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Public Capital Stock and Interstate Variations in Manufacturing Efficiency

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Abstract

This study focuses on the role of public capital stock in contributing to interstate differences in productive efficiency in manufacturing. Our motivation is to assess the role of public capital as a source of persistent regional variations in efficiency, and thus determine if public infrastructure policies might reasonably be expected to alter the competitive environment of a state's manufacturing sector. We use a stochastic frontier production function model, explicitly incorporating infrastructure capital, to examine the relative performance of the aggregate manufacturing sector across states and over time. We calculate an index of productive efficiency and estimate the determinants of statewide variations in it. These results show that variations in per capita infrastructure stocks significantly affect manufacturing efficiency. Considering the behavior of public capital as both a direct and an indirect input furthers our understanding of the role of infrastructure and its implications for regional economic development policies.

INTRODUCTION

There continues to be considerable uncertainty and debate about the role of infrastructure investment in stimulating productivity and economic growth. Several researchers have investigated the causal link between public capital formation and measures of economic performance in regions, but the existing literature is decidedly mixed concerning infrastructure's importance in generating gains in productivity. This lack of consensus is particularly unsettling to policymakers searching for approaches to enliven their economies in view of persistent interstate and regional differentials in income levels, productivity, and growth rates. The research reported here provides detailed evidence on the relationship between infrastructure capital and productive efficiency within state manufacturing sectors. Estimates of a stochastic frontier production function model are used to calculate two indices of manufacturing efficiency. One estimate explicitly considers the stock of public capital as an

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"unpaid" input in the underlying technology and the other does not. Public capital stock per capita has a positive effect on both indices. Thus, it contributes directly to manufacturing output and indirectly by increasing the efficiency of state manufacturing production. This approach enables a better understanding of the linkage between publicly provided inputs and the nature of the manufacturing production process. Knowing these linkages is imperative if public policy formulation is to move beyond broad generalizations.

The next section of this article provides some important background and explains our motivation in the process of discussing the most pertinent literature. This is followed by a methodology section which also contains a brief description of data sources. A discussion of the statistical results and major findings is then presented. A final section summarizes our research and draws some conclusions with policy implications.

BACKGROUND AND PROBLEM DEFINITION

The nature of spatial economic restructuring in the United States is not adequately understood, particularly as it pertains to manufacturing activities. Recently, much attention has been focused on the relative efficiency of manufacturing activities across regions, states, and urbanized areas. For example, Beeson and Husted [1989] find substantial and persistent variation in manufacturing efficiency across states which might help explain shifting patterns of jobs and economic activities. Crandall [1993] argues that the patterns of manufacturing migration evident over the past few decades are likely to continue because dynamic differentials such as regional wage rates have not been eliminated. Although there have been attempts to explain interstate differences on the basis of industrial structure, labor force, and other characteristics in order to guide remedial action, the effectiveness of policy intitiatives in this area remains an unsettled issue. Specifically, the suggested link between public capital and private sector productivity implies a need to explore how, or if, the performance of the manufacturing sector responds to infrastructure investment.

The existing estimates of state-level manufacturing efficiency ignore the potential impact of infrastructure stocks on output and, in turn, relative productivity performance. The present research seeks to fill this void so that we may better assess the impact of specific policy levers in improving manufacturing efficiency and regional economic welfare. Also, a different perspective on the broader relationship between infrastructure and productivity is provided by focusing on a specific economic sector whose production technology can be more realistically modeled within a production function framework.

Existing estimates of the productivity of public capital are sensitive to methodology, data sources, and econometric specification. Regional- or state-level analyses have yielded novel insights and been of particular interest to policymakers. Yet there are wide disparaties in results regarding the magnitude and significance of the estimated output elasticites. Munnell [1990], Garcia-Mila and McGuire [1992], Moomaw, Mullen, and Williams [1995], and others find empirical support for the hypothesis that public infrastructure enhances regional economic performance. Moomaw and Williams [1991] corroborate this effect as it relates to total factor productivity growth within states' manufacturing sectors. Other studies, however, are either inconclusive or not generally supportive of a prominent role for infrastructure. For example, Hulten and Schwab [1991], using growth-accounting techniques, conclude that infrastructure investment has had little impact on regional economic growth. Holtz-Eakin [1994] and Holtz-Eakin and Schwartz [1994] argue that the measured positive effect of infrastructure disappears once state-specific effects are accounted for, although the interaction between these effects and public capital formation is largely ignored. Evans and Karras [1994] also exploit a panel of states in addressing the inherent endogeneity problem exhibited by the production function approach in "explaining" aggregate output. They find no evidence that government capital is productive, in contrast to their findings for government-provided educational services.

