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And yet they Co-move! Public capital and productivity in OECD

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Abstract

In this paper, we add to the debate on the public capital–productivity link by applying very recent developments in the panel time series literature that take into account cross sectional correlation in non-stationary panels. In particular, we evaluate the productive effect of public capital by estimating various production functions on a panel of 21 OECD countries over the period 1975–2002. Our results suggest that public capital has a positive long run impact on output, with elasticities that range between 0.05 and 0.15, depending on model specification. These findings are robust to the existence of spillover effects from public capital investments in other countries and to the inclusion of other productivity determinants, like human capital, the stock of patents and R&D capital. Finally, we do not find any important effect of public capital on GDP in the short run: this suggests that public infrastructure investments might not be a powerful countercyclical policy instrument.

JEL classification: C33; C15; H54; O47

Keywords: Public capital; Productivity; Panel cointegration; Cross-section dependence

1. Introduction

The role of public expenditure as a countercyclical economic policy instrument has been the object of a lively debate among both academics and policymakers, at least since the burst of the

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2008–9 recession and the announcement of the fiscal package stimulus by the Obama's administration. In particular, the U.S. Congress approved in 2009 the $787 billion American Recovery and Reinvestment Act, whereof approximately two-thirds amounted to direct government expenditure and transfers. Most of the recent theoretical and applied macroeconomic literature focuses on quantifying the economic impact of the fiscal stimulus and, more generally, on estimating the magnitude of the fiscal multiplier (see Hall (2009) among the others).

However, a large fraction of the Obama's fiscal package (approximately $130 billion) has been devoted to infrastructure expenditure, which not only may be used as a countercyclical tool, but it also might have a more lasting long run effect on the productive potential of an economy: this is the issue we focus on in this study.

Since the Aschauer (1989)'s seminal paper, several contributions have highlighted that public infrastructures are important inputs that contribute to economic growth. Improvements in public infrastructures (e.g. better and more extensive transport networks) might impact TFP in a number of ways, e.g. by increasing the productivity of private inputs like physical capital and R&D or by reducing production and transport costs, thereby fostering greater specialization, more intensive competition and in general by providing those public goods that are crucial for economic growth.[1]

The relevant empirical literature has developed along a number of strands according to differences in the type of sample, theoretical approach and econometric methodology. Most studies estimated production functions, while others relied on the estimation of cost functions where public capital is assumed to be a quasi-fixed input (Cohen and Morrison (2004)); in turn, some authors included public investment as an additional explanatory variable in growth convergence equations (Esfahani and Ramirez (2003)). As far as the sample choice is concerned, most contributions are based on aggregate data at either country (Pereira and Roca-Sagales (2001) and Pina and Aubyn (2005)) or regional level (Bronzini and Piselli (2009)), with a minority focusing on industry level data (Bottasso and Conti (2010)) or cross country data (Canning and Pedroni (2008)). Turning to the econometric methodology, recent studies on aggregate (single) country data adopted VAR techniques, which investigate the relationship between public capital, GDP and private inputs without imposing a theoretical structure, and generally found positive effects of public capital on GDP; however, purely time series studies are often plagued by small sample problems linked to the short time span of the data. For this reason, many authors have turned to conducting studies based on cross country or regional level data: while "first generation" panel studies simply estimated either fixed or random effects models, "second generation" studies tackled the endogeneity problems that plague the estimation of production functions more seriously by using instrumental variable techniques, such as the Arellano-Bond GMM estimator. Only recently issues stemming from the non-stationarity nature of panel data have been addressed by some authors in order to avoid possible biases associated to the presence of unit roots. However, these studies do not account for unobservable time varying heterogeneity associated to unobserved common shocks which might affect each country or region to a different extent, thus generating cross sectional correlation: this is likely to be the case when analyzing macro panel data, where cross section dependence can be due to a variety of factors, such as omitted unobserved common factors, spatial spillover effects, trade linkages, global economic cycles, etc. Indeed some authors (e.g. Calderon, Moral-Benito, and Serven (2011)) have addressed cross sectional correlation by removing unobserved common factors through a demeaning of the variables: this procedure works insofar as

[1] Other papers have investigated the impact of public capital on employment and labor market outcomes (see, among the others, Everaert & Heylen, 2004)).

unobserved common factors have the same impact on individual countries, which is however a very restrictive assumption. The presence of cross-sectional dependence may affect the validity of commonly used panel unit root and cointegration tests, since stationarity and no-cointegration tests that assume independence might have substantial size distortions when this assumption does not hold. In particular, Banerjee, Marcellino, and Osbat (2004) show that neglecting the cross-section dependence arising from a common factor structure may have quite drastic effects on cointegration testing, while Urbain and Westerlund (2008) find that the presence of cross-sectional dependence is likely to lead to substantial bias for various pooled estimators. To date, only few studies apply cointegration analysis techniques which account for cross sectional dependence and there does not yet seem to exist a consensus about successful modeling strategies in such framework.

In this paper we make use of some recent developments in the panel time-series literature and we assume that cross sectional dependence can be successfully modeled within the framework proposed by the PANIC representation of Bai and Ng (2004) who adopt a common factor structure for the series investigated. This assumption is maintained both in the analysis of the statistical properties of the data and in the cointegration analysis which is based on Gengenbach, Palm, and Urbain (2006). This kind of analysis has never been conducted before and adds to the literature on the productive impact of public capital by providing more accurate and robust estimates of the long run elasticity of GDP with respect to public capital for the most important OECD countries.

