Public Capital and Private Production in Australia

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I. Introduction

Recent empirical research by Ratner [31], Aschauer [4; 5] and others seeks to demonstrate that publicly provided capital should enter as a complementary input to private production.[1] The general approach of these studies is to specify and estimate a functional form for private production which is dependent upon both private inputs to production and some measure of public capital. With few exceptions, see for instance Holtz-Eakin [18], across a variety of economies and at different levels of aggregation, public capital is shown to be a significant input to private production.

This paper uses a newly constructed quarterly data set for the Australian economy to examine two specific criticisms levelled at many of the previous empirical studies. The first is the possibility that the estimated relations reflect a spurious correlation between variables with purely coincidental low frequency movements; that is, the variables are non-stationary and the regressions are spurious. This criticism is addressed by Aschauer [3] where he argues that his previous results are valid even if the variables are non-stationary. More recently, Clarida [10] and Lynde and Richmond [21] use techniques suitable for non-stationary variables to confirm Aschauer's [4] results for the United States as well as for a number of European countries. We also employ co-integration techniques, those developed by Phillips and Hansen [29] and Hansen [17], to estimate a production function for the private economy in Australia during the post-war period.

The second issue we address is the question of causality between public capital and private production. Aschauer [4] argues that causality runs from public capital to private output (or private factor productivity). However, an equally plausible interpretation of the estimated relationships is that they reflect the response of public investment to private production; in other words, public capital stocks are possibly endogenous. This endogeniety may arise from various political constraints on public expenditure or, alternatively, the relationship may reflect aspects of the public decision-making process where public investment decisions are a response to private output growth or possibly private investment. To examine this second criticism, we employ vector

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1. These studies, at various levels of aggregation, are now numerous. Munell [22] provides a discussion of the main issues and survey of the literature. More recent studies include Lynde and Richmond [21] and Holtz-Eakin [18] for the United States and Otto and Voss [25] for Australia.

autoregression (VAR) techniques associated with Sims [33] to examine the effect of public capital on private sector variables and also to examine whether there is any feedback from the private sector variables to the public capital stock. Unlike previous studies, which use annual data, our quarterly data provides us with a sample size large enough to effectively use these techniques.

The paper proceeds as follows. Section II presents the basic representation of private production used in the empirical analysis. In section III, single equation cointegration techniques are used to estimate the long run relationship between private production and private and public inputs. A VAR analysis is used in section IV to examine the dynamic interactions among the variables of interest. Section V concludes.

II. Representations for Private Production

Private production is represented using a Cobb-Douglas production function. We allow for three inputs: private capital, public capital, and private labor. The justification for treating public capital as an input to private production is discussed by Arrow and Kurz [2]; the Cobb-Douglas functional form is employed for its suitability in empirical analysis. Four alternative specifications for private production are considered:

$$y - k = \alpha_{n1}(n - k) + a_1 \tag{1}$$

$$y - k - g = \alpha_{n2}(n - k - g) + a_2 \tag{2}$$

$$y - k = \alpha_{n3}(n - k) + \alpha_{g3}(g - k) + a_3 \tag{3}$$

$$y - k = \alpha_{n4}(n - k) + \alpha_{g4} \cdot g + a_4 \tag{4}$$

where y is private sector output, n is private sector labor, k is private sector capital, g is public capital and $a_i, i = 1 \ldots 4$, is a non-observable measure of technology. All variables are in logarithms.

These four models represent alternative ways of treating public capital as an input to private production. Model (1) assumes constant returns to scale (CRS) between private inputs and no role for public capital. Model (2) includes public capital as an input to production but as a perfect substitute for private capital; the model assumes CRS across all inputs. In many studies of aggregate production, for example the real business cycle studies initiated by Kydland and Prescott [20], it is common to aggregate public and private capital stocks, effectively treating them as perfect substitutes. As far as private production is concerned, comparison of model (2) with the other specifications provides some information about the suitability of doing so. Models (3) and (4) treat public capital as a complementary input. Model (3) assumes CRS across all inputs and, necessarily, decreasing returns to scale in private inputs. This representation is suitable if public capital is not a pure public good. Model (4) assumes CRS between private labor and capital but allows increasing returns to scale (IRS) across all three inputs. This representation is suitable if public capital is a non-rival input in private production.

The latter two models form the basis of much of the empirical literature on public capital provision. Ratner [31] estimates model (3) for the United States and obtains an estimate for the output elasticity of public capital of 0.06. Aschauer [4] estimates both (3) and (4) for the United States, finding greatest support for the CRS specification. His elasticity estimates are in the order of 0.40, notably larger than Ratner's estimates. Otto and Voss [25] estimate models (3) and (4) for the Australian economy using annual data and obtain results very similar to Aschauer's.

A feature of Aschauer's results, and by implication our results for Australia, is that the elasticity estimates imply very high rates of return to additional investment in public capital compared to estimates obtained from project based cost-benefit studies. Consequently, a number of authors, for example Winston and Bosworth [34], have argued that the public capital marginal productivity estimates from aggregate production functions are implausibly large.[2]

Thus we are interested in seeing if these large elasticities are robust to the use of econometric techniques which can account for non-stationary series.

