

# **Essays on the Contribution of Public Infrastructure to Private Production and its Political Economy**

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—*To Angelika*—

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# Thesis Abstract

This thesis consists of five self-contained essays on two central subjects. The first subject is the relationship between infrastructure and private productivity. The second subject is the determinants of regional infrastructure investment allocation. Whereas the first and the second essay in this thesis mainly focus on the first subject, i.e. the contribution of infrastructure to private production; the essays three to five study besides the productivity effects also the determinants of infrastructure investment allocation.

In the first essay, using time-series cross-section data from the manufacturing sector of the 11 German 'Bundesländer' (federal states) from 1970 to 1996, we examine the impact of public capital on private production. Our econometric analysis explicitly takes into account four of the most frequent specification issues in the context of time-series cross-section data analysis: serial correlation, groupwise heteroscedasticity, cross-sectional correlation and nonstationarity of data. For all approaches and tested specifications, we find that public capital is a significant input for production in the manufacturing sector. Moreover, we find that differences in public capital endowment can explain long-term differences in productivity across the Bundesländer.

The second essay uses a partial adjustment approach to measure the contribution of public infrastructure to local private production. In the first step of the empirical analysis we apply a principal component analysis in order to construct 2 new infrastructure indicators from an array of 7 measures of transport and human capital infrastructure. In the second step the output of different sectors is regressed on private factor inputs and on these 2 infrastructure indicators. Our main finding is that expected long-run equilibrium output in an area of local government will be higher, the better it is endowed with both transport and human

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capital infrastructure. Moreover, transport and human capital infrastructure appear to be complementary, i.e. raising only transport infrastructure will not yield an increase in private production at the local level.

The third essay proposes a simultaneous-equation approach to the estimation of the contribution of transport infrastructure accumulation to regional growth. We model explicitly the political-economy process driving infrastructure investments; in doing so, we eliminate a potential source of bias in production-function estimates and generate testable hypotheses on the forces that shape infrastructure policy. Our empirical findings on a panel of France's regions over 1984-91 suggest that influence activities were, indeed, significant determinants of the cross-regional allocation of transportation infrastructure investments. Moreover, we find little evidence of concern for the maximisation of economic returns to infrastructure spending, even after controlling for pork-barrel and when imposing an exogenous preference for convergence in regional productivity levels.

The fourth essay applies a simultaneous-equation approach to the estimation of the contribution of infrastructure accumulation to private production. A political economy model for the allocation of public infrastructure investment grants is formulated. Our empirical findings, using a panel of large German cities for the years 1980, 1986, and 1988, suggest that cities ruled by a council sharing the State ('Bundesland') government's current political affiliation were particularly successful in attracting infrastructure investment grants. With regard to the contribution of infrastructure accumulation to growth, we find that public capital is a significant factor for private production. Moreover, at least for the sample studied, we find that simultaneity between output and public capital is weak; thus, feedback effects from output to infrastructure are negligible.

The fifth essay describes the different institutional frameworks for infrastructure policy in Germany and France. The economic effects of infrastructure are estimated econometrically for German and French regions. We find evidence that regional road infrastructure has a significant impact on regional output. Moreover, we find evidence that for Germany the priority of promoting equal living conditions throughout the regions is an important determinant of regional infrastructure policy.

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The bottom line of our research is that throughout the essays evidence of a positive impact of infrastructure on private productivity is found. This finding holds also for different levels of aggregation; the essays 1, 3 and 5 are based on data at the regional level of the Bundesländer and the French regions respectively, whereas essay 2 is based on data at the local level of the German counties and essay 4 is based on data at the local level of large self-administrated German cities. Furthermore, we find only little evidence throughout the studies for a simultaneity between output and infrastructure investment. Thus, reverse causality running from output to infrastructure investment appears—at least for the various samples studied here—not to be significant. Regarding the determinants of infrastructure investment, we find that lobbying and political affiliation matter for the regional allocation of infrastructure investment, whereas expected returns to infrastructure investment do not seem to have an impact.

# Part I

## 1.1 Introduction

It is conventional wisdom that investment in infrastructure capital is a necessary condition for long-run growth in industrial countries and, *a fortiori*, in developing ones. *How much* infrastructure investment actually contributes to private productivity is still, in spite of a long-standing debate, a largely unsettled question.

Economists have long considered public capital to play an important role in regional economic development. Hirschmann (1958), for instance, defined ‘social overhead capital’ (SOC) as comprising of those basic services without which primary, secondary, and tertiary productive activities cannot function. Moreover, according to Hirschmann’s definition the services are provided in practically all countries by public agencies or by private agencies subject to some public control.

Similarly, Hansen (1965) divided public capital<sup>1</sup> into two groups: ‘social’ overhead capital (SOC) and ‘economic’ overhead capital (EOC). While SOC is devoted to investment in human capital like education and health, EOC is primarily geared toward promoting directly productive activities. Examples for EOC are highways and streets, gas and electricity facility, water and sewer systems.

Hansen advanced the theory that the potential effectiveness of public capital investment will vary across types of regions: congested, intermediate and lagging. While the potential benefits of investments in EOC are large in intermediate regions, investment in EOC in lagging regions would have little impact on economic activity. In congested regions, any marginal social benefit that might accrue from further infrastructure investment would be outweighed by the marginal social cost of pollution and congestion resulting from increased economic activity.

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.<sup>2</sup> Con-

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<sup>1</sup>Throughout this thesis we use the terms ‘infrastructure’ and ‘public capital’ interchangeably. However, strictly speaking public capital refers to infrastructure services that are owned by the public sector, whereas ‘infrastructure’ applies also to services which are owned by the private sector.

<sup>2</sup>For an excellent survey on this topic, see Sturm (1998).

sequently, public investment as a share of GDP has declined considerably in most OECD countries during the last two decades.

Figure 1.1 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 1997, while the share of public sector's investment has declined from its peak of about 5 percent in 1964 to about 1.5 percent in 1997. After the German Reunification, public investment as a share of GDP increased for a short period from 2.0 to 2.5 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 East German states ('Neue Bundesländer').

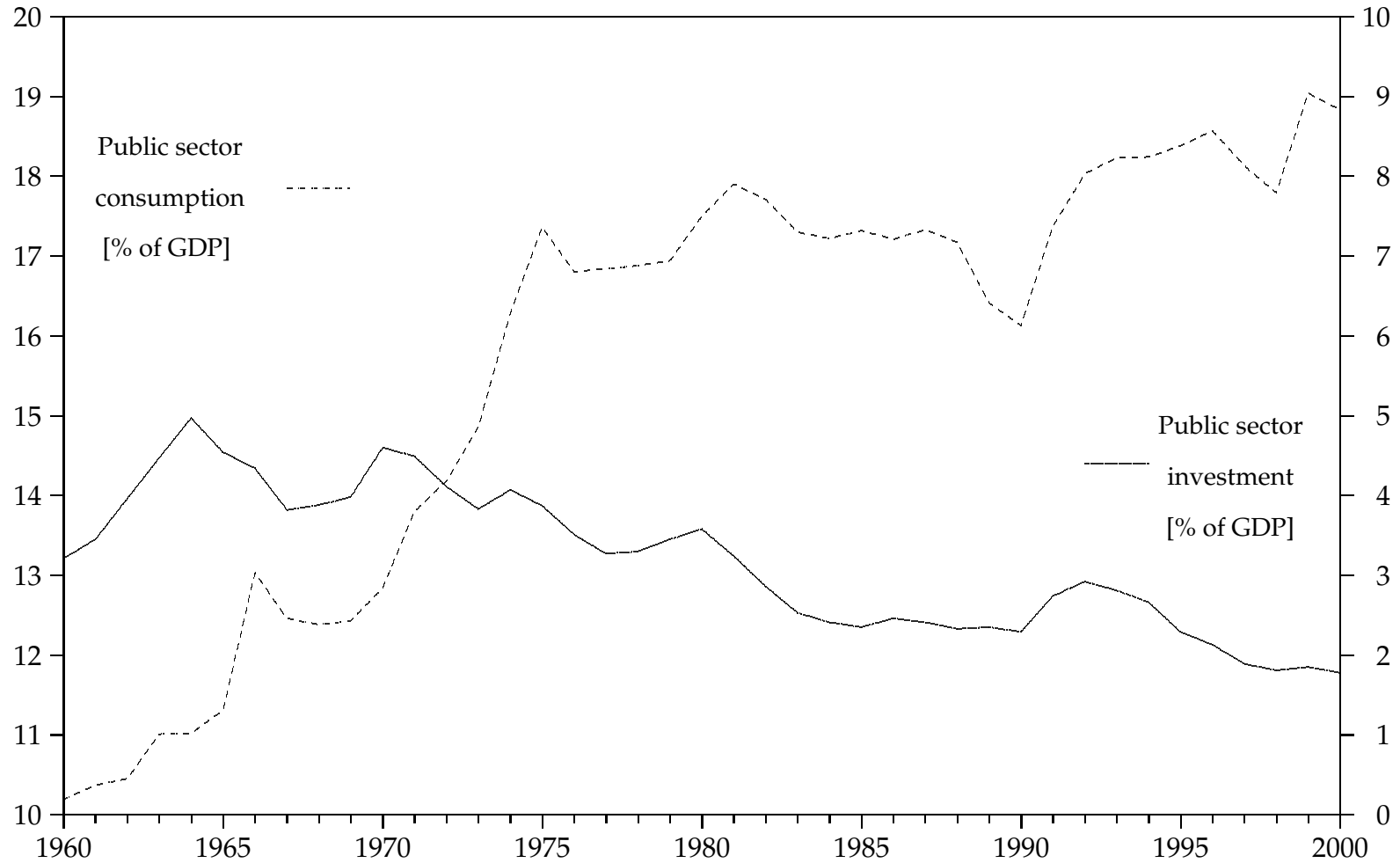
A number of researchers such as Aschauer (1989*a*; 1989*b*; 1989*c*) or Munnell (1990*a*; 1990*c*; 1992) have documented also empirical evidence for a strong correlation between public capital and private sector performance. Furthermore, these authors have hypothesised that the decrease of governmental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

If the 'Public Infrastructure Hypothesis' is of empirical relevance for Germany, then at least a part of the productivity gap between East and West Germany can be attributed to the still existing differences in infrastructure endowments between the East and West German regions. In fact, recent studies (Komar, 2000; Seidel and Vesper, 2000) show that infrastructure endowment in the East German federal states is still only two thirds of that in the West German federal states.

Turning to the hypothetical effects of infrastructure, Aschauer (1995), 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 labour and capital inputs more or less productive. The indirect effect arises because an increase in public capital stock will affect the marginal products of labour and private capital, which in turn influence the chosen quantities of private inputs.

In Germany, the priority of transport infrastructure projects is evaluated on the basis of cost-benefit studies, e.g. the federal transport infrastructure plan

**Figure 1.1:** Public Sector Consumption and Investment in the Federal Republic of Germany from 1960-2000



Source: Federal Statistical Office of Germany, series 18, *National Accounts*  
(after 1990 for both East and West Germany)

(‘Bundesverkehrswegeplanung’). If investment in transport 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 infrastructure 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.<sup>3</sup>

Specifically, these cost-benefit studies do not solely rely on expected economic returns from infrastructure projects, but on environmental impact evaluations as well. However, since a project is not undertaken if it does not yield positive returns, and if the ex-ante evaluations of the returns from transport infrastructure projects are in most cases correct, then in principle it should be possible to find from the ex-post perspective empirical evidence of an impact of infrastructure on private output. This is particularly the case if the empirical analysis aggregates over individual projects, i.e. focus on the average return of transport infrastructure projects.

The purpose of this thesis is twofold. The first aim is to analyse the contribution of infrastructure to private production. The second aim is to examine the politico-economic determinants of regional infrastructure investment allocation. However, these two subjects are not independent but interrelated. By formulating political-economy models which explain the allocation of infrastructure investment to regions, we are able to estimate the productivity effects of infrastructure from structural models in which public capital investment is endogenised. We employ this approach in essays 3 and 4.

Throughout the essays of this thesis we treat infrastructure according to Meade (1952) as an unpaid factor of private production.<sup>4</sup> Hence, infrastructure capital, i.e. EOC, is modelled to enter the production function of private firms.

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<sup>3</sup>This reasoning is, however, based on the assumption that capital is freely transferable between an economy’s private and public sectors.

<sup>4</sup>In contrast to pure public goods (or ‘atmosphere’ goods in the terminology of Meade) unpaid factors of production such as roads or sewer systems are also characterised by congestion or locational aspects (Duffy).

On the other hand, the input factor 'public capital' is not purchased by the firms at the market as are labour and private capital, but is supplied by governmental institutions. Moreover, private firms can allocate resources to influence the level of public capital provision by working through the political process. In other words, firms can lobby in order to obtain the desired level of public capital from the governmental institutions.

Consequently, 'pork-barrel' politics due to the influence of firms on the allocation of investment may determine the regional allocation of infrastructure investment. However, we also consider distortions in the allocation of infrastructure due to the political affiliation of governments at different levels. These influences can give rise to allocation outcomes that might depart substantially from an optimal allocation as a result of maximising social welfare.

Our approach for measuring the politico-economic determinants of infrastructure investment allocation in essays 3 and 4 is based on a general framework for political-economy analysis which views economic policy decisions as being a result of the maximisation of objective functions by incumbent politicians under constraints that are primarily political (e.g., Dixit, 1996). Thus, rather than treating policy as exogenous or chosen by the mythical 'social planner' public policy is viewed as the result of some interaction between citizens-voters and policy-makers within institutions having certain characteristics (Alt and Alesina, 1996).

Our approach is also related to the economics of rent-seeking, which is interested in the rational behaviour of special interest groups and politicians. For instance, early studies on this topic (e.g., Krueger, 1974) pointed out when individuals can gain from government policies, they have an incentive to expend resources up to the expected value of that gain in order to get the benefits, thereby creating substantial welfare losses in the process. The policy outcome depends on the possibilities of the interest groups to influence the budget process according to their wishes (Ekelund and Tollison, 2001).

The literature on the determinants of infrastructure policies so far has mainly discussed optimal rules for the provision of infrastructure at different levels of government in a federal system (e.g., Hulten and Schwab, 1997). However, it remains an open question whether and how much infrastructure policies in real-

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ity are designed according to such efficiency considerations. To our knowledge, only a few studies so far have discussed the politics of infrastructure (e.g. Crain and Oakley, 1995). In this perspective, the thesis bridges the gap between the infrastructure and political-economy literature.

In essays 3 and 4 we test the hypothesis that the number of manufacturing firms in a given region is decisive for infrastructure spending. This prediction is derived from a common-agency model whereby incumbent politicians sell infrastructure investments to local lobbies who bid for them through campaign contributions. The underlying assumption is that firms have sunk investments and therefore a vested interest in the quality and maintenance of the infrastructure where they are located. Politicians—on the other hand—are assumed to be responsive to the lobbying efforts by business, for instance in anticipation of potential campaign contributions from firms, or in anticipation of the expected loss of tax revenues and/or employment opportunities if firms move to another location.

Essays 3 and 4 also test the idea that party affiliation between higher and lower-tier governments matters for the outcome of infrastructure policies. Grossman (1994), for example, argues that the ‘political capital’ of a region is of higher value to a higher-tier government if the party affiliation is the same. Accordingly, the distribution of infrastructure investments is driven by self-interests of politicians with the purpose to enhance reelection chances. In essay 4, for the specific case of large German cities our interpretation of the significance of party affiliation is that the identity of political colour shortcuts the bargaining process between lower and higher-tier governments and thereby favours certain municipalities by means of party loyalty.

Recently, it has also been suggested in the literature that redistributive politics are used as tactical (electoral politics) instruments for buying support of marginal voters (‘swing voter’ approach, e.g. Dixit and Londregan, 1998). Using this framework, we furthermore test in essay 4 the hypothesis that cities will receive more grants if they are politically powerful, i.e. if there is a large number of voters who are indifferent between the two parties and therefore potentially could be influenced by pork barrel politics.

In essay 5 we consider further economic and institutional determinants for the allocation of infrastructure investments in Germany and France. Following the approaches of de la Fuente and Vives (1995) or Martin (1999b) we find that—surprisingly—neither for France nor for Germany efficiency considerations matter for the allocation of infrastructure investment across regions. However, we find that in Germany, in contrast with France, the priority of promoting the regional convergence of living conditions is important for infrastructure policy in Germany.

## 1.2 Literature survey on empirical infrastructure studies based on production functions

In this section we provide a short overview of results of empirical infrastructure studies based on the production function approach, because throughout the essays our analysis will be based on this approach. For more comprehensive surveys including studies employing the dual cost or profit function approach see for instance Gramlich (1994), Pfähler, Hofmann and Bönnte (1997) or Sturm (1998).

Table 1.1 shows a summary of results of empirical infrastructure studies at the *regional* level for the US. One study which is closely related to our own research in essay 1 is that from Hulten and Schwab (1991). This study, like essay 1, 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 1.1 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  $\epsilon_{YG}$  differs considerably.

Another interesting insight from Table 1.1 is that the degree of the estimation's econometric sophistication also varies substantially among these studies. While

**Table 1.1: Production Function Studies at the Regional Level for the US**

Study	Data	Specification	Results, output elasticity $\epsilon_{YG}$
Kelejian and Robinson (1997)	US, panel, 48 states, 1970-1986, public capital stock from Munnell (1990a) <sup>a</sup>	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) <sup>a</sup>	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) <sup>a</sup>	translog	$\epsilon_{YG}=0.11$ , highways, water & sewer systems significant, other types not
Baltagi and Pinnoi (1995)	US, panel, 48 states, 1970-1986, public capital stock from Munnell (1990a) <sup>a</sup>	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) <sup>a</sup>	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) <sup>a</sup>	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 <sup>a</sup> for public capital	Cobb-Douglas	$\epsilon_{YG}=0.14$

<sup>a</sup> Three different types of public capital: 1. Highways, 2. Water and sewer systems, 3. Other types (primarily buildings)

**Table 1.1:** (cont.) *Production Function Studies at the Regional Level for the US*

<b>Study</b>	<b>Data</b>	<b>Specification</b>	<b>Results, output elasticity <math>\epsilon_{YG}</math></b>
Garcia-Milà and McGuire (1992)	US, panel, 48 states, 1969-83, highway capital, education expenditures	Cobb-Douglas, time effects	$\epsilon_{YG}=0.04$ for highways, education significant
Carlino and Voith (1992)	US, panel, 48 states, 1963-1986, highway density, educational attainment	CES, fixed and random effects	$\epsilon_{YG}=0.22-1.00$ for highways, education significant
Moonaw and Williams (1991)	US, panel, 48 states, manufacturing, 1959-76, TFP growth highway density		$\epsilon_{YG}=0.17$
Hulten and Schwab (1991)	US, panel, manufacturing, 9 regions, public capital stock from Munnell (1990a) <sup>a</sup>	TFP growth, time effects	not significant
Munnell (1990a)	US, panel, 48 states, 1970-1986, capital outlays from <i>Government Finances</i> <sup>a</sup>	Cobb-Douglas, translog	$\epsilon_{YG}=0.16$ 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 <i>Government Finances</i>	translog	$\epsilon_{YG}=0.19-0.26$ , 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

<sup>a</sup> Three different types of public capital: 1. Highways, 2. Water and sewer systems, 3. Other types (primarily buildings)

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.

Table 1.2 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. Utilising 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, utilising 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  $\epsilon_{YG}$  between 0.08 and 0.19.

Next, we discuss some studies at the *national* level. An overview of the results of these studies is shown in Table 1.3. 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 emphasised 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 all regional spill-overs from

**Table 1.2: Production Function Studies at the Regional Level for Other Countries**

<b>Study</b>	<b>Data</b>	<b>Specification</b>	<b>Results, output elasticity <math>\epsilon_{YG}</math></b>
Prud'Homme (1996)	France, panel, 21 regions, 1970-1990, transportation infrastructure	Cobb-Douglas, TFP growth	$\epsilon_{YG}=0.08$
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	$\epsilon_{YG}=0.08$ to 0.19
de la Fuente and Vives (1995)	Spain, panel, 17 regions, 1981, 1986, 1990, transportation infrastructure, education	Cobb-Douglas, translog, time effects	$\epsilon_{YG}=0.21$
Picci (1995)	Italy, panel, 20 regions, 1970-1991, public capital stock	Cobb-Douglas, fixed and random effects, 1. diffs.	$\epsilon_{YG}=0.08-0.43$ , 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 et al. (1987) and Mera (1973)	translog, fixed-effects, SUR estimation	$\epsilon_{YG}=0.46-0.58$ for Japan, $\epsilon_{YG}=0.20$ for US
Mera (1973)	Japan, panel, 9 regions, 1954-63, social overhead capital	Cobb-Douglas	$\epsilon_{YG}=0.12-0.22$

**Table 1.3: Production Function Studies at the National Level**

<b>Study</b>	<b>Data</b>	<b>Specification</b>	<b>Results, output elasticity <math>\epsilon_{YG}</math></b>
Everaert and Heylen (2001)	Belgium, time-series 1953-96, public capital stock	multifactor productivity, cointegration	significant, $\epsilon_{YG}=0.29$
Batina (1999)	US, time-series, different proxies for public infrastructure	aggregate production function, dynamic OLS	significant
Duggal, Saltzman and Klein (1999)	US, time-series, public infrastructure	non-linear 'S-shaped' production function, 2SLS	significant
Denny and Guiomard (1997)	Ireland, time-series, manufacturing, 1951-1994, stock of roads & highways	Cobb-Douglas, AR(1)	$\epsilon_{YG}=0.92$
Fernald (1999)	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	$\epsilon_{YG}=0.27-0.42$

**Table 1.3:** (cont.) *Production Function Studies at the National Level*

<b>Study</b>	<b>Data</b>	<b>Specification</b>	<b>Results, output elasticity <math>\epsilon_{YG}</math></b>
Bajo-Rubio and Sosvilla-Rivero (1993)	Spain, time-series, 1964-88, public capital	Cobb-Douglas, cointegration, Hausman exogeneity test	$\epsilon_{YG}=0.18$ , public capital exogenous
Berndt and Hansson (1992)	Sweden, time-series, 1964-88, public infrastructure	Cobb-Douglas	$\epsilon_{YG}=-1.66-0.369$ , 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 (1990c)	US, time series, 1948-87, public capital	Cobb-Douglas	$\epsilon_{YG}=0.34-0.37$
Aschauer (1989a)	US, time series, 1949-85, non-military public capital	Cobb-Douglas	$\epsilon_{YG}=0.38-0.56$
Ratner (1983)	US, time series, 1949-73, non-military public capital	Cobb-Douglas	$\epsilon_{YG}=0.06$

**Table 1.4: Production Function Studies at the International Level**

<b>Study</b>	<b>Data</b>	<b>Specification</b>	<b>Results, output elasticity <math>\epsilon_{YG}</math></b>
Aschauer (1995)	OECD, 12 countries, panel, infrastructure capital from Ford and Poret (1991)	TFP growth, fixed country and time effects, 4-year average	$\epsilon_{YG}=0.33-0.55$
Nourzad and Vrieze (1995)	OECD, 7 countries, panel, 1963-88, public investment (data sources not given)	Cobb-Douglas, energy input, 1. diffs., random effects	$\epsilon_{YG}=0.05$
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 privately 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	$\epsilon_{YG}=0.34-0.73$

infrastructure are fully captured (Munnell, 1993).

However, the result of Aschauer's study has led to considerable scepticism in literature (Gramlich, 1994; Jorgenson, 1991; Tatom, 1991a; Tatom, 1991b; Tatom, 1993a). 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.

Finally, Table 1.4 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 summarise this short review, the published results on the productivity effects of infrastructure so far are rather ambiguous. Moreover, this literature survey 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.

### 1.3 Outline of the thesis

Essay 1 examines the impact of public capital on private production in manufacturing at the regional level of the 11 German 'Bundesländer' (federal states) in the period from 1970 to 1996. Essay 2 uses a partial adjustment approach to measure the contribution of public infrastructure to private production at the local level of the 327 German counties ('Kreise'). Public infrastructure endowment is measured by means of indicators and the analysis is carried out for 3 different sectors of the local economy: service, trade & transport and manufacturing. Essay 3 employs a simultaneous-equation approach to the estimation of the contribution of transport infrastructure accumulation to regional growth. The model is based on two equations. The first equation describes the regional production function and the second consists of a policy function which determines the regional allocation of infrastructure investment. In essay 4 we test the political-economy approach of essay 3 with a data set of large self-administrated German cities. Essay 5 investigates the productivity effects of infrastructure using a pooled data set of German and French regions. Furthermore, essay 5 analyses the institutional determinants of infrastructure policies in Germany and France.

## **Part II**

### **Essays**

## Essay 1

# Assessing the Contribution of Public Capital to Private Production: Evidence from the German Manufacturing Sector

### *Abstract*

*Using time-series cross-section data from the manufacturing sector of the 11 German 'Bundesländer' (federal states) from 1970 to 1996, we examine the impact of public capital on private production. Our econometric analysis explicitly takes into account four of the most frequent specification issues in the context of time-series cross-section data analysis: serial correlation, groupwise heteroscedasticity, cross-sectional correlation and nonstationarity of data. For all approaches and tested specifications, we find that public capital is a significant input for production in the manufacturing sector. Moreover, we find that differences in public capital endowment can explain long-term differences in productivity across the 'Bundesländer'.*

## 1.1 Introduction

Our study is motivated by the controversy that has developed recently about the contribution of public capital—e.g. highways, mass transits, water and sewer systems, etc.—to private production. This controversy has been stimulated by the large elasticity of output with respect to public capital found in the pioneering work of Aschauer (1989*a*; 1989*b*). Aschauer's findings suggest that part of the productivity slowdown observed in the 1970s and 80s in the United States and in other OECD countries may be due to an underinvestment in public capital. This has become known in the literature as the 'public capital hypothesis'.

A number of follow-up studies have been spurred by this controversy some of which have been supportive to the public capital hypothesis (Berndt and Hansson, 1992; Fernald, 1999; Morrison and Schwartz, 1996; Munnell, 1990*b*; Munnell, 1992; Otto and Voss, 1994; Ram and Ramsey, 1989) while others have not been supportive (Baltagi and Pinnoi, 1995; Garcia-Milà et al., 1996; Erber, 1995; Evans and Karras, 1994*a*; Holtz-Eakin, 1994; Hulten and Schwab, 1991; Tatom, 1991*a*; Tatom, 1993*a*).<sup>1</sup> The usual approach taken in these studies is to regress some measure of output e.g. gross domestic product (GDP) or value added on an array of factor inputs and a measure of public capital.<sup>2</sup>

The purpose of this essay is to examine the significance of the 'public capital hypothesis' for Germany. One major finding that emerges from our empirical investigation is that public capital appears to be a significant determinant for private production in the manufacturing sector. Thus, our empirical results are in line with other studies for Germany e.g. Seitz (1993), Licht and Seitz (1994), Seitz (1994) or Schlag (1997). However, we stress that our study (i) uses a different methodology which is not based on the cost but on the less restrictive production function approach, (ii) focuses on the manufacturing sector at the regional level of the Bundesländer and (iii) incorporates several important econometric issues in the statistical analysis which have been neglected in previous studies.

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<sup>1</sup>For comprehensive surveys on this literature, see for instance Gramlich (1994), Sturm, Kuper and de Haan (1996) or Pfähler et al. (1997).

<sup>2</sup>Berndt and Hansson (1992), Erber (1995) and Morrison and Schwartz (1996) are based on a dual cost function instead of a production function. Hulten and Schwab (1991) and Fernald (1999) use total factor productivity (TFP) growth as the dependent variable in the analysis.

Thus, our study addresses some important methodological concerns raised about previous studies. For example, as pointed out by Aaron (1990), Jorgenson (1991) and Tatom (1991a; 1993a) most of the time series employed for the examination of the relationship between public capital and private output are likely to be nonstationary and thus they advise estimating the model in first differences if the variables are not cointegrated. Following this advice, for instance, Tatom (1991a) or Garcia-Milà et al. (1996) find the elasticity of output with respect to public capital to be insignificant for the US. This highlights the importance of an appropriate modelling of stochastic or deterministic trends in variables. In our empirical analysis this matter is particularly considered.

Another important motivation of our study is that we intend to shed some light onto the nature of the positive correlation between public capital and private output. Thus, we analyse the underlying structure of the data that gives rise to this correlation. The question is whether the correlation results from the variation between cross-sections (Bundesländer) or from the variation over time, i.e. from the ‘within’ variation. Moreover, we also investigate whether this correlation is manifested in the short-run or in the long-run trends in the data.

The remainder of this essay is organised as follows. Section 1.2 outlines the specification used in the empirical analysis. Section 1.3 presents the results and considers several econometric specification issues. Section 1.4 summarises and concludes the essay.

## 1.2 Specification

This section considers the specification for our econometric approach to assess the contribution of public capital to private production.

Suppose that production of value-added output  $Q_{it}$  in the manufacturing sector in Bundesland  $i = 1, \dots, B$  at time  $t = 1, \dots, T$  depends on inputs of private capital  $K_{i,t-1}$  and labour  $L_{i,t-1}$ . Our assumption that output at time  $t$  is only produced using factor inputs which were already installed in period  $t - 1$  avoids a potential simultaneity between output and factor inputs.

We assume that output  $Q_{it}$  also depends on the Hicks-neutral level of technology  $A_i()$ , which is a function of time  $t$  and the level of the non-rival public input  $G_{i,t-1}$ . Suppose  $A_i()$  takes the functional form  $A_i = A_{i0} G_{i,t-1}^{\beta_g} \exp(\lambda t)$ , where  $A_{i0}$  is the initial level of technology at time  $t = 0$  in Bundesland  $i$  and  $\lambda$  is the exogenous rate of technology growth. The exogenous technology growth rate  $\lambda$  is therefore restricted to be the same for all Bundesländer whereas the initial level of technology  $A_{i0}$  can vary across the Bundesländer.

Now, specifying a Cobb-Douglas functional form we get the estimating equation in logarithms as

$$\ln Q_{it} = \ln A_{i0} + \lambda t + \beta_g \ln G_{i,t-1} + \beta_k \ln K_{i,t-1} + \beta_l \ln L_{i,t-1} + \varepsilon_{it}, \quad (1.1)$$

$$i = 1, \dots, B, \quad t = 1, \dots, T,$$

where  $\varepsilon_{it}$  denotes an error term which reflects contemporaneous exogenous shocks to logarithmic output  $\ln Q_{it}$ . We assume that  $\varepsilon_{it}$  is an i.i.d. random variable with variance  $\sigma_\varepsilon^2$ . Note that in (1.1) the estimate  $\hat{\beta}_j, j \in \{g, k, l\}$ , gives the elasticity of output with respect to factor  $j$ . Note also that in contrast to many previous studies we refrain from including measures of capacity utilisation in (1.1). All measures of capacity utilisation we are aware of (e.g. the quite often used unemployment rate) incur the risk of simultaneity with output, since the actual level of output directly or indirectly influences these measures. Instead we assume that over- or underutilisation of factors in producing output  $Q_{it}$  are reflected in the error term  $\varepsilon_{it}$ .

Even if the Cobb-Douglas functional form is restrictive because the elasticities of substitution of input factors are restricted to equal one, it is a first order approximation to any arbitrary production function in the neighbourhood where the factor input vector  $X = (G, K, L)$  is  $(1, 1, 1)$ .<sup>3</sup> It is worth stressing that (1.1) does not put any restriction on the technology with respect to returns to scale.

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<sup>3</sup>In order to capture the second order effects we also estimated flexible functional forms for the production function e.g. translog (Christensen, Jorgenson and Lau, 1971; Christensen, Jorgenson and Lau, 1973) in the empirical analysis. However, it turned out that the estimation of these specifications suffered from a strong multicollinearity problem. This problem arises from extremely high correlations of the single factor inputs with the quadratic and the cross effect terms.

It should be mentioned that instead of a production function it would have been possible to specify a dual cost function approach with public capital entering as a quasi-fixed unpaid factor of production. However, at such a fairly high level of aggregation the behavioural assumption of the cost function approach that costs are endogenous and determined by choosing cost minimizing quantities of factor inputs given a certain exogenous quantity of output seems to be unrealistic (Berndt, 1991, p. 457). Furthermore, factor prices are quite often not directly observed but have to be calculated using some (restrictive) assumptions which are likely to introduce further sources of measurement error in the data. The production function approach on the other hand does neither require a behavioural (minimizing or maximising) assumption nor does it require data on factor prices (Chambers, 1988).

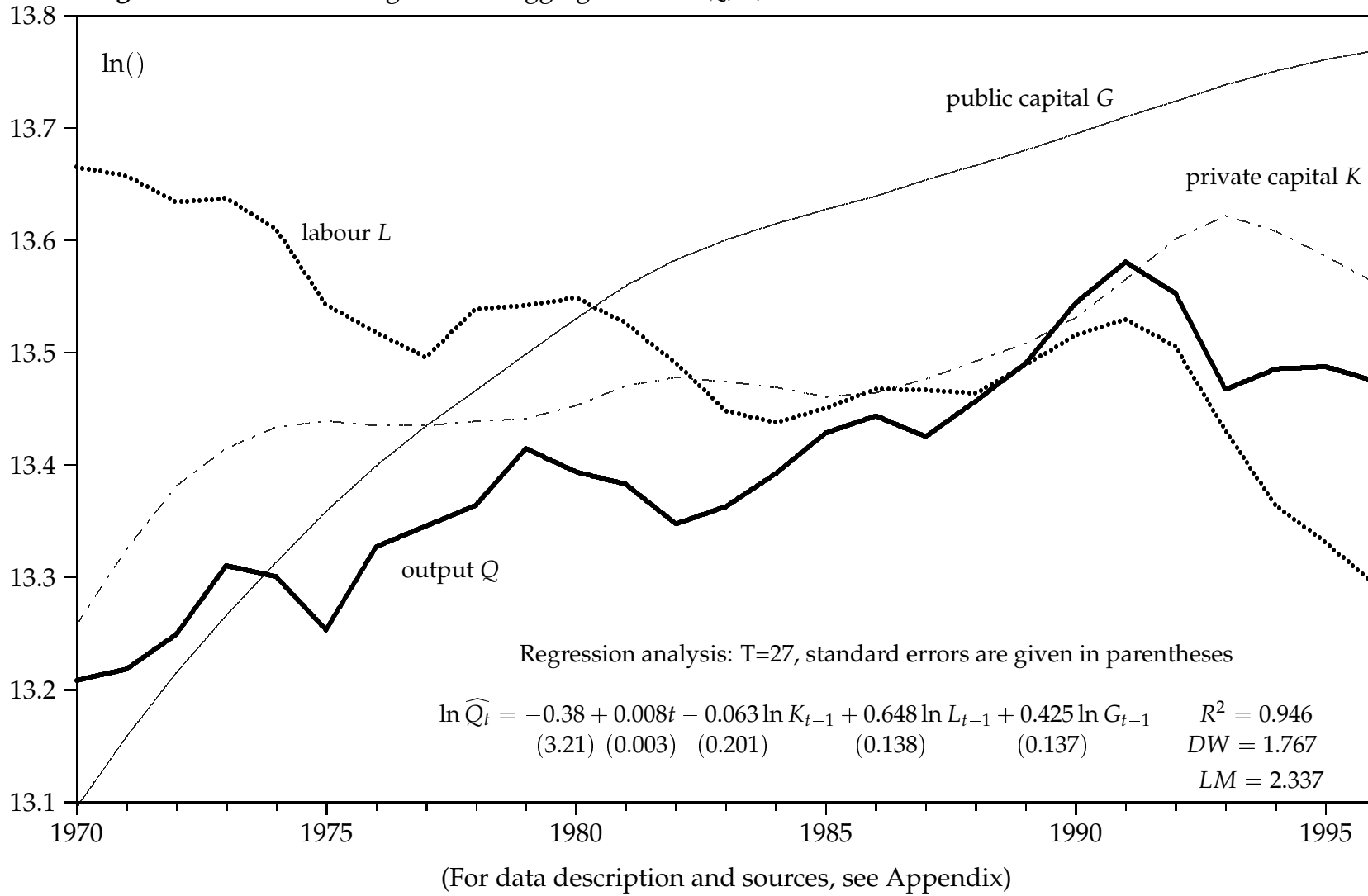
### 1.3 Econometric issues and results

The data used in the analysis cover the manufacturing sector of the 11 German Bundesländer ( $B = 11$ ) from 1970 to 1996 ( $T = 27$ ). A comprehensive description of the data is given in the Appendix.

Figure 1.1 graphs the aggregate series of  $Q$ ,  $L$ ,  $K$  and  $G$  over the period 1970-1996. Growth of the aggregate public capital stock was particularly high during the period from 1970 to 1981. After 1983 the growth rate of public capital slightly declined compared to the previous period, but was still positive and relatively constant.

