Investigaciones Regionales

Public Capital and Regional Economic Growth: a SVAR Approach for the Spanish Regions

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ABSTRACT: Recently, a significant share of the empirical analysis on the impact of public capital on regional growth has used multivariate time-series frameworks based on vector autoregressive (VAR) models. Nevertheless, not as much attention has been dedicated to the analysis of the long-run determinants of regional growth processes using multi-region panel data and applying panel integration and co-integration techniques. This paper estimates the dynamic domestic effects of public infrastructure using a structural vector autoregressive (S-VAR) methodology for the Spanish regions. From a methodological point of view, the paper contains several features that can be viewed as a contribution to the existing empirical literature. First, the important issues of the stationarity of the data and the existence and estimation of cointegrating relationships in the long-run are addressed in the context of the analysis of panel data. Secondly, the long-run cointegrating production function is embedded within structural vector error correction (S-VEC) short-run models to produce consistent estimates of impulse responses, contrary to many researchers who have estimated unrestricted VAR models in levels or VAR models in first differences. The estimates reveal new results with respect to the previous empirical evidence.

JEL Classification: C32; E62; H54; R53.

Keywords: Public capital, regional growth, VAR methodology, Spain.

Capital Público y Crecimiento Económico Regional: un enfoque SVAR para las Regiones Españolas

RESUMEN: Recientemente, un porcentaje significativo de los estudios empíricos que analizan el impacto del capital público sobre el crecimiento económico regional ha utilizado series temporales multivariantes basadas en modelos de vectores autoregresivos (VAR). En este contexto, no se ha prestado demasiada atención al

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Received: 6 april 2011 / Accepted: 12 june 2011.

análisis de los determinantes a largo plazo de los procesos de crecimiento regional utilizando paneles de datos multi-regionales y aplicando técnicas de integración y cointegración para paneles. Este trabajo estima los efectos domésticos dinámicos de las infraestructuras públicas utilizando una metodología de vectores autorregresivos estructurales (S-VAR) para las regiones españolas. Desde el punto de vista metodológico, el trabajo contiene distintas características que pueden ser vistas como una contribución a la literatura empírica existente. Primero, las importantes cuestiones de la estacionaridad de los datos y de la existencia y estimación de relaciones de cointegración en el largo plazo son abordadas en el contexto del análisis de los datos de panel. En segundo lugar, en los modelos de corto plazo de vectores de corrección de error estructurales (S-VEC) se tiene en cuenta la función de producción de cointegración en el largo plazo para producir estimaciones consistentes de las respuestas a impulsos; esto es contrario a lo que hasta ahora han hecho muchos investigadores, que han estimado modelos VAR sin restringir en niveles, o modelos VAR en primeras diferencias. Las estimaciones muestran resultados nuevos con respecto a la evidencia empírica previa.

Clasificación JEL: C32; E62; H54; R53.

Palabras clave: Capital Público; Crecimiento Regional; Metodología VAR; España.

1. Introduction

The role of public capital investment has been a critical component of the policy agenda focused on enhancing regional growth. Permanent changes in public capital investment could have important effects on regional economic activity. The theoretical arguments pointing to the role of public capital on economic development are embodied in many of the «New Growth Theory» (NGT) and «New Economic Geography» (NEG) models. These models challenge traditional Neo-Classical Growth Models, which predicted regional convergence without a specific theoretical consideration of the role of public capital: steady-state income per capita is assumed to be independent of the initial conditions, no matter the size of the inherited differences in capital stock.

In contrast, endogenous growth theory was based on the existence of increasing returns and positive externalities (Romer, 1986, 1990; Lucas, 1988, 1993), where the existence of increasing returns could be explained by an intensive investment in knowledge, human capital or infrastructure (e. g., Barro, 1990). In this theoretical context, Barro and Sala-i-Martin (1992) analyzed the growth effects of the flow of productive government spending, while Turnovsky (1997) and Aschauer (2000) considered the growth effects of the stock of public capital. Therefore, the stock of public infrastructures could be among the significant variables conditioning the level and growth of regional productivity, and thus government policy —through its expenditure programs on public capital over space— would have the potential to affect the long-run growth rate of a regional economy.

On the other hand, in the early 1990s, the NEG models provided explanations for the formation of a large variety of economic agglomerations in geographical space (Fujita and Krugman, 2004). This new line of research emphasizes the interaction among increasing returns to scale, transportation costs (broadly defined) and the movement of productive factors. According to Fujita and Thisse (2002), public expenditure is fundamental in both the reduction of transport costs and in the supply of local public goods, playing a key role in the critical trade-off between increasing returns and transport costs. The general belief is that public capital could increase the productivity of private factors, thereby generating a significant impact on growth. Accordingly, it becomes essential (from a policy evaluation point of view) to have a quantifiable measure of the impact of public investment on the growth performance of receiving economies.

There exist a number of studies (see, among others, Kamps, 2005, and Roca-Sagalés and Sala, 2010, for a comprehensive review) in the literature that documents the effects of public capital on economic growth. Initially, earlier studies (Aschauer, 1989, and Barro, 1990) and the following set of studies have concentrated mostly on country case studies. Lately, a second set of studies (with earlier work from Munnell, 1990) has focused on regions within a country. These econometric studies have shown the importance of spillover effects as potential factors that may affect regional growth. However, an overwhelming amount of research has focused on the measure of spillover effects in the analysis of the aggregate effects of the public capital provision at the regional level (see, for example, Holtz-Eakin and Schwartz, 1995; Boarnet, 1998; and Pereira and Roca-Sagalés, 2003). Adopting this perspective, spillover effects, understood as positive or negative externalities derived from the impact of the public capital provision in a region, would have to be considered when investigating the effects of public capital in one region on the production of other regions.

In sum, the evaluation of the aggregate effects of public capital should contemplate the existence of both direct (domestic) and indirect (spill over) effects. For a region, domestic effects are the effects derived from public capital installed in the region itself, while than spillover effects are derived from public capital installed outside that region. Even then, the issue of domestic effects has been ignored the recent contributions try to improve the measurement of the spillover effects of public capital. Empirical results and policy implications from the existing literature based on spillover effects to regional economies should be complemented, taking into account the own specificities and constraints of such regions derived from the analysis of domestic effects.

