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Stephan, Andreas

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Andreas Stephan

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ABSTRACT

The Impact of Road Infrastructure on Productivity and Growth: Some Preliminary Results for the German Manufacturing Sector

by Andreas Stephan*

Using time-series cross-section data from the manufacturing sector of the 11 Bundesländer from 1970 to 1993, we examine the impact of road infrastructure on private production applying three different approaches; i.e., a Cobb-Douglas production function, a translog production function and a growth accounting approach. Our econometric analysis explicitly takes into account four of the most frequent problems in the context of time-series cross-section analysis: serial correlation, groupwise heteroscedasticity, cross-sectional correlation and nonstationarity of data. For all approaches and tested specifications, we find that road infrastructure is significant for production in the manufacturing sector. Moreover, we find that variations between the Bundesländer are more important for explaining infrastructure's contribution to production than variations across years.

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ZUSAMMENFASSUNG

Der Einfluß von Straßeninfrastruktur auf Wachstum und Produktivität: Einige vorläufige Ergebnisse für das produzierende Gewerbe in Deutschland

Das Arbeitspapier untersucht den Einfluß von Straßeninfrastruktur auf Produktivität und Wachstum im verarbeitenden Gewerbe in Deutschland für den Zeitraum 1970-1993. Den Hintergrund bildet die in der Literatur kontrovers diskutierte sogenannte Hypothese über das "Defizit an öffentlichem Kapital", wonach der seit Anfang der siebziger Jahren zu beobachtende allgemeine Rückgang des Produktivitätswachstum auf die in diesem Zeitraum ebenfalls zurückgegangenen öffentlichen Infrastrukturinvestitionen zurückgeführt werden kann. Die empirische Analyse wird mit einem Paneldatensatz auf Ebene der 11 westdeutschen Bundesländer durchgeführt. Zwei Wirkungsweisen von Infrastruktur auf die private Produktion werden unterschieden. Erstens können Firmen öffentliche Infrastruktur quasi wie ein kostenloses "Zwischenprodukt" für die Produktion nutzen. Zweitens kann Infrastruktur ähnlich wie technischer Fortschritt die Produktivität der privaten Produktionsfaktoren erhöhen. Im vorliegenden Papier werden beide Wirkungen von Infrastruktur erörtert. Im ersten Teil wird eine einfache Cobb-Douglas Produktionsfunktion geschätzt, in welcher Straßeninfrastruktur als ein Faktorinput in die Produktion eingeht. Es wird gezeigt, daß auf Basis der Cobb-Douglas Funktion die beiden Wirkungen von Infrastruktur empirisch nicht unterschieden werden können. Im zweiten Teil wird der Einfluß von Infrastruktur auf die Wachstumsrate der totalen Faktorproduktivität im verarbeitenden Gewerbe untersucht. Im dritten und abschließenden Teil wird mit einer Translog-Produktionsfunktion die Beziehung zwischen Infrastruktur und den privaten Inputfaktoren analysiert. Insgesamt zeigen die Ergebnisse eine starke Korrelation zwischen dem Produktivitätswachstum im verarbeitenden Gewerbe und der Straßeninfrastruktur. Allerdings ist diese Korrelation im Querschnittsvergleich der Bundesländer stärker ausgeprägt als innerhalb der einzelnen Zeitreihen. Die Hypothese über "das Defizit an öffentlichem Kapital" wird daher für das verarbeitende Gewerbe in Deutschland im Rahmen dieser Studie nicht bestätigt.

1 Introduction

During the 1970's and 80's, many OECD countries experienced a serious decline in output and productivity growth. Rising unemployment, increasing social transfers and public debt constrained public investment in many countries. Consequently, public investment as a share of GDP has declined considerably in most OECD countries during the last two decades.

{Insert Figures 1 and 2 about here}

Figure 1 shows, that the share of non-military public sector consumption increased in Germany from 10.5 to 18.5 percent during the period from 1960 to 1995, while Figure 2 shows, that the share of public sector's investment has declined from 3.3 to 2.0 percent from 1960 to 1997. After the German Reunification, public investment as a share of GDP increased for a short period from 2.5 to 3 percent, but afterwards has continued on a general decline since the mid 70's. This is even more surprising if one considers the still relatively high demand for infrastructure projects in the new Bundesländer.

Recently, a number of researchers such as Aschauer (1989a, 1989b, 1989c) or Munnell (1990a, 1990b, 1992) have documented empirical evidence for a strong correlation between public capital and private sector performance. Furthermore, these authors have hypothesized that the decrease of governmental investment in the US and other countries may be crucial for explaining the observed decline in productivity growth. This argumentation has been popularized as the "Public Infrastructure Hypothesis" in literature.

If the "Public Infrastructure Hypothesis" is of empirical relevance to Germany, then at least a part of the productivity gap between East and West Germany could also be attributed to the still existing differences in infrastructural endowments between east and west German regions. In fact, road infrastructure in the East German Bundesländer is still only two thirds of that in West German Bundesländer.

Turning to the hypothetical effects of infrastructure, Aschauer, for example, postulates that public capital can have both a direct and indirect effect on private output. The direct effect arises because changes in public capital stock alter the level of output by making private labor and capital inputs more or less productive. The indirect effect arises because an increase in public capital stock will affect the marginal products of labor and private capital, which in turn influence the chosen quantities of private inputs.

Spurred by the work of Aschauer, an increasing number of papers have examined the relationship between infrastructure and output growth.[1] While some of these studies (Garcia-Milà and McGuire 1992, Moonaw and Williams 1991, Carlino and Voith 1992) find similar positive and significant effects from infrastructure, others find only negligible or insignificant effects (Hulten and Schwab 1991, Tatom 1991a).

[1] For comprehensive literature surveys, see Pfähler, Hofmann, and Bönte (1997) or Sturm, Kuper, and de Haan (1996b).

In Germany, the priority of infrastructural projects is evaluated on the basis of cost-benefit studies (e.g., Bundesverkehrswegeplan 1992, Ministry of Transport). If investment in road infrastructure is efficiently allocated by governments on the basis of cost-benefit studies, then we would expect it to have a positive and significant impact on private production. Moreover, from the theoretical work of Arrow and Kurz (1970) it is known that if a government's infrastructural investment program is optimal, then the rate of return on infrastructure projects should equal the rate of return on private capital. Otherwise it would be beneficial to increase investments in infrastructure even if this would result in less investment in the private sector. (However, this reasoning is based on the assumption that capital is freely transferable between an economy's private and public sectors.)

Specifically, these cost-benefit studies do not solely rely on economic returns from infrastructure projects, but on environmental impact evaluations as well. However, since a project will not be undertaken if it does not have an expected positive return, and if the ex-ante evaluations of the returns from road infrastructural projects are correct, then in principle it should be possible to find ex-post empirical evidence on the impact of road infrastructure on private output, in particular if the empirical analysis aggregates over individual projects.

This is the first objective of this paper. Our research provides an estimation of road infrastructure's impact on production in the manufacturing sector from an ex-post perspective. Our focus on road infrastructure, which is often referred to as being a part of the "core" public capital on the one hand, and the manufacturing sector on the other hand, reflects the view, that if road infrastructure has any effect on private production, this effect is most likely to be found in the manufacturing sector than in others.

The second objective of this paper is to treat the econometric issues of estimation seriously. Specifically, our estimations take autocorrelation, heteroscedasticity and cross-sectional correlation into account, and we consider also nonstationarity of data. It is worth mentioning that our paper does not explicitly deal with issues of causality. That is, we do not ask whether infrastructure causes higher output or whether higher output leads to a higher demand for infrastructure. This question, which is of particular importance for interpreting the results, has been examined in a recent strand of empirical infrastructure literature applying the concept of Granger causality (Seitz 1995, Ehrenburg and Wohar 1995, Schlag 1997). The results of these studies are so far rather ambiguous. For example, Schlag (1997) finds a bi-directional causality (feedback) between infrastructure capital and output for the German Bundesländer, while Seitz's (1995) results from a panel of 99 cities in Germany indicate a strong uni-directional causality running from public capital to output, and a bi-directional causality between public and private investments.

The remainder of this paper is organized as follows. In section 2, we discuss some of the related literature in more detail. Section 3 explores the concept of our analysis. In particular, we elaborate upon the idea that infrastructure enters the production of private firms. In section 4 we describe the empirical implementation and the obtained results. At the end of section 4, the findings from the analysis are summarized. Finally, the implications of our findings are discussed in section 5.

2 Related Literature

In this section we provide an overview of empirical infrastructure studies based on the production function approach. Table 1 shows a summary of studies at the *regional* level for the US.

{Insert Table 1 about here}

One study which is closely related to our own research is that from Hulten and Schwab (1991). This study, like ours, also focuses on regional manufacturing. The main finding of Hulten and Schwab is that public infrastructure does not have an effect on regional total factor productivity (TFP) growth in U.S. manufacturing.

The picture emerging from Table 1 is that the results of the various studies are rather diverse. While some studies find positive and significant effects of infrastructure, others find only negligible or insignificant effects. Furthermore, the size of the estimated output elasticity of infrastructure capital ϵ_{YG} differs considerably.

Another interesting insight from Table 1 is that the degree of the estimation's econometric sophistication also varies substantially among these studies. While most of the newer studies take the data's panel structure explicitly into account (by including fixed or random effects), some of the older studies have ignored this potential source of bias. Additionally, more recent studies also consider the data's time series properties, for example by taking first differences or providing unit root tests. It should also be noted that most of the newer studies use the same data from Munnell (1992) for public capital stock. Considering this, similar patterns of findings in different studies become less surprising, for example that sewer and water systems are significant, but highways are not.

{Insert Table 2 about here}

Table 2 shows an overview of studies at the *regional* level for countries *other* than the US. Again, while some studies find positive and significant effects others do not. However, these studies are only to a very limited extent comparable since different definitions of public infrastructure or even different levels of regional aggregation are used. For example, Hofmann (1996) has examined the impact of infrastructure on Hamburg's business sector. In this study, Hofmann specifies a Cobb-Douglas production function, which is estimated as a dynamic error correction model. Utilizing data from 1970 to 1992, Hofmann finds an output elasticity of public capital that appears either to be insignificant or to be significant with a negative sign. This result turns out to be rather robust with regards to variations in the econometric specification. In another study at the regional level, utilizing data from 99 German cities from 1980 to 1989, Seitz (1995) finds a positive and significant contribution from infrastructure to private output, with an estimated output elasticity ϵ_{YG} between 0.08 and 0.19.

{Insert Table 3 about here}

Next, we discuss some studies at the *national* level. An overview is shown in Table 3. As mentioned above, the most prominent study at the national level is from Aschauer (1989a). Aschauer estimates an output elasticity of public capital with a value between 0.38 and 0.56, which implies a marginal productivity of public capital of more than 100 percent. This is in sharp contrast to more moderate finding in an earlier study by Ratner (1983), who found the output elasticity of infrastructure capital to be 0.06.

It has been emphasized by several authors that the productivity effects from infrastructure found at the national level might be larger than the effects found at the regional level, because only at the national level are all regional spill-overs from infrastructure fully captured (Munnell 1993).

However, the result of Aschauer's study has led to considerable sceptism in literature (Gramlich 1994, Jorgenson 1991, Tatom 1991a, Tatom 1991b, Tatom 1993). It is argued that the elasticity found by Aschauer is too high to be plausible. Tatom (1991a), for example, points out that the econometric analysis of Aschauer is not appropriate since it neglects the data's time series properties. Specifically, Tatom shows that the times series used by Aschauer are nonstationary. Rerunning the regression from Aschauer (1989a) with variables in first differences and including an energy price variable to control for oil price shocks, it turns out that infrastructure capital no longer appears to be significant.

{Insert Table 4 about here}

Finally, Table 4 provides a summary of findings from studies at the *international* level. One of the earliest studies is from Aschauer (1989c), who finds significant and positive effects from infrastructure for the G7 countries for the period 1966-1985. Extending his study, Aschauer (1995) estimates the productivity effects from infrastructure for 12 OECD countries. It turns out that the effects are significant with an output elasticity between 0.33 and 0.55. Another study by Ford and Poret (1991) on 12 OECD countries takes the data's time series characteristics explicitly into account and obtains mixed results. Only the estimates for 5 countries, that is the US, Germany, Canada, Belgium and Sweden are significant. In a more recent study on 7 OECD countries, taking both the data's time series and panel data structure into account, Nourzad and Vrieze find a relatively low, but significant output elasticity for infrastructure with a value of 0.05.

To summarize this short review, the published results on the productivity effects of infrastructure so far are rather ambiguous. Moreover, this overview has shown that early studies in general have used rather simplistic and inappropriate econometric techniques to study the productivity effects of infrastructure, and that these results turned out to be spurious when applying more appropriate econometric techniques. However, a substantial number of studies exist using appropriate econometric techniques and documenting positive and significant effects from infrastructure.