III. The Long-Run Elasticities

In this section, we use techniques suitable for non-stationary data, specifically integrated processes, to estimate each of the four models of private production. We first establish that the variables of interest are indeed integrated of order one, $I(1)$, and then seek to determine if the input variables and private production are cointegrated. If these variables are cointegrated then we can reject the argument that previously estimated relationships in levels are spurious and in addition obtain consistent output elasticity estimates from the low-frequency components of the data.

Each model is estimated using quarterly time series data for Australia over the period 1959:3 to 1992:2. The data we use is not directly available from published sources. Measures of output, capital stock and hours worked for the private sector are constructed along with a series for government capital, which includes both general government capital and the capital stock of public trading enterprises. Full details on primary data sources and the method of construction are provided in an appendix.

Since the technology variable a_i is itself not observable, we follow standard practice and treat this as part of the disturbance term. This creates the additional possibility that evidence of no cointegration between the other variables may be indicative of an integrated technology process and not a rejection of the model itself.[3] Pursuit of these matters, however, is beyond the scope of this paper and we are limited to simply recognising the possibility.

The first step is to perform tests for unit roots on all observable variables in equations (1) to (4), both in levels and in first-differences, to determine if these variables are $I(1)$. Table I presents augmented Dickey-Fuller (ADF) tests. The evidence suggests that the variables $y-k$, $y-k-g$, $n-k$, and $n-k-g$ are $I(1)$ variables. The results of the ADF test for the variables $g-k$ and g are not as sharp. Whether we can view these two variables as being stationary in first-differences is sensitive to the choice of lag length in the ADF test regression. Of the two series, the latter seems least likely to be first-difference stationary. In fact, a two-sided ADF test leads to the rejection of the null hypothesis that g is $I(1)$ in favor of the alternative that it follows an explosive autoregressive process. This indicates that g behaves more like an $I(2)$ than an $I(1)$ process and leads us to suspect that model (4), the IRS specification, is unlikely to perform well on the quarterly data set.[4]

Given these results which suggest that most of the variables in models (1) to (4) are $I(1)$,

2. Munnell [22] provides a general discussion of this issue.

3. An $I(1)$ technology shock is not an uncommon assumption in equilibrium models of business cycle fluctuations. See for example Christiano and Eichenbaum [9].

4. As a general rule, we might expect measures of the capital stock to behave as an I(2) series. The equation for capital accumulation is: $K_t = (1-\delta)K_{t-1} + I_t$. So if gross investment I_t is an $I(1)$ variable, then as the depreciation rate δ goes to zero, the capital stock will approximate an $I(2)$ series.

Table I. Augmented Dickey-Fuller Tests

Variable	Lags(m):	1	2	3	4
$y-k$		−2.83	−3.13	−3.14	−3.06
$n-k$		−0.86	−0.93	−1.18	−1.15
$g-k$		−3.05	−2.70	−2.34	−2.02
$y-k-g$		−2.57	−2.72	−3.07	−2.44
$n-k-g$		−0.51	−0.63	−0.97	−0.84
g		0.04	−0.47	−0.15	0.11
$\Delta(y-k)$		−7.57	−5.93	−6.51	−5.65
$\Delta(n-k)$		−8.07	−5.78	−5.64	−5.71
$\Delta(g-k)$		−2.43	−3.02	−3.32	−2.79
$\Delta(y-k-g)$		−7.73	−6.01	−6.63	−5.75
$\Delta(n-k-g)$		−7.52	−5.39	−5.23	−5.22
Δg		−2.85	−2.43	−3.69	−3.72

The ADF test regression includes a constant, a time trend and up to m lags of the dependent variable. Critical values are from Fuller [16, 381–82]. The one, five and ten per cent, critical values are −3.96, −3.41, and −3.12 respectively. All models are estimated using quarterly data for the period 1959:3 to 1992:2.

we now estimate each of the models and test whether any of them represent valid cointegrating relationships. To estimate these models we embed them in a triangular representation for cointegrated systems of the following form:

$$Z_t = \mu + \beta t + \mathbf{X}_t'\alpha + u_{1t} \tag{5}$$

$$\Delta\mathbf{X}_t = \delta + u_{2t} \tag{6}$$

where Z_t is the dependent variable in models (1) to (4) and $\mathbf{X}_t$ is a vector of the stochastic regressors. A set of conditions on the vector of error terms $u_t = (u_{1t}, u_{2t})'$ for (5) and (6) to be a valid cointegrating system are given by Park and Phillips [28]; these are assumed to be satisfied here.

Conditional on the assumption that each of our models represents a valid cointegrating relationship, we can obtain both consistent and (asymptotically) efficient estimates of the model parameters using Phillips and Hansen's [29] "fully modified" (FM) estimator. The version of this estimator used in this study is the one developed in Hansen [17]. This involves the use of a pre-whitened kernel estimator with a plug-in bandwidth to estimate the long-run covariance matrices; for further details see Hansen. The results obtained for models (1) to (4) are reported in Table II. We are interested in discovering whether one (or more) of the models containing public capital provides relatively strong evidence of a cointegrating relationship.