In the present paper, the effects of public capital for the 17 regions that make up Spain are measured using a «structural» VAR (S-VAR) approach. The dynamic effects will be considered from a domestic perspective [1]. From a methodological point of view, the paper contains several innovative features that can be viewed as a contribution to the existing empirical literature. First, the important issues of the stationarity of the data and the existence and estimation of cointegrating relationships in the long-run are addressed in the context of the new tools proposed recently

[1] This article is complementary to Márquez *et al.* (2010), where the spillover effects of one-time innovations in the public capital installed in a given region on the economic growth of the other Spanish regions (cross-border effects) are estimated by using «bi-regional models».

for the analysis of panel data [2]. In this sense, to date, none of the existing studies of the impact of public capital investment on the economic growth performance using multi-region panel data has applied panel integration and cointegration techniques to analyze the long-run determinants of regional growth processes. Secondly, based on the integration and cointegration results, the long-run cointegrating production function is embedded within structural vector error correction (S-VEC) short-run models to produce consistent estimates of impulse responses, in contrast to many researchers who have estimated unrestricted VAR models in levels or VAR models in first differences. These models might produce inconsistent estimates of the impulse response functions.

The results could assist in formulating economic policies, complementing the approach shown in Márquez *et al.* (2010), where it is possible to identify the regions where the spillover effects originate. From these findings, the regions that are able to generate spillover effects on other regions are determined, deepening the understanding of the spatial and temporal dynamics of the location of further public investment.

The results on the impact of public capital on regional economic growth in the present paper are somehow unexpected in comparison to previous research findings on the Spanish regional economies. A main determinant of these results is the inclusion in the short-run regional models of an error correction term derived from the estimation of a joint steady-state relationship for the Spanish regional system. The use of the pooled mean group methodology to obtain the estimation of the production function of the regional economic system as one cointegrating vector allowed for cross-section specific heterogeneity in the coefficients of the short-run parameters of the regional VAR models (see Pesaran *et al.*, 1999). Thus, the stability of the regional models in the short-run is ensured by means of an error correction mechanism that takes into account the information of the joint regional equilibrium in the long run.

Departing from the standard method used until now, the application of this empirical approach would be helpful in simulating the domestic effects generated by regional public capital investment in a region on output, employment, and private performance in the same region. The results that were obtained involve both positive and negative domestic effects from public capital. Another contribution derives from the analysis of the spatial distributions of the estimated domestic effects: the long run effects of public capital on private capital show a strong geographic pattern and reveal the presence of positive spatial dependence.

In section 2, a succinct review of the theoretical and empirical literature on public capital and economic growth is presented, with special reference to the Spanish regional case. In section 3, a brief description of the data properties is provided and the empirical results are reported and discussed. The final section summarizes the paper's major findings and offers some policy prescriptions.

[2] To separate the long run behaviour from the short run dynamics it is necessary that the variables under consideration are nonstationary [typically integrated of order one, $I(1)$], so that the errors from the long-term cointegrating relationships could be stationary.

2. Public capital and regional economic growth

Public capital has been considered an important instrument of regional policy (see de la Fuente and Vives, 1995). Previous research about the role of public capital in economic growth could be systematized considering different perspectives (see Romp and de Haan, 2007 for a survey of the extensive literature): the definition and scope of the public capital variable; the division between country and regional level studies; the main approaches (production functions, cost functions and VAR/VECM models); and the level of aggregation of the data (data over specific sectors or data over all sectors).

Authors like Aschauer (1989), García-Mila and McGuire (1992) and Munnell (1992), among others, have applied neoclassical production functions. Their findings provide a diversity of results, making it difficult to obtain any definitive conclusions. Further, several inconsistencies have been reported. The single-equation regression model used by Aschauer has potential econometric problems like spurious regression due to non-stationarity of the data, possible misspecification of the production function, endogeneity and/or the direction of causality from public capital to productivity. With respect to the problem of the spurious regression, cointegration theory provides a means of approaching this problem, taking into account the non-stationarity problem. The missing variables problem makes reference to the possible omission of relevant variables like those indicated by NGT (e. g., knowledge, human capital, R&D investment, etc.). Finally, the direction of causality, that is, the possible influence from economic growth on public capital, causing a problem of endogeneity, is one of the main drawbacks of the production function approach.

Alternatively, the cost function approach (see, for example, Ezcurra *et al.,* 2005 for the Spanish case) measures the impact of public capital on economic growth in terms of cost-savings benefits. This approach evaluates whether costs decrease with public capital provision. The cost-function approach is more flexible than the production-function approach, and this is its main advantage. Nevertheless, the requirement of data for the cost-function approach is greater than in the case of the production-function approach.

More recently, in the context of the VAR models, the impulse response analysis has been used as a fundamental tool to simulate the effect that an unexpected change of the public capital would have on another variable, for example, on the value of regional production. The use of the VAR approach to test the significance of the dynamic effects of public capital on economic growth presents some advantages. According to Kamps (2005), this approach allows for the existence of indirect links between the variables under investigation. In addition, if the number of long-run (cointegrating) relationships are tested and estimated consistently, the vector error correction (VEC) models would produce consistent estimates of impulse response functions. With respect to the empirical literature where the VAR methodology has been used to simulate the effects of unexpected changes in the public capital on regional macroeconomic variables for the case of the Spanish regions, a few studies like Pereira and Roca-Sagalés (1999, 2001) can be found. Further, Pereira and Roca-Sagalés (2003) and Roca-Sagalés and Sala (2006) have investigated the existence of regional spillover effects of public capital formation in the economic regional system of Spain.

Regional economic growth could be affected by public capital through different mechanisms. The most direct way is the consideration of public capital as a factor of production (see Sturm, 1998). The effects derived from the interactions between public capital and private capital would be another way. In this sense, the existence of a positive effect of public investment on private capital accumulation was obtained by Martínez-López (2006) for the Spanish regions over the period 1965-1997. On the other hand, the new economic geography (Krugman, 1991; Fujita *et al.,* 1999) suggests that public capital may affect regional economic growth through its influence on transport costs. More public capital (specially transport infrastructure) could have an important impact on market access (see, for example, Redding and Venables, 2004, or Head and Mayer, 2004). Good access to large markets (high market access) may prove to be critical in the explanation of regional economic performance.