3 Methodological Framework

Our analysis of the productivity effects of road infrastructure is based on the production function approach. In principle, it would also be possible to investigate the productivity effects with a cost or profit function approach (see for example Berndt and Hansson (1992) or Seitz (1994)). The concept of duality outlined in the classical study of Shephard (1953) and further exploited by Diewert (1974) implies that the production, cost and profit function approach should yield equivalent results. However, these approaches differ in their empirical implementation. While the production function approach treats inputs as given and output as endogenous, the cost function approach treats output and factor prices as exogenous, whereas factor inputs and costs are treated as endogenous. Moreover, the cost function approach also assumes that factor markets are competitive and that firms behave as cost minimizers, given the specific level of output which they want to produce. We argue here that the production function approach is more appropriate because it relies on less restrictive assumptions. Furthermore, the basic assumption of the cost function approach that output is given and exogenous, whereas factor inputs are endogenous determined by cost minimizing behaviour of firms, is not particularly realistic if one considers the relatively high level of aggregation over firms and industries we use in this study.[2]

Our analysis using the production function approach is based on two major assumptions. First, it assumes that road infrastructure enters the production function of a firm as a public intermediate input. Second, it assumes that production in the manufacturing sector can be described by an aggregate production function. In the following two sections we explore these basic assumptions in further detail.

3.1 Infrastructure as a Public Input to Private Production

The manufacturing sector uses many goods and services in production. Besides private inputs such as labor and capital, a manufacturing firm also uses publicly provided inputs, such as highways, roads, water and sewer facilities. For example, highways and roads are important for transporting intermediate and finished goods as well as for employees for commuting to work. An increase in the stock of highways and roads increases the quantity of transportation services available to firms, thus leading to lower transportation costs. In addition, efficiency in private factor inputs utilization increases as well as the firm's productivity (Deno 1988).

In contrast to private inputs, which are purchased on the market, public inputs are provided by Federal, State or local governments. Since most of the infrastructure projects face long construction periods, they can only slowly adjust to changes in desired levels. Thus, from the viewpoint of a manufacturing firm, the amount of public input is fixed in the short-run. However, it is very likely that in the long run the supply of public input is variable to firms, since they can influence the

[2]On this, see also Berndt (1991, p. 457).

allocation of infrastructure investments, for example through "lobbying" activities or other direct involvements in the political process (Eberts 1990).

The theoretical treatment of public inputs which enter private production goes back to Kaizuka (1965) and Sandmo (1972). These papers show, in analogy to the famous paper by Samuelson (1954) dealing with a pure public good for consumers, that resources are allocated efficiently when the sum of private producer's marginal rate of substitution between a public good and the labor service is equal to the marginal rate of substitution between two goods in the public production of the public good.

Furthering this idea, Negishi (1973), for example derives conditions for the optimal supply of public inputs to firms. The optimal level of public inputs is achieved if the government supplies a level of public inputs that maximizes the joint net profit of industries.

To summarize, the concept of public inputs entering private production is theoretically well developed. How relevant this concept is to road infrastructure as a public input for the manufacturing sector is an empirical question, which is the focus of our paper.

3.2 Specification of the Production Function

We begin our analysis by defining a firm's production function f_i. Meade (1952) identifies two channels through which public input can affect privately produced output. First, public input may act as an "environmental factor" which enhances productivity in a way similar to technical progress. We denote this type of public input as "atmosphere" goods. Second, infrastructure may directly enter the production function of a firm as an "unpaid factor of production".

We first model the case where firm i uses public input as an "unpaid factor of production". Then, a firm's production function f_i can be written as

$$q_i = A_i(t) f_i(g_i, k_i, l_i), \tag{1}$$

where q_i denotes firm's i output, $A_i(t)$ shifts in the production function due to technical progress, k_i private capital, l_i labor and g_i the public input (e. g. roads, highways, airports, water and sewer systems, etc.). Assuming that the production technology used by the firm can be described by the simple Cobb-Douglas function, equation (1) then becomes

$$q_i = A_i(t) g_i^{\beta_{g_i}} k_i^{\beta_{k_i}} l_i^{\beta_{l_i}}. \tag{2}$$

The Cobb-Douglas function is both homothetic and strongly separable. It is possible to show using the general aggregation theorem provided by Chambers (1988, pp. 192) that the aggregate production function $F\left(\sum_{i=1}^{n} f_i(g_i, k_i, l_i)\right)$ of the $i = 1 \ldots n$ firms exists for a homethetic and strongly separable production function. To make this more explicit, write the aggregate production function of the manufacturing sector as

$$q = A(t) g^{\beta_g} k^{\beta_k} l^{\beta_l}, \tag{3}$$

where aggregate output is given as $q = \sum_{i=1}^{n} q_i$, and aggregate factor inputs are given as $g = \sum_{i=1}^{n} g_i$, $k = \sum_{i=1}^{n} k_i$ and $l = \sum_{i=1}^{n} l_i$.

Now consider the second case, where public input is an "atmosphere" good and enhances productivity of private factors in a way similar to technological progress. For simplicity, we first assume that a change in the quantity of public input acts like a Hicks-neutral shift of the production function. A single firm's i production function then becomes

$$q_i = A(g,t) f_i(k_i, l_i). \tag{4}$$

Again, the general aggregation result implies that the aggregate production function exists for the Cobb-Douglas. This aggregate function can be written as

$$q = A(g,t) k^{\beta_k} l^{\beta_l}. \tag{5}$$

The distinction of these types of externalities created through public input has important implications for returns to scale at the aggregate level. In the case of the "unpaid factor of production" doubling output requires the doubling of all inputs including public input. In the case of "creation of atmosphere" doubling output only requires doubling of private factor inputs excluding the public input. Thus, in the case of "unpaid factor of production", returns to scale are constant both at the single firm and aggregate levels, while in the case of "atmosphere goods" returns to scale at the single firm level are constant, but returns to scale at the aggregate level are increasing.

Unfortunately, it is not possible to distinguish between the "unpaid factor of production" and the "creation of atmosphere" cases within the Cobb-Douglas function framework. To make this explicit, we first differentiate both sides of (5) with respect to t and find

$$\dot{q} = \dot{A}(g,t) + \epsilon_k \dot{k} + \epsilon_l \dot{l}, \tag{6}$$

where a dot over a variable denotes the logarithmic derivative with respect to time (e. g. $\dot{y} = d \ln y / dt$), and ϵ_k (ϵ_l) denotes the elasticity of output with respect to capital (labor). Furthermore, following the framework of Hulten and Schwab (1991), $\dot{A}(g,t)$ can be divided into two components

$$\dot{A}(g,t) = \dot{A}^* + \gamma \dot{g}, \tag{7}$$

where $\dot{A}^*$ is the growth rate of the "true" Hicksian efficiency term and γ is the elasticity of $A(g,t)$ with respect to g.

Now, let $\dot{A}(g,t) \equiv \dot{\text{TFP}}$, where $\dot{\text{TFP}}$ denotes total factor productivity growth. If each input is paid the value of its marginal product, the elasticities in equation (6) are equivalent to cost shares, that is, $s_k = p_k\, k / p_q\, q$ and $s_l = p_l\, l / p_q\, q$, where p_k and p_l are factor prices of capital and labor and p_q is the price of output. Using the income shares s_k and s_l, total factor productivity growth is given by

$$\dot{\text{TFP}} = \dot{q} - s_k \dot{k} - s_l \dot{l}. \tag{8}$$

Thus, estimation of γ could be based on the following equation

$$\dot{\text{TFP}} = \dot{A}^* + \gamma\dot{g} + \varepsilon, \tag{9}$$

where ε denotes an iid error term. However, if g also enters the production function as an unpaid factor of production, γ cannot be estimated using equation (9). To see this, note that equation (6) has the additional term $\epsilon_g\,\dot{g}$, and becomes

$$\dot{q} = \dot{A}(g,t) + \epsilon_g\dot{g} + \epsilon_k\dot{k} + \epsilon_l\dot{l}. \tag{10}$$

From (10), equation (9) becomes

$$\dot{\text{TFP}} = \dot{A}^* + (\gamma + \epsilon_g)\dot{g} + \varepsilon. \tag{11}$$

Therefore, if we estimate the parameter β_g within a TFP framework it captures both effects of $\dot{g}$, that is $\beta_g = \gamma + \epsilon_g$.

3.3 Translog Production Function Specification

From the last section it follows that in order to establish road infrastructure's impact on private factor productivity, we have to employ a more general production function framework, one that allows us to measure the "creation of atmosphere" effect from infrastructure on a single private factor's productivity. This can be achieved by using a production function with a flexible functional form, which is a functional form that does not place a priori restriction on elasticity of substitution.

A very popular choice for this in applied research is the *transcendental logarithmic* (for short: translog) function (Christensen, Jorgenson, and Lau 1971, Christensen, Jorgenson, and Lau 1973). In its general form, the translog function can be defined as

$$\ln y(t,x) = \beta_0 + \beta_t t + \sum_{i=1}^{n} \beta_i \ln x_i + \frac{1}{2}\sum_{i}^{n}\sum_{j}^{n} \beta_{ij} \ln x_i ln x_j, \tag{12}$$

where $x = (x_1, x_2, \ldots, x_n)$ denotes a vector of inputs. The translog is interpretable as a numerical second order approximation to an arbitrary function in the neighborhood of $x_0 = (1, 1, \ldots, 1)$.

The effect from public input g on private factor productivity, i.e. $\partial y/\partial k$ and $\partial y/\partial l$, can be derived from an estimation of equation (12) as

$$\frac{\partial^2 \ln y}{\partial \ln g \partial \ln k} = \hat{\beta}_{gk}, \qquad \text{and} \qquad \frac{\partial^2 \ln y}{\partial \ln g \partial \ln l} = \hat{\beta}_{gl}, \tag{13}$$

from which $\partial y/\partial k$ and $\partial y/\partial l$ can be calculated as

$$\frac{\partial^2 y}{\partial g \partial k} = \hat{\beta}_{gk}\frac{y}{g\,k}, \qquad \text{and} \qquad \frac{\partial^2 \ln y}{\partial \ln g\ \ln l} = \hat{\beta}_{gl}\frac{y}{g\,l}. \tag{14}$$

Within a translog function framework several restrictions on production technology can be tested. If technology is homogeneous, then the sum of the coefficients of the squared terms and the cross-effects will be zero:

$$\sum_{i}^{n}\sum_{j}^{n}\beta_{ij} = 0. \tag{15}$$

In addition, linear homogeneity requires the above condition plus that the sum of the linear terms equals one:

$$\sum_{i}^{n}\beta_{i} = 1. \tag{16}$$

It is worth noting that the translog function cannot represent a flexible separable technology, that is without putting further restrictions on its parameters. Again, only if technology is separable and homothetic then it also is consistent with aggregation.

3.4 Allen Partial Elasticity of Substitution

In order to measure the elasticity of substitution, we apply a concept from Allen (1938). The Allen partial elasticity of substitution, σ_{ij}, is defined as

$$\sigma_{ij} = \frac{\sum_i x_i f_i}{x_i x_j}\frac{F_{ij}}{F} \tag{17}$$

where F is the bordered Hessian determinant

$$F = \begin{vmatrix} 0 & f_1 & f_2 & \dots & f_n \\ f_1 & f_{11} & f_{12} & \dots & f_{nn} \\ f_2 & \dots & 1 & \dots & f_{nn} \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ f_n & f_{1n} & \dots & \dots & f_{nn} \end{vmatrix}, \tag{18}$$

and f_{ij} denotes partial derivates of f with respect to inputs i, j, and where F_{ij} is the cofactor associated with f_{ij}. The Allen partial elasticity of substitution refers to the degree of substitutability between input pairs. It measures the curvature of a production isoquant, and hence measures the substitutability of input pairs while holding output fixed. Inputs are *complements*, if the elasticity of substitution is negative, and *substitutes*, if the elasticity is positive (Chambers 1988, pp.33).

3.5 Outline of the Empirical Analysis

In the first part of the analysis, assuming that $A(t)$ can be described as $A_0 \exp(\lambda t)$ and taking logarithms on both sides of (3), we estimate the empirical counterpart of (3) as

$$\ln q = \beta_0 + \lambda t + \beta_g \ln g + \beta_k \ln k + \beta_l \ln l + \varepsilon, \tag{19}$$

where β_i measures the elasticity of output with respect to factor i. It is worth noting that the Cobb-Douglas function in its logarithmic form is interpretable as a first-order numerical approximation of an arbitrary production function in the local neighborhood of $x_0 = (1, 1, \ldots, 1)$. The limitation of the Cobb-Douglas function in empirical research is that it restricts the substitution elasticities of input pairs to equal one.