Studies based on time-series data at the national level display similarly checkered results. Output elasticities reported by these studies tend to be larger than those based on regional analyses, perhaps owing to the difficulties of capturing all the benefits of infrastructure investment at this level of disaggregation. Moreover, the methodology and econometric techniques employed in these studies have been attacked as flawed [e.g., see Jorgenson, 1991]. Although unsettling, the disparate results and lack of consensus within the literature can be attributed partially to differences in the focus of the analysis and/or scope of activities under investigation. Greater insight is possible by narrowing the focus to specific sectors and to specific activities rather than concentrating on overall economic output.

The role of public infrastructure in explaining interstate variations in manufacturing efficiency has yet to be adequately addressed. Nadiri and Mamuneas [1994] recently report results suggesting that public capital significantly affects the cost structure and productivity of the manufacturing sector, based on a time-series analysis of two-digit industries on the national level. However, they find the magnitudes of these effects to be much smaller than generally suggested in the literature; furthermore, the favorable impacts are shown to vary dramatically by industry. This suggests that there ought to be differential *regional* impacts from infrastructure investment even if they arise solely from spatial variations in industrial mix. It is important to determine whether or not persistent regional differences in manufacturing efficiency can be influenced by public infrastructure stocks. Such evidence is of practical importance to policymakers for several reasons. States compete among themselves for private investment, and their ability to provide an attractive manufacturing environment may be linked to their public capital budgets. Also, the financial crisis facing many state governments has made the choice between maintenance of existing facilities and additional capacity via new infrastructure an even more important one. These policy issues justify more focused research initiatives such as the one presented herein.

METHODOLOGY AND DATA

The debate over the effectiveness of public capital in stimulating output or productivity growth has been centered around measuring the degree to which infrastructure matters to the *average* industry or region. Yet it is likely that public capital displays a wide array of impacts depending on, for example, the nature of the industry or the social/economic characteristics of the region.

Certain industries may simply not require "external" inputs of this type, while others may be strongly dependent on existing infrastructure in attempts to remain competitive with other producers. Similarly, some regions may benefit enormously from a relatively small infusion of public capital, while others may have largely exhausted those gains which come from additional "economic overhead capital." Thus it would be useful for policy purposes to know how public capital may affect *relative* economic performance across various industries or regions. Our investigation uses a frontier production function to examine how public capital differentially affects efficiency within state manufacturing sectors.

The estimated production function defines the maximum output that can be produced by a given set of inputs. The use of a frontier production function allows the development of a measure of inefficiency. This measure calculates how far any one observation's output falls below the maximum (as defined by the frontier). The unit that produces on the frontier is deemed to be 100 percent efficient, while those below the frontier display varying degrees of inefficiency.[1] The production function to be estimated is:

$$\ln Q = \ln X^j + v - u \tag{1}$$

where Q is output, X^j is a vector of j inputs, and v and u are a two-part disturbance term. The first part of the disturbance term, v, is the usual two-sided stochastic error term, reflecting random events and statistical noise. The second part, u, represents the technical inefficiency of the particular unit in production. The value of u is constrained to be greater than or equal to zero, so that output lies below (or on) the frontier. In general, estimation of such a model is problematic owing to "inappropriate" assumptions concernign the "u" component of the error term. Panel data, however, can be exploited to yield estimates of technical inefficiency for cross-sectional units without making distributional assumptions. Further, this approach can be adopted without invoking the assumption that inefficiency is independent of the X^j. In practice, a number of estimators may be employed depending on the validity of the assumptions concerning the nature of technical inefficiency for those observations which constitute a specific application. For example, if it is appropriate to treat the u_i as fixed over time, the "within" or least squares dummy variable (LSDV) estimator can be used; it estimates a different intercept for each cross-sectional unit. An alternative to this fixed effects model is to treat the u_i as random and uncorrelated with the regressors, whereupon GLS becomes the most likely choice of an estimator [Mundlak, 1978].

The LSDV estimator is most appropriate here because it allows us to measure the "fixed" technical inefficiency of each cross-sectional unit (state) with a minimum of restrictive assumptions. More specifically, we need not assume that inefficiency is independent of the choice of inputs themselves (as regressors in the production function model) because it is simply captured via a separate intercept term for each state. Thus we proceed initially by employing an F-test to consider, and ultimately reject, the hypothesis that the constant

[1] The following discussion draws heavily upon Schmidt and Sicles' [1984] technique for estimating these production functions with panel data. Also, the study by Beeson and Husted [1989] is most helpful.

terms are all equal.[2] These results prompt us to continue by estimating the following stochastic production function model:

$$\ln Q_{it} = \alpha + \sum^{j} \beta^{j} X^{j}_{it} + v_{it} - u_i (i = 1, \ldots, N; t = 1, \ldots, T) \quad (2)$$

We assume that the u_i are fixed over time for each state. This allows the u_i term to be contained within the intercept as $\alpha_i = \alpha - u_i$. The measure of fixed technical inefficiency for each state is derived by exploiting estimates of the u_i which are defined as $\alpha - \alpha_i$ where $\alpha = \max(\alpha_i)$. The index of efficiency (IE) is then computed as:

$$\mathrm{IE} = 100e^{-u_i} \quad (3)$$

The state producing on the frontier is considered 100 percent efficient, with each other state's efficiency index being directly calculable from (3).