On the other hand the aggregate private capital stock in manufacturing grew with a relatively high rate from 1970 to 1975, but in the period 1976-1985 the growth rate of the capital stock wore off. Note that changes in aggregate private capital appear to follow changes in output with a lag of about 2 to 3 years. For example the decrease in output during the years 1979 to 1982 seems to have had an effect on the formation of private capital after 1982. Hence, at least at the aggregate level there is some evidence that the private capital formation is likely to follow the development in output and not vice versa. Similarly, from 1993

**Figure 1.1:** Manufacturing Sectors' Aggregate Series  $Q$ ,  $L$ ,  $K$  and  $G$  for West German Bundesländer from 1970-1996



to 1996 we can observe a decline in the stock of private manufacturing capital whereas a sharp decline in output occurred already from 1991 to 1993.

The aggregate labour series shows a clear downward trend over the total period. This can be ascribed to the structural change in the German economy where the share of the manufacturing sector's employment in the total economy is declining.

Note also that the series of output and labour show rather high annual fluctuations due to the business cycles of the economy, whereas the series for capital, and in particular for public capital, are more smooth. One reason for this is that planning and decisions in public investment are long-term oriented, sometimes with a horizon of 5 up to 15 years. Therefore, annual fluctuations in output, i.e. fluctuations due to business cycles, do not appear to have an impact on the short-run formation of public capital. However, in the long-run, business cycles are likely to influence the formation of public capital due to the effects of the business cycles on tax revenues.

In addition, Figure 1.1 also presents the results of a regression analysis where output  $Q$  is regressed on inputs  $L$ ,  $K$ ,  $G$  and a linear trend  $t$ . The basic specification for the estimation is an autoregressive model of order one (AR(1)), which has been estimated by applying the iterated Prais-Winsten method (Greene, 2000, p. 547). Note that the Durbin-Watson statistic (DW) as well as the Godfrey's Lagrange multiplier statistic (LM) which is distributed chi-square with one degree of freedom indicate that the AR(1) is the appropriate specification (against the alternative of higher order ARMA specifications).

We find that the linear time trend  $t$ , as well as labour  $\ln L_{t-1}$  and public capital  $\ln G_{t-1}$  are statistically significant at a 5 percent level, whereas private capital  $\ln K_{t-1}$  is not. The fit of this preliminary regression with a  $R^2$  of about 0.95 is remarkably high. It is worth mentioning that the estimate for labour with a value of 0.648 is in line with the share of wages in value added of the manufacturing sector in our sample, which is about 0.55.

To begin with the main part of the empirical analysis based on the time-series cross-section data, we first present results for the model where only private inputs are included in (1.1), i.e.  $K_{i,t-1}$  and  $L_{i,t-1}$ . This preliminary step is undertaken in

order to evaluate the changes in results due to the inclusion of the public capital input  $G_{i,t-1}$  in the production function (1.1). In the second step we therefore present estimation results for the model with all inputs, including  $G_{i,t-1}$ .

Ordinary Least Squares (OLS) estimation of the Cobb-Douglas production function as specified in (1.1) with private inputs  $K_{i,t-1}$  and  $L_{i,t-1}$  using the pooled time-series cross-section data yields the following estimates:<sup>4</sup>

$$\widehat{\ln Q_{it}} = \text{Länder-effects}^* + 0.014^* t + 0.151 \ln K_{i,t-1} + 0.645^* \ln L_{i,t-1} \quad (1.2)$$

$F(10|283)=142.7 \quad (0.002) \quad (0.084) \quad (0.072)$

N: 297 ( $G = 11, T = 27$ )  $R^2 : 0.9931 \quad \hat{\rho} : 0.868$

**Diagnostic test**

Test for serial correlation:  $DW : 0.262^*, \rho_{LM} = 209.8^* \sim \chi_{df=1}^2$

Test for groupwise heteroscedasticity:  $LM = 181.4^* \sim \chi_{df=10}^2$

Test for cross-sectional correlations:  $\lambda_{LM} = 478.1^* \sim \chi_{df=55}^2$

Test for random walk of residual:  $R_p : 0.243$

Hausman test: 2.30

Multicollinearity: condition number = 416.4

Notice that in (1.2) the included dummy variables for the Bundesländer ('Länder' effects) correspond to the term  $\ln A_{i0}$  in (1.1). The displayed  $F$ -test indicates that these Bundesländer effects are highly significant. The value of 2.30 of the Hausman test favours a random effects model against the fixed effects model. Furthermore, labour is significant with a value of 0.645. However, the estimate of private capital is not significant. Note, that the fit of the regression with  $R^2$  equal to 0.9931 is remarkable high.

A frequent observation in the empirical analysis of time-series data is the presence of autocorrelation. Also, it is very likely that heteroscedasticity will be observed as the Bundesländer in our sample have different sizes. Furthermore, macroeconomic factors affecting one region will also affect other regions, thus the errors across the Bundesländer are likely to be correlated.

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<sup>4\*</sup> denotes statistical significance at a 5 percent level. Robust panel corrected standard errors (PCSEs) are given in parentheses. PCSEs are estimated by the square root of the diagonal of  $(X'X)^{-1}X(\Sigma \otimes I_T)X(X'X)^{-1}$  where  $\Sigma$  is a  $N \times N$  matrix of cross-sectional variances and covariances. A consistent estimate of  $\Sigma$  is given by  $E'E/T$ , where  $E$  denotes  $T \times i$  matrix of OLS residuals from equation (1.2) (Beck and Katz, 1995; Beck and Katz, 1996).

In order to explore such econometric specification issues, several diagnostic checks are shown in (1.2).<sup>5</sup> First, to test for the presence of autocorrelation, the value of the Durbin-Watson (DW) statistic which is 0.262 has been calculated from the residuals of the OLS estimation according to Bhargava, Franzini and Narendranathan (1982) as

$$DW = \frac{\sum_{i=1}^G \sum_{t=2}^T (\tilde{u}_{it} - \tilde{u}_{i,t-1})^2}{\sum_{i=1}^G \sum_{t=1}^T \tilde{u}_{it}^2},$$

where  $\tilde{u}_{it}$  are the residuals from the fixed effects model (1.2). The Durbin-Watson statistic can be used to test the null hypothesis that the serial correlation is  $\rho = 0$  against the alternative that  $|\rho| < 1$ . The *exact* critical value for the DW statistic is 1.810 and has been found by using the Imhof (1961) routine.<sup>6</sup> Thus, the null that the errors of the OLS estimation are serially independent is rejected.

This finding is also confirmed by the value of the Lagrange-Multiplier test statistic  $\rho_{LM} = 209.8$ .<sup>7</sup> This statistic is distributed  $\chi^2$  with 1 degree of freedom ( $\chi_{crit,0.05,df=1}^2 = 3.84$ ), hence we can reject the null hypothesis of serial independence at a 5 percent level by this test.

Second, in order to test for groupwise heteroscedasticity the following Lagrange multiplier (LM) test has been calculated as (Greene, 2000, p. 596)

$$LM = T/2 \sum_1^B \left[ \frac{s_i^2}{s^2} - 1 \right]^2,$$

where  $s^2$  is the pooled OLS residual variance and  $s_i^2$  is the estimated unit-specific residual variance from groupwise regressions. The LM statistic has a limiting  $\chi^2$  distribution with  $B - 1$  degrees of freedom. The reported value of 181.4 from the LM statistic leads to a rejection of the null hypothesis of no groupwise heteroscedasticity ( $\chi_{crit,0.05,df=10}^2 = 18.3$ ).

Third, in order to test for cross-sectional correlations the residuals obtained from (1.2) are used to compute the following Lagrange multiplier statistic

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<sup>5</sup>In all these cases, OLS estimation still yields consistent parameter estimates. However, estimates of standard errors will be biased and inconsistent.

<sup>6</sup>The Imhof routine is implemented in SHAZAM 8.0.

<sup>7</sup>The Lagrange multiplier statistic is found by regressing  $\tilde{u}_{it}$  on  $\tilde{u}_{i,t-1}$  and the other regressors. The statistic  $\rho_{LM}$  is then defined as the  $R^2$  obtained from this auxiliary regression multiplied with the number of observations.

**Table 1.1:** Cross-sectional Correlation and Variance/Covariance<sup>1</sup> Matrix for the 11 Bundesländer Based on Residuals from Equation (1.3)

	BaW	Bay	Ber	Bre	Ham	Hes	Nie	NRW	RhP	Saa	SHo
BaW	<b>1.46</b>	0.90	1.78	-0.22	-1.86	1.75	0.79	1.38	1.10	2.09	-0.89
Bay	0.68	<b>1.19</b>	1.37	0.77	-0.35	1.35	1.01	1.13	1.37	1.27	-0.31
Ber	0.74	0.63	<b>3.91</b>	-1.50	-4.30	2.27	1.00	2.02	1.63	2.43	-1.63
Bre	-0.08	0.34	-0.36	<b>4.45</b>	3.73	0.10	0.94	0.06	1.15	-0.37	0.85
Ham	-0.41	-0.09	-0.57	0.46	<b>14.36</b>	2.47	4.51	8.67	7.09	6.87	21.03
Hes	0.89	0.76	0.70	0.03	-0.46	<b>2.64</b>	1.45	1.99	1.83	2.86	-2.13
Nie	0.50	0.70	0.39	0.34	0.07	0.68	<b>1.73</b>	1.21	1.39	1.35	-1.16
NRW	0.86	0.78	0.77	0.02	-0.39	0.93	0.70	<b>1.75</b>	1.60	2.30	-1.71
RhP	0.63	0.87	0.57	0.38	-0.22	0.78	0.74	0.84	<b>2.09</b>	1.71	-1.51
Saa	0.87	0.59	0.62	-0.09	-0.43	0.90	0.52	0.89	0.60	<b>3.87</b>	-2.55
SHo	-0.26	-0.10	-0.30	0.15	0.66	-0.47	-0.32	-0.47	-0.37	-0.47	<b>7.64</b>

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

<sup>1</sup>Variances / covariances [ $10^{-3}$ ], correlations are given below, variances on the diagonal

(Greene, 2000, p. 452)

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

where  $r_{ij}^2$  is the squared  $ij$ th correlation coefficient of residuals between Bundesland  $i$  and  $j$ . The large-sample distribution of this statistic is chi-square with  $B(B - 1)/2$  degrees of freedom. Hence, this statistic with a value of 478.1 is highly significant, indicating the presence of substantial cross-sectional correlations between the Bundesländer ( $\chi_{crit,0.05,df=55}^2 = 73.3$ ).

Table 1.1 shows the correlations  $r_{ij}$  and variances / covariances of residuals between the Bundesländer. The variances of the residuals of the Bundesländer are given in bold print on the diagonal of the matrix. Covariances are given in the upper half of Table 1.1. The ratio of the largest variance with 14.36 ('Hamburg') to the smallest with 1.19 ('Baden-Württemberg') is about 12, which confirms the high degree of groupwise heteroscedasticity in the data. Similarly, some of the correlations shown in the lower half of Table 1.1 are remarkably high, for instance between 'Baden-Württemberg' and 'Hessen' with a value of about 0.9.

Fourth, since the first glance at Figure 1.1 revealed that the (aggregate) se-

ries exhibit some (random or deterministic) trends, the  $R_p$  statistic<sup>8</sup> according to Bhargava et al. (1982) for testing the null that the residuals from (1.2) follow a random walk, i.e.  $\rho = 0$  against  $|\rho| > 0$ , is also presented. Small values of  $R_p$  favour the null hypothesis. The exact critical value for this statistic again can be found by using the Imhof routine. In our case the critical value for  $R_p$  at a 5 percent level is 0.336<sup>9</sup>, therefore the null hypothesis of a unit root is not rejected.<sup>10</sup> Since the error is nonstationary, the variables appear not to be cointegrated.

Finally, a recent paper by Ai and Cassou (1997) points out that the findings of some studies for the US using fixed effects models in the analysis of productivity effects of public capital based on state level data, e.g. Holtz-Eakin (1994) or Evans and Karras (1994a), should be interpreted with some caution because of the high correlation between the public capital stocks and the fixed effects. This multicollinearity problem arises because there is not enough variation in the public capital series to disentangle the effect of public capital from the state specific effect, i.e. the public capital series do not have enough ‘within’ variation. Thus, to get some indication whether multicollinearity matters for our estimations we also report the condition number<sup>11</sup> which has a value of 416.4. Judge, Griffiths, Hill, Lee and Lütkepohl (1985, p. 902) suggest that values exceeding 30 reveal potential multicollinearity problems. Thus, the occurrence of poor or imprecise estimates can be a result of the high degree of multicollinearity in the data.

Estimating (1.1) for all inputs, i.e.  $K_{it}$ ,  $L_{it}$  and  $G_{it}$  we obtain the following results:

$$\widehat{\ln Q_{it}} = \text{Länder-effects}^* + 0.001 t - 0.139 \ln K_{it} + 0.805^* \ln L_{it} + 0.781^* \ln G_{it}$$

$F(10 282)=159.2$	(0.002)	(0.077)	(0.062)	(0.078)
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(1.3)

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<sup>8</sup>The  $R_p$  statistic is calculated as  $R_p = e'e/e'F^*e$ , where  $e$  are the OLS residuals from estimating (1.2) in first differences,  $F^* = (I_G \otimes F)$ , and  $F$  is a  $(T - 1) \times (T - 1)$  symmetric matrix with elements of the form  $F_{jk} = (T - j)k/T$  if  $j \geq k$  and  $F_{jk} = F_{kj}$ .

<sup>9</sup>The 10 percent critical value is 0.307, the 1 percent critical value for  $R_p$  is 0.398 ( $B = 11, T = 27$ ).

<sup>10</sup>For alternative approaches of testing for unit roots with panel data, see for instance Baltagi and Kao (2000) or Maddala and Kim (1998).

<sup>11</sup>The condition number is defined as the square root of the ratio of the largest to the smallest Eigenvalue of  $S(X'X)S$ , where  $S$  is a diagonal matrix with the  $k$ th diagonal element  $1/\sqrt{x'_k x_k}$ .

$N: 297 (G = 11, T = 27) \quad R^2 : 0.9959 \quad \hat{\rho} : 0.783$

**Diagnostic tests**

Test for serial correlation:  $DW : 0.432^*, \rho_{LM} = 178.2^* \sim \chi^2(df = 1)$

Test for groupwise heteroscedasticity:  $LM = 121.1^* \sim \chi^2(df = 10)$

Test for cross-sectional correlations:  $\lambda_{LM} = 496.3^* \sim \chi^2(df = 55)$

Test for random walk of residuals:  $R_p : 0.347^*$

Hausman test: 21.64\*

Multicollinearity: condition number = 507.1

Again, we find that the coefficient of labour input is significant, whereas the coefficient of private capital is not. In contrast to this, the estimate of the coefficient of public capital input is highly significant. Here, the value of 21.64 of the Hausman test favours the fixed effects model against a random effects model. Also, from the increase in the Hausman test statistic from 2.30 in (1.2) to 21.64 in (1.3) we infer that public capital appears to be correlated with the Bundesländer effects.

The displayed diagnostic tests reveal that all the specification issues for estimation such as serial correlation, groupwise heteroscedasticity and cross-sectional correlation are present as before. One difference is, however, that the null hypothesis of a random walk of the residuals is rejected at a 5 percent, since the  $R_p$  statistic exceeds the critical value of 0.336.

Our further estimation strategy is therefore as follows. From the reported  $R_p$  statistics in (1.2) and (1.3) respectively it is difficult to judge whether a trend stationary or difference stationary model is more appropriate. In the former case the estimation can be carried out in levels, whereas for the latter case the estimation should be based on variables in first differences. The two presented regressions in (1.2) and (1.3) seem to require different treatment of trends according to the reported  $R_p$  statistics. For the model with only  $L$  and  $K$  as inputs the difference stationary model seems to be the appropriate specification whereas for the model with inputs  $L$ ,  $K$  and  $G$  the trend stationary model appears to be adequate. Therefore, we will present estimation results both for the specification in levels and for the specification in first differences. This also allows us to check the robustness of obtained results.

Additionally, instead of calculating robust PCSEs as in (1.2) and (1.3) another estimation strategy is to apply Feasible Generalised Least Squares (FGLS) in order to take serial correlation, groupwise heteroscedasticity and/or cross-sectional correlation properly into account. Feasible Generalised Least Squares (FGLS) estimation in the context of time-series cross-section models is also known as ‘Kmenta’ or ‘Parks’ method (Kmenta and Oberhofer, 1974; Kmenta, 1986; Parks, 1967). In two recent papers Beck and Katz (1995) have argued that one should be aware of the fact that although FGLS might be more efficient when cross-sectional correlations or groupwise heteroscedasticity are very significant, the standard errors obtained by the FGLS estimation do not correctly reflect the sampling variability of parameter estimates, because in samples of small size the cross-sectional correlations or variances obtained in the first step of FGLS are likely to be very poor estimates of the underlying ‘true’ variances. Thus, as Beck and Katz (1995) have shown by Monte-Carlo simulations standard errors from FGLS estimation in small samples have a tendency to be too small, they are ‘overconfident’. Beck and Katz recommend applying OLS estimation with consistent and robust panel corrected standard errors (PCSE) instead of FGLS if the ratio of number of time periods to the number of cross-sections is smaller than 3. This is the case for our sample, since the ratio of  $T$  to  $B$  is 2.45. Thus, there is a risk that standard errors obtained from FGLS are ‘overconfident’. Therefore, we present results both for FGLS estimation as well as for OLS with PCSEs.

For the AR(1) models a consistent estimate of the autocorrelation parameter  $\rho$  was obtained from residuals of equation (1.2) and (1.3) respectively as  $\hat{\rho} = 1 - DW/2$ . Using this estimate, the first step AR(1) correction has been carried out by employing the Cochrane-Orcutt transformation (Greene, 2000, p. 546). As such, the first observation in each group is lost.<sup>12</sup> In the second step, we use two estimation variants. The first variant—which is due to the AR(1) correction in the first step also a FGLS estimation—is based on OLS estimation in the second step with robust panel corrected standard errors (PCSEs) of the trans-

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<sup>12</sup>Note that in the fixed-effects model, the Prais-Winsten transformation (Greene, 2000, p. 546) is not an appropriate choice for an AR(1) correction, because the ‘within’ transformation, that is forming deviation from group means, will not remove the heterogeneity if the Prais-Winsten transformation is used for the first observation.

formed variables. The second variant is based on FGLS estimation in both steps ('Kmenta' method)—in the first step an AR(1) correction is performed and in the second step the FGLS estimation which takes groupwise heteroscedasticity and cross-sectional correlation into account. Also, for the model in first differences we apply both estimation methods, i.e. (i) OLS with PCSEs and (ii) FGLS ('Kmenta' method).

Table 1.2 summarises the results of the estimations. The upper half (I) contain the results for inputs  $K$  and  $L$ , and in the lower half (II) the results for inputs  $K$ ,  $L$  and  $G$ . Columns 1 and 2 present the results for the AR(1) models, whereas columns 3 and 4 display the results for variables in first differences. Note that only the AR(1) models includes the Bundesländer dummy variables (fixed effects), since the dummy variables are removed when taking first-differences. Similarly, only the AR(1) models includes a time trend  $t$ , because the time trend becomes a constant when taking first-differences.

The usual  $F$ -test for OLS relies on homoscedasticity. If this is not an appropriate assumption one can use a Wald test instead.<sup>13</sup> Both the  $F$ -tests as well as the Wald tests show that the Bundesländer effects are highly significant. Also, the null hypothesis of constant returns to scale (CRS) is rejected in almost all specifications, but not in (II) for all inputs  $K$ ,  $L$  and  $G$ .

By contrast with the low values for the  $DW$  statistics reported for the previous estimations (1.2) and (1.3), both the AR(1) and the model with variables in first differences generate  $DW$  statistics of about 2, indicating that autocorrelation and also stationarity of residuals are not problematic for the estimations. This is further confirmed by the Lagrange-Multiplier statistic  $\rho_{LM}$ , which does not reject the null hypothesis of serial independence for most specifications at a 5 percent level. However, in (II) for the model in first differences we find a value for  $\rho_{LM}$  of 3.80, which corresponds to a  $p$ -value of 0.051. This indicates that the observed serial dependence may result from taking first differences of a trend stationary model ('overdifferencing'). Therefore, in (II) the trend stationary AR(1) model is the preferred specification.

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<sup>13</sup>The Wald statistic  $W$  is defined as  $W = (\mathbf{R}\hat{\mathbf{f}} - \mathbf{q})' (\mathbf{R}(\text{Var}(\hat{\mathbf{f}})\mathbf{R}')^{-1} (\mathbf{R}\hat{\mathbf{f}} - \mathbf{q}))$ , where  $\mathbf{R}\hat{\mathbf{f}} = \mathbf{q}$  imposes a set of restrictions on the parameter vector  $\hat{\mathbf{f}}$  (Greene, 2000, p. 361).

**Table 1.2: Production Function Estimates (G=11, T=27)**

(I) Dependent variable $\ln Q_{it}$ , factors of production $K_{i,t-1}, L_{i,t-1}$								
	AR(1), $\rho = 0.869$				first differences			
	FGLS (PCSE)		FGLS (Kmenta) (het., corr.)		OLS (PCSE)		FGLS (Kmenta) (het., corr.)	
<i>const</i>	Länder-dum.*		Länder-dum.*		0.018*	(0.007)	0.021*	(0.004)
<i>t</i>	0.009	(0.005)	0.012*	(0.004)	—	—	—	—
$\ln K_{i,t-1}$	-0.174	(0.181)	-0.145*	(0.060)	-0.018	(0.185)	-0.041	(0.067)
$\ln L_{i,t-1}$	0.664*	(0.156)	0.678*	(0.066)	0.635*	(0.165)	0.698*	(0.068)
$R^2$	0.917		—		0.185		—	
N	286		286		286		286	
<b>Diagnostic tests</b>								
	F tests:		Wald tests $\chi^2$ :		F tests:		Wald tests $\chi^2$ :	
fixed effects	10.1*		463.9*		—		—	
CRS <i>K, L</i>	18.4*		6.20*		11.7*		2.96	
<i>DW</i>	2.056		—		2.073		—	
<i>LM</i> test: $\rho_{LM}$	0.60		—		0.67		—	
cond.-number	267.8		—		1.98		—	

(II) Dependent variable $\ln Q_{it}$ , factors of production $K_{i,t-1}, L_{i,t-1}, G_{i,t-1}$								
	AR(1), $\rho = 0.783$				first differences			
	FGLS (PCSE)		FGLS (Kmenta) (het., corr.)		OLS (PCSE)		FGLS (Kmenta) (het., corr.)	
<i>const</i>	Länder-dum.*		Länder-dum.*		0.006	(0.004)	0.012	(0.008)
<i>t</i>	0.003	(0.005)	0.007	(0.004)	—	—	—	—
$\ln K_{i,t-1}$	-0.176	(0.162)	-0.104	(0.067)	-0.221	(0.183)	-0.118	(0.076)
$\ln L_{i,t-1}$	0.663*	(0.145)	0.698*	(0.065)	0.669*	(0.156)	0.741*	(0.072)
$\ln G_{i,t-1}$	0.651*	(0.257)	0.549*	(0.120)	0.593*	(0.273)	0.417*	(0.161)
$R^2$	0.967		—		0.227		—	
N	286		286		286		286	
<b>Diagnostic tests</b>								
	F tests:		Wald tests $\chi^2$ :		F tests:		Wald tests $\chi^2$ :	
fixed effects	15.8*		727.9*		—		—	
CRS <i>K, L</i>	20.8*		7.88*		22.1*		6.7*	
CRS <i>K, L, G</i>	0.43		0.27		0.07		0.02	
<i>DW</i>	1.953		—		2.200		—	
<i>LM</i> test: $\rho_{LM}$	0.01		—		3.80		—	
cond.-number	571.1		—		3.84		—	

\* denotes statistical significance at a 5 % level, standard errors are given in parentheses

The parameter estimate of private capital in the upper half (I) of Table 1.2 is neither in the first differences nor in the AR(1) models significant, except in column 2 for the FGLS 'Kmenta' method. The labour input is significant in all specification with values of about 0.65. This is a reasonable estimate considering again that the average (wage) share of labour in output in our sample is about 0.55. Notice also the decrease of the condition number from the AR(1) to the specification in first differences. Hence, for the specification in first differences multicollinearity is not longer problematic for the estimations.

The results in (II) shows that in contrast to private capital the parameter for public capital appears to be significant in all specifications with values ranging between 0.42 and 0.65. Since the ratio of output  $Q$  to public capital  $G$  varies between 1.12 in year 1970 and 0.69 in year 1996, these estimated elasticities imply a marginal productivity of  $G$  between 47 and 73 percent in 1970 and between 29 and 45 percent in 1996.<sup>14</sup> The differences in  $R^2$  between (I) and (II) are about 0.04. Hence, in our model the public capital input can explain about 4 percent of the differences in observed output across the Bundesländer.

Finally, we provide several tests on the stability of parameters both over (i) cross-sections ('testing for poolability of the data') and (ii) over time ('testing for occurrence of structural breaks in the data'). In order to test (i) we perform a Chow test (Baltagi, 1995, chap. 4.1) on the null hypothesis that the parameters (including the intercept) across the Bundesländer are equal, i.e.  $H_0 : \beta_i = \beta$ ,  $i = 1 \dots B$ . To accomplish this, based on the model in first differences from table 1.1, an observed  $F$ -value of 0.473 is obtained which is distributed as  $F(40, 242)$  under the null. This does not reject poolability across the Bundesländer. Similarly, for testing (ii),  $H_0 : \beta_t = \beta$ ,  $t = 1 \dots T$ , an observed  $F$ -value of 3.845 is obtained. Note, that a structural break, i.e. a change of the parameter vector over time, can only be significant if at least one of the parameter vectors  $\beta_t$  differ from  $\beta$ . Based on the central  $F(100, 182)$  distribution, the null that the parameters across time can be pooled is rejected at the 1 percent level. However, if we are willing to trade some bias for a reduction in variance, some weaker criteria can be used (Baltagi, 1995, p. 54). The null hypothesis is then that the restricted model is better

<sup>14</sup>The rate of return  $r_G$  is obtained from the estimated elasticity as  $r_G = \beta_G Q/G$ .

than the unrestricted model in terms of the trade-off between bias and variance. As a criterion for this test we use the noncentrality parameter  $\lambda$  (Baltagi, 1995, p. 55). From the observed  $\lambda$  value of 1.05, the null hypothesis is neither rejected by the first and second ‘weak’ MSE criterion (Wallace, 1972) nor by the ‘strong’ MSE criterion (Toro-Vizcarrondo and Wallace, 1968).

As the final step of our empirical analysis, in order to shed some light on the underlying structure of the positive correlation between public capital and output, we consider a very simple regression where the growth rate of output in the period 1970-1996, denoted by  $\hat{q}_i$ , is regressed on the growth rates of inputs denoted by  $\hat{k}_i, \hat{l}_i, \hat{g}_i$ , over the same period.

The first regression with only inputs  $\hat{k}_i$  and  $\hat{l}_i$  yields the following result:

$$\hat{q}_i = \begin{matrix} 0.584 & -0.423\hat{k}_i & +0.568\hat{l}_i \\ (0.187) & (0.378) & (0.291) \end{matrix} \quad (1.4)$$

N: 11  $R^2$  : 0.369  $F$  : 2.34

The second regression with inputs  $\hat{k}_i, \hat{l}_i$  and  $\hat{g}_i$  gives:

$$\hat{q}_i = \begin{matrix} 0.075 & -0.416\hat{k}_i & +0.768\hat{l}_i & +0.867\hat{g}_i \\ (0.196) & (0.251) & (0.202) & (0.259) \end{matrix} \quad (1.5)$$

N: 11  $R^2$  : 0.758  $F$  : 7.29\*

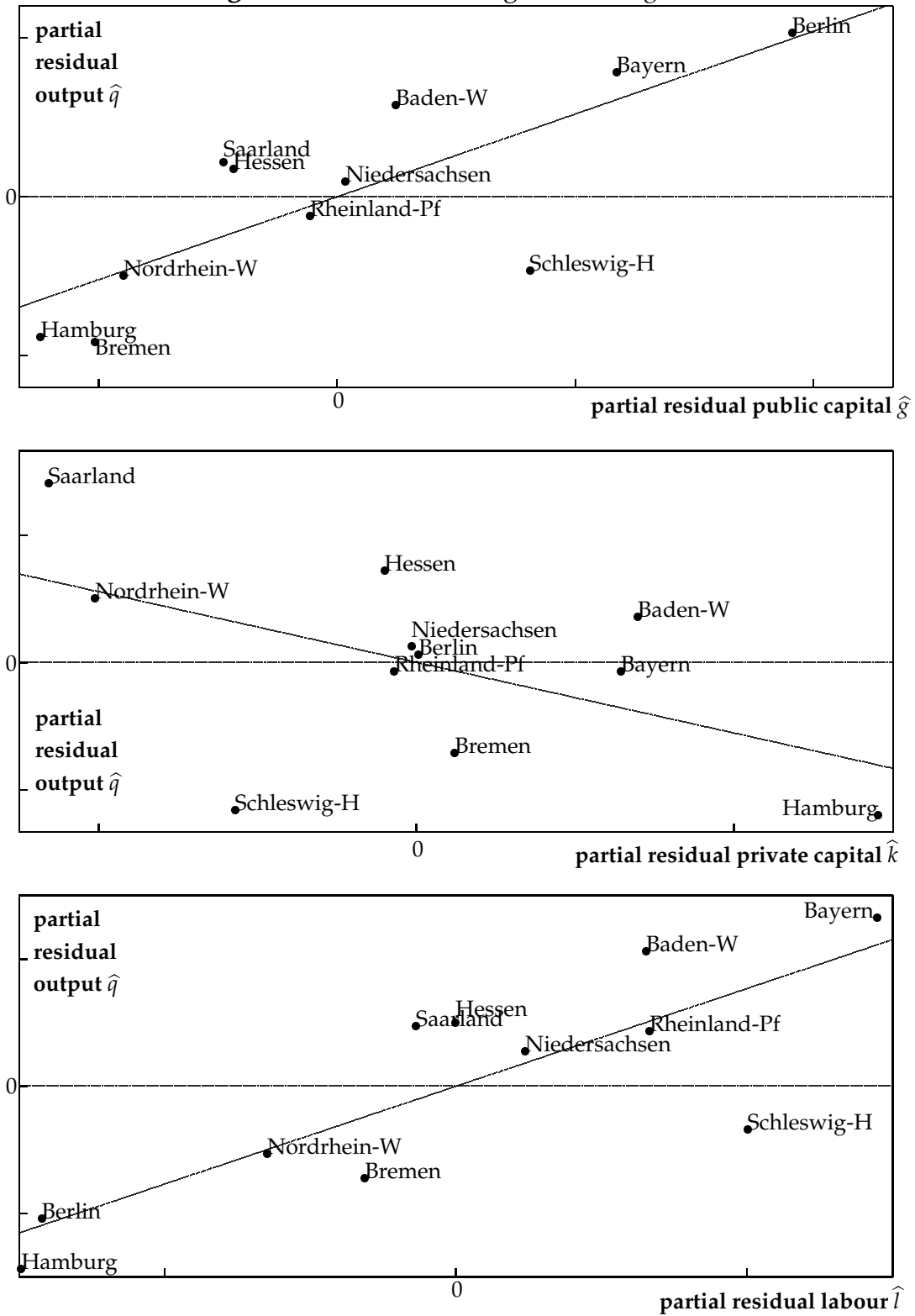
The two regressions show that long-term changes in public capital and labour (in the ‘between’ Bundesländer dimension) are associated with long-term changes in manufacturing sector’s outputs. The coefficients of labour and public capital are in line with the previous results, although the estimates are somewhat higher. This regression also supports the view that the negative coefficient of private capital is not due to the omission of capacity utilisation in (1.1), since over such a long period capacity utilisation is negligible for realised output.

Figure 1.2 presents the partial leverage plots for regression (1.5). Two reference lines are displayed in the plots. One is the horizontal line where the partial residual of  $\hat{q} = 0$ , and the other is the fitted regression of the partial residual of  $\hat{q}$  against the partial residual of the  $\hat{q}$  respective input.<sup>15</sup> The latter has an

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<sup>15</sup>The partial residual of  $\hat{q}$  is obtained by regressing  $\hat{q}$  on  $\hat{k}_i$  and  $\hat{l}_i$ . The partial residual of an input is obtained by regressing this input on the other inputs.

Figure 1.2: Partial Leverage Plots for  $\hat{g}$ ,  $\hat{k}$  and  $\hat{l}$



intercept of 0 and a slope equal to the parameter estimate associated with the explanatory variable in the model.<sup>16</sup> The partial leverage plots reveal that the results of the regression (1.5) are not driven by single influential observations. Except for Schleswig-Holstein all observations contribute positively to the partial correlation between  $\hat{q}$  and  $\hat{g}$ . Also, it can be seen from Figure 1.2 that the insignificance of  $\hat{k}$  is not determined by single influential observations. Interestingly, the Bundesländer Saarland and Nordrhein-Westfalen which experienced the most intense structural change in the manufacturing sector with strongly declining heavy industries during the last 2 decades as well as Hamburg do not fit in a hypothetical positive partial correlation between private capital and output. This hints that the insignificance of private capital could be driven by the structural change in the manufacturing sector which made large parts of the private capital stock obsolete.

We performed also a further regression which is not reported here where the average level of output was regressed on the average levels of the inputs over the period 1970-1996. Thus, the number of observations for this 'between' regression is again 11. It turned out that parameters of all inputs were insignificant. Hence, from this evidence we conclude that differences in levels of public capital or in public capital intensity, defined as the ratio of public capital to labour, do not matter for differences in productivity across the Bundesländer. This is not a surprising finding considering that the level of public capital endowment for each Bundesland also depends on the geographical characteristics of the Bundesland.<sup>17</sup>

From the econometric analysis of this section the following 3 key findings of this study can be recorded. First, and most important, the stylised finding of this study is that public capital is significant for production in the manufacturing sector. This holds for all tested econometric models and specifications. For vari-

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<sup>16</sup>The leverage plot also shows the changes in the residuals for the model with and without the explanatory variable. For a given data point in the plot, its residual without the explanatory variable is the vertical distance between the point and the horizontal line; its residual with the explanatory variable is the vertical distance between the point and the fitted line.

<sup>17</sup>Note also that 3 of the Bundesländer, i.e. Berlin, Bremen and Hamburg, are agglomerated urban Bundesländer which have very different public capital intensities compared to the territorial Bundesländer.

ables in levels, this result is mainly driven by the 'within' variation whereas the 'between' variation does not contribute to it. Thus, differences in public capital intensity can not explain differences in observed levels of output, but differences in changes of public capital can explain differences in changes of output. Furthermore, this correlation between changes of public capital and output holds both in the short-run and in the long-run dimension.

Second, differences in public capital growth can explain about 4 percent of the differences in the manufacturing sector's output growth across Bundesländer over the period 1970 to 1996.

Third and finally, at least for the sample studied here the factor inputs and output appear not to be cointegrated series. For the model with all inputs, i.e. labour  $L$ , private capital  $K$  and public capital  $G$ , the trend stationary model seems to be the appropriate specification.

## 1.4 Conclusions

The starting point of this essay has been Aschauer's (1989a,1989b) public capital hypothesis, which states that the decline in government's infrastructure spending in the US and other major OECD countries during the 1970s and 80s can explain a major part of the observed decline in productivity growth over the same period.

Several methodological improvements to related studies have been incorporated into the analysis in this essay. We have explicitly taken into account four of the most frequent specification issues in the context of time-series cross-section analysis: serial correlation, groupwise heteroscedasticity, cross-sectional correlation and nonstationarity of the data. Furthermore, we have used a specification in the analysis that has avoided a potential simultaneity problem between output and factor inputs. Finally, we have provided tests on the poolability of data and the stability of parameters over time.

In summary, we find a strong positive and significant correlation between public capital and the manufacturing sector's output at the regional level of the 'Bundesländer' in all of the tested specifications.

One tentative conclusion that can be drawn from this finding is that differences in public capital endowment might also explain a part of the still existing productivity gap between manufacturing in East and West Germany. Recent studies (Komar, 2000; Seidel and Vesper, 2000) report that the gap in public capital endowment on a per capita basis between East and West German regions is still about 30 percent,<sup>18</sup> while at the same time productivity of firms located in East Germany is only about two-third of the productivity of firms located in the West. Thus, at least a part of the productivity differences might be also attributed to differences in public capital endowment.

Given the significance of public capital for private production, one potential economic policy implication is whether the process of convergence in public capital endowment between East and West German regions should be accelerated over the next years. At this point, however, we emphasise that the existence of positive effects of public capital on private production is a necessary, but not a sufficient condition for drawing the conclusion that public investment should be boosted in the future. To make this inference, the costs of the public capital provision have to be included in the analysis as well. For instance an increase in public investment may only be possible if tax revenues are also increased. This in turn can give rise to distortions bearing additional costs for the economy. Similarly, if higher public investment is financed by higher governmental debt this may also imply other kinds of additional costs e.g. higher interest rates on capital markets. In this respect, our study has focused only on the necessary condition for increasing the supply of public capital, i.e. the existence of significant and positive effects of public capital on private production.<sup>19</sup>

The obtained estimates of the output elasticity of public capital between 0.42 and 0.65 imply rate of returns between 29 (minimum) and 72 (maximum) percent for our sample. Since these are measures for the return of public capital only for

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<sup>18</sup>The public capital definition in this study excluded public capital funded by the federal government e.g. highways and waterways.

