Analysing the impact of public capital stock using the NEG wage equation: a panel data approach

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Abstract:

This paper examines the relationship between the level of public infrastructure and the level of productivity using panel data for the Spanish provinces over the period 1985-2004, a period which is particularly relevant due to the substantial changes occurring in the Spanish economy at that time. The underlying model used for the data analysis is based on the wage equation, which is one of a handful of simultaneous equations which when satisfied correspond to the short-run equilibrium of New Economic Geography (NEG) theory. This is estimated using various spatial panel models with either fixed or random effects to allow for individual heterogeneity. Using these models we find consistent evidence that productivity depends directly on the public capital stock endowment of each province, but also there is evidence of negative spillover effects from changes in capital stock in neighbouring Provinces.

Keywords: Public capital, New Economic Geography, spatial econometrics.

JEL Code: C23, R11, R12.

1. INTRODUCTION

The World Bank has referred to public capital as one of the “wheels of economic growth”. This claim appears to be one that could be the basis for much empirical research, for it is evident that nowadays the role of public investment is very much in the spotlight as a possible way out of the current global economic downturn. Since the beginning of the current economic crisis in August, 2007, many renowned economists and institutions have suggested the need for an expansive fiscal policy to alleviate the worldwide economic recession.

From a theoretical point of view there are several channels through which public investment affects regional per capita income. Macroeconomists typically emphasize three “conventional” channels through which public infrastructure may affect growth. Public investment has a direct productivity effect on private production inputs and a complementarity effect on private investment. However, an increase in the stock of public capital in infrastructure may have an adverse effect on activity, to the extent that it displaces (or crowds out) private investment. So, despite the direct and complementarities effects, the net effect of an increase in public infrastructure may well be to hamper, rather than foster, economic growth. The importance of each effect might depend on the initial stock of the economy, the diversity of productive structures and the degree of maturity of infrastructure systems.

The empirical literature on the effects of public infrastructure is inconclusive; studies are divided on both the magnitude and direction of the net effect of infrastructure on economic growth. The first author to detect a positive relationship between public infrastructure and productivity was Ratner (1983), though it was Aschauer (1989) who established that a decline in the rate of public investment in infrastructure during the 60´s could be one of the causes of the productivity slowdown in the United States in 1970s and 1980s. More recent studies have partially discredited the results obtained in early research. Many researchers agree that the apparently positive impact of public capital stock might be due to inadequate model specifications which cause spurious relations or fail to appropriately control for region or country heterogeneity.

Most empirical analyses use neoclassical production functions to quantify the public infrastructure effect on economic activity. Using this approach, some papers established a positive effect, including Munell (1990, 1993), Ford and Poret (1991), Bajo-Rubio and Sosvilla-Rivero (1993), Otto and Voss (1994), Mas *et al.* (1996) and Cantos *et al.* (2005). Others established a negative effect, such as McMillin and Smyth (1994), Otto and Voss (1996) and Voss (2002) and others find no significant effect of public investment on economic activity; this is true of Tatom (1991), Batina (1999), Evans and Karras (1994), Baltagi and Pinnoi (1995), García-Mila and McGuire (1992) and Gómez-Antonio and Fingleton (2008), among others.

Another approach in the infrastructure literature examined the impact of infrastructure on employment. Studies such as Munell (1990), Haughwout (1999) and Dalenberg and Partridge (1995) have found that public infrastructure is positively associated with the level of employment. However, under this approach we are unable to identify whether greater employment is caused by an increase in labour demand because firm production has increased, or by an increase in labour supply because there are more household amenities. Dalenberg and Partridge (1997) use wages as a dependent variable to identify whether infrastructure plays a greater role as an unpaid factor for firms than household amenities. They find that highway infrastructures dominate.

Many researchers had for a long time insisted that increasing returns were essential for a proper understanding of spatial disparities in economic development, but this was given new impetus by the development of a formal theoretical framework, based on a monopolistic competition market structure model (Dixit and Stiglitz, 1977). This development of increasing returns based on micro-economic foundations, by proponents of New Economic Geography (NEG) and related models in Urban Economics, led to the integration of increasing returns models within mainstream economics. However, while a few papers use urban economics as a theoretical basis for analysing the relationship between public investment and economic growth, like Martin and Rogers (1995) and Martin (1999), to our knowledge, no other study has been based on NEG theory.