For each model we perform two tests for cointegration. The first is a residual-based test of the null hypothesis of no cointegration, using the ADF test and the FM coefficient estimates. In addition, we perform a variable addition test due to Park, Ouliaris and Choi [27] which sets up cointegration as the null hypothesis. This test can be viewed as a general test for a misspecified cointegrating relationship. The variables used to augment each model are powers of the time trend: t^2, t^3, t^4. Under the null hypothesis of no misspecification in the cointegrating model, these trend variables will have no additional explanatory power. A Wald test (adjusted for the FM estimate of the long-run error variance) is used to test the joint significance of these additional regressors. The test statistic is denoted H(1,4) and has a limiting chi-squared distribution with three degrees of freedom under the null hypothesis, see Ogaki and Park [23] for an application.

Table II. Long Run Private Sector Production Functions

Coefficient	Model:	(1)	(2)	(3)	(4)
μ		−2.7179	−3.0595	−2.6112	−6.7571
		(0.1645)	(0.0905)	(0.1274)	(2.4839)
β		0.0010	0.0016	0.0011	0.00005
		(0.0005)	(0.0003)	(0.0004)	(0.0006)
α_n		0.4990	0.4425	0.4375	0.6840
		(0.0558)	(0.0442)	(0.0427)	(0.1424)
α_g				0.1675	0.2961
				(0.0796)	(0.1786)
ADF	(1)	−3.37	−4.01*	−4.07*	−3.68
	(2)	−3.46	−4.17*	−4.29*	−3.67
	(3)	−3.86*	-4.77^{+}	-4.85^{+}	−4.05*
	(4)	−3.27	−4.04*	−4.03*	−3.35
H(1,4)		14.52^{+}	9.21*	11.41*	154.93^{+}

These models are estimated using Phillips and Hansen's fully modified procedure as described in Hansen [17]. A * indicates a test statistic is significant at the five per cent level while a $^{+}$ indicates significance at the one per cent level. The H(1,4) statistic is distributed as a chi-squared with four degrees of freedom under the null hypothesis. The critical values for the ADF test of no cointegration are taken from Phillips and Ouliaris [30]. All models are estimated using quarterly data for the period 1959:3 to 1992:2.

The ADF statistics in Table II indicate that the strongest evidence against the null of no cointegration arises for models (2) and (3), both of which include public capital as an input and impose CRS. This supports the argument that public capital does play a role in private production; however, the similarity of the results for these two models does not allow us to discriminate between a model that treats public and private capital as separate inputs and one that treats public and private capital as perfect substitutes. For models (1) and (4), evidence against the no cointegration hypothesis is much weaker and is sensitive to the choice of the lag length for the ADF test. The absence of cointegration for model (1) is consistent with the view that public capital is an important input to the private production process. The lack of cointegration for model (4) is likely to be a reflection of our previous findings relating to the time series properties of the variable g.

Turning to the variable addition test we find significant H(1,4) statistics for all four models. Strictly speaking this suggests that none of the models is a completely adequate specification of long run private sector production. However, the strongest evidence against the null of cointegration arises from models (1) and (4) which is consistent with the ADF results above.

Although the H(1,4) test statistic does point to the need for some caution, we view the results in Table II as providing further support of the importance of public capital for private production. In addition, the parameter estimates of the output elasticities associated with either model (3) or (4) are quite reasonable. Consider model (3) which is the least restrictive representation of private production. The associated output elasticities are: private labor, 0.44; private capital, 0.39 and public capital, 0.17. An interesting feature of these results is that the point estimate for the public capital elasticity is less than half that obtained by Aschauer [4] and Otto and Voss [25]. The estimate is, however, very similar to that obtained by Lynde and Richmond [21], which uses similar statistical techniques for the United States. We note, however, that there is considerable uncertainty about the point estimate: the 95 percent confidence interval is 0.01 to 0.32.

IV. Short-Run Dynamics

While the preceding single equation approach provides information about the long-run relationships, it does not tell us anything about the short-run dynamic relationships between public capital and the private sector variables. To examine these relationships, we specify a vector autoregression (VAR) model and use the variance decomposition and impulse response techniques developed by Sims [33]. We motivate our investigation of these short run relationships from three perspectives, each of which focuses on the relationship between public capital levels and one of the other variables considered in the estimated production function.

First, and of primary interest, is the possibility that the estimated long run relationship of the preceding section does not represent a description of private sector production but rather arises because public investment responds to levels of private production. Aschauer [4] suggests that this reverse causation might occur if public expenditure is a superior good, rising more than proportionally with increases in per capita income, an example of Wagner's Law. Alternatively, reverse causation may be motivated from a political economy perspective. For example, during recessionary periods of the business cycle, with a consequent fall in tax revenues, governments may be unwilling to finance public investment, especially if the returns are not immediately visible.[5] During expansionary periods of the business cycle, public investment is more easily funded and possibly in higher demand by the private sector. If this adequately describes the public decision making process then we might expect a positive relationship between levels of public capital and private production. In this case, it is difficult to know how to properly interpret the estimated coefficients of the long run analysis.

The causality issue is considered elsewhere but not in a complete manner. Aschauer [3] argues, from a number of perspectives, that the causality argument cannot fully explain the empirical results he and other authors obtain. While generally convincing, he does not offer any direct empirical evidence. Holtz-Eakin [19], Munell [22], and Easterly and Rebelo [12] provide some preliminary evidence for aggregate economies although none of these authors provides any conclusive evidence.