<sup>19</sup>In a more rigorous fashion, the sufficient condition for increasing public investment is that the social net benefit—defined as the sum of social gross benefits (consumer and producer surpluses, positive externalities e.g. spillover effects, etc.) minus the sum of social costs (costs of provision, negative externalities e.g. environmental effects, etc.)—has to be positive.

manufacturing, but do not capture the returns for other economic sectors, they appear to be too high to be a plausible estimate of the 'true' returns of public capital for manufacturing. On the other hand these magnitudes are in line with other studies which have been also conducted for the manufacturing sector e.g. Morrison and Schwartz (1996).

A fundamental problem of both our study and related ones is that there is no certainty whether or not other factors that might also positively contribute to the manufacturing sector's output have been omitted from the analysis. If these factors are positively correlated with public capital but excluded in the regression equation then the expected value of the parameter of public capital will be upward biased. Such a factor could be for instance the stock of knowledge or of the available technology in the manufacturing sector. However, it is very difficult to find plausible measures for these intangible stocks, since these stocks are not directly observed.

One startling result of our study is that private capital turns out to be not significant in most of the tested specifications. From the analysis we could exclude the possibility that this finding is a result of a collinearity between the private and the public capital stock, since the insignificance of private capital holds even if only private capital is included in the regression. As the first glance at Figure 1.1 in section 1.3 revealed, it is very likely that the movements in private capital follow the movements in output with a lag of 2 to 3 years, thus private capital at this aggregate level appears not to be exogenously given, but endogenous determined by developments in output. Another potential explanation for the insignificance of private capital is the intense structural change and technical shift in the manufacturing sector during the last 3 decades which have made large parts of the existing capital stock obsolete.

This puzzling insignificance of private capital in the German manufacturing sector is an interesting starting point for future research. One possibility to address this issue would be to formulate a simultaneous equation model where both private and public capital are endogenously determined. Another promising line for future research is to compare the outcomes of the production, dual cost and profit function approaches as in Vijverberg, Vijverberg and Gamble (1997) who

use time-series data for the US and do not find much agreement between the three approaches. The main advantage of this research strategy is that it opens the avenue to study whether the obtained results are robust with respect to the applied (dual) methodology.

# Appendix Essay 1.

## 1A. Data

**Output Q** Output is measured as gross value-added at market prices of the manufacturing sector in 1991 constant prices aggregated over industries in year  $t$ . These data have been obtained from the series '*National accounts for the Bundesländer*' (engl. transl.), issue 30: 'Entstehung des Bruttoinlandsprodukts in den Ländern der Bundesrepublik Deutschland 1970 bis 1996', which is provided by the Statistical Office of Baden-Württemberg. For years 1991-1996 we obtained updated figures (in mid 2000) from the Statistical Office of Baden-Württemberg. These updated figures have not been published yet.<sup>20</sup>

**Public capital (G)** Public capital is measured as the public gross stock of fixed assets at the ground level ('Bruttoanlagevermögen öffentlicher Tiefbau') at the end of year  $t - 1$ . It is given in 1991 constant prices. It includes capital formation of all levels of government in Germany, i.e. the local governments, the Federal States ('Bundesländer') governments and the Federal Government. The main parts of this stock are roads and highways (about 50 percent), bridges and railways, but also water and sewer systems, dikes and ports, etc. Note that these stocks are measured according to international convention in gross terms because of the very low depreciation rate for these types of fixed assets.

The figures for the public gross stock of fixed assets have been provided by the Statistical Office of Baden-Württemberg from the study group of the 'National accounts of the Bundesländer' and have not been published yet.

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<sup>20</sup>The data used in the analysis are available from the authors upon request.

**Private capital (K)** Private capital is measured as the net stock of fixed assets in the manufacturing sector at the end of year  $t - 1$  in constant prices of 1991. It includes machinery, equipment and construction, and is taken from '*National accounts for the Bundesländer*', issue 29: '*Anlageinvestitionen, Anlagevermögen und Abschreibungen in den Ländern der Bundesrepublik Deutschland 1970 bis 1995*'. This statistical report is also provided by the Statistical Office of Baden-Württemberg from the study group of the '*National accounts of the Bundesländer*'. For years 1991-1996 we obtained revised and updated figures from the Statistical Office of Baden-Württemberg which have not been published by the Statistical Office Baden-Württemberg until now (mid 2000).

**Labour (L)** Labour is measured as the number of employees in the manufacturing sector at the regional level of the Bundesländer. These data have been drawn from the series '*Statistical Yearbook for the Federal Republic of Germany*' published by the Federal Statistical Office in Wiesbaden. These figures are measured each year after the first quarter on 1st April, thus in contrast to private and public capital they do not exactly measure the amount of labour at the end of year  $t - 1$ . Alternatively to this labour input measure, we have also estimated the production function with the number of working hours (only of blue-collar employees, also given for 1st April) as the labour input which we obtained from the same publication mentioned above. The differences in the obtained parameter estimates are rather small, therefore we have refrained from reporting these results.

**Wages (L)** Wages cover both blue- and white-collar employees in the manufacturing sector at the regional level of the Bundesländer. The date of reference is the 1st of April for each year. These data have been obtained from the series '*Statistical Yearbook for the Federal Republic of Germany*' published by the Federal Statistical Office in Wiesbaden. In the empirical analysis, wages are only used to compute the average share of labour in total income. For our sample, this share is about 55 percent.

## Essay 2

# The Contribution of Transport and Human Capital Infrastructure to Local Private Production: A Partial Adjustment Approach

### *Abstract*<sup>1</sup>

*This essay uses a partial adjustment approach to measure the contribution of public infrastructure to local private production. In the first step of the empirical analysis we apply a principal component analysis in order to construct 2 new infrastructure indicators from an array of 7 measures of transport and human capital infrastructure. In the second step the output of different sectors is regressed on private factor inputs and on these 2 infrastructure indicators. Our main finding is that expected long-run equilibrium output in an area of local government will be higher, the better it is endowed with both transport and human capital infrastructure. Moreover, transport and human capital infrastructure appear to be complementary, i.e. raising only transport infrastructure will not yield an increase in private production at the local level.*

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<sup>1</sup>A previous version of this essay was published in Review of Regional Science (2001), vol. 21, 91-108.

## 2.1 Introduction

This study examines the role of publicly provided infrastructure for economic development at the local level of the 327 German counties ('Kreise'). Our essay aims at testing empirically the following two ideas. The first postulates that because the main part (about 60-70 percent) of infrastructure is provided by local governments, the main benefits from infrastructure might emerge at the local rather than at the regionally aggregated level (see also Seitz, 1995). Secondly, empirical studies using infrastructure stock measures in monetary terms implicitly assume that infrastructure stocks are homogenous across regions.

However, this assumption is quite often not particularly plausible. Consider, for instance, two regions of the same geographical size, the same population, economy, etc. Suppose both have accumulated a transport infrastructure stock worth 1 billion Euro. However, should one region be geographically flat while the other is mountainous then the productivity of a 1 billion infrastructure stock might be higher in the flat region than in the mountainous one. Thus, in this case public capital can not be regarded as homogenous and therefore comparable across regions. In contrast to this, if transport infrastructure is measured in terms of accessibility, e.g. travel distance to the nearest motorway from a given region, such a measure is comparable across regions even if regions have different geographic characteristics.

Another contribution of this study to the existing empirical literature on the effects of infrastructure is that we simultaneously consider both transport and human capital infrastructure. The importance of the latter type of infrastructure for economic development is stressed in models of 'new growth theories' (e.g. Barro and Sala-I-Martin, 1995; Lucas, 1988).

Previous studies such as Bröcker (1989) which used similar data, i.e. infrastructure indicators at the local level, have either not been based on a local production function or have omitted important factor inputs in the production function such as private capital (e.g. Biehl, 1986). The latter approach is problematic due to a potential omitted variable bias (Greene, 2000, p. 334).

In our empirical analysis we use an array of 7 infrastructure measures, describing the availability of transport and human capital infrastructure at the local

level of the counties. From this set of 7 infrastructure measures, using Principal Component Analysis (PCA), we construct 2 new infrastructure indicators as linear combinations of the original 7 infrastructure measures. These 2 new indicators explain about 64 percent of the variation in infrastructure endowment across the counties. Moreover, specifying a Cobb-Douglas production function within a partial adjustment framework we regress output of different sectors, e.g. *manufacturing*, *services* and *trade & transport*, on private factor inputs and on these 2 infrastructure indicators.

We find that differences in output across counties can be explained by differences in overall endowment with infrastructure. This finding is most pronounced with regard to the *trade & transport* and the *service* sectors. Unexpectedly, we do not find evidence of an effect of infrastructure on production in *manufacturing*. Moreover, for regions well endowed with transport but with poor human capital infrastructure we do not find significant effects of infrastructure on output. Thus, it appears that transport infrastructure alone is not sufficient for higher output in a given county. We interpret this finding as an indication that human capital and transport infrastructure are complementary infrastructure components, at least for the sample studied here.

Finally, tests for spatial independence of residuals in the empirical analysis are performed. However, it turns out that spatial dependence of residuals is not significant, thus the usual econometric techniques such Ordinary Least Squares (OLS) or Seemingly Unrelated Regression (SUR) are applicable to our estimation problem.

The remainder of this essay is organized as follows. Section 2.2 describes the partial adjustment model which builds the basis of our approach. Section 2.3 presents the results of the empirical analysis. Section 2.4 concludes.

## 2.2 Partial Adjustment Model

To begin with, let us assume that production  $Q_{it}$  in county  $i$  at time  $t$  can be described as

$$Q_{it} = f(A(t, INFRA_{it}), K_{it}, L_{it}), \quad i = 1 \dots N, \quad t = 1 \dots T, \quad (2.1)$$

where  $Q_{it}$  denotes output,  $K_{it}$  private capital,  $L_{it}$  labour input and  $A(t, INFRA_{it})$  denotes a technical efficiency parameter depending both on time  $t$  and an index of the public infrastructure stock denoted by  $INFRA_{it}$ . Specifying a Cobb-Douglas functional form for the production function (2.1) and assuming a Hicks-neutral form for  $A(\dots)$ , we get

$$Q_{it} = A(t, INFRA_{it}) L_{it}^{\alpha_L} K_{it}^{\alpha_K}, \quad (2.2)$$

where  $\alpha_L$  and  $\alpha_K$  denote the elasticities of output  $Q$  with respect to  $L$  and  $K$ .

Finally, assuming that  $A(t, INFRA_{it})$  has the following functional form

$$A(t, INFRA_{it}) = A_0 \exp(\alpha_t t) INFRA_{it}^{\alpha_{INFRA}}, \quad (2.3)$$

where  $A_0$  is the initial value of  $A(\dots)$  at time  $t_0$ , and dividing (2.2) by  $L_{it}$ , we get

$$q_{it} = A_0 \exp(\alpha_t t) INFRA_{it}^{\alpha_{INFRA}} k_{it}^{\alpha_K} L_{it}^{\tilde{\alpha}_L}, \quad (2.4)$$

where small capitals denote variables in terms of the labour input  $L$  and  $\tilde{\alpha}_L$  is defined as  $\tilde{\alpha}_L = \alpha_L + \alpha_K - 1$ . Note that  $\tilde{\alpha}_L$  will equal zero if returns to scale are constant with respect to inputs  $L$  and  $K$ . This approach has the advantage that it *a priori* does not put on (2.2) any restriction with respect to returns to scale.

Our empirical approach is based on a partial adjustment model. Suppose that long-run equilibrium output  $q_{it}^*$  in county  $i$  is given by (2.4). Taking logarithms of (2.4) we obtain

$$\ln q_{it}^* = \alpha_0 + \alpha_{INFRA} INFRA_{it} + \alpha_K \ln k_{it} + \alpha_K \ln L_{it} + \epsilon_{it}, \quad (2.5)$$

where  $\alpha_0 = \ln A_0$  and  $\epsilon_{it}$  is an i.i.d. random variable with variance  $\sigma_\epsilon$ .

The adjustment process can be described by the following equation (Greene, 2000, p. 722)

$$\ln q_{it} - \ln q_{it_0} = (1 - \lambda)(\ln q_{it}^* - \ln q_{it_0}), \quad (2.6)$$

where  $q_{it_0}$  denotes initial output at time  $t_0$ .

Solving (2.6) for  $\ln q_{it}$  and inserting (2.5) for  $q_{it}^*$  we obtain the baseline model for our empirical analysis

$$\begin{aligned} \ln q_{it} = & \lambda \ln q_{it_0} + \alpha_0(1 - \lambda) + \alpha_{INFRA}(1 - \lambda) INFRA_{it} + \alpha_K(1 - \lambda) \ln k_{it} \\ & + \tilde{\alpha}_L(1 - \lambda) \ln L_{it} + (1 - \lambda)\epsilon_{it}. \end{aligned} \quad (2.7)$$

This equation can also be estimated without restrictions as

$$\ln q_{it} = \lambda q_{it_0} + \alpha_0' + \alpha_{INFRA} INFRA_{it} + \alpha_K' \ln k_{it} + \tilde{\alpha}_L' \ln L_{it} + \epsilon_{it}'. \quad (2.8)$$

From (2.8), the short-run elasticities can be obtained from estimates of  $\alpha_{INFRA}'$ ,  $\alpha_K'$ , and  $\alpha_L'$  (from  $\tilde{\alpha}_L' - \alpha_K' + 1$ ), whereas long-run elasticities can be calculated either from these estimates as  $\alpha_{INFRA} = \alpha_{INFRA}' / (1 - \lambda)$ ,  $\alpha_K = \alpha_K' / (1 - \lambda)$ , and  $\alpha_L = \alpha_L' / (1 - \lambda)$  or can be obtained directly from (2.7) by using nonlinear methods.

This partial adjustment specification proves to be particularly useful for our analysis. Suppose that there is some unobserved heterogeneity in output  $q_{it}$  across counties, for instance due to the particular locations of counties, or due to different manufacturing technologies, etc. If panel data are available, one can control for this unobserved heterogeneity by including fixed or random individual effects. In our case with cross-sectional data, however, we can presume that if this unobserved individual county-specific attribute was already present at time  $t_0$ , then it might be reflected in output  $q_{it_0}$  as well. Thus, including the lagged dependent variable  $q_{it_0}$  as a right-hand side variable allows us to control for such time-invariant unobserved heterogeneity.

## 2.3 Empirical Implementation

### 2.3.1 Description of the Data

Our sample comprises of the 327 counties ('Kreise') in West Germany. A county itself usually contains several townships ('Gemeinden'). The next higher regional level above counties is the 166 'labour market regions' ('Arbeitsmarktregionen'). Indicators we use in order to describe local public infrastructure endowment are only available at the level of these 'labour market regions'. Thus, on average, a labour market region consists of about 2 counties. We have merged these two data sets with different regional levels. In the final data set the observations for the variables are at the county level. However, on average 2 counties will have the same values for the infrastructure indicators.

*Table 2.1: Labels of Variables*

<b>Label</b>	<b>Short Description</b>
<i>Transport Infrastructure</i>	
<i>Motorway</i>	accessibility of motorways, 1989
<i>FreightTr</i>	accessibility of freight transfer railway stations, 1989
<i>Airport</i>	accessibility of regional airports, 1989
<i>ICTrain</i>	accessibility of inter-city express trains, 1988
<i>Human Capital Infrastructure</i>	
<i>VocTrain</i>	vocational training in future-oriented branches, 1988
<i>Coll&amp;Uni</i>	students at colleges & universities, 1988
<i>ScienceP</i>	availability of science parks, 1988
<i>Production Function Variables</i>	
<i>Q</i>	Output measured as gross value added at factor costs, different sectors, 1980, 1988
<i>K</i>	private capital stock of manufacturing sector, 1988
<i>L</i>	number of employees, different sectors, 1987

Table 2.1 shows a list of variables used in the analysis. Output ( $Q$ ) in counties is measured as gross value added at *factor costs* in 1988 and has been drawn from a publication of the Statistisches Landesamt Baden-Württemberg (1995). The difference between gross value added at market prices and at factor costs is, that the latter is calculated from the former by subtracting the difference between indirect production taxes and governmental subsidies. The difference is on average only about 1-2 percent.

Table 2.2 displays some descriptive statistics of the infrastructure variables. Note that for some of the variables, e.g. *Highway*, *FreightTr* or *Airport*, the median is substantially different from the mean, thus the distribution of these variables appears to be skewed.

We estimate the contribution of public infrastructure to local private production also separately for 3 sectors, i.e. *manufacturing*, *trade & transport*, and *services*. Note, that the output measure of *all* sectors also includes the *agriculture, forestry & fishing* as well as the *governmental* sector.

Output for the 3 sectors *manufacturing*, *trade & transport*, and *services* is also measured as gross value added at *factor costs* in 1988 prices. In the publication mentioned above, however, only gross value added at *market prices* is reported for the single sectors. Thus, we computed the difference between gross value added at *factor costs* and *market prices* for each sector from the difference given for total gross value added, and allocated this difference according to the share of each sector in total gross value added.

Unfortunately, our measure for labour ( $L$ ), given as numbers of employees, is only available for the year 1987 and not for 1988. It has been drawn from the joint publication of the Federal States Statistical Offices in Germany titled 'Erwerbstätigenrechnung des Bundes und der Länder, Erwerbstätige in den kreisfreien Städten und Landkreisen in der Bundesrepublik Deutschland 1980, 1987, 1990-1993', Heft 2.

The private capital stock ( $K$ ) of the *manufacturing* sector in 1988 at the county level has been obtained from Deitmar (1993). We have measures neither for the total capital stock in counties nor for the capital stocks of the *trade & transport* or *service* sectors. However, since *manufacturing* is the main part of the total stock we

**Table 2.2:** *Descriptive Statistics of Infrastructure Variables*

	Mean	Std.	C.V.	Min	Max	Median
<i>Motorway</i>	90.49	22.767	25.16	0	100	0
<i>FreightTr</i>	81.67	27.905	34.17	0	100	4.8
<i>Airport</i>	56.55	43.385	76.71	0	100	19.4
<i>ICTrain</i>	65.36	27.723	42.42	0	100	75.6
<i>VocTrain</i>	105.13	13.275	12.63	73	133	105
<i>Coll&amp;Uni</i>	145.4	123.4	84.86	0	568	168
<i>ScienceP</i>	4.60	5.52	119.78	0	20	2

presume that it is a reasonable approximation for the latter. We also approximate the private capital stock of the *trade & transport* or *service* sectors with the capital stock of manufacturing. This allows us to conclude whether or not output of these sectors are related to the *manufacturing* sector.

The indicators for public infrastructure endowments of the labour market regions are taken from Gatzweiler, Irmen and Janich (1991). As shown in Table 2.1, our first 4 indicators describe counties' endowment with transport infrastructure. We employ these indicators in order to describe the accessibility of a county by means of transport. A short description of the 4 indicators is also provided in Table 2.1.

The variable *Motorway* measures the percentage of employees in a given 'labour market region', whose places of work are located in a county closer than 30 minutes travel by car to the nearest motorway (or similar long-distance road). Variable *FreightTrans* measures the percentage of manufacturing sector firms in a 'labour market region' located in a township closer than 45 minutes travel by lorry to the nearest freight transfer railway station. Variable *Airport* measures the percentage of firms in a given 'labour market region' which are located in a township closer than 45 minutes travel by car to the nearest regional airport. Variable *ICTrain* gives the percentage of people in a labour market region which have access to inter-city express train stations within a travel distance of 30 minutes by car.

Furthermore, the data also contain measures of counties' infrastructure with

regard to human capital. Variable *VocTrain* is a combined indicator which is based both on the availability of vocational training in general and on the number of training opportunities in future-oriented industries such computing, biotechnology, etc. Variable *Coll&Uni* is a combined indicator both for the availability of colleges & universities and for the percentage of students at colleges & universities in a given region studying engineering, computing, mathematics or natural sciences. Finally, variable *ScienceP* is a combined indicator for the availability of science parks and science & technology transfer service centres in a given region. For further details how these indicators are constructed, see Gatzweiler et al. (1991).

### 2.3.2 Analysis

The structure of the empirical analysis is as follows. First, we analyse the relationships between the various infrastructure indicators using principal component analysis (PCA). In a second step, we apply the PCA to construct 2 new indicator variables, i.e. we use the first two principal components as new indicators. Finally, in a third step we regress output of several sectors on private factor inputs and on these 2 principal components.

**Table 2.3:** *Correlations within Transport Infrastructure Variables*

	<i>Motorway</i>	<i>FreightTr</i>	<i>Airport</i>	<i>ICTrain</i>
<i>Motorway</i>	1.000	0.495	0.308	0.442
<i>FreightTrans</i>	0.495	1.000	0.424	0.407
<i>Airport</i>	0.308	0.424	1.000	0.407
<i>ICTrain</i>	0.442	0.407	0.407	1.000

**Table 2.4:** *Correlations within Human Capital Infrastructure Variables*

	<i>VocTrain</i>	<i>Coll&amp;Uni</i>	<i>ScienceP</i>
<i>VocTrain</i>	1.000	0.417	0.573
<i>Coll&amp;Uni</i>	0.417	1.000	0.572
<i>ScienceP</i>	0.573	0.572	1.000

**Table 2.5:** Correlations between Transport and Human Capital Infrastructure

Variables	<i>VocTrain</i>	<i>Coll&amp;Uni</i>	<i>ScienceP</i>
<i>Motorway</i>	0.349	0.330	0.294
<i>FreightTrans</i>	0.485	0.433	0.409
<i>Airport</i>	0.507	0.345	0.549
<i>ICTrain</i>	0.462	0.528	0.510

Table 2.3 displays the correlations between the various indicators for transport infrastructure. It appears that all indicators are positively correlated. The correlation is highest with about 0.5 between the variables *Motorway* and *Freightransfer*, and lowest with about 0.3 between the variables *Motorway* and *Airport*.

Table 2.4 gives the correlations between the indicators for human capital infrastructure. Indicators *VocTrain*, *Coll&Uni*, and *ScienceP* are all positively correlated. The correlation is highest with about 0.57 between *VocTrain* and *ScienceP*.

Table 2.5 presents the correlations between *transport* and *human capital* indicators. Again, we find that all correlations are positive. We observe the lowest correlations between *transport* and *human capital* indicators for the variable *Motorway*, and the highest for the variable *ICTrain*.

**Table 2.6:** Eigenvalues of Principle Components Analysis of Infrastructure Variables

	Eigenvalue	Difference	Proportion	Cumulative
<i>Infra<sub>1</sub></i>	3.655	2.799	0.522	0.522
<i>Infra<sub>2</sub></i>	0.855	0.134	0.122	0.644
<i>Infra<sub>3</sub></i>	0.721	0.176	0.103	0.748
<i>Infra<sub>4</sub></i>	0.545	0.067	0.077	0.826
<i>Infra<sub>5</sub></i>	0.478	0.064	0.068	0.894
<i>Infra<sub>6</sub></i>	0.415	0.087	0.059	0.953
<i>Infra<sub>7</sub></i>	0.328	.	0.047	1

**Table 2.7:** Eigenvectors of Principle Components Analysis of Infrastructure Variables

	<i>Infra<sub>1</sub></i>	<i>Infra<sub>2</sub></i>	<i>Infra<sub>3</sub></i>	<i>Infra<sub>4</sub></i>	<i>Infra<sub>5</sub></i>	<i>Infra<sub>6</sub></i>	<i>Infra<sub>7</sub></i>
<i>Motorway</i>	0.322	0.732	0.050	0.303	0.045	0.513	0.036
<i>FreightTr</i>	0.376	0.380	0.255	-0.615	0.122	-0.432	-0.263
<i>Airport</i>	0.367	-0.297	0.532	0.315	0.538	-0.111	0.304
<i>ICTrain</i>	0.390	0.064	-0.411	0.532	-0.189	-0.585	-0.119
<i>VocTrain</i>	0.397	-0.193	0.305	-0.114	-0.767	0.109	0.316
<i>Coll&amp;Uni</i>	0.377	-0.121	-0.617	-0.367	0.258	0.156	0.487
<i>ScienceP</i>	0.411	-0.419	-0.084	-0.014	0.056	0.400	-0.697

This particular pattern of correlations between the infrastructure indicators lends itself to a principle component analysis in order to reduce the complexity of information within the total set of indicators. Table 2.6 and Table 2.7 give the results for the principal components analysis (PCA) of the correlation matrix for all 7 indicators. In Table 2.6, the Eigenvalues (characteristic roots) of the PCA are presented, and in 2.7 the associated Eigenvectors (characteristic vectors) are displayed.

The PCA is based on the following decomposition (Greene, 2000, p. 36)

$$\mathbf{V}'\mathbf{X}'\mathbf{X}\mathbf{V} = \mathbf{\Delta},$$

where  $\mathbf{V}'$  is a  $(k \times k)$  matrix of the  $(\mathbf{v}_1, \dots, \mathbf{v}_k)$  Eigenvectors of  $\mathbf{X}'\mathbf{X}$ , where  $\mathbf{X}$  is a  $(n \times k)$  data matrix (with  $n$  observations on  $k$  variables) and  $\mathbf{\Delta}$  is a  $(k \times k)$  diagonal matrix of associated Eigenvalues. The  $j^{th}$   $(n \times 1)$  principal component  $\mathbf{p}_j$  of the  $(n \times k)$  matrix  $\mathbf{P} = \mathbf{X}\mathbf{V}$  of principal components is thereby defined as

$$\mathbf{p}_j = \mathbf{X}\mathbf{v}_j, \quad j = 1, \dots, k.$$

It is worth noting, that the first Eigenvector in Table 2.7, associated with the first Eigenvalue in Table 2.6, can already explain 52.2 percent of the variation within the infrastructure variables. Moreover, the second Eigenvector can explain 12.2 percent of the total variation. Hence, the first two Eigenvectors together can explain about 64 percent of the total variation within all infrastructure indicators.

The coefficients of the Eigenvectors in Table 2.7 reflect the contribution of each single indicator to a corresponding principal component. Thus, all indicators contribute with a positive sign to the first component. Consequently, counties with high values on these indicators will also have a high score for the first component.

With respect to the second principal component, the variables *Motorway* and *FreightTrans* contribute with a positive sign, whereas the variables *Airport*, *VocTrain*, *CollTrain* and in particular variable *ScienceP* contribute with a negative sign. Thus, counties with high values for the variables *Motorway* and *FreightTrans* but with relatively low or close to zero values for the variables *Airport*, *VocTrain*, *CollTrain* and *ScienceP* will have a high positive score on this second principal component.

**Table 2.8:** Correlations of Infrastructure Variables and Principal Components with Output

	Correlations of Output <sup>1)</sup> q <sub>i</sub>			
	All sectors	Manufac- turing	Trade & Transp.	Servi- ces
<i>with original infrastructure variables</i>				
<i>Highway</i>	0.241***	0.105**	0.197***	0.160***
<i>FreightTr</i>	0.308***	0.183***	0.218***	0.140**
<i>Airport</i>	0.368***	0.162***	0.196***	0.210***
<i>ICTrain</i>	0.459***	0.328***	0.210***	0.155***
<i>VocTrain</i>	0.452***	0.184***	0.322***	0.330***
<i>Coll&amp;Uni</i>	0.322***	0.228***	0.186***	0.142***
<i>ScienceP</i>	0.459***	0.249***	0.328***	0.255***
<i>with Principal Components</i>				
<i>Infra<sub>1</sub></i>	0.521***	0.288***	0.330***	0.278***
<i>Infra<sub>2</sub></i>	-0.114**	-0.052	-0.043	-0.076

<sup>1)</sup> measured as gross value added.

significant \* at 10 %, \*\* at 5 %, \*\*\* at 1 %.

On the other hand, counties with high values for the variables *Airport*, *VocTrain*, *CollTrain* and *ScienceP* but low values for the variables *Motorway* and/or *FreightTrans* will have a high negative score for the second component. We interpret this as an indication that the second component reflects the contrast of counties with either good quality transport infrastructure (except airports) but with a relatively low human capital infrastructure or vice versa.

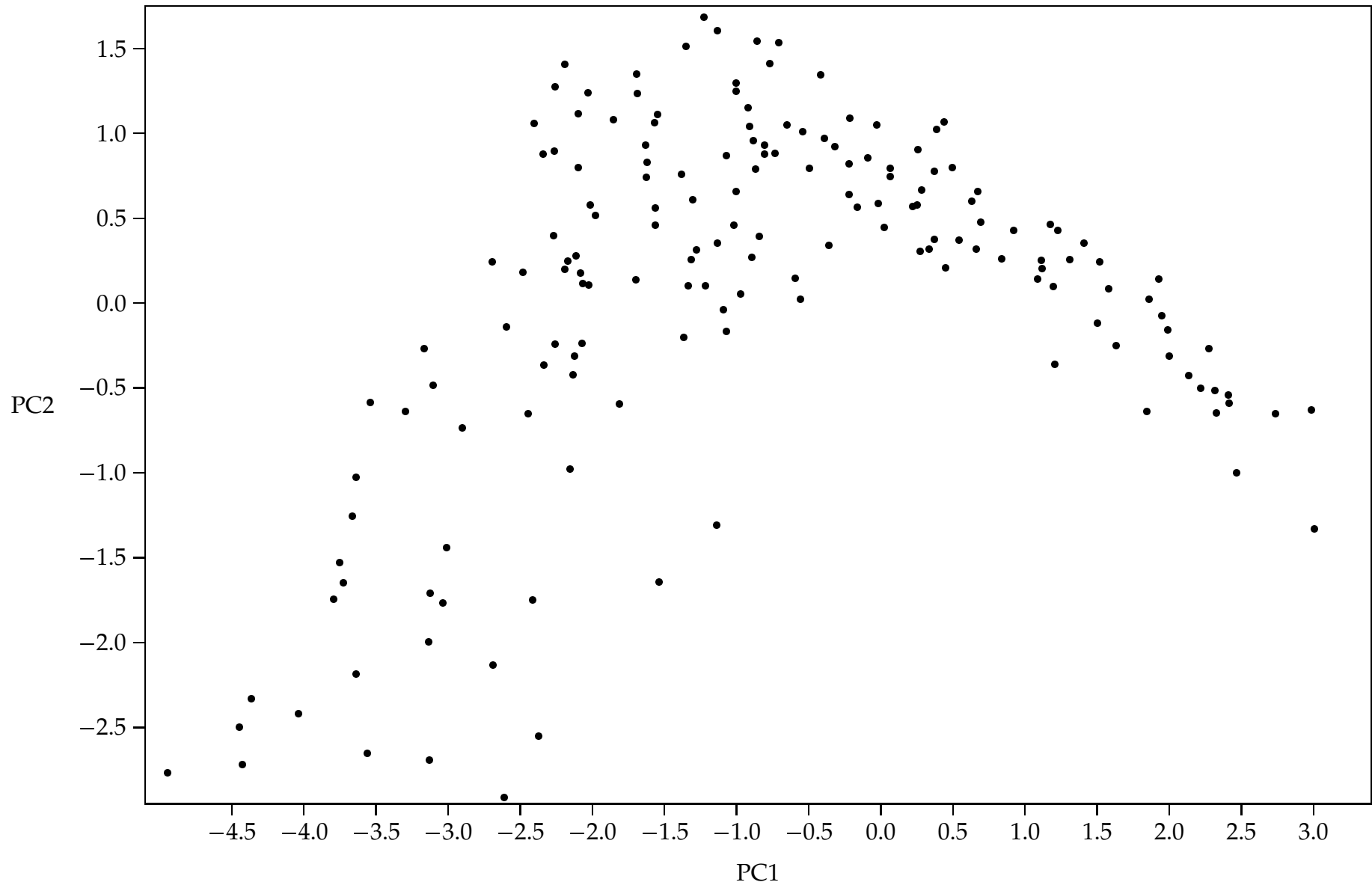
Figure 2.1 graphs the first principal component versus the second. The single observations in this graph are the labour market regions. Note, that by construction the principal components are uncorrelated. Labour market regions well endowed with infrastructure will have high values on the first component, thus they are located on the right-hand side of the graph. Observations with high values on the second components are located in the upper-half of the graph.

Table 2.8 gives the correlations of the original indicator variables and the principal components with output, measured for the different sectors. All infrastructure variables are significant and positively correlated with these output measures. This pattern also justifies the application of PCA in the regression analysis because due to this collinearity between infrastructure variables it would be difficult to get precise estimates of the contribution of single variables. The first principal component, which we label as *Infra<sub>1</sub>*, is significantly correlated with output measures for all sectors, whereas the second principal component, which we label as *Infra<sub>2</sub>*, is not.

Table 2.9 displays the results of the regression analysis. The estimations have been carried out using LIMDEP 7.0. Equation (2.8) has been estimated both in the unrestricted and the restricted form, where the long-run parameters are directly estimated. The unrestricted specification has been estimated with linear OLS. The restricted specification has been estimated both with nonlinear Ordinary Least Squares (OLS) and with nonlinear Seemingly Unrelated Regression (SUR). The correlations of the residuals across equations are shown in Table 2.10. Some of these correlations, e.g., between equations *total* and *manufacturing*, and *total* and *services* are positive and quite significant. Thus, we expect a gain in efficiency from using SUR compared to OLS.

We have added both Länder dummy variables and a dummy variable indicat-

*Figure 2.1: Plot of Principal Components for Labour Market Regions*



**Table 2.9: Regression Results for Different Sectors**

	OLS <sup>1)3)</sup>		nonlinear OLS <sup>1)3)</sup>		nonlinear SUR <sup>1)</sup>	
<i>Dependent Variable: Output <math>q_i</math></i>						
<b>All Sectors</b>						
<i>Dummy var.<sup>2)</sup></i>	<i>Länder***</i>		<i>Länder***</i>		<i>Länder***</i>	
<i>DCity</i>	0.023	(1.69)	0.062	(1.945)	0.013	(0.26)
<i>Intercept</i>	1.413	(6.69)	3.807	(22.62)	4.102	(14.21)
$\ln q_{it_0}$	0.629	(10.63)	0.629	(10.63)	0.782	(28.98)
$\ln k_i$	0.047	(3.91)	0.120	(3.84)	0.184	(4.44)
$\ln L_i$	0.045	(4.80)	0.123	(4.63)	0.150	(3.67)
<i>Infra<sub>1</sub></i>	0.009	(3.47)	0.024	(3.05)	0.029	(2.42)
<i>Infra<sub>2</sub></i>	-0.005	(-0.96)	-0.013	(-0.97)	-0.013	(-0.56)
$R^2$	0.759		0.759		0.750	
White $\chi^2$ (76) <sup>2)</sup>	91.5		91.5		94.0**	
<b>Manufacturing Sector</b>						
<i>Dummy var.<sup>2)</sup></i>	<i>Länder***</i>		<i>Länder***</i>		<i>Länder***</i>	
<i>DCity</i>	0.041	(2.01)	0.093	(2.27)	0.080	(1.44)
<i>Intercept</i>	0.970	(3.92)	2.191	(6.58)	2.183	(6.27)
$\ln q_{it_0}$	0.557	(8.75)	0.557	(8.75)	0.695	(23.74)
$\ln k_i$	0.169	(5.65)	0.382	(4.74)	0.376	(6.25)
$\ln L_i$	0.072	(4.66)	0.163	(4.51)	0.261	(5.64)
<i>Infra<sub>1</sub></i>	-0.009	(-1.92)	-0.020	(-1.71)	-0.041	(-2.79)
<i>Infra<sub>2</sub></i>	0.041	(0.32)	0.006	(0.32)	0.010	(0.38)
$R^2$	0.705		0.709		0.699	
White $\chi^2$ (76) <sup>2)</sup>	136.8***		136.8***		124.1***	
<b>Trade &amp; Transport Sector</b>						
<i>Dummy var.<sup>2)</sup></i>	<i>Länder***</i>		<i>Länder***</i>		<i>Länder***</i>	
<i>DCity</i>	0.011	(0.54)	0.024	(0.56)	0.016	(0.35)
<i>Intercept</i>	1.948	(5.45)	4.210	(19.28)	4.224	(15.74)
$\ln q_{it_0}$	0.537	(5.56)	0.537	(5.56)	0.601	(14.95)
$\ln k_i$	-0.023	(-1.83)	-0.050	(-1.74)	-0.032	(-0.90)
$\ln L_i$	0.041	(3.08)	0.088	(2.83)	0.070	(2.11)
<i>Infra<sub>1</sub></i>	0.021	(3.56)	0.045	(2.95)	0.057	(4.27)
<i>Infra<sub>2</sub></i>	-0.007	(-0.95)	-0.016	(-0.97)	-0.017	(-0.87)
$R^2$	0.534		0.534		0.530	
White $\chi^2$ (76) <sup>2)</sup>	146.2***		146.2***		145.8***	
<b>Service Sector</b>						
<i>Dummy var.<sup>2)</sup></i>	<i>Länder***</i>		<i>Länder***</i>		<i>Länder***</i>	
<i>DCity</i>	-0.080	(-4.88)	-0.276	(-3.49)	-0.264	(-3.77)
<i>Intercept</i>	1.679	(6.11)	5.829	(12.13)	5.613	(13.87)
$\ln q_{it_0}$	0.712	(11.37)	0.712	(11.37)	0.724	(21.44)
$\ln k_i$	0.029	(2.48)	0.101	(2.30)	0.175	(3.54)
$\ln L_i$	-0.007	(-0.57)	-0.023	(-0.58)	-0.087	(-1.92)
<i>Infra<sub>1</sub></i>	0.018	(4.20)	0.063	(3.96)	0.073	(4.48)
<i>Infra<sub>2</sub></i>	0.002	(0.21)	0.006	(0.21)	0.004	(0.15)
$R^2$	0.603		0.603		0.596	
White $\chi^2$ (76) <sup>2)</sup>	57.4		57.4		58.7	
Condition-num.	126.7		52.4		191.4	

Number of observations: 327 for each equation

1) Asymp. t-values are given in parentheses.

