.06	0.13	1.2
95 percent confidence interval	[0.19, 0.31]	[0.15, 0.30]	[0.20, 0.30]	[0.28, 0.50]	[0.09, 0.19]	[0.08, 0.26]	[0.18, 0.27]

Source: Based on the overview of studies presented in Table 2.

1/ Studies for selected OECD countries.
2/ The average production elasticity derived from studies employing first differences drops to 0.23 if the study of Ford and Poret (1991) is excluded.

Table 4. Composition of the Portuguese Capital Stock, 1975 and 1995

	Percent of Total 1975	Percent of Total 1995
Total capital stock	100.0	100.0
Total private capital stock	87.2	85.1
Equipment and transport material	25.4	29.6
Construction	61.8	55.5
Residential	33.1	30.7
Nonresidential	28.7	24.8
Total public capital stock	12.8	14.9
Equipment and transport material	0.6	1.4
Construction	12.3	13.5
Buildings	3.0	4.9
Other (including core infrastructure)	9.2	8.6
Public-private capital stock ratio	0.15	0.18

Source: Historical Series of the Bank of Portugal.

Table 5. Unrestricted Estimates of the Production Function in Levels, 1965–95

$$\ln(Y) = \underset{(1.07)}{1.43} + \underset{(16.06)^{**}}{0.374}\ln(K) + \underset{(5.10)^{**}}{0.667}\ln(L) + \underset{(3.15)^{**}}{0.186}\ln(G) + \underset{(6.97)^{**}}{0.682}\ln(CU) - \underset{(-3.96)^{**}}{0.02^{*}}D_{7585}$$

R^2 ajd. = 0.998, D-W = 0.74, E-G = -2.49 1/

	F-statistic	Probability
Wald test on the restriction: $\alpha + \beta + \gamma = 1$ 2/	1.52	0.229

$$\ln(Y) = \underset{(3.70)^{**}}{0.337}\ln(K) + \underset{(7.93)^{**}}{0.193}\ln(G) + \underset{(32.11)^{**}}{0.901}\ln(L)$$

R^2 ajd. = 0.992, D-W = 0.62, E-G = -2.38

1/ Engle-Granger (E-G) procedure which tests whether the $I(1)$ variables in the equation are cointegrated. The Davidson and MacKinnon (1993) asymptotic critical values for cointegration are -4.10 (-4.64) for the 5 percent (and 1 percent) level, respectively.

2/ Tests whether production can be characterized by a constant returns to scale specification.

Table 6. Estimates of the Production Function in Levels, 1965–95 1/

	I		II		III	
	OLS	C-O	OLS	C-O	OLS	C-O
Constant	3.061 (79.26)	3.075 (8.46)	2.482 (17.14)	2.337 (7.69)	3.043 (113.94)	3.211 (11.33)
Ln(*L*/*K*)	0.405 (59.31)	0.395 (11.48)	0.321 (15.46)	0.314 (8.72)	0.390 (71.50)	0.391 (16.58)
Ln(*G*/*K*)	0.215 (7.73)	0.273 (4.26)	0.256 (5.35)	0.205 (1.94)	0.282 (12.45)	0.371 (5.65)
Ln(*CU*)	0.775 (12.21)	0.807 (13.43)			0.820 (18.51)	0.883 (15.45)
Ln(*U*)			-0.07 (-5.16)	-0.07 (-3.88)		
D_{7585}	-0.03 (-5.48)	-0.01 (-1.94)			-0.688 (-5.78)	-0.546 (-3.26)
D_{7585} ln(*G*/*K*)					-0.355 (-5.50)	-0.288 (-3.18)
ρ 2/		0.876 (7.48)		0.656 (4.40)		0.787 (5.071)
R^2 adj.	0.997	0.998	0.986	0.992	0.998	0.999
D-W	0.63	1.34	0.68	1.42	1.14	1.88
E-G 3/	-2.14 (0)		-2.95 (4)		-3.30 (0)	

1/ t-statistics in parentheses.
2/ Coefficient of the C-O estimation procedure.
3/ Engle-Granger (E-G) test for cointegration. The number in parentheses refers to the number of lags—determined by Akaike's Information Criterion included in the unit root test on the estimated residuals.

Table 7. Disaggregated Estimates of the Production Function 1/

	All Three Components	Core Infrastructure	Buildings	Transport Material and Equipment
Constant	0.440	0.344	-1.070	2.535
	(0.11)	(1.25)	(-0.39)	(1.60)
Ln(*K*)	0.506	0.403	0.367	0.374
	(6.15)	(18.85)	(3.54)	(15.01)
Ln(*GC*)	-0.018	0.185		
	(-0.06)	(4.71)		
Ln(*GB*)	-0.147		0.078	
	(-1.61)		(0.80)	
Ln(*GE*)	0.139			0.104
	(0.75)			(4.63)
Ln(*L*)	0.930	0.751	1.195	0.720
	(3.76)	(3.53)	(3.61)	(3.28)
Ln(*CU*)	0.468	0.651	0.462	0.608
	(2.92)	(6.55)	(2.77)	(6.41)
D_{7585}	-0.03	-0.032	-0.029	-0.029
	(-3.62)	(6.55)	(-2.66)	(-3.70)
R^2 adj.	0.999	0.998	0.996	0.998
D-W	1.04	0.78	0.87	0.72
E-G 2/	-2.95	-2.65	-4.00**	-2.46

1/ t-statistics in parentheses.
2/ Asterisks denote significance at the 1 percent level.