In the second part, the results of the estimation of equation (11) are presented. As discussed above, it is not possible to measure using equation (11) whether the theoretical case of "unpaid factor of production" or "creation of atmosphere" is more relevant to the manufacturing sector. However, equation (11) has the advantage over (19) in that it might not suffer from simultaneous equation bias between output and infrastructure. If simultaneity bias would be present in equation (19), then a substantial difference in the parameter estimates of (11) and (19) should be observed.

In the third part of the empirical analysis, we present the estimation results for the translog production function of equation (12), which is specified as

$$\begin{aligned}\ln q = \beta_0 &+ \lambda t + \beta_g \ln g + \beta_k \ln k + \beta_l \ln l \\ &+ \tfrac{1}{2}\left(\beta_{tt} t^2 + \beta_{gg} \ln^2 g + \beta_{kk} \ln^2 k + \beta_{ll} \ln^2 l\right) \\ &+ \beta_{gk} \ln g \ln k + \beta_{gl} \ln g \ln l + \beta_{kl} \ln k \ln l \\ &+ \beta_{tg}\, t \ln g + \beta_{tk}\, t \ln k + \beta_{tl}\, t \ln l + \varepsilon.\end{aligned} \tag{20}$$

Based on these translog function estimates, we calculate the marginal productivity of road infrastructure and the Allen partial elasticity of substitution for each Bundesland. The empirical analysis concludes with a summary of findings.

4 Econometric Implementation and Results

We utilize annual data from the 11 German Bundesländer ($G = 11$) for the period from 1970 to 1993 ($T = 24$). The data set is fully described in Appendix A. Ordinary Least Squares (OLS) estimation of the Cobb-Douglas production function, as specified in (19) using the pooled time-series cross-section data yields the following result:

$$\begin{array}{llllll}\ln q = & \textit{Länder-effects}^* & +0.012^*\, t & +0.734^* \ln g & +0.119^* \ln k & +0.809^* \ln l \\ & F(10|\ 249){=}286^* & (0.002) & (0.060) & (0.059) & (0.055)\end{array} \tag{21}$$

R^2: 0.99997 N: 264 (G=11, T=24) SE: 0.728 DW: 0.442

Groupwise heteroscedasticty: LM= 104.7*
Cross-sectional correlations: $\lambda_{LM} = 222.9^*$

(* significant at a 5 percent level, standard errors are given in parentheses)

Note that in the empirical implementation of equation (19) we have included dummy variables for the Bundesländer (fixed effects). The shown F-test indicates that the Bundesländer effects are highly significant. Thus, the Bundesländer dummy variables should be included in the regression equation. Furthermore, labor is significant with a value of about 0.8. In addition, the estimate of private capital is significant with a value of 0.12. Moreover, road infrastructure is significant with a parameter of 0.73. At the mean data points, the ratio of private capital to output is about 1, while the ratio of private capital to road infrastructure is about 2. Accordingly, the ratio of output to road infrastructure is about 2. Thus, the estimated output elasticity of road infrastructure implies a marginal productivity of 146 percent.

4.1 Autocorrelation, Heteroscedasticity and Cross-Sectional Correlation

A frequent problem in the empirical analysis of time-series data is the presence of autocorrelation. Since we do not want to impose a constant returns to scale restriction on (11) by dividing it by labor l, it is very likely that heteroscedasticity will be observed, since the Bundesländer in our sample have different sizes. Furthermore, since macroeconomic factors affecting one region will also affect other regions, errors across the regions are likely to be correlated.

With all these shortcomings present, OLS estimation would still yield consistent parameter estimates. However, estimates of standard errors would be biased and inconsistent. Thus, in order to check for the presence of autocorrelation, groupwise heteroscedasticity, and cross-sectional correlation, we report additional test statistics in (21).

First, in order to test for autocorrelation, the Durbin-Watson (DW) statistic is displayed. The DW statistic for this fixed effects models has been computed according to Bhargava, Franzini, and Narendranathan (1982) as

$$DW = \frac{\sum_{i=1}^{T}\sum_{t=2}^{T}(\varepsilon_{it} - \varepsilon_{it-1})^2}{\sum_{i=1}^{T}\sum_{t=1}^{T}\varepsilon_{it}^2}. \tag{22}$$

The shown low value of 0.442 in (21) for the DW statistic indicates that the errors of the OLS estimation are indeed not serially independent.

Second, in order to test for groupwise heteroscedasticity, a Lagrange multiplier (LM) test has been employed. This LM test statistic is defined as (Greene 1993, pp. 448),

$$LM = T/2\sum_{i}\left[\frac{s_i^2}{s^2} - 1\right]^2, \tag{23}$$

where s^2 is the pooled OLS residual variance and s_i^2 is the estimated unit-specific residual variance from groupwise regression. The LM statistic has a limiting chi-squared distribution with $i-1$ degrees of freedom. The reported value of 104.7

from the LM statistic leads to a rejection of the null hypothesis of no groupwise heteroscedasticity ($\chi_{krit,0.05}(10df) = 18.3$).

Third, in order to test for cross-sectional correlation, the following Lagrange multiplier statistic λ_{LM} has been computed (Greene 1993, pp. 452):

$$\lambda_{LM} = T \sum_i \sum_{j<i} r_{ij}^2, \tag{24}$$

where r_{ij}^2 is the squared ijth residual correlation coefficient. The large-sample distribution of this statistic is chi-squared with $i(i-1)/2$ degrees of freedom. Hence, this statistic is significant, indicating the presence of cross-sectional correlations between the Bundesländer ($\chi_{krit,0.05}(55df) = 73.3$).

{Insert Table 5 about here}

Table 5 shows the cross-sectional correlation matrix and the groupwise variance/covariance matrix of the Bundesländer, which have been calculated from the residuals of the OLS estimation in (21). The variances of the residuals for the Bundesländer are given in bold print on the diagonal of the matrix. The ratio of the largest variance with 81 ("Hessen") to the smallest with 4.7 ("Nordrhein-Westfalia") is about 17, again showing the high degree of groupwise heteroscedasticity.

As already indicated by the λ_{LM} statistic, some of the displayed residual correlations are substantially different from zero. Moreover, some of the correlations are remarkably high, for instance between "Baden-Würtemberg" and "Bayern" with a value of 0.74 or between "Rheinpfalz" and "Hessen" with a value of 0.79. On the other hand, a few correlations are surprisingly low; for example between the 3 metropolitan states in sample, "Hamburg", "Bremen" and "Berlin" ("Hamburg" and "Bremen", -0.2; "Hamburg" and "Berlin", 0.35; "Berlin" and "Bremen", 0.04). In summary, the reported high correlations for some of the Bundesländer stress the need for correcting the variance/covariance matrix of the parameters estimates.

4.2 Kmenta Method versus OLS/PCSE

Basically, if the variance/covariance matrix of errors were known, it would be possible to correct for autocorrelation, groupwise heteroscedasticity, and cross-sectional correlations with Generalized Least Squares (GLS) estimation. However, in the practical empirical analysis the variance/covariance matrix of errors has to be estimated.

In the context of time-series cross-section (TSCS) data analysis this Feasible Generalized Least Squares (FGLS) estimation is also known as "Kmenta" (1974, 1986) or "Parks" (1967) method. In the rest of this paper we shall refer to this method as the "Kmenta" method. A short description of the Kmenta method is given in Appendix C.

In two recent articles based on evidence from Monte-Carlo experiments, Beck and Katz (1995, 1996) have pointed out that although FGLS might be more efficient when cross-sectional correlations or groupwise heteroscedasticity are very significant, the standard errors obtained by the Kmenta method do not correctly reflect the sampling variability of parameter estimates. A problematic property of the Kmenta method in small samples is that the cross-sectional correlations or variances obtained in the first step might only be very poor estimates of the underlying "true" variances. This is not properly taken into account when standard errors of the parameter estimates in the second step are calculated. Thus, as Beck and Katz (1995) have shown in their study, standard errors from the Kmenta method in small samples have a tendency to be too small, they lead to "overconfidence". Beck and Katz recommend applying OLS estimation with with panel corrected standard errors (PCSE) if the ratio of T to G is smaller than 3.

In our case this ratio is about 2.2, therefore we present both results obtained from Kmenta and from OLS estimation with PCSE. In Appendix D we give a brief motivation for OLS with PCSE.

4.3 Nonstationarity

Another important issue in the context of time-series analysis is the existence of common trends in data, leading to spurious correlation (Granger and Newbold 1974, Granger 1981). For example, if two wholly unrelated measures have similar time trends, they can exhibit an apparent, statistically significant relationship between themselves when, in fact, no economic relationship exists. First-differencing typically renders the data stationary and removes the problem of justifying or explaining the existence of a deterministic trend or trends. Note, that by differencing variables the parameters are unaffected. Hence, if the parameters of equation (19) are viewed as the appropriate long-run parameters in a levels estimate, they will remain so in first differences (Tatom 1991a, Tatom 1991b).

The Durbin-Watson statistic of Bhargava, Franzini, and Narendrananthan (1982) for panel data with fixed effects can also be used for testing residuals from OLS being generated by a Gaussian random walk, that is the autocorrelation parameter ρ is one. In this case, the expected value of the Durbin-Watson statistic is zero. Thus, the Durbin-Watson statistic can be used to test for both serial independence of errors and residuals being generated by a random walk.

4.4 Results of the Cobb-Douglas Production Function Estimation

Table 6 summarizes the results of the Cobb-Douglas production function from (19) for the pooled time-series cross-section data.

{Insert Table 6 about here}

It is worth noting that the correction for autocorrelation can be done in the first step independent of the second step where Kmenta or OLS/PCSE estimation is applied.

A consistent estimate of the autocorrelation parameter ρ has been obtained from $\rho = 1 - \mathrm{DW}/2$. Using this estimate, an AR(1) correction has been carried out by employing the Cochrane-Orcutt transformation (Greene 1993, pp. 431). As such, the first observation in each group is lost. Note that in the fixed-effects model, the Prais-Winston transformation (Greene 1993, pp. 431) is not an appropriate choice for AR(1) correction, because the "within" transformation, that is forming deviation from group means, will not remove the heterogeneity if the Prais-Winston transformation is used for the first observation.

Table 6 displays results for both the AR(1) corrected variables and variables in first differences. Moreover, the upper half of Table 6 shows the results of the Cobb-Douglas function with three inputs; infrastructure g, private capital k and labor l, while the lower half contains the results for the Cobb-Douglas function with inputs k and l. Standard errors are given in parentheses.

Both the Kmenta and OLS/PCSE results indicate that the group effects are highly significant. The joint test of their significance can no longer be carried out with the usual F-Test for OLS/PCSE, as it relies on homoscedasticity. We can, however, use a Wald test. The Wald statistic W is given as (Greene 1993, p. 188)

$$W = (\mathbf{Rb} - \mathbf{q})' \left(\mathbf{R}(\mathrm{Var}(\widehat{\mathbf{b}})\mathbf{R}'\right)^{-1} (\mathbf{Rb} - \mathbf{q}). \tag{25}$$

Note, that only the AR(1) model includes the Bundesländer dummy variables, since the dummy variables are removed when taking first-differences. In addition, only the AR(1) model includes a time trend t, because this becomes a constant when taking first-differences.

By contrast with the low value of 0.4 for the DW statistic reported above, both the AR(1) corrected and the model with variables in first differences generate DW statistics of about 2, indicating the absence of autocorrelation and of nonstationarity.

The parameter of private capital in the upper half (1) of Table 6 is neither in the Kmenta nor in the OLS/PCSE specification significant, which holds for both the AR(1) and the model in first differences. In contrast to this the estimated parameters for infrastructure appear to be significant in all specifications, with a value ranging from 0.40 to 1.13. These parameter estimates imply a marginal productivity for road infrastructure between 80 and 226 percent. The estimated parameter for labor is in all specifications significant with values about 0.60, which is a reasonable estimate if one considers that the average share of labor in output for our data is about 0.55.

Part (2) of Table 6 shows the results of the Cobb-Douglas function with only private factor inputs, k and l. Without g, the estimate of the output elasticity of private capital becomes more precise. Here, the estimated output elasticity of private capital ranges from 0.16 to 0.22. As before, the estimated output elasticity of labor ranges from 0.56 to 0.64.

Finally, the test for constant returns to scale (CRS) is rejected in almost all specifications.

4.5 Results for Total Factor Productivity Growth (TFP)

In this section we describe the empirical implementation of equation (11), where the growth of total factor productivity ($\dot{\mathrm{TFP}}$) is regressed on the growth of the road infrastructure stock, $(\dot{g})$.

