The relationship between regional value added and public capital in Finland: What do the new panel econometric techniques tell us?*

Pellervo Hämäläinen[†] Tuomas Malinen[‡]

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Abstract

This paper applies new panel estimation techniques to the estimation of the elasticity of private production with respect to public capital in a regional setup. We use the widely applied production function approach and regional data from Finland in the period of 1975-2004. In contrast to many previous studies about the productivity of public capital, we focus especially on panel estimation techniques. We show that the results from commonly applied fixed effects OLS are probably biased and sensitive to change of an estimator. To get more reliable results, we use the panel DOLS and panel DSUR estimators.

JEL Codes: C33, H11, H54
Keywords: Production function, panel unit root tests, panel DOLS, panel DSUR

1 Introduction

The productivity of public investments has been in the research agenda for 20 years. Since Aschauer's (1989) article number of studies using different approaches and data sets have been made.[1] Researchers using aggregate level data have generally ended up with the conclusion that the impact of public capital on the private sector's productivity is positive, but much smaller than Aschauer's original estimate.

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[†]Department of Economics, University of Turku. email: pellervo.hamalainen@utu.fi. Financial support from the Finnish Cultural Foundation and Yrjö Jahnsson Foundation is greatly acknowledged.

[‡]Department of Economics, University of Helsinki. email: tuomas.malinen@helsinki.fi. Wishes to thank Henri Nyberg and Leena Kalliovirta for comments.

[1]For theoretical background look, e.g. Arrow & Kurz (1970), and for older empirical paper e.g. Ratner (1983).

The aims, developments and previous results of this literature are extensively discussed in literature reviews written, for example, by Romp and de Haan (2007), Strum, Kuper and de Haan (1996) and also by Ligthart and Bom (2008), who take a meta-analysis approach to sum up results from previous studies.

The direction of the research on the productivity of public capital has recently changed toward regional panel analyses. The tightest restriction in regional analyses has been, and still is, the availability of suitable capital stock data. Therefore, regional studies have been made only in few countries outside the US.

The purpose of this paper is to apply new methods of panel data econometrics to Finnish regional data and estimate the elasticity of private production with respect to public capital. Our analysis departs from previous literature in the following ways: (i) We are using, in addition to traditional panel unit root tests, also tests that allow for spatial dependence across regions and breaks in the tested series. (ii) In addition to the basic fixed effects OLS-estimator, we are using also panel DOLS and panel-DSUR-estimators. Panel DOLS have been used previously only in Okubo (2007). To our knowledge, panel DSUR has not been used before in this branch of literature. (iii) The Finnish regional data has not been used before and studies using data from any of the Nordic countries have been rare.

The history of regional productivity analyses focusing especially on the productivity of public capital and applying the production function approach could be thought to begin from Mera's (1973) research with the regions of Japan. Most of the regional studies on the 90's used the data from the US states (e.g. Munnell (1990), Garcia-Mila & McGuire (1992), Evans & Karras (1994), Holtz-Eakin (1994) and Garcia-Mila, McGuire & Porter (1996)). Some researchers found positive and some insignificant effect of public capital or infrastructure on the private sector productivity.[2]

More recently, regional capital stock data have been available increasingly also for European countries, which has led to an increasing extent of regional studies using European data. Positive effect of public capital on private production or TFP is identified, at least to some extent, for West Germany (Stephan 2003), Italy (Destefanis and Sena 2005), France (Cadot, Röller and Stephan 2006) and Spain (Moreno and López-Bazo 2007 and Salinas-Jimenez 2004). In regional setup, spill-over effects of public capital have also received a lot of attention (see e.g. Pereira & Roca-Sagalés 2003).

Econometric methodology has varied a lot in previous studies although basic fixed effects OLS is the most commonly used estimator, especially in the older literature. Stephan (2003) applies feasible generalized least squares (FGLS) estimator, Salinas-Jiminez (2004) uses fixed effects instrumental variable -estimator using lagged values as instruments and Moreno and López-Bazo (2007) use fixed effects OLS with time dummies. Destefanis and Sena (2005) use t-bar (Im *et al.* 2003) test to examine possible unit roots. Results indicate that TFP would be trend-stationary, but authors take it as I(1)-series and proceed to cointegration analyses and testing the possible long-run relationships. They apply also free disposal hull to get non-parametric estimates.

Among these four studies, for example, results from unit root studies have been reported only in Destefanis and Sena (2005) although cointegration has been tested in three of them. Especially in some older studies, econometric specification does not get much attention. In the matter of fact, studies that mainly focus on the econometric problems are clearly in a minority. This is one motivation to the point of view of our article.

[2]In the 21st century US state data has been used mainly in cost function studies (e.g. Cohen & Morrison (2004).

Kelejian and Robinson (1997) and Okubo (2007) are exceptions by focusing on empirical problems in regional panel data analyses. Kelejian and Robinson (1997) noted that, if all econometric problems are ignored and production function is estimated with basic OLS, the results are in line with previous studies. However, the picture is quite different, when they correct, for example, autocorrelation, variable endogeneity and heteroscdasticity. They conclude, based on estimation results for the US data, that public capital gets positive and significant elasticity estimate only in the specifications that ignore econometric problems. Okubo (2007) shows that commonly found negative elasticity estimate for public capital changes to positive, if panel DOLS estimator is used for regional data of Japan. Okubo argues that negative estimate is previously found, because of endogeneity bias, which can be corrected by using panel DOLS.

In this paper, we show that all unit root tests suggest that our dependent variable, value added, would be a trend-stationary process in the period tested. This strongly contradicts the results obtained in previous studies, where value added series is usually found or assumed to be a I(1)-process. We argue that this peculiar finding may result from the fact that we have both stationary and nonstationary regional specific value added series in our data. Furthermore, we will show that the results of commonly used fixed effects OLS and panel DSUR differ substantially in sub-samples. It seems, based on the evidence got with the Finnish data, that fixed effects OLS produces unreliable results in a regional setup. In addition, considerable differences in the behavior of regional specific series may lead to problems in inference of the results estimated using the data from all regions.