2) significant \* at 10 %, \*\* at 5 %, \*\*\* at 1 %.

3) White's (1980) heteroscedasticity robust t-values.

**Table 2.10: Cross Equation Correlations from OLS Table 2.9**

	All sectors	Manu- facturing	Trade & Transp.	Ser- vices
All sectors	1.000	0.703	0.339	0.529
Manufacturing	0.703	1.000	-0.075	0.006
Trade & Transport	0.339	-0.075	1.000	0.262
Services	0.529	0.006	0.262	1.000

ing whether or not a county is a self-administrated city to all equations. White's (1980) heteroscedasticity test has been applied to all regressions. Homoscedasticity of residuals is rejected except for the *trade & transport* equation. Thus, White's (1980) heteroscedasticity robust standard errors have been used for calculating the  $t$  values for the OLS estimations. The reported condition numbers with values greater than 20 may indicate a potential problem of multicollinearity for the estimations (Judge et al., 1985, p. 902).

The fit as indicated by  $R^2$  is remarkably good for all equations. Several key findings emerge from Table 2.9. First of all,  $Infra_1$  is significant for *all*, the *service* and *trade & transport* sectors, but surprisingly not for the *manufacturing* sector. In contrast to this,  $Infra_2$  is not significant for private production with regard to all sectors. Second, the estimates of the adjustment parameter  $\lambda$  are positive and significant for all equations. Values of  $\lambda$  of about 0.6 to 0.7 imply that the halfway ( $\lambda = 0.5$ ) between the actual value and the long-run equilibrium value of output was reached within the 8 years of observation. Third, the Länder dummy variables are significant for all equations. Hence, there are systematic differences in output of industries and branches across the Bundesländer. Fourth, it turns out that our measure for private capital approximated as the private capital stock of the manufacturing sector, is related to *all*, *manufacturing* and *service* sectors output, but not to output of the *trade & transport* sector. Fifth, it is worth noting that the city dummy variable is positive and significant for *manufacturing*, but negative and significant for the *service* sector. Sixth and finally, the positive and significant coefficients for  $\ln L$  show that economies of scale and/or agglomeration

economics are important. However, this does not apply for the *service* sector.

**Table 2.11:** Tests on normality of residuals from OLS estimations, Table 2.9

	All sectors	Manu- facturing	Trade & Transp.	Ser- vices
Shapiro-Wilk <i>W</i>	0.982	0.976**	0.935***	0.929***
Jarque-Bera <i>JB</i>	98.1***	86.1***	2172.3***	1014.3***

significant \* at 10 %, \*\* at 5 %, \*\*\* at 1 % .

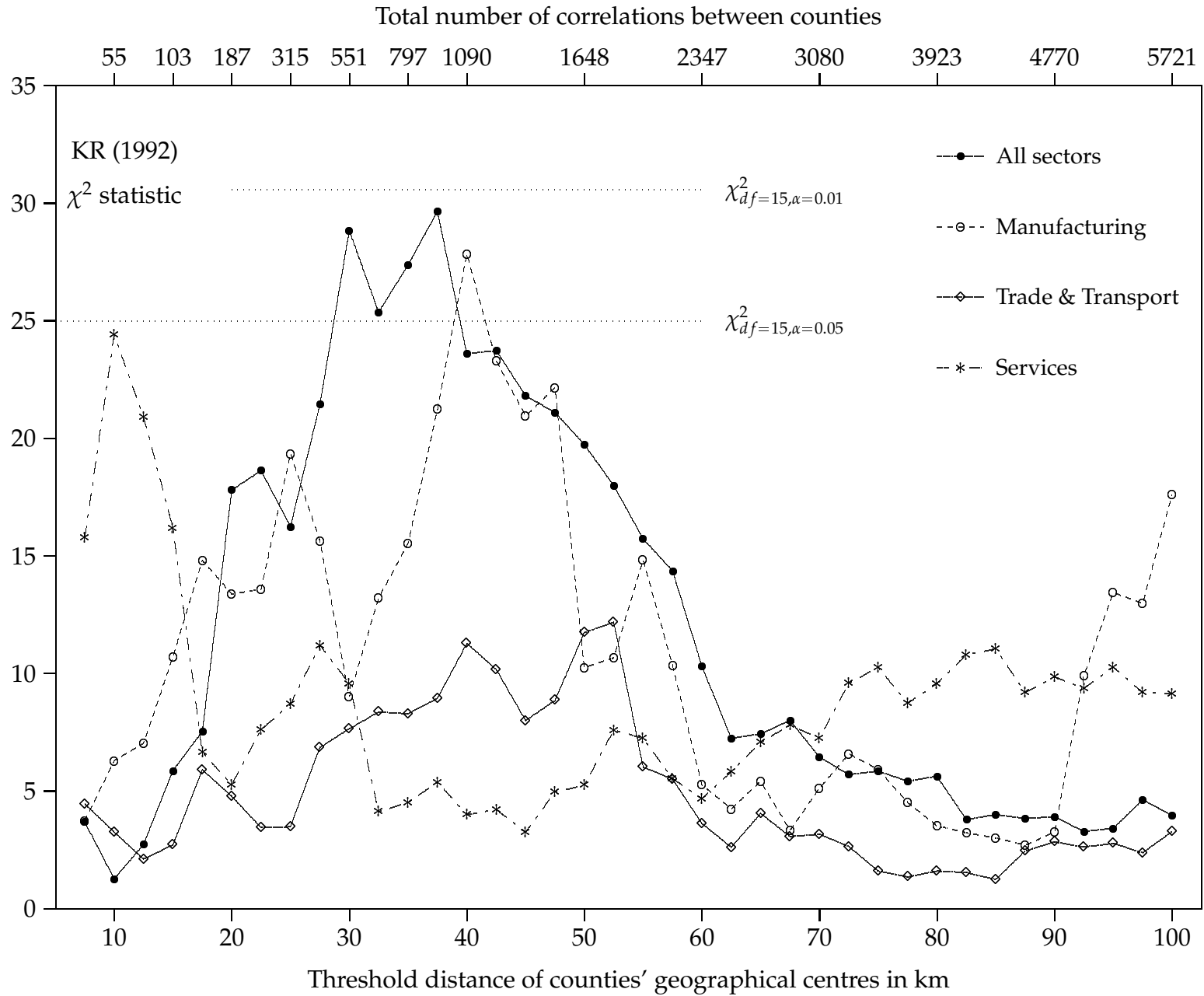
To conclude the empirical analysis, we finally examine whether or not the residuals of the estimations exhibit evidence of spatial dependence. As outlined in Schulze (1998), the first step in the analysis of spatial dependence should consist of a test on the normality of residuals. Table 2.11 provides the results both of the Shapiro-Wilk and the Jarque-Bera tests for all equations. The Jarque-Bera *JB* statistic is distributed  $\chi^2$  with 2 degrees of freedom, thus the critical value for  $p=0.01$  is 9.21. The null hypothesis of normality is rejected for all equations by the Jarque-Bera test as well as by the Shapiro-Wilk statistic, except for the *all* sectors equation.

Hence, in this case the test procedure for spatial dependence suggested by Moran (1950), and extended by Cliff and Ord (1972), appears not to be appropriate. However, as an alternative the *KR* test proposed by Kelejian and Robinson (1992) is still applicable. In contrast to Moran’s *I* this test neither requires the model to be linear nor the disturbance terms to be normally distributed.

Applying this test it turns out—as it would also be the case with Moran’s *I*—that the outcome of the *KR* statistic depends on the specification of correlations between regions, i.e. on the specification of the binary spatial weight matrix (see also Cliff and Ord, 1973; Cliff and Ord, 1981).

This point is illustrated in Figure 2.2. It shows the outcome of the *KR* statistic depending on the specification of threshold Euclidian distance between counties. If the distance between geographical midpoints of regions is larger than the

Figure 2.2: Dependence of the KR  $\chi^2$ -statistic on the threshold distance



threshold distance, then this corresponds to a zero in the spatial weight matrix. Hence, in this case these 2 counties are not considered in the computation of the  $KR$  statistic.

The lower axis in Figure 2.2 gives the threshold distance in kilometers, the upper axis gives the number correlations between counties which are taken into account in the computation of  $KR$ .

The  $KR$  statistic is distributed  $\chi^2$  and in our case with 15 degrees of freedom. Thus, the critical value at  $p=0.05$  is 24.99 and at  $p=0.01$  is 30.57. These two critical values are plotted as lines in Figure 2.2.

It emerges that for *all* and the *manufacturing* sectors the  $KR$  statistic reaches a maximum when the threshold distance is between 30 and 50 km. At this maximum, the total number of correlations between counties being considered in the calculation of  $KR$  ranges between 500 and 1500. This means, that the maximum of the  $KR$  statistic is reached when, for each county, between 2-5 correlations with the nearest neighbouring counties are taken into account.

From this explorative analysis, we conclude that spatial dependence is not significant in our case, since for no equation is the maximum of  $KR$  larger than the given critical value from the  $\chi^2$  distribution for  $p=0.01$ , which is 30.57. Hence, the estimation and inference based on the usual econometric techniques e.g. OLS or SUR remain valid.

## 2.4 Summary and conclusions

Overall, we find that long-run equilibrium output in a county will be higher, the better it is endowed with infrastructure. Thus, our essay substantiates the findings of other studies e.g. Aschauer (1989a), Biehl (1986), Munnell (1992) or Seitz (1993; 1994; 1995) that infrastructure contributes positively to private production. However, our approach is an extension of previous studies in several aspects. The main difference to most studies in this field is that we focus on effects of infrastructure at the local level. This appears to be reasonable since the main part of infrastructure is supplied at the local level, hence one can expect its benefits to be seen particularly at a local level.

Furthermore, in contrast to most previous studies we use indicators for infrastructure instead of public capital stocks measured in monetary terms. This allows us to overcome the unrealistic assumption that infrastructure stocks can be regarded as homogenous and thereby as comparable across regions.

Our main finding is that counties better endowed with both transport and human capital infrastructure have also a higher level of expected long-run total output. However, one surprising result of our study is that we do not find effects of infrastructure endowment on long-run output of the *manufacturing* sector at the local level. Hence, other factors than infrastructural endowment seem to determine the choice of location and production of manufacturing firms. On the other hand, we find that the contribution of infrastructure to local private production is most pronounced for the *service* and the *trade & transport* sectors. Moreover, we find that both transport and human capital infrastructure are important for total output, i.e. these two types of infrastructure appear to be complementary.

## Essay 3

# A Political Economy Model of Infrastructure Allocation: An Empirical Assessment

with

Olivier Cadot

Lars-Hendrik Röller

### *Abstract*

*This essay proposes a simultaneous-equation approach to the estimation of the contribution of transport infrastructure accumulation to regional growth. We model explicitly the political-economy process driving infrastructure investments; in doing so, we eliminate a potential source of bias in production-function estimates and generate testable hypotheses on the forces that shape infrastructure policy. Our empirical findings on a panel of France's regions over 1984-91 suggest that influence activities were, indeed, significant determinants of the cross-regional allocation of transportation infrastructure investments. Moreover, we find little evidence of concern for the maximisation of economic returns to infrastructure spending, even after controlling for pork-barrel and when imposing an exogenous preference for convergence in regional productivity levels.*

### 3.1 Introduction

If there is little doubt that investment in public infrastructure capital is a necessary condition for long-run growth in industrial countries and, *a fortiori*, in developing ones, *how much* infrastructure investment actually contributes to growth is still, in spite of a long-standing debate, a largely unsettled question. Disagreement over the magnitudes involved has persisted in spite of a massive amount of research sparked by the influential work of Aschauer (1989a, 1989b). Using aggregate data for the US between 1949 and 1985, Aschauer found that the elasticity of output to a broad measure of public infrastructure capital was significant and quantitatively very large; other studies using aggregate data (Ram and Ramsey (1989), 1988, and Munnell, 1990a) also found public capital to influence productivity significantly. At a time of widespread concern about the slowdown in US productivity growth, these findings suggested that a decline in the rate of public-capital accumulation was “a potential new culprit” (Munnell, 1990a, p. 3).

However, the early studies were fraught with logical and econometric difficulties, the most important of which are discussed in Gramlich’s 1994 review essay. Among the econometric problems, it was pointed out that the direction of causation was unclear (see Eisner, 1991; Tatom, 1993b; or Holtz-Eakin, 1994). For instance, Holtz-Eakin remarked that “it is tempting to infer a causal relationship from public-sector capital to productivity, but the evidence does not justify this step. It is just as easy to imagine the reverse scenario in which deteriorating economic conditions reduce capital-stock growth” (1994, p. 12). Disagreement over the meaning of elasticity estimates was not limited to time-series studies. Munnell (1990b) and Garcia-Milà and McGuire (1992) also found positive elasticities of output to public capital using panel data at the state level, but state-level evidence was vulnerable to similar criticism: quoting again Holtz-Eakin (1994, p. 13), “[b]ecause more prosperous states are likely to spend more on public capital, there will be a positive correlation between the state-specific effects and public sector capital. This should not be confused, however, with the notion that greater public capital leads a state to be more productive”. Holtz-Eakin’s own approach consisted of introducing fixed effects in the specification of the error

structure in order to control for unobserved state characteristics. But, as he himself remarked (p. 13), “in doing so the investigator ignores the information from cross-state variation in the variables”, which is of course unfortunate given that in a panel of short duration a substantial part of the information comes, precisely, from the data’s cross-sectional variation. Moreover, if the endogeneity of public infrastructure investments is a serious problem, the best way to address it is probably the most direct one, that is, to use simultaneous-equation methods (see Hulten, 1995 for a discussion; see also Tatom, 1993). A few authors followed this approach, e.g. Duffy-Deno and Eberts (1991) or de Frutos and Pereira (1993), and nevertheless found significant elasticities of output to infrastructure capital. But the key question, if one believes that the endogeneity issue matters, is how infrastructure investment decisions should be modelled.

Clearly, the “second equation” should be grounded in a theory of how public infrastructure investment decisions are made, i.e. on some explicit view about what drives policy-making. Indeed, Gramlich (1994) rightly points out that the primary interest of the infrastructure debate is not so much in the battle over elasticity estimates as in the implied policy debate. In his words, “rather than asking whether there is a shortage, it seems more helpful to ask what, if any, policies should be changed” (p. 1190). What Gramlich suggests is to focus explicitly on policy choices and the institutional context in which they are made in order to assess, from a normative perspective, whether they are socially optimal or not. This presupposes that institutions and policy choices are designed to maximise social welfare. But are they? A growing literature, at the frontier of economics and political science, views economic-policy decisions as resulting from the maximisation by incumbent politicians of objective functions that may depart considerably from social welfare, under constraints that are primarily political (see Dixit, 1996, for a survey). This literature approaches from a positive angle questions that used to be the realm of the normative, taking policy variables to be endogenous rather than control variables. Ultimately, normative considerations are likely to reappear, e.g. in the form of prescriptions in favour of rules or institutions mitigating policy capture by special interests; but the literature’s key message is that irrespective of what politicians ought to do, economists need to

understand what they actually do and why. For instance, if public infrastructure investment decisions are influenced by pork-barrel politics, pork-barrel politics should be part of the model. We believe that this provides a useful starting point for a discussion of what the “second equation” should be.

In this perspective, the present essay is a first attempt to formally bridge the gap between the infrastructure and political-economy literature. More specifically, we apply a common-agency model to the allocation of infrastructure investment in France. In principle, there exist other political economy mechanisms that may be able to explain how infrastructure investment is allocated in a particular country. However, the centralist institutional context in France lends itself towards an analysis under the common-agency setup. To put it differently, an empirical test of the common-agency model in political economy is ideally suited for the French environment.

The initial assumption is that firms have sunk investments giving them vested interests in the quality of the infrastructure in regions where they have production units (henceforth called “establishments”). We also posit that a firm with a large establishment in a given region should be expected to lobby harder than other firms for the maintenance and upgrading of that region’s infrastructure, for three reasons. First, large establishments produce, on average, for more distant markets (as higher volumes must be absorbed by wider geographical areas); as a result, they use highways and railways more intensively than others and are consequently more concerned about their maintenance and upgrading. Second, large establishments are typically owned by firms with headquarters in Paris; those firms are likely to be in a better position to effectively reinforce local lobbying by direct access to national policy-makers. Third, although we do not deal explicitly with collective-action problems in mobilising local political resources, such problems are likely to be easier overcome by a few firms with large stakes, such as Michelin in Auvergne or Citroën in Bretagne, than for a host of small or medium-sized local firms. For all these reasons, we assume that the number of large establishments in a region has a positive influence on the intensity of that region’s lobbying for transport infrastructure investment.

The form of the lobbying is very simple: we suppose that firms offer campaign

contributions to incumbent politicians in return for additional spending. At the margin, these contributions reflect the firms' willingness to pay for additional infrastructure—that is, they reflect the infrastructure's marginal contribution to firm value, both on the supply side, through the infrastructure's contribution to productivity in all sectors, and on the demand side for the construction industry itself.<sup>1</sup> The political process is as follows. Local politicians (we focus on regional presidents, whose power of influence has increased after administrative reforms enacted in 1982) act as contribution-collectors, providing their affiliated parties' headquarters with promises of locally-generated campaign contributions. Their own role is to assemble credible public-works projects for their region (as per their constituents' demands) and submit them to Paris for approval. Final decisions are made at the national level, either in the Ministry of Transport for relatively minor projects, or at Cabinet meetings for larger ones; *ceteris paribus*, local politicians who (i) share the national executive's political obedience and (ii) generate substantial campaign financing, are rewarded with a larger slice of the cake; this is what we call "pork-barrel politics". Thus, the process can be viewed as an auction whereby incumbent politicians sell infrastructure investments to local lobbies who bid for them through campaign contributions. Of course, in reality the mechanism through which lobbying pressures are transmitted from the local to the national level is neither as frictionless nor as transparent as it is portrayed here. But the central idea that local politicians are more effective voices for their constituents' demands if they happen to share the national executive's current political obedience is a plausible one and is supported by the data.

In order to capture this idea, we approximate lobbying by two variables: the number of large firms in each region and—perhaps more importantly—a dummy variable equal to one when a regional council's majority and the national government are of the same obedience (left wing or right wing) and zero otherwise. We find both measures of lobbying influence to be significant, statistically and in magnitude. Since lobbying takes place over investment levels (i.e. flows) and

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<sup>1</sup>Our panel covers a period (1985-91) immediately preceding a new law on political party financing. Ample anecdotal evidence suggests that, prior to that reform, a substantial part of the financing of mainstream parties came from contributions linked to public-works contracts.

output is determined by stocks, one may suspect that the simultaneity bias in the output equation, may not be very large. As it turns out, the simultaneity bias from estimating a production function alone is negligible, as single equation elasticity estimates are almost identical to those obtained by simultaneous estimation of both equations. It appears that stocks are too large relative to investments for feedback influences to be a real source of concern over a sample period of less than a decade. We also find that over our sample period (during which left-wing parties were in power for five years and right-wing ones for two), the French government did not seem to be significantly concerned by the maximisation of the *economic* returns to infrastructure spending. This result, which is robust to changes in the functional form of the government's objective function, reinforces our conclusion that pork-barrel considerations were important—if not primary—policy drivers in the sample period. Finally, we carry out an exercise in which we compare the actual allocation of infrastructure investments across regions with a simulated socially optimal value. Interestingly, we find that most of the cross-regional variation in investment levels is attributable to pork barrel, suggesting that uniform allocation would be a good rule of thumb to reduce policy capture by lobbies.

The remainder of this essay is organised as follows. In section 2, we state general conditions for the efficient provision of a public input and derive conditions under which influence activities lead to inefficient provision in a political-economy model. In section 3, we report the results of empirical testing of the model's structural equations on a French data set. Section 4 concludes.

## 3.2 Theory

Basic results on the optimal provision of public inputs were derived by Kaizuka (1965), Sandmo (1972), and Negishi (1973). We briefly review these results in the following section in order to provide a benchmark against which inefficiencies arising from influence activities can be assessed.

### 3.2.1 Efficient provision of a public input

Consider an economy producing  $m$  final goods for consumption, with technologies  $Q_i = F^i(K_i, L_i, X)$ , for  $i = 1, \dots, m$ , where  $K_i$  and  $L_i$  are respectively the amounts of capital and labour used up in the production of good  $i$  and  $X$  is a pure public input. Following Negishi (1973), we take the latter to be of the “unpaid input” type, meaning that the function  $F^i$  is linearly homogenous in  $K_i$ ,  $L_i$  and  $X$ .<sup>2</sup> When such is the case, owners of capital, which are residual claimants, appropriate the rents generated by the public input if the latter is not priced at the value of its marginal product. This is a source of potential inefficiency in capital-allocation decisions; but for simplicity (and for reasons that will become clear in the empirical part of the essay) we will limit the analysis to a short-run case where capital is fixed; the theory can be easily extended to a long-run case.

Let the public input be produced with labour only according to a technology  $G(L_X)$ , and let  $L = \sum_{i=1}^m L_i + L_X$  be the economy’s total endowment of labour. For reasons that will become clear later on, we will assume that the social utility function is quasilinear (necessary conditions for the aggregation of individual preferences into a social utility function are assumed to hold); thus,  $U(Q_1, \dots, Q_m) = Q_1 + \sum_{i=2}^m u(Q_i)$  where the function  $u$  is increasing and concave. Given this, the problem of a social planner is:<sup>3</sup>

$$\begin{aligned} & \max_{L_1, \dots, L_m, L_X} Q_1 + \sum_{i=2}^m u(Q_i) \\ & \text{s.t.} \\ & Q_i = F^i(K_i, L_i, X), \quad i = 1, \dots, m, \\ & X = G(L_X), \\ & L = \sum_{i=1}^m L_i + L_X. \end{aligned} \tag{3.1}$$

<sup>2</sup>The alternative formulation is to assume that the production function is linearly homogeneous in  $K_i$  and  $L_i$  alone and has increasing returns in all factors including  $X$ . This alternative formulation is generally seen as appropriate for publicly-provided R&D, whereas the classical example of the former formulation is, according to Sandmo (1972) and Negishi (1973), transport infrastructure.

<sup>3</sup>Transport infrastructure is used as an input not only by firms, but also by households; so a complete statement of the problem should include a household production function. We will abstract from such considerations and treat transport infrastructure as a “pure input”.

Letting subscripts denote partial derivatives (so  $F_L^i = \partial F^i / \partial L_i$  and  $F_X^i = \partial F^i / \partial X$ ), solving (3.1) and rearranging the resulting first-order conditions gives the basic condition for the efficient provision of  $X$ ; namely,

$$\sum_{i=1}^m \frac{F_X^i}{F_L^i} = \frac{1}{G'} \quad (3.2)$$

Condition (3.2), which closely parallels Samuelson's condition for the optimal provision of public goods, was initially derived by Kaizuka (1965). It states that the sum over industries (firm-level production functions can be aggregated within each industry because the production function is homogenous) of the rates of technical substitution between labour and the public input must be equated to the marginal cost of the public input's provision.

Whereas the maximisation of social utility under technology and factor endowment constraints is the most natural way of deriving (3.2), this efficiency condition can also be derived from the maximisation of firm profits. Let good 1 be the numeraire,  $p_i$  the price of good  $i$  in terms of good 1, and  $w$  the wage rate, and fix all prices and the wage rate at the levels obtained implicitly from the solution of problem (3.1). Suppose that, at these exogenously given prices and wage, firms make profit-maximising decisions contingent on  $X$ ; let also  $H$  be the inverse function of  $G$  so that  $L_X = H(X)$ . A government maximising firm profits by choice of  $X$  will solve:

$$\begin{aligned} \max_X \quad & F^1(K_1, L_1, X) + \sum_{i=2}^m p_i F^i(K_i, L_i, X) - w \sum_{i=1}^m L_i - wH(X) \\ \text{s.t} \quad & \\ & F_L^1 = p_2 F_L^2 = \dots = p_i F_L^i = w. \end{aligned} \quad (3.3)$$

It is easily checked that the solution of (3.3) satisfies first-order condition (3.2) and consequently yields the same level of provision of  $X$ . Although straightforward, this result is very important for our purposes. To see why, consider a simple influence-activity game in which firms offer monetary contributions to an incumbent politician in exchange for the public input's provision, and suppose that the incumbent maximises the sum of those contributions net of the input's cost. If, at the margin, contributions reflect the willingness of firms to pay for the input, the influence-activity game's unique equilibrium is the solution to (3.3). In other

words, if  $C^i(X)$  is industry  $i$ 's offer of a monetary contribution to the government and  $\pi^i = p_i F^i(K_i, L_i, X) - wL_i$  (with  $p_i = 1$  when  $i = 1$ ) is its profits, whenever  $\partial C^i / \partial X = \partial \pi^i / \partial X$ , a government maximising  $\sum_i C^i(X) - wH(X)$  will maximise (3.3) and consequently provide  $X$  according to (3.2). Thus, influence activities *by themselves* do not imply inefficient provision of the public input.

This result—namely, that the existence of influence activities is not a sufficient condition for an inefficient policy outcome—is simply a restatement of Bernheim and Whinston's (1986a) result according to which, if influence activities can be represented as a "menu auction" and if special-interest groups bid for policy according to their marginal valuation, the resulting "truthful" equilibrium is Pareto-efficient (see also Bernheim and Whinston, (1986a), for parallel efficiency results in a common-agency context). This result also appears in a trade-policy context in Grossman and Helpman (1994) who show that in a small open economy, if all agents are represented in one lobby or another, the resulting equilibrium is free trade. We now turn to conditions under which influence activities do lead to inefficient policy choices.

### 3.2.2 Influence activities and inefficient policies

We have established that, in the case of a pure public input, a “policy auctioneer” implements the same policies that a social planner would, provided that all firms have access to the bidding process and bid according to their marginal willingness to pay. It follows that inefficiencies can come from only two sources. First, some firms may not have access to the bidding process, or may choose to free-ride. For instance, small firms may keep out of lobbying because it entails an entry fee that is prohibitive for them. This kind of incomplete coverage of the bidding process may lead to under-provision of the public input, quite like a standard collective-action problem. If infrastructures are specific to geographical entities, like regions or states, and the number of large firms varies across these entities, distortions will also appear in the spatial allocation of the public input.<sup>4</sup> Second, incumbent politicians may pursue an agenda of their own; that is, instead of simply maximising the sum of the lobbies’ transfers as a pure auctioneer would, they may maximise a composite function in which lobbying and non-lobbying arguments enter as substitutes. Non-lobbying arguments—such as priority development of some types of regional infrastructures—may entail choices which, although desirable from the incumbent’s perspective, deviate from the first-best allocation of the public input. But they may also reflect economic-efficiency considerations, as opposed to pork barrel (this is the case considered by Grossman and Helpman, 1994, in which the government maximises a linear combination of social welfare and contributions from lobbies).

For instance, suppose that the incumbent government maximises a convex combination of social utility  $U(\cdot)$  and a monetary contribution  $C^k(X)$  from some non-numeraire industry  $k$ . Again, the economy is in a competitive equilibrium as far as consumption and the allocation of labour across industries are concerned, the government’s only problem being the provision of public input  $X$ .

Suppose that the government now maximises a linear combination of industry  $k$ ’s contribution and social utility, the resource constraint being represented as

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<sup>4</sup>Note that the existence of lobbying implies that collective-action problems are at least partially overcome. If collective-action problems were so severe as to hamper *any* lobbying, there would be no distortion in the state’s provision of the public input.

in problem (3.3) by the term  $wH(X)$ . Although we are aggregating money and “utils”, this poses no particular problem as long as preferences are quasilinear. Letting  $a$  be the weight on social utility the government now solves

$$\begin{aligned}
 & \max_X (1-a)C^k(X) + a [Q_1 + \sum_{i=2}^m u(Q_i)] - wH(X) \\
 & \text{s.t.} \\
 & Q_i = F^i(K_i, L_i, X), \quad i = 1, \dots, m, \\
 & F_L^1 = p_2 F_L^2 = \dots = p_m F_L^m = w, \\
 & u' = p_i \quad \forall i = 2, \dots, m, \\
 & C^{k'}(X) = \partial \pi^k / \partial X = p_k F_X^k,
 \end{aligned} \tag{3.4}$$

with first-order condition

$$(1-a)p_k F_X^k + a \left( F_X^1 + u' \sum_{i=2}^m F_X^i \right) - wH' = 0,$$

after substitution of the relevant constraints, this becomes

$$\frac{F_X^k}{F_L^k} + a \sum_{i \neq k} \frac{F_X^i}{F_L^i} = H'. \tag{3.5}$$

Thus, efficiency condition (3.2) is now violated; as the left-hand side of (3.5) is a decreasing function of  $X$  whereas its right-hand side is an increasing one, the public input is underprovided in (3.5) compared to (3.2). However, underprovision follows from the choice of a convex combination of social utility and industry  $k$ 's contribution in the objective function; non-convex linear combinations could yield overprovision. In the empirical part of this essay, we will not impose convexity.

Given that (3.5)'s departure from optimality comes from the fact that sector  $k$  and only sector  $k$  lobbies, it can be eliminated in two ways. First, the distortion shrinks as  $a$  increases; in the limit, when  $a = 1$ , (3.5) reduces to (3.2). That is, the departure from optimality disappears if the government's valuation of sector  $k$ 's contribution goes to zero. Second, if all industries lobby, (3.5) reduces to (3.2) irrespective of the value of  $a$  in  $[0, 1]$ , because by maximising a convex combination of social utility and the profits of competitive firms, the government in effect maximises twice the same thing.

Although simple, this theoretical framework provides a useful starting point for our empirical exploration of the effect of lobbying on the allocation of transport infrastructure investments. Whether or not there is underprovision of the public input as implied by (3.5) is a very important question because underinvestment in infrastructure is a subject of recurrent concern, in particular in the US.<sup>5</sup> A second important implication of (3.5) is that if the intensity or effectiveness of influence activities varies across states or regions, distortions in the overall level of infrastructure investments will be compounded by distortions in their spatial allocation. We now turn to an estimable model of regional infrastructure allocation building on these foundations.

### 3.2.3 A model of regional infrastructure allocation

#### Production function

Let  $Q_{it}$  be the aggregate output of region  $i$  at time  $t$ ,  $L_{it}$  the level of regional employment,  $K_{it}$  the region's aggregate (non-infrastructure) capital stock, and  $X_{it}$  its stock of transport infrastructure.<sup>6</sup> All regions have identical aggregate Cobb-Douglas production functions  $F$ :

$$Q_{it} = F(A_{it}, L_{it}, K_{it}, X_{it}) = A_t L_{it}^{\alpha_L} K_{it}^{\alpha_K} X_{it}^{\alpha_X}, \quad (3.6)$$

where  $A_t$  is a technical-change parameter common to all regions. Note that this formulation rules out cross-regional externalities in the productivity of transport infrastructure; while this assumption is obviously an oversimplification, Gramlich (1990) and Holtz-Eakin (1994) argued on the basis of US data that such externalities are unlikely to be a major problem, as most traffic, even on interstate highways, is local. Moreover, relaxing it would require the estimation of a large

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<sup>5</sup>It should be noted, however, that the political model implicit in (3.4) is a representative-democracy one, whereas in the US about 20% of new state and local construction must be approved by referendum (see Gramlich, 1994). Peterson (1991) showed that under this partial direct-democracy mechanism the political economy of infrastructure construction is also likely to lead to underprovision of the infrastructure, as risk-averse politicians undertake projects only when assured of an overwhelming majority (the average approval percentage in referenda was close to 70% over 1948-90).

<sup>6</sup>We abstract from non-transport infrastructure like schools, hospitals, and so on.

number of parameters relative to our sample size. Dividing through by  $L_{it}$ , (3.6) becomes

$$q_{it} = A_t k_{it}^{\alpha_K} x_{it}^{\alpha_X} L_{it}^{\tilde{\alpha}_L}, \quad (3.7)$$

where  $q_{it}$  is labour productivity,  $k_{it}$  is the capital-labour ratio and  $x_{it}$  is the stock of transport infrastructure per worker (we will henceforth use lower-case letters to designate per-worker variables),<sup>7</sup> and  $\tilde{\alpha}_L = \alpha_L + \alpha_K + \alpha_X - 1$ . Note that  $\tilde{\alpha}_L = 0$  if returns to scale are constant.

As policy decisions are concerned with infrastructure investments rather than stocks, for future purposes we need to establish the formal link between the two. The law of motion of region  $i$ 's real transport infrastructure stock  $X_{it}$  is given as

$$X_{it} = \gamma X_{i,t-1} + Z_{it}, \quad (3.8)$$

where  $Z_{it}$  denotes real gross investment in transport infrastructure and  $1 - \gamma$  is the rate of depreciation of the infrastructure stock, so that

$$\frac{\partial X_{it}}{\partial Z_{it}} = 1. \quad (3.9)$$

### Policy function

We model lobbying as a common agency game. Although the problem should formally be treated in an explicit intertemporal context (see Bergemann and Valimaki, 1998), for the sake of simplicity we will reduce it to a succession of static games. In each period, region-specific lobbies indexed by  $i = 1, \dots, n$  (the principals) simultaneously face the government with monetary transfer offers  $C_{it}(\mathbf{Z}_t)$  conditioned on the vector of transport infrastructure investments  $\mathbf{Z}_t = (Z_{1t}, \dots, Z_{nt})$ . These transfers can be interpreted, depending on the context, as political campaign contributions or outright bribes. The government then chooses a value  $\mathbf{Z}_t^*$  of the policy vector  $\mathbf{Z}_t$  that maximises a suitably defined objective function  $\mathcal{V}[\mathbf{Z}_t, \sum_i C_{it}(\mathbf{Z}_t)]$ . Finally, lobbies make transfers  $C_i(\mathbf{Z}_t^*)$  as promised. In order to be consistent with the framework of the previous section, keeping the same notation let  $\mathcal{V}[\mathbf{Z}_t, \sum_i C_{it}(\mathbf{Z}_t)] = au(\mathbf{Z}_t) + (1 - a) \sum_i C_{it}(\mathbf{Z}_t) - H(\mathbf{Z}_t)$ ; the nature of the functions  $u$  and  $H$  will be explained later on. Let also  $\Pi_{it}(\mathbf{Z}_t)$  be the

<sup>7</sup>Using per-worker variables reduces heteroscedasticity due to unequal region sizes.

value of  $\mathbf{Z}_t$  to lobby  $i$ . The game's unique "truthful" equilibrium is characterised by the following equations:

$$\left. \frac{\partial C_{jt}(\mathbf{Z}_t)}{\partial Z_{it}} \right|_{\mathbf{Z}_t^*} - \left. \frac{\partial \Pi_{jt}(\mathbf{Z}_t)}{\partial Z_{it}} \right|_{\mathbf{Z}_t^*} = 0, \quad i, j = 1, \dots, n; \quad (3.10)$$

$$a \frac{\partial u}{\partial Z_{it}} + (1 - a) \sum_{j=1}^n \frac{\partial C_{jt}}{\partial Z_{it}} - \frac{\partial H}{\partial Z_{it}} = 0, \quad i = 1, \dots, n. \quad (3.11)$$

Equations (3.10) are "truthfulness" conditions whereas (3.11) is the government's first-order condition. Bernheim and Whinston (1986b, Theorem 2) state a number of sufficient conditions under which the common-agency game's unique equilibrium maximises the joint surplus of the agent and principals, i.e. under which it collapses to a single principal-agent problem which, in the absence of hidden action, generates no inefficiency. These conditions do not apply here, because small firms do not lobby, whence transfer functions are distorted. Thus, efficiency does not hold.