{Insert Table 7 about here}

Total factor productivity growth has been calculated, as described above, according to the definition given in equation (8). The cost shares have been computed as an average of the periods t and $t-1$. The income share of private capital, s_k, has been computed by using the constant returns to scale assumption, that is $s_k = 1 - s_l$.

Table 7 shows the results for 7 different econometric specifications. Model (1) displays the results for the OLS estimation of (11), where the parameter of g is given by $\beta_g = \gamma + \epsilon_g$. Note, that although the reported R^2 is only 0.034, the overall F statistic is significant at a 5 percent level. This indicates that the model can explain a significant part of the variance in total factor productivity growth across the Bundesländer.

Model (2) provides estimates for OLS/PCSE. As expected, this specification yields standard errors which are larger than those from OLS. The null hypothesis, that the parameter of $\dot{g}$ is zero, is not rejected at a 5 percent level (the t-value is 1.58). However, at the 10 percent level this parameter estimate would be significant.

Model (3) provides the results for the Kmenta method. As shown, both groupwise heteroscedasticity and cross-sectional correlations are significant. As expected, Kmenta standard errors are lower than the computed PCSE. The Kmeta method yields a parameter estimate for $\dot{g}$ (0.325) that is of comparable size to the estimate from OLS (0.366).

Specifications (4) and (5) contain the results for one-way fixed and random-effects panel data models. In contrast to the usual practice, we model the time effects in the one-way specification. The reason for this is that, by taking first differences the group effects become insignificant, as expected from our considerations above. Table 7 shows that unlike the group effects, the time effects are highly significant for both one-way models (random and fixed). In the two-way fixed (6) and random-effects models (7), the Bundesländer dummy variables are included. However, as before, the Bundesländer effects turn out not to be significant (the reported value of the F-test is 1.34).

In order to test whether the fixed or random effects model is the appropriate specification, the last column of (5) and (7) shows the results of a Hausman test. The Hausman statistic tests the null hypothesis that the difference of the GLS and the within estimator is zero. Under the null hypothesis both the GLS and the within estimator are consistent, but the GLS is more efficient, while under the alternative hypothesis the GLS becomes inconsistent. Thus, large values of the Hausman test statistic weigh in favor of the fixed effects model. In our case, the null hypothesis is not rejected for both the one-way and the two-way models, indicating that random effects are the appropriate specification.

In summary, the parameter estimates for $\dot{g}$ in the panel data models (4) to (7) are larger than the estimates from OLS and the Kmenta method, but of a comparable size with Table 6. Despite the significance of $\dot{g}$ holding for all specifications, the explained variance, expressed in terms of R^2, is relatively low, ranging from 0.032 to

0.034. (The exception being the fixed-effects models, where time dummy variables are included).

Thus, one result which emerges from Table 7 is that the growth rate of road infrastructure can explain about 3.3 percent of the observed differences in total factor productivity growth across the Bundesländer during the period from 1970 to 1993.

On the basis of the panel data analysis it is possible to gain further interesting insights. While the fixed-effects estimation captures only the within groups variation and neglects the between groups variation, the random effects estimation takes both the within and the between groups variation into account. Since the described one-way fixed effects model includes time effects, the within groups variance is given in our case by the variation between the Bundesländer, while the between groups variance is determined by the variation across years (taking the average for the Bundesländer for each year.)

{Insert Figure 3 about here}

Figure 3 shows a graph of the average (over the Bundesländer) annual growth rate of TFP versus the average annual growth rate of road infrastructure ($\dot{g}$). The average growth of road infrastructure was about 4 percent from 1971 to 1980, while the average growth of TFP was about 1.9 percent. From 1981 to 1990, the average growth of road infrastructure was 0.8 percent, while the average growth of TFP was 1.3 percent.

Supplementary to this Figure, the results of a between years regression are shown. The results show that the explanatory power of ($\dot{g}$) in explaining the observed pattern of TFP is very limited as the reported overall F-test is not significant. Similarly, the estimated parameter of $\dot{g}$ is also not significant. However, it becomes obvious from Figure 3 that the positive relationship between TFP and road infrastructure growth would appear much stronger if the years 1970 to 1974 were not in the sample.

One could argue, however, that in the within time-series case capacity utilization could have an important role for explaining the growth of total factor productivity, because it captures the observed fluctuations in output due to business cycles. As a test for this using the aggregate capacity utilization (CU) rate as a approximation for the "average" utilization rate, we find the following results for the between years regression:

$$\begin{array}{lccc} \dot{\mathrm{TFP}} = & 0.751 & +0.351\dot{g} & +58.9^{*}\,\dot{\mathrm{CU}}. \\ & (0.541) & (0.189) & (11.5) \end{array}$$

R^2: 0.59 N: 253 (G=11, T=23) DW: 2.57

As shown, the parameter of $\dot{g}$ is still not significant at the 5 percent level (but would be at the 10 percent level). However, the increase in the *DW* statistic from 2 to 2.6 indicates that the model is now mis-specified.

{Insert Figure 4 about here}

Next, we turn to the "within" years variation analysis of our panel data study. We shall refer to the "within" years variation as the "between" variation of the Bundesländer, which corresponds more to the usual terminology.

Figure 4 graphs the average growth rate of TFP (over years) versus the average growth of road infrastructure. It is part of the "folk-wisdom" of applied econometrics, that the cross-sectional "between" regression measures the "long-run relationship" of the variables (Sevestre and Trognon 1996).

As expected, the results for the between Bundesländer regression are quite similar to the results for the one-way fixed effects model in Table 7. However, Figure 4 shows that the results for the between Bundesländer regression depends strongly on whether or not the three Bundesländer Hamburg, Bremen and Saarland are included in the analysis. It can be stated that without these three Bundesländer the relation between average $\dot{\text{TFP}}$ and average $\dot{g}$ would not be significant.

In summary, Figures 3 and 4 suggest that the between Bundesländer variation is more important in explaining the relationship between road infrastructure and Total Factor Productivity growth than the between years variation.

4.6 Results for the Translog Production Function

In this section we describe the results for the estimation of the translog production function as specified in (20). As mentioned above, the aim of the translog production estimation is to measure road infrastructure's effects on private factor productivity. By OLS estimation of (20) we get the following results:

$$\begin{array}{llllll}
\ln q = & \textit{Länder-effects}^* & +0.023^*\, t & +0.302^* \ln g & +0.154 \ln k & +0.683^* \ln l \\
 & F(10|239){=}286^* & (0.009) & (0.143) & (0.088) & (0.113) \\
 & & & & & \\
 & +0.001\, t & -0.138 \ln^2 g & +0.846^* \ln^2 k & +0.650^* \ln^2 l & -0.084 \ln g \ln k \\
 & (0.001) & (0.166) & (0.391) & (0.320) & (0.199) \\
 & & & & & \\
 & -0.147 \ln g \ln l & -0.671^* \ln k \ln l & +0.003\, t\ \ln g & -0.018\, t\ \ln k & +0.020\, t\ \ln k. \\
 & (0.116) & (0.302) & (0.004) & (0.010) & (0.011)
\end{array} \tag{26}$$

R^2: 0.997 N: 264 (G=11, T=24) SE: 0.659 DW: 0.572

Groupwise heteroscedasticity: LM= 74.8^*
Cross-sectional correlations: $\lambda_{LM} = 223.6^*$

(* significant at a 5 percent level)

Again, the reported low Durbin-Watson (DW) statistic shows that the errors are not serially independent, leading least squares to be inefficient and inference based on least squares estimates to be adversely affected. Additionally, as in the case of the Cobb-Douglas production function estimation both groupwise heteroscedasticity and cross-sectional correlations turn out to be significant. Therefore, as before, both the Kmenta and OLS/PCSE have been estimated using AR(1) corrected variables and in first-differences.

{Insert Table 8 about here}

Table 8 summarizes the results of the translog production function estimation. The reported Wald and F statistics indicate, that the Länder-effects are both in the Kmenta and OLS/PCSE setup highly significant. Similarly to the the Cobb-Douglas function estimation, the DW statistic for the AR(1) corrected and first differences variables is about 2, indicating that errors are serially independent and stationary.

The estimate of g appears to be significant in all specifications. In contrast to this, the estimate of k turns out to be insignificant for all specifications. The estimate of $(\ln g)^2$ is for the AR(1) model with the Kmenta method significant. The negative sign indicates decreasing returns to scale in g. However, since this result does not emerge in any other specification, we suspect that it is not a very robust result.

Overall, the results of the translog estimation are rather unsatisfactory. A reason for this could be the high correlation between the linear and the quadratic and cross terms in the translog specification. Due to this multicollinearity it might be difficult to achieve more precise estimates of single terms.

{Insert Table 9 about here}

To get an impression which of the terms of the translog might be important, Table 9 summarizes several tests based on the translog estimates. As before, a Wald test is for OLS/PCSE and a F-test is for the Kmenta method applied. The hypothesis tested is that a subset of the translog, say all terms containing g, are zero.

In addition, the lower half of Table 9 shows the results for the hypothesis, that the linear terms (that is t, $\ln g$, $\ln k$, $\ln l$), cross terms (that is $\ln g \times \ln k$, $\ln \times g \ln l$, $\ln k \times \ln l$, $t \times \ln g$, $t \times \ln k$, $t \times \ln l$), or quadratic terms (that is $t \times t$, $\ln g \times \ln g$, $\ln k \times \ln k$, $\ln l \times \ln l$) are zero[3]. Note, that degrees of freedom are different for the AR(1) and the variables in first differences, since in the latter the linear time trend t is removed.

We find for all specifications that labor and time are significant. Furthermore, we find in 3 out of 4 specifications that road infrastructure is important, while private capital appears to be not significant. The linear and cross terms are import in all specifications, whereas the quadratic terms are only significant in the AR(1) Kmenta estimation.

{Insert Table 10 about here}

[3]The critical values for the χ^2 at a 5 percent level for 3, 4, 5 and 6 degrees of freedom are 7.81, 9.48, 11.07 and 12.59, and for F 2.64, 2.40, 2.25 and 2.13.

Finally, in Table 10 both the marginal factor productivity and the Allen partial elasticity of substitution for each Bundesland are shown. We emphasize, that these Figures should be interpreted with some caution and are only for a descriptive purpose, because they have rather large variances (not reported here) which stem from the imprecise parameter estimates of the translog function. Table 10 shows that for all of the Bundesländer road infrastructure and private capital and similarly road infrastructure and labor are Allen substitutes, whereas private capital and labor are Allen complements.

4.7 Summary of Empirical Results

Our analysis has explicitly taken into account four of the most frequent problems in the context of time-series cross-section analysis: serial correlation, groupwise heteroscedasticity, cross-sectional correlation and nonstationarity of data. In summary, we find a strong positive and significant correlation between road infrastructure and the manufacturing sector's output in all of the tested specifications.

Specifically, within the Cobb-Douglas production function framework, we find support for the idea that infrastructure enters into the production of firms as a public input. However, the implied marginal productivity of road infrastructure seems to be too high to be a plausible estimate of the "true" productivity of road infrastructure.

Moreover, for the regression of TFP growth on road infrastructure growth we find that the variation between Bundesländer is more important for explaining the contribution of road infrastructure to output than the variation between years. As the major result of this section, we state that road infrastructure can explain about 3.3 percent of the observed variance in total productivity growth across the Bundesländer.

Finally, the translog function estimation yields less satisfactory results. Although the joint test of cross-terms suggests that these are important, the single terms appear not to be significant. Therefore, since the estimates are rather imprecise, it is not possible to conclude whether or not infrastructure creates an "atmosphere" and thereby enhances private factor productivity.

5 Concluding Remarks

The starting point of this paper has been Aschauer's "Public Capital Hypothesis", which states that the decline in government's infrastructural spending in the US and other major OECD countries since the mid 70's can explain a major part of the observed decline in productivity growth during the same period.

In summary, our findings indicate a strong correlation between road infrastructure and output in German manufacturing at the regional level of the Bundesländer. One conclusion we draw from this is that differences in road infrastructure might explain a part of the existing productivity gap between manufacturing in east and west German regions.

Turning to the evidence over time, we find that the explanatory power of road infrastructure in explaining the observed pattern in TFP growth over time is rather limited. We conclude from this that other exogenous factors might be more relevant than road infrastructure for explaining the development in TFP growth.

A limitation of this study is that the estimated output elasticities of road infrastructure ranging from 0.325 to 1.130 in different specifications are too high to be plausible estimates of the "true" returns on road infrastructure. One reason for this result could be the omission of other relevant variables. Suggestions of possible candidates are given below in the discussion on future research. Another explanation could be that road infrastructure is not "truly" exogenous, which in turn would also lead to biased estimates.