One caution about the terminology is in order. We are using public net capital stock, which is a quite broad concept compared to infrastructure capital. This choice is rationalized by the lack of proper data of the whole infrastructure capital stock in Finnish regions. Public capital includes part of the infrastructure, but there are also a great many other items (such as public buildings, for example) included. It should also be noticed that in the National Accounts part of the infrastructure capital is included in the private sector's accounts.

2 Model, data and tests

Theoretical framework

We assume that every region has the following general form production function

$$Y_t = A(K2_t) \cdot F(L_t, K1_t, K2_t), \tag{1}$$

where Y is private output, L is private labor, $K1$ is private capital stock and $K2$ is public capital stock of the region.

More precisely, we define $F(\bullet)$ to be Cobb-Douglas type, as it is commonly assumed

$$F(\bullet) = L_t^{\beta_1} \cdot K1_t^{\beta_2} \cdot K2_t^{\beta_3}, \tag{2}$$

Parameter β_i measures the elasticity of private output with respect input $i = \{L, K1, K2\}$.

In addition, Hicks-neutral technological progress A of the region is specified as follows[3]

$$A = Ce^{\delta \cdot t} \cdot K2_t^{\beta_4}, \tag{3}$$

[3] Actually, in the case of Cobb-Douglas, the choice between Hicks, Harrod or Solow neutral technological progress does not matter.

where C is a constant describing initial level of productivity in the region and t is a linear time trend.

Combining equations (2) and (3) yields

$$Y_t = Ce^{\delta \cdot t} \cdot L_t^{\beta_1} \cdot K1_t^{\beta_2} \cdot K2_t^{(\beta_3+\beta_4)}, \tag{4}$$

which is the standard specification used in regional analyses (see e.g. Mas, Maudos, Pérez & Uriel (1996) or Stephan (2003)).

The limitations and problems of this production function approach are well-known and discussed e.g. in Romp and de Haan (2007) or Destefanis and Sena (2005). However, we are interested in estimating the elasticity of private production with respect to public capital in a robust way for the panel data from Finnish regions. Therefore, we are taking this standard framework as given, which makes our paper, in this respect, also comparable to previous researches.

Data

The data consists of yearly observations from 77 Finnish sub-regional units in the period of 1975-2004. Private sector's regional production is measured as value-added at factor prices. Private labor consists of number of workers in each region. Regional net capital stocks are taken from Salmela (2008), who has constructed those series using the current National Account standards. Variables are measured as constant prices at 2000 and the regional division corresponds to the situation in the year 2005. [4] The variables are described in more detail in the data appendix.

Unit root tests

Some of the previous studies made on the topic have proceeded to cointegration analyses without testing the unit roots at all (e.g. Moreno & López-Bazo (2007)). This is an odd method of analysis because many of the panel cointegration tests are residual based, i.e. they test is the residual nonstationary or stationary (e.g. Kao (1999), McCoskey and Kao (1998), Pedroni (2004)). If the dependent variable is not nonstationary, residual based cointegration tests can give flawed results. Many previous studies have also relied on the so called traditional panel unit root tests that assume independence of cross-sections. This is a very restrictive assumption when testing includes regions within a country. In a testing setup, where different sub-regions' series of value added are tested, it is very likely that majority of the different series are correlated and/or cointegrated with each other.

The traditional panel unit root tests are usually based on the following regression:

$$\triangle y_{it} = \rho_i y_{i,t-1} + \delta_i + \eta_i t + \theta_t + \varepsilon_{it}, \tag{5}$$

where δ_i are the individual constants, $\eta_i t$ are the individual time trends, and θ_t are the common time effects. Tests rely on the assumption that $E[\varepsilon_{it}\varepsilon_{js}] = 0 \; \forall \; t,s$ and $i \neq j$, which is required for the calculation of common time effects. Thus, if the different series are correlated and/or cointegrated, the last assumption is violated. Despite of this restriction, some tests are found to be consistent under cross-sectional cointegration (Banerjee *et al.* 2005).

[4]The names and the locations of sub-regional units can be found from appendix

The traditional panel unit root tests used in this study are based on the regression presented in equation (5). The null hypothesis is that $H_0 : \rho_i = 0 \ \forall \ i$. Tests have different assumptions about the heterogeneity of the unit root process. Levin, Lin and Chu (2002) (LLC), and Breitung's (2000) tests assume that the unit root process is common to all cross-sections and Im, Pesaran, and Shin (2003) test (IPS) and Fisher type ADF and PP tests, presented by Maddala and Wu (1999), allow for an individual unit root processes. The inclusion of individual constants and time trends is optional, although Breitung's test requires that individual trends are included. The alternative hypotheses also differ between tests. Under alternative hypothesis Levin *et. al* test and Fisher type ADF and PP tests assume that all series are stationary, whereas Im *et al.* test allows some of the series to be nonstationary. Breitung's test assumes that under alternative hypothesis all cross-sections are trend-stationary.

Table 1 presents the results of five traditional panel unit root tests.[5] First test includes individual constants and the second test includes individual constants and deterministic trends.

Table 1: Traditional panel unit root tests

variable	LLC	Breitung	IPS	ADF	PP
value added (c)	-4.1863	-	-0.068	173.81	178.12
	(<.0001)		(0.473)	(0.109)	(0.073)
value added (c&t)	-8.717	-6.036	-8.139	303.73	309.41
	(<.0001)	(<.0001)	(<.0001)	(<.0001)	(<.0001)
labor (c)	-2.885	-	2.274	93.27	74.392
	(0.0020)		(0.989)	(1.000)	(1.000)
labor (c&t)	1.631	-2.132	1.618	117.49	53.021
	(0.949)	(0.017)	(0.947)	(0.983)	(1.000)
private capital (c)	-11.814	-	-7.479	328.48	591.72
	(<.0001)		(<.0001)	(<.0001)	(<.0001)
private capital (c&t)	-7.816	2.124	0.493	139.64	356.42
	(<.0001)	(0.983)	(0.689)	(0.755)	(<.0001)
public capital (c)	-24.468	-	-13.911	515.56	560.41
	(<.0001)		(<.0001)	(<.0001)	(<.0001)
public capital (c&t)	3.849	17.767	16.821	24.289	22.539
	(0.999)	(1.000)	(1.000)	(1.000)	(1.000)

All variables are tested in logarithms. (c) denotes that individual constants and (c&t) that individual constants and trends have been included in the test. Probabilities of the test statistics are presented in parentheses.