The second motivation is the relationship between public and private levels of capital. While not directly affecting the interpretation of the long-run results, the direction of causality between these variables is of considerable interest. The first possibility, similar to the reverse causality discussion, is that public investment responds to private investment. In this case, the interest is the nature and length of adjustment process. The second possibility is that public capital provision can initiate private investment and hence private production and employment. Given the results of the previous section which indicate that the marginal product of private capital is an increasing function of the level of public capital, this seems intuitive. Baxter and King [6] demonstrate this argument formally using a neo-classical growth model with production similar to that considered here. In their model, the result on private investment and output is unambiguously positive because of the increased marginal productivity of private capital; the effect on private employment, however, is ambiguous because of off-setting wealth effects. Empirical evidence of these sorts of effects might provide support to arguments for the use of public infrastructure programmes to initiate recovery from periods of recession. Finally, it is possible that public capital crowds out

5. The lack of immediate and visible benefits of public investment has been identified as a possible bias against its efficient provision by Rogoff [32]. This argument has been used by Alesina, Gruen and Jones [1] to explain the reductions in public investment in Australia in the late 1980s. It also has considerable support at an institutional level, see for example Berne and Stiefel [7].

private capital; this might be the case if private and public capital are highly substitutable and both are, on average, near optimal levels. These latter two arguments, potentially offsetting, are the focus of Aschauer [5]. He finds evidence of both effects for the United States with the suggestion that the crowding out effect is dominated by private investment responding to increased public capital investment.

This issue of causality between public and private capital is also addressed by Eberts and Fogarty [13]. The authors consider municipal data for the U.S. and find evidence of causality in both directions, depending upon the time period and location. While interesting in its own right, this evidence does not directly extend to the issue of causality at the aggregate level because of the degree of capital mobility within the U.S. While local public investment may attract private investment, if it does so at the expense of private investment elsewhere, then on aggregate, little or no increase in private investment occurs. To this end, the issue of causality at an aggregate level must be settled using aggregate data.[6]

The third motivation is the relationship between public capital and employment. While not generally presented as an argument in the literature on public investment, there is considerable evidence that some policy makers view public investment as a suitable means of reducing unemployment during recessionary periods of the business cycle. Whether or not such considerations are likely to lead to socially efficient levels of public investment is simply beyond the analysis here; we are, however, able to provide evidence on the efficacy of such programs by considering the response of private employment to changes in the level of public capital.

To pursue these issues, we consider an unrestricted vector autoregression model using the four variables of interest in levels: (y, k, g, n). Such a model does not impose the long-run restrictions identified in the previous section but does allow for these restrictions to be satisfied asymptotically. Although this is not as efficient as a model which imposes these restrictions, we use the unrestricted model to minimize problems associated with the imposition of invalid long-run restrictions.[7]

The results reported are based on a VAR model with four lags of each variable and a linear time trend. The VAR(4) specification reflects the use of quarterly data; the findings are not qualitatively sensitive to marginal changes in the lag length. Nor are the results qualitatively sensitive to re-estimation without a time trend.

The residual correlation matrix for the VAR model is reported Table III. The largest correlations are among the residuals of the private sector variables with the correlations between the public sector innovations and the other private sector variable innovations very close to zero (the largest is that between public capital and private labor input, a value of -0.14). This result should not be too surprising; one might expect private inputs, determined by decentralized behaviour, to experience innovations quite different from the collectively determined stock of public capital.

The low correlations between public and private capital presented in Table III also provide further support for the treatment of these variables as separate inputs to private production. If one interprets the innovations on the private capital stock as the response of investment to technological shocks (as is true in a simple stochastic growth model with Hicks neutral technological

6. Munell [22] and others cite this paper when discussing the issue of causality between private production and public capital. Since private production and private investment are obviously highly correlated, the relationship between public and private investment may provide indirect information in this regard. The Eberts and Fogarty study, however, is still not strictly appropriate because of its disaggregated nature. There is an obvious extension of these ideas to the international location of investment; Clarida [10] investigates this issue in a neo-classical growth model.

7. A complete set of results, including those for the vector error-correction model, are contained in Otto and Voss [26].

Table III. Residual Correlation Matrix for VAR(4) Model

	Variable			
	g	k	n	y
g	1.00	−0.02	−0.15	−0.13
k	0.02	1.00	0.25	0.42
l	−0.14	0.25	1.00	0.38
y	−0.06	0.40	0.38	1.00

The correlations above the diagonal are from a VAR model excluding a linear time trend while those below are from a VAR including a time trend.

change), then the evidence that the two capitals have uncorrelated innovations suggests that the two capital stocks are determined quite differently and should not be treated as an aggregate in a general equilibrium model.