In this study we consider a panel of 21 OECD countries observed over the period 1975–2002 and we estimate different production functions in order to investigate the short and long run relationship between public capital and GDP, also taking into account the role played by other important productivity determinants such as human capital and innovative activity. We focus on a large sample of the richest OECD countries for two reasons: first, we want to understand the productive effect of public capital investment in the case of high income countries, given the revival of interest in public infrastructure investments among policymakers in both the US and the EU; secondly, by considering countries with similar institutions and levels of development, we should alleviate concerns of parameter heterogeneity in the production function, although still allowing for heterogeneity in productivity levels and growth rates as well as in the effects of unobserved common shocks.

Integration and cointegration analysis show that the non-stationarity of the variables entering our production function is entirely due to unobserved common components and that our series are cointegrated along the cross-sectional dimension. In order to tackle estimation and inference issues we apply the Continuously Updated Estimator suggested by Bai, Kao, and Ng (2009), which is an extension of the two stage fully modified estimator proposed by Bai and Kao (2006) for the case of non-stationarity of the unobserved common components and in presence of cross-member cointegration. Overall results suggest that the long run elasticity of GDP with respect to public capital ranges between 0.05 and 0.15, depending on model specification: this result is in line with the past literature which found average public capital elasticities of about 0.1–0.2, as shown in the meta-analysis conducted by Bom and Ligthart (2009). Another interesting result is provided by Granger causality tests which suggest that public capital and the stock of patents might Granger cause GDP while the opposite does not seem to occur. These findings, taken together, have important policy implications as they suggest that public capital investments can be used by policy makers to affect a country's productive potential in the long run; however, since such result is not confirmed for the short run, public capital investment might be a poor countercyclical policy instrument. Secondly, we provide some weak evidence of the existence of possible spillover effects associated to public capital stocks in neighboring countries. Such result suggests that public infrastructures might be underprovided given that individual countries do not internalize the full benefits arising from such investments.

The remainder of the paper is organized as follows. In Section 2 we discuss our empirical model, while Section 3 describes the data. The statistical properties of time series as well as the cointegration analysis can be found in Sections 4 and 5 describes long run analysis, while Section 6 is devoted to analyzing short run dynamics. Finally, Section 7 concludes.

2. The empirical model

In this paper we estimate the impact of public capital on productivity adopting a production function framework. In particular, we assume that GDP is produced according to the following Cobb-Douglas technology:

$$Y_{it} = TFP_{it} K_{it}^{\alpha} (HL)_{it}^{\beta} \tag{1}$$

where Y_{it} is the GDP in country i at time t; K_{it} and HL_{it} are the associated private capital stock and human capital augmented labor and TFP_{it} represents total factor productivity while α and β are the output elasticity of private capital and labor, respectively. The economic literature has identified many possible determinants of TFP, with the firms' innovative activity being as one of the most important. Hence, TFP can be represented by the following equation:

$$TFP_{it} = G_{it}^{\upsilon} P_{it}^{\eta} \exp(u_{it}) \tag{2}$$

where G_{it} and P_{it} represent the public capital stock and the stock of patents, respectively and u_{it} is an error term accounting for other determinants of productivity.

It should be noted that controlling for human capital and innovation can be important (even if it has rarely been done before: a recent exception is Bronzini and Piselli (2009) if we remember the role played by knowledge in new growth theories. Moreover, our data suggest that countries with a high public capital to output ratio tend also to have a highly educated population and a large stock of patents. Substituting Eq. (2) into (1) and taking logs, we get (where lower case variables denote natural logs):

$$y_{it} = \alpha k_{it} + \beta hl_{it} + \upsilon g_{it} + \eta p_{it} + u_{it} \tag{3}$$

Turning to the error term u_{it}, we can decompose it as follows:

$$u_{it} = e_i + a_i \text{trend}_t + \lambda_i F_t \tag{4}$$

where e_i is a country fixed effect accounting for persistent differences in TFP levels across countries; trend_t is a set of country specific trends (with associated parameters a_i) which account for unobserved shocks that drive the evolution of each country's TFP linearly and, finally, $\lambda_i F_t$ is a vector of (possibly non-stationary) unobserved common factors with country specific factor loadings λ_i that proxy for global macroeconomic shocks to TFP, like global changes in economic policy, oil shocks, or spatial spillovers that may generate cross sectional correlation in the data. It is important to highlight that the assumption of country specific factor loadings allows for a differential impact of the same global shock on TFP across countries.

In order to verify and quantify the existence of spatial spillover effects associated to public capital, in some regression specifications we augment equation with the stock of public capital in other countries (goth_{it}). Finally, we can note that in equation we are not imposing constant returns to scale neither for private nor for all inputs and we do not make any assumption of perfect

Table 1
Descriptive statistics.

Variable	Year	
	1975	2002
GDP	522.6	1092
	(88.3)	(203)
K	1171.6	2582
	(1807)	(4363)
G	324.4	633
	(61.3)	(1237.4)
H	8.76	11.06
	(1.88)	(1.76)
P	49,676	65,477
	(164,613)	(468,793)
G/Y	0.55	0.50
	(0.21)	(0.20)
L	27,258	32,789
	(40,161)	(56,399)

Note: St. Dev in parenthesis. Y, K, G, billions of $; H, number of years; P, number of patents (in 1975 and 1999); L, number of hours.

competition. However, for robustness check, we have also estimated a version of equation after imposing constant returns to scale in private inputs: this assumption in turn, together with that of perfect competition in input and output markets, allows us to use income shares as proxies for private inputs' elasticities and therefore to reformulate Eq. (3) as:

$$tfp_{it} = \upsilon g_{it} + \eta p_{it} + v_{it} \tag{5}$$

where tfp_{it} was computed residually as $y_{it} - \alpha k_{it} + (1 - \alpha) l_{it}$ assuming a constant capital share of one third. The estimation of Eqs. (3)–(5) – raises a set of significant econometric challenges which we address in the following sections.