One thing is clearly visible in the results of table 1. The results of the tests crucially depend on the inclusion of individual trends. If there are no deterministic trends included in the test, they may give flawed results. This is because the inclusion of individual deterministic trend does not alter the test. It just removes a trend in the series, if there is a trend. That is why we concentrate only on the results of those tests which include both the individual constants and trends.

According to all tests the series of value added is trend-stationary. All the other variables are non-stationary according to all tests except the series of private capital which

[5]All the tests have been done with Eviews 6. Lag lengths have been determined using Schwarts information criterion, spectral estimation has been conducted with Bartlett kernel and bandwidth has been selected using Newey-West method.

is stationary according to LLC and the series of labor, which is stationary according to Breitung's test. However, it is likely that most of the tested series are correlated or even cointegrated across sub-regional units. This would violate the assumption of uncorrelated residuals among cross-sections, i.e. $E[\varepsilon_{it}\varepsilon_{js}] = 0 \; \forall \; t,s$ and $i \neq j$. Banerjee *et al.* (2005) have studied the effect of the violation of the assumption of no cross-unit cointegration to rejection frequencies of the null hypothesis. Their results show that in the presence of cross-unit cointegration ADF, PP, and IPS tests grossly overreject the null hypothesis of unit root with small T and relatively large n dimension of data. As all tests accept the null hypothesis of unit root in labor and public capital series, they seem to be unit root processes. Results for private capital are inconclusive.

The dependence between cross-sections may go further than the one studied by Banerjee *et al.* (2005). Different regions of a country are likely to be *spatially dependent* as they (usually) lie in the same geographical area. This would violate the assumption of independence of error processes, but the different spatially dependent statistical units need not to be statistically correlated or integrated (Baltagi *et al.* 2007).

Many Nordic countries experienced a severe economic downturn at the beginning of the 1990s. In Finland, one of the most important factors that contributed to this rapid downturn was financial crisis that stemmed from reckless lending by banks after credit restrictions were eased in the late 1980s. In the aftermath, one of the major banks in Finland went bust and Finland as a country was driven on the verge of bankruptcy. Cause of bursting property and equity bubbles and aggressive cutbacks in lending, the downturn was very rapid (GDP growth was +5,4% in 1989, +0,1 in 1990 and -6,2% in 1991 followed by two years of contraction). This *structural shift* is clearly visible in the Finnish GDP series. It also likely that in the span of 30 years almost all countries in the world have experienced a recession. That is why the possibility of structural breaks should be taken into account in unit root tests.

To account for spatial dependence in the tested series we conduct Phillips and Sul's (2003) (PS) panel unit root test. Baltagi *et al.* (2007) found that it performed robustly in the presence of spatial dependence compared to traditional panel unit root tests. Phillips and Sul's test is based on the regression

$$\triangle y_{it} = \rho y_{i,t-1} + \eta_i t + \alpha_i + \delta_i \theta_t + \varepsilon_{it}, \tag{6}$$

where α_is are the individual constants, $\eta_i t$ are the individual time trends, θ_t is the common time effect whose coefficients, δ_i, are assumed to be non-stochastic, measure the impact of the common time effects of series i, $\varepsilon_{it} \sim i.i.d.N(0,\sigma^2)$ over t, and ε_{it} is independent of ε_{js} and θ_s for all $i \neq j$ and s,t. Cross-sectional dependence is allowed through the common time effects which are proxied by the cross-section mean of $y_i t$ ($\bar{y}_t = N^-1\Sigma_{j=1}^{n} y_{jt}$) and its lagged values, $\bar{y}_{t-1}$, $\bar{y}_{t-2}$, etc. The null hypothesis is that $H_0 : \rho_i = 0 \; \forall \; i$ and the alternative hypothesis is that majority of the series are stationary.

To account for possible structural breaks in the tested series we use Im *et al.* (2005) (ILT) panel unit root test that allows for structural shifts in the tested series. Im *et al.* test assumes the following data generating process:

$$\begin{aligned} y_{it} &= z_{it} + x_{it} \\ z_{it} &= \gamma_{1i} + \gamma_{2i} t + \delta_i D_{it}, \\ x_{it} &= \phi_i x_{i,t-1} + \varepsilon_{it} \end{aligned} \tag{7}$$

where

$$D_{it} = \begin{cases} 0 & t \leq T_{B,i} \\ 1 & t \geq T_{B,i} + 1 \end{cases},$$

where $T_{B,i}$ is the time period of structural shift in the ith series and $\varepsilon_{it} \sim i.i.d.N(0, \sigma^2)$. After rearranging, equation (7) becomes

$$\triangle y_{it} = \beta_i y_{i,t-1} - \beta_i \gamma_{1i} + [1 - (\beta_i + 1)(t-1)]\gamma_{2i} + \varepsilon_{it}, \tag{8}$$

where $\beta_i = -(1 - \phi_i)$. The null hypothesis is that $H_0 : \phi_i = 0 \; \forall \; i$ and the alternative hypothesis is that $H_1 : \phi_i < 0$ for some i.

We run two versions of each test. For Phillips and Sul's (2003) test we first run a test including only individual constants and then a test that includes both individual constants and deterministic trends. First Im *et al.* (2005) test allows for no breaks in the tested series and the second one allows for one common break in the series.