Another implausible finding of our study is that private capital turns out in almost all specifications not to be significant when road infrastructure is included at the same time in the regression. An explanation for this could be the relatively high correlation between public and private capital stocks, as well the growth rates of the stocks are highly correlated. A solution to this problem would be to use a cost or profit function approach, but as we have already mentioned before, this would be at the cost of more restrictive (and in turn less realistic) assumptions than have been used for our production function approach.

Finally, we consider three modifications to our analysis for future research. First, as has been suggested by Tatom (1993) and Sturm and Kuper (1996a), energy input of the manufacturing sector should be included in the production function. Second, since the utilization of private capital changes over time due to business cycles, regional manufacturing capacity utilization rates should be included to capture these short-run fluctuations. Third, since the growth of road stock does not necessarily correspond to the growth of effective capacity, the capacity utilization of roads should be taken into account as well.

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Appendix

A Data Description

A.1 Public Sector Investment and Consumption

Public Sector Consumption Both time series "Public Sector Consumption" and "Non-Military Public Sector Consumption" have been obtained from the annual publication "Statistical Yearbook for the Federal Republic of Germany" by the Federal Statistical Office in Wiesbaden, chapter "Volkswirtschaftliche Gesamtrechnung". Public sector consumption includes public expenditures for administrative purposes, health, education, etc. In addition, it also incorporates the social transfer payments from the public sector to private households.

Public Sector Investment Total public sector investment has been taken from the "Statistical Yearbook for the Federal Republic of Germany" by the Federal Statistical Office in Wiesbaden, chapter "Volkswirtschaftliche Gesamtrechnung". It embraces non-military investment in machinery and equipment, and investment in residential and non-residential construction. Public sector road and highway investment has been obtained from the annual yearbook "Verkehr in Zahlen" by the "Deutsches Institut für Wirtschaftsforschung" (DIW), Berlin. For a further description of this data, also see the paragraph "Road infrastructure (g)" in the next section.

A.2 Production Function Estimation

The data used in our empirical analysis covers the manufacturing sector of the 11 German Bundesländer from 1970 to 1993. We utilize the following variables:

Output (q) Output is measured as gross value-added of the manufacturing sector in 1991 constant prices aggregated over industries. These data have been obtained from "Volkswirtschaftliche Gesamtrechnung der Länder, Heft 30: Entstehung des Bruttoinlansprodukts in den Ländern der Bundesrepublik Deutschland 1970 bis 1996", which is provided by the "Statistisches Landesamt Baden-Würtemberg".

Note, that for the translog function estimation we have calculated indices of output and input measures by dividing the measure of each year by the measure of the base year 1970 (1970=100). The use of indices is not required for the Cobb-Douglas function in logarithms, since parameter estimates are not affected by using the "rough" output and input measures in 1991 constant prices.

Labor (l) Labor is measured in terms of working hours in the manufacturing sector. These data are taken from the series "Statistical Yearbook for the Federal

Republic of Germany" published by the Federal Statistical Office in Wiesbaden. Working hours are only a measure for firms with more than 10 employees.

Alternatively to this labor input measure, we have also estimated the production function with the number of employees as the labor input. The Number of employees in the manufacturing sectors of the German Bundesländer is published in "Volkswirtschaftliche Gesamtrechnung der Länder, Heft 30: Entstehung des Bruttoinlansprodukts in den Ländern der Bundesrepublik Deutschland 1970 bis 1996", Statistisches Landesamt Baden-Würtemberg. The differences in the obtained parameter estimates are rather small, therefore we have refrained from reporting these results.

Private Capital (k) Private capital is measured as the net capital stock in the manufacturing sector. It includes machinery, equipment and buildings, and is taken from "Volkswirtschaftliche Gesamtrechnung der Länder, Heft 29: Anlageinvestitionen, Anlagevermögen und Abschreibungen in den Ländern der Bundesrepublik Deutschland 1970 bis 1995". This statistical report is also provided by the "Statistisches Landesamt Baden-Würtemberg, Arbeitskreis Volkswirtschaftliche Gesamtrechnung der Länder".

Capacity Utilization Rate (CU) Capacity utilization rate of private capital at the national level was obtained from "IFO Schnelldienst", 31, 1997, published by the IFO Institut, München. Note, that CU is only included in the estimation on page 16, because this capacity utilization measure is not available at the regional level of the Bundesländer. However, since we use working hours and not the number of employees for labor input in the production function, we expect that fluctuations in output due to business cycles will be at least partially captured by this measure of labor input.

Road Infrastructure (g) The road infrastructure is the main part of the total transportation infrastructure stock (about 75 percent), which additionally also includes airports, water transport, mass transit systems, railways, etc. Road infrastructure stock is a measure of the capital stock of roads, bridges and highways. We use the *net* capital stock measure. Figures for the *aggregated* net capital stock of road infrastructure have been obtained from "Verkehr in Zahlen" published by the "Deutsches Institut für Wirtschaftsforschung" (DIW), Berlin. However, road infrastructure stock data at the regional level of the Bundesländer are not reported in this publication.

Thus, in order to obtain the regional infrastructure stock data, we have gathered data on roads, bridges and highways investments in the Bundesländer from various statistical sources:

- Ministry of Transport: *Istausgaben der Bundesfernstraßen in den Jahren 1971 bis 1990*, internal report, Bonn, 1991.

- Ministry of Transport: *Straßenbaubericht*, annual report for the German parliament, Bonn, 1971-1994.
- *Government budget plans of the Bundesländer*, 1970-1995.
- *Selected statistical yearbooks of the Bundesländer*, 1970-1995.

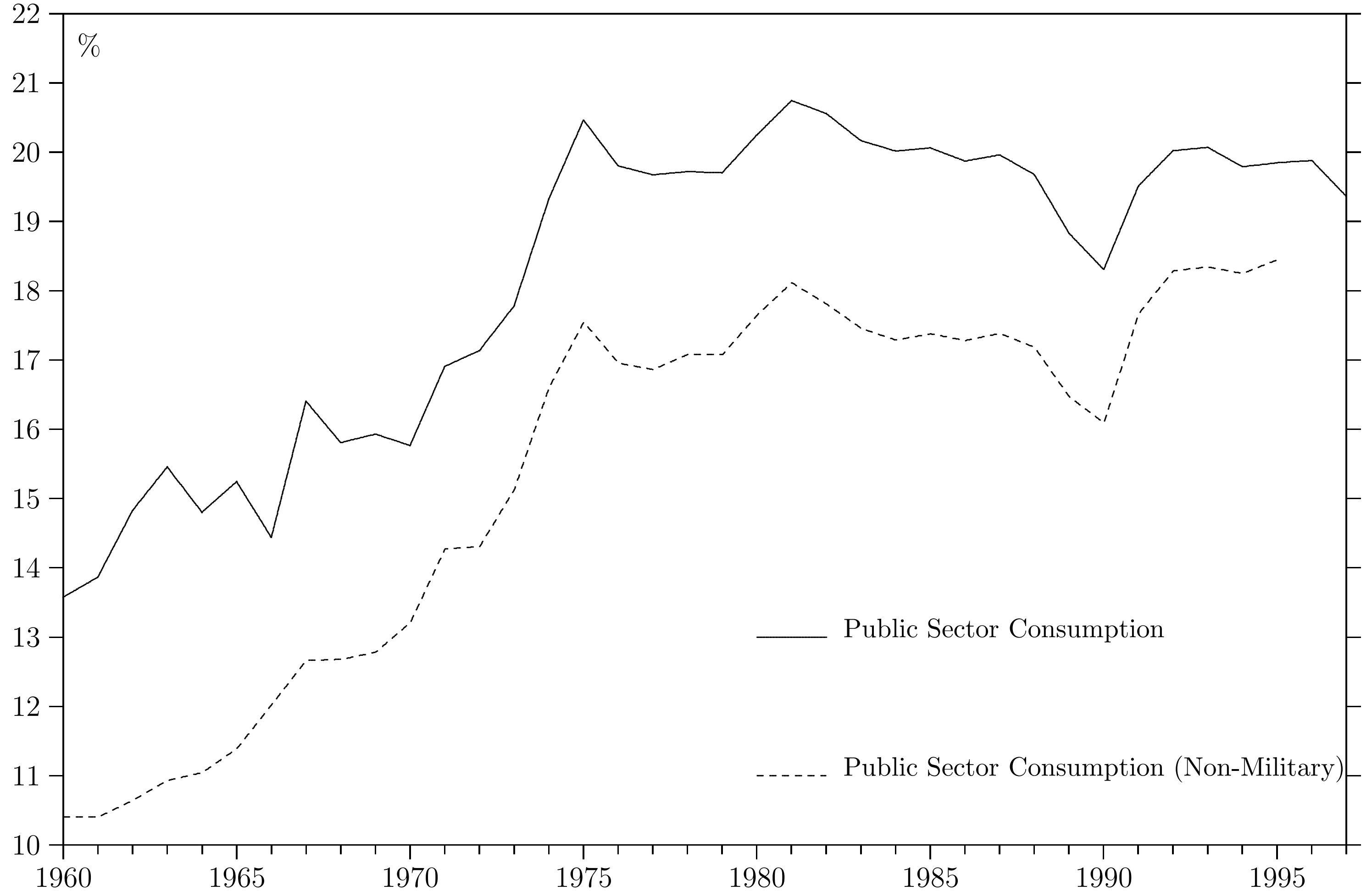
Our investment data includes investments undertaken by the Federal, State, and local governments. The investment figures refer to construction of new roads as well as maintainance costs and extensions to existing roads. To check the consistency of the gathered data, we have compared our regional investment data with the aggregate investment figures from the DIW. On the average, our aggregated investment figures are about 8 percent smaller than the national figures from the DIW. We didn't find an explanation for this. However, we assume, that the aggregate figures from the DIW are more reliable than the regional figures.

From these regional investment data we have calculated regional road infrastructure stocks using the Perpetual Inventory Method. Since we require the regional estimates to be consistent with the aggregate capital stock estimate from the DIW, we have used a restriction in the Perpetual Inventory Method to achieve the value of the aggregated infrastructure stock from the DIW. Applying a goal seeking analysis, we have determined for every year a (positive) depreciation rate so that our aggregated regional stock equals the national stock reported by the DIW. (To make this tractable, we had to assume that the depreciation rate is the same for each region in that year). While the implicit depreciation rate from the DIW figures for the national stock is on average about 2.5 percent, our depreciation rates range between 0.5 to 1.5 percent. Moreover, starting values for the Perpetual Inventory Method for regional capital stocks have been obtained from the publication "Regionale Verkehrsinfrastruktur in der BRD" by Bernd Bartholmai, Heft 26, 1973. This publication contains regional estimates for regional road infrastructure stocks at the level of the Bundesländer for the year 1970.

To summarize our method, we have obtained these estimates for the regional stocks by the applying Perpetual Inventory Method, and we haven chosen a depreciation rate that ensures that the aggregated value of our regional stocks equals the aggregate stock from DIW for every year.