Table 8. Augmented Dickey-Fuller Tests for Nonstationarity 1/

variable	ADF	Lags 2/	variable	ADF	Lags	Variable	ADF	lags
$\ln(Y/K)$	-4.68**	7	$\Delta\ln(Y/K)$			$\Delta^2 \ln(Y/K)$		
$\ln(L/K)$	-2.59	1	$\Delta\ln(L/K)$	-1.70	0	$\Delta^2 \ln(L/K)$	-7.28**	0
$\ln(G/K)$	-3.97**	5	$\Delta\ln(G/K)$			$\Delta^2 \ln(G/K)$		
$\ln(GC/K)$	-4.73**	6	$\Delta\ln(GC/K)$			$\Delta^2 \ln(GC/K)$		
$\ln(GB/K)$	-0.25	7	$\Delta\ln(GB/K)$	-2.79	5	$\Delta^2 \ln(GB/K)$	-7.03**	0
$\ln(GE/K)$	-0.69	1	$\Delta\ln(GE/K)$	-2.15	1	$\Delta^2 \ln(GE/K)$	-4.05**	0
$\ln(Y)$	-1.96	1	$\Delta\ln(Y)$	-4.00**	2	$\Delta^2 \ln(Y)$		
$\ln(K)$	-2.60	2	$\Delta\ln(K)$	-3.79**	7	$\Delta^2 \ln(K)$		
$\ln(L)$	-0.74	1	$\Delta\ln(L)$	-3.72**	0	$\Delta^2 \ln(L)$		
$\mathrm{Ln}(G)$	-0.47	2	$\Delta\ln(G)$	-2.66	1	$\Delta^2 \ln(G)$	-4.39**	0
$\mathrm{Ln}(GC)$	0.25	1	$\Delta\ln(GC)$	-2.09	0	$\Delta^2 \ln(GC)$	-5.18**	0
$\mathrm{Ln}(GB)$	-2.61	1	$\Delta\ln(GB)$	-1.93	4	$\Delta^2 \ln(GB)$	-7.44**	0
$\mathrm{Ln}(GE)$	-0.65	3	$\Delta\ln(GE)$	-3.84**	2	$\Delta^2 \ln(GE)$		
$\mathrm{Ln}(CU)$	-1.37	1	$\Delta\ln(CU)$	-3.76**	0	$\Delta^2 \ln(CU)$		
$\mathrm{Ln}(u)$	-5.51**	2	$\Delta\ln(u)$			$\Delta^2 \ln(u)$		

1/ The tests are conducted with a constant, ϕ, included in the following equation:

$\Delta y_t = \phi + \rho y_{t-1} + \sum_{s=1}^{n} \Omega_s \Delta y_{t-s} + \varepsilon_t$, where y_t is the relevant time series, ε_t is an i.i.d. sequence of random variables.

2/ The number of autoregressive lags is chosen so as to minimize Akaike's (1969) Information Criterium. The null hypothesis is that the variable under investigation has a unit root (i.e., $\rho = 1$) against the alternative that it does not. A value of the augmented Dickey-Fuller (ADF) statistic exceeding the critical value for the specific lag length leads to a rejection of the null hypothesis.

* Significant at the 5 percent level.

** Significant at the 1 percent level.

Table 9. Estimates of the Production Function in First Differences, 1965–95 1/

	Total Stock		Core Infrastructure		Buildings		Equipment and Transport Material	
Constant	0.010 (1.03)	0.000 (0.13)	0.007 (1.68)	0.006 (1.34)	0.013 (3.43)	0.014 (3.41)	0.011 (1.70)	0.004 (0.72)
$\Delta\ln(L/K)$	0.481 (4.51)	0.347 (3.14)	0.489 (4.81)	0.449 (4.22)	0.673 (13.54)	0.687 (12.85)	0.639 (7.74)	0.537 (6.40)
$\Delta\ln(G/K)$	0.199 (2.03)	0.371 (3.30)						
$\Delta\ln(GC/K)$			0.175 (2.05)	0.216 (2.35)				
$\Delta\ln(GB/K)$					0.022 (0.36)	0.035 (0.56)		
$\Delta\ln(GE/K)$							0.034 (0.54)	0.188 (2.33)
Interaction Dummy		-0.415 (-2.54)		-0.184 (-1.16)		-0.081 (-0.75)		-0.185 (-2.67)
$\Delta\ln(CU)$	0.825 (14.35)	0.893 (15.18)	0.829 (14.26)	0.854 (13.83)	0.770 (14.11)	0.773 (14.01)	0.762 (13.82)	0.770 (15.45)
R^2 adj.	0.947	0.954	0.944	0.945	0.936	0.935	0.943	0.948
D-W	1.51	1.50	1.50	1.46	2.11	2.25	2.00	2.04
E-G 2/	-4.22 (0)		-4.19 (0)		-4.63 (1)		-5.43 (0)	