Finally, it is important to highlight that the distinction between short- and long-run effects of public capital is important in regional economic analysis. There is no reason to believe that public capital has the same spatial impact whether in terms of sign or magnitude of its effects in both the long- and the short-run. In this sense, and with respect to the long-run effects of public investment, Baxter and King (1993) note that an unexpected (permanent) increase in public investment will induce a response of output. This long run response will be both direct and indirect (derived from the supply-side effect generated by private capital and labor). On the other hand, considering the short run effects of public investment, Baxter and King (1993) declare that an unexpected (permanent once it occurs) shock in the stock of public capital will imply a transition of the economy to the new steady state. During this transition, the stock of public capital accumulates, increasing the output. This accumulation involves a governmental absorption of resources that could generate some interactions. As a result, the rising stock of public capital will alter the stock of private capital and labor through the change of the marginal product. Obviously, this theoretical difference between short- and long-run effects has important empirical implications as demonstrated example, by Moreno *et al.* (2002) who determined the short- and long-run effects of public infrastructure in the context of manufacturing industries in the Spanish regions using aggregated cost functions. In summary, one might venture to say that public capital could be a complement or substitute with respect to private capital and employment, conditioning the pattern of the output responses; further, the response could be different in the long- and short-run.

As documented in the literature on the effects of public infrastructure, although there is a general consensus of the need for a certain level of public capital, the results obtained are inconclusive. The studies analyzing the impact of public capital on regional output and regional productivity generally point to the effectiveness of public capital as a tool for regional policy; some examples are provided in order to reveal the different conclusions that have been derived to date. Destefanis and Sena (2005), in studying the Italian case, concluded that public capital had positive effects, at least in some Italian regions. Karada *et al.* (2004) used a vector autoregression (VAR) model to estimate long run accumulated elasticities of private sector variables with respect to public capital in the seven geographical regions of Turkey. These authors showed evidence of the positive effects of public capital on private output in five of the seven regions. However, for some regions, public capital crowds out private sector inputs. Sloboda and Yao (2008) analyzed interstate spillovers of private capital and public spending in the United States; they detect crowding out effects among the 48 contiguous states for the period 1989-2002.

For the Spanish economy, the general perception is the existence of positive effects such as Cantos *et al.* (2005), Ezcurra *et al.* (2005), Moreno *et al.* (2002), Boscá *et al.* (2002), Mas *et al.* (1996). Other studies such as Gorostiaga (1999) and González-Páramo and Martínez (2003) do not show significant effects of public capital stock on economic growth. In the literature, it is argued that the non significant effect of public investment in economic growth is due to the existence of spillover effects. Thus, Salinas-Jiménez (2004), obtains positive effects for the Spanish case, but only if spillover effects were taken into account.

3. The dynamic domestic effects of public capital on the Spanish regions: new evidence from structural VAR models

This section describes an empirical application analyzing the domestic effects of public capital for the Spanish regions. This empirical section is organized as follows. First, the Spanish data used to implement the S-VAR approach are presented. Secondly, panel integration tests are applied to this data set, and the results of the unit roots analysis are reported. Next, panel cointegration tests are employed to test for cointegration, and the results on the estimation of the long-run equilibrium cointegrating relationship are presented. Finally, individual S-VEC short-run models are first presented and then estimated, and the results of an impulse response analysis based on a set of identifying assumptions are shown.

3.1. Spanish regions and data

Spain is composed of 17 regions and Ceuta and Melilla —two Spanish North African cities— that constitute the so-called Autonomous Communities [3]. In the present work, only the 17 regions in Spain are analyzed (see Figure 1). The Spanish regional system has a marked economic core-periphery pattern, with an unequal

[3] The Autonomous Communities have achieved the status of self-governed territories, sharing governance with the Spanish central government within their respective territories.

economic geography. Traditionally, the peninsular economic periphery is comprised of Castilla-León, Castilla-La Mancha and Extremadura while Madrid, País Vasco, Cataluña and Valencia make up the economic core. Galicia, Andalucía, Murcia, Islas Baleares and Islas Canarias are also considered as «peripheral» regions; while Navarra, La Rioja, and Aragón may be considered as «core» regions. Finally, Asturias and Cantabria are historical «core» regions, but currently experiencing significant industrial restructuring processes.

Figure 1. Spanish Regions

Accordingly, the panel data-set contains 17 regions over the period 1972-2000; for each region, the variables used are the public net productive capital stock (*PK*), the private net capital stock (*K*), the number of employed persons (*E*), and the real Gross Added Value (*Y*). The regional series for *Y* have been drawn from the Instituto Nacional de Estadística (INE) of Spain and from the Hispadat database (see Pulido and Cabrer, 1994, and Cabrer, 2001) and the time series for *PK*, *K* and *E* have been taken from the Instituto Valenciano de Investigaciones Económicas (IVIE) of Spain. The regional public capital stock comprises public capital owned by the local, regional and national administrations, including transport infrastructures (roads, ports, airports and railways), water and sewage facilities and urban structures.

Table 1 displays selected summary indicators for the 17 Spanish regions, presenting some relevant data about the geographical distribution of the aforementioned variables for the (approximately) three decades comprising the database (1972-1980, 1981-1990 and 1990-2000). As the table shows, there are clear regional disparities in the geographical distribution of output, employment, and private and public capital stocks. These sharp disparities could be shown, for example, in the case of two regions like Madrid and Extremadura. Madrid has an area corresponding to 1.6% of the Spanish regional system. During the first (third) sub-period, Madrid produced 15.7% (16.6%) of the aggregate output, with 12.1% (13.7%) of the total employment, 15.4% (15.3%) of the private capital stock and 10.6% (10.0%) of public capital stock of Spain. Conversely, Extremadura, with 8.3% of the total area, during the first (third) sub-period accounted for only for 1.7% (1.8%) of the Spanish output, with 2.7% (2.3%) of the total employment, 1.8% (1.9%) of private capital and 3.1% (3.3%) of public infrastructures of Spain.