Im *et al.* (2005) test estimates the time of the break in the different series and then uses a common time dummy to control for the break. Im *et al.* test is consistent only when there is a break in the series. That's why we only report the results of ILT test with break when there seems to be a one structural break in the series. In the individual time series of value added, there is a clear break point visible in the value added series in the year 1990. This is also the same year that the ILT test estimates as a break point. Labor series seems to have two break points: In the late 1970s and around 1990. ILT estimates the break point to be in the year 2000, which is clearly off. In the private capital series, there is a clear break visible in 1990-1991. ILT estimates that break point is in 1991. In the case of public capital series, there is no break visible. Table 2 presents the results of Im *et al.* and Phillips and Sul's (2003) tests.[6]

Table 2: Panel unit root tests allowing for cross-sectional dependence/structural shifts

variable	ILT (no break)	ILT (1 break)	PS (c)	PS (c&t)
log(value added)	-8.384	-8.718	100.98	319.49
	(<.0001)	(<.0001)	(0.999)	(<.0001)
log(labor)	-5.476	-	70.450	223.44
	(<.0001)		(1.000)	(<.0001)
log(private capital)	0.158	-33.990	151.43	266.99
	(0.934)	(<.0001)	(0.452)	(<.0001)
log(public capital)	-0.771	-	199.79	97.301
	(0.656)		(0.004)	(0.999)

P-values of the test statistics are presented in parentheses. The values presented without brackets denotes the value of *z*-statistics.

Results of Phillips and Sul's (2003) test support the findings of traditional tests, i.e. the inclusion of individual trends in the test alters the results significantly. On the other hand, according to Im *et al.* (2005) test, value added and private capital series are stationary, if we allow for one break point in the tested series.

Thus, both tests seem to enforce the result of the traditional tests, i.e. that the value added series would be stationary $AR(p)$ process in the tested period. However, all of these tests have their reservations. Traditional panel unit root tests and ILT test assume cross-sectional independence. Phillips and Sul's (2003) test allows for cross-sectional correlation, but may be inconsistent in the presence of cross-sectional cointegration. If tested series are cross-sectionally cointegrated, the common trends present in the data

[6]Tests have been conducted with Gauss. Gauss code for ILT and PS tests were provided by Im *et. al* (2005) and Phillips and Sul's (2003).

may be identified as common factors in equation (6) and removed from the analysis (Breitung & Pesaran 2005). In this case, if the remaining idiosyncratic component is stationary, the test has tendency to present the time series as stationary when panel units are actually nonstationary. [7]

The trend-stationarity of value added series may also result from large number of stationary value added series in the panel. The economic reason for this finding could be the highly diverse economic development in the Finnish sub-regional units. In Finland, especially, population is concentrated on few rapidly growing areas, and most of the value added growth comes from these few heavily populated areas. Thus, we may have several depressing sub-regions, whose value added growth is slightly upward sloping or stays more or less constant. In these sub-regions, the value added series may be more like a trend-stationary series, whereas few heavily populated rapidly developing regions have clearly more dynamic, integrated value added series growth.

So, the somewhat surprising result of trend-stationarity of the regional value added series could result from a strong cross-sectional cointegration and/or from large number of stationary series in the panel. Series could also have *local unit processes*, which could bias the results of our panel unit root tests. For these reasons we have also run Pedroni's (2004) panel cointegration tests on our variables. Table 3 summarizes the results of Pedroni's panel cointegration tests.[8]

Table 3: Summary of the results of Pedroni's panel cointegration tests

Dependent variable: log(value added)	Number of test statistics that reject H_0 at 5% level
log(labor)	10/11
log(private capital)	9/11
log(public capital)	10/11
all variables	10/11

Null hypothesis is that there is no cointegration between the variables.

As is visible in table 3, results of Pedroni's (2004) panel unit root tests clearly support the hypothesis of cointegration between value added and all explanatory variables. But, as Pedroni's (2004) test is residual based, it requires that the dependent variable (value added) is a nonstationary process. As almost all our panel unit root tests conclude that the value added series is trend-stationary, the results of Pedroni's panel cointegration tests have to be taken cautiously. Nonetheless, we are left with one choice. Try to take the possibility of cointegration between the value added series and some or all series of explanatory variables into account in estimation. Assuming a $I(0)$ dependent variable with $I(1)$ regressors would also create problems in estimation cause such a setup could result to spurious regressions (Stewart 2007).

[7] Panel unit root tests that allow for different forms of cross-sectional dependence, including cross-sectional cointegration, have been developed, but, to our knowledge, all these tests require panels with large dimensions of T and n (eg. Bai & Ng 2004). Our panel has a relatively large cross-sectional dimension (76), but the time dimension is relatively small (29 observations).

[8] Detailed results are available upon request.

3 Estimation

Some previous studies have used standard panel estimators to estimate variables that are found or assumed to be cointegrated. Unfortunately, many standard panel estimators are not consistent or asymptotically unbiased in panel cointegrated data. For example, GMM estimator is, by definition, inconsistent in panel cointegrated data. OLS is also not asymptotically unbiased, if panel includes cointegrating relations between the dependent and explanatory variables (Kao & Chiang 2000). We account for the possible cointegration in the panel by using the panel dynamic OLS estimator, which is a consistent estimator in cointegrated panel data, and that accounts for possible endogeneity present in the model. However, panel DOLS does not fully account for the possible correlation and/or cointegration between statistical units of the panel. That's why we also use the panel dynamic seemingly unrelated regressors estimator, which accounts for this correlation/cointegration.

To make a comparison, we first estimate our production function with traditional panel estimators. We estimate a model:

$$\begin{aligned} log(valueadded)_{it} = \alpha + \beta_1 log(labor)_{it} + \beta_2 log(publiccapital)_{it} \\ + \beta_3 log(privatecapital)_{it} + dummy(1991-1993) \\ + trend + \varepsilon_{it}. \end{aligned} \tag{9}$$

On ε_{it} we assume following error structure:

$$\varepsilon_{it} = \mu_i + v_{it} \tag{10}$$

i.e. we assume that error process is one-way. Here, the disturbance term v_{it} is assumed to be *i.i.d.*

We also include a dummy variable to account for the severe economic downturn in 1991-1993. As a reference point the equation (9) is estimated also only for the private sector and without cross-section specific constants. Table 4 reports the results. In a way, second and third equation serve as a benchmark cases estimated in the most of previous studies. In addition, we also include a variable that combines public capital in each region to public capital in neighboring regions. This variable is supposed to take spill-over effects into account and it is constructed similarly as in Mas *et al.* (1996).