It is possible to extend the LSDV approach to capture a time-specific fixed effect as well. This type of extension is unwarranted by the the focus of our model which is not concerned with the timewise performance of the error term. We explicitly include a time trend, however, as a proxy for technical change.

Two principal data sources are drawn on in estimating the stochastic production function. The first of these is the U.S. *Census of Manufactures* (COM) which reports data on value added (as a measure of net output) and production labor-hours for the manufacturing sector of each state for the 1972, 1977, 1982, and 1987 years.[3] The second major source provides annual data from 1969–1988 on the public infrastructure stock and manufacturing capital stock for each state. These data were produced at the Federal Reserve Bank, Boston, and have been used extensively by Munnell [1990] and others concerned with a variety of issues surrounding public infrastructure capital. The public capital stock series was created by utilizing annual state public investment data (provided by the Census Bureau's *Governmental Finances*) together with discard/depreciation schedules to apportion public capital totals for the nation (as published by the Bureau of Economic Analysis). The resulting series is notable in that it details *total* public capital stocks for each state, but also provides a breakdown classified as water and sewer, highways, and "other" infrastructure stocks. The private capital stocks in manufacturing for each state were similarly created by apportioning national totals (BEA) on the basis of measures of each state's manufacturing activity.[4]

We make specific use of the private and *total* public capital stocks to represent, together with labor-hours, the primary production function inputs. For estimation purposes, our panel set of data consists of 48 cross-sectional units (Alaska and Hawaii are excluded) and four separate time periods (1972, 1977,

[2] Such an F-test is described in detail in Greene [1990, pp. 484–485]. The resulting F-value, constructed from the R^2 values of the unrestricted and pooled (single intercept) models, is $F(n - 1, nT - n - k) = F(47, 130) = 226.81$. This figure vastly exceeds the critical F-value so that we can strongly reject the hypothesis that the state effects are identical.

[3] The *Annual Survey of Manufactures* does report (less reliable) information on the relevant variables for all of the intervening (non-Census) years with the exception of the 1979–1981 period. Our analysis is confined to the 5-year Census periods.

[4] See Appendix A in Munnell [1990] for greater detail on how the state-by-state estimates of public and private capital stocks were created.

1982, 1987).[5] All dollar figures are converted to 1982 values using the GNP implicit price deflator [*Survey of Current Business, 1989*]. The estimation of the production function model with "fixed effects" is followed by the construction of an index of efficiency for each state. Our final task involves explaining the variations in this efficiency index. Variable definitions, data sources, and summary statistics for this additional empirical exercise are reported in the Appendix.

ESTIMATION AND RELATED FINDINGS

An application of the stochastic frontier technique described above requires an explicit specification of the production function. The translog specification is perhaps the most flexible as it allows for nonconstant returns to scale, does not restrict the elasticity of substitution between inputs, and allows for both neutral and factor-augmenting technical change. Thus we rely on the translog production function to estimate output frontiers. This approach assumes that the translog is broadly representative of the technology underlying production within the manufacturing sector of each state. For comparative purposes, we estimate these stochastic production frontiers both with and without the explicit inclusion of public capital stock. The basic equation (with public capital included) is presented in (4) below.

$$\begin{aligned}\ln Y_{it} = \alpha &+ \beta_L \ln L_{it} + \beta_K \ln K_{it} + \beta_G \ln G_{it} + \beta_T T \\ &+ \beta_{LL}(0.5\ln L_{it}^2) + \beta_{KK}(0.5\ln K_{it}^2) + \beta_{GG}(0.5\ln G_{it}^2) + \beta_{TT}(.5T^2) \\ &+ \beta_{LK}(\ln L_{it}\ln K_{it}) + \beta_{LG}(\ln L_{it}\ln G_{it}) + \beta_{KG}(\ln K_{it}\ln G_{it}) \\ &+ \beta_{LT}(\ln L_{it})T + \beta_{KT}(\ln K_{it})T + \beta_{GT}(\ln G_{it})T - u_i + v_{it} \qquad (4)\end{aligned}$$

where Y_{it} = value added in manufacturing within state i at time t, L_{it} = labor input (in manufacturing) in state i at time t, K_{it} = private capital (in manufacturing) in state i at time t, G_{it} = public capital stock in state i at time t, T = time trend, u_i = state-specific technical inefficiency, v_{it} = random disturbance term, and α, β_{ij} = parameters to be estimated. The error term, v_{it}, is assumed to be normally distributed and independent of the u_i. The u_i are assumed to be correlated with the regressors and therefore are representative of fixed effects across the states. Dummy variables are used for each state, so equation (4) is estimated without an intercept term. The results of the two estimates are presented in Table 1.