Table 1. Basic data for Spanish regions

Regions	*Area*	*GAV*			*Employment*			*Private Capital*			*Public Capital*		
	% km²	*1972-1980*	*1981-1990*	*1990-2000*	*1972-1980*	*1981-1990*	*1990-2000*	*1972-1980*	*1981-1990*	*1990-2000*	*1972-1980*	*1981-1990*	*1990-2000*
AN	17.36	14	13.55	13.86	14.43	14.07	14.61	11.83	12.41	13.04	14.95	15.56	17.21
AR	9.45	3.35	3.46	3.31	3.36	3.37	3.25	3.16	3.18	3.15	5.57	5.01	4.09
AS	2.1	3.19	2.9	2.44	3.28	3.05	2.55	3.28	2.92	2.5	3.29	3.35	3.3
BA	0.99	2.12	2.27	2.27	1.77	1.96	2.14	2.36	2.52	2.94	1.46	1.47	1.56
CB	1.04	1.28	1.28	1.24	1.45	1.39	1.29	1.68	1.48	1.34	1.32	1.49	1.59
CL	18.59	6.61	6.36	5.81	7.11	6.95	6.29	6.11	6.41	6.07	10.25	9.05	7.95
CM	15.74	3.85	3.72	3.57	4.34	4.25	4.11	3.43	3.8	3.93	5.67	5.4	5.52
CN	1.48	2.77	3.46	3.72	3.18	3.53	3.93	2.88	3.14	3.49	3.63	3.89	4.05
CT	6.36	18.63	18.07	18.86	16.7	16.61	17.65	21.13	20.06	19.3	14.98	13.73	13.5
CV	4.61	9.52	9.88	9.79	9.7	10.02	10.32	10.02	11.03	11.43	8.43	8.77	9.03
EX	8.25	1.74	1.85	1.81	2.74	2.49	2.32	1.84	2.07	1.87	3.07	2.96	3.33
GA	5.86	5.75	5.91	5.57	9.52	9.13	7.4	5.31	5.45	5.32	5.77	6.32	6.85
MA	1.59	15.68	15.95	16.64	12.05	12.89	13.67	15.44	14.96	15.34	10.61	10.63	9.99
MU	2.24	2.14	2.3	2.33	2.35	2.48	2.64	2.16	2.3	2.51	1.69	2.15	2.39
NA	1.94	1.68	1.69	1.67	1.39	1.43	1.51	1.38	1.36	1.5	1.93	2.04	1.96
PV	1.4	7.09	6.61	6.34	5.9	5.66	5.61	7.42	6.25	5.56	6.3	6.79	6.71
RI	1	0.61	0.74	0.76	0.73	0.73	0.7	0.58	0.66	0.7	1.09	1.38	0.96
SPAIN	100	100	100	100	100	100	100	100	100	100	100	100	100

3.2. Testing for panel unit roots and cointegration, and estimation of the long-run equilibrium production function

The empirical analysis begins with an evaluation of the stationarity of the four variables of the database using panel unit root tests starts [4]. All panel tests used are based on the null hypothesis of the presence of a unit root in the series, with the exception of Hadri's (2000) test, whose hypothesis is that the series are stationary. The tests differ from each other in the restrictions imposed on the autoregressive process of each of the panel series. Thus, the tests of Levin, *et al*. (2002), Breitung (2000) and Hadri (2000) impose a common persistence parameter to all the series. Therefore, if the null were rejected, the alternative would be that all the series are simultaneously stationary for the first two tests and non-stationary for the latter. Alternatively, the tests of Im, *et al*. (2003) and the Fisher-type tests suggested by Maddala and Wu (1999) allow for the autoregressive parameter to change freely among the different regional variables under consideration. Therefore, the alternative hypothesis in these cases is the presence of a non-null proportion of stationary series of the total. The latter set of tests seem more appropriate from an empirical point of view as they impose less restrictions on the data generating process.

A general overview of the statistics, presented in Table 2, shows the evidence to clearly favor the hypothesis that the four basic variables considered behave as non-stationary variables, with a unit root at least for a non-negligible fraction of the 17 regions of the panel. Indeed, only for the variable *K,* in logs, do the test statistics show

Table 2. Unit root tests for log *Y*, log *E*, log *K* and log *PK*

	Log Y	*Log E*	*Log K*	*Log PK*
Null: Unit root (assumes common unit root process)				
Levin-Lin-Chu	2.201	8.162	–3.785 ***	3.445
Breitung	–2.424 ***	8.341	–2.487 ***	3.078
Null: Unit root (assumes individual unit root process)				
Im-Pesaran-Shin	0.026	8.381	–4.560 ***	0.993
Maddala-Wu ADF-Fisher	31.217	0.659	91.173 ***	26.392
Maddala-Wu PP-Fisher	40.984	0.971	96.893 ***	17.617
Null: No unit root (assumes common unit root process)				
Hadri	3.790 ***	9.371 ***	7.306 ***	6.634 ***

Notes: 1) Probabilities for Fisher tests were computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality; 2) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values; 3) Exogenous variables: Individual effects, individual linear trends; 4) Automatic selection of lags based on MAIC criterion: 0 to 4; 5) Newey-West bandwidth selection using Bartlett kernel.

[4] The use of panel unit root tests is justified by the results from recent studies [see Banerjee (1999), Baltagi and Kao (2000) or Breitung and Pesaran (2008), among others], which suggest that unit root tests based on panel data are more powerful than those based on individual data.

evidence favorable to the hypothesis of stationarity of the corresponding time series (in Table 2, a deterministic linear trend is included in all the specifications, but if not, the unit root hypothesis is clearly not rejected in this particular case). Since the test results generally support the unit root hypothesis, from now it is assumed that all time series under consideration (all in log values) are integrated of order one. This makes it possible to distinguish between short-run and long-run relations, and to interpret the long-run relations as cointegrating relationships.

To analyze the existence of cointegration between the four variables considered, three panel tests were applied. Two of them, those of Pedroni (1999, 2004) and Kao (1999), are residual-based tests that assume a single cointegrating vector; while the third test, of Maddala and Wu (1999), allows for multiple cointegrating relationships [5]. On the other hand, not all the tests used assume the same degree of individual heterogeneity; while the Pedroni and Maddala-Wu statistics allow the coefficients of each cointegration relation to vary freely for each region, the Kao approach assumes panel homogeneity.

The estimates of the various cointegration statistics are presented in Tables 3, 4 and 5. As a general assessment of the values presented in these tables, one can deduce that there is considerable evidence pointing to the existence of cointegration between the real GAV and the input-production variables for the panel of 17 Spanish regions. Thus, in the case of the Pedroni statistics, all the three versions of the PP and ADF statistics strongly reject the non-cointegration hypothesis. The Fisher type and Kao statistics also corroborate the existence of a stable long-run relationship. Therefore, the overall evidence is consistently in favor of the existence of an aggregate production function as a long-run equilibrium relationship [6].