The main aim of this paper is to test Aschauer's hypothesis and quantify the impact of public capital stock on productivity, something which, as far as we know, has not been attempted in the context of NEG theory. This paper differs from previous literature by being based not on a strictly neoclassical production function but on the theoretical arguments of New Economic Geography. It is worth noting that our model specification allows us to estimate the effects of spillovers operating across geographical space. Most of the literature on public capital has focused on whether or not public infrastructure has positive productivity effects, but relatively little attention has been paid to the fact that the presence of public capital may shift economic activities from one location to another.

The Spanish provinces[1] provide an interesting case study as it has undergone a sustained period of growth in the last 40 years, together with a large increase in public investment. The period analysed in this paper is 1985-2004, which is particularly relevant due to substantial changes in the Spanish economy during this time. At the beginning of the period, the level of government capital endowment and economic activity in the Spanish regions were far below those of other European economies. Since Spain joined the European Union, however, there has been a very intensive period of capital investment by the Spanish government with, until the current slump, no perceptible slowing of investment due to economic cycles.

The results show that changes in provincial productivity are positively associated with changes in public investment within the same province but negatively associated with such changes in other regions. This is consistent with possible predictions of NEG theory. As is stated in Puga (2002), one should not forget that roads generally have lanes going both ways. A better connection between two regions with different development levels not only gives firms in a less developed region better access to the inputs and markets of more developed regions. It also makes it easier for firms in richer regions to supply poorer regions at a distance, and can thus harm the industrialisation prospects of less developed areas. Likewise, using NEG theory, Roberts et. al. (2010) find that road transport investment in China has thus far mainly benefitted the larger cities and more developed regions. Previous papers, like Boarnet (1998) for the Californian counties and Moreno *et al.* (2007) for the Spanish provinces, also found a negative spillover effect of public infrastructure investment under different theoretical approaches.

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[1] Spanish provinces correspond to level 3 of the Nomenclature of Territorial Units for Statistics (NUTS) of EUROSTAT, the Statistical Office of the European Union. The average surface of a representative province is 10,120 km^2 (range 1,980 km2 to 21,766 km^2).

The paper is organized as follows. Section 2 briefly sketches the theoretical model; section 3 is concerned with issues related with the data. Section 4 details the empirical model and its estimation, and finally section 5 gives some concluding remarks.

2. THE THEORETICAL MODEL

The theory used here is set out by Fujita *et al.* (1999), which is the basic two-region two-sector core-periphery model. The NEG model reduces to five simultaneous non-linear equations[2]. The first one, and the one we use in this paper, involves a simple relationship between the industrial sector M nominal wage level (w_r^M) and the market potential variable (P_r).

$$w_r^M = \frac{\bar{W}_r^M}{E_r^M} = \left[\sum Y_r (G_r^M)^{\sigma-1} (\bar{T}_{ir})^{1-\sigma} \right]^{1/\sigma} = P_r^{\ 1/\sigma} \tag{1}$$

In equation (1) r denotes region, $\bar{W}_r^M$ is area r's total M wage bill, E_r^M is the M workforce, and the summation is over the set of regions including r. The transport cost between region i and r is denoted by $\bar{T}_{ir}$, G_r^M is a price index in M, Y_r is the income and σ is the elasticity of substitution for M varieties. Following Fujita *et al.* (1999), the M price index is given by

$$G_r^M = \left[\sum \lambda_r (w_r^M \bar{T}_{ir})^{1-\sigma} \right]^{1/(1-\sigma)} \tag{2}$$

in which the number of varieties produced in region r is captured by λ_r , which is equal to the share in region r of the total supply of M workers. Income is given by

$$Y_r = \theta \lambda_r w_r^M + (1-\theta)\phi_r w_r^C \tag{3}$$

[2] This model is described in detail in the cited references so we do not include technical details. However an annexe is available upon request to the authors.