The estimates of individual coefficients for the translog are not easily interpretable but can be exploited to provide a number of insights. For example, the estimated β_{KG} is negative, suggesting that private and public capital are substitutes in production. Nadiri and Mamuneas [1994] report a similar finding in their analysis of U.S. manufacturing industries, although other evidence on this relationship is mixed. A substitutional relationship here

[5] Ideally, a larger number of time periods (T) would allow us to consistently estimate the individual intercepts and compare relative efficiencies across states (N). However, our case is similar to the usual panel data literature; as such, the (nonintercept) coefficient estimates via the "within" approach do remain consistent for small T and large N.

Table 1. Parameter estimates for translog production function.[a]

Variable	Without public capital	With public capital
$\ln L$	−2.0658	−1.8562
	(1.47)	(1.25)
$\ln K$	2.5400[b]	3.0612[b]
	(1.97)	(2.06)
$\ln G$	—	−2.1080[b]
		(1.67)
T	−0.5968[b]	−0.4329[b]
	(2.72)	(1.83)
$\ln L \ln K$	−0.3325[b]	−0.3086
	(1.77)	(1.63)
$\ln L \ln G$	—	0.1539
		(0.93)
$\ln K \ln G$	—	−0.1169
		(0.87)
$T \ln L$	0.0819	0.0633[b]
	(2.44)	(1.78)
$T \ln K$	−0.0725[b]	−0.0492
	(2.06)	(1.32)
$T \ln G$	—	−0.0086
		(0.32)
$0.5 \ln L^2$	0.3735[b]	0.2494
	(1.91)	(1.21)
$0.5 \ln K^2$	0.3306[b]	0.3510[b]
	(1.75)	(1.77)
$0.5 \ln G^2$	—	0.0893
		(0.54)
$0.5 T^2$	0.0342[b]	0.0323[b]
	(2.74)	(2.46)
$\overline{R}^2$	0.999	0.999
ε_L	0.792	0.818
ε_K	0.298	0.272
ε_G	—	0.006
ε_T	0.075	0.076
Returns to scale	1.09	1.096

[a] *t*-values in parentheses.

[b] Indicates coefficient is significantly different from zero at the 10 percent level;—indicates that particular variable does not appear in the equation.

implies that an increase in infrastructure lowers the marginal productivity of private capital inputs. In a similar fashion, the estimated β_{LK} implies substitutability between labor and private capital. Interestingly, the estimated β_{LG} implies that labor and public capital are complementary inputs. Although the empirical evidence concerning this latter relationship also is inconclusive, a number of researchers report similar findings [see Berndt and Hansson, 1991; and Shah, 1992]. This result is consistent with the idea that public capital may have an indirect effect on the productivity of private inputs [Nadiri, 1993]. In particular, public capital formation stimulates the demand for labor inputs and lessens the demand for private capital. Taxpayers are sometimes reluctant to support public financing of infrastructure projects. Evidence that such projects increase the demand for labor may make them more acceptable.

Next, we calculate the output elasticities (at mean values of the variables); these are presented at the bottom of the table. The overall degree of returns to scale may be derived by summing the output elasticities for the specific inputs. Returns to scale are estimated to be 1.09 in Regime I (excluding public capital) and slightly higher (1.096) when G is directly included as an input (Regime II).[6] These output elasticities are close to those reported by Beeson and Husted. We are especially interested in the output elasticity of public capital. This value is estimated to be very small at 0.006. Such a small value for this parameter is not surprising when analyzing data at a disaggregated regional level. In fact, Nadiri [1993] points out that estimates based on state-level data generally indicate a much smaller infrastructure contribution to output than suggested by national level results. The generally smaller range for this elasticity measured on a regional level is partially attributable to the inability of the data to capture spillover or "system" effects. Still, the small output elasticity value here is not vastly beyond the low-end range of estimates in the literature. Moreover, because we focus on the manufacturing sector alone, the output elasticity value is expected to be smaller than it would be if we were to focus on all industries. If the benefits of infrastructure extend more broadly to other industries, it would be necessary to "gross up" the marginal benefits of public capital in some fashion to make them comparable to other studies. Although the output elasticity value suggests that public capital has only a minor role as a direct input, we examine its function as an indirect input later in the study.