3. Data

The data employed in this study are derived from different sources and are referred to 21 OECD countries, namely Austria, Belgium, Denmark, Spain, France, Italy, UK, The Netherlands, Sweden, Finland, Germany, Austria, Greece, Canada, Japan, New Zealand, Norway, USA, Switzerland, Portugal and Ireland observed over the period 1975–2002.

Output Y is taken from the OECD Analytical Database and is defined as GDP at 1995 constant prices, converted in purchasing power parities using OECD PPP exchange rates. The labor input L is defined as the total annual hours worked and it is sourced from the total Economy Database of the University of Groningen. Private (K) and public (G) capital stocks are taken from Kamps (2004) to whom we refer for details. Summary statistics reported in Table 1 suggest that in both 1975 and 2002 the countries with the highest public capital to output ratio were New Zealand, Japan and The Netherlands; in turn, Portugal, Belgium and Spain were the countries with the lowest ratio in 1975 and Ireland, Belgium and Canada those with the lowest ratio in 2002. In general, the cross country average slightly fell over time from about 0.56 to about 0.5 possibly because of a decline in government investment or because of the privatization process that reduced the scope of government intervention in some areas of the economy especially after 1995.

Table 1 reports summary statistics for human capital stock, which is proxied by the average number of schooling years for the population aged 25 or more (S) and it is taken from Cohen and Soto (2007) to whom we refer for details. As these data have only been computed at ten year intervals since 1960–2010, we have derived information for the missing years by linear interpolation. We then have followed Bils and Klenow (2001) in order to build a human capital augmented labor input as $HL_{it} = L_{it} * \exp[f(mq_i) * S_{it}]$, where $f(mq_i)$ represents a country specific piecewise concave function of the mincerian return to education of one additional schooling year corrected for the quality of the educational system, taken from Cohen and Soto.

The stock of patents P has been computed as in Bottazzi and Peri (2007) by accumulating past patents using the perpetual inventory method. In particular, we have taken information on the number of patents from the University of Groningen Patent Database[2] and we refer for an accurate description of the procedure we have followed to the Working Paper version of this study. Data reported in Table 1 confirm both the substantial differences in the stock of patents even between countries with similar levels of GDP, but also the surge in the stock of patents granted by the USPTO that occurred between 1975 and 1999, with the exceptions of the US and the UK. In some model specifications we will need an estimate of the stock of public infrastructure of the other countries in the sample in order to capture the spatial spillovers associated to the infrastructure activity carried out in the rest of the world. The stock of public capital of the rest of the world for country i in year t is defined as $GOTH_{it} = \sum_{j \neq i} G_{jt} W_{ij}$, where the row standardized weight matrix W_{ij} is based on the inverse of distance between the capital of country i and that of each of the other 20 countries in the sample.

As a robustness check we have used physical indicators as proxies for the stock of public infrastructure, namely the Km of motorways and railways lines taken from the Database of World Infrastructure, Eurostat and country level sources. Finally, in some specifications we need the stock of R&D capital which as been calculated from the EUKLEMS database and from OECD data (details on the construction of the variable are available in the Working Paper version of this study).

4. Statistical properties of time series

4.1. Integration analysis

In this section we conduct a thorough investigation of stationarity properties of our data by applying recent panel data unit root tests. The most commonly used panel unit roots tests are those proposed by Levin, Lin, and Chu (2002) and Im, Pesaran, and Shin (2003), which however have been found to poorly perform in the presence of cross sectional dependence and when the number of cross section increases Larsson and Lyhagen (2000b). For this reason we prefer to employ the PANIC unit root tests proposed by Bai and Ng (2004) which take into account the presence of cross sectional dependence in the data.[3] In particular, the PANIC framework assumes the following factor structure for observed panel data:

$$Y_{it} = \lambda_i' F_t + e_{it} \tag{6}$$

[2] For the patent stock variable we lose information on New Zealand and Switzerland. Moreover, the period covered is 1975–1999.

[3] Nevertheless, we have performed the Levin et al. (2002) and Im et al. (2003) tests and we have found that for our variables in no case we can reject the null of non-stationarity at standard levels of significance; in turn first differences of time series resulted to be stationary.

Table 2
Bai and Ng (2004) unit root tests.

Variable[a]	# of factors	Z_e^τ	ADF^τ	MQ_f^τ	# of factors[*]
GDP	1	5.96	−2.81		
LHC	1	6.2	−3.15		
K	1	8.19	−3.72		
G	1	6.89	−1.97		
P	2	7.91		−8.96	2

Note: The suffices τ for the statistics Z, ADF and MQ_f indicate the intercept and liner trend case.
The PC3 criterion of Bai and Ng (2002) was used to estimate the number of unobserved common factors.
[*] Number of factors estimated by the MQ_f statistic.
[a] The abbreviation for the variables are presented in Section 3.

where F_t is a $(k \times 1)$ vector of common factors, λ_i is a vector of factor loadings and e_{it} is an idiosyncratic error component. The series may be non-stationary if either F_t or e_{it} (or both) are non-stationary and each hypothesis can be separately tested. In order to test for individual unit roots on the idiosyncratic component, e_{it}, the authors propose pooled tests (Z_e) for the hypothesis that all e_{it} are non-stationary, which are based on Fisher-type statistics and converge to a standard normal distribution for $(N, T) \rightarrow \infty$. As far as the presence of unit roots in the common component is concerned, Bai and Ng (2004) suggest the following strategy: if a single common factor is estimated, it can be applied an ADF test whose limiting distribution coincides with the Dickey-Fuller distribution; if more than one common factor is estimated, authors provide an iterative procedure to select the number of independent stochastic trends, which is similar to the Johansen trace test for cointegration. They suggest two modified statistics, MQ_d and MQ_f, where the former uses a non-parametric correction to account for additional serial correlation, while the latter employs a parametric correction. In Table 2 below we present the results for the MQ_f statistic only, but the results are robust to the application of the MQ_d test whose results are not reported for reasons of space.