The value of the F -test for the joint significance of sub-region dummy coefficient indicates that sub-regions dummies are jointly significant. Thus, normal cross-sectional or pooled estimation would suffer from omitted-variables bias.

Basic production function estimates (table 4) seem economically reasonable in sign and size. Estimates for private inputs are highly statistically significant also when both capital stocks are included. Public capital gets an estimate of 0.09, which is well in line with previous results from regional studies done for European countries. The combined variable is statistically significant and the estimate is larger than the estimate for public capital solely. This suggests that there are some spill-over effects present. As results from joint significance test already point out, fixed effects specification is necessary in regional setup. In table 4 White period method is used to correct standard errors for serial correlation. Overall, results look quite similar to previous studies.

The general problem in production function estimation is the possible endogeneity of regressors. To account for this, and the possible cointegration between dependent and some or all explanatory variables, we use the panel dynamic OLS estimator developed by Mark and Sul (2003). Mark and Sul's estimator accounts for cross-sectional

Table 4: OLS production function estimates

Dependent variable: log(value added)				
variable	Pooled	One-way FE	One-way FE	One-way FE
constant	-3.489****	-	-	-
	(0.1760)			
log(labor)	0.5413****	0.7529***	0.7675***	0.7800***
	(0.0560)	(0.0520)	(0.0519)	(0.0473)
log(private capital)	0.5721****	0.2789***	0.2355***	0.2192**
	(0.0626)	(0.0642)	(0.0665)	(0.0751)
log(public capital)	-0.0144	-	0.0924*	-
	(0.0287)		(0.0406)	
log(combined)	-	-	-	0.1533**
				(0.0581)
dummy	-0.0686****	-0.025**	-0.031***	-0.035***
(1991-1993)	(0.0112)	(0.009)	(0.009)	(0.008)
trend	0.0189****	0.0259***	0.0249***	0.0242
	(0.0018)	(0.0016)	(0.0016)	(0.0016)
joint	-	41.01	41.59	41.59
significance		(<0.0001)	(<0.0001)	(<0.0001)
sub-regions	76	76	76	76
years	29	29	29	29
observations	2204	2204	2204	2204

* = p<.05, ** = p<.01, *** = p<.001. Standard errors are presented in parentheses except in the test for joint significance where it denotes the p-value of rejecting H_o. Joint significance gives the value of F -statistics for the test of equal sub-regional dummy coefficients. Standard errors are corrected for serial correlation using White period method.

dependence by introducing a common time effect. Wagner and Hlouskova (2007) have compared the performance of different estimators for panel cointegrated data. They found that DOLS system estimator (panel DOLS) performs robustly in the presence of cross-unit correlation or cointegration compared to several other estimators developed for panel cointegrated data.[9]

Mark and Sul's (2003) estimator assumes that observations on each individual i obey the following triangular representation

$$y_{it} = \alpha_i + \lambda_i t + \theta_t + \gamma' x_{it} + u_{it}, \tag{11}$$

where $(1, -\gamma')$ is a cointegrating vector between y_{it} and x_{it}, which is identical across individuals, α_i is an individual-specific effect, $\lambda_i t$ is a individual-specific linear trend, θ_t is a common time-specific factor, and u_{it} is a idionsyncratic error that is independent across i, but possibly dependent across t. The model (11) allows for a limited form of cross-sectional correlation where the equilibrium error for each individual is driven in part by θ_t.

Panel DOLS eliminates the possible endogeneity between explanatory variables and dependent variable by assuming that u_{it} is correlated at most with p_i leads and lags of $\triangle x_{it}$. This endogeneity can be controlled for by projecting u_{it} onto these p_i leads

[9]The tested estimators included FM-OLS presented by Phillips and Moon (1999), DOLS presented by Kao and Chiang (2000) and Mark and Sul (2003), one-step VAR, and two-step VAR presented by Breitung (2005)

and lags:

$$u_{it} = \sum_{s=-p_i}^{p_i} \delta'_{i,s} \triangle x_{i,t-s} + u_{it}* = \delta'_i z_{it} + u_{it}* \tag{12}$$

The projection error $u_{it}*$ is orthogonal to all leads and lags of $\triangle x_{it}$ and the estimated equation becomes

$$y_{it} = \alpha_i + \lambda_{it} + \theta_t + \gamma' x_{it} + \delta'_i z_{it} + u_{it}*, \tag{13}$$

where $\delta'_i z_{it}$ is a vector of projection dimensions. The consistent estimation of (13) is based on sequential limits, i.e. as $T \to \infty$ then $n \to \infty$.

Previously, panel DOLS has been used in regional analysis only by Okubo (2007), who argued that panel DOLS eliminates the endogeneity bias in nonstationary and cointegrated panels. In the case of Japan, Okubo (2007) showed that the results of traditional LSDV-estimator do not hold, if the equation is estimated with panel DOLS. When panel DOLS was used, the negative elasticity estimate for public capital, a result generally found in previous studies, changed to positive.

Table 5 presents the results of dynamic OLS fixed-effects estimations of equation (9).[10] DOLS estimation uses leads and lags of 1 to account for possible correlation between equilibrium error and $\triangle x_{jt},\ j = 1, ..., n$. DOLS estimator uses Andrew and Mohanan's pre-whitening method to account for possible autocorrelation. DOLS estimations include individual constants and individual trends, but their values are not presented in table 5.