Figure 1 presents the impulse response functions and associated standard errors for the four variables of the system based on Doan [11]. The ordering for the variables is (g, k, n, y). The effects of a one standard deviation shock to the orthogonalized innovation of each equation are traced out for 40 quarters. Since each variable is in logarithms, the vertical axes indicate (approximate) percentage changes. A key feature of the results is that the one standard deviation bounds on the estimated impulse response functions are relatively large, suggesting there must be a non-trivial degree of uncertainty about our conclusions.[8]

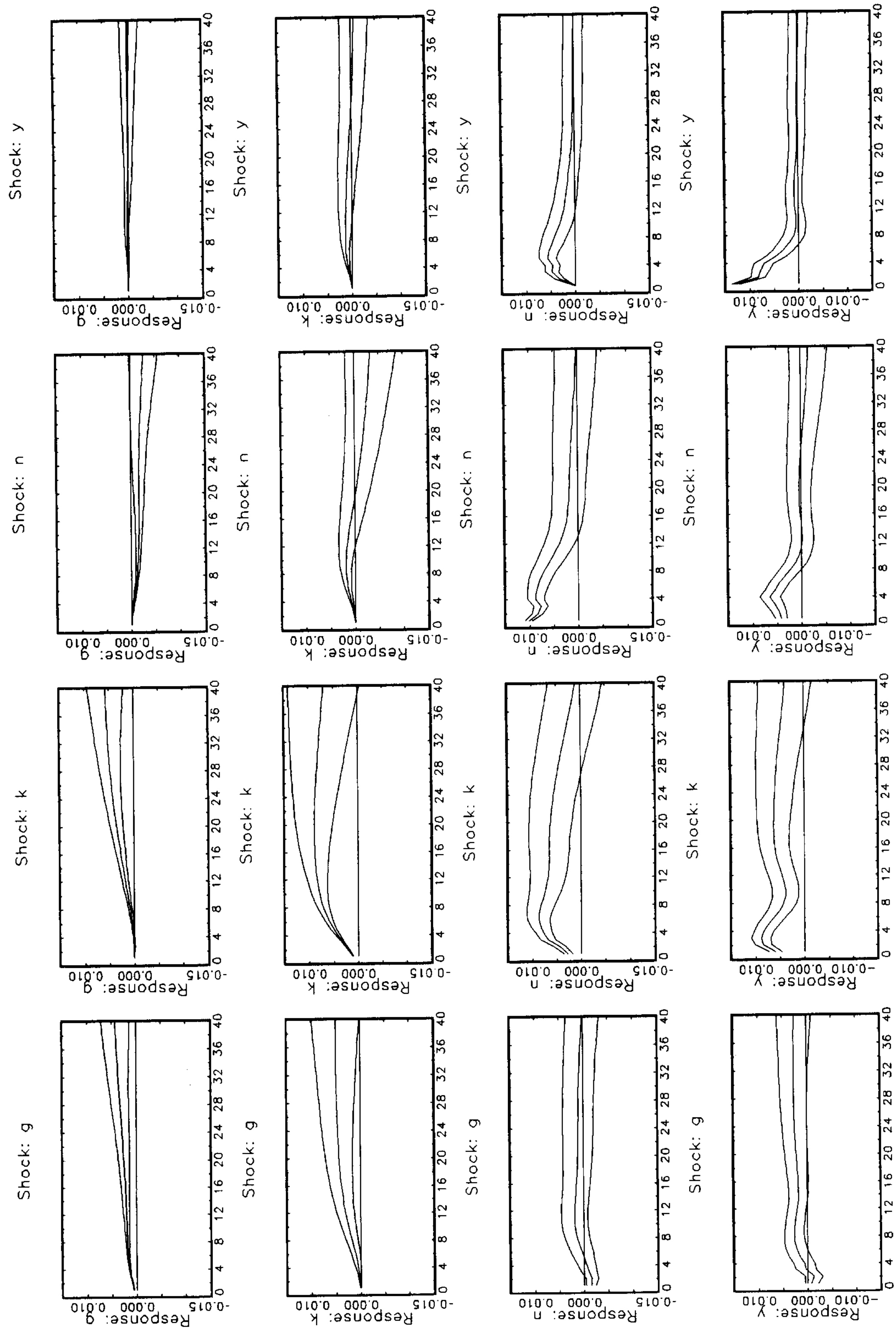
In general, the strongest interactions occur among the private sector variables. Innovations to public capital have very little effect on private sector hours and output. To underscore this result, compare the effect on these variables of a shock to public capital to that of a shock to private capital. For both, the response is much larger in the case of the private capital innovation; indeed, in response to the public capital shock there is only weak evidence of a significant response in hours and output and this is negative. The main impact of a shock to the stock of public capital is on private capital. A positive shock to public capital tends to have a positive lagged effect on private capital. This is consistent with the arguement that public and private capital are complementary inputs to private production. This positive response provides some evidence for the view that public capital provision can induce private investment in aggregate, as suggested by Aschauer [5] and Baxter and King [6].[9]

We may also consider what evidence there is for the reverse causation argument. In this regard, the impulse response functions indicate quite dramatically that public capital shows almost no response to innovations in private output. Similarly, public capital does not respond to innovations to private hours. However we do observe, as we might expect from two complementary inputs, that public capital responds to private capital innovations. These results provide evidence against the reverse causality argument that has been raised concerning the relationship between public capital and private output/productivity. An additional conclusion available from the impulse response functions is that the absence of any significant response of public capital to a shock to private output suggests that public investment has not been used to pursue counter-cyclical fiscal policy.

To ensure that choice of the ordering for the variables in the VAR is not responsible for

8. Such uncertainty is typical of unrestricted VAR models, see Eichenbaum [14].

9. The results concerning output and employment, however, are inconsistent with the predictions of Baxter and King. This may reflect the fact that we have not imposed any long-run restrictions on the estimated VAR.



Shock: g
Shock: k
Shock: n
Shock: y
Response: g
Response: k
Response: n
Response: y

Figure 1.