The Z_e tests reject the null of unit root for the estimated idiosyncratic components for all the analyzed variables so that it becomes important to verify if possible non-stationarity of the series rest in the unobserved common components, i.e. if the non-stationarity in the observed data is due to a pervasive source. When estimating a single common factor for Human capital augmented labor, Public capital, Private capital and GDP, the *ADF* tests does not reject the unit root hypothesis except for the Private capital (the critical value ADF^τ is −3.41); estimating 2 common factors for Patents, both MQ_f^c and MQ_f^τ cannot reject the null hypothesis that there are 2 independent stochastic trends.[4] The critical values for the statistic is −31.356 (Bai & Ng, 2004). The remainder of our analysis proceeds on the assumption, supported by the tests performed above, that most log level variables are $I(1)$ processes, while all log differenced variables follow stationary, $I(0)$, processes.[5]

[4] We use the BIC3 criterion of Bai and Ng (2002) to estimate the number of unobserved common factors and we allow for at most 6 factors. Since the cross section and time series dimensions of the panel are approximately of the same magnitude, the BIC3 criterion tends to be superior over the alternatives. However, the results shown are robust to using other selecting criterions and selecting a different maximum numbers of allowed common factors.

[5] As a robustness check we have also performed the Pesaran (2007) panel unit root test, which is robust to the presence of cross sectional dependence, and our results are confirmed.

4.2. Cointegration analysis

In this section we describe the cointegration analysis conducted through different econometric approaches. First, we apply the panel cointegration tests derived by Pedroni (1999) and Pedroni (2004), who proposed seven different statistics for testing the presence of a single cointegration relationship under the assumption of cross-sectional independence.

Results support the hypothesis of a cointegration relationship between our variables, since four out of seven tests reject the null hypothesis of no cointegration at the 10% significance level. The statistics that fail to reject the no cointegration hypothesis are however undersized in small panels (Pedroni, 2004). As an additional test for cointegration, we apply the LR-bar and the PC-bar tests proposed by Larsson, Lyhagen, and Lothgren (2001) and Larsson and Lyhagen (2000a), respectively. The LR-bar statistic suggests that it does exist a common cointegration rank in the panel, or at least a common largest rank of 2, while the PC-bar test rejects a minimum cointegrating rank of 2: hence we cannot infer that there is a common cointegrating rank for all countries in the panel.

However the panel multivariate cointegration methods proposed by Larsson and Lyhagen (2000a), Larsson et al. (2001) and by Pedroni (1999) do not take into account the presence of cross sectional dependence; in particular Gengenbach et al. (2006) demonstrate that the panel cointegration tests proposed by Pedroni (1999) are inconsistent when the data present a common factor structure and show a consistent size distortion which increases with the cross-sectional dimension N.

In the presence of cross sectional dependence, cross-unit cointegration might arise. In the case of cross sectional cointegration, standard panel multivariate cointegration analysis might provide misleading results and might fail to detect any cointegration relationship among data. In order to take into account the possible existence of cross sectional cointegration, we follow the approach proposed by Gengenbach et al. (2006) which focus on testing for no-cointegration when the cross-sectional dependence in the panel is modeled with the PANIC approach of Bai and Ng (2004). In particular, Gengenbach et al. (2006) addresses the issue of no-cointegration within three different possible frameworks: (1) testing for idiosyncratic components no-cointegration when the observed non-stationarity in the series originates from idiosyncratic stochastic trends only, (2) testing for common factors no-cointegration when the non-stationarity is due to cross-sectional common trends only, (3) testing for panel no cointegration when there are both cross-sectional common and idiosyncratic stochastic trends. As discussed in section, the integration analysis has shown that the non-stationarity in our panel is entirely due to a reduced numbers of common stochastic trends: in this case cointegration between the dependent variable and the regressors can only occur if the common factors for Y_{it} cointegrate with those of X_{it}. Hence, we have to test for common factor no-cointegration (case 2 listed above). In this case Gengenbach et al. (2006) suggests to test the null of no-cointegration between the factors using the Johansen likelihood ratio test.

Table 3 presents the results of the Johansen trace test for cointegration between the six estimated common factors. Results suggest the existence of a single cointegrating relationship, which in turn allows us to interpret the long run relation in equation as a conventional production function. To the best of our knowledge this is the first paper seeking to estimate the productive effect of public capital in a panel non-stationary environment that tests for the number of cointegrating vectors, as previous studies simply assumed the existence of a unique contegrating relationship.

Table 3
Gengenbach et al. (2006) cointegration test.

Trace test statistics	Critical value	Cointegration rank
102.77	94.15	0
55.36**	68.52	1
34.93	47.21	2
23.03	29.68	3
11.64	15.41	4
4.73	3.76	5

* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

5. Long run analysis

In the previous analysis we have found evidence that the variables entering the production function are non-stationary and cointegrated. In this section we discuss parameter estimates of the augmented production function presented in Eq. (3). It is well known that estimating it by OLS is not appropriate if regressors are endogenous and residuals are serially correlated: in such a case the estimator is inefficient and the bias in the cointegration parameters is of order T. As a matter of fact, the estimation of production functions is plagued by the risk of bias due to endogeneity because inputs and outputs are jointly determined; moreover, the presence of cross sectional dependence needs to be properly taken into account.