We now consider the statewide fixed effects emerging from the regression results via the "within" estimator. As noted above, this procedure involves using dummy variables for each state to measure technical inefficiency for each cross-sectional observation. By exploiting the panel data methodology described in general terms in the previous section, we proceed to calculate an index of efficiency (IE) for each of our stochastic production functions (estimated under two separate regimes by either including or excluding public capital as a direct input). The results are presented in Tables 2 and 3.

Not surprisingly, the findings presented in Table 2 parallel those of Beeson and Husted for an earlier time period (1959–1973); correlations of our rankings of states (in descending order of manufacturing efficiency) with their rankings are positive and significant.[7] This finding simply suggests that the pattern of differences in manufacturing efficiency across states has persisted over time. Some small differences do emerge in the ranking of individual states however. For example, Utah and Wyoming have seen their IE rankings slip from 2nd and 4th to 13th and 28th place, respectively. In the other direction, Arizona and Iowa leap from 8th and 10th to 1st and 2nd respectively. Although there have been significant changes in the composition of individual state's manufacturing sectors that can account for these differences, other state-specific influences may also have been operative. Our purpose here is not so much in elucidating those dynamic forces, but rather in

[6] We do not impose (or test the validity of) restrictions which often are placed on the value of these parameters in conforming to the assumption of linear homogeneity.

[7] The Spearman rank correlation coefficient of 0.85 is highly significant as is the Pearson correlation coefficient (0.73).

Table 2. State rankings of manufacturing efficiency.

State	Public capital excluded		Index of efficiency (% terms)
	Region[a]	Parameter estimate	
1. Arizona	MTN	7.0814	100.0
2. Iowa	WNC	7.0534	97.2
3. Colorado	MTN	7.0514	97.0
4. Minnesota	WNC	7.0472	96.6
5. Nebraska	WNC	7.0438	96.3
6. Kansas	WNC	7.0410	96.0
7. Nevada	MTN	7.0187	93.9
8. Delaware	SA	7.0106	93.2
9. North Dakota	WNC	7.0050	92.6
10. South Dakota	WNC	6.9929	91.5
11. Missouri	WNC	6.9737	89.8
12. Kentucky	ESC	6.9618	88.7
13. Utah	MTN	6.9614	88.7
14. Maryland	SA	6.9309	86.0
15. Connecticut	NE	6.9286	85.8
16. Washington	PAC	6.9283	85.8
17. New Jersey	MA	6.9257	85.6
18. Vermont	NE	6.9152	84.7
19. Oklahoma	WSC	6.9094	84.2
20. New Hampshire	NE	6.9067	84.0
21. Idaho	MTN	6.9014	83.5
22. Massachusetts	NE	6.8967	83.1
23. New York	MA	6.8736	81.2
24. Wisconsin	ENC	6.8718	81.1
25. Oregon	PAC	6.8632	80.4
26. California	PAC	6.8374	78.3
27. New Mexico	MTN	6.8305	77.8
28. Wyoming	MTN	6.8281	77.6
29. Illinois	ENC	6.7982	75.3
30. Florida	SA	6.7952	75.1
31. West Virginia	SA	6.7748	73.6
32. Michigan	ENC	6.7548	72.1
33. Virginia	SA	6.7532	72.0
34. Ohio	ENC	6.7475	71.6
35. Rhode Island	NE	6.7428	71.3
36. Louisiana	WSC	6.7231	69.9
37. Indiana	ENC	6.7092	68.9
38. Texas	WSC	6.7012	68.4
39. Montana	MTN	6.6981	68.2
40. Arkansas	WSC	6.6737	66.5
41. Tennessee	ESC	6.6583	65.5
42. Georgia	ESC	6.6488	64.9
43. Mississippi	ESC	6.6156	62.8
44. Pennsylvania	MA	6.6116	62.5
45. Maine	NE	6.5720	60.1
46. Alabama	ESC	6.5242	57.3
47. North Carolina	SA	6.5140	56.7
48. South Carolina	SA	6.4578	53.6

[a] The United States in broken into nine regions conforming to a common Census Bureau configuration: mountain (MTN); west north central (WNC); South Atlantic (SA); east south central (ESC); northeast (NE); Pacific coast (PAC); middle Atlantic (MA); east north central (ENC); west south central (WSC).

Table 3. State rankings of manufacturing efficiency.