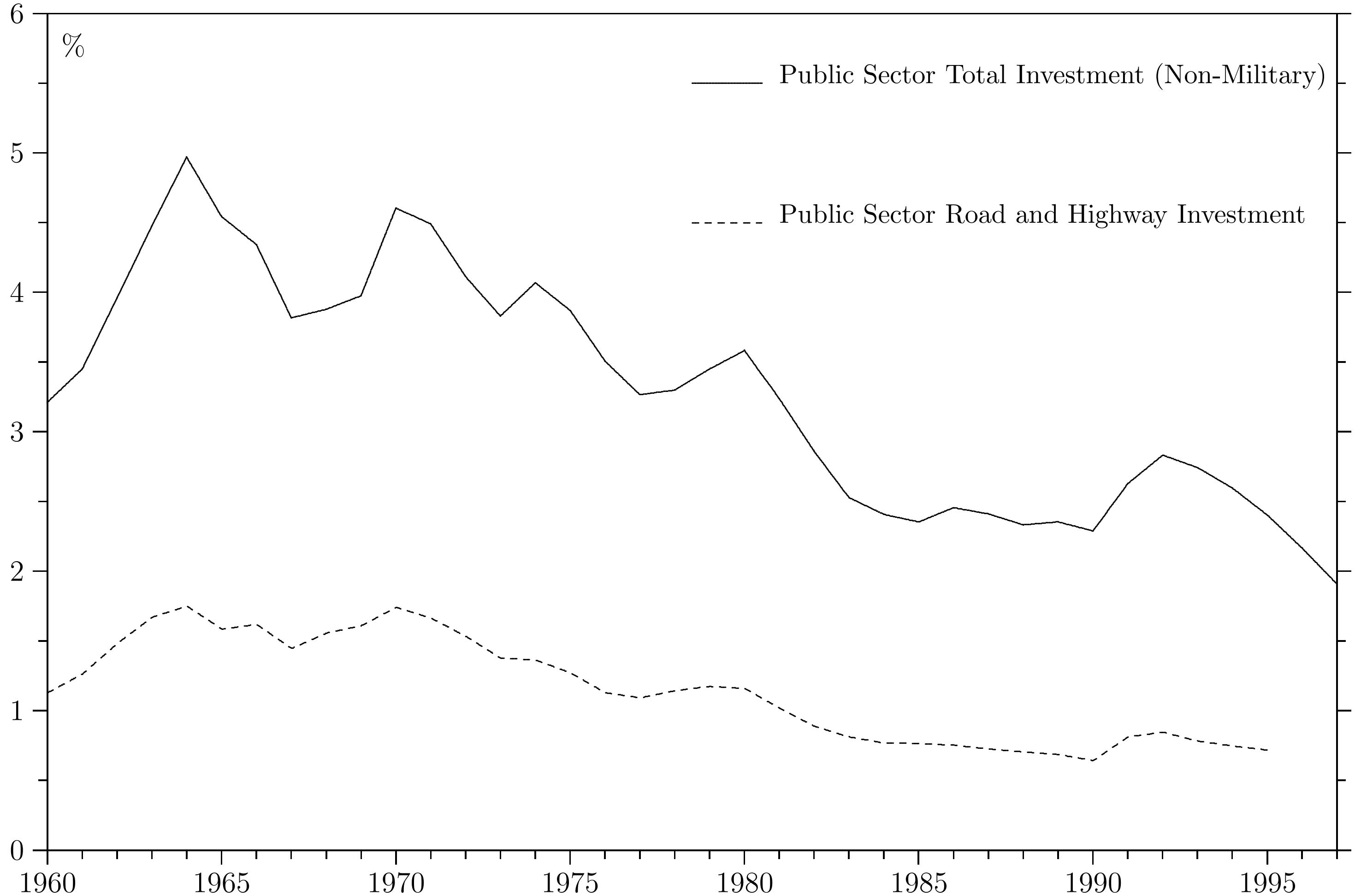
B Tables and Figures

Figure 1: Public Sector Consumption in the Federal Republic of Germany from 1960-1997 as a Share of GDP [%]



(Description of Data in Appendix A, after 1990 for both East and West Germany)

Figure 2: Public Investment in the Federal Republic of Germany from 1960-1997 as a Share of GDP [%]



(Description of Data in Appendix A, after 1990 for both East and West Germany)

Table 1: Production Function Studies at the Regional Level for the US

Study	**Data**	**Specification**	**Results, output elasticity ϵ_{YG}**
Kelejian and Robinson (1997)	US, panel, 48 states, 1970-1986, public capital stock from Munnell (1990a)[a]	Cobb-Douglas, spatial correlation, AR(1), heterosc., spillover-effects	not significant, spatial correlation model
Garcia-Milà, McGuire, and Porter (1996)	US, panel, 48 states, 1970-1983, public capital stock from Munnell (1990a)[a]	Cobb-Douglas, fixed and random state effects, 1. diffs	not significant, revised estimation from Garcia-Milà and McGuire (1992)
Moonaw, Mullen, and Williams (1995)	US, panel, 48 states, 1970, 1980, 1986, public capital stock from Munnell (1990a)[a]	translog	ϵ_{YG}=0.11, highways, water & sewer systems significant, other types not
Baltagi and Pinnoi (1995)	US, panel, 48 states, 1970-1986, public capital stock from Munnell (1990a)[a]	Cobb-Douglas, fixed and random state effects, IV estimation	highways not significant, water & sewer significant
Pinnoi (1994)	US, panel, 48 states, 4 industries, 1970-1986, public capital stock from Munnell (1990a)[a]	translog, fixed and random state effects	water & sewer systems, other types negative, standard errors not reported
Evans and Karras (1994a)	US, panel, 48 states, 1970-1986, public capital stock from Munnell (1990a)[a]	Cobb-Douglas, fixed and random state effects, AR(1), heterosc.	not significant
Holtz-Eakin (1994)	US, panel, 48 states, 1970-1986, revised public capital from Munnell (1990a)	Cobb-Douglas, fixed & random state, time effects, IV estim.	not significant
Munnell (1993)	US, panel, 48 states, diff. industries, 1970-1990, revised data[a] for public capital	Cobb-Douglas	ϵ_{YG}=0.14

[a] Three different types of public capital: 1. Highways, 2. Water and sewer systems, 3. Other types (primarily buildings)

Table 1: (cont.) Production Function Studies at the Regional Level for the US

Study	**Data**	**Specification**	**Results, output elasticity ϵ_{YG}**
Garcia-Milà and McGuire (1992)	US, panel, 48 states, 1969-83, highway capital, education expenditures	Cobb-Douglas, time effects	ϵ_{YG}=0.04 for highways, education significant
Carlino and Voith (1992)	US, panel, 48 states, 1963-1986, highway density, educational attainment	CES, fixed and random effects	ϵ_{YG}=0.22-1.00 for highways, education significant
Moonaw and Williams (1991)	US, panel, 48 states, manufacturing, 1959-76, highway density	TFP growth	ϵ_{YG}=0.17
Hulten and Schwab (1991)	US, panel, manufacturing, 9 regions, public capital stock from Munnell (1990a)[a]	TFP growth, time effects	not significant
Munnell (1990a)	US, panel, 48 states, 1970-1986, capital outlays from *Goverment Finances*[a]	Cobb-Douglas, translog	ϵ_{YG}=0.16 for Cobb-Douglas, public capital and private capital substitutes
Da Silva Costa, Ellson, and Martin (1987)	US, cross-section, 48 states, 1972, capital outlays from *Goverment Finances*	translog	ϵ_{YG}=0.19-0.26, labor and public capital complementary, diminishing returns in public capital
Eberts (1986)	US, panel, 38 SMSA, 1958-78, public capital stock metropolitan area	translog	0.03

[a] Three different types of public capital: 1. Highways, 2. Water and sewer systems, 3. Other types (primarily buildings)

Table 2: (cont.) Production Function Studies at the Regional Level for Other Countries

Study	**Data**	**Specification**	**Results, output elasticity ϵ_{YG}**
Prud'homme (1996)	France, panel, 21 regions, 1970-1990, transportation infrastructure	Cobb-Douglas, TFP growth	ϵ_{YG}=0.08
Hofmann (1995), (1996)	Germany, Hamburg, time-series, 1970-1992, locally provided infrastructure services	Cobb-Douglas, 1. diffs., error correction model	not significant or implausible (negativ)
Seitz (1995)	Germany, panel, 99 cities, 1980-89, public capital	Cobb-Douglas, translog	ϵ_{YG}=0.08 to 0.19
De La Fuente and Vives (1995)	Spain, panel, 17 regions, 1981, 1986, 1990, transportation infrastructure, education	Cobb-Douglas, translog, time effects	ϵ_{YG}=0.21
Picci (1995)	Italy, panel, 20 regions, 1970-1991, public capital stock	Cobb-Douglas, fixed and random effects, 1. diffs.	ϵ_{YG}=0.08-0.43, short-run effects, long-run effects not signif.
Merriman (1990)	48 US states, 1972, 9 Japanese regions, 1954-63, public capital from Da Silva Costa, Ellson, and Martin (1987) and Mera (1973)	translog, fixed-effects, SUR estimation	ϵ_{YG}=0.46-0.58 for Japan, ϵ_{YG}=0.20 for US
Mera (1973)	Japan, panel, 9 regions, 1954-63, social overhead capital	Cobb-Douglas	ϵ_{YG}=0.12-0.22

Table 3: Production Function Studies at the National Level

Study	**Data**	**Specification**	**Results, output elasticity** ϵ_{YG}
Denny and Guiomard (1997)	Irland, time-series, manufacturing, 1951-1994, stock of roads & highways	Cobb-Douglas, AR(1)	ϵ_{YG}=0.92
Fernald (1993)	US, time-series, 35 sectors, 1948-1985, stock of roads & highways	TFP growth	significant, explains half of the observed decline in productivity growth
Christodoulakis (1993)	Greece, time-series, manufacturing, 1963-1990, public infrastructure (roads, railways, electricity, communication, etc.)	Cobb-Douglas, cointegration	ϵ_{YG}=0.27-0.42
Bajo-Rubio and Sosvilla-Rivero (1993)	Spain, time-series, 1964-88, public capital	Cobb-Douglas, cointegration, Hausman exogeneity test	ϵ_{YG}=0.18, public capital exogenous
Berndt and Hansson (1992)	Sweden, time-series, 1964-88, public infrastructure	Cobb-Douglas	ϵ_{YG}=-1.66-0.369, results implausible
Tatom (1991a)	US, time series, 1949-85, public capital data from Aschauer (1989a)	Cobb-Douglas, including energy prices, 1. diffs.	not significant
Munnell (1990b)	US, time series, 1948-87, public capital	Cobb-Douglas	ϵ_{YG}=0.34-0.37
Aschauer (1989a)	US, time series, 1949-85, non-military public capital	Cobb-Douglas	ϵ_{YG}=0.38-0.56
Ratner (1983)	US, time series, 1949-73, non-military public capital	Cobb-Douglas	ϵ_{YG}=0.06

Table 4: Production Function Studies at the International Level

Study	**Data**	**Specification**	**Results, output elasticity** ϵ_{YG}
Aschauer (1995)	OECD, 12 countries, panel, infrastructure capital from Ford and Poret (1991)	TFP growth, fixed country and time effects, 4-year average	ϵ_{YG}=0.33-0.55
Nourzad and Vrieze (1995)	OECD, 7 countries, panel, 1963-88, public investment (data sources not given)	Cobb-Douglas, energy input, 1. diffs., random effects	ϵ_{YG}=0.05
Evans and Karras (1994b)	OECD, 7 countries, panel, 1963-88, public capital	Cobb-Douglas, 1. diffs.	not significant
Neusser (1993)	G7 countries, manufacturing, 1970-87, public capital from Ford and Poret (1991)	TFP growth, cointegration techniques, long-run effects	unstable and unreliable results
Taylor-Lewis (1993)	G7-countries, panel, sector specific, public capital from Ford and Poret (1991), indicators of physical infrastructure	Cobb-Douglas	not significant
Ford and Poret (1991)	OECD, 12 countries, time series, 1960-1988, non-military public capital stock, broad definition includes also privatly provided infrastructure services	TFP growth, AR(1), AR(2)	only significant for US, Germany, Canada, Belgium and Sweden
Aschauer (1989c)	G7-countries, panel data, 1966-85, public investments from OECD national accounts	Cobb-Douglas	ϵ_{YG}=0.34-0.73

Table 5: Cross-Sectional Correlation and Variance/Covariance Matrix for the Bundesländer Based on Residuals from Equation (19)

	BaW	Bay	Ber	Bre	Ham	Hes	Nie	NRW	RhP	Saa	SHo
BaW	**4.68**	3.77	1.68	4.44	6.51	4.40	2.79	3.53	4.12	3.55	5.48
Bay	0.74	**5.62**	0.75	3.51	5.85	4.94	3.92	3.23	4.65	4.44	1.67
Ber	0.18	0.07	**19.39**	0.8	11.03	3.40	5.41	3.48	6.88	3.59	5.77
Bre	0.45	0.32	0.04	**21.04**	-6.46	8.21	6.09	3.91	8.76	8.28	-3.93
Ham	0.43	0.35	0.35	-0.20	**50.15**	2.47	4.51	8.67	7.09	6.87	21.03
Hes	0.69	0.71	0.26	0.61	0.12	**80.73**	6.29	4.36	8.28	5.71	2.83
Nie	0.36	0.46	0.34	0.37	0.18	0.59	**13.01**	3.61	5.40	2.95	-0.83
NRW	0.76	0.63	0.37	0.40	0.57	0.68	0.46	**4.65**	5.19	4.78	5.45
RhP	0.54	0.56	0.44	0.54	0.28	0.79	0.42	0.68	**12.47**	6.90	3.19
Saa	0.48	0.54	0.24	0.52	0.28	0.56	0.24	0.64	0.57	**11.88**	1.46
SHo	0.45	0.13	0.23	-0.15	0.53	0.17	-0.04	0.45	0.16	0.08	**31.58**

BaW=Baden-Würtemberg, Bay=Bayern, Ber=Berlin, Bre=Bremen, Ham=Hamburg, Hes=Hessen, Nie=Niedersachsen, NRW=Nordrhein-Westfalen, RhP=Rheinland-Pfalz, Saa=Saarland, Sho=Schleswig-Holstein

Table 6: Cobb-Douglas Production Function Estimates (G=11, T=23)