In order to tackle the econometric issues raised by simultaneity and cross sectional dependence we apply two different techniques, namely the Dynamic OLS (DOLS) proposed by Mark and Sul (2001), which corrects for the possible endogeneity of the non-stationary regressors but does not take into account the cross-equation dependence in the equilibrium errors, and the Continuously Updated estimator introduced by Bai et al. (2009), which accounts for the presence of cross sectional dependence in the data.

5.1. Econometric issues

The estimation of the Eq. (3) by DOLS involves adding past and future values of the first differences of the explanatory variables as additional regressors, so that all nuisance parameters, which represent short run dynamics, are $I(0)$ and uncorrelated with the error term (by construction). This procedure corrects for the possible endogeneity of the non-stationary regressors; however, this estimator sacrifices asymptotic efficiency because it does not take into account the cross-section dependence; furthermore, it may fail to precisely identify parameter estimates in the presence of important spillover effects. Nevertheless we apply DOLS technique for robustness results.

Given the presence of cross sectional dependence in our data we also apply the estimator proposed by Bai et al. (2009) who consider the problem of estimating the cointegrating vector for a panel with unobserved non-stationary common factors. The authors consider the following model,

$$y_{it} = x_{it}^{'}\beta + \lambda_i^{'}F_t + \varepsilon_{it} \tag{7}$$

when F_t is a (rx1) vector of non-stationary unobserved common factors and x is a vector of regressors possibly including country trends and country specific fixed effects.

Bai et al. (2009) propose the CupBC (continuously-updated and bias-corrected) and the CupFM (continuously-updated and fully-modified) estimators for β. Both estimators are asymptotically unbiased and normally distributed and are valid when there are mixed stationary and non-stationary factors, as well as when the factors are all stationary. The same authors propose an iterative solution in the same line of that proposed by Bai (2009) and Bai and Kao (2006). The CUP estimators of Bai et al. (2009) minimize the following concentrated least square function:

$$CLS(\beta, F) = \sum_{i=1}^{N}\sum_{t=1}^{T}\frac{1}{NT^2}(y_{it} - x_{it}'\beta + \lambda_i' F_t)^2$$

where the function has been already minimized over λ_i and F_t, treated as parameters. λ_i and F_t are subject to the following identification constraints: $T^{-2}\sum_{t=1}^{T} F_t F_t' = I$ and $\Lambda'\Lambda$ is positive definite where $\Lambda = (\lambda_1', \cdots, \lambda_N')'$. The continuous updated estimator ($\hat{\beta}_{CUP}, \hat{F}_{CUP}$) is the solution to the following two nonlinear equations:

$$\hat{\beta} = \left(\sum_{i=1}^{N} x_i' M_{\hat{F}} x_i'\right)^{-1} \sum_{i=1}^{N} (x_i' M_{\hat{F}} y_i) \tag{8}$$

$$\hat{F} V_{NT} = \left[\frac{1}{NT^2}\sum_{i=1}^{N}(y_i - x_i\hat{\beta})(y_i - x_i\hat{\beta})'\right]\hat{F} \tag{9}$$

where x_i is a $(T \times k)$ matrix of regressors, y_i is a $(T \times 1)$ vector of dependent variables and V_{NT} is the diagonal matrix of the r largest eigenvalues of the matrix inside the brackets, arranged in decreasing order. The estimator is obtained by iteratively solving for $\hat{\beta}$ and $\hat{F}$ using (8) and (9). An estimate of Λ can be obtained as: $\hat{\Lambda} = T^{-2}\hat{F}'(Y - X\hat{\beta})$.

While the CUP estimator of β is consistent, there is an asymptotic bias arising from endogeneity and serial correlation, and thus the limiting distribution is not centered around zero. Bai et al. (2009) consider two fully-modified estimators which correct the asymptotic bias. The first one, the CupBC estimator, does the bias correction only once, at the final stage of the iteration while the second one, the CupFM estimator, corrects the bias at every iteration. While the CupFM estimator is computationally more costly it may have better finite sample properties.

5.2. Estimation results

In Table 4 we report empirical estimates of equation obtained with both DOLS and the CUP-FM estimator of Bai et al. (2009). In both cases, we allow for country-specific fixed effects and time trends while, in the case of the CUP-FM estimator, we also allow for a set of unrestricted common factors with heterogenous factor loadings. Estimates show that all parameters are statistically significant at the 1% level, with very similar values across estimation methods, with the notable exception of private capital elasticity. Moreover, estimates of the elasticity of public capital are similar with a value of about 0.13, well in line with previous empirical evidence. For instance, Calderon et al. (2011) employed the Pesaran and Smith (1995)'s pooled mean group estimator to a large panel of countries and obtained output elasticities of infrastructure in the range 0.07–0.10 while Canning and Bennathan (2000) find output elasticities of paved roads in the range 0.05–0.08 for a panel of world countries. Finally, the elasticity of GDP with respect to patents turns out to be about 0.10, i.e. in the lower range identified by Madsen (2007) who estimated with DOLS the long run cointegrating relationship between TFP and the domestic stock of patents for a sample

Table 4
Estimates of the long run cointegration relation.