Table 3. Pedroni panel cointegration tests (Null Hypothesis: No cointegration)

	v – stat	*ρ – stat*	*PP – stat*	*ADF – stat*
Alternative hypothesis: common AR coefs. (within-dimension)				
Unweighted panel stats	0.964	–0.907	–5.187 ***	–5.353 ***
Weighted panel stats	–1.426	–0.684	–5.635 ***	–6.453 ***
Alternative hypothesis: individual AR coefs. (between-dimension)				
Group-mean stats		0.795	–4.525 ***	–4.542 ***

Notes: 1) All of the panel and group statistics have been standardized by the means and variances given in Pedroni (1999) so that all reported values are distributed as $N(0,1)$ under the null hypothesis of no cointegration; 2) The panel-stats weighted statistics are weighted by long run variances (Pedroni, 1999, 2004); 3) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values (1.28, 1.64 and 2.33, respectively); 4) For the semiparametric *PP* tests the Newey-West (1994) rule for truncating the lag length for the kernel bandwidth has been used, and for the parametric *ADF* tests a step-down procedure starting from $K = 2$ has been used; 5) The residuals have been estimated using the least squares estimator.

[5] See Gutiérrez (2003) for a Monte Carlo analysis of the statistical properties of these tests.

[6] With respect to the Maddala-Wu results, it is known that the Johansen tests —the kernel of the Maddala-Wu statistics— for the second and subsequent cointegrating vector suffer from substantial size distortions and tend to find multiple cointegrating vectors when the ratio of data observations to the number of parameters is relatively small (Maddala and Kim, 1998). This might explain the non rejection of the hypothesis of the presence of two cointegrating vectors both in maximal eigenvalue and trace statistics.

Table 4. Kao panel cointegration test (Null Hypothesis: No cointegration)

	t – stat
ADF	–4.347 ***

Notes: 1) Probability has been computed assuming asymptotic normality; 2) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values; 3) Trend assumption: No deterministic trend; 4) Lag selection: Automatic 2 lags by SIC with a max lag of 2; 5) Newey-West bandwidth selection using Bartlett kernel; 6) The residuals have been estimated using the least squares estimator.

Table 5. Maddala and Wu Fisher-type panel cointegration tests [Null Hypothesis: number (r) of cointegration relationships]

	Trace – stat	*Max.eigen. – stat*
$r = 0$	221.10 ***	185.00 ***
$r \leq 1$	76.26 ***	56.99 ***
$r \leq 2$	44.76	40.22
$r \leq 3$	44.96 *	44.96 *

Notes: 1) Probabilities have been computed using asymptotic Chi-square distribution; 2) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values; 3) Trend assumption: Linear deterministic trend; 4) Lags interval (in first differences): 1 to 1.

The next step is to estimate the parameters of the detected long-run equilibrium production function. The estimated steady-state relationship has the following expression:

$$y_{it} = \beta_{0,i} + \beta_1 e_{it} + \beta_2 k_{it} + \beta_3 pk_{it} + \beta_4 t + v_{it} \tag{1}$$

where $y = \log Y$, $e = \log E$, $k = \log K$ and $pk = \log PK$. As shown, long-run homogeneity of input elasticities is assumed [7], fixed-region effects ($\beta_{0,i}$) are permitted in order to control for time-invariant regional heterogeneity, and a temporal trend (t) is introduced to take into account the time evolution of the technical progress [8]. Given the homogeneity of slopes hypothesis assumed in the above specification, the estimated relation must be interpreted as an average long-run equilibrium production function for the panel of 17 Spanish regions.

With respect to the technique chosen to estimate the equilibrium relationship, and given that ordinary least squares (OLS) estimates of the long-run model would

[7] We also perform the long-run analysis on a region-by-region basis using the Johansen approach. Not surprisingly (due to the short span of data available at the single-region level), the Johansen individual-estimates of the long-run parameters were mixed and noisy, with some coefficients appearing as implausible. The poor results obtained in this case compels us to impose the homogeneity assumption in the estimation of the long-run equilibrium production function [see, among others, the works of Pesaran *et al.* (1999) and Baltagi *et al.* (2000) that consider the issue of pooling in detail, asking the question «To pool or not to pool?»].

[8] Also, introducing a trend in the long-run relation ensures that the deterministic trend properties of the VEC models estimated later remain invariants to the cointegrating rank assumptions (Pesaran *et al.*, 2000).

suffer from asymptotic bias (Kao and Chiang, 2000), the so-called Dynamic Seemingly Unrelated Cointegrating Regressions (DSUR) method proposed by Mark *et al.* (2005) was used. This method allows for the efficient simultaneous estimation of panel cointegrating relationships with correlated disequilibrium errors, working with panel data in which, as in our case, the cross-sectional dimension is small or about the same order with respect to the length of the time series.

The results of the DSUR estimation of the average long-run production function are presented in Table 6. According to these results, the elasticity of employment is around 0.35. Private capital and public capital show elasticities estimated to be 0.32 and 0.10, respectively. In terms of statistical significance, magnitude and theoretical plausibility, the estimates obtained from the DSUR are very consistent, and are well within the range of estimates obtained by other authors. In this sense, one could point to the work of Kamps (2005) and Romp and de Haan (2005), among others, who have summarize information on international studies that have analyzed the dynamic effects of public capital, while Boscá *et al.* (2004) and Mas and Maudos (2005) present surveys of the Spanish experience about this topic.

Table 6. DSUR estimates for $y_{it} = \beta_{0,i} + \beta_1 e_{it} + \beta_2 k_{it} + \beta_3 pk_{it} + \beta_4 t + v_{it}$

$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	$\hat{\beta}_4$
0.348 *** [0.025]	0.315 *** [0.029]	0.102 *** [0.022]	0.010 *** [0.001]

Notes: 1) Cross-section SUR standard errors are given in brackets; 2) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate p-values.

3.3. Region-specific and short-run S-VEC models

In the empirical strategy, an explicit distinction is made between the long-run properties of the regional economies (associated in our case with the cointegrating production function suggested by the economic theory and tested and estimated in the previous sub-section) and the short-run dynamics of the regional system. In this sense, the modeling approach assumes that macroeconomic or regional economic theories are not explicit enough to propose specific relationships that might exist between the input and output regional variables over short time horizons. Hence, a parsimonious and flexible econometric specification is used that attempts to account for the complex dynamic relationships that drive the short-run regional behavior. Specifically, the short-run dynamics of each region are modeled within a VAR framework using S-VEC models that drive the dynamics of adjustment of the input and output variables of each region to the long-run equilibrium across-regions.

These hypotheses allow estimation and testing of the domestic properties of the different region-specific models, analyzing the dynamics of the transmission of shocks from public capital to the rest of state variables (private capital, employment and output).