(1) Factors of Production $\mathbf{g}, \mathbf{k}, \mathbf{l}$

	AR(1), ρ=0.778				*Vars. in first diffs.*			
	Kmenta (Het., Corr.)		**OLS** (PCSE)		**Kmenta** (Het., Corr.)		**OLS** (PCSE)	
const	Dum.*	—	Dum.*	—	0.018*	(0.005)	0.010	(0.008)
t	0.009*	(0.003)	0.005	(0.006)	—	—	—	—
$\ln g$	1.045*	(0.125)	1.130*	(0.393)	0.402*	(0.119)	0.607*	(0.219)
$\ln k$	0.080	(0.051)	0.123	(0.145)	0.045	(0.061)	0.019	(0.169)
$\ln l$	0.622*	(0.053)	0.622*	(0.111)	0.613*	(0.055)	0.595*	(0.110)
Länder-	F test:		Wald χ^2:		F test:		Wald χ^2:	
effects	159.1*		1602.9*		—		—	
CRS g, k, l[a]	27.2*		4.26*		62.5*		0.59	
CRS k, l[b]	23.2*		2.33		83.2*		3.52	
N	253		253		253		253	
R^2	0.99971		0.99973		0.383		0.312	
SE	-0.001		-0.001		-0.104		-0.104	
DW	1.955		1.955		2.137		2.137	

(2) Factors of Production $\mathbf{k}, \mathbf{l}$

	AR(1), ρ=0.857				*Vars. in first diffs.*			
	Kmenta (Het., Corr.)		**OLS** (PCSE)		**Kmenta** (Het., Corr.)		**OLS** (PCSE)	
const	Dum.*	—	Dum.*	—	0.023*	(0.004)	0.020*	(0.007)
t	0.015*	(0.004)	0.011*	(0.005)	—	—	—	—
$\ln k$	0.052	(0.059)	0.031	(0.164)	0.183*	(0.060)	0.214	(0.186)
$\ln l$	0.624*	(0.061)	0.634*	(0.109)	0.559*	(0.059)	0.559*	(0.120)
Länder-	F test:		Wald test χ^2:		F test:		Wald test χ^2:	
effects	57.7*		591.9*		—		—	
CRS k, l[c]	18.6*		3.01		60.4*		1.0	
N	253		253		253		253	
R^2	0.99916		0.99928		0.332		0.224	
SE	-0.059		-0.059		-0.016		-0.016	
DW	2.049		2.049		1.957		1.957	

* significant at a 5 % level, standard errors are given in parentheses

[a] degrees of freedom, AR(1): $F(1 \mid 239)$, Wald: $\chi^2(1)$, 1.diffs: $F(1 \mid 249)$, Wald: $\chi^2(1)$

[b] degrees of freedom, AR(1): $F(1 \mid 239)$, Wald: $\chi^2(1)$, 1.diffs: $F(1 \mid 249)$, Wald: $\chi^2(1)$

[c] degrees of freedom, AR(1): $F(1 \mid 240)$, Wald: $\chi^2(1)$, 1.diffs: $F(1 \mid 250)$, Wald: $\chi^2(1)$

Table 7: Results for TSCS and Panel Data Models

	Model	$\dot{g}$	*const*	*dum.-var.*	R^2	Hypothesis Tests
	Dependent variable: Total factor productivity growth $\dot{\text{TFP}}$, (G=11, T=23)					
(1)	**OLS**	0.366* (0.122)	0.005 (0.004)	—	0.034	Model: $F(1, 251) = 8.94^*$
(2)	**OLS** (PCSE)	0.366 (0.231)	0.005 (0.008)	—	0.034	Heteroscedasticity LM $\chi^2(10) = 62.9^*$
(3)	**Kmenta** (Het,Corr)	0.325* (0.128)	0.012* (0.005)	—	—	Cross-sect. corr. λ_{LM}: $\chi^2(55) = 297.3^*$
(4)	**Fixed Effects**	0.957* (0.280)	-0.032* (0.011)	Time	0.404	Time: $F(22, 229) = 6.46^*$
(5)	**Random Effects**	0.616* (0.198)	-0.001 (0.008)	Time	0.037	Hausman:$\chi^2(1) = 2.94$
(6)	**Fixed Effects**	0.755* (0.355)	-0.026* (0.013)	Time & Group	0.439	Time: $F(22, 219) = 6.40^*$ Group: $F(10, 219) = 1.34$
(7)	**Random Effects**	0.580* (0.201)	0.001 (0.008)	Time & Group	0.032	Hausman:$\chi^2(1) = 0.36$

* significant at a 5 % level, standard errors are given in parentheses

Table 8: Translog Production Function Estimates (G=11, T=23)

	AR(1), ρ=0.714				*Vars. in first diffs.*			
	Kmenta (Het., Corr.)		**OLS** (PCSE)		**Kmenta** (Het., Corr.)		**OLS** (PCSE)	
const	Dum.*	—	Dum.*	—	0.038*	(0.015)	0.037	(0.029)
t	0.025	(0.014)	0.022	(0.022)	—	—	—	—
$0.5\ t^2$	-0.001	(0.001)	0.001	(0.001)	-0.002	(0.001)	-0.001	(0.002)
$\ln g$	0.558*	(0.210)	0.473	(0.382)	0.175	(0.183)	0.275	(0.375)
$\ln k$	0.042	(0.107)	-0.068	(0.198)	0.074	(0.102)	-0.029	(0.187)
$\ln l$	0.653*	(0.106)	0.885*	(0.218)	0.624*	(0.084)	0.781*	(0.186)
$0.5 \ln^2 G$	-0.427*	(0.176)	-0.358	(0.298)	-0.269	(0.169)	0.181	(0.308)
$0.5 \ln^2 K$	0.167	(0.183)	0.617	(0.398)	-0.033	(0.176)	0.549	(0.400)
$0.5 \ln^2 L$	0.299	(0.193)	0.604	(0.480)	0.193	(0.147)	0.457	(0.421)
$\ln g \ln k$	0.024	(0.135)	-0.300	(0.269)	0.135	(0.145)	-0.301	(0.301)
$\ln g \ln l$	-0.045	(0.100)	-0.071	(0.187)	-0.040	(0.080)	-0.044	(0.143)
$\ln k \ln l$	-0.188	(0.159)	-0.423	(0.388)	-0.048	(0.127)	-0.256	(0.358)
$t \ln g$	0.006	(0.004)	0.006	(0.006)	0.006	(0.006)	-0.001	(0.010)
$t \ln k$	-0.006	(0.006)	-0.005	(0.014)	-0.001	(0.007)	0.001	(0.016)
$t \ln l$	0.010	(0.007)	0.014	(0.016)	0.003	(0.007)	0.011	(0.019)
	F test:		Wald χ^2:		F test:		Wald χ^2:	
Länder-effects[a]	19.3*		87.6*		—		—	
Linear homogeneity[b]	10.9*		31.0*		4.1*		13.9*	
Translog vs. Cobb-Douglas[c]	5.1*		39.8*		2.8*		16.5*	
N	253		253		253		253	
R^2	0.9835		0.9955		0.4653		0.3477	
SE	0.033		0.033		-0.091		-0.091	
DW	1.886		1.886		2.124		2.124	

* significant at a 5 % level, standard errors are given in parentheses

[a] degrees of freedom, AR(1): $F(10 \mid 228)$, Wald: $\chi^2(10)$

[b] degrees of freedom, AR(1): $F(4 \mid 228)$, Wald: $\chi^2(4)$, 1.diffs: $F(4 \mid 240)$, Wald: $\chi^2(4)$

[c] degrees of freedom, AR(1): $F(10 \mid 228)$, Wald: $\chi^2(10)$, 1.diffs: $F(10 \mid 240)$, Wald: $\chi^2(10)$

Table 9: Tests Based on the Translog Production Function Estimates

	df	$AR(1)$, ρ=0.714		*Vars. in first diffs.*	
		Kmenta	**OLS**	**Kmenta**	**OLS**
	$AR(1)$|*1.diffs.*	(Het., Corr.)	(PCSE)	(Het., Corr.)	(PCSE)
Test for Factor		F test:	Wald test χ^2:	F test:	Wald test χ^2:
g	5|5	6.8*	26.8*	2.6*	6.2
k	5|5	0.9	5.7	1.1	4.1
l	5|5	22.3*	62.3*	26.8*	63.3*
t	5|4	9.6*	36.5*	4.1*	12.0*
Test for Effects		F test:	Wald test χ^2:	F test:	Wald test χ^2:
Linear	4|3	31.9*	41.6*	30.0*	21.7*
Cross	6|6	6.5*	28.9*	3.1*	13.7*
Quadratic	4|4	2.6*	6.6	2.3	3.5

* significant at a 5 % level

Table 10: Marginal Productivities and Allen Partial Elasticities of Substitution Based on Translog Function Estimates of Table 8, Column 1

	Marginal Productivity			Elasticity of Substitution		
Land	f_g	f_k	f_s	σ_{gk}	σ_{gs}	σ_{ks}
Baden-Würtemberg	0.443	0.084	0.729	1.099	0.582	-1.980
Bayern	0.348	0.076	0.729	1.017	0.564	-1.534
Berlin	0.836	0.091	0.759	1.255	0.764	-2.920
Bremen	0.875	0.073	0.701	1.243	0.829	-2.792
Hamburg	0.849	0.097	0.800	1.258	0.773	-2.952
Hessen	0.437	0.077	0.721	1.096	0.646	-1.847
Niedersachsen	0.386	0.076	0.693	1.080	0.619	-1.800
Nordrhein-Westfalen	0.385	0.073	0.731	1.011	0.559	-1.498
Rheinland-Pfalz	0.445	0.086	0.700	1.155	0.653	-2.252
Saarland	0.634	0.061	0.609	1.198	0.792	-2.384
Schleswig-Holstein	0.482	0.084	0.699	1.178	0.711	-2.346