	CupFM	DOLS
LH	0.168***	0.188***
K	(0.019)	(0.057)
	0.246***	0.748***
	(0.038)	(0.036)
G	0.129***	0.127***
	(0.036)	(0.032)
P	0.092***	0.108***
	(0.023)	(0.024)
Country trends	X	X
Country fixed effects	X	X

Note: sample includes 19 countries for the period 1975–2000. Standard errors in parenthesis
* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

of 16 OECD countries observed over the period 1870–2004; moreover, they are fully in line with the results of Coe, Helpman, and Hoffmaister (2009) who in fact found an elasticity of about 0.10 between a country's TFP and the stock of domestic R&D capital for a panel of OECD countries observed over the period 1970–2004.

In Table 5 we probe the results of our baseline specification reported in Table 4 along a number of ways. In column 1 we run a regression with TFP as the dependent variable and empirical results

Table 5
Estimates of the long run cointegration relation. CupFM estimator Bai et al. (2009).

Dep var	TFP	GDP	GDP	GDP	GDP	GDP	GDP
LH	–	0.199***	0.09***	0.17***	–	–	–
		(0.017)	(0.015)	(0.02)			
L	–	–	–	–	0.57***	0.58***	0.40***
					(0.029)	(0.027)	(0.029)
G	0.12***	0.15***	0.15***	–	0.06***	0.09***	0.13***
	(0.028)	(0.029)	(0.028)		(0.024)	(0.026)	(0.029)
K	–	0.38***	0.34***	0.71***	0.62***	0.34***	0.63***
		(0.032)	(0.037)	(0.029)	(0.029)	(0.027)	(0.03)
P	0.21***	–	–	–	–	–	0.05***
	(0.017)						(0.015)
R&D	–	0.05***	–	0.03***	–	0.09***	–
		(0.011)		(0.011)		(0.009)	
GOTH	–	–	0.40***	–	–	–	–
			(0.064)				
CONG POP	–	–	–	0.05*	–	–	–
				(0.029)			
Country fix. effects	X	X	X	X	X	X	X
Country trends	X	X	X	–	X	X	–

Note: sample includes 19 countries for the period 1975–2000. Standard errors in parenthesis.
* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

show that both public capital and the stock of patents have a positive and statistically significant coefficient: in particular, the elasticity of the former is barely altered with respect to that reported in Table 4, while the elasticity of the stock of patents about doubles, although remaining within the range of estimates one can find in Madsen (2007), Coe et al. (2009) and in the short literature review contained in Eberhardt and Teal (2011) who report, for studies conducted on panel of countries, elasticities of TFP with respect to R&D capital between 0.05 and 0.23.

In column 2 we use the stock of R&D as a measure of a country knowledge capital. As we can see, there are no major changes with respect to our baseline specification: in particular, R&D coefficient is positive and significant, although quite low.

Given the possibility that public infrastructures can generate spillovers related to network effects mainly associated to transport infrastructures, in column 3 we include in our regression specification the stock of public capital in other countries: to the best of our knowledge, this is the first paper to shed some light on this issue within an econometric framework that duly takes into account cross sectional correlation. Parameter estimates show that the elasticity of own public capital is barely affected by the inclusion of the stock of public capital in other countries. In turn, the latter enters significantly in the production function with an elasticity of about 0.4. This effect is larger than that reported by Cohen and Morrison (2004), for a panel of US states; however, it is also notably smaller than that reported by Bronzini and Piselli (2009) for a sample of Italian regions. Our findings of a large spillover effect for public capital however appears to be quite sensitive to the specification of the production function. In fact, if we include in equation the stock of R&D, the stock of public capital in other countries becomes marginally insignificant. This weak evidence of positive cross-country spillovers suggests that public capital might be underprovided because individual countries might not be internalizing the full benefits arising from it: this finding might lend some empirical support to the role played by EU institutions in financing major European public infrastructure projects, such as the so called Trans-European Networks.

There is evidence (Bottasso and Conti (2010)), that congestion might significantly reduce the productive effects of public infrastructures; therefore, in column 4 we include the public capital stock after dividing it by country population (CONGPOP) in order to take into account possible congestion effects. Empirical results confirm that public capital still enters significantly in the production function, but with a smaller coefficient of about 0.05.

A possible pitfall of the specifications estimated so far is the possibility that human capital corrected labor has been estimated with error; for this reason we have run some versions of Eq. (3) with raw labor. In column 5 we estimate a production function augmented only by the stock of public capital: we note a low public capital elasticity of about 0.06, statistically significant at 1%. In column 6 we augment the previous regression with the stock of R&D and we find public capital and R&D elasticities of about 0.09. Finally, in column 7 we estimate our baseline specification by including the stock of patents and main results are again broadly confirmed.[6]

On average, the econometric estimates displayed in Tables 4 and 5 suggest an elasticity of GDP with respect to public capital in the range of 0.05–0.15, with an average of 0.11. The

[6] As we mentioned in the data section, the use of public capital stock in monetary terms might be criticized on a number of grounds, namely the differences across countries in building project costs associated to both government efficiency and corruption; differences in timing of privatization of government assets; etc. We decided to assess the robustness of our results by running a series of regressions using alternatively railways and motorways kms instead of the public capital stock. We generally found that both motorways and railways have a positive and statistically significant impact on GDP, with elasticities of 0.02 and 0.17, respectively.

importance of these results for policymakers could be better appreciated with the following thought experiment: what would happen to GDP if a country could increase its public capital stock by about 40% – which corresponds to a movement from the 25th percentile of the public capital to population ratio distribution (Australia) to the 75th percentile (Norway)-? In this case, GDP would increase in the long run by about 4.4% (equivalent to an additional 0.2 percentage points per annum over a 20 year horizon) if we assume an elasticity of 0.11; but the increase could be as low as 2% (equivalent to slightly less than 0.1 percentage points per annum over a 20 year period) if we instead consider the most conservative of our elasticity estimates. These are no trivial effects and lend some support to those who argue that public spending cuts should hinge relatively more on current expenditure instead of investment expenditure; given that many EU economies have been currently undertaking fiscal consolidation processes, policy makers should bear in mind the importance of public investments even if the magnitude of the effects found in this paper also suggest that even important public capital accumulation programmes are unlikely to play a major role in sustainably increase the slow rates of growth of some OECD countries.