Y_r depends on the nominal M sector wage rate w_r^M, the competitive sector C wage rate w_r^C, the number of M varieties λ_r and the equivalent number of C varieties ϕ_r. It also depends on θ which is the total number of M workers adding across all regions, although $0 < \theta < 1$ since it is measured on a scale such that the overall number of workers in the economy is equal to 1. Consequently the total number of C workers in the economy is $1 - \theta$.

Taking logs to equation 3 the basic wage equation is

$$\ln w_r^M = \frac{1}{\sigma} \ln P_r \tag{4}$$

We will assume that wage rates also depend on factors other than market potential, namely the level of efficiency of workers (A_r). Our initial assumption is that efficiency depends on the extent of upper level education, on the public capital stock endowment and on private capital within each province. Introducing these additional variables means that we have a departure from the traditional reliance on purely pecuniary externalities.

The extended wage equation becomes[3]:

$$w_r^M = \frac{\overline{W}_i^M}{E_i^M} = P_r^{1/\sigma} A_r \tag{5}$$

In addition, we also assume that the level of efficiency within a given province is related to the level in other provinces that are nearby, due to spillover effects across space. This means that the level of efficiency is dependent on covariates X and on 'nearby' efficiency levels denoted by the matrix product *Wln(A).* Written in general matrix notation, the vector for efficiency level is

[3] Other researchers have also considered additional variables to extend the basic NEG equation, as for example Combes and Lafourcade (2005). This equation, as shown by Head and Mayer (2006), can be obtained from micro assumptions, by introducing a labour quality adjusted production function for the firm.

$$\begin{aligned} &\ln A = \rho W \ln A + Xb + \xi \\ &\xi \sim N(0, \Omega^2) \\ &\ln A = (I - \rho W)^{-1}(Xb + \xi) \end{aligned} \tag{6}$$

In which X is an n by k matrix of exogenous variables (with columns equal to 1, and variables *lnH*, *lnK* and *lnV*), b is a k by *1* vector of coefficients, the matrix product *Wln(A)* is an n by *1* vector with scalar coefficient ρ, and vector ξ represents excluded variables which behave as random shocks. Variable H is a measure of human capital, variable K denotes public capital stock whereas V is a measure of private capital. A higher level of human and private capital will enhance the efficiency of labour, and likewise it is assumed that labour efficiency will be increased by superior public capital stock in the form of better transport infrastructure, and publicly provided services such as water, electricity and health services. The endogenous variable *Wln(A)* represents the additional contribution to efficiency which is assumed to be due to 'nearby' provinces. The hypothesis is that regions with high labour efficiency levels occurring in neighbouring regions will also incur higher labour efficiency than would otherwise be the case, and vice versa for regions with lower labour efficiency. The implication of this is that labour efficiency in distant regions will have less impact, so that "who your neighbours are" is important. We considered different alternatives when selecting the matrix W, but preferred to adopt a first order binary spatial contiguity matrix in which the elements are one when provinces share a common border, and zero otherwise[4].

[4] With rows summing to 1 and the elements of the main diagonal are set to zero by convention. Other definitions, based on the quantitative 'distance' between the different provinces, as well as the (economic) size of provinces were also tried, but these involved major assumptions also and resulted in less well-fitting models.

On taking logs of equation (5), ignoring subscripts, writing $a_1 = \frac{1}{\sigma}$, substituting for *lnA* and rearranging terms we obtain,

$$\begin{aligned} &\ln w^M = \rho W \ln w^M + a_1(\ln P - \rho W \ln P) + a_0 + a_2 \ln H + a_3 \ln K + a_4 \ln V + \xi + (I - \rho W)\omega \\ &\xi \sim N\left(0, \Omega^2\right) \end{aligned} \qquad (7)$$

in which the log of *M* sector productivity is denoted by $\ln w^M$, which is an *n x 1* vector at time *t*, *W* is the *n x n* standardised contiguity matrix, so that on multiplication the resulting *n x 1* vector $W \ln w_t^M$ is the spatial lag of $\ln w^M$, P_t is market potential, and $W \ln P_t$ is the spatial lag of market potential. Also the *n x 1* vector $\ln H_t$ is the log of our measure of human capital, and equivalently $\ln K_t$ denotes log public capital stock and $\ln V_t$ is log private capital stock. Finally the specification includes an *n x 1* vector of errors ξ_t plus the moving average error process[5] given by $(I - \rho W)\omega_t$. The constant is retained to allow for autonomous wage growth due to unexplained productivity increases that are constant across provinces.