The reference individual S-VEC model for the region i ($i = 1,2,...,17$) is given by:

$$A_i(L)\Delta X_{it} + C_i Z_{it} = E_{it} \quad , \quad A_{i0} E_{it} = B_i U_{it} \tag{2}$$

where $X_{it} = (pk_{it}, k_{it}, e_{it}, y_{it})'$ is the vector of endogenous variables; $Z_{it} = (1, \hat{\upsilon}_{i,t-1})'$ is the vector of predetermined variables, given in the empirical application by an intercept and the lagged estimated error correction term corresponding to the equilibrium relationship presented in Table 6; $E_{it} = (e^{pk}_{it}, e^{k}_{it}, e^{e}_{it}, e^{y}_{it})$' is the canonical errors vector from the reduced form; and $U_{it} = (u^{pk}_{it}, u^{k}_{it}, u^{e}_{it}, u^{y}_{it})'$ is the structural errors vector [9]. Matrix $A_i(L) = \sum_k A\_K_i L^k$ includes in our application a maximum of four lags, the optimal lag determined by the standard selection criteria AIC, HQ and SC, where the higher lag order is chosen based on these three information statistics.

With respect to the identification of the structural innovations, a standard recursive Cholesky-type decomposition scheme was used assuming that the relation between the canonical errors and the structural disturbances is given by the equation $A_{i0}E_{it} = B_i U_{it}$, where:

$$A_{i0} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ a^i_{21} & 1 & 0 & 0 \\ a^i_{31} & a^i_{32} & 1 & 0 \\ a^i_{41} & a^i_{42} & a^i_{43} & 1 \end{bmatrix} \qquad B_i = \begin{bmatrix} b^i_{11} & 0 & 0 & 0 \\ 0 & b^i_{22} & 0 & 0 \\ 0 & 0 & b^i_{33} & 0 \\ 0 & 0 & 0 & b^i_{44} \end{bmatrix} \tag{3}$$

This identification scheme has the following implications: i) innovations in public investment affect contemporaneously private capital, employment and real output, but the reverse is not true, ii) shocks to private capital affect contemporaneously the employment and real GAV, but the reverse is not true, and iii) unanticipated changes in employment affect contemporaneously the real GAV, but employment does not react contemporaneously to shocks in regional output. Therefore, the identified shocks are not subject in any case to the reverse causation problem.

3.4. Are there significant domestic effects of public capital formation in the Spanish regional system?

Tables 7 and 8 show summary information about the domestic effects of shocks in public capital installed inside each region displaying, respectively, the short-run and long-run elasticities of private capital, employment and real GAV obtained from

[9] To facilitate the interpretation of the estimated impulse responses, the endogenous variables (in logs) of the structural VEC models have been multiplied by 100. In this case, the accumulated impulse responses provide the percentage change in the level of the respective variable.

the seventeen regional S-VEC models considered [10]. These estimates generate respectively the 0 year point and 25 year point percentage change in private capital, employment, and output per one-percentage point (impact or long-run) change in public capital. Each point estimate in the tables is marked (or not) with an asterisk depending on the corresponding 68% confidence interval that does not include zero [11].

Table 7. Short-run effects of public capital (individual region models)

Region	*Private capital*	*Employment*	*Real GAV*
Andalucía	0.12 *	0.59 *	0.59 *
Aragón	0.34 *	0.49 *	–0.27 *
Asturias	–0.10 *	–0.25	–0.49 *
Baleares	0.01	0.45 *	0.46 *
Cantabria	–0.21 *	–0.01	–0.09
Castilla-León	–0.23 *	0.18	–1.05 *
Castilla-La Mancha	0.09 *	0.35 *	0.93 *
Canarias	0.37 *	0.60 *	0.73 *
Cataluña	–0.14 *	–0.21 *	0.32 *
Comunidad Valenciana	0.06	0.07	0.29 *
Extremadura	0.05	–0.21	0.11
Galicia	0.10 *	–0.10	0.32 *
Madrid	0.05	0.23	0.58 *
Murcia	0.01	0.41 *	–0.01
Navarra	–0.07 *	0.02	–0.16 *
País Vasco	–0.02	–0.21 *	–0.10
La Rioja	0.04 *	0.52 *	0.26 *

Note: A (*) denotes that the corresponding 68% Hall percentile confidence interval does not include zero. The confidence intervals for individual regions are computed using a bootstrap procedure with 1,000 replications.

[10] They are obtained by dividing the impact or long-run response of private capital, employment, and real GAV to a shock to public capital, respectively, by the impact or long-run response of public capital to a shock to public capital. In the computations, we set the response horizon $T = 25$ (since from the simulations it was possible to verify that for all regions the impulse responses converged to their long-run levels before 15 years) to ensure that for all regions the impulse responses have converged to their long-run levels.

[11] The confidence intervals have been computed using Hall's percentile interval bootstrap procedure described in Breitung *et al.* (2004), and are based on 1,000 bootstrap replications.

Table 8. Long-run effects of public capital (individual region models)

Region	*Private capital*	*Employment*	*Real GAV*
Andalucía	–0.04 *	0.27 *	0.31 *
Aragón	0.32 *	0.01	–0.31 *
Asturias	–0.87 *	–0.65 *	–1.92 *
Baleares	0.66 *	0.20	–0.14 *
Cantabria	–0.15 *	–0.08	0.48 *
Castilla-León	–0.28 *	0.57 *	–0.09
Castilla-La Mancha	–0.15 *	0.02	0.12
Canarias	0.62 *	0.35 *	0.11
Cataluña	0.05	–0.52 *	0.32 *
Comunidad Valenciana	0.18 *	0.48 *	0.59 *
Extremadura	–0.55 *	0.34 *	0.04
Galicia	–0.42 *	–0.32 *	–0.47
Madrid	0.28 *	–0.17	–0.07
Murcia	0.27 *	0.51 *	0.83 *
Navarra	0.11	0.15 *	0.15 *
País Vasco	–0.44 *	–0.46 *	–0.43 *
La Rioja	0.15 *	0.32 *	0.20 *

Note: A (*) denotes that the corresponding 68% Hall percentile confidence interval does not include zero. The confidence intervals for individual regions are computed using a bootstrap procedure with 1,000 replications.

Overall, the estimated effects suggest a highly significant pattern of responses of regional private capital, employment and output to innovations in public capital located in the region itself. The regional effects of innovations in public infrastructures on output, employment and private capital are now considered.

Starting from the effects on output, the short-run real GAV effects of public capital (Table 7) show significantly positive responses in nine of the seventeen cases. This output response is statistically significant and negative in four regions located in the medium-upper zone of Spain (Aragón, Asturias, Castilla-León and Navarra), whereas four regions have no significant output responses (Cantabria, Extremadura, Murcia and País Vasco). For these regions exhibiting negative output responses, a possible explanation is that labor and private capital are altered by the rising stock of public capital. In other words, public capital and private capital could be substitutes in the short run, crowding out employment.