Table 5: Dynamic OLS production function estimates

Dependent variable: log(value added)		
variable	FE DOLS	FE DOLS
log(labor)	0.5464***	0.5705***
	(0.0510)	(0.0593)
log(private capital)	0.2477***	0.2140**
	(0.0506)	(0.0668)
log(public capital)	-0.0335	-
	(0.0470)	
log(combined)	-	0.0620
		(0.0800)
sub-regions	76	76
years	29	29
observations	2204	2204

* = p<.05, ** = p<.01, *** = p<.001. Standard errors are presented in parentheses. Standard errors are estimated using Andrews and Monahan's Pre-whitening method. Estimation includes individual constants and trends.

DOLS estimates suggest that the coefficient of public capital is negative and insignificant if possible spillover effects are not taken into account. The combined variable gets positive, but still statistically insignificant estimate. Elasticity estimates for private inputs remain statistically significant and reasonable in size. Nevertheless, the results of panel DOLS estimation on the effect of public capital differ substantially on

[10] Estimation is conducted with Gauss. Gauss code has been provided by Mark and Sul (2003).

OLS estimations. This implies that some or all of the explanatory variables appearing in the model may be endogenous and/or there are cointegrating relations between the dependent variable and some or all explanatory variables.

However, the results of panel DOLS estimation may have been affected by endogeneity if leads and lags of 1 have not been enough to remove the correlation between equilibrium error and first differenced explanatory variables.[11] It is also possible that the common time-effect included in panel DOLS estimation has not captured all the cross-sectional correlation present in the data. This is a problem especially, if there remains correlation between equilibrium error and leads and lags of other cross-sections $\triangle x_{jt}$, $j = 1, ..., n$. In this case the panel DOLS exhibits the same form of second order asymptotic bias as pooled OLS (Mark & Sul 2003). To account for this, panel DSUR estimator is used, which controls for the endogeneity between equilibrium errors and cross-equations (Mark *et al.* 2005). Panel DSUR estimates a long-run covariance matrix that is used in the estimation. This actually makes panel DSUR more efficient the more the cross-sections are correlated across the panel. Endogeneity is controlled by including leads and lags of first differenced explanatory variables into the regression as in panel DOLS estimator.

The drawback of panel DSUR is that estimation of the long-run covariance matrix requires large time series dimension compared to cross-sectional dimension (Mark *et al.* 2005). In our case, the panel can include up to 10 cross-sections.[12] As mentioned in the previous section, some of the Finnish sub-regions have grown progressively whilst some have stagnated. It is thus reasonable to analyze these two categories in our restricted estimation. To do this, we select 10 sub-regions that have increased their value added the most and 10 sub-regions that have increased their value added the least between 1976 and 2004. Differences in the growth rate between these two samples are quite large. 10 fastest growing sub-regions of Finland have grown with the annual rate of 3,8% in average while the slowest growing sub-regions have grown with the annual rate of 0,55% in average.

According to Levin *et al.* (2002) , Im *at al.* (2003), ADF, and PP panel unit root tests, the value added series is $I(1)$ in the sample of 10 fastest growing sub-regions. In the sample of 10 slowest growing sub-regions, all the traditional tests presented previously find the value added series to be $I(0)$. This implies that there would be some non-stationary and some stationary series of valued added in the panel. Table 6 presents the summary of the results of Pedroni's (2004) panel cointegration test for 10 fastest growing sub-regions.

According to the results presented in table 6, the series of value added and labor, and value added and public capital seem to be cointegrated. Only 3 out of 11 Pedroni's (2004) test statistics find the value added and private capital to be cointegrated. However, the results of table 6 needs to be interpreted cautiously because of size distortions in Pedroni's test with small dimensions on n and T (Banerjee *et al.* 2005).

Table 7 presents the results of panel DSUR estimation of equation (9).[13] Panel DSUR includes common time effects, individual constants, and individual trends. A parametric correction is used to account for possible autocorrelation. As a reference we have also estimated a simple random sample drawn from the remaining 56 sub-regions.

All elasticity estimates estimated by panel dynamic SUR are highly statistically

[11] We also estimate our model using leads and lags of 2 using only three explanatory variables, private and public capital and labor, and leads and lags of 3 using only public capital as explanatory variable. There were no major changes in the values or standard errors of parameter estimates of public capital.

[12] If the cross-sectional size is increased beyond this point, panel DSUR fails to converge.

[13] Estimation is conducted with Gauss. Gauss code was provided by Mark *et al.* (2005)

significant. Public capital gets an elasticity estimate of 0.11 in the sample of 10 fastest growing regions. Labor gets quite high elasticity estimate compared to the results of panel DOLS. However, the size of these estimates is still reasonable.

The results for the sample of 10 fastest growing regions are the most reliable from the panel econometric viewpoint. In this sample, all variables are nonstationary and cointegrated according to our tests. Thus, the possible problem of spurious regression, which may have been present in the previous estimations done with the whole data, disappears. In addition, panel dynamic SUR is not only consistent when regions are correlated or cointegrated with each other, but is also more efficient when this is the case.

When results of 10 fastest growing sub-regions are compared to FE-OLS estimates presented in table 5, they surprisingly seem to be somewhat in line with each other. Despite of this it should be remembered that OLS is *not asymptotically unbiased* estimator of panel cointegrated data (Kao & Chiang 2000). For comparison we have estimated the three groups presented above using fixed-effects OLS. Table 8 presents the FE-OLS results of estimation of equation (9) assuming one-way error process on the three groups explained above.

Although results of table 5 and 7 indicate, that controlling for the endogeneity by

Table 6: Summary of the results of Pedroni's panel cointegration tests for 10 sub-regions that have grown the fastest

Dependent variable: log(value added)	Number of test statistics that reject H_0 at 5% level
log(labor)	10/11
log(private capital)	3/11
log(public capital)	8/11
all variables	8/11

Null hypothesis is that there is no cointegration between the variables.