The first step in taking (3.10) and (3.11) to the data consists of parameterizing the  $u$  and  $H$  functions. We define  $u$  as a nested function of  $\mathbf{Z}_t$ ; i.e.  $u$  is a function of productivities  $q_{it}$ , themselves functions of  $\mathbf{Z}_t$  through (3.7). Specifically, let  $u(\mathbf{Z}_t) = 2 \sum_i L_{it} q_{it}(\mathbf{Z}_t)^{1/2}$ . This formulation reflects the twin assumptions that the government values convergence in per-capita incomes (hence the concave form),<sup>8</sup> and that a given departure from optimal productivity levels receives more weight, *ceteris paribus*, if it affects a more populous region (hence the multiplicative term  $L_{it}$ ).

The costs of these investments are captured by the function  $H$ . Some of the spending is financed by corporate taxes whose impact is reflected in the firms' willingness to lobby (see details below). The rest is financed by other taxes which, although they do not directly affect the profits of local firms, affect local welfare and are therefore of concern to the government. Accordingly, let  $H(\mathbf{Z}_t) = \sum_i L_{it} z_{it}^2 / 2$ . The quadratic form reflects a rising marginal distortion

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<sup>8</sup>Other functional forms, e.g. logarithmic, were tried in the estimation and found to yield similar estimates.

cost of individual tax burdens;<sup>9</sup> in general, convexity (quadratic or other) of infrastructure investments costs can reflect non-financial considerations as well as financial ones. For instance, in a pork-barrel context, it may be politically important for the incumbent government to appear even-handed in the distribution of favours. Using these functional forms and differentiating the non-lobbying terms of (3.11) with respect to  $Z_{it}$  gives

$$a \frac{\partial u}{\partial Z_{it}} - \frac{\partial H}{\partial Z_{it}} = a L_{it} q_{it}^{-1/2} \frac{1}{L_{it}} \frac{\partial Q_{it}}{\partial X_{it}} \frac{\partial X_{it}}{\partial Z_{it}} - z_{it},$$

where  $z_{it} = Z_{it}/L_{it}$ . Using (3.6) and (3.9), this simplifies to

$$a \frac{\partial u}{\partial Z_{it}} - \frac{\partial H}{\partial Z_{it}} = a F_X q_{it}^{-1/2} - z_{it}, \quad (3.12)$$

where  $F_X = \partial F / \partial X_{it}$ .

We now turn to the lobbying term. In a transport-infrastructure allocation problem, it is natural to suppose that lobbying is organised along regional lines, with industrial firms playing an important role in the process. As already discussed, we will treat transport infrastructure as a pure input, so that only firms lobby for it, and we will assume, in addition, that firms do not lobby in regions where they have only small production units (this is the “small-firms-out” assumption). In the absence of cross-regional externalities (discussed in section 2.3.1), firms having establishments in multiple regions make separate lobbying decisions for each one of their establishments, so we can treat the latter, without loss of generality, as separate firms. Accordingly, suppose that in any region  $i$ ,  $N_{it}$  identical, large “firms” are active in lobbying. Although large, these firms are price-takers, and we will assume that they all produce a single manufactured good priced at  $p_{it}$ ; as all variables are measured in constant 1992 francs in the empirical part, we set  $p_{it} = 1$  for all  $i$  and  $t$ . As transport infrastructure investments, in particular on highway maintenance and construction, are typically financed out of composite packages combining local and national budgets, we model their impact on local taxes through a tax function  $\mathcal{T}_{it}^\ell(Z_{it})$ . On the other hand, we assume that the *use* of transport infrastructure is free. In order to include regional

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<sup>9</sup>The tax burden can alternatively be expressed as a percentage of regional GDP; however, such a formulation turns out to yield awkward functional forms with difficult-to-interpret parameters.

employment and private capital stocks as right-hand side variables in the production function, we assume that they are taken by the representative firm as fixed; finally, we denote by  $F^\ell$  the production function of a representative large firm. The profit of a representative large firm in region  $i$  at time  $t$  is then

$$\pi_{it}^\ell = p_{it}F^\ell(A_t, K_{it}^\ell, L_{it}^\ell, X_{it}) - w_{it}L_{it}^\ell - r_{it}K_{it}^\ell - \mathcal{T}_{it}^\ell(Z_{it}), \quad (3.13)$$

where  $w_{it}$  and  $r_{it}$  are the wage rate and the rental rate of capital in region  $i$  at time  $t$ . Suppose that firms pay local taxes in proportion to their employment in the region;<sup>10</sup> then, letting  $L_{it}^\ell$  be the total number of employees in large establishments in region  $i$  at time  $t$ , the tax function facing a representative large firm in region  $i$  is

$$\mathcal{T}_{it}^\ell(Z_{it}) = \lambda \frac{Z_{it}}{N_{it}} \frac{L_{it}^\ell}{L_{it}}, \quad (3.14)$$

for some (unknown) parameter  $\lambda$ . Substituting (3.14) into (3.13), aggregating over  $N_{it}$  identical large firms (i.e. multiplying by  $N_{it}$ ) and differentiating with respect to  $Z_{it}$  gives

$$\frac{\partial \Pi_{it}}{\partial Z_{it}} = N_{it} \frac{\partial \pi_{it}^\ell}{\partial Z_{it}} = N_{it} \left[ F_X^\ell - \mathcal{T}_{it}^{\ell'}(Z_{it}) \right] = N_{it} \left( F_X^\ell - \lambda \frac{L_{it}^\ell}{N_{it}} \right), \quad (3.15)$$

where  $F_X^\ell \equiv \partial F^\ell / \partial X_{it}$  and  $l_{it}^\ell = L_{it}^\ell / L_{it}$ . We will henceforth assume that (3.15) is positive; that is, that the marginal local-tax burden does not swamp the marginal benefit of infrastructure investments (since otherwise there would be no lobbying). Finally, using (3.10) and making use of the no-externality assumption,

$$\frac{\partial C_{jt}}{\partial Z_{it}} = \frac{\partial \Pi_{jt}}{\partial Z_{it}} = \begin{cases} F_X^\ell N_{it} - \lambda l_{it}^\ell & \text{if } j = i \\ 0 & \text{otherwise.} \end{cases} \quad (3.16)$$

The government's first-order condition (3.11) is found by adding (3.16) to (3.12) and setting their sum equal to zero. Finally, solving for  $z_{it}$  yields

$$z_{it} = a F_X q_{it}^{-1/2} + (1 - a)(F_X^\ell N_{it} - \lambda l_{it}^\ell). \quad (3.17)$$

<sup>10</sup>The largest local tax in France is the *taxe professionnelle* which is proportional to employment. As a robustness check, we also tried empirically an alternative formulation whereby the tax burden on local companies was proportional to their sales; it gave similar results.

Together, (3.7) and (3.17) form a system of two equations which we will estimate simultaneously, yielding consistent estimates of the contribution of transport infrastructure investments to GDP and of the extent of political interference with these investment decisions.

### 3.3 Empirical Implementation

#### 3.3.1 Data and Summary Statistics

We use a panel data set covering 21 of France's 22 regions (we excluded Corsica because of its poor data) from 1985-91. Table 3.1 provides a brief description of the variables and a list of the relevant regions. All figures are in 1992 Francs. Output  $Q$  is measured as value added at factor cost and has been obtained from the Eurostat database 'New Cronos' (June 1999). Regional employment  $L$  is also taken from 'New Cronos' and covers all private sectors of the economy. The private capital stock  $K$  is constructed by the *Laboratoire d'Observation Economique et des Institutions Locales* (OEIL) using national data from INSEE's *Compte de Patrimoine* and allocating the national stock to the regions on the basis of corporate tax rates.

The transport infrastructure stock  $X$  is constructed as follows. As stock data was not available at the regional level, we construct the stock from investment data using the perpetual inventory method (PIM). In order to obtain a benchmark stock level for the initial period, we allocate the national stock, for which data is given by the *Fédération Nationale des Travaux Publics* (FNTP, see also Laguarrigue, 1994) across the 21 regions in proportion to their average investment share over the first three years of the sample period. The relatively slow rate of depreciation of infrastructure capital implies that our stock converges slowly to the true one. In order to reduce possible biases in the calculation of the infrastructure stock we use infrastructure investment data going back to 1975. Aggregating our regional stock data to the national level and comparing it with national data obtained from INSEE yields only marginal differences.

The transport infrastructure investment data come from several sources. Railway figures were provided directly by SNCF, the national railway company.

*Table 3.1: Variable Description and Regions*

Variable	Description
$Q$	Regional GDP, million 1992 Francs
$L$	Regional employment, million individuals
$K$	Non-residential private capital stock, million 1992 Francs
$X$	Transport infrastructure stock, million 1992 Francs
$Z$	Transport infrastructure net investments, million 1992 Francs
$N$	Number of establishments with more than 500 employees
$D_{it}$	Dummy variable equal to 1 when regional council and national parliament have same political majority.
<b>Regions</b>	
Alsace	Champagne-Ardennes    Midi-Pyrenées
Aquitaine	Franche-Comté    Nord-Pas de Calais
Auvergne	Haute-Normandie    Pays de Loire
Basse-Norm.	Ile-de-France    Picardie
Bourgogne	Languedoc-Roussillon    Poitou-Charentes
Bretagne	Limousin    Provence-Alpes-Côte d'Azur
Centre	Lorraine    Rhône-Alpes

Highway figures, which are reported for the year in which the work is done (rather than for the year of budget allocation—there is a delay between the two) have been collected by the OEIL from data generated by the FNTP (see Fritsch and Prud'homme, 1994, for details). The FNTP's data are based on reports by the Federation's member companies. Finally, investment data for waterways was taken directly from the FNTP's statistical yearbook. Although airport construction data, which we had collected from the *Direction Générale de l'Aviation Civile* (DGAC), would have been a natural inclusion in the study, we found that they were not sufficiently reliable and consequently eliminated them from this study.

The number of industrial establishments with more than 500 employees ( $N_{it}$ ) is taken from various issues of *L'Industrie dans les Régions*, a yearly statistical publication of the Ministry of Industry. From our model it is natural to suppose that the region Paris plays a specific role in lobbying process for infrastructure investment, therefore we defined two new variables as  $N_{it}(1 - D_{Paris})$  and  $N_{it}D_{Paris}$ , i.e. for the former the observations from Paris are excluded whereas for latter all other observations except from Paris are excluded. Finally, the partisan dummy variable ( $D_{it}$ ) is equal to 1 when the majority in a Regional Council (and hence the affiliation of the region's President) and that of the national parliament (and hence of the current government) are either both right-wing or both left-wing, and zero otherwise.<sup>11</sup> As our sample includes one regional election (in 1986) and two national legislative elections (in 1986 and 1988 respectively),  $D_{it}$ , which was constructed using press sources, varies both across regions and across time. We lagged it by one year to take account of budget delays.

Table 3.2 shows descriptive statistics for these variables. In 1992 Francs, over the sample period, average infrastructure investment amounted to 1396 Francs per worker, or roughly 0.54 percent of GDP; the infrastructure *stock* amounted to 50,920 Francs per worker, or 19.8 percent of GDP. The value of the highway infrastructure stock was about 5 times that of the railway stock and 70 times that of the waterways infrastructure stock.

<sup>11</sup>“Right wing” was defined in the sample as RPR, UDF, Front National, and “Divers Droite”. “Left wing” was defined as Parti Socialiste, Parti Communiste, Mouvement des Radicaux de Gauche, various environmentalist parties, and “Divers Gauche”. The “Divers Gauche” and “Divers Droite” categories classify independent individuals according to their voting patterns.

**Table 3.2: Summary Statistics**

Variable	Mean	Std.Dev.	C.V.	Minimum	Maximum
$Q/L$	256723	27380.40	0.106	195921	357617
$K/L$	298142	62624.98	0.210	188442	484980
$X/L$	50920	9412.91	0.185	35453	70935
$Q/X$	5.166	0.8416	0.163	3.849	7.405
$Z/L$	1396	628.15	0.450	412	4934
$N * (1 - D_{\text{Paris}})$	35.27	21.66	0.614	5	113
$N * D_{\text{Paris}}$	7.19	32.42	4.505	0	170
$D_{it}$	0.435	0.50	1.143	0	1
$X_{\text{HIGHWAY}}/X_{\text{RAIL}}$	5.18	2.36	0.456	1.74	13.68
$X_{\text{HIGHWAY}}/X_{\text{WATER}}$	68.8	80.17	1.165	2.27	290

Total number of observations: 147

### 3.3.2 Baseline estimates

Several further adjustments are needed before (3.6) and (3.17) can be taken to the data. First, we drop the assumption that the weights on lobbying and non-lobbying terms in the government's objective function ( $1 - a$  and  $a$  respectively) add up to one. As these weights are arbitrary, we will simply call them  $a_1$  and  $a_2$ . Second, we approximate the marginal product of infrastructure capital for the representative large firm in one region by its aggregate value in that region; using the fact that, under technology (3.6),  $F_X^\ell \approx \alpha_X Q_{it}/X_{it} = \alpha_X q_{it}/x_{it}$  and simplifying, (3.17) becomes

$$\begin{aligned}
 z_{it} &= a_1 \alpha_X (q_{it}/x_{it}) q_{it}^{-1/2} + a_2 \left[ \alpha_X (q_{it}/x_{it}) N_{it} - \lambda l_{it}^\ell \right] \\
 &= \theta_{\text{PROD}} q_{it}^{1/2}/x_{it} + \theta_{\text{LOBBY}} q_{it} N_{it}/x_{it} + \theta_{\text{TAX}} l_{it}^\ell,
 \end{aligned} \tag{3.18}$$

where  $\theta_{\text{PROD}} = \alpha_X a_1$ ,  $\theta_{\text{LOBBY}} = \alpha_X a_2$ , and  $\theta_{\text{TAX}} = \lambda a_2$ . Third, as (3.18) is nonlinear in  $x_{it}$ , using (3.8) to substitute for  $x_{it}$  does not yield a closed form for  $z_{it}$ . Therefore we take care of the endogeneity of  $x_{it}$  by instrumenting it with its lagged

value. Fourth, we include time dummies<sup>12</sup> and regional dummies (for Ile-de-France in the production function, and for Ile de France, Nord-Pas-de-Calais in 1991, and Centre between 1986 and 1990 in the policy function; the first because it contains Paris and the last two because of large-scale Eurotunnel and TGV construction). Fifth, in the policy function we include as a separate regressor the “partisan” dummy  $D_{i,t-1}$ . Finally, we assume an AR(1) structure for the error term of both equations. Denoting fixed time-effects by  $\alpha_t$  and  $\beta_t$ ,  $t = 1 \dots T$ , the system to be estimated is thus:

$$\ln q_{it} = \alpha_t + \alpha_K \ln k_{it} + \alpha_X \ln(x_{i,t-1} + z_{it}) + \tilde{\alpha}_L \ln(L_{it}) + \alpha_{PARIS} D_{Paris} + v_{1it}, \quad (3.19)$$

$$\begin{aligned} z_{it} = & \beta_t + \theta_{PROD} q_{it}^{1/2} / x_{i,t-1} + \theta_{LOBBY} q_{it} N_{it} / x_{i,t-1} (1 - D_{Paris}) \\ & + \theta_{LOBBY\_PARIS} q_{it} N_{it} / x_{i,t-1} D_{Paris} + \theta_{TAX} l_{it}^{\ell} + \theta_{PARTY} D_{i,t-1} \\ & + \beta_{NORD} D_{Nord} + \beta_{PARIS} D_{Paris} + \beta_{CENTRE} D_{Centre} + v_{2it}, \end{aligned} \quad (3.20)$$

where  $v_{kit} = \rho_k v_{ki,t-1} + \varepsilon_{kit}$ ,  $k = 1, 2$ , and  $\varepsilon_{kit}$  are i.i.d normal variables with mean zero and variance  $\sigma_k$ . The estimation procedure is as follows. We estimate (3.19) and (3.20) simultaneously by non-linear Full-Information Maximum Likelihood (FIML),<sup>13</sup> using a Prais-Winsten transformation which avoids omitting observations for  $t = 1$ , (Greene, 1997, p. 601). For the non-linear OLS estimation we obtain the autocorrelation parameters  $\rho_k$ ,  $k = 1, 2$ , by minimizing the Sums of Squares Errors (SSE) for each equation, whereas for the non-linear FIML the autocorrelation parameters  $\rho_k$  are jointly estimated with the other parameters. The results are reported in Table 3.3.

Several specification tests are performed. In order to test the AR(1) specification against the alternative of an AR(2) specification, we employ the Godfrey

<sup>12</sup>Instead of fixed time effects we could also use linear time trends both for the policy equation and the production function, as supposed in (3.6). While estimating the model with linear time trends does not change the main results, using time dummies stresses also the cross-sectional variation between regions.

<sup>13</sup>Estimations have been carried out using PROC MODEL, SAS 6.12.

**Table 3.3: Estimation Results**

	(a)	(b)	(c)	(d)
	OLS	FIML	FIML	FIML
<b>Production Function</b> Dependent Variable: $\ln q_{it}$				
$\alpha_{85}$	11.24 (23.50)	11.18 (20.91)	11.19 (20.94)	11.18 (20.89)
$\alpha_{86}$	11.27 (23.52)	11.21 (20.93)	11.22 (20.96)	11.21 (20.92)
$\alpha_{87}$	11.28 (23.53)	11.22 (20.94)	11.23 (20.96)	11.23 (20.92)
$\alpha_{88}$	11.30 (23.54)	11.24 (20.95)	11.25 (20.98)	11.24 (20.94)
$\alpha_{89}$	11.32 (23.58)	11.26 (20.99)	11.27 (21.01)	11.26 (20.97)
$\alpha_{90}$	11.31 (23.55)	11.25 (20.95)	11.26 (20.98)	11.25 (20.94)
$\alpha_{91}$	11.31 (23.52)	11.25 (20.93)	11.26 (20.95)	11.25 (20.91)
$\alpha_K$	0.189 (5.44)	0.181 (4.77)	0.182 (4.80)	0.182 (4.82)
$\tilde{\alpha}_L$	0.025 (1.60)	0.025 (1.43)	0.025 (1.44)	0.025 (1.44)
$\alpha_X$	0.097 (2.30)	0.101 (2.14)	0.100 (2.12)	0.101 (2.15)
$\alpha_{\text{Paris}}$	0.218 (5.20)	0.219 (4.27)	0.219 (4.57)	0.219 (4.58)
AR(1) $\rho_1$	0.867	0.904	0.904	0.904
Godfrey LM	1.764	3.366	3.293	3.278
Shapiro-Wilk $W$	0.979	0.985	0.985	0.985
$R^2$	0.9539	0.9537	0.9537	0.9537
<b>Policy Function</b> Dependent Variable: $z_{it}$				
$\beta_{85}$	848.7 (2.59)	879.8 (2.52)	997.2 (2.85)	1043.6 (2.57)
$\beta_{86}$	1159.8 (2.92)	1186.8 (3.40)	1304.4 (3.71)	1362.6 (3.36)
$\beta_{87}$	1199.4 (3.02)	1226.1 (3.58)	1340.8 (3.90)	1392.8 (3.50)
$\beta_{88}$	1539.4 (3.03)	1564.0 (4.56)	1730.1 (5.09)	1734.0 (4.37)
$\beta_{89}$	1433.3 (3.64)	1457.2 (4.28)	1622.6 (4.82)	1632.5 (4.15)
$\beta_{90}$	1949.9 (4.34)	1975.9 (6.03)	2039.6 (6.13)	2145.9 (5.61)
$\beta_{91}$	1767.7 (4.36)	1794.6 (5.67)	1853.3 (5.77)	1955.9 (5.28)
$\theta_{\text{Prod}}$	-28165 (-0.79)	-31545 (-1.08)	-33253 (-1.12)	-31319 (-0.89)
$\theta_{\text{Lobby}}$	1.334 (2.07)	1.456 (2.79)	1.388 (2.61)	— (—)
$\theta_{\text{Lobby\_Paris}}$	-1.775 (-0.89)	-1.762 (-0.96)	-1.914 (-1.03)	— (—)
$\theta_{\text{Party}}$	120.5 (1.98)	123.8 (2.11)	— (—)	117.9 (2.01)
$\theta_{\text{Tax}}$	-1126.8 (-0.44)	-1473.5 (-0.71)	-1906.4 (-0.91)	376.1 (0.16)
$\beta_{\text{Paris}}$	1971.0 (0.99)	1993.9 (1.07)	2100.6 (1.11)	-30.7 (-0.11)
$\beta_{\text{Nord}_{91}}$	2373.3 (9.85)	2422.1 (10.23)	2457.3 (10.19)	2373.3 (9.98)
$\beta_{\text{Centre}_{86-90}}$	708.5 (3.97)	779.1 (4.53)	771.4 (4.43)	636.9 (3.92)
AR(1) $\rho_2$	0.783	0.718	0.720	0.785
Godfrey LM	1.250	0.661	0.647	0.985
Shapiro-Wilk $W$	0.982	0.980	0.974	0.983
$R^2$	0.8593	0.8586	0.8546	0.8483
<b>System</b>				
Henze-Zirkler $T$	1.198	0.817	1.112	0.899

Estimated asymptotic t-values are given in parentheses

Lagrange multiplier tests for non-linear regression models (Godfrey 1988, p. 117; White, 1992). The test statistic has a critical value of 3.84 at a 5 percent level, which implies acceptance of the AR(1) process for all our specifications (see Table 3.3). It is also comforting that normality of the error structure is accepted for both single equation tests (Shapiro-Wilk) as well as for a system test (Henze-Zirkler) for all specifications.

Table 3.3 reports three different specifications of the policy function, labelled respectively (b), (c) and (d), depending on which lobbying variable is used. In (b) both  $N_{it}$  and  $D_{i,t-1}$  are included together; in (c), only  $N_{it}$ , the number of large establishments, is included; in (d), only  $D_{i,t-1}$ , the partisan dummy, is included. The estimated AR(1) parameters  $\rho_1$  and  $\rho_2$  are about 0.90 and 0.72 respectively.

Two preliminary remarks on Table 3.3's results are in point. First, the proportion of the variability in regional infrastructure investments explained by the policy equation is high (the  $R^2$  is about 0.86), given that the equation includes only  $D_{i,t-1}$  and three regional dummies as out-of-model explanatory variables. Second, the reported parameter estimates turn out to be fairly robust across estimation procedures (OLS and FIML) as well as with respect to changes in the lobbying variable. This remarkably good fit of the policy equation can also be seen by a comparison of the actual values of transport infrastructure investment from Table 3.4 with the predicted values from Table 3.5.

The results reported in Table 3.3 suggest that lobbying, as we proxy it, exerts a statistically significant and quantitatively non-negligible influence on the allocation of infrastructure investment across regions. Their primary interest is qualitative—namely, that lobbying matters.<sup>14</sup> Quantitative estimates are, of course, sensitive to model specification (although the estimate of  $\hat{\theta}_{LOBBY}$  proved remarkably stable) but they nevertheless provide a rough estimate of the orders of magnitude involved, and it is instructive to take a look at them, albeit a very cautious one. *Ceteris paribus*, an additional “representative” large establishment in a region brings that region 1.46 Francs of additional infrastructure investment

<sup>14</sup>We do not directly interpret  $\theta_{LOBBY}$ , but the weight of lobbying by firms defined as  $a_2 = \theta_{LOBBY}/\alpha_X$ . Statistical tests of the null hypothesis that  $a_2$  equals zero yield the following results: Likelihood Ratio (LR) 5.86, Wald 2.90 and Lagrange Multiplier (LM) 4.90. These tests are distributed as  $\chi^2(1)$ , thus at a 10 percent level the null hypothesis is rejected by all tests.

*Table 3.4: Infrastructure Investment Allocation Across Regions and Years*

<b>Actual Values of Transport Infrastructure Net Investment Per Worker, 1985-1991, in 1992 Francs</b>							
	1985	1986	1987	1988	1989	1990	1991
Alsace	1085	1454	1306	1708	1291	1685	1723
Aquitaine	688	1106	1090	1598	1533	1984	1496
Auvergne	613	1297	1543	2026	1949	2334	1358
Basse-Normandie	526	873	620	938	878	1402	1250
Bourgogne	529	944	1115	1665	1235	1697	1424
Bretagne	412	818	871	1432	1527	1904	1449
Centre	1357	2106	2613	3003	3492	3343	2373
Champagne-Ardenne	552	965	868	1281	1128	1892	2158
Charentes Franche-Comte	542	675	580	1052	836	1451	1074
Haute-Normandie	416	617	635	865	1127	1575	1662
Ile de France	455	834	1148	1621	1532	1730	1747
Languedoc-Roussillon	630	685	822	1391	1439	1735	1574
Limousin	981	1408	1515	1288	1407	1592	1432
Lorraine	1096	1086	860	1398	1232	1693	1681
Midi-Pyrenées	811	1074	1127	1342	1402	1864	1608
Nord-pas de Calais	863	1766	1367	1983	1790	2936	4934
Pays de la Loire	868	1017	1111	1560	1255	1243	995
Picardie	1504	1625	1896	1732	1433	1586	1499
Poitou	888	1294	1095	1500	1186	1693	1276
Provence-Alpes-Côte d'Azur	505	836	739	1189	1312	2077	2240
Rhône-Alpes	1086	1338	1602	2218	1739	2142	2486

per worker each year; or, with an average of 1,022,000 workers, a total of 1.5 million francs for the representative region, a relatively small amount (the number of large establishments per region varies between 6 in Limousin and 113 in Rhône-Alpes). A region with a president sharing the current national executive's political obedience will attract 123.8 Francs more of infrastructure investment per worker than one having a president of the 'wrong' obedience; for the average region, this means an additional 126.5 million Francs, or 8.8 percent of average infrastructure investment. Moreover, lobbying by firms having large establishments in the region and the political orientation of the region's president slightly reinforce each other, as expected from our two-stage lobbying model whereby firms first approach local politicians, who then take up their demands to the relevant Ministry. Abundant anecdotal evidence<sup>15</sup> suggests that our results do capture a phenomenon widely perceived as important. Several caveats are in point, however. First, in the model of section 2, lobbying comes from *users* of transport infrastructure, whereas in reality, the construction industry itself is one of the most active lobbyists as far as new highway construction projects are concerned. But the construction industry is composed of a few giants such as Bouygues for whom location across regions is irrelevant, and a host of small firms many of which are below our cutoff of 500 employees (a construction lobbyist recently boasted that the industry association has "52,000 members, practically one in each commune").<sup>16</sup> This type of lobbying activity is not or only imperfectly picked up in our framework. Second, region presidents are not the only local politicians involved; members of parliament are also important relays of local lobbying pressure.

If the positive results concerning lobbying activity were to be expected—although perhaps not as clear-cut as they turned out to be—the insignificance of the productivity term, which picks up the government's concern to allocate in-

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<sup>15</sup>See for instance the cover story of the magazine *Capital* (June 18, 1998) entitled "100 lobbies qui font la loi en France"; in particular pp 92–ff. According to the magazine, the construction industry is a major political-campaign contributor and a powerful force behind highway construction projects, although lobbying by French firms is expected to decline as a result of a Brussels directive imposing open bidding procedures (and therefore diluting the return to lobbying).

<sup>16</sup>*Capital*, 18 June 1998, p. 92.

infrastructure investments to where their marginal product is highest (and to foster regional convergence, since the postulated functional form is concave), is more puzzling.<sup>17</sup> Although it is certainly possible that the government simply doesn't care about the efficient allocation of resources, this conclusion is probably a strong one to draw from such limited evidence and given the scope for misspecification in a simple political-economy model. Moreover, the variety of state-aid schemes aimed at fostering stronger growth in backward regions suggests that European governments, including the French one, do care about convergence—unless, of course, these state-aid schemes are themselves driven by lobbying forces. An obvious alternative for the square-root form used in the first term of the function  $U$ , namely a log form, gave very similar results. A convex form, being implausible since implying preference for divergence, was also tried with inconclusive results. It is therefore fair to say that, as far as this study is concerned, government objectives in the allocation of transport infrastructure investment are unclear once lobbying is controlled for.

Production-function estimates are significant and have the expected sign. Constant returns to scale are not rejected. The estimated elasticity  $\hat{\alpha}_K$  of private capital is 0.181 and that of infrastructure  $\hat{\alpha}_X$  is 0.101; both estimates are significant at the 5 percent level, and remarkably stable across estimation procedures: the OLS infrastructure elasticity estimate is about 0.099, suggesting, as noted in the introduction, that the simultaneous-equation bias from OLS estimation of the production function is negligible. Our estimate of the infrastructure share is much lower than Aschauer's (1989) estimate on US aggregate data (0.39) but the two are not directly comparable since Aschauer's infrastructure variable was a broad aggregate of public capital whereas ours is limited to transport infrastructure. Munnell's (1990) estimate, which was more directly comparable to ours in that she used state-level data, was 0.14, whereas de la Fuente and Vives' (1995) estimate on Spanish regional data was somewhat higher than ours. Although plausible, our estimate should nevertheless be interpreted cautiously, as  $\hat{\alpha}_X$ , in all

<sup>17</sup>As before, we do not directly interpret  $\theta_{PROD}$ , but the weight of productivity concerns by the governments defined as  $a_1 = \theta_{PROD}/\alpha_X$ . Statistical tests of the null hypothesis that  $a_1$  equals zero are not rejected at a 10 percent level by the Likelihood Ratio, Wald and Lagrange Multiplier tests.

likelihood, picks up not only the supply-side effects of infrastructure investments (what it is meant to measure) but also their demand-side or Keynesian effects; it is in fact possible that the latter dominates the former. Moreover, a common drawback of the production-function approach is that it takes the private capital stock as fixed, which can be a valid approximation of reality only in the short run (see de la Fuente and Vives, 1995, for a discussion and alternative formulation); the same is true of employment, although inter-regional labour mobility is arguably a lesser problem than interregional capital mobility. Thus, our estimates are best construed as short-term ones. Finally, we have not included human capital for lack of reliable data; although this is, in general, a potentially serious omission, systematic cross-regional variation in educational levels also may not be a serious a problem given France's relatively egalitarian education system.

As the rates of return on infrastructure capital implied by production-function estimates have been a subject of intense debate in the US (see e.g. Office, 1988, or Gramlich, 1994), it is instructive to calculate the rates of return implied by our estimates for private and infrastructure capital. Let  $r_K$  be the rate of return on private capital; in a competitive environment the unconstrained demand for private capital is given by  $r_K = \hat{a}_K Q/K$ . Assuming that the short-run stock of private capital is at its long-run equilibrium level and using national aggregates of  $Q$  and  $K$  averaged over our sample period, the implied rate of return is 0.156, which is lower than estimates from US data (see e.g. Munnell, 1990b) but nevertheless plausible. As for infrastructure, the implied rate of return, using again national aggregates averaged time-wise, is  $r_X = \hat{a}_X Q/X = 0.522$ ; this is slightly higher than the upper bound of the range of values reported by the US Congressional Budget Office, which vary between 0.35 for highway maintenance projects and 0.05 for new rural highway projects (see Gramlich, 1994, Table 4). Thus, the high rate of return on infrastructure capital implied by our elasticity estimate suggests that in France's case there is some ground to the claim that, overall, transport infrastructure is underprovided; in fact, using our elasticity estimates, the value of the infrastructure stock that would bring its rate of return down to the rate of return on private capital would be 115,221 Francs per worker, or 2.3 times the current one. However, the difference in rates of return between private and in-

*Table 3.5: Predicted Values*

<b>Predicted Values of Transport Infrastructure Allocation from Table 3.3 (b), in 1992 Francs</b>							
	1985	1986	1987	1988	1989	1990	1991
Alsace	772	1318	1410	1726	1549	1766	1566
Aquitaine	847	1042	1142	1451	1471	1886	1750
Auvergne	676	941	1254	1752	1743	2165	1987
Basse-Normandie	572	864	916	1182	981	1402	1317
Bourgogne	848	945	1030	1470	1537	1705	1551
Bretagne	595	791	867	1358	1343	1870	1676
Centre	808	2306	2073	2876	2757	3548	2188
Champagne-Ardenne	725	910	1026	1376	1246	1614	1684
Charentes Franche-Comte	471	839	756	1115	1033	1335	1320
Haute-Normandie	796	835	784	1223	953	1636	1461
Ile de France	504	810	930	1588	1521	1926	1754
Languedoc-Roussillon	770	971	828	1244	1292	1811	1552
Limousin	651	1201	1328	1592	1168	1979	1465
Lorraine	818	1333	1132	1426	1367	1706	1567
Midi-Pyrénées	784	1105	1133	1483	1261	1787	1645
Nord-pas de Calais	1116	1225	1702	1582	1762	2350	4934
Pays de la Loire	821	1188	1055	1593	1485	1735	1232
Picardie	935	1639	1550	2055	1586	1813	1487
Poitou	807	1171	1278	1453	1396	1620	1536
Provence-Alpes-Côte d'Azur	878	908	976	1210	1171	1740	1823
Rhône-Alpes	1216	1474	1354	2039	2101	2165	1944

frastructure capital should not be overplayed, as rates of return are very sensitive to elasticity estimates, which are themselves fairly imprecise.<sup>18</sup> Moreover, France was, during our sample period, in the middle of a major effort of transport infrastructure construction, both for highways and for high-speed railway lines. The picture might be different a decade later.

*Table 3.6: Simulated Solution Values*

Actual, Predicted & Optimal Values of Transport Infrastructure Net Investment Allocation, Averages of Years 1985-1991, in 1992 Francs				
	Actual	Predicted	Method I	Method II
Alsace	1465	1444	1299	2077
Aquitaine	1356	1370	1275	2492
Auvergne	1589	1503	1464	2783
Basse-Normandie	927	1034	983	2586
Bourgogne	1230	1298	1205	2535
Bretagne	1202	1214	1150	2831
Centre	2613	2365	2266	3564
Champagne-Ardenne	1264	1226	1168	2542
Charentes Franche-Comte	887	982	982	2688
Haute-Normandie	985	1098	1020	2354
Ile de France	1295	1290	1968	3744
Languedoc-Roussilon	1183	1210	1169	2587
Limousin	1375	1341	1299	3005
Lorraine	1292	1336	1244	2575
Midi-Pyrenées	1318	1314	1214	3269
Nord-pas de Calais	2234	2096	1933	3261
Pays de la Loire	1150	1301	1171	3088
Picardie	1611	1581	1480	2991
Poitou	1276	1323	1241	3045
Provence-Alpes-Côte d'Azur	1271	1244	1179	2452
Rhône-Alpes	1802	1756	1524	2949

<sup>18</sup>In fact, the difference between  $r_X$  and  $r_K$  is statistically not significant at a 10 percent level.

Pork barrel (of which we found evidence) distorts not only the overall level of infrastructure investments, but also their spatial allocation. In order to assess the size of these distortions, we perform two experiments. In the first, we calculate predicted optimal values of infrastructure investments,  $\hat{z}'_{it}$ , using the estimated coefficients of the policy function's "non-lobbying" terms. That is,

$$\hat{z}'_{it} = \hat{\beta}_t + \hat{\theta}_{PROD} q_{it}^{1/2} / x_{i,t-1} + \hat{\beta}_{NORD} D_{Nord} + \hat{\beta}_{PARIS} D_{Paris} + \hat{\beta}_{CENTRE} D_{Centre}. \quad (3.21)$$

The resulting values of  $\hat{z}'_{it}$  are reported in the second column of Table 3.6. The major drawback of this approach is that  $\hat{\theta}_{PROD}$  is a very imprecise estimate (indeed, not significantly different from zero). Thus, in the second experiment, using the fact that the infrastructure stock is 0.43 times what it would take to bring its rate of return down to the rate of return on private capital (15.6 percent), we simply assume that the aggregate (nationwide) level of predicted investment is also 0.43 times its optimal value when averaged over the sample period; i.e.  $\sum_t \sum_i \hat{z}_{it} = 0.43 \sum_t \sum_i z_{it}^*$ , and solve the equation

$$2.3 \sum_t \sum_i \hat{z}_{it} = n \sum_t \hat{\beta}_t + \theta \sum_t \sum_i (q_{it}^{1/2} / x_{i,t-1}) + \sum_t (\hat{\beta}_{NORD} D_{Nord} + \hat{\beta}_{PARIS} D_{Paris} + \hat{\beta}_{CENTRE} D_{Centre}), \quad (3.22)$$

for the unknown parameter  $\theta$ . The solution  $\tilde{\theta}$  of (3.22) is then used in place of  $\hat{\theta}_{PROD}$  in (3.21) to recalculate optimal regional investments. The resulting values are reported in the third column of Table 3.6.

Whereas the coefficient of variation of actual investments (averaged over time) is 45 percent, the coefficient of variation of optimal investments is 45.2 percent according to method 1 but only 33.8 percent according to method 2. Thus, our experiment suggests that a least a part of the observed cross-regional variability in infrastructure investments comes from pork-barrel terms. The reason for this is apparent: the optimal investment rule calls for equalisation of the term  $\alpha_X q_{it}^{1/2} / x_{it}$  (after adjustment with regional dummies); as long as output per head ( $q_{it}$ ) and the infrastructure stock per head ( $x_{it}$ ) do not vary too much, the optimal allocation is nearly uniform. This result has an important practical consequence: whenever political distortions to the allocation of infrastructure are a source of

concern—as they are in our sample—the uniform rule, which is simple to administer and monitor, is a good rule of thumb. Of course, there is a caveat; namely that if the ratio  $q_{it}^{1/2}/x_{it}$  does not vary excessively in an industrial country with a large existing infrastructure stock, the same is not necessarily true in developing countries with patchy infrastructure stocks, where the uniform rule could be misleading.