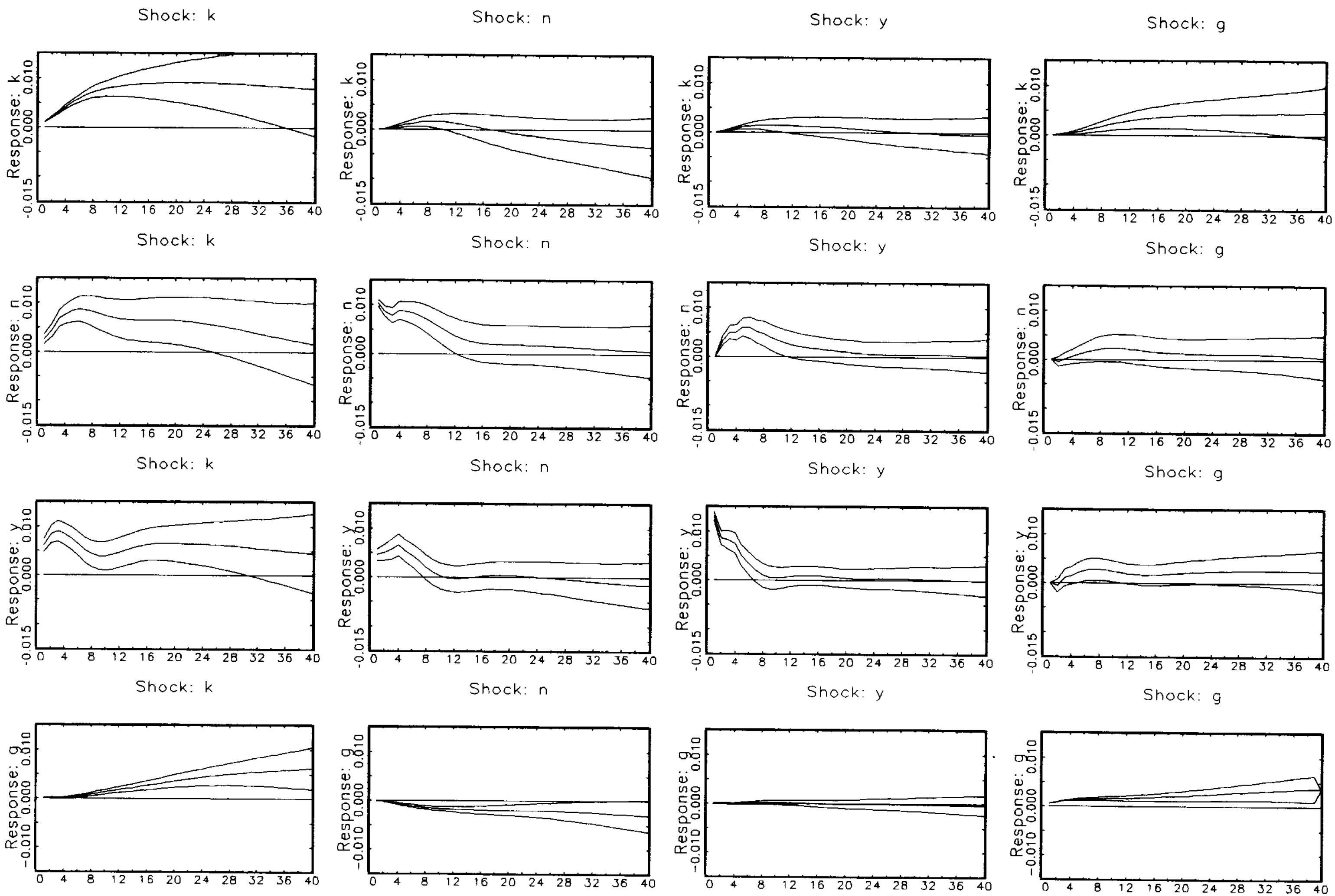

Shock: k
Shock: n
Shock: y
Shock: g
Response: k
Response: n
Response: y
Response: g

Figure 2.

Table IV. Variance Decomposition, Ordering (g, k, n, y)

Quarters	g		k		n		y	
Forecast variance of g (percent) explained by shock to:								
1	100	(0)	0	(0)	0	(0)	0	(0)
4	93	(4)	0	(1)	7	(4)	0	(1)
12	58	(13)	18	(11)	23	(11)	1	(3)
20	40	(15)	41	(17)	18	(13)	0	(3)
28	34	(15)	53	(18)	13	(13)	0	(4)
40	31	(16)	57	(20)	12	(14)	0	(4)
Forecast variance of k (percent) explained by shock to:								
1	0	(1)	100	(1)	0	(0)	0	(0)
4	1	(3)	94	(4)	2	(2)	3	(2)
12	6	(8)	86	(10)	4	(6)	3	(4)
20	12	(11)	84	(13)	2	(6)	2	(5)
28	16	(13)	81	(15)	2	(7)	1	(5)
40	19	(14)	75	(18)	5	(10)	1	(5)
Forecast variance of n (percent) explained by shock to:								
1	2	(3)	6	(4)	92	(5)	0	(0)
4	1	(3)	26	(9)	62	(9)	11	(5)
12	2	(4)	41	(14)	44	(13)	13	(9)
20	2	(6)	49	(17)	37	(14)	12	(9)
28	2	(7)	53	(17)	34	(15)	11	(9)
40	2	(8)	54	(18)	33	(15)	10	(9)
Forecast variance of y (percent) explained by shock to:								
1	0	(2)	16	(6)	8	(4)	75	(6)
4	0	(2)	35	(9)	16	(7)	49	(9)
12	4	(6)	42	(12)	17	(9)	36	(10)
20	5	(7)	52	(13)	14	(9)	29	(10)
28	7	(8)	57	(14)	11	(9)	25	(10)
40	10	(10)	58	(15)	11	(10)	21	(10)