Regarding the long-run responses of output to a shock to public capital installed inside the regions (Table 8), the general pattern is similar to the short-run responses: the results show that seven responses are significant and positive, four responses are significant and negative (Aragón, Asturias, Baleares and País Vasco), and six cases are not significant. The new steady state shows that, as in the case of the short-run, Aragón and Asturias have negative responses on output.

The results reported in Tables 7 and 8 also show that all the significant and positive short- and long-run output elasticities are smaller than 1, indicating that an increase in public capital of a one percent will imply a less than one short- or long-run increase in the real GAV. The more than proportional negative output effects of public capital in Castilla-León (in the short term) and Asturias and País Vasco (in the long run) may be explained by the substitution effect of public capital on private output in these regions, accompanied by a negative elasticity of employment in the last two regions.

As general conclusion, the results would indicate that public capital is productive for most regions, indicating that public capital and private capital are complements in the long-run. Comparing our estimates with those (long term) reported in Pereira and Roca-Sagalés (2003), and considering both significance and sign of the elasticities, the present study only has the same results in 7 of the 17 cases; specifically in the cases of Andalucía, Asturias, Cantabria, Cataluña, Comunidad Valenciana, Galicia and Murcia. This lack of consensus between these results could be explained by two factors: the use of a different sample (1970-1995 in the case of the cited reference and 1972-2000 in the present paper) and a different methodology (in this paper VEC models in levels are used to produce consistent estimates of impulse responses, whereas in Pereira and Roca-Sagalés VAR models in first differences are used which might produce —due to the non consideration of cointegration properties in the estimated systems— inconsistent estimates of impulse response functions).

As regards the short-run responses of employment to a shock to public capital (Table 7), there are only two regions for which the short-run effects of public capital are negative and significant: Cataluña and País Vasco. In the rest of the regions, seven regions have significant and positive short-run effects, while eight regions have no significant effects. In the long run (Table 8), the results indicate that public capital and employment are complements (significant and positive effects) for eight regions and present substitute characteristics for four regions, while the rest (five regions) have no significant effects.

The estimates for private capital elasticities are less conclusive, since in the short-run they are positive for six regions and negative in the case of five regions. For the rest of the regions, these short-run measures are not statistically significant. In the long-run, the pattern is similar: significant and positive elasticities in the case of seven regions, significantly negatives in the case of eight regions, and no statistically significance in the rest of the remaining two regions. This would indicate that private capital and public capital could act as both complements and substitutes in the long-run.

In summary to this point, the long-term effects of public capital formation installed inside the Spanish regional system could lead to an increase in the long-run in both the regional real GAV and the regional employment. Nevertheless, if the aim is to increase private capital in the long-run, there is no empirical evidence that an increase in public capital would generate the required response from the private sector.

3.5. Discussion

From the empirical literature, the impact of public capital on private capital is complex and uncertain. From a theoretical perspective, and in the framework of a production function where the public capital stock is introduced as a separated input, it is possible to establish three different relationships between public capital and private capital. In short, they could be complementary, independent or direct substitutes (see, for example, Ramírez, 2000).

If public capital is complementary to private capital, an increase in public capital will increase output directly. In addition, public capital will increase private capital investment directly while public capital will increase output indirectly (stimulating positively the marginal productivity of the private capital stock). Finally, as public capital increases the amount of both private and public capital per worker, the marginal productivity of labor increases, increasing output.

In the case where public capital and private capital are independent, an increase in public capital will generate a positive effect on output and the marginal productivity of labor in the public sector only.

If public and private capital are direct substitutes, an increase in public capital formation will raise output directly. Nevertheless, there will exist a negative effect on the marginal productivity of private capital and labor that could counterbalance the positive effects.

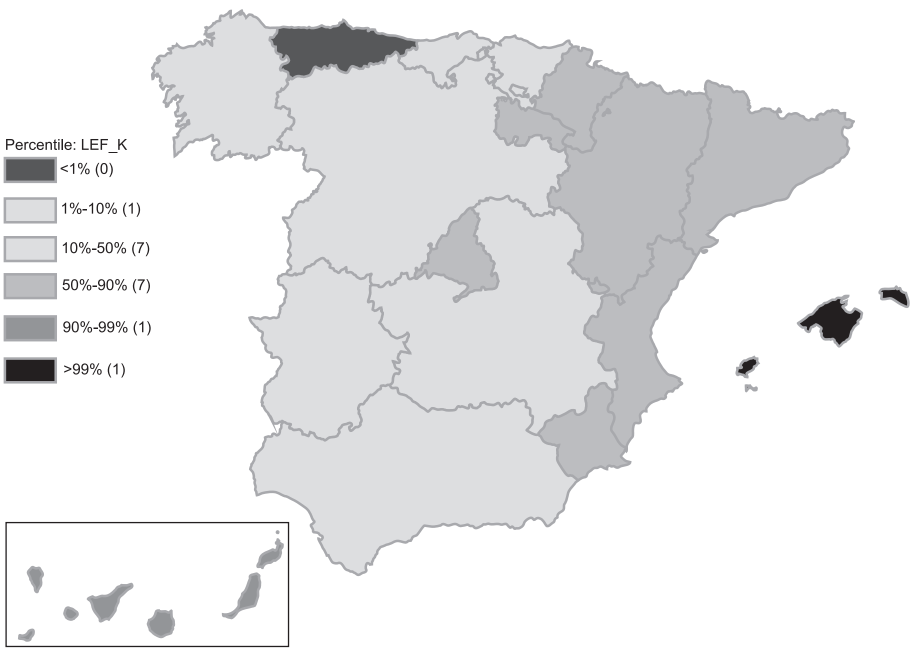
Under the aforementioned relationships, we can say that public capital is complementary to private capital in 3 Spanish regions (Comunidad Valenciana, Murcia, and Rioja); public capital is independent to private capital in 2 Spanish regions (Cataluña and Navarra), and there is a direct substitution effect for the case of 8 Spanish regions (Andalucía, Asturias, Cantabria, Castilla-León, Castilla-La Mancha, Extremadura, Galicia and País Vasco). Finally, for the rest of the Spanish regions (Aragón, Baleares, Canarias and Madrid), from our results it is not possible to classify the type of relationships between public and private capital.