Table 7: Panel dynamic SUR production function estimates

Dependent variable: log(value added)			
variable	10 fastest	10 slowest	10 "normal"
log(labor)	0.9324***	0.6788***	0.9327***
	(0.0156)	(0.0253)	(0.0121)
log(private capital)	0.2017***	0.0763*	0.2807***
	(0.0146)	(0.0376)	(0.0227)
log(public capital)	0.1094***	0.1278***	0.1888***
	(0.0239)	(0.0219)	(0.0131)
sub-regions	10	10	10
years	29	29	29
observations	290	290	290

* = p<.05, ** = p<.01, *** = p<.001. Standard errors are presented in parentheses. Standard errors are calculated using parametric correction. Estimation includes individual constants and trends. 10 fastest includes sub-regions whose value added series have grown the fastest between 1976-2004. 10 slowest includes sub-regions whose value added series have grown the slowest between 1976-2004. 10 normal includes a simple random sample of 10 sub-regions on the remaining 56 sub-regions.

Table 8: OLS production function estimates II

Dependent variable: log(value added)			
variable	10 fastest	10 slowest	10 "normal"
log(labor)	0.6550***	0.7105***	0.5741***
	(0.0786)	(0.0525)	(0.0408)
log(private capital)	0.1531	-0.0557	0.4134***
	(0.1683)	(0.0395)	(0.0532)
log(public capital)	0.0392	0.5477***	-0.0436
	(0.2158)	(0.0913)	(0.0397)
dummy	-0.0832**	-0.0297**	-0.0462**
(1991-1993)	(0.030)	(0.0099)	(0.0147)
trend	0.0367***	0.0147***	0.0194***
	(0.0056)	(0.0019)	(0.0012)
sub-regions	10	10	10
years	29	29	29
observations	290	290	290

* = p<.05, ** = p<.01, *** = p<.001. Standard errors are presented in parentheses. Standard errors are corrected for serial correlation using White period method. Estimations is with fixed-effects. 10 fastest includes sub-regions whose value added series have grown the fastest between 1976-2004. 10 slowest includes sub-regions whose value added series have grown the slowest between 1976-2004. 10 normal includes a simple random sample of 10 sub-regions on the remaining 56 sub-regions.

removing unobserved sub-regional effects in OLS estimation would result to reliable estimates, results of table 8 tell a different story. It seems that the results presented in table 5 are just averages of different, and probably biased, parameter estimates. Only explanatory variables whose coefficients remain somewhat stable are labor, dummy, and trend. Parameter estimates of public and private capital experience wild swings from positive to negative. Thus, results of table 8 compared with the results of the Table 7 imply that one should be extra cautious, when using OLS estimation in panels that may include cointegrating relations between dependent and explanatory variables and/or cross-sectional correlation.

4 Conclusions

In this paper, we have focused on the econometric aspects of regional productivity analysis of public capital using panel data from Finnish regions. Our results imply that a national panel of sub-regions may include both stationary and nonstationary series of value added. In panel unit root testing, this may result to flawed conclusion that the whole panel would be stationary. This should be taken into account in estimation cause basic versions of OLS, and many traditional panel estimators, are either biased or inconsistent in panel cointegrated data. We also argue that regional panels are likely to suffer from strong cross-sectional correlation, which is likely to cause bias in traditional OLS estimation.

We have shown that the results may differ substantially, if we use panel dynamic SUR -estimator instead of fixed effects OLS. Panel dynamic SUR -estimator controls for endogeneity and is efficient when cross-sections of the panel are correlated. The results of panel DSUR suggest that the elasticity estimate of private production with respect to public capital is 0.11 in the sample of 10 fastest growing regions of Finland. Unfortunately, current data restricts the maximum sample size to 10 regions in panel DSUR estimation. If the same sub-sample is estimated with basic fixed effects OLS, results differ substantially and, for example, private capital is not statistically significant. Thus, it seems that the commonly used fixed effects OLS may lead to false conclusions in the production function setup with regional data and may be useful only as a reference point for other estimators.

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Data Appendix

General notices: The data consists of yearly observations from 77 Finnish sub-regional units (seutukunta in finnish)[14] in the period of 1975-2004.[15]

Variables are measured as constant prices at 2000 and the regional division corresponds to the situation in the year 2005 (see appendix). All industries are included.

All three capital stock series are measured at the end of the year. Therefore, for the year t we have used $t-1$ values of the capital stocks in the estimated production functions. Due to this correction our sample in estimation is 1976–2004.

Regional output Y: Private sector's regional production is measured as value-added at factor prices. The data is taken from official statistics complied by Statistics Finland after wide revision of National Account statistics finalized in the spring 2006.

Labor L : Private labor consists of number of workers. The number of hours would be better variable, but it is not available at sub-regional level prior to 1995. The data is taken from official statistics complied by Statistics Finland.

Private capital $K1$: Private capital is measured as private net capital stock. Net capital stocks are taken from Salmela (2008) and they are constructed using the current National Account standards.

Public capital $K2$: Public capital is measured as public net capital stock, which includes both central and local governments capital stocks. Net capital stocks are taken from Salmela (2008) and they are constructed using the current National Account standards. Public capital stock is used instead of some infrastructure capital measurement. The use of the whole public capital stock can be justified by the restrictions in the availability of more proper data. Recently, there have been been attempts also in Finland to construct variables for infrastructure capital (see Uimonen 2007,2008). Until now, this is done only for roads and railroads.

OECD's current recommendation considering the productivity studies is to use effective capital, which is a flow variable measured as a volume index of serviced provided by capital. These kind of variables are available in official statistics only in three countries (the United States, Canada and Australia).

Combined public capital and spill-over This variable is constructed for region i by adding up public capital in region i and public capital in neighboring regions. Neighbor region is defined similarly as above.

More detailed description of the data is available upon request.

[14]Sub-regional units do not enter in the NUTS (Nomenclature des Unités Territoriales Statistiques) classification, which became effective in 2003 as a European Union's regulation. They are one step lower than the NUTS level 3, which would be regions in the Finnish case.

[15]However, Porvoo is excluded due to data problems. For instance, value added drops 75 % from 1985 to 1986. Actually, this is not a data error. The development of petrochemical industry in Porvoo has been highly volatile and thus the use of the series in economic analysis is not meningful.