State	Public capital included: Region[a]	Public capital included: Parameter estimate	Index of efficiency (% terms)
1. Arizona	MTN	13.1936	100.0
2. Colorado	MTN	13.1614	96.8
3. Iowa	WNC	13.1606	96.8
4. Nebraska	WNC	13.1540	96.1
5. Nevada	MTN	13.1461	95.4
6. Kansas	WNC	13.1433	95.1
7. North Dakota	WNC	13.1321	94.0
8. Minnesota	WNC	13.1207	93.0
9. South Dakota	WNC	13.0811	89.4
10. Kentucky	ESC	13.0657	88.0
11. Missouri	WNC	13.0652	88.0
12. Utah	MTN	13.0595	87.5
13. Connecticut	NE	13.0366	85.5
14. Washington	PAC	13.0344	85.3
15. Maryland	SA	13.0330	85.2
16. Delaware	SA	13.0161	83.7
17. Oklahoma	WSC	13.0132	83.5
18. Wyoming	MTN	13.0107	83.3
19. New Jersey	MA	13.0100	83.2
20. Massachusetts	NE	12.9717	80.1
21. Oregon	PAC	12.9697	79.9
22. Wisconsin	ENC	12.9646	79.5
23. New Hampshire	NE	12.9595	79.1
24. New Mexico	MTN	12.9539	78.7
25. Idaho	MTN	12.9231	76.3
26. Louisiana	WSC	12.8818	73.2
27. West Virginia	SA	12.8707	72.4
28. Vermont	NE	12.8613	71.7
29. Virginia	SA	12.8395	70.2
30. Indiana	ENC	12.8307	70.0
31. Rhode Island	NE	12.8292	69.6
32. Florida	SA	12.8250	69.2
33. Illinois	ENC	12.8162	68.6
34. Montana	MTN	12.8085	68.0
35. Michigan	ENC	12.8083	68.0
36. Ohio	ENC	12.7954	67.2
37. Arkansas	WSC	12.7818	66.2
38. Tennessee	ESC	12.7541	64.4
39. Texas	WSC	12.7513	64.3
40. Georgia	ESC	12.7409	63.6
41. Mississippi	ESC	12.7226	62.4
42. New York	MA	12.6917	60.5
43. California	PAC	12.6543	58.3
44. North Carolina	SA	12.6518	58.2
45. Alabama	ESC	12.6358	57.2
46. Pennsylvania	MA	12.6353	57.2
47. Maine	NE	12.5935	54.9
48. South Carolina	SA	12.5861	54.5

[a] See footnote in Table 2.

isolating the role of public capital in these IE rankings more generally. As such, we are more interested in comparing the Table 2 and Table 3 results.

Tables 2 and 3 both show that the most efficient states are generally located in the mountain (MTN) and west north central (WNC) states. Alternatively, the states in the South—east south central (ESC) and west south central (WSC)—are consistently found to be near the bottom of the efficiency rankings. Overall, there is a wide disparity between the most and least efficient state regardless of index chosen. For example, South Carolina is measured as the least efficient state in both sets of results, weighing in with IEs of 54.5 and 53.6 percent. Moreover, a cursory examination of these results suggests a fairly uniform distribution of efficiency across states in the sample. Although the efficiency rankings that emerge from these two sets of results are similar (Spearman rank correlation coefficient is 0.93 and significant), they still allow us to discern the impact of ignoring the contributions of public capital on relative performance. Most notably, New York and California both drop substantially in their respective IE rankings once public capital is directly considered. Specifically, New York and California fall from 23rd and 26th to 42nd and 43rd, respectively. These two states, both of which are public capital intensive, appear to be much more efficient when public capital inputs are ignored. The implication here is that infrastructure is the source of their loftier positions in the efficiency rankings of Table 2. New Jersey, Massachusetts, Ohio, Illinois, and Pennsylvania all drop slightly in the efficiency rankings once public capital is explicitly considered within the production function. Although changes in ranking for other states are less dramatic, these findings suggest an important role for infrastructure in influencing productive efficiency. Admittedly, this phenomenon may be more pronounced in select states, perhaps those with traditionally large manufacturing sectors or those with a certain industrial mix.

The above findings depict a pattern of substantial variation in manufacturing efficiency across states that persists over time, despite the forces of convergence. Also, our results hint at an important role for public capital in generating estimates of efficiency performance within manufacturing. To provide stronger empirical evidence about the sources of regional variations in efficiency we regress the efficiency index (IE) on a number of potential explanatory variables. Besides ad hoc and theoretical considerations, this exercise draws upon the findings of a similar analysis conducted by Beeson and Husted. Refinements of certain variables and improvements in measuring specific influences yield additional insights. A rationale for the inclusion of specific variables is presented below.

Given our preoccupation with infrastructure, we include a variable which measures per capita public capital (PUBCAP) within each state. This variable should positively impact manufacturing efficiency if public capital stock acts as a direct but unpaid factor, or if it performs as an indirect input that somehow enhances the productivity of traditional factors. Spatial variations in manufacturing efficiency also are posited to be influenced by labor-force characteristics. Those states with better trained and educated workforces are likely to display greater productivity and/or efficiency. We hypothesize that IE should be influenced by average levels of educational attainment as proxied by a variable which measures the percentage of the state's population with at least a high school education (HSGRAD). Similarly, the labor relations atmosphere within the state may contribute to measured differences in IE.