The empirical findings of this paper would suggest that increases in public capital in core Spanish regions would raise the marginal productivity of private capital thereby inducing higher rates of private investment spending. On the other hand, public capital investment in peripheral regions can be substituted directly for private capital investment. These results for peripheral regions could retard future regional economic growth. Effectively, the detected crowding out effects could act as a penalty in peripheral regions if they operate in key sectors of the regional economy such as basic industries and agriculture.

Another additional goal of this discussion is to enlarge the empirical analysis of the detected effects by means of the consideration of the spatial dimension. In this sense, the geographic dimension of the different estimated effects were explored by using an exploratory spatial data analysis (ESDA) approach. This analysis will help with the identification of the type of spatial pattern present in the distribution of regional effects. All computations were carried out by using SpaceStat 1.91 (Anselin, 2002), GeoDA (Anselin, 2003) and ArcView GIS 3.2 (ESRI, 1999) software packages. First, global spatial autocorrelation was tested by using Moran's I statistic (Cliff and Ord, 1981), $I = \frac{N}{S_0}\frac{z'Wz}{z'z}$, where N is the number of regions, $S_0 = \sum_i \sum_j w_{ij}$, z_{it} is the effect of public capital in region i for the t cases considered in deviation from the mean, W was defined expressing for each region (row) those regions (columns) that belong to its neighborhood. Formally, $w_{ij} = 1$ if regions i and j are neighbors, and $w_{ij} = 0$ otherwise. This simple contiguity matrix ensures that interactions between regions with common borders are considered. For ease of economic interpretation, a row-standardized form of the W matrix was used. Thus, the spatial lags terms represent weighted averages of neighboring values.

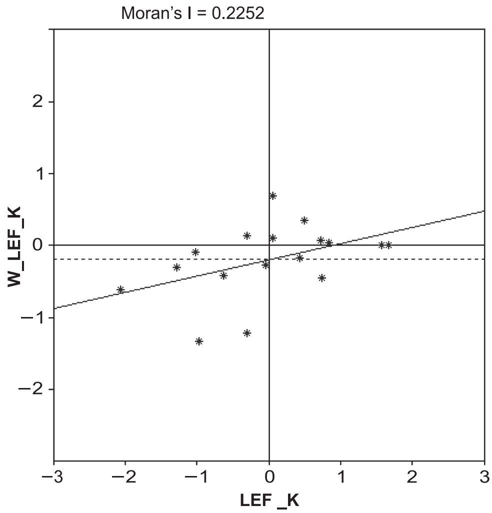
The values of I for five of the six different effects were well below the expected value for this statistic under the null hypothesis of no spatial correlation. It appears that these effects are not spatially correlated, since their statistics are not significant. Nevertheless, for the case of the long-run effects of public capital on private capital, the Moran's I reveals the existence of a strong and statistically significant degree of positive spatial dependence in the distribution of regional effects. Figure 2 shows the spatial distribution of long-run effects of public capital on private capital. Figure 3 provides a clearer view of the spatial autocorrelation in these regional effects through the Moran scatterplot, showing a strong geographic pattern and revealing the presence of positive spatial dependence.

Figure 2. Long-run regional effects from public capital on private capital



Note: LEF_K denotes long-run regional effects from public capital on private capital.

Figure 3. Morans' I of long-run regional effects from public capital on private capital



Note: LEF_K denotes long-run regional effects from public capital on private capital; W_LEF_K denotes the spatial lag of LEF_K. For the calculated Moran's I, *p*-value = 0.024.

Both figures show a strong geographic pattern, revealing the presence of spatial heterogeneity in the form of two spatial clusters of rich and poor regions, with the rich regional economies' cluster including the regions within the triangle area comprising the axis País Vasco-Cataluña, Cataluña-Valencia and Valencia-País Vasco plus the capital, Madrid, and the islands (Baleares and Canarias); whereas the rest of the regional system could be characterized as the Spanish «periphery» with less economic activity and a much lower level of per capita income.

4. Summary and conclusions

The effects of public capital on economic growth have received a great deal of attention in the recent economic literature. Within the approaches that have been applied to assess the impact of public infrastructures, this paper estimates the dynamic domestic effects of innovations in public capital using a structural vector autoregressive (S-VAR) methodology for the Spanish regions.

From a methodological point of view, the work contains different innovative features with respect to the previous studies using S-VAR models. First, recently developed panel integration and cointegration tests are used to examine the long-run determinants of aggregate regional production. Thereafter, using a two-step approach (*a la* Engle and Granger, 1987) the detected cointegrating relation is first estimated and then the residuals from the long term relationship are used to estimate individual region-specific structural vector error-correction (S-VEC) models. Thus, the domestic dynamic properties of the estimated S-VEC models are investigated via impulse response functions that portray the effects of shocks to the public capital installed in one region on the rest of variables of the region. As a general conclusion, the long-term effects of public capital formation installed inside the Spanish regional system could lead to an increase in the long-run in both regional real GAV and employment. Nevertheless, if the aim is to increase private capital in the long-run, there is no empirical evidence about the appropriateness of stimulating private capital through an increase in public capital as an adequate policy measure. In the short-run, private capital and public capital could act as both complements and substitutes, although employment seems to receive a predominantly positive stimulus in the short-run from public capital formation.

From these estimates, the direct substitution effects prevail for the peripheral regions. Thus, more precise indications for policy-making can come from further research on the underlying reasons as to why these effects happen. The findings in this paper suggest that regional policy makers would have to implement regional measures where the increases of public capital do not imply negative effects on private capital.

Finally, this paper considers that there exists cross-sectional independence, which probably is not the case. Further analysis on this issue could be conducted in the future using extended versions of the class of VAR models applied in the present work. The natural extension would be to formulate a «Global VAR» model for the Spanish regional system. This would combine all the S-VAR models in a global specification in which the state variables of each region would be related to the state variables of the rest of the regions (see Pesaran *et al.,* 2004, and Dees *et al.,* 2007). Related to this, if there exists spatial dependence in the data, it would be more appropriate to use a «Second Generation» approach in the unit-roots and cointegration analysis of section 3, which assumes the existence of cross-sectional dependence (Breitung and Pesaran, 2008).

Also, as stated by a referee, another extension for the future could be to split public capital into its two main components: i) transport infrastructure and ii) the rest.

Acknowledgements

The comments of the referees and the editor on earlier versions are greatly appreciated. The authors acknowledge and appreciate the funding received from the Ministry of Science and Innovation of Spain through the project ECO2009-12506.

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