These results are based on VAR(4) model including a linear time trend. Standard errors are given in brackets and were computed using the procedure described in Doan [11] with 500 draws from the posterior distribution of the VAR coefficients.

the above results, Figure 2 presents the impulse response functions for the alternative ordering (k, n, y, g). With this ordering, we allow public capital to respond to contemporaneous innovations in the other three variables; that is, public capital is treated as the 'most endogenous' of the four variables. As is evident from Figure 2 the alternative ordering does not produce qualitatively different results to those reported above.

To get a clearer indication as to the quantitative importance of the above effects, we report the variance decompositions associated with the above models. These indicate the proportion of the variance of the k-step ahead forecast error of a variable that is due to its own innovation and to the innovations of the variables in the system. The variance decompositions corresponding to Figures 1 and 2 are reported in Tables IV and V respectively.

Consider the ordering (g, k, n, y). From the first cell of Table IV it is apparent that over longer horizons (about 5 to 10 years) a significant percentage of the forecast error variance for public capital is due to private capital shocks; specifically, over 50 percent of the forecast error

Table V. Variance Decomposition, Ordering (k, n, y, g)

Quarters	k		n		y		g	
Forecast variance of k (percent) explained by shock to:								
1	100	(0)	0	(0)	0	(0)	0	(0)
4	95	(4)	2	(2)	3	(2)	1	(1)
12	87	(10)	3	(6)	6	(5)	7	(7)
20	85	(13)	2	(6)	2	(5)	12	(10)
28	82	(14)	3	(8)	1	(5)	14	(11)
40	77	(17)	6	(11)	1	(5)	16	(11)
Forecast variance of n (percent) explained by shock to:								
1	6	(4)	94	(4)	0	(0)	0	(0)
4	26	(10)	63	(10)	11	(6)	0	(1)
12	41	(15)	43	(14)	13	(10)	2	(6)
20	49	(17)	36	(15)	12	(10)	3	(7)
28	53	(18)	33	(15)	10	(10)	3	(7)
40	54	(18)	32	(16)	10	(9)	3	(8)
Forecast variance of y (percent) explained by shock to:								
1	16	(6)	8	(4)	75	(6)	0	(0)
4	35	(10)	16	(8)	49	(10)	1	(1)
12	43	(13)	16	(9)	36	(11)	5	(6)
20	52	(14)	13	(9)	29	(11)	6	(7)
28	58	(15)	11	(9)	24	(11)	7	(8)
40	59	(15)	11	(10)	21	(11)	9	(9)
Forecast variance of g (percent) explained by shock to:								
1	0	(1)	2	(3)	0	(1)	98	(3)
4	0	(2)	16	(8)	0	(2)	84	(8)
12	20	(11)	34	(14)	0	(3)	47	(12)
20	43	(16)	26	(15)	0	(4)	31	(12)
28	55	(17)	18	(14)	0	(4)	26	(12)
40	59	(19)	17	(15)	0	(5)	24	(13)

As in Table IV.

variance in public capital is due to private capital. This result reinforces our conclusion from the impulse response functions that public capital responds to changes in private capital. In contrast to the impulse response function results, however, the variance decompositions indicate that private capital tends to be largely exogenous, with its own innovations explaining about 75 percent of its forecast error variance, even after 10 years. This argues against the conclusions that public investment can motivate private investment. The variance decompositions also indicate that shocks to private output account for almost none of the variation in public capital.

The second feature of Table IV is that it confirms the previous result that public capital shocks do not explain a quantitatively important amount of the variation in either private hours or private output. Rather it is shocks to private capital that explain most of the longer-term forecast error variance in private hours and output.[10]

10. For completeness, the variance decomposition results for the alternative ordering (k, n, y, g) are presented in Table V. These results do not substantially alter the above conclusions.

One final feature of the VAR results merits comment. As noted, the response of the private sector to a shock to public capital is largely confined to private capital and production; there is very little evidence of a significant response of private hours. Consequently, private labor productivity must rise because of the increase in public capital. One possible explanation for this is that a dominant component of public investment, in terms of contributing to increased private production, are those investments which improve the effective labor force. At this level of aggregation, there is simply insufficient information to extend this interpretation any further. However, it does suggest possible directions for further consideration. First, to explicitly model the role for public investment to improve the effective labor force. Arrow and Kurz [2] considers some of these issues. Second, to consider a disaggregated public capital stock in an attempt to quantify the contributions of various components of public investment to private production. This is pursued to some extent in Easterly and Rebelo [12] using cross-country data as well as in Finn [15].

V. Conclusions

Aschauer's [4] study of private production and the role of public capital has received considerable attention and criticism. The analysis of the Australian economy presented here, by focusing on both the long-run and short-run components of the data, provides further evidence of the nature of private production and its dependence upon publicly provided capital.