Table 4. Regressions explaining variations in manufacturing efficiency indices.[a] Dependent variable = index of efficiency (IE) in manufacturing.

Explanatory variables	Regime I (without public capital)	Regime II (with public capital)
Constant	0.0377	0.1069
	(0.29)	(0.72)
UNION	−0.1573	−0.3013[b]
	(1.27)	(2.16)
HSGRAD	0.6135[b]	0.4770[b]
	(2.50)	(1.73)
DUR	0.2240[b]	0.2299
	(1.81)	(1.65)
EMPCH	0.1834	0.1208
	(1.16)	(0.68)
DENS	0.000067	0.000043
	(1.00)	(0.57)
PUBCAP	0.0421[b]	0.0462[b]
	(3.14)	(3.06)
$\bar{R}^2$	0.468	0.398
F-value	7.60[b]	5.96[b]

[a] Numbers in parentheses are t-ratios.
[b] Statistically significant at the 10 percent level.

We include a variable (UNION) which measures the unionization rate in manufacturing to test this effect.

A state's population density (DENS) is included as an explanatory variable in order to capture the strength of aggolmeration economies. Presumably, those states with higher density should have more pronounced agglomeration effects and greater manufacturing efficiency. Alternatively, we experiment with the percentage of the state's population living in metropolitan areas (METRO), but do not include both variables in the same equation. Ideally, these regressor variables would either reflect "average" levels over the period of analysis of values for a midpoint year. The Appendix provides summary statistics and complete descriptions for these variables.

A final group of explanatory variables attempts to account for the influence of industrial structure on efficiency performance. Differences in industrial composition and associated variations in the underlying technology of production across states could result in misleading estimates of technical inefficiency. We include a variable which captures the relative importance of durable goods industries within the manufacturing sector of each state over the sample period. DUR is defined as the (average) percentage of total output accounted for by the durable goods industries (SIC codes 24, 25, 32–39) for the Census (COM) years. Additionally, we include a variable which measures the percentage change in employment (EMPCH) within the manufacturing sector over the 1982–1987 period. This variable (readily available from the COM) reflects the overall presence of employment in growing versus declining manufacturing industries within the state. States with a stronger presence of growth industries can be expected to display superior efficiency performance.

The results of this empirical exercise are in Table 4. Separate regressions are reported for the alternate measures of IE (exclusive and inclusive of public capital as a direct input in the translog production function). In general, the results are plausible and consistent with initial expectations, and the explanatory power is quite strong. The entire relationship is statistically significant, and the individual variables behave consistently across both regimes. The performance of specific variables in these equations warrants further discussion.

The public infrastructure stock variable (PUBCAP) has a positive and strongly significant coefficient under both regimes. Higher levels of (per capita) public capital have a strong impact on efficiency performance. This variable performs more consistently than other regressors and attains the highest degree of statistical significance. It is noteworthy that the importance of public capital to the productive efficiency of the manufacturing sector is established regardless of how the frontier is measured. Apart from any measurable effect that public capital has in the production process as a direct input, these results suggest that it has an indirect impact on the efficiency of private sector inputs. The idea that public capital may enter the manufacturing production process indirectly is not unique, but evidence pertaining to this effect has been sparse. The results here are consistent with our finding, noted above, that infrastructure and labor inputs behave in a complementary fashion. A further implication is that simply treating infrastructure as a direct input is insufficient to fully capture its broader impacts within the production process, at least for the manufacturing sector.

The influence of educational attainment on efficiency is as hypothesized. The coefficient of HSGRAD is positive and significant in both equations. The agglomeration variables have a positive but insignificant influence on the index of manufacturing efficiency. We report results solely for the equations which include DENS, but an alternate format with METRO as a regressor yields very similar findings. The unionization variable has a negative coefficient (statistically insignificant in one case however), suggesting that more highly unionized work forces are detrimental to manufacturing efficiency performance. Such a finding contradicts that of Beeson and Husted, although this descrepancy might be traced to alternative ways in which this variable is measured. We rely on Census figures, [U.S. Bureau of the Census, 1991], which report unionization rates among nonagricultural production workers, whereas their findings are based upon *estimates* of private sector unionism from a different source. Substantial differences between these two measures across individual states are evident, perhaps accounting for the incongruous results.