Our elasticity estimates allow us also to compute the gross rates of return of public capital. Considering the public capital stock to GDP ratio in 2002, this yields a gross rates of return of public capital of about 0.23, with most countries in the range 0.15–0.25, with the true exceptions being Japan and Finland on the lower and upper tails of the returns distribution, respectively. Such rates of return are quite large, although notably smaller than the 100% value implied by Aschauer (1989) empirical estimates. If we instead take the most conservative value for the public capital stock elasticity, namely 0.05, these gross rates of return would be about halved, with most countries following in the range 0.08–0.15. It might be interesting to compare these gross rates of return of public capital to the user cost of public infrastructure in order to compute net rates of return (as of 2002): we find, for an elasticity of 0.11, that all countries in our sample might have a positive net rate of return of public investment, perhaps with the exception of Japan.[7] However, for an elasticity of 0.05 we find rates or return very close to zero or even negative for a few countries, namely Japan, Austria, New Zealand, Germany and France. These results suggest that, for some countries, policymakers should be extremely careful before deciding large plans of public investments, because their net returns could well be negative.

6. Short run dynamics

Given our finding that the production function in Eq. (3) represents a long run cointegrating relationship, we re-parametrize it in the Error Correction Form (ECM) in order to analyze short term dynamics and to formally test for Granger Causality between GDP and the explanatory variables in Eq. (3), both in the short and in the long run. In particular, we consider the following

[7] The user cost of public capital has been computed using the Jorgenson approach as $ucpc = (p_I/p)(\delta - \hat{p}_g + r)$, where p is the GDP deflator, p_I is the deflator of investment, r is the long run real interest rate on ten year government bonds, $\hat{p}_g$ is the rate of change of the deflator of investment (averaged over the period 1997–02) and δ is the depreciation rate, assumed to be 4.5%. It is important to remember that in this empirical analysis we do not include in *ucpc* the distortions arising from non-lump sum taxation. On the other hand, we do not consider the benefits of public capital accruing to households (e.g. in terms of lower travel time). Data for real interest rates are taken from the EU AMECO database, while the GDP and investment deflators are from the Penn World Tables.

panel ECM:

$$\begin{aligned}\Delta y_{it} &= \varphi_{1i} + \eta_1\hat{e}_{it-1} + \delta_{11}\Delta y_{it-1} + \delta_{12}\Delta lh_{it-1} + \delta_{13}\Delta k_{it-1}\\ &\quad + \delta_{14}\Delta g_{it-1} + \delta_{15}\Delta p_{it-1} + u_{it}\end{aligned} \tag{10}$$

$$\begin{aligned}\Delta lh_{it} &= \varphi_{2i} + \eta_2\hat{e}_{it-1} + \delta_{21}\Delta y_{it-1} + \delta_{22}\Delta lh_{it-1} + \delta_{23}\Delta k_{it-1}\\ &\quad + \delta_{24}\Delta g_{it-1} + \delta_{25}\Delta p_{it-1} + u_{it}\end{aligned} \tag{11}$$

$$\begin{aligned}\Delta k_{it} &= \varphi_{3i} + \eta_3\hat{e}_{it-1} + \delta_{31}\Delta y_{it-1} + \delta_{32}\Delta lh_{it-1} + \delta_{33}\Delta k_{it-1}\\ &\quad + \delta_{34}\Delta g_{it-1} + \delta_{35}\Delta p_{it-1} + u_{it}\end{aligned} \tag{12}$$

$$\begin{aligned}\Delta g &= \varphi_{4i} + \eta_4\hat{e}_{it-1} + \delta_{41}\Delta y_{it-1} + \delta_{42}\Delta lh_{it-1} + \delta_{43}\Delta k_{it-1}\\ &\quad + \delta_{44}\Delta g_{it-1} + \delta_{45}\Delta p_{it-1} + u_{it}\end{aligned} \tag{13}$$

$$\begin{aligned}\Delta p_{it} &= \varphi_{5i} + \eta_5\hat{e}_{it-1} + \delta_{51}\Delta y_{it-1} + \delta_{52}\Delta lh_{it-1} + \delta_{53}\Delta k_{it-1}\\ &\quad + \delta_{54}\Delta g_{it-1} + \delta_{55}\Delta p_{it-1} + u_{it}\end{aligned} \tag{14}$$

where Δ represents the first difference operator, $\hat{e}_{it}$ he residual of the production function in Eq. (3) estimated with the CUP-FM estimator and η measures the speed of adjustment to the equilibrium in the above model, while u is an error term and the φ_i s are a set of country fixed effects.[8] For the variables in Eqs. (10)–(14) to represent a long run cointegrating relationship, the Engle–Granger representation theorem requires at least one of the φ_i s to be significantly different from zero. Moreover, the sign of the ηs, as well as that of the δ coefficients, can be used to test for the existence of short and long run Granger causality.[9] As far as the estimation strategy is concerned, the presence of the lagged dependent variables as well as simultaneity concerns associated to production function inputs lead us to prefer the GMM-System estimator of Arellano and Bover (1995) to OLS.[10]

Parameter estimates in Table 6 show that in all equations there is no evidence of second order serial correlation and both Hansen and Difference in Hansen test statistics suggest that we do not fail to reject the null hypothesis that instruments are uncorrelated with the error terms. Turning to parameter estimates, the coefficient of the error correction term (η_1) in the GDP equation is negative, as required for the system to be stable, and statistically significant at 10% level,[11] confirming that output is caused in the long run by the variables in Eq. (3); however, an F test on the joint significance of lagged variables fails to reject the hypothesis that they are jointly equal to

[8] The inclusion of country fixed effects in the system of Eqs. (10)–(14) is due to the existence of a set of country trends in the cointegrating relationship (3).