### 3.4 Concluding Remarks

The primary interest of our results is that they highlight the importance of the pork-barrel dimension of policy-making, suggesting that explicit modelling of the political-economy processes driving economic-policy decisions is interesting in its own right, irrespective of whether its omission would or would not introduce a simultaneity bias in regressions where policy variables are treated as exogenous. Commenting on the high rates of return on infrastructure investments estimated by Aschauer, Gramlich (1994) remarked, “If public investment really were as profitable as claimed, would not private investors be clamouring to have the public sector impose taxes or float bonds to build roads, highways, and sewers to generate these high net benefits? [...] Very little such pressure seems to have been observed, even when the implied econometric rates of return were allegedly very high” (p. 1187). We find that, in the absence of a loud clamour, the quiet whisper of lobbies can, indeed, be heard. The interest of our political-economy approach is that it can provide indications—however rough—both on the departure of policy from the social optimum and on the extent of special-interest influence. As far as policy implications are concerned, our results contain good news and bad news. The bad news is that influence activities appear to be significant drivers of infrastructure-investment decisions, whereas non-lobbying governmental objectives, if any, are unclear. The good news, however, is that the resulting distortions appear to be relatively minor. First, feedback effects on production-function estimates are weak, and the marginal product of infrastructure capital does not vary tremendously across regions, so that departures from the first-best allocation of infrastructure across regions are fairly inconsequential. Second, in rich industrial

countries, transport infrastructure investments are small compared to the level of the existing stocks, so that political distortions in the amounts and spatial allocation of investments are unlikely to make themselves felt on GDP before a while. But one should not be excessively optimistic about this. First, if investment decisions have *always* been made on the basis of pork-barrel politics, the stock levels should themselves be severely distorted. So our results beg the question: when did things start getting seriously bad? In France's case, the answer seems to be fairly recently. The conventional wisdom among political scientists is that corruption has vastly expanded in the 1980s, largely as a result of administrative reforms enacted in 1982 (see e.g. Mény, 1992; Borraz and Worms, 1996; or Service Central de Prevention de la Corruption, 1994).<sup>19</sup> Second, if pork barrel is prevalent in infrastructure-investment decisions (although de la Fuente and Vives (1995) found little trace of political influence in Spanish infrastructure investment decisions), developing countries are likely to be less robust to the ensuing distortions simply because the stocks are so much smaller relative to the investments. Under such conditions, political distortions in the allocation mechanisms are unlikely to be innocuous.

If, as our positive analysis suggests, political distortions ought to be taken seriously, at least in the long run, one should be able to offer normative guidance for the design of rules or institutions that could mitigate those distortions. The second interesting aspect of our results is that they provide just such a rule. Given our functional forms (alternative ones give similar rules) the first-best allocation of infrastructure equalises the term  $\alpha_X q_{it}^{1/2} / x_{it}$  across regions. Provided that neither productivity levels ( $q_{it}$ ) nor infrastructure stocks per worker ( $x_{it}$ ) vary too much across regions (our data suggests that they don't), uniform allocation is a good enough rule of thumb. Even if the ratio varies, it is not a very difficult one to compute, so the more sophisticated rule is itself not excessively demanding. Of course, if the rule is clear, how it should be implemented is not as clear, since rational politicians are unlikely to abide by a rule. What mixture of centralised vs. decentralised decision-making is least conducive to pork barrel is a question

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<sup>19</sup>We are grateful to Jean-Louis Briquet, from the Institut d'Etudes Politiques de Paris, for a useful conversation on this and for attracting our attention to the relevant political-science work.

that we leave open; only careful international comparisons will shed light on it. What is clear from our work is that France does not yet seem to have the answer.

## Essay 4

# Political Economy of Infrastructure Investment Allocation: Evidence from a Panel of Large German Cities

with

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### *Abstract*<sup>1</sup>

*This essay applies a simultaneous-equation approach to the estimation of the contribution of infrastructure accumulation to private production. A political economy model for the allocation of public infrastructure investment grants is formulated. Our empirical findings, using a panel of large German cities for the years 1980, 1986, and 1988, suggest that cities ruled by a council sharing the State ('Bundesland') government's current political affiliation were particularly successful in attracting infrastructure investment grants. With regard to the contribution of infrastructure accumulation to growth, we find that public capital is a significant factor for private production. Moreover, at least for the sample studied, we find that simultaneity between output and public capital is weak; thus, feedback effects from output to infrastructure are negligible.*

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<sup>1</sup>A revised version of this essay is forthcoming in *Public Choice*, 2002

## 4.1 Introduction

This essay examines the role of public capital in private production and provides empirical evidence on the political-economy determinants of the allocation of public infrastructure investments. From this perspective, our study links the literature on the productivity effects of infrastructure with the literature on the political-economy of fiscal federalism.

Since Aschauer published an influential series of papers (1988; 1989*a*; 1989*b*; 1989*c*) about the effects of public infrastructure investment for long-run growth and productivity in the U.S. and other major countries, there has been an ongoing debate about the role of public infrastructure in generating national welfare. Aschauer (1989*a*), for example, using a production function approach with aggregate time-series data for the U.S. from 1949 to 1985, found that the elasticity of output with respect to a broad measure of public infrastructure was significant and of a remarkable magnitude. At a time of widespread concern about the slowdown of U.S. productivity growth in the 1970's and 1980's this finding suggested that the general decline in public infrastructure spending in the U.S. since the 1970's could at least partly explain the observed slowdown in productivity growth.

However, the magnitude of the estimated elasticity of infrastructure capital in Aschauer (1989*a*; 1989*b*; 1995) and other studies (Garcia-Milà and McGuire, 1992; Munnell, 1990*a*; Munnell, 1990*c*; Munnell, 1992; Munnell, 1993) is still a matter of discussion (for an overview, see Gramlich, 1994). The main focus of the so-called 'infrastructure' debate is on the interpretation of results and the appropriate empirical methodology (Aaron, 1990; Holtz-Eakin, 1994). For example, it is argued that the direction of causation is unclear, i.e., whether causality runs from infrastructure to output or from output to infrastructure (Tatom, 1991*a*; Tatom, 1993*a*). In order to address the problem of causality econometrically several studies have suggested simultaneous-equation-approaches with public infrastructure investment as an endogenous variable (e.g., Cadot, Röller and Stephan, 1999; Duffy-Deno and Eberts, 1991; de Frutos and Pereira, 1993).

Similar to Duffy-Deno and Eberts (1991) or Cribfield and Pangebean (1995)

our study estimates the contribution of public capital to private production at the local level. This approach seems to be justified by the fact that about 60 percent of public infrastructure is provided by local governments and not by the federal or states governments (Seitz, 1995).

Infrastructure investments at the municipal level in Germany usually consists of two parts: autonomous investment and matching investment grants from higher-tier governments. The increasing weight of investment grants for the realisation of local investment projects in Germany suggests to model both parts (grants and autonomous investments) separately within our simultaneous equation approach. Whereas the former is a matter of decision for the municipal councils the latter is predominantly provided by the federal states ('Bundesländer').

The literature on the role of fiscal federalism for infrastructure policies so far has mainly discussed optimal rules for the provision of infrastructure at different levels of government (e.g., Hulten and Schwab, 1997). However, it remains an open question whether infrastructure policies in reality are designed according to such efficiency considerations. Therefore, the main contribution of our essay is that we empirically shed light on other potential determinants of infrastructure policies and test them against traditional efficiency arguments.

In this essay we adopt the approaches of Cadot, Röller and Stephan (1999) and Crain and Oakley (1995) in that we analyse the politics of infrastructure. What we suppose as politico-economic determinants of local infrastructure investment decisions are (i) 'pork-barrel' infrastructure policies due to the influence of firms on the allocation of investments or (ii) distortions in allocation of intergovernmental infrastructure investment grants due to the political affiliation of governments at different levels (iii) distortions in allocation of investment grants due to the strategic advantage of heavily contested constituencies ('swing voter' approach). All in all, these potential influences may give rise to outcomes of local infrastructure investment decisions that might depart substantially from an optimal allocation as a result of maximising social welfare.

With our empirical model we test these different ideas on a panel data set consisting of 87 German cities for the years 1980, 1986 and 1988. We use a simultaneous-equations approach to estimate the relationship between infrastructure

investments, investment grants, local manufacturing output, policy and lobbying variables. The main findings of our analysis are (i) the contribution of local public capital to private production in cities is positive and significant (ii) political affiliation, measured by the coincidence of party colour between state and local government, is decisive in explaining the distribution of investment grants across cities (iii) cities with a prevalence of ‘marginal voters’ do neither spend more on public infrastructure nor receive more investment grants from higher-tier governments (v) the larger the majority of government in city council the higher is local infrastructure spending (vi) investment grants do not induce higher autonomous infrastructure spending of cities, i.e. there is no evidence of a complementary relationship between matching investment grants and infrastructure spending (vii) efficiency considerations do not seem to determine the observed intergovernmental grant allocation across cities whereas redistributive concerns of higher-tier governments matter.

The remainder of this essay is organised as follows: In section 4.2, we discuss the determinants of local infrastructure policies in Germany. Section 4.3 elaborates the hypotheses and presents the structure of the empirical model. Section 4.4 describes the empirical implementation and reports estimation results. Section 4.5 provides conclusions.

## 4.2 Determinants of Local infrastructure policies in Germany

The resources for infrastructure investments in Germany are usually by means of mixed financing between two or more levels of German governments. There are two different financial sources for infrastructure investments: autonomous investments by municipalities and investment grants<sup>2</sup> provided by other institutions, the central government (‘Bund’), the states (‘Bundesländer’) the ERP or horizontal fiscal exchange mechanisms.

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<sup>2</sup>These grants or subsidies from higher levels are called ‘Finanzzuweisungen’ (financial assignments). One major example in the infrastructure context are GVFG-funds that were created to promote transportation infrastructure.

The majority of investment grants from states to municipalities takes the form of matching funds. Nevertheless, fixed matching ratios between state level grants and local investments are rarely found as a result of planning problems and changing investment costs (Statistisches Bundesamt, 1986).

The procedure of starting a new infrastructure investment project is a complex arrangement between the local government, which makes proposal in the first stage of the project planning, and the state administration that grants an investment subsidy. Because of the growing fiscal tension in the local budgets (e.g. Pohlan, 1997), the role of investment subsidies in Germany has risen all over the 1980's.<sup>3</sup> In 1980 the ratio between investment subsidies from the federal government ('Bund'), states and the ERP to total investment in transport infrastructure was about 24 percent whereas in 1988 this ratio had risen to 46 percent. The municipalities' dependency on investment grants makes it also difficult for them to plan investment projects autonomously. A first reason is the overall increase of insecurity in the planning process, as local decision-makers cannot anticipate the correct amount of future transfer payments (Statistisches Bundesamt, 1986, p. 913).

On the one hand, mixed financing of local infrastructure project is usually justified as a means of internalising positive externalities from infrastructure projects (e.g., Oates, 1999). On the other hand, the political cost of mixed financing of infrastructure projects is that local political autonomy is undermined. An example might illustrate this point: Schmals and Siewert (1982) elaborates a case study about public transportation in Munich in the 1970's. Two alternative plans to improve public transportation existed. The first plan proposed the construction of a network of underground railways to alleviate inner-city traffic. The majority of city council members favoured this project. The construction and improvement of a municipal railway system, the second proposal, was backed by Bavarian government. Because the Bavarian government linked an investment grant with the realisation of the second project, the city council had to give in. Thus, in this

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<sup>3</sup>There are, of course, alternative views which state that the growing dependence on investment grants is endogenous to institutional change (Inman, 1988, p. 56). But the overwhelming part of literature on German fiscal federalism insists on the growing dependency of lower level governments on grants.

case investment grant prospects had a decisive impact on the bargaining power between the two governmental levels.

The amount of investment subsidies granted to local infrastructure projects formally depends on such external factors e.g. as length of the existing road infrastructure, expected impacts on the local economy, environmental effects, etc. But as Garlich (1986) shows for the case of infrastructure funds for highways, the actual amount of money is a matter of intense bargaining between all lower level governments and the higher level. An iron quotas system, which is more than often in German politics the result of bargaining processes or of even legally settled principles such as unanimity, is likely to create further distortions.

The assumption of a simple (Westminster) representative democracy is severely violated in the German case. One of the reasons for this is the peculiar nature of German federalism which has been characterised in the literature as 'unitary federal state'. In Germany spheres of competence and control as well as the financing of investments is not as separated as it is for example in the case of the U.S. federalism but it is overlapping and mutually dependent. Therefore, although the states ('Bundesländer') and the municipalities and autonomous cities are exclusively responsible for the main part of public infrastructure in a formal sense, investment projects also depend on the amount of public subsidies that either come from the federal government, the states or from the European Union.

Moreover, German federalism is constitutionally obliged to balance local autonomy and the uniformity of living conditions throughout the German territory. Humplick and Moini-Araghi (1996) argue that this often results in a less efficient provision of public infrastructure. As they put it 'the equity objective overrides the efficiency objective' (Humplick and Moini-Araghi, 1996, p. 32). These obligations of German federalism create the need for a network of horizontal and vertical bargaining institutions that are coordinating the interests of the several governmental levels.

In our political-economy framework we explicitly test political variables that might influence the allocation of investment subsidies. We argue that because of the complex federal system which can be described by the notion of 'intertwined politics' it is likely that those local governments whose political 'colour'

corresponds to that of the state ('Bundesland') government get more investment subsidies, because this lowers the transaction costs of information transmission between governments. The identity of political colour shortcuts this bargaining process and favours certain municipalities by means of party loyalty.

### 4.3 Hypotheses and structure of the model

Our simultaneous equation model is based on 3 equations, which we label as (i) production function  $Q_{it} = f()$ , (ii) infrastructure investment function  $INV_{it} = f()$  and (iii) grant allocation function  $GRANT_{it} = f()$ .

#### 4.3.1 Production function

To begin with the specification of the production function, we assume that production  $Q_{it}$  of the manufacturing sector can be described as

$$Q_{it} = f(t, K_{it}, L_{it}, G_{it}), \quad i = 1 \dots N, \quad t = 1 \dots T, \quad (4.1)$$

where  $t$  denotes time,  $Q_{it}$  output,  $K_{it}$  private capital,  $L_{it}$  labour input and  $G_{it}$  denotes the infrastructure stock in city  $i$ . In addition, city  $i$ 's infrastructure stock  $G_{it}$  is defined as

$$G_{it} = (1 - \gamma)G_{i,t-1} + INV_{it} + GRANT_{it}, \quad (4.2)$$

where  $\gamma$  denotes the depreciation rate of public capital,  $INV_{it}$  denotes infrastructure investment, and  $GRANT_{it}$  denotes infrastructure investment grants given to city  $i$  from higher-tier governments. Therefore, total infrastructure investment in city  $i$  is defined as  $INV_{it} + GRANT_{it}$ .

Assuming a Cobb-Douglas functional form for the manufacturing sector's production function in city  $i$  at time  $t$  we get

$$Q_{it} = A_0 \exp(\alpha_t t) L_{it}^{\alpha_L} K_{it}^{\alpha_K} G_{it}^{\alpha_G}, \quad (4.3)$$

where  $\alpha_X$  denotes the elasticity of output  $Q$  with respect to input  $X$ , and  $X \in \{L, K, G\}$ . Dividing by  $L_{it}$ , (4.3) becomes

$$q_{it} = A_0 \exp(\alpha_t t) k_{it}^{\alpha_K} g_{it}^{\alpha_G} L_{it}^{\tilde{\alpha}_L}, \quad (4.4)$$

where lower-case capitals denote variables in terms of the labour input  $L$  and  $\tilde{\alpha}_L$  is defined as  $\tilde{\alpha}_L = \alpha_L + \alpha_K + \alpha_G - 1$ .

Note that  $\tilde{\alpha}_L$  will equal zero if returns to scale are constant with respect to *all* inputs, i.e.,  $L$ ,  $K$  and  $G$ ; and  $\tilde{\alpha}_L - \alpha_G$  will equal zero if returns to scale are constant with respect to private inputs  $L$  and  $K$ .

### 4.3.2 Infrastructure investment function

The increasing weight of investment grants for the realisation of infrastructure projects suggests to model both parts (grants and autonomous investments) separately within our simultaneous equation approach. Accordingly, to describe the simultaneous determination of investments and grants properly, besides the local production function our model is based on two additional equations: one which describes autonomous investment decisions of the cities and one which describes the level of investment grants the cities receive from higher-tier governments. Furthermore, autonomous investments enter the grants equation and, vice versa, grants enter the investment equation.

Our hypotheses regarding the determinants of city's autonomous infrastructure spending can be summarised as follows. The first hypothesis we are able to test with our model is with regard to the relationship between grants and autonomous investments. Though the major part of grants a city in Germany receives is matching funds it nevertheless is an open question whether these matching grants have a complementary, substitutive or neutral relation to the autonomously financed infrastructure investments of cities.

The reason is that even in the case of matching grants the relationship between grants and investment is not necessarily positive and therefore complementary, since the local government can reduce its own efforts on financing infrastructure projects by taking into account the amount of grants it will receive for a project from higher-tier governments.

Accordingly, grants and autonomous infrastructure spending is only complementary if grants do not lead to a reduction of financing efforts by local governments. Therefore, if the relationship between grants and investments is comple-

mentary a local government which receives grants will *autonomously* finance more infrastructure projects than a government which does not receive any grants.

On the other hand, if local governments plan their infrastructure projects irrespective on the amount of future matching grants, then the relationship between autonomous spending and grants can be labelled as neutral. This implies that the local government will neither reduce nor increase its own financing efforts when anticipating the matching grants it receives. Consequently, the local government's own financing efforts are independent from the amount of matching grants.

The second hypothesis we test with our framework is that local infrastructure spending should also reflect the preferences of a city's residents. For instance cities with relatively more cars are likely to spend more on transport infrastructure.

The third hypothesis we test is whether a local government's spending on infrastructure is more responsive to increases in intergovernmental grant receipts than it is to increases in own city's tax revenues. The regularly finding of various previous studies on this topic is that local government's spending is more responsive to intergovernmental grant receipts has been dubbed in literature as the 'flypaper effect'—money sticks where it hits (e.g. Oates, 1999; Oulasvirta, 1997).

Moreover, following an idea proposed by Cadot et al. (1999) we test the hypothesis that the number of manufacturing firms is decisive for local infrastructure spending. The main motivation for this presumption is that particularly manufacturing firms have sunk investments and therefore have a vested interest in the quality and maintenance of the infrastructure where they have their production located. Local politicians—on the other hand—are assumed to be sensitive for the lobbying efforts by business, for instance in anticipation of potential campaign contributions from firms, or in anticipation of the expected loss of trade tax revenues and/or employment opportunities for their city if firms move to another location.

The fifth hypothesis regarding determinants of local infrastructure investment is the role of the stability of the government majority in the city council. If local governments want to buy the support of the local swing voters, one would expect

that the smaller its majority in the city council the larger is its spending on local infrastructure projects.

Finally, the sixth hypothesis we test is the presumption that local governments might take the expected productivity effects of infrastructure spending on the local industry into account. Because of this, if local politicians indeed care about the efficiency of infrastructure projects we would observe a positive effect from the expected productivity effect of these infrastructure projects on actual the amount of infrastructure spending. Hence if expected productivity effect is higher in a given city spending of the local government should be higher as well.

### 4.3.3 Grant allocation function

The first hypothesis we can test with our model is the empirical relevance of the traditional main topic on intergovernmental grant allocation, i.e. the question whether or not grant allocation policies are based on efficiency and/or equity criteria. Accordingly, we include in our model both a measure for expected productivity effects from infrastructure projects (efficiency) as well as income as a measure for redistributive concerns (equity).

However, a recent strand of literature discusses alternative politico-economic influences on the intergovernmental grant distribution. For instance, Grossman (1994) hypothesizes that the distribution of grants is driven by the self-interest of grant givers. The assumption is that politicians from higher-level governments are likely to allocate grants for the purpose of enhancing their reelection chances. In the words of Grossman higher governmental level politicians uses grants to 'purchase political capital' to be used to influence the voting decisions of the local residents.

Accordingly, the second hypothesis we test with our model is that party affiliation between higher and lower-tier governments matters for the outcome of grant allocation. Grossman (1994) states that the political capital is of higher value to grant-givers if the party affiliation with grant-receivers is the same. However, in the specific case of German cities, our interpretation why party affiliation matters for grant-givers is that the identity of political colour shortcuts the bargaining

process between lower and higher-tier governments and thereby favours certain municipalities by means of party loyalty.

Recently, it has also been suggested in the literature that grants are used as tactical (electoral politics) instruments for buying support of marginal voters ('swing voter' approach, e.g. Dixit and Londregan, 1998; Johansson, 1999). Using this framework, the third hypothesis we test is that cities will receive more grants if they are politically powerful, i.e. if there is a large number of voters who are indifferent between the two parties and therefore potentially could be influenced by pork barrel politics. Following Johansson we proxy political powerfulness as closeness between the major two blocks, Social Democrats (SPD) and Christian Democratic Union (CDU), in the last election for the city council. Hence, we expect that if there is evidence of political powerfulness as a determinant for the distribution of grants, the closer the last election results between the two major blocks, the larger the amount of grants a city receives from higher-tier governments.

Finally, the fourth hypothesis we test is that the number of manufacturing firms is decisive for grant-givers when allocating grants across regions. The reason is again the expectation that also higher-tier governments are sensitive to business interests. If business interests indeed matter for the outcome of infrastructure policies, then a priori it is not clear at which level of government lobbying by firms or business associations takes place. For this reason the number of manufacturing firms is included both in the investment and the grant allocation function.

## 4.4 Empirical implementation

### 4.4.1 Data

We use a panel data set consisting of 87 German cities and three years (1980, 1986, 1988). Table 4.1 provides a brief overview of the variables used in the analysis.

Most of the data is taken from the 'Statistical Yearbook of German Cities and Municipalities'.<sup>4</sup> For reasons of data availability only 87 large cities are in-

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<sup>4</sup>Original title: 'Statistisches Jahrbuch der Städte und Gemeinden'.

*Table 4.1: Variable Description and Cities*

<b>Variable</b>	<b>Description</b>
<i>Q</i>	Value added, manufacturing sector, million 1980 DM
<i>L</i>	Hours worked in manufacturing sector, million hours
<i>K</i>	Capital stock in manufacturing, million 1980 DM (from Deitmar, 1993)
<i>G</i>	Public infrastructure stock, million 1980 DM, (from Seitz, 1995)
<i>INV</i>	Public infrastructure investment, million 1980 DM
<i>GRANT</i>	Infrastructure investment grants, million 1980 DM
<i>DEBT</i>	Total debt of city, million 1980 DM
<i>TAX</i>	Trade tax ('Gewerbsteuer') revenues of city <i>i</i> , million 1980 DM
<i>CARS</i>	Number of registered motor vehicles (business and private)
<i>NFIRMS</i>	Number of manufacturing firms in city <i>i</i>
<i>DMIN</i>	Dummy variable equal to 1 when mining industry is present in city <i>i</i>
<i>PARTISAN</i>	Percentage of members in city council with the same political affiliation as the federal state ('Bundesland') government
<i>MAJORITY</i>	Percentage difference of the 2 large parties SPD (Social Democrats ) and CDU (Christian Democratic Union) in last city council election, values rank transformed from 1 (largest) to 261 (smallest difference)

cluded in the sample. All of these cities are predominantly self-administered (autonomous) at the local level ('kreisfreie Städte'). Because of this, from the fiscal federalism perspective, these cities are highly comparable. Table 4.2 displays the names of cities in our sample.

Output ( $Q$ ), measured as gross value added of a city's manufacturing sector,<sup>5</sup> is taken from a joint publication of several German federal states statistical offices.<sup>6</sup> These data are not available for each year, which restricts our sample to three years, 1980, 1986, and 1988.

The private capital stock ( $K$ ) of the manufacturing sector is taken from Deitmar (1993). It is measured in 1980 prices and has been corrected for the territorial reforms that occurred in the 1970's in Germany.<sup>7</sup> The infrastructure capital stock ( $G$ ), which includes investments both for construction and equipments, is taken from Seitz (1994) and is also measured in 1980 prices. Transport infrastructure is the largest part (about 30 percent) of local infrastructure (Bach, Gorning, Stille and Voigt, 1994).

Annual investment in infrastructure ( $INV$ ) has been obtained from the statistical yearbook mentioned above. From the same source we have also the following variables: labour input ( $L$ ), operationalised by the number of working hours in the manufacturing sector; special grant-in-aids ('Finanzzuweisungen') for investments ( $GRANT$ ) from 'Bundesländer', 'Bund' or ERP; several measures of the financial situation of a city like the cumulated debt ( $DEBT$ ) or trade taxes revenues ( $TAX$ ) which are levied at the local level of cities, the number of (four-wheel) motor vehicles (private and business) ( $CARS$ ), and the number of manufacturing firms ( $NFIRMS$ ) in a city.

Furthermore, we constructed a political variable denoted as *PARTISAN* to measure the congruence between the local city government and the state ('Bundesland') government. It gives the percentage of seats in the city council with the same political affiliation as the 'Bundesland' government where the city is lo-

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<sup>5</sup>This includes also the mining industries.

<sup>6</sup>'Volkswirtschaftliche Gesamtrechnung der Länder, Bruttowertschöpfung der kreisfreien Städte, der Landkreise und der Arbeitsmarktregionen in der Bundesrepublik Deutschland', Heft 26, Statistisches Landesamt Baden-Württemberg, 1995.

<sup>7</sup>For further details, see Deitmar (1993).

*Table 4.2: Cities in Panel*

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<b>Cities in Panel</b>		
1 Aachen	30 Hamm	59 Neustadt/Weinstraße
2 Amberg	31 Hannover	60 Nürnberg
3 Ansbach	32 Heidelberg	61 Oberhausen
4 Aschaffenburg	33 Heilbronn	62 Offenbach/Main
5 Augsburg	34 Herne	63 Oldenburg
6 Baden-Baden	35 Hof	64 Osnabrück
7 Bamberg	36 Ingolstadt	65 Passau
8 Bayreuth	37 Kaiserslautern	66 Pforzheim
9 Bielefeld	38 Karlsruhe	67 Pirmasens
10 Bochum	39 Kassel	68 Regensburg
11 Bonn	40 Kaufbeuren	69 Remscheid
12 Bottrop	41 Kempten/Allgäu	70 Rosenheim
13 Braunschweig	42 Kiel	71 Saarbrücken
14 Coburg	43 Koblenz	72 Salzgitter
15 Darmstadt	44 Köln	73 Schwabach
16 Delmenhorst	45 Krefeld	74 Schweinfurt
17 Dortmund	46 Landau/Pfalz	75 Solingen
18 Duisburg	47 Landshut	76 Speyer
19 Düsseldorf	48 Leverkusen	77 Straubing
20 Erlangen	49 Lübeck	78 Stuttgart
21 Essen	50 Ludwigshafen	79 Trier
22 Flensburg	51 Mainz	80 Ulm
23 Frankenthal/Pfalz	52 Mannheim	81 Weiden/Oberpfalz
24 Frankfurt/Main	53 Memmingen	82 Wiesbaden
25 Freiburg/Breisgau	54 Mönchengladbach	83 Wilhelmshaven
26 Fürth	55 Mülheim/Ruhr	84 Worms
27 Gelsenkirchen	56 München	85 Wuppertal
28 Göttingen	57 Münster/Westfalen	86 Würzburg
29 Hagen	58 Neumünster	87 Zweibrücken

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**Table 4.3: Descriptive Statistics of Variables**

Variable	Mean	Std.Dev.	C.V.	Minimum	Maximum
<i>Q</i>	2099.1	2500.3	119.1	144.3	15718.8
<i>G</i>	2468.8	2834.5	114.8	302.5	18176.1
<i>K</i>	4087.7	5007.6	122.5	252.0	25714.9
<i>L</i>	30.74	29.08	94.6	2.4	168.2
<i>INV</i>	93.6	123.8	132.3	8.1	1040.4
<i>GRANT</i>	32.8	44.7	136.3	0.8	266.1
<i>DEBT</i>	407.9	509.1	124.8	14.3	3066.7
<i>TAX</i>	135.6	210.4	155.2	7.1	1314.6
<i>CARS</i>	88921	91046	102.4	14845	635888
<i>NFIRMS</i>	124.0	101.1	81.5	21	637
<i>DMINING</i>	0.126	0.333	263.4	0	1
<i>PARTISAN</i>	45.9	8.0	17.5	29.0	68.2
<i>MAJORITY</i>	131	75.5	57.6	1	258.5
Total number of observations: 261					

cated. All cities had at least one city council election during the period 1980 to 1987, some cities had also 2 city council elections in this period.

In a first step, the variable *MAJORITY* was constructed as percentage difference of the 2 major parties, which are parties SPD and CDU in Germany, from the results of last city council election. In a second step, in order to smooth the highly skewed distribution of this variable and also to make it less correlated with the variable *PARTISAN*<sup>8</sup> a simple monotonic rank transformation has been performed which assigns the variable *MAJORITY* rank numbers from 1 for the observation with smallest to number 261 for the observation with largest difference in majority.

Table 4.3 displays descriptive statistics of the variables. Note, for instance that grants are on average about one-third of autonomous investments. Annual infrastructure investment undertaken by cities is on average about 3.8 percent of the existing infrastructure capital stock. The mining industry is present in about 13 percent of cities in our sample. The partisan variable is on average 45.9 percent,

<sup>8</sup>The correlation between *MAJORITY* without and *MAJORITY* with rank transformation is -0.97. The correlation between *MAJORITY* without rank transformation and *PARTY* is 0.47, the correlation between *MAJORITY* with rank transformation and *PARTY* is -0.40.

with a minimum of 29.0 and a maximum of 68.2 percent.

Our simultaneous model is based on the following 3 equations for city  $i$ ,  $i = 1, \dots, N$ , in year  $t$ ,  $t = 1980, 1986, 1988$ .

$$\begin{aligned} & \textbf{Production function} & (4.5) \\ \ln Q_{it}/L_{it} & = \alpha_0 + \alpha_{BL} + \alpha_t \text{ (+) } \alpha_G \ln((G_{i,t-1} + INV_{it} + GRANT_{it})/L_{it}) \\ & \text{ (+) } \alpha_K \ln(K_{it}/L_{it}) + \tilde{\alpha}_L \ln(L_{it}) + \alpha_{MINING} DMIN_i + u_{1it}, \end{aligned}$$

$$\begin{aligned} & \textbf{Local autonomous investment function} & (4.6) \\ INV_{it}/L_{it} & = \beta_0 + \beta_{BL} + \beta_t + \beta_{GRANT} GRANT_{it}/L_{it} + \beta_{NFIRMS} NFIRMS_{it} \\ & \text{ (+) } \beta_{MAJOR} MAJORITY_{it} \text{ (+) } \beta_{CARS/L} CARS_{it}/L_{it} \\ & \text{ + } \beta_{G/L} G_{i,t-1}/L_{it} \text{ (+) } \beta_{PROD} \alpha_G Q_{it}/G_{i,t-1} \text{ (-) } \beta_{DEBT/L} DEBT_{it}/L_{it} \\ & \text{ (+) } \beta_{TAX/L} TAX_{it}/L_{it} + \beta_{INCOME} Q_{it}/L_{it} + \beta_{MINING} DMIN_i + u_{2it}, \end{aligned}$$

$$\begin{aligned} & \textbf{Grant allocation function} & (4.7) \\ GRANT_{it}/L_{it} & = \gamma_0 + \gamma_{BL} + \gamma_t + \gamma_{INV/L} INV_{it}/L_{it} \text{ (+) } \gamma_{SWINGV} MAJORITY_{it} \\ & \text{ (+) } \gamma_{PARTISAN} PARTISAN_{it} \text{ (-) } \gamma_{REDISTRIB} Q_{it}/L_{it} \\ & \text{ (-) } \beta_{G/L} G_{i,t-1}/L_{it} \text{ (+) } \gamma_{PROD} \alpha_G Q_{it}/G_{i,t-1} \\ & \text{ (+) } \gamma_{NFIRMS} NFIRMS_{it} + \gamma_{MINING} DMIN_i + u_{3it}. \end{aligned}$$

Equation (4.5) refers to the production function of the manufacturing sector in city  $i$  described in section 4.3.1. Equation (4.6) is derived from the hypothesis discussed in section 4.3.2 and describes the autonomous infrastructure investments undertaken by city  $i$ . Equation (4.7) corresponds the hypotheses discussed in section derived from the hypothesis discussed in section 4.3.3 and describes investment grants from higher-tier governments which city  $i$  receives. We add a dummy variable  $DMIN$  to all equations indicating whether or not the mining industry is present in city. If a coefficient has an expected sign it is displayed in parentheses.

From the Cobb-Douglas production function, marginal productivity of infrastructure capital is defined as  $\partial Q_{it}/\partial G_{it} = \alpha_G Q_{it}/G_{it}$ . We include this measure of the expected productivity effects of infrastructure both in the investment and the grant allocation function. Since  $G_{it}$  also contains current investment  $INV_{it}$ , we replaced it with its lagged value  $G_{i,t-1}$ .

Parameters  $\alpha_{BL}$ ,  $\beta_{BL}$  and  $\gamma_{BL}$ ,  $BL = 1, \dots, 8$ , refer to fixed effects for the states ('Bundesländer') and  $\alpha_t$ ,  $\beta_t$  and  $\gamma_t$ ,  $t = 1, 2, 3$ , refer to fixed effects for years.

For disturbances we assume a one-way error-component model with  $u_{kit} = \mu_{ki} + v_{kit}$  for equation  $k = 1, 2, 3$ , where  $\mu_{ki} \sim IID(0, \sigma_{k\mu}^2)$  reflects random individual effects of cities and  $v_{kit} \sim IID(0, \sigma_{kv}^2)$  residual errors (Krishnakumar, 1995).

#### 4.4.2 Results

The results of the estimations are presented in Table 4.4. Unobserved heterogeneity of cities is modelled as random error components (EC) for each equation. By modelling the individual effects as random, it is possible to add fixed effects for states and mining industry to the equations.<sup>9</sup> In addition, the endogeneity of  $Q_{it}$ ,  $INV_{it}$  and  $GRANT_{it}$  as right-hand side variables as well as the correlation of errors across equations is taken into account by using Full-Maximum-Likelihood (FIML) for the estimation of the simultaneous system.

Column 1 of Table 4.4 reports the results for the single equation estimation with Error Components Generalised Least Squares (GLS) (e.g., (Baltagi, 1995), 1995). Columns 2 and 3 contain the results of simultaneous system estimations with Error Components FIML.<sup>10</sup> The specification of column 3 differ from column 2 in that *NFIRMS* is excluded from the investment equation and *MAJORITY* is excluded from the grant equation.

Overall, the fit of the 3 equations is remarkable high with  $R^2$  ranging between 0.71 and 0.81. Turning first to the results for the production function, we find that local public capital is a productive input for local manufacturing. The estimated coefficient which is the elasticity of output with respect to infrastructure is positive and statistically significant with a value of about 0.17. This coefficient is remarkably stable with respect to the different estimation methods and specifications. As the estimates for the infrastructure coefficient do not vary much between single equation and simultaneous equation estimation, the econometric evidence for an endogeneity of infrastructure capital in the production function

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<sup>9</sup>This would not be possible if the error components where modelled as fixed. The main reason, however, why we model unobserved heterogeneity of cities as random is that (i) the random effects model is more parsimonious in parameters (ii) more importantly, our sample does not have sufficient 'within' variation, which is due to the fact that there are only 3 distinct years of observation for each city.

<sup>10</sup>The estimations have been carried out using the PROC MODEL procedure in SAS V8.