1/ t-statistics in parentheses below the coefficients.
2/Engle-Granger (E-G) test for cointegration. The number in parenthesis is the number of lags included in the unit root test on the estimated residuals.

Table 10. Johansen's Cointegration Analysis of Portuguese Production 1/

	Null hypothesis for test statistics			
Statistic 2/	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
Eigenvalue	0.727	0.477	0.362	0.124
$\lambda_{\max}$	37.6**	18.2	13.0	3.9*
$\lambda^a_{\max}$	27.3*	13.6	9.4	2.8
95 percent critical value	27.1	21.0	14.1	3.8
λ_{trace}	73.3**	35.7**	16.9*	3.9*
λ^a_{trace}	53.1*	25.8	12.2	2.8
95 percent critical value	47.2	29.7	15.4	3.8

1/ The VAR includes two lags on each variable and is estimated over the period 1967–95.

2/ The statistics $\lambda_{\max}$ and λ_{trace} are the maximum eigenvalue and trace eigenvalue statistics ($\lambda^a_{\max}$ and λ^a_{trace} adjust for the degrees of freedom) for testing cointegration in the Johansen procedure. The null hypothesis is defined in terms of the cointegration rank, r. The critical values are taken from Osterwald and Lenum (1992): (**) denotes significance at the 1 percent level whereas (*) indicates significance at the 5 percent level.

Table 11. Likelihood Ratio Tests to Determine the Lag Length 1/

Lag	Loglikelihood	F-test 2/
1	477.10	308.89**
2	515.57	4.43**
3	537.42	1.43
4	562.22	1.03
5	603.60	0.77

1/ F-form of the likelihood ratio test. At a 1 percent level, the restrictions involved in moving to a lag length of 2 cannot be rejected.

2/ The asteriks indicate signficance at the 1 percent level.

Table 12. Granger-Causality Tests

	G		*K*		*L*		*Y*	
Direction of causality:	sum 1/	F-stat. 2/	sum	F-stat.	sum	F-stat.	sum	F-stat.
$G \rightarrow$	0.972		-0.039	1.13	0.066	3.95*	0.139	3.30*
$K \rightarrow$	-0.024	1.20	0.926		-0.036	1.66	0.277	3.74*
$L \rightarrow$	-0.238	1.21	0.233	1.97	0.438		0.856	0.74
$Y \rightarrow$	0.121	1.54	0.057	8.45**	0.079	7.47**	0.202	
R^2	0.99		0.99		0.97		0.99	

1/ The sum of the coefficients in the VAR with common lag length is included as a rough indicator of the sign of the relationship between two variables.

2/ The reported F-statistic are the Wald statistic for the null hypothesis that lagged values of variable *x* cannot improve the explanation of the variation in variable *y*. Significant *x* coefficients—which means that *y* is Granger caused by *x*—are indicated by asteriks: (*) denotes significance at the 5 percent level and (**) indicates significance at the 10 percent level.

Table 13. Variance Decompositions 1/

Equation	Innovation	1	3	5	7	10	15
G	g	100.0	93.6	88.0	86.7	85.7	81.3
	k	0.0	2.2	5.0	6.1	7.0	10.6
	l	0.0	0.4	0.3	0.2	0.2	0.3
	y	0.0	3.8	6.7	7.0	7.1	7.8
k	g	22.2	11.8	9.5	8.7	8.2	7.5
	k	77.8	66.8	60.8	59.6	60.4	61.8
	l	0.0	112.5	23.1	26.4	26..5	25.7
	y	0.0	7.7	6.5	5.3	4.9	5.0
l	g	0.5	3.9	9.6	15.2	19.0	19.9
	k	9.1	26.2	25.2	24.1	22.4	23.4
	l	90.4	56.0	52.7	48.7	47.3	44.6
	y	0.0	13.9	12.5	12.0	11.3	12.0
y	g	0.5	0.7	2.6	5.9	9.2	10.3
	k	41.0	44.0	42.7	40.7	38.7	39.7
	l	18.1	30.7	33.8	32.7	32.3	30.5
	y	40.4	24.6	20.9	20.6	19.8	19.5

1/ The columns contain the percentage of forecast variance of a variable in time period t (=1,..,15) explained by a shock to one of the four variables (g, k, l, or y). The following variable ordering is used: g, k, l, y.

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