Appendix: Sub regional units

Finnish sub-regional units in 2005. Source: Statistics Finland (2005).

011	Helsinki	093	Imatra	152	Vaasa
012	Lohja	101	Mikkeli	153	Sydösterbottens region
013	Tammisaari	102	Juva	154	Jakobstadsregionen
021	Turunmaa	103	Savonlinna	161	Kaustinen
022	Salo	105	Pieksämaäki	162	Kokkola
023	Turku	111	Ylä-Savo	171	Oulu
024	Vakka-Suomi	112	Kuopio	173	Oulunkaari
025	Loimaa	113	Koillis-Savo	174	Raahe
041	Rauma	114	Varkaus	175	Siikalatva
043	Pori	115	Sisä-Savo	176	Nivala-Haapajärvi
044	Pohjois-Satakunta	122	Joensuu	177	Ylivieska
051	Hämeenlinna	124	Keski-Karjala	178	Koillismaa
052	Riihimäki	125	Pielisen Karjala	181	Kehys-Kainuu
053	Forssa	131	Jyväskylä	182	Kajaani
061	Luoteis-Pirkanmaa	132	Joutsa	191	Rovaniemi
062	Kaakkois-Pirkanmaa	133	Keuruu	192	Kemi-Tornio
063	Etelä-Pirkanmaa	134	Jämsä	193	Torniolaakso
064	Tampere	135	Äänekoski	194	Itä-Lappi
068	Lounais-Pirkanmaa	138	Saarijärvi-Viitasaari	196	Tunturi-Lappi
069	Ylä-Pirkanmaa	141	Suupohja	197	Pohjois-Lappi
071	Lahti	142	Seinäjoki	201	Porvoo
072	Heinola	143	Eteläiset seinä-naapurit	202	Loviisa
081	Kouvola	144	Kuusiokunnat	211	Mariehamns stad
082	Kotka-Hamina	145	Härmänmaa	212	Ålands landsbygd
091	Lappeenranta	146	Järviseutu	213	Ålands skärgård
092	Länsi-Saimaa	151	Kyrönmaa		

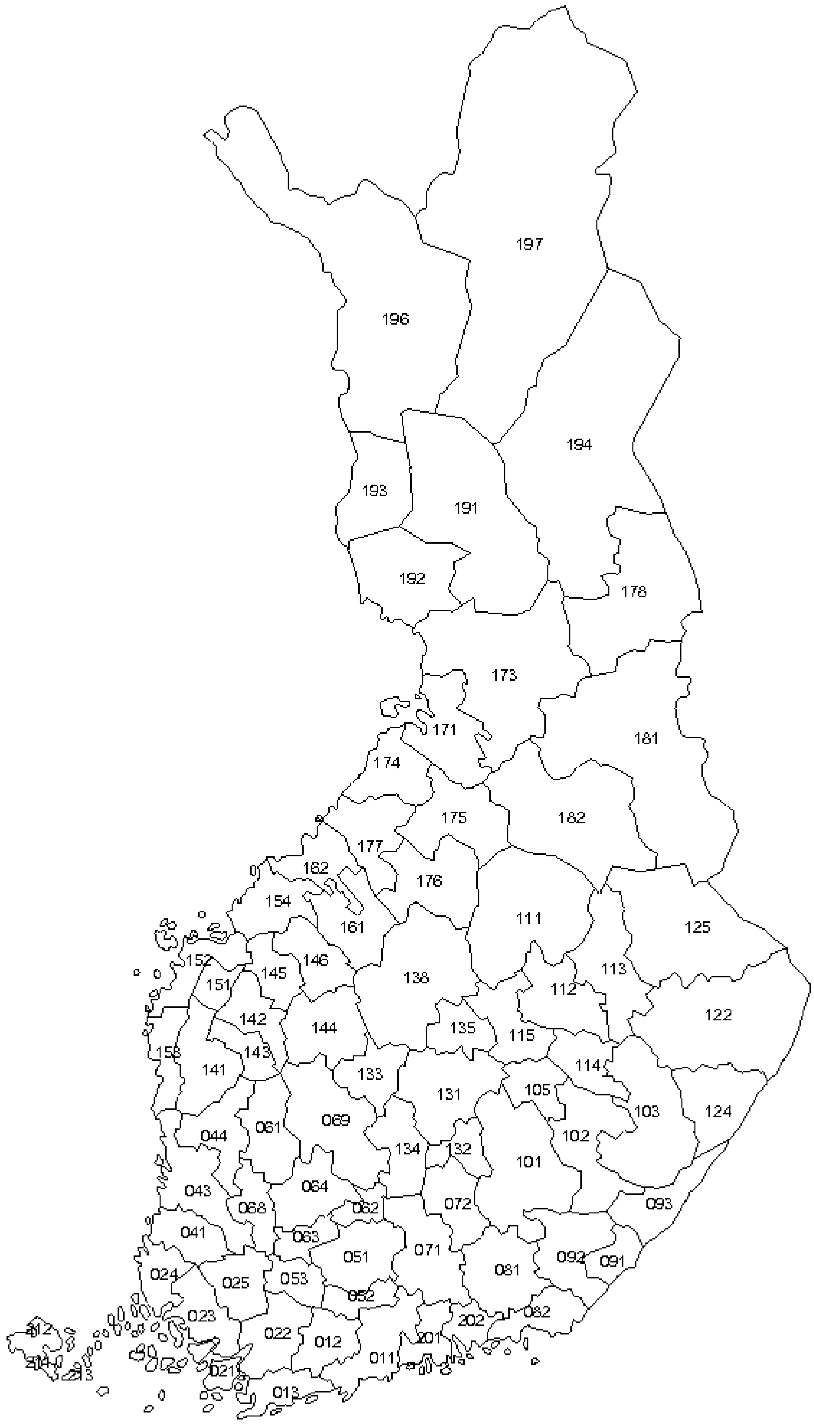

Figure 1: Map of Finnish sub-regional units in 2005. Source: Statistics Finland.