Figure 3: Between Years Regression

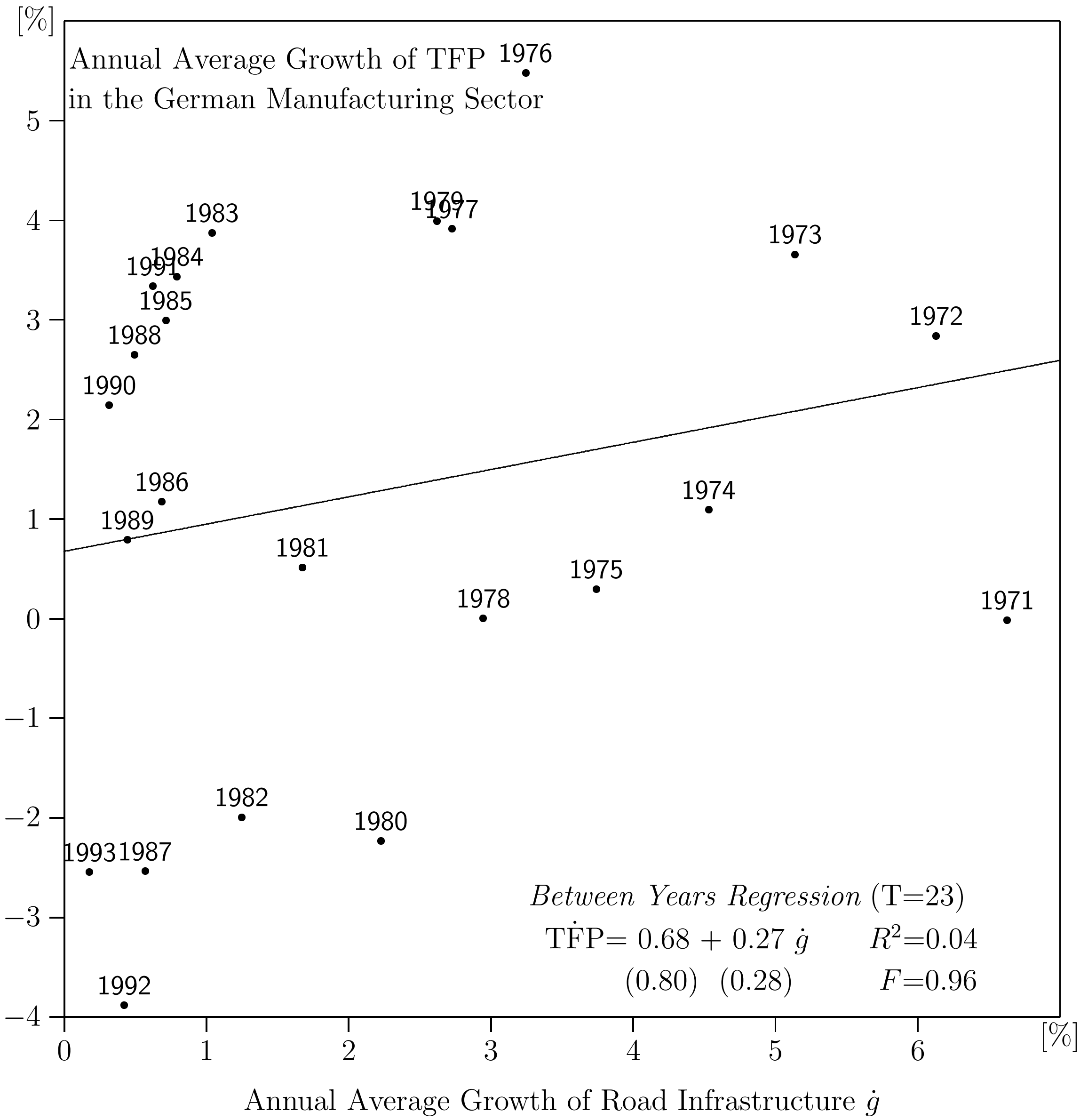
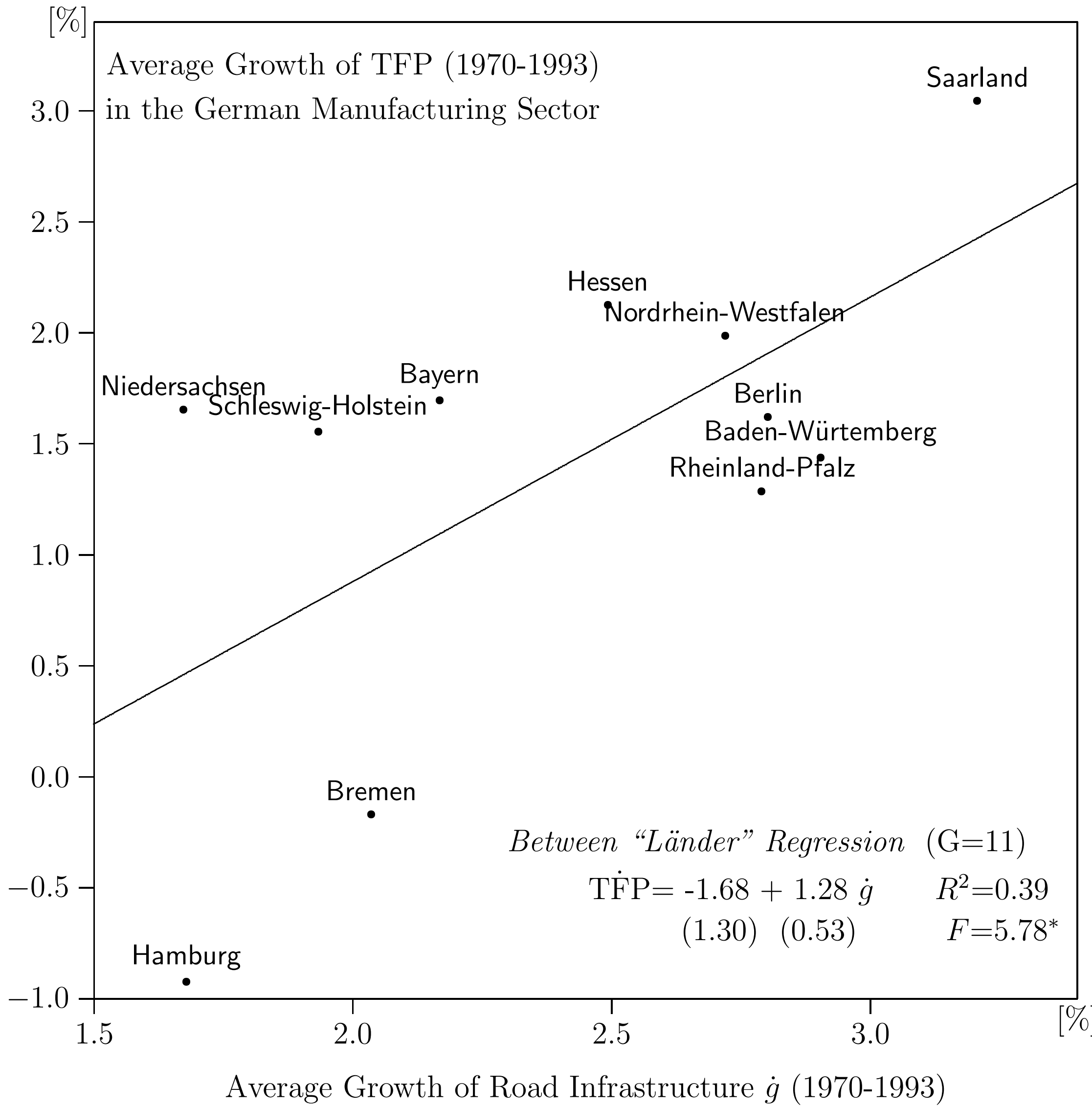


Figure 4: Between Bundesländer Regression



C A Brief Description of the Kmenta Method

The "Kmenta"-method (Kmenta and Oberhofer 1974, Kmenta 1986), which is sometimes also referred as the "Parks" method (Parks 1967), is based on Generalized Least Squares and can correct for temporally correlated errors and panel heteroscedasticity as well as contemporaneously correlations across individuals. Thus, it is assumed that the random errors u_{it}, $i = 1, 2, \ldots, N$, $t = 1, 2, \ldots, T$ have the structure

$$\begin{aligned}
\mathrm{E}(u_{it}^2) &= \sigma_{ii} && \text{(groupwise heteroscedasticity)} \\
\mathrm{E}(u_{it}u_{jt}) &= \sigma_{ij} && \text{(cross group correlation)} \\
u_{it} &= \rho_i u_{i,t-1} + \epsilon_{ij} && \text{(within group autocorrelation)}
\end{aligned}$$

where

$$\begin{aligned}
\mathrm{E}(\epsilon_{it}) &= 0 \\
\mathrm{E}(u_{it}\epsilon_{it}) &= 0 \\
\mathrm{E}(\epsilon_{it}\epsilon_{jt}) &= \Phi_{ij} \\
\mathrm{E}(\epsilon_{it}\epsilon_{js}) &= 0 && (s \neq t) \\
\mathrm{E}(u_{i0}) &= 0 \\
\mathrm{E}(u_{i0}u_{j0}) &= \sigma_{ij} = \Phi_{ij}/(1 - \rho_i\rho_j)
\end{aligned}$$

(Note, that in our empirical analysis we have assumed a common ρ for all groups).

In this model, the covariance matrix for the vector of random errors $\mathbf{u}$ can be expressed (Greene 1993, pp.447,457)

$$\mathrm{E}\left(\mathbf{u}\mathbf{u}'\right) = \mathbf{V} = \begin{bmatrix} \sigma_{11}\mathbf{P}_{11} & \sigma_{12}\mathbf{P}_{12} & \ldots & \sigma_{1N}\mathbf{P}_{1N} \\ \sigma_{21}\mathbf{P}_{21} & \sigma_{22}\mathbf{P}_{22} & \ldots & \sigma_{2N}\mathbf{P}_{2N} \\ \vdots & \vdots & \vdots & \vdots \\ \sigma_{N1}\mathbf{P}_{N1} & \sigma_{N2}\mathbf{P}_{N2} & \ldots & \sigma_{NN}\mathbf{P}_{NN} \end{bmatrix} \tag{27}$$

where

$$\mathbf{P}_{ij} = \begin{bmatrix} 1 & \rho_j & \rho_j^2 & \ldots & \rho_j^{T-1} \\ \rho_i & 1 & \rho_j & \ldots & \rho_j^{T-2} \\ \rho_i^2 & \rho_i & 1 & \ldots & \rho_j^{T-3} \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ \rho_i^{T-1} & \rho_i^{T-2} & \rho_i^{T-3} & \ldots & 1 \end{bmatrix}. \tag{28}$$

The matrix $\mathbf{V}$ is estimated by a two-stage procedure, and β is then estimated by Feasible Generalized Least Squares (FGLS). The first step in estimating $\mathbf{V}$ involves the use of ordinary least squares to estimate β and obtaining the fitted residuals, as follows:

$$\hat{\mathbf{u}} = \mathbf{y} - \mathbf{X}\widehat{\beta}_{OLS}. \tag{29}$$

A consistent estimator of the first-order autoregressive parameter is then obtained in the usual manner, as follows:

$$\widehat{\rho}_i = \left(\sum_{t=2}^{T} \widehat{u}_{it}\widehat{u}_{i,t-1}\right) \Big/ \left(\sum_{t=2}^{T} \widehat{u}_{i,t-1}^2\right) \qquad i = 1, 2, \ldots, N. \tag{30}$$

Finally, the data's autoregressive characteristics can be removed (asymptotically) by the usual transformation of taking weighted differences. That is, for $i = 1, 2, \ldots, N$, and $t = 2, \ldots, T$

$$y_{it} - \widehat{\rho}_i y_{i,t-1} = \sum_{k=1}^{p} \left(X_{itk} - \widehat{\rho}_i X_{i,t-1}\right) \beta_k + u_{it} - \widehat{\rho}_i u_{i,t-1} \tag{31}$$

while for $i = 1, 2, \ldots, N$, and $t = 1$

$$y_{it}\sqrt{1 - \widehat{\rho}_i^2} = \sum_{k=1}^{p} X_{i1k}\sqrt{1 - \widehat{\rho}_i^2}\beta_k + u_{it}\sqrt{1 - \widehat{\rho}_i^2}. \tag{32}$$

This system can be written as

$$y_{it}^* = \sum_{k=1}^{p} X_{itk}^* \beta_k + u_{it}^*. \tag{33}$$

The second step in estimating the covariance matrix V is to apply ordinary least squares to the preceding transformed model, obtaining

$$\widehat{\mathbf{u}}^* = \mathbf{y}^* - \mathbf{X}^* \widehat{\beta}_{OLS}^* \tag{34}$$

from which the consistent estimator of σ_{ij} is calculated:

$$s_{ij} = \frac{\widehat{\Phi}_{ij}}{(1 - \widehat{\rho}_i \widehat{\rho}_j)} \tag{35}$$

where

$$\widehat{\Phi}_{ij} = \frac{1}{(T-p)} \sum_{t=1}^{T} \widehat{u}_{it}^* \widehat{u}_{jt}^*. \tag{36}$$

FGLS then proceeds in the usual manner,

$$\widehat{\beta}_{FGLS} = (\mathbf{X}'\widehat{\mathbf{V}}^{-1}\mathbf{X})^{-1}\mathbf{X}'\widehat{\mathbf{V}}^{-1}\mathbf{y}. \tag{37}$$

Note, that for models which do not correct for autocorrelation Maximum Likelihood estimates can be obtained by iterating the described procedure to convergence (Kmenta and Oberhofer 1974).

D A Short Motivation of OLS with Panel Corrected Standard Errors (PCSE)

Two recent articles (Beck and Katz 1995, Beck and Katz 1996) argue that ordinary least squares with panel corrected standard errors (PCSE) is superior to the Kmenta generalized least squares approach when estimating time-series cross-section (TSCS) models using small samples. This proposition is based on evidence from Monte-Carlo experiments. Although the estimates of the Kmenta method might be more efficient in terms of root mean square error than OLS/PCSE in situations where cross-sectional correlations and groupwise heteroscedasticity are significant, standard errors obtained from the Kmenta method have a tendency to be too small (in finite samples), thus, they are "overconfident" and do not correctly reflect the "true" sampling variability of the parameter estimates.

While OLS is not efficient in the presence of non-spherical errors, it does yield consistent estimates. OLS standard errors will be inaccurate in the presence of non-sperical errors in that they do not provide good estimates of the sampling variability of the OLS parameter estimates. Panel corrected standard errors are a direct extension of White's (1980) heteroscedasticity-consistent standard errors. However, since PCSEs take into account the panel structure of the data, they perform even better for TSCS data than White's heteroscedasticity-consistent standard errors do.

PCSEs are estimated by the square root of the diagonal of

$$(\mathbf{X}'\mathbf{X})^{-\mathbf{1}}\mathbf{X}\,(\mathbf{\Sigma}\otimes\mathbf{I_T})\,\mathbf{X}(\mathbf{X}'\mathbf{X})^{-\mathbf{1}}, \tag{38}$$

where $\mathbf{\Sigma}$ is a $N \times N$ matrix of cross-sectional variances and covariances. A consistent estimate of Σ is given by $\mathbf{E}'\mathbf{E}/T$, where $\mathbf{E}$ denote $T \times i$ matrix of OLS residuals from equation (34). PCSEs are consistent estimates of the standard errors of $\widehat{\beta}_{OLS}$.