[9] For instance, if (η_4) in the equation for public capital is not significantly different from zero, then one might say that public capital is not Granger-caused in the long run by the other variables in the system and can therefore be considered as weakly exogenous; in turn, if also the coefficients of the lagged differentiated variables are jointly equal to zero, then public capital could be considered strongly exogenous.

[10] As a robustness check we have also estimated the panel ECM by OLS, fixed effects and GMM-DIFF (as well as by the GMM-SYS with different lags for the instruments) finding similar results.

[11] The coefficient of 0.6 implies that the system returns to its log run equilibrium following a shock in less than two years.

Table 6
Estimates of the ECM model.

Dep. Var.	$\Delta\ln Y$	$\Delta\ln LHC$	$\Delta\ln K$	$\Delta\ln G$	$\Delta\ln P$
$\Delta\ln Y_{t-1}$	0.745***	0.961	0.124***	0.027	0.036
	(0.238)	(0.615)	(0.589)	(0.072)	(0.297)
$\Delta\ln LH_{t-1}$	−0.114	−0.533**	0.034	0.002	0.329
	(0.126)	(0.254)	(0.027)	(0.013)	(0.312)
$\Delta\ln K_{t-1}$	0.100	−0.372	0.769***	−0.008	0.069
	(0.269)	(0.597)	(0.079)	(0.069)	(0.349)
$\Delta\ln G_{t-1}$	−0.121	0.213	0.013	0.889***	−0.156
	(0.145)	(0.227)	(0.031)	(0.038)	(0.18)
$\Delta\ln P_{t-1}$	0.091	−0239	−0.0154	0.023	0.869***
	(0.11)	(0.209)	(0.011)	(0.019)	(0.131)
$\hat{e}_{t-1}$	−0.674*	−0.238	0.019	−0.047	−0.688
	(0.363)	(0.209)	(0.098)	(0.016)	(0.663)
Country fix. effects	X	X	X	X	X
M1 (p val.)	0.038	0.40	0.23	0.25	0.11
M2 (p val.)	0.38	0.07	0.35	0.65	0.39
Hansen (p val.)	0.50	0.64	0.57	0.25	0.72
Diff. Hansen (p val.)	0.96	0.60	0.73	0.59	0.93

Note: GMM-SYS estimates; standard errors are two step robust with the Windmeijer correction; M1 and M2 are Arellano-Bond tests for first and second order serial correlation; Hanse J test is an over-identification test statistics, Diff-Hansen is a test for the validity of the extra moment conditions for the level equation. Instrument used are y, lh, k, g, p all dated T-5, T-6, T-7 for the level equation and the same variables differenced for the level equation; instruments have been collapsed to avoid over-fitting problems associated to the proliferation of instruments when T is large.

* $p<0.1$.
** $p<0.05$.
*** $p<0.01$.

zero, implying that there is not short run impact of the regressors on GDP. This result suggests that public capital investments might not be an effective countercyclical instrument, while it might be a valid tool for increasing GDP in the long run. Moreover, we cannot reject the hypothesis that the coefficients of the error correction terms are equal to zero in all the other equations; this, together with the fact that a series of F tests suggest that lagged differenced variables are jointly statistically significant in the case of the private capital equation only, leads us to conclude that public capital, the stock of patents and human capital-augmented labor are strongly exogenous, while private capital is only weakly exogenous.

7. Conclusions

In this paper we consider a panel of 21 OECD countries observed over the period 1975–2002 and we estimate different production functions in order to investigate the short and long run relationships between public capital and GDP, while controlling for other productivity determinants such as human capital and patents. The novelty of our study rests in the adoption of the most recent econometric methodologies which control for the presence of cross sectional dependence and cross sectional cointegration in a panel time series framework. Main results show that our series are not stationary and that cross sectional cointegration does exist among estimated common factors. We estimate the long run relationship among GDP and explicative variables with appropriate estimation techniques which account for such data characteristics. On average, the econometric estimates suggest a long run elasticity of GDP with respect to public capital in the range of 0.05–0.15, thus suggesting that public capital investments might have positive (*but relatively modest*) productive effects in the long run; in turn, the average elasticity of GDP with respect to patents turns out to be about 0.11. Overall result are confirmed when we estimate the impact of public capital on TFP and when we augment the baseline model in order to take into account congestion and spill-over effects stemming from public investments in neighboring countries. Short run analysis however does not confirm the existence of a significant impact of public capital on GDP: given that our findings point toward the existence of a productive effect of public capital which holds only for the long run, it might be the case that the confidence put into public infrastructure expenditure, during the last recession, as a countercyclical policy instrument, might have been too optimistic. Furthermore, in our empirical model we do not consider the issues related to the financing of public infrastructure investments (tax versus debt) and the general equilibrium effects they generate: while for some countries the latter might be negligible, given the almost zero interest rates they have been recently paying on their public debt, for other countries financing costs might be so large (either because interest rates on government bonds are high or because their tax rates are so high that the shadow costs of taxation are large as well) that net returns of public investment might become negative not only in the short, but also in the long run.

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