**Table 4.4: Estimation Results**

Nonlinear	EC GLS <sup>(a)</sup>		EC FIML <sup>(b)</sup>		EC FIML <sup>(b)</sup>	
<i>Production function: <math>\ln(Q/L)_{it}</math></i>						
$\alpha_{BL}$	Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>	
$\alpha_t$	Fixed effects		Fixed effects <sup>**</sup>		Fixed effects <sup>**</sup>	
$\alpha_0$	0.784	(3.06)	0.828	(4.58)	0.829	(4.59)
$\alpha_G$	0.169	(3.23)	0.170	(4.63)	0.169	(4.60)
$\alpha_K$	0.569	(10.68)	0.555	(14.91)	0.558	(14.95)
$\tilde{\alpha}_L$	0.044	(1.55)	0.045	(2.25)	0.043	(2.16)
$\alpha_{MINING}$	-0.497	(-7.12)	-0.494	(-10.01)	-0.495	(-10.01)
$R^2$ <sup>(c)</sup>	0.814		0.811		0.814	
<i>Infrastructure investment function: <math>(INV/L)_{it}</math></i>						
$\beta_{BL}$	Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>	
$\beta_t$	Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>	
$\beta_0$	-10.77	(-4.40)	-9.77	(-5.10)	-10.88	(-5.25)
$\beta_{GRANT/L}$	0.530	(5.38)	-0.344	(-0.58)	-0.422	(-0.96)
$\beta_{NFIRMS}$	-0.262	(-0.38)	0.223	(0.37)	—	(—)
$\beta_{MAJOR}$	-0.002	(-2.00)	-0.004	(-2.60)	-0.004	(-4.27)
$\beta_{CARS/L}$	1.646	(5.36)	1.594	(6.67)	1.733	(6.52)
$\beta_{G/L}$	0.014	(4.01)	0.025	(3.41)	0.025	(4.43)
$\beta_{PROD}^{(d)}$	3.081	(1.81)	2.252	(1.80)	2.475	(1.94)
$\beta_{DEBT/L}$	-0.048	(-4.04)	-0.046	(-5.11)	-0.047	(-4.91)
$\beta_{TAX/L}$	0.025	(1.67)	0.026	(2.34)	0.029	(2.42)
$\beta_{INCOME}$	-6.199	(-1.76)	-12.31	(-3.34)	-12.80	(-3.73)
$\beta_{MINING}$	-1.643	(-7.29)	-1.61	(-7.86)	-1.55	(-7.00)
$R^2$ <sup>(c)</sup>	0.811		0.753		0.743	
<i>Grant allocation function: <math>(GRANT/L)_{it}</math></i>						
$\gamma_{BL}$	Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>		Fixed effects <sup>***</sup>	
$\gamma_t$	Fixed effects		Fixed effects <sup>*</sup>		Fixed effects <sup>*</sup>	
$\gamma_0$	-0.709	(-1.62)	-0.247	(-0.66)	-1.047	(-3.44)
$\gamma_{INV/L}$	0.171	(5.24)	0.018	(0.27)	0.133	(2.23)
$\gamma_{SWINGV}$	-0.001	(-2.30)	-0.002	(-3.83)	—	(—)
$\gamma_{PARTISAN}$	0.014	(2.34)	0.013	(2.91)	0.020	(5.13)
$\gamma_{REDISTRIB}$	-3.758	(-1.95)	-4.707	(-2.97)	-5.321	(-3.52)
$\gamma_{G/L}$	0.009	(6.25)	0.013	(7.29)	0.011	(6.73)
$\gamma_{PROD}^{(d)}$	0.236	(0.33)	-0.121	(-0.22)	0.317	(0.61)
$\gamma_{NFIRMS}^{(d)}$	0.498	(1.25)	0.388	(1.28)	0.273	(0.97)
$\gamma_{MINING}$	0.293	(2.00)	0.064	(0.46)	0.281	(2.28)
$R^2$ <sup>(c)</sup>	0.738		0.715		0.730	

Notes: t-values in parentheses, <sup>\*\*\*</sup>10 %, <sup>\*\*</sup>5 %, <sup>\*</sup>1 % significant.

EC=Error Components Model, <sup>(a)</sup> Single Equation, <sup>(b)</sup> Simultaneous System.

<sup>(c)</sup> Based on GLS residuals, not bounded [0,1], <sup>(d)</sup> [10<sup>-3</sup>].

is weak. This can be attributed to the fact that infrastructure investment is relatively small compared to the infrastructure stock, thus replacing investment with predicted values from instrumental variables has therefore only a small impact on the estimated parameter for  $G_{it}$ . The ratio of output to public infrastructure stock is on average about 0.95, which implies a rate of return of infrastructure of about 16 percent.

In addition, private capital is significant with a value of about 0.55. The ratio of output to private capital stock is on average about 0.6, which implies a rate of return of private manufacturing capital of about 33 percent. From the value of t-statistic for labour input  $L$  it can be inferred that for the single-equation estimation constant returns are not rejected at a 10 percent level, whereas for the simultaneous equation estimations constant returns to scale are rejected at a 5 percent level. Cities where mining industry is located have a lower expected output.

Turning second to the results for the infrastructure investment function, we find that from the positive and significant coefficient for  $GRANT/L$  in the first column it appears as if grants and local public investments are *complementary*, i.e. grants stimulate further infrastructure projects. However, if the endogeneity of grants is taken into account by applying simultaneous system estimation methods, it turns out that the relationship between grants and local public investments appear to be *neutral*. Thus, the receipt of future grants is not taken into account by local governments when fixing their amount of autonomous spending. Assessing this result, the 'good news' is that cities do not reduce their own efforts in anticipating the receipt of future matching grants from higher-tier governments. The 'bad news' from this result is that cities do not increase their own spending efforts even in the prospect of matching grants for infrastructure projects. Thus, the incentive created by grants for expanding own infrastructure investments are rather low.

With regard to the second hypothesis that local infrastructure spending should also reflect preferences and demand of city's residents and business we find that the coefficient of  $CARS/L$ , which is measured as number of (four-wheel) motor vehicles (business and private) per labour, is positive and highly significant. Thus, cities with a high intensity of cars indeed invest more in infrastruc-

ture.

Turning to economic factors that might determine a city's infrastructure spending, we find that the higher debt ( $DEBT/L$ ) of a city is the lower is its infrastructure spending. This corroborates our initial presumption that the financial room for manoeuvre is decisive for local infrastructure investments. On the other hand, local infrastructure spending is higher the higher trade tax revenues of a city. Thus, in our case there is no evidence of the 'flypaper' effect described above.

Furthermore, we find that labour productivity ( $Q/L$ ) from the manufacturing sector is negatively related to infrastructure spending. Thus, cities where labour productivity of manufacturing is lower spend more on infrastructure. This evidence turns out to be even stronger if the endogeneity of output  $Q$  and  $GRANTS$  is taken into account in columns 2 and 3. At a first glance this finding suggests some kind of catching-up in infrastructure spending of economically underdeveloped cities. We also conclude from this finding that the argument of reverse causality meaning in our case that more prosperous cities are likely to spend more on infrastructure is empirically not supported.

However, we also find that the coefficient for infrastructure endowment ( $G/L$ ) is positive and significant. This does not support the expectation of catching-up of economically weak cities, because cities which have already a good infrastructure endowment spend more than cities which a poor infrastructure endowment. Hence, at least for our sample no convergence of cities' infrastructure endowments can be expected in the long-run.

In addition, expected productivity effects from infrastructure ( $\beta_{PROD}$ ) appear to matter for local investments. However, the statistical reliability of this result is relatively weak at a 10 percent significance level.

Finally, turning to the political-economy determinants of infrastructure investment we find that the number of manufacturing firms ( $\beta_{NFIRMS}$ ) is not decisive for local infrastructure spending. However, this result should not be interpreted as evidence of non-existence of lobbying efforts of business at the local level.<sup>11</sup>

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<sup>11</sup>Numerous anecdotal evidence on this can be found in newspapers.

The reason is the difficulty to find plausible and observable measures for the lobbying efforts of firms. For instance, it can be argued that the number of manufacturing firms is not an adequate proxy for potential lobbying strength of business, since one dominating big firm might have a stronger influence on policy decisions than many small firms. For this reason, we also tried a dummy variable in our regressions indicating whether one or more headquarters of large stock companies are located in a city. However, this alternative measure of potential lobbying power turned out to be insignificant as well.

On the other hand, we find that the size of majority of the government in the city council is decisive for infrastructure spending. However, since the coefficient  $\beta_{\text{MAJOR}}$  is negative it implies that spending is higher the bigger the majority of the local government. As a consequence, this evidence does not support the hypothesis that local governments spend more on infrastructure if the majority is more unstable. One explanation for the positive coefficient  $\beta_{\text{MAJOR}}$  is that controversial infrastructure projects are likely to be prevented by the opposition in city council if the majority of the government is only small. Furthermore, the larger *MAJORITY* the less likely it is that a city's government is formed on the basis of a party coalition.

A similar finding holds also for the grant equation. The coefficient  $\gamma_{\text{SWINGV}}$  is negative and significant but as before for the support of the swing voter hypothesis a positive coefficient is expected. This result means that cities where the majority of local government is small (i.e. more unstable) receive less grants whereas the swing voter hypothesis predicts that these cities will receive more grants in order to buy the support of swing voters. Thus, the negative coefficient for  $\gamma_{\text{SWINGV}}$  corresponds to the findings for  $\beta_{\text{MAJOR}}$  in the investment equation that cities where government majority is more stable spend more on infrastructure and receive more grants.

Moreover, the estimate for the partisan variable (*PARTISAN*) is significant, which means that the expected level of grants is higher the larger the correspondence of political affiliation between the local city council and the state ('Bundesland') government. At the mean data points, one percentage point increase in political affiliation correspondence between city's and state government gives on

average 1.25 percent more investment grants. This is a considerable amount with regard to the fact that political affiliation in our sample varies from 29.0 to 68.2 percent. This finding is as an indication that self-interests of grant-givers indeed matter for the allocation of grants. Local governments which have a higher 'political capital' (in terms of votes) to sell, are rewarded 'with a larger slice of the cake'. Thus, party affiliation of government is used as a shortcut for ideology, which allows politicians to target grants to those cities with the highest payoff.

Turning to the determinants of grant distribution that are not tactically but benevolently motivated, the negative and significant coefficient  $\gamma_{\text{REDISTRIB}}$  indicates that redistributive concerns are important. Hence, the lower the labour productivity ( $Q/L$ ) of manufacturing in a city (i.e. the more economically underdeveloped it is) the more grants it gets. On the other hand in contrast to what is expected, cities which have already a good endowment with infrastructure (i.e. a high infrastructure intensity  $G/L$ ) get more grants. If infrastructure intensity  $G/L$  is also related to economic development of a city, we would expect a negative sign.

Expected productivity effects ( $\gamma_{\text{PROD}}$ ) of infrastructure investment appear not to matter for the allocation of investment grants. This can be explained by the fact that because investment grant decisions of state governments in Germany are based on consensus with all local level governments, this approach is prone to produce decisions that carefully skirt all areas of conflict. In terms of economic efficiency, the bargaining process will often lead to outcomes which are from a welfare perspective not optimal, so there is no guarantee that the money is being put to its most productive use.

Finally, for the specification in column 3 of Table 4.4 we find that investments have a positive impact on grants. The question arises whether this is a contradiction to the finding for the investment function where no effect of grants on investments is found. Our explanation for this evidence is that local governments on the one hand are not responsive to changes in the amount of grants from higher-tier governments when fixing their level of autonomous investment spending, but higher-tier governments on the other hand are responsive to increased autonomous investment spending. This probably simply reflects the fact that the

relation between local investments and matching grants is relatively fixed, i.e. if autonomous spending increases grants increase as well. The underlying mechanism is that autonomous investments determine the amount of grants, but grants do not determine the amount of autonomous investment spending. In addition, it is also interesting to note that the coefficient  $\gamma_{INV/L}$  with a value of 0.133 is significantly lower than one could expect from the average relation between matching grants and investments in our sample of 0.333. Once other factors that explain the distribution of grants are taken into account, the coefficient of investments gets significantly lower. Finally, cities where mining industry is present spend less on infrastructure investment but receive on average more grants.

## 4.5 Summary and Conclusions

In this study we estimated a system of equations comprising of a production function, an infrastructure investment function and an investment grant function using a panel data set of large German cities. Overall, our empirical results highlight the significance of political factors both for local infrastructure spending and the intergovernmental distribution of grants.

Several key empirical findings emerge from our analysis (i) public capital is a significant input for local production (ii) cities where the city council's majority has the same political affiliation as the state ('Bundesland') government receive more grants (iii) cities where 'marginal voters' are decisive for the outcome of city council elections neither spend more on public infrastructure nor receive more investment grants from higher level governments but the larger majority of government the higher is spending (iv) redistributive concerns of higher-tier governments matter for the allocation of grants, whereas efficiency considerations (i.e. putting the money to its most productive use) appear to be less important.

From a normative perspective our findings support the view that pork barrel politics are indeed important determinants for intergovernmental grant allocation which might give rise to policy outcomes that depart substantially from optimal policies that maximise social welfare.

Finally, many studies on the productivity effects of public capital have treated infrastructure as an exogenous factor of production and neglected the politico-economic factors that shape infrastructure policy. However, the good news from our study is that evidence of endogeneity bias of infrastructure capital estimates in a production function framework as well as evidence of reverse causality running from output to infrastructure investments is weak.

## Essay 5

# Regional Infrastructure Policy and its Impact on Productivity: A Comparison of Germany and France

### *Abstract*<sup>1</sup>

*This essay describes the different institutional frameworks for infrastructure policy in Germany and France. The economic effects of infrastructure are estimated econometrically for German and French regions. We find evidence that regional road infrastructure has a significant impact on regional output. Moreover, we find evidence that for Germany the priority of promoting equal living conditions throughout the regions is an important determinant of regional infrastructure policy.*

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<sup>1</sup>A previous version of this essay is forthcoming in *Applied Economics Quarterly* (2001), vol. 47, 274-303.

## 5.1 Introduction

At the end of the 1960's, economic research began to deal increasingly with the importance of infrastructure for economic development. Early studies e.g. Frey (1970), Jochimsen (1966), Jochimsen and Simonis (1970) or Simonis (1977) dealt especially with the theoretical aspects of infrastructure provision and infrastructure's conceptual basis.

Since the end of the 1980's, there has been greater interest also in empirical infrastructure research. By using production function approaches, the direct and indirect effect of improvements in infrastructure on private productivity have been estimated. The studies by Aschauer (1988; 1989a; 1989b) have not only raised the attention of scientists but have also had an effect on economic policy. As a result, spending on public infrastructure in the US increased considerably during President Clinton's first term of office (Gramlich, 1994).

For the period 1949-1985, Aschauer (1989a) reports a significant elasticity of output with respect to public non-military capital between 0.38 and 0.56 for the US using aggregated time series data. Thus the estimated marginal productivity of public capital in this study considerably exceeded that of private capital. This finding implies that returns resulting from public investment were higher than those arising from private investment projects.

Turning to the hypothetical effects of infrastructure, Aschauer (1995), 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 labour and capital inputs more or less productive. The indirect effect arises because an increase in public capital stock will affect the marginal products of labour and private capital, which in turn influence the chosen quantities of private inputs.

Furthermore, Aschauer advances the theory that the up to 60 percent of the decrease of productivity growth in the USA during the 1970's and 1980's can be attributed to the cut-back in public infrastructure investment during this period.

However, some economists have voiced doubts about the plausibility of the results of Aschauer's studies (Aaron, 1990; Gramlich, 1994; Hulten and

Schwab, 1991; Jorgenson, 1991; Tatom, 1991a). One criticism is the high degree of aggregation of the data used by Aschauer. Therefore, more recent research works have examined the effects of infrastructure on regionally more disaggregated levels.

Yet the results of these studies are not unequivocal. Whereas for instance Munnell (1990b; 1992; 1993) confirm the hypothesis formulated by Aschauer, the studies by Baltagi and Pinnoi (1995), Garcia-Milà et al. (1996), Holtz-Eakin (1994), or Holtz-Eakin and Schwartz (1995) find no evidence of a significant influence of infrastructure on productivity for the US.

Authors such as Baltagi and Pinnoi (1995) or Holtz-Eakin (1994) point out that regional effects should be taken into consideration in econometric examinations on a disaggregated level. Differences between the regions regarding the geographical location (centre versus periphery), climate or factor endowments can be captured by econometric techniques using fixed or random cross-sectional effects. Our study takes this criticism into account in that the econometric methods applied here are able to estimate such regional-specific influences.

In the literature, when referring to the differentiation of individual areas of infrastructure,<sup>2</sup> a distinction is made between household-related infrastructure and business / or business-related infrastructure. Household infrastructure covers healthcare and educational, leisure and cultural institutions. Road infrastructure can be placed under the heading of business related infrastructure. Not only transport belongs under this heading, also energy and water provision and telecommunications infrastructure are business-related (Frey, 1978). This study focuses on road infrastructure because comparable data for both the French and German regions for this section of infrastructure are available.

The purpose of this essay is threefold. First, we survey the institutional framework of infrastructure policy as an instrument of regional policy in Germany and France. Second, we study the effects of infrastructure on private productivity.

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<sup>2</sup>Throughout the essay we use the terms 'infrastructure' and 'public capital' interchangeable. Note, however, that in a more rigorous fashion public capital refers to infrastructure services that are solely publicly financed, whereas the more general term 'infrastructure' applies also to privately financed services.

Third, we investigate the determinants of regional infrastructure investment allocation.

One contribution of our study to the existing empirical infrastructure literature is that it simultaneously refers to both French and German regions. The main advantage of pooling the data for Germany and France is that the database is expanded and therefore we are able to obtain more reliable estimates of the parameters of the model. A further contribution is that our study implements some methodological improvements in comparison with previous investigations in that the estimations of the effects of infrastructure on French regions are carried out by taking regional-specific effects into account. Another important aspect of our study, which has seldom been dealt with in the existing literature, is that it highlights the different institutional frameworks under which infrastructure policy is carried out in Germany and France. Finally, our study discusses and implements a new framework for studying empirically the determinants of regional infrastructure investment allocation.

The further development of this essay is as follows: In the next subsection, we provide an overview on related studies. In the second section we compare regional policies in Germany and France. In addition, some theoretical aspects of the rationale of regional policies are collated and discussed. In the third section we present the results of the empirical analysis of the effects of infrastructure on productivity. Furthermore, an investigation on the empirical determinants of the allocation of road infrastructure is carried out. In the fourth and final section, the results of the study are summarised and discussed.

### **5.1.1 Related literature**

A number of investigations into the effects of infrastructure on private production in Germany have already been carried out (Conrad and Seitz, 1992; Conrad and Seitz, 1994; Erber, 1995; Hofmann, 1996; Licht and Seitz, 1994; Schlag, 1997; Seitz, 1993; Seitz, 1994; Seitz, 1995; Stephan, 1997). However, as far as we know only two studies have been undertaken with regard to the impact of infrastructure on regional development in France (Fritsch, 1995; Prud'Homme, 1996).

The majority of the previous studies with regards to Germany apply a cost function approach, only a few studies are based on a production function approach. The regional and sectoral levels of reference of the respective studies are sometimes very different, making a comparison of results very difficult. Furthermore, also different measures and definitions of infrastructure and public capital respectively are used in these studies.

For example Licht and Seitz (1994) examine the economic importance of infrastructure at the level of the 11 West German federal states. The method of investigation is based on a cost function approach. The estimated cost elasticity for public capital is significant and ranges from -0.01 to -0.36. Also, the studies by Conrad and Seitz (1992), Seitz (1993), and Conrad and Seitz (1994) confirm the evidence of cost effects arising from infrastructure at the aggregated level of West Germany. Erber's study (1995), which performs the analysis for both Germany and the US finds only for 4 of 26 branches an influence of the public capital stock on costs.

In Seitz (1995), 85 self-administrated cities in Germany serve as the regional level of reference. Significant effects on the cost of private production are found in this infrastructure study. Hofmann (1996) analyses the effects of public infrastructure on productivity applying various econometric methods for Hamburg. However, no plausible results are obtained so that it is not possible to make conclusions about the importance of public infrastructure for Hamburg.

In Stephan (1997), the influence of road infrastructure on production in the manufacturing industry in the 11 West German federal states for the period 1970-1995 is examined based on a production approach. Using this method, significant effects of infrastructure are found for almost all specifications.

Schlag (1997) studies the causality link between the public infrastructure capital and the output of the total business sector in Germany at an aggregate level. Cointegration analysis is applied and error correction models for time series and panel data results from Granger causality tests are presented. The results indicate bi-directional causality (feedback) between the public infrastructure capital and the output.

Fritsch (1995) estimates a significant effect of infrastructure on productivity in 21 French regions. Also Prud'Homme (1996), whose study is likewise based on the 21 French regions, provides evidence of a significant effect. However, regional specific effects are not specified in the econometric estimations in both of these studies.

In sum, so far no clear-cut evidence of the effects of infrastructure on private productivity emerges from empirical studies carried out for Germany or France. Whereas a number of studies find significant effects of infrastructure others do not. However, it is worth pointing out that the results of the different approaches for testing the significance of infrastructure are hardly comparable due to the different levels of regional or sectoral reference as well as due to the different definitions of infrastructure capital used in the studies. In the following section, we highlight and discuss the differences between infrastructure policies in Germany and France. Before that, we provide evidence on the development of regional income disparities in Germany and France.

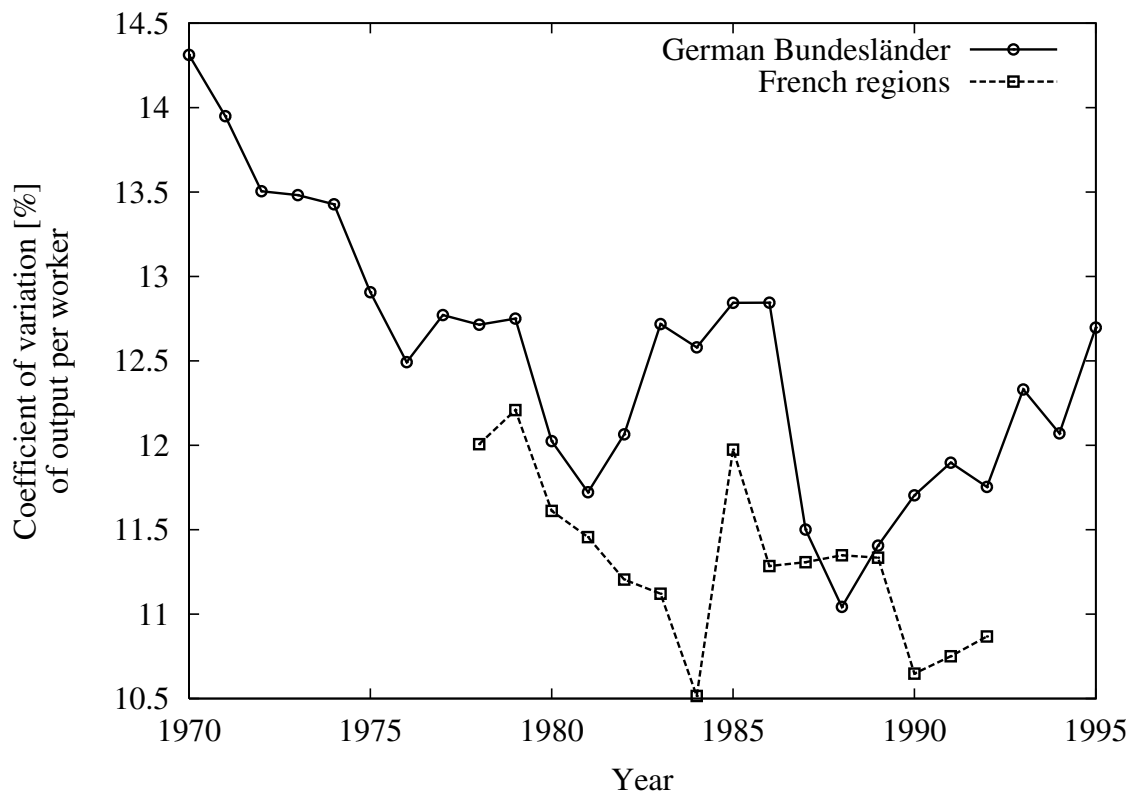
## **5.2 A comparison of infrastructure policies in Germany and France**

Infrastructure is often used as an instrument for regional economic policies in order to reduce regional disparities in income. Proponents of active regional economic policies maintain that without these state support, disparities in income will increase between the regions.

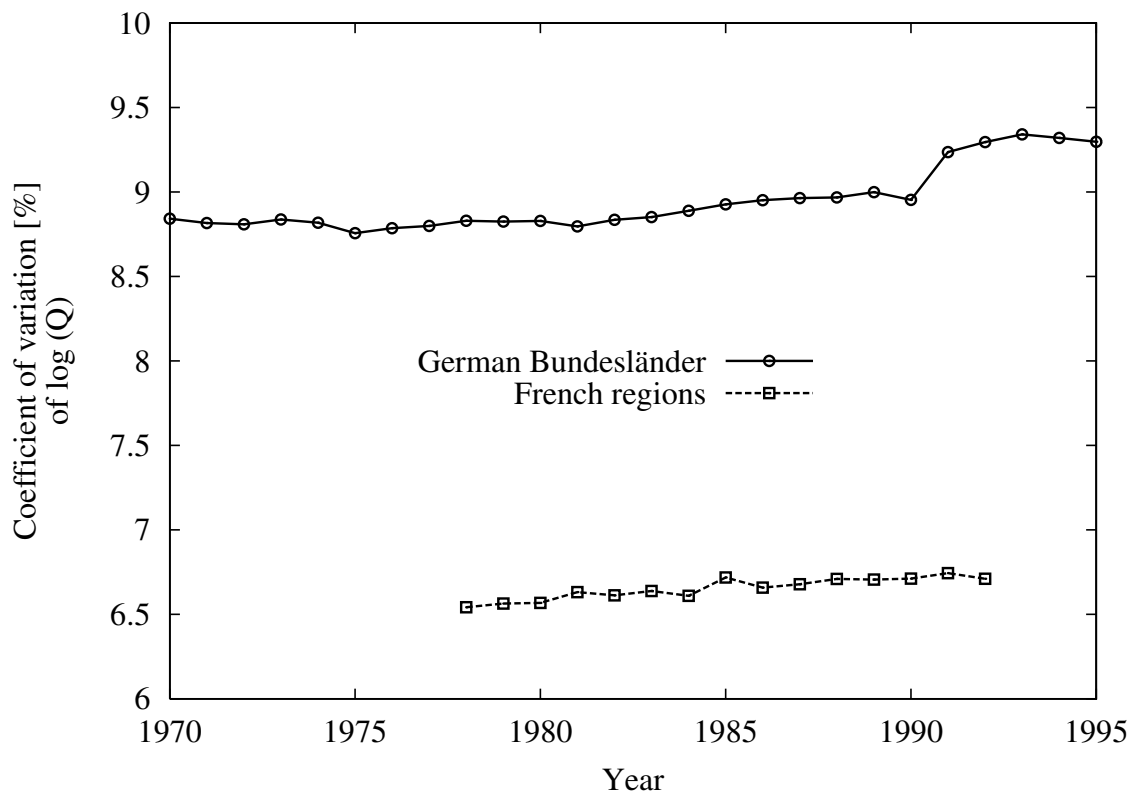
Figure 5.1 shows the development of the regional differences in productivity in Germany and in France. Regional productivity is measured as gross value added per worker. For each year, we have calculated the coefficient of variation for this variable. Note that the coefficient of variation is a unitless measure of relative variability. It is defined as the ratio of the standard deviation to the mean expressed as a percentage. If a convergence of regional productivities occurs, the coefficient of variation will decrease over the course of time.

Figure 1 reveals that although regional differences in productivity decreased during the period 1970-81, they increased slightly from 1982-86, and following a

**Figure 5.1:** Regional Productivity Differences for 11 West German Bundesländer 1970-95, 21 French regions 1978-92



*Figure 5.2: Regional Differences of Value Added for 11 West German Bundesländer 1970-95, 21 French regions 1978-92*



further fall during the period 1987-88, have increased again since 1989. Therefore, at least for the period 1976-95, it is not possible to determine a convergence of regional productivity in Germany. Note, that this result for the 11 West German states, which emerges from a rather descriptive analysis, is in line with the findings of studies using more sophisticated analytical tools e.g. unit root tests for panel data (Bohl, 1998; Funke and Strulik, 1999).

A similar picture is found for France. Although regional differences decreased during the period 1979-84 and 1986-90, increases can be seen in the years 1984-85 as well as since 1991. This result that regional convergence can neither be found for France nor for Germany is also confirmed by Figure 5.2. Here the coefficients of variation are calculated based on the logarithms of regional value added. By taking logarithms of the levels of output the regional absolute differences are transformed into relative percentage differences. Again, we expect a decrease in the coefficient of variation if a convergence in regional outputs occurs.

In contrast to Figure 5.1, where in the coefficient of variation the effects of the development of regional labour force are also included, only the development of the relative regional output differences is captured in Figure 5.2 independent of the development in the labour force (for instance due to labour migration between regions). Apparently, in the period under investigation neither in France nor in Germany convergence of regional income is observed.

As a result of this section we state that a decrease in regional disparity in income can neither be observed in Germany nor in France during the period 1970-1995 and 1978-1991 for France respectively. However, when no regional convergence can be found in the two countries the question arises whether regional policies were efficient or whether the disparities would have been even more marked without active regional policies. The following section describes some important institutional differences in infrastructure and regional policy in Germany and France.

### 5.2.1 The different frameworks of infrastructure policy in Germany and France

The geographical differences (density of population 104 inhabitants per square km in France and 223 inhabitants per square km in Germany) highlight the differing conditions for the forming of infrastructure policy in Germany and in France. In Germany in 1992 there were a total of 11,000 km of motorway ('Bundesautobahnen'), 42,000 km A roads ('Bundesfernstraßen'), 170,140 km B roads (85,200 km 'Landesstraßen' and '84,940 Kreisstraßen') and 413,000 km smaller roads and streets ('Gemeindestaßen') (Source: German Ministry of Transport, Building and Housing, 1995).

In France there are similarly 5 categories of road: 'autoroutes' (motorways), 'routes nationales' (A roads), 'routes départementales' and 'voies communales' (B roads) and 'chemins ruraux' (smaller roads and streets).

In 1992 the length of the French motorway network was 9,081 km and is therefore comparable with the extent of the German system. The length of the 'routes nationales' consisting of 27,500 km is around half the length of the German A roads. On the other hand, there are 365,600 km of B roads and in addition a network of smaller roads and streets ('voies communales') which, with its 579,000 km is clearly longer than in Germany. This can be attributed to the larger geographical area of France (Centre National de Documentation Pédagogique, 1998).

However, not only the geographical differences but also history and politics have contributed to forming differing infrastructure policies. French infrastructure policy is on the one hand marked by strong regional policy considerations and on the other by the emphasis given to individual large infrastructure projects (Kistenmacher, Marcou and Clev, 1994).

Due to its tradition of political centralism in the first years of the post second world war, a markedly interventionist regional policy was approved. For a long time Paris and the greater Paris area (Ile-de-France) stood in the foreground of regional development. However, in the following years a policy of de-concentration and the development of industrial centres outside of the area around Paris was increasingly pursued. The central regional planing institution

DATAR (=Délégation à l'aménagement du territoire et à l'action régionale), set up in 1963 for this purpose, has wide-reaching decision making powers. In order to improve the carrying out of proactive centrally controlled regional policy DATAR was put directly under the control of the Prime Minister.

Since the beginning of the 1980's it has also been possible to observe an increasing tendency towards the decentralisation of planning in France. The hierarchical control of the 1960's and 1970's has been superseded in the form of a contractual agreement between the central state and the regions (Neumann and Uterwedde, 1994).

Furthermore, the private building and management of motorways plays an important role. Private or non-profit making firms manage the majority of the motorway network (6,490 km of a total of 9,081 km) and charge the road users tolls. Through concessions, the public authorities grant the private firms specific rights. Not only the maintenance and management of the motorways are financed by tolls, but the building of further motorways is also financed this way (Ministère de L'Equipement, des Transports et du Logement, 1998).

The 'Direction des Routes' is responsible for the financing and planning of the 'routes nationales' and the state motorways. This is subordinate to the ministry responsible (Ministère de l'Equipement, des Transports et du Logement). On the one hand, the 'Direction des Routes' sets the targets for road and motorway construction according to a certain scheme (Schéma directeur routier national). On the other hand the 'Direction des Routes' enacts laws which determine building, maintenance and management. Additionally, the determination of the necessary means for finance is undertaken by the 'Direction des Routes'. The 22 regions are therefore represented by the 'Ministère de l'Equipement, des Transports et du Logement' (DRE) and are subordinate to the 'Direction des Routes'.

In Germany, in contrast to France, several regional metropolitan areas of equal rank have developed over the years (polycentric development). The principle of federalism played an important role in the forming of regional policy in Germany during the post-war period. The constitutional law promoting the convergence of living conditions throughout the regions above all represents an important target of regional policy.

As a result of the federal structure, the states have legislative competence for B roads and smaller roads and streets. On the other hand the federal government is the owner of and responsible for motorways and A roads; these are built and administered upon commission from the federal government by the federal states (German Ministry of Transport, Building and Housing, 1995).

In the immediate post-war period, the intention was that the federal government would only play a role in setting out the conditions in the planning of regional policies. With the passing of the regional planning law ('Raumordnungsgesetz') in 1965, the importance of the federal government within the federal system was strengthened. Through the federal transport infrastructure plan ('Bundesverkehrswegeplanung'), a transport policy program was introduced which was supposed to co-ordinate all federal transport (federal roads, federal rail and federal waterways). Federal transport infrastructure plans were drawn up by the Federal Cabinet and regularly reformulated. Similar long term transport planning did not exist in France at the time.

To sum up, we can determine that infrastructure and regional policy in Germany was accompanied by the aim of having similar living standards throughout the regions, whereas in France during the last few decades, importance was placed, above all, on decentralisation and the relief of the concentrated area around Paris. The differences which still exist between infrastructure policy in Germany and France will, in the future, become less prominent due to the European integration (Kistenmacher, Marcou and Clev, 1996). In the next section some theoretical considerations regarding the efficiency of regional infrastructure policies are collated and discussed.

### **5.2.2 Can regional policies work in principle? Some preliminary considerations**

The following reflections should serve as a compilation of several arguments based on economic theory for or against an active role of regional policy. Regional disparities in income often prompt governments to make efforts in order to achieve a more evenly balanced regional economic development. Tradition-

ally, public infrastructure policy has been an instrument for regional economic support. The intention of this policy is to minimise competitive disadvantages in the regions, and promote private investment through the improvement of regional infrastructure.

According to standard neoclassical growth theory built on the assumption of decreasing returns to reproducible factors on the other hand, income disparities arising from differences in regional capital/labour ratios will diminish over time: both trade and factor flows tend to equalise factor prices. A convergence of income in the regions would therefore also take place without active regional policy. Similarly, also the so-called 'rule of thumb' of a 2 percent convergence rate according to empirical studies based on neoclassical growth theory leaves no scope for an active role of regional policy (for an overview, see Barro and Sala-I-Martin, 1995).

Apparently, the presence of externalities (spill-over effects) can give objective reasons for regional economic support. For instance, if negative 'crowding' externalities exist in economically better developed regions, these can be mitigated by regional policies which provide infrastructure to economically less developed regions, so that labour migration from the less to the better developed regions is prevented and externalities thus reduced.

However, according to Homburg (1993) even in the absence of externalities active regional distribution policies can be justified. To analyse the consequences of regional distribution policies, Homburg's model assumes a neoclassical production function  $Q_i = f(G_i, K_i, \bar{L}_i)$ ,  $i = 1 \dots N$ , linear homogenous of degree one, where  $Q_i$  is region  $i$ 's output,  $G_i$  is the *stock* of public capital in region  $i$ ,  $K_i$  is the *stock* of private capital in region  $i$ , and  $\bar{L}_i$  is an immobile factor of production in region  $i$ , e.g. land endowment, geographical characteristics etc.

The steady state of the model is described by the two conditions: (i)  $\tau Q_i = \delta G_i$ , where  $\tau$  denotes the tax rate,  $\delta$  denotes the depreciation rate of public capital, and (ii)  $(1 - \tau)F_K^i = r + \delta$ , where  $F_K^i$  denotes the marginal productivity of private capital for region  $i$  and  $r$  the exogenous rate of interest. Condition (i) states that in steady state taxes are only used to finance replacement of  $G_i$ , and condition (ii)

states that the marginal after-tax productivity of private capital equals the rental price of private capital  $r + \delta$ .

Homburg (1993) shows that a *spatial efficient* allocation of infrastructure can be regarded as a maximisation of the joint total output  $\sum Q_i$  given the sum of stocks of private capital  $\sum K_i$  and given the sum of endowments with infrastructure  $\sum G_i$ . The spatial efficient allocation of infrastructure solves the following maximisation problem

$$\max! \sum Q_i, \text{ given } \sum K_i = \bar{K} \text{ and } \sum G_i = \bar{G}. \quad (5.1)$$

It can be shown that the condition for the spatial efficient allocation applies exactly when the marginal productivities of infrastructure  $F_G^i$  and private capital  $F_K^i$  in all regions are given as

$$F_G^i = \mu \text{ and } F_K^i = \theta \text{ for all } i.$$

This means that the marginal productivities of private and public capital are equal for all regions. If public and private capital is also homogeneous in a neo-classical sense, then furthermore  $\mu = \theta$  should also apply. From this result we can state that under neoclassical assumptions it would be optimal to allocate infrastructure investment in such a way across regions that the marginal productivity of infrastructure is equal in all regions.