The cointegration analysis points to an estimated elasticity of public capital of approximately 0.17, one half of that of private capital. This result, similar to Lynde and Richmond [21], is about one half of our own previous estimates and seems to be a much more plausible number. By implication our estimate of the marginal (private) product of public capital does not imply implausibly large returns to public capital. Care must be taken when interpreting these numbers, however, as there seems to be a fair degree of uncertainty about the parameter estimates.

The short-run analysis provides some very useful additional information about private production. First, we are able to provide some evidence against the claim that the long-run relationship arises solely because public capital is endogenous, directly related to private production. We find no evidence of causality from private production to public capital stocks. The short-run analysis also confirms that the two capital stock are highly complementary with the strongest evidence indicating that public investment is highly responsive to private investment. Somewhat weaker evidence suggests that private investment is responsive to public investment, a result consistent with Aschauer [5].

Data Appendix

This appendix details the approach used to construct measures of private sector production and the inputs to private production, both privately and publicly owned. In Otto and Voss [25], we construct annual data for the private sector of the Australian economy by accumulating data for a number of sectors. This strategy is limited, however, because while production in some sectors of the economy (for example, mining, agriculture, manufacturing, wholesale and retail trade, and recreation and personal services) is predominantly done by the private sector, other sectors (for example, transport, storage and communication) encompass both public and private production and could not be included in our measure of private sector output; as a result, our previous measure of private sector output accounted for just under 50 percent of constant price GDP. In an effort to improve on this number we adopt a different approach to separating the aggregate economy into its public and private components.

All of the series constructed below are seasonally adjusted and cover the period 1959:3 to 1992:2.

Private and Public Sector Output

We follow Carmichael and Dews [8] in separating aggregate output into its private and public components. Total output is measured by the Australian Bureau of Statistics (ABS) income-based measure of GDP. To obtain that part of this measure which accrues to the private sector, we sum the following components, all of which are quarterly and in current prices. Private sector GDP equals private sector wages, salaries and supplements plus gross operating surplus of companies and unincorporated enterprises plus dwellings owned by persons and private financial enterprises less private sector imputed bank charges plus net indirect taxes paid by the private sector. This series is then converted to constant 1989/90 prices using the implicit price deflator for the ABS income-based measure of GDP.

Private and Public Hours Input

The measure of aggregate labor input used is total hours of the private sector. The ABS does not publish this data so we construct it using published data on total and government employment and average hours worked.

Total employment (excluding defence) is available on a quarterly basis from the Australian NIF Database. Government employment can be obtained on an annual basis from 1959/60 to 1991/92. To create a quarterly series from this annual series, we assume that government employment grows smoothly within each year. Private employment is then simply total employment less government employment. To obtain an estimate of total hours of the private sector per quarter we use the average weekly hours series from the NIF Database. This is available quarterly from 1966:3 to 1992:2. For the period 1959:3 to 1966:2, we set average weekly hours worked equal to its value in 1966:3. Private sector total hours per quarter are calculated as private employment times average weekly hours times twelve.

Private and Public Capital

The ABS publishes annual constant price estimates of the net capital stocks for both private and public sectors for the years 1966/67 to 1991/92. We measure both public and private capital as the sum of non-dwelling construction and equipment. We construct three quarterly series: the private capital stock, the general government capital stock and the capital stock of public trading enterprises.

We begin by constructing an annual net capital stock for each of these definitions for the period 1959/60 to 1991/92. Taking the ABS capital stock estimates from 1966/67 to 1991/92, we need estimates for the years 1959/60 to 1965/66. These figures are obtained by iterating backwards using the capital accumulation equation:

$$K_t = (1 - \delta)K_{t-1} + I_t$$

where K_t is the net capital stock (at end-June), I_t is gross fixed capital expenditure and δ is the rate of depreciation. An estimate of the annual depreciation rate for the years 1966/67 to 1991/92 can be obtained using a method suggested by Carmichael and Dews [8]. Let D be annual capital consumption (depreciation); then

$$\delta = D_t/K_t.$$

Using ABS data for capital consumption and net capital stocks we obtain an annual estimate of the depreciation rate. Setting δ equal to the estimate for the year 1966/67 and using data for gross fixed capital expenditure allows us to use the accumulation equation to compute annual capital stock estimates for the years 1959/60 to 1965/66.

We now have annual capital stock estimates for the period 1959/60 to 1991/92, which we wish to make quarterly; more specifically, suppose we have capital stock estimates for June 1991 and June 1992 and we wish to interpolate the interceding quarters. We do this in the following manner. We take the change in the yearly net capital stock,

$$I_t^n = K_t - K_{t-1}.$$

which gives annual net investment and divide this by annual gross fixed capital expenditure. This gives a fraction, which can be negative and is for the year 1991/92. Assuming this fraction is constant for each quarter of the year and using the resulting quarterly series to multiply the quarterly series for gross fixed capital expenditure provides a quarterly estimate for net investment. Since

$$K_t = K_{t-1} + I_t^n,$$

we can construct quarterly estimates for the net capital stock. One advantage of this procedure over a simpler strategy of straightforward accumulation is that June stocks for our quarterly series from 1966/67 onwards are consistent with the published ABS data.

The data used in this study are available on request.

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