The final variables in the empirical model are those intended to measure industry mix. The relative size of the durable goods industries (as reflected by DUR) has a positive and statistically significant impact on efficiency. Differences in production technologies between these and the nondurable goods industries presumedly account for the generally stronger efficiency performance where the former industries are more dominant. This outcome also conflicts with earlier findings (Beeson and Husted). The alternate variable intended to reflect industrial composition (EMPCH) fails to attain statistical significance, although it has the expected positive sign. This variable may be capturing overall regional economic growth trends as much as it is reflects industrial composition.

The above analysis and results confirm the hypothesized effects of certain causal forces, although unsettled issues remain and await further extensions. For example, alternative measures of industrial mix might be developed to better capture differences in underlying production technologies. On a broader level, it would be interesting to compare the efficiency performance of the manufacturing sector with that of the rest of the private business economy within each state. Specifically, this latter investigation might focus on the relative importance of public infrastructure on efficiency performance across these sectors.

SUMMARY AND IMPLICATIONS

This research has focused on measuring the impact of public capital stock on technical inefficiency within states' manufacturing sectors. A translog specification of a stochastic frontier production function is used to generate an index of manufacturing efficiency for each state. We exploit a panel data approach which facilitates econometric estimation of the fixed inefficiency of each cross-sectional observation. Doing so allows us to conclude that, in general, states within the west north central and mountain regions are more efficient in producing manufacturing output than their regional counterparts. Our results broadly corroborate estimates generated for an earlier time period, but demonstrate that these findings are sensitive to the treatment of infrastructure stocks in developing the index of manufacturing efficiency. At a minimum, these results are suggestive of an important role for public capital stocks in productive efficiency within the manufacturing sectors of certain states. Overall, the output elasticity of public capital is measured as being positive but quite small. Its magnitude appears reasonable, however, in consideration of the disaggregate nature of the analysis and our focus solely on manufacturing output.

An additional exercise here attempts to explain variations in manufacturing efficiency across states on the basis of, among other determinants, public infrastructure stocks. Our regression results underscore the importance of public capital as having an influence on productive efficiency. Per capita infrastructure stocks are positively and significantly related to measures of manufacturing efficiency regardless of how the production frontier is estimated. Thus we provide evidence that the aggregate stock of public capital within a state helps to strengthen the competitiveness of its manufacturing sector. Further, our results imply that government capital also may operate as an indirect input or "environmental" factor which strengthens the productivity of traditional inputs. Evidence of this phenomenon is important as policymakers seek a better understanding of the role that infrastructure plays in regional growth. The failure to assign an important quantitative role to public capital in some previous studies may be reflective of their emphasis on the "direct input" approach. Our results indicate that it would be premature to conclude that infrastructure has little or no measurable impact on output or productivity performance within a regional context. On the contrary, the evidence presented here shows that public capital formation serves to "ratchet up" productive efficiency within the manufacturing sector.

Additional evidence suggests that labor force characteristics exert a significant influence on efficiency performance within the manufacturing sector. Specifically, a better educated workforce serves to increase manufacturing efficiency, but greater unionization tends to have the opposite effect. A strong presence of agglomeration economies is not obvious from our findings. The inconclusive results pertaining to the influence of industrial composition suggests that this area of inquiry be given greater consideration in future work. Undoubtedly, industrial structure influences efficiency measures based on (state-level) aggregate manufacturing sector data, assuming the nature of the underlying production technology is variable. More disaggregated analyses, such as those based on two-digit or three-digit industry classes, may be necessary to yield additional insights here.

A number or related issues and analyses await further investigation. For example, it would be interesting to examine the alternative impacts of specific categories of public capital in explaining variations in manufacturing performance. A need exists to explore the relative merits of spending on new infrastructure versus maintaining existing stocks in an effort to enhance the competitive environment for manufacturing. Another important issue concerns the role of infrastructure stocks in contributing to variations in the relative economic performance of the manufacturing versus other business sectors of states' economies. Such analyses should yield specific policy implications for guiding public sector investments aimed at improving economic performance within targeted sectors.

Appendix. Variable descriptions and summary statistics (sources of efficiency differences).

Variable name	Variable description	Mean	Standard deviation	Minimum	Maximum
UNION[a]	Percent unionization rate in manufacturing, 1987	19.1	12.1	3.1	51.9
HSGRAD[a]	Percent of population with high school minimum educational attainment, 1980	67.1	7.2	51.9	80.3
DUR	Percent of manufacturing-output in durable good industries, 1972–1987 average	50	13.9	11.8	78.4
EMPCH	Percentage change in manufacturing employment, 1982–1987	1.09	9.71	−23.4	22.9
DENS[a]	Population density (per sq. mile), 1980	159.3	225.2	4.8	986.2
PUBCAP	Per capita public capital stock, 1982	$5967	$1425	$4199	$10,534

[a] Supplementary data derived from *State and Metropolitan Area Data Book* [1982, 1991] and the 1980 Census of Population (general social and economic characteristics).

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