An important result of this model is that the spatial efficiency criterion is valid whether the total endowment with infrastructure is optimal or sub-optimal. If the assumptions of the model apply, then an efficient spatial allocation of infrastructure is observed in the steady-state equilibrium even without governmental subsidies. However, as Homburg shows in his further analysis, the adjustment processes to this efficient steady state equilibrium is characterised by an inefficient spatial allocation. This implies that the adjustment path can be improved upon by means of using intergovernmental grants.

As the main result therefore we can state that if the initial allocation of infrastructure stocks of infrastructure across the regions is unbalanced, the joint national product can be increased by a regional infrastructure policy. The target of such a policy should be to balance the ratio of output  $Q_i$  to infrastructure stock  $G_i$

in all regions, i.e. to equalise the marginal productivities of infrastructure across regions.

However, it should be noted that studies such as Martin (1998; 1999a) or Ottaviano and Thisse (1999) arrive at different results regarding the possible effects of regional infrastructure policy. These models of the 'New Economic Geography' predict that the consequences of a policy which targets at the achievement of a balanced spatial allocation of economic activity can result even in greater regional disparities. The mechanism behind this result is that the reduction of transport costs, for example by means of improved transport infrastructure, can have negative effects on the economic development of poorer regions. This will happen if companies from the poorer regions move to take advantage of the agglomeration and scale economics in centrally located regions while at the same time, however, they can maintain their sales outlets in the poorer regions due to the reduced transport costs.

In the next section we analyse empirically whether investment in road infrastructure has a positive effect on economic development. For this purpose, an econometric analysis is carried out based on a production function with panel data for the French regions and the German federal states.

## 5.3 Empirical analysis

The first part of the empirical analysis deals with the productivity effects of infrastructure. In the second part, we investigate the determinants of the regional allocation of infrastructure investment.

### 5.3.1 Productivity effects of regional road infrastructure in Germany and France

The central hypothesis to be examined empirically is that infrastructure increases private output or reduces respectively the costs for a given unit of output. From a theoretical point of view, this can be the case when infrastructure either directly

exerts a positive effect on private factor productivities or indirectly exerts a positive influence on private factor productivities which in turn increases the demand for private factor inputs (Aschauer, 1995).

In the following section, the effects of road infrastructure on productivity are examined using two different approaches

1. Cobb-Douglas production function
2. Transcendental logarithmic (translog) production function according to Christensen et al. (1971; 1973).

For the estimations, we employ econometric methods from panel data analysis by specifying fixed cross-sectional effects (Baltagi, 1995; Hsiao, 1986). For the first approach, our empirical model is based on the production function for region  $i$ ,  $i = 1 \dots N$ , in year  $t$ ,  $t = 1 \dots T$ ,

$$Q_{it} = A_{it}(t)F(t, X_{1it}, \dots, X_{Mit}), \quad (5.2)$$

where  $Q_{it}$  describes output,  $A_{it}(t)$  technical efficiency (or the Hicks-neutral technical progress) and  $X_{1it}, \dots, X_{Mit}$  describe the  $M$  factors of production. Assuming a Cobb-Douglas production technology and with factor inputs labour  $L_{it}$ , private capital  $K_{it}$  and road infrastructure  $G_{it}$  and after taking logarithms and dividing by  $L_{it}$ , we obtain the following empirical model is obtained which forms the basis of our empirical assessment

$$\ln q_{it} = \ln A_{0i} + \alpha_t t + \alpha_k \ln k_{it} + \alpha_g \ln g_{it} + \tilde{\alpha}_L \ln L_{it} + u_{it}, \quad (5.3)$$

$$u_{it} = \rho u_{i,t-1} + \epsilon_{it} - \gamma \epsilon_{i,t-1},$$

$$\text{and } \tilde{\alpha}_L = \alpha_k + \alpha_g + \alpha_L - 1,$$

where  $u_{it}$  follows an autoregressive moving average process ARMA(1,1) and  $\epsilon_{it}$  denotes normal i.i.d. distributed random innovations. In addition, we assume that  $A_{it}(t) = A_{0i} \exp(\alpha_t t)$ . Note that variables in lower-case letters in (5.3) are defined as  $x = X/L$ . The parameters  $\alpha_k, \alpha_g, \alpha_L$  describe the elasticity of the output  $Q_{it}$  with respect to inputs  $K_{it}$ ,  $G_{it}$  and  $L_{it}$ .

The advantage of this specification for the production function is that by dividing (5.2) by  $L_{it}$  the problem of heteroscedasticity for the empirical estimation

is reduced. Notice also that no 'a priori' restrictions are placed on (5.3) regarding returns to scale. If the parameter  $\tilde{\alpha}_L$  is significantly different from zero, then the null hypothesis of constant returns to scale is rejected.

Table 5.1 contains the results for the Cobb-Douglas production function which has been estimated by using the procedure PROC MIXED in SAS V8. A detailed description of the data used in this analysis can be found in the Appendix.

Maximum likelihood estimation (MLE) is applied to all specifications based on a total of 596 observations (281 for Germany and 315 for France). The main benefit of pooling the data for France and Germany is that the analysis can be based on a larger data set and therefore more reliable estimates of the parameters of the model are obtained.

In column (1), the model is estimated assuming heterogenous parameters for German and French regions. Note, that in contrast to Ordinary Least Squares (OLS) estimation the parameters in column (1) are different due to the specified covariance structure from parameters that would be obtained by running two separate regressions for German and French regions.

In column (2), we assume parameter homogeneity for German and French regions except for the covariance parameters  $\rho$  and  $\gamma$ . That means that the specification in column (2) can be deduced from column (1) by imposing restrictions with respect to parameter homogeneity on the specification of column (1). Furthermore, in column (3), except for labour and the covariance parameters  $\rho$  and  $\gamma$ , parameter homogeneity across German and French regions is assumed.

Note, that fixed cross-section effects are added to all specifications (1)-(3) of Table 5.1. The results of likelihood ratio (LR) tests not reported here imply that these fixed cross-section effects are significantly different from zero. Also, because a linear time trend  $t$  is included in eq. (5.3), it is not possible to estimate additional time effects due to the resulting singularity.

The ARMA(1,1) parameters  $\rho$  and  $\gamma$  for both Germany and France are significantly different from zero for all specifications. The model selection criteria AIC and SBC, which we describe below indicate that these specifications are preferred compared with AR(1) alternatives not reported here. The displayed 'null model likelihood ratio test' checks the model without covariance parameters against

**Table 5.1: Regression Results for the Productivity Effects of Road Infrastructure**

Maximum-Likelihood Estimations (MLE)						
	(1)		(2)		(3)	
Intercept	<i>fixed effects</i> ***		<i>fixed effects</i> ***		<i>fixed effects</i> ***	
$t$	0.0114 <sup>1</sup>	(6.88)***	0.0151	(12.93)***	0.0133	(10.67)***
	0.0141 <sup>2</sup>	(7.74)***				
$\ln k$	0.2310 <sup>1</sup>	(3.03)***	0.1162	(3.55)***	0.1457	(4.28)***
	0.1254 <sup>2</sup>	(3.36)***				
$\ln g$	0.0828 <sup>1</sup>	(1.32)	0.0837	(2.12)**	0.1120	(2.79)***
	0.1282 <sup>2</sup>	(2.13)**				
$\ln L$	0.3455 <sup>1</sup>	(3.56)***	-0.0588	(-0.94)	0.2784 <sup>1</sup>	(3.18)***
	-0.2932 <sup>2</sup>	(-3.39)***			-0.3337 <sup>2</sup>	(-4.14)***
AR(1) $\rho$	0.7857 <sup>1</sup>	(16.79)***	0.8370 <sup>1</sup>	(20.96)***	0.8172 <sup>1</sup>	(18.99)***
	0.5888 <sup>2</sup>	(7.08)***	0.5257 <sup>2</sup>	(6.15)***	0.6303 <sup>2</sup>	(8.29)***
MA(1) $\gamma$	0.8477 <sup>1</sup>	(29.35)***	0.8823 <sup>1</sup>	(34.55)***	0.8678 <sup>1</sup>	(31.59)***
	0.6431 <sup>2</sup>	(12.89)***	0.5977 <sup>2</sup>	(12.08)***	0.6686 <sup>2</sup>	(13.69)***
Null Model LR Test $\chi^2$	447.9***		504.1***		506.4***	
Log- Likelihood	1531.2		1514.8		1528.7	
AIC	1485.2		1472.8		1485.7	
SBC	1451.5		1442.0		1454.2	
Observ.	281 <sup>1</sup>		281 <sup>1</sup>		281 <sup>1</sup>	
	315 <sup>2</sup>		315 <sup>2</sup>		315 <sup>2</sup>	

Parameters for <sup>1</sup>German regions, <sup>2</sup>French regions, otherwise for both regions (homogenous).

Approx. t-statistics are given in parentheses. Significance levels: \*10 %, \*\*5 %, \*\*\*1 %.

Dependent variable is the natural logarithm of (regional value added / labour).

the alternative of the specified ARMA(1,1) covariance structure. Thus, the 'null model' without covariance structure parameters is rejected for each of the three specifications (1)-(3).

In column (1) of Table 5.1, the estimates of the parameters of the input factors  $k$ ,  $g$  and  $L$  are statistically significant for the French regions. For the German 'Bundesländer' (federal states), estimates of the parameters of input factors  $k$  and  $L$  are statistically significant, however not for  $g$ . Note, that constant returns to scale are rejected both for German as well as for French regions.

In column (2), a decrease in the value of the log likelihood from 1531.2 in column (1) to 1514.8 in (2) can be observed. Indeed, the LR test for the restriction of parameter homogeneity,  $-2[1514.8 - 1531.2] = 32.8 \sim \chi^2(4)$ , is highly significant, thus the null hypothesis of parameter homogeneity across German and French regions is rejected. This is also reflected in the decrease of the Akaike Information Criterion (AIC) (Akaike, 1974), which has been computed as  $AIC = l(\hat{\theta}) - d$ , where  $l(\hat{\theta})$  is the maximised log likelihood and  $d$  is the effective number of parameters (fixed effects and covariance parameters). It can be used to compare different models; the model with the largest AIC is deemed best. Similarly, Schwarz's Bayesian Criterion (SBC) (Schwarz, 1978) has been computed as  $SBC = l(\hat{\theta}) - \frac{1}{2}d \log N$ , where  $N$  equals the number of valid observations for maximum likelihood estimation. Again, models with larger SBC are preferred, but note also that SBC penalises models with a greater number of parameters more than AIC does, that means it will lean toward a simpler model. Therefore, the specification of column (1) is preferred compared to (2) by both criteria.

However, the rejection of specification (2) is mainly driven by the heterogeneity of the parameter for labour between German and French regions. Therefore, in column (3) we allow for this heterogeneity, whereas the other parameters (except the covariance parameters) are restricted to be equal across German and French regions. In contrast to column (2), this specification is not rejected by the LR test,  $-2[1528.7 - 1531.2] = 5 \sim \chi^2(3)$ . Furthermore, also the AIC and SBC criteria are higher than for column (1). Thus, we conclude that column (3) contains the results of the preferred specification of the empirical model. We find that the time trend  $t$  with a value of 0.0133 is significant, and private and public capital is

significant with values of 0.1457 and 0.1120 respectively. In sum, the main finding of the performed analysis is that road infrastructure is significant for private production at the regional level.

However, the Cobb-Douglas production function approach restricts the elasticities of input substitution to equal one. In order to overcome this limitation, our second approach is based on a translog production function

$$\begin{aligned} \ln Q_{it} &= \ln A_{0it} + \alpha_t t + \alpha_k \ln K_{it} + \alpha_g \ln G_{it} + \alpha_L \ln L_{it} \\ &+ \alpha_{kg} \ln K_{it} \ln G_{it} + \alpha_{kl} \ln K_{it} \ln L_{it} + \alpha_{gl} \ln G_{it} \ln L_{it} \\ &0.5 \left[ \alpha_{kk} \ln^2 K_{it} + \alpha_{gg} \ln^2 G_{it} + \alpha_{ll} \ln^2 L_{it} \right] + \epsilon_{it}, \end{aligned} \quad (5.4)$$

$$u_{it} = \rho u_{i,t-1} + \epsilon_{it} - \gamma \epsilon_{i,t-1}.$$

Again, we assume that  $u_{it}$  follows an autoregressive moving average process ARMA(1,1). The effect from public input  $G$  on private factor productivities, i.e.  $\partial^2 Q / \partial K \partial G$  and  $\partial^2 Q / \partial L \partial G$ , can be derived from the estimates of equation (5.4) as

$$\hat{\alpha}_{kg} = \frac{\partial^2 \ln Q}{\partial \ln Q \partial \ln K}, \quad \text{and} \quad \hat{\alpha}_{gl} = \frac{\partial^2 \ln Q}{\partial \ln G \partial \ln L}, \quad (5.5)$$

from which  $\partial^2 Q / \partial K \partial G$  and  $\partial^2 Q / \partial L \partial G$  can be computed as

$$\frac{\partial^2 Q}{\partial K \partial G} = \hat{\alpha}_{kg} \frac{Q}{KG}, \quad \text{and} \quad \frac{\partial^2 Q}{\partial G \partial L} = \hat{\alpha}_{gl} \frac{Q}{GL}. \quad (5.6)$$

Since the ratios  $Q/KG$  and  $Q/GL$  are positive, we can infer from the signs of  $\hat{\alpha}_{kg}$  and  $\hat{\alpha}_{gl}$  whether the effect of  $G$  on private factor productivities is positive or negative respectively.

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

$$\sum_p^m \sum_l^m \hat{\alpha}_{pl} = 0, \quad (5.7)$$

where  $p, l \in \{K, G, L\}$ ,  $m = 3$ . In addition, linear homogeneity requires the above condition plus that the sum of the linear terms equals one (Chambers, 1988)

$$\sum_p^m \hat{\alpha}_p = 1. \quad (5.8)$$

We obtain the following results for the translog production function approach

$\widehat{\ln Q}_{it}$  = fixed effects\*\*\*

+0.014 <i>t</i> (11.70)***	−0.445 ln <i>K</i> <sub><i>it</i></sub> (−1.16)	+0.740 ln <i>G</i> <sub><i>it</i></sub> (2.20)***	−0.659 ln <i>L</i> <sub><i>it</i></sub> (−1.53)
−0.106 ln <i>K</i> <sub><i>it</i></sub> ln <i>G</i> <sub><i>it</i></sub> (−1.19)	−0.362 ln <i>K</i> <sub><i>it</i></sub> ln <i>L</i> <sub><i>it</i></sub> (5.29)***	−0.143 ln <i>G</i> <sub><i>it</i></sub> ln <i>L</i> <sub><i>it</i></sub> (−2.31)***	
+0.096 ln <sup>2</sup> <i>K</i> <sub><i>it</i></sub> (−0.83)	+0.176 ln <sup>2</sup> <i>G</i> <sub><i>it</i></sub> (2.00)***	−0.198 ln <sup>2</sup> <i>L</i> <sub><i>it</i></sub> (−2.16)***	

*N* : 596 Log-Likelihood: 1536.9 AIC: 1488.9 SBC: 1453.8

The value of the LR statistic,  $-2[1514.8 - 1536.9] = 44.2 \sim \chi^2(6)$ , which is highly significant, shows that due to the addition of cross and quadratic terms the translog model is preferred compared with the Cobb-Douglas specification in column (2) of table 5.1 (however, it is not preferred according to the SBC criterion). Again, we find that infrastructure  $G_{it}$  is significant. Moreover, with respect to marginal productivities inputs  $G_{it}$  and  $L_{it}$  are substitutes ( $\alpha_{gl} = -0.143$ ), whereas  $G_{it}$  and  $K_{it}$  appear not to affect each other ( $\alpha_{kg}$  is insignificant).

However, it should be mentioned that the results of the translog specification should be interpreted with some caution due to the high correlation of the single with the quadratic and the cross terms. The correlations not reported here between the single and the cross and quadratic terms are greater than 0.8 for most of the cross and quadratic terms. Due to this high degree of multicollinearity between the explanatory variables, imprecise or even estimates with implausible signs can result (Judge et al., 1985, chap. 22). Finally, note that by applying LR tests, linear homogeneity is rejected for the estimated translog production function.

It can be summarised that our empirical analysis finds evidence that regional road infrastructure has a significant impact on regional output. The specification with heterogenous parameters between German and French regions appears to indicate that the effect of road infrastructure on productivity is significant only for France. However, by pooling the data for Germany and France we obtain more efficient and reliable parameter estimates. Thus, the model with heterogenous parameters is rejected and the specification where only for labour heterogeneity

of parameters is assumed is deemed best. In the following section, the determinants of the regional allocation of road infrastructure in Germany and France are analysed.

### 5.3.2 Empirical determinants of the regional allocation of road infrastructure investment

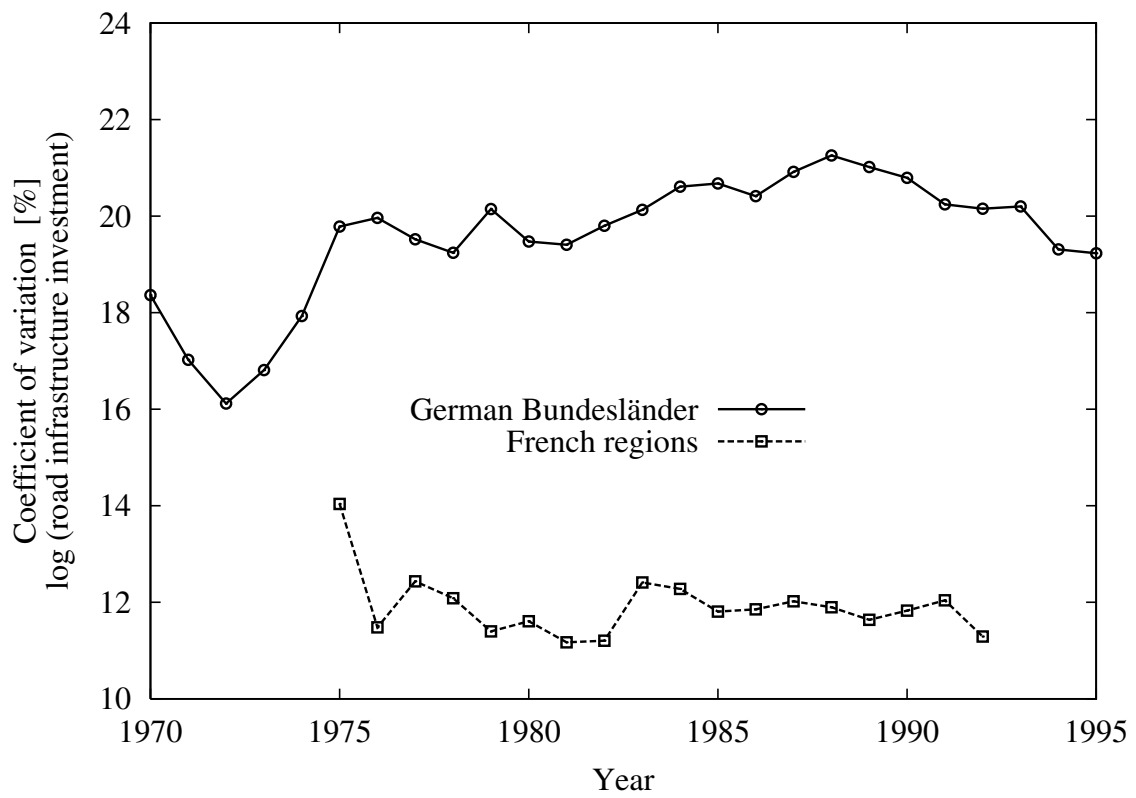
Figure 5.3 shows the variation in the allocation of regional road infrastructure investment in Germany and in France. Again, the development of the variation is expressed in terms of the coefficient of variation calculated for each year. For Germany, the variation increased relatively constant in the period 1972-89, but decreased after 1990. This is probably a result of the German reunification after which priority was given to improvements of infrastructure in East German regions, whereas differences of infrastructure investment in the West German regions are levelled out due to budget constraints.

With regards to France, neither a constant increase nor decrease can be observed. This means that the variation of the investment allocation remains relatively constant with the course of time. Considered on the whole, no decrease in the variation of the allocation can be identified for both countries.

According to expectations from neoclassical theory this finding is surprising since in the long-run infrastructure endowments across regions should become more balanced and therefore the variation in the regional allocation of infrastructure should decrease in the course of time. However, this only applies if the government actually aims at equalising the marginal productivities of infrastructure across regions. Therefore, in the following we examine the empirical determinants of the allocation of infrastructure investment across regions.

de la Fuente and Vives (1995) identify three criteria which could be of relevance for the regional allocation of public investment in road infrastructure. If the government strives for equal living standards in all regions, then according to the first criterion, labelled *equality*, investment in infrastructure should be directed to where the regional per capita income or productivity respectively is below average. The aim of such a policy would be to reduce competitive disadvantage in

Figure 5.3: Regional Differences of Road Infrastructure Investment in Germany and France



a region by improving the public infrastructure and thereby to attracting private investment. In order to operationalise the *equality* criterion empirically, we use the labour productivity  $Q_i/L_i$  as a measure for regional income differences.

According to the second criterion, which is labelled as *efficiency* criterion, infrastructure investment should flow where the marginal productivity of investment is highest. Thus the objective of this criterion would be to maximise the sum of regional incomes. In order to obtain an empirical measure for this criterion, by assuming a linear homogenous production function  $F$  of degree one the marginal productivity  $\partial F/\partial G_i$  of the infrastructure capital stock in region  $i$  is proportional to the ratio of the output  $Q_i$  to the infrastructure stock  $G_i$ , i.e.

$$\frac{\partial F}{\partial G_i} \sim \frac{Q_i}{G_i}.$$

Thus, we use the ratio  $Q_i/G_i$  as an operational measure for a regional infrastructure policy criterion according to the *efficiency* criterion.

A third criterion for the allocation of investment in infrastructure can be labelled as *neutrality*. The rationale for this criterion is that the state should ensure that differences in public capital stocks do not give an unfair advantage or disadvantage to any region. The goal of this policy would be to equalise the infrastructure endowments across regions. In practice, this criterion would be met when for example  $G_i/L_i$ , i.e. the capital intensity of infrastructure (or any other regionally comparable measure for infrastructure endowment), is equal in all regions.

If a government intends to apply these criteria to the decision process of regional investment allocation, in most cases these three criteria cannot be fulfilled simultaneously. On the contrary, they will quite often lead to conflicting priorities regarding the ranking of infrastructure investment projects. For the Spanish regions, for example, de la Fuente and Vives (1995) find a conflict between the *efficiency* criterion on the one hand, and the *neutrality* criterion and *equality* criterion on the other hand.

Table 5.2 shows the results of a regression of  $I_{it}/L_{it}$ , i.e. investment per labour as a measure for the regional allocation of infrastructure investment, on the measures for the three criteria *equality*, *efficiency* and *neutrality* described above. Furthermore, we have also added a measure for private capital intensity, i.e.  $K_{it}/L_{it}$ , which is another potential determinant of regional infrastructure investment.

Note, that infrastructure investment is measured in *net* figures in order to reduce the size effect in the allocation of investment, due to the fact that higher stocks also require higher maintenance investment which is reflected in higher *gross* investment figures.

One important issue for the implementation is that at least for Germany, investment of different levels of governments, i.e. the Federal government, the governments of the federal states and the local governments of the counties, are included. One could argue that the autonomous investment decisions at lower governmental levels are unlikely to reflect especially the *equality* and *neutrality* criteria, whereas the *efficiency* criterion should be also relevant for investment undertaken by lower levels of government. However, investment by the federal government is not only the main part of total infrastructure investment, but the federal government can also influence investment decisions at lower government levels via its investment grant policy. Thus, we argue that this investment figures including all levels of government are appropriate for the problem of regional investment allocation we study here.

In case the government pursues a regional infrastructure policy according to the *neutrality* criterion, we expect a negative correlation between the investment per person in work ( $I_{it}/L_{it}$ ) and the infrastructure intensity ( $G_{it}/L_{it}$ ). A negative correlation between ( $I_{it}/L_{it}$ ) and ( $Q_{it}/L_{it}$ ) is expected if the government allocates investment according to the criterion of *equality*. Hence, regions with lower income will receive more investment. Finally, if the government pursues a regional infrastructure policy according to the criterion of *efficiency*, then we expect the correlation between ( $I_{it}/L_{it}$ ) and ( $Q_{it}/G_{it}$ ) to be positive. Thus, regions where the expected returns of infrastructure investments are higher would obtain more investment.

The estimation of the regression in Table 5.2 has been carried out separately for the German and the French regions by again using MLE with an ARMA(1,1) specification for the covariance structure. Also, fixed cross-section and time-effects were added to the specification, which turned out to be highly significant.

From the results reported in Table 5.2, we find that only the criterion of *neutrality* can explain differences in the amount of infrastructure investment the regions

**Table 5.2:** *Determinants of Regional Infrastructure Investment Allocation in Germany and France*

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*Results of the regression analysis:*

Dependent variable:  $I/L$

	(1) Germany	(2) France
Independent variables:		
<i>Equity</i> $Y/L$	(0)	(0)
<i>Efficiency</i> $Q/G$	(0)	(0)
<i>Neutrality</i> $G/L$	(-) <sup>***</sup>	(+) <sup>**</sup>
<i>Private Capital</i> $K/L$	(0)	(0)

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Significance levels: \*\*\*1%, \*\*5%, \*10%

(0) not significant, (+) with positive sign, (-) with negative sign

receive, but only for Germany it has the expected sign. This finding fits well into the institutional framework of infrastructure policy in Germany we have described in section 5.2.1 where priority is given to the convergence of living conditions throughout all regions in Germany. For France, a potential explanation for the positive sign of the criterion of *neutrality* is that in the process of deconcentration of the Paris region, political priorities are given to the development of certain regions, but not to the development of *all* regions with equal priority.

On the other hand, surprisingly neither the criterion of *efficiency* nor the private capital intensity are significant for Germany or France. Thus, governments do not seem to anticipate the expected returns of infrastructure investments in the decision process. Moreover, the *equity* criterion is not significant. As a by-product of this result we can also infer that—contrary to what is often presumed in the literature—simultaneity between infrastructure investment and output is negligible, i.e.  $Q_{it}/L_{it}$  does not determine  $I_{it}/L_{it}$ . It is worth noting that if simultaneity matters we expect a positive sign for  $Q_{it}/L_{it}$ , because prosperous states are financially more capable of infrastructure spending than poorer ones.

As a result of this section we can determine on balance that, contrary to France, the criterion of *neutrality* does appear to play a role in the allocation of

public infrastructure investment in Germany. Public investment in Germany has flowed above all into regions with a below average initial endowment of road infrastructure, thus in Germany infrastructure policy is used as an instrument of regional policy to minimise the competitive disadvantages of economically underdeveloped regions.

## 5.4 Summary and conclusions

In the first part of this study, we described the differences in infrastructure and regional planning policies between Germany and France. In France, for example, a dominance of Paris and the surrounding region compared to the other French regions can be observed. A further basic difference is that regional planning in Germany is divided hierarchically between the regional authorities and is conceived in the medium to long-term. In contrast to this, regional planning in France is based on so called 'planning contracts' between the state and the regions, in which the individual regional authorities have equal rights and planning is conceived rather in the medium to short-term.

Following the description of the institutional concepts of regional infrastructure policy in Germany and France, the effects of road infrastructure on productivity were examined for the German and French regions in the second part of this study. For that purpose, production functions were estimated using the data of an 'unbalanced panel' consisting of the 21 French regions for the period 1978-92 and the 11 West German federal states for the period 1970-95.

On the whole it can be concluded that regional road infrastructure has a significant impact on regional output. In addition, we find evidence that the direct effect arising from infrastructure, i.e. increasing the marginal productivities of private factors, is more important than the indirect effects, i.e. the positive effects on the demand for private factor inputs. As a caveat, however, the results regarding the indirect effects of infrastructure should be interpreted with some caution due to the strong correlation between the explanatory variables in the translog model.

In addition, we cannot observe a decrease in income disparities in the period of investigation either in Germany or France. This finding is at odds with the predictions of neoclassical economic theories. On the contrary, economic counter-effects could have emanated that worked against regional convergence as assumed by models of the 'New Economic Geography' due to the removal of transport barriers. The explanation of this puzzling evidence is a challenging issue for future research.

Finally, the determinants of the allocation of infrastructure investment in Germany and France were examined empirically by applying a new approach. Surprisingly, neither in France or Germany do efficiency considerations matter for the allocation of infrastructure investment across regions. However, it could be shown that in Germany, in contrast with France, the criterion of *neutrality* plays a role in the allocation of public investment in infrastructure, which means that public investment flows above all to those regions which have a below average endowment with public capital. Despite the principal difficulty to link the institutional differences in infrastructure policies in Germany and France with this evidence of the determinants of infrastructure investment allocation, we interpret this finding as a reflection of the priority of promoting the regional convergence of living conditions for infrastructure policy in Germany.

## Appendix Essay 5.

### 5A. Data

The German-French regional data include 21 of the 22 French regions (Corsica was not included due to incomplete statistical information) for the period 1978-1992 and 11 West German federal states for the period 1970-1995, although in the case of West Berlin, data is only available for the period 1970-1990.<sup>3</sup> All values have been converted into ECU at constant 1991 prices. For investment in transport infrastructure, we are able to differentiate in France between roads, rail and inland waterways. The infrastructure data for France are also described in Fritsch (1995) and Cadot et al. (1999). Note, that road infrastructure investment in France includes both public investment for all road categories and private investment for licensed motorways.

In case of Germany, with respect to transport infrastructure only investment data for *road* infrastructure are available at the regional level of the Bundesländer. Therefore, the empirical analysis focused on road infrastructure in order to allow a comparison of the infrastructure data between Germany and France.

For Germany, we are able to differentiate investment between categories of road (A roads, B roads and smaller roads and streets). An internal report of the German Ministry of Transport, Building and Housing was used as a source for the investment made by the Federal Government and the Federal States ('Straßenbaubericht 1996'). Thus, this report gives the allocation of Federal investment for motorways and A roads across the Bundesländer in the period 1970-1995.

The information regarding investment made by the State and local governments in B and smaller roads was taken from a publication issued by the Federal

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<sup>3</sup>The data used in this analysis are available from the authors upon request.

Statistical Office Wiesbaden '*Rechnungsergebnisse des öffentlichen Gesamthaushalts*', series 14, section 3.1. It contains the road investment of the different bodies at the regional level of the Bundesländer.

The regional capital stocks of road infrastructure in Germany and France were determined from the regional investment series (French regions 1975-1992, German Bundesländer) using the 'Perpetual Inventory Method' (PIM). Different procedures were used for both Germany and France. The problem in the case of France was to determine the initial capital stock for each region for 1975. Thus, the aggregated transport infrastructure stocks in France as given by the Federation Nationale des Travaux Publics (FNTP) have been allocated proportionally to the individual regions in accordance with the investment proportion of the individual regions. The calculated value was then used as the initial stock for the Perpetual Inventory Method (PIM). For the linear depreciation rate, we assumed a value of 2.5 percent. As a control for the capital stocks of road infrastructure arrived at by using this method, the sum over the individual regions was computed and compared with the aggregated value reported by FNTP. It became apparent that the deviation between the sums of the regional and the aggregated stock was only between 1 and 2 percent.

In contrast to this method applied for France, with regards to Germany it is possible to use a study carried out by the German Institute for Economic Research (DIW), in which the regional capital stock of road infrastructure was estimated for the West German federal states for the year 1970 (Bartholmai, 1973). In order to update the initial stocks for 1970 over the period 1971-1995, the Perpetual Inventory Method was used. The publication series '*Verkehr in Zahlen*' (Transport in figures) of the DIW also gives the aggregated stock of road infrastructure for the period 1970-1995. Therefore it was possible to apply a restriction for the calculation of the regional stock. Contained within this restriction was the assumption that the sum of the stocks in the regions equals the aggregated value for Germany given by DIW.

Furthermore, the majority of regionally specific measures were obtained from official statistics such as value added as a measure for output we use in the analysis. In the case of Germany, the majority of this data originates from the series

of 'National Accounts for the Bundesländer' which is published by the Statistical Office of Baden-Württemberg ('Statistisches Landesamt Baden-Württemberg').

Gross value added (for all areas at market prices) is taken from the publication 'Entstehung des Bruttoinlandsprodukts in den Ländern der Bundesrepublik Deutschland 1979-1996' ('The Origin of Gross National Product in the Federal States of the Federal Republic of Germany, 1970-1996'), vol. 30, and is used as a measure for the output  $Q$  of the federal states for the period 1979-1995.

For France, the measure for labour has been taken from the EUROSTAT data base 'New Cronos', edition June 1999. The regional value added data at market prices for the years 1980-1992 were drawn from the EUROSTAT publications 'Regional and Statistical Yearbook, Series 1A, 1993, 1995.' The values for 1979 and 1978 were extrapolated using the information of the development of gross domestic product (GDP) for the years 1978 and 1979.

The data relating to the regional stock of private capital in France for the period 1978-1991 were provided by Professor Remy Prud'Homme of the University of Paris. A description of these data can be found in Prud'Homme (1996). The stocks for the year 1992 were computed by applying the Perpetual Inventory Method from the stocks in 1991 by adding regional gross investment in 1992 for all industries taken from the 'New Cronos' data base and assuming a linear depreciation rate of 10 percent.

## **Part III**

### **Concluding Remarks**

The bottom line of this thesis is that throughout the essays evidence of a positive impact of infrastructure on private productivity is found. This evidence holds also for different levels of aggregation studied in this thesis; whereas the Essays 1, 3 and 5 are based on data at the regional level of the Bundesländer and the French regions respectively, Essay 2 is based on data at the local level of the German counties and Essay 4 is based on data at the local level of large self-administrated German cities.

Also, different measures of public capital are used in the Essays. The analysis of Essay 1 is based on a broad measure of public capital including transport infrastructure, water and sewer systems, pipelines, etc. In Essay 2, infrastructure is measured using indicators describing both human capital and transport infrastructure endowment at the local level. Essay 3 focuses only on transport infrastructure, whereas Essay 4 uses again a broad measure of public capital. Finally, the analysis in Essay 5 is based on road infrastructure capital stocks in German and French regions.

Moreover, the analyses of the infrastructure effects in the Essays are performed for different economic sectors. Essays 1 and 4 perform the analysis for the manufacturing sector, whereas Essay 2 conducts the analysis for several sectors and Essays 3 and 5 conduct the analysis for the total economy including all sectors.

Some caveats, however, are in point here. As stated in Essay 1, the existence of positive effects of public capital on private production is a necessary, but not a sufficient condition for drawing the conclusion that public investment should be boosted in the future. To make such an inference, the costs of the public capital provision have to be included in the analysis as well.

For instance an increase in public investment may only be possible if tax revenues are also increased. This in turn can give rise to distortions bearing additional costs for the economy. Similarly, if higher public investment is financed by higher governmental debt this may also imply other kinds of additional costs e.g. higher interest rates on capital markets. In this respect, this thesis has focused only on the necessary condition for increasing the supply of public capital,

i.e. the existence of significant and positive effects of public capital on private production.

From a more theoretical point of view, the sufficient condition for increasing public investment in the future is that the social net benefit—defined as the sum of social gross benefits (consumer and producer surpluses, positive externalities e.g. spillover effects, etc.) minus the sum of social costs (costs of provision, negative externalities e.g. environmental effects, etc.)—is positive. Consequently, the appropriate instrument to conduct such an investigation is social cost-benefit analysis.

In addition, a word of caution pertains in extrapolating these findings of positive effects of infrastructure in the past to the future. Even if infrastructure investment was productive in the past, it might not be productive in the future. While building the first road to a remote region might be very productive, once a network of roads is already established the addition of further roads might induce no gains in private productivity.

With respect to the simultaneity between output and infrastructure investment, we find only little evidence in the Essays 3, 4 and 5 that output is a determinant for infrastructure investment. Thus, in contrast to what is often claimed in the literature, reverse causation running from output to infrastructure investment appears—at least for the various samples studied here—not to be significant.

Finally, regarding the predominant politico-institutional determinants of infrastructure investment, we find that lobbying and political affiliation matter for the regional allocation of infrastructure investment, whereas efficiency considerations appear not to have an influence on the regional allocation of public investment. In conclusion, the observed allocation of regional infrastructure investment may depart substantially from a socially optimal allocation.

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## **Erklärung**

Hiermit erkläre ich, dass ich außer der im Literaturverzeichnis angeführten Literatur keine weiteren Hilfsmittel benutzt habe. Außer den unter 'Acknowledgements' genannten Personen habe ich keine weitere Hilfe erhalten. Ich bezeuge durch meine Unterschrift, dass meine Angaben über die bei der Abfassung meiner Dissertation benutzten Hilfsmittel, über die mir zuteil gewordene Hilfe sowie über frühere Begutachtungen meiner Dissertation in jeder Hinsicht der Wahrheit entsprechen.

Berlin, den

Andreas Stephan