3. DATA

For each year we represent w_r^M by province *r's* gross value added (GVA) in industry sectors (including building and energy activities) divided by *r's* industrial employment. Data which were provided by FBBVA (*La Renta Nacional de España y su Distribución Provincial*) until 1997, and thereafter by Fundación de las Cajas de Ahorro Confederadas (FUNCAS)[6] as documented in "*Balance Económico Regional"*[7].

[5] In practice, for simplicity, we will drop the assumption of a moving average error process in favour of independent identically distributed errors, assuming they are equal to $u \sim N(0, \sigma^2 I)$.

[6] In order to make the Gross Value Added and employment series homogeneous we took the rates of growth of the variable in FUNCAS database and applied it to the variable produced by FBBVA. Previously

Our human capital variable (H) is the proportion of people in each province with higher education, data published in "*Human Capital in Spain and its distribution by provinces (1964-2004)"* by *Instituto Valenciano de Investigaciones Económicas* (IVIE).

The public capital stock (K) was taken from the publication "*Series Historicas de Capital Publico en España y su distribucion territorial 1900-2005"* which detailed work done by FBBVA in collaboration with IVIE. And the private capital variable (V) was taken from "*El stock de capital en España y su distribución territorial 1964-2002*" which detailed work done by FBBVA[8].

The Market Potential variable (P) is constructed using the kernel of equation (1), which we repeat here with the assumed trade cost function in place, hence

$$\begin{aligned} \overline{T} &= e^{\tau D_{ir}} \\ P_r &= \sum_r Y_r (G_r^M)^{\sigma-1} (e^{\tau D_{ir}})^{1-\sigma} \end{aligned} \tag{8}$$

This is a function of the M sector price index, which with the trade cost function is $G_r^M = [\sum_r \lambda_r (w^M e^{\tau D_{ir}})^{1-\sigma}]^{\frac{1}{1-\sigma}}$. In order to calculate trade cost $e^{\tau D_{ir}}$, we use straight line distances between Provincial capitals, denoted for origin i and destination r by D_{ir}, scaled by τ which is equal to 0.001. Distances within provinces are estimated using the convention that

we had to transform the valued added into constant euros of 2000 using the Implicit Index Prices facilitated by both organisations.

[7] GVA is measured in thousands of constant (2000) Euros and we attribute the difference between wages and GVA per worker to measurement error, which is represented by ω in equation (7).

[8] Data are in constant 1990 units. Due to the lack of data for the last four years of the period we assumed that private capital remained at its 2002 level over these four years. Alternatively we could have assumed that private capital grew at the same rate as public capital, or at the rate at which it had been growing in recent years, however in each case we would have been introducing measurement error. We test the significance of private capital using instrumental variables below, and are satisfied that the conclusions we reach would not have been different under different assumptions for these omitted years of data.

$$D_{ii} = \frac{2}{3}\sqrt{\frac{area_i}{\pi}} = \frac{2}{3}R \tag{9}$$

where $area_i$ is the number of square km in province *i*. This is 2/3rds of the radius *R* of a circle with area equals to that of province *i*, and is equal to the average distance from the centre of all points within a circle. The trade cost function produces non-negative scalars greater that 1, which are multiplied by the wage rates to allow for the additional cost of transport within and between provinces in calculating the price index.

In order to calculate the income given by equation (3), we also need the wage w_r^C, which represents the average wage in non-industrial *C* sectors. Since the *C* sector incurs no trade costs, this is constant across provinces in any one year. It is obtained as the sum of non-industrial GVA divided by sum of non-industrial employment, in both cases summing across provinces in any one year. Also the total number of *M* workers in the whole economy θ is taken as the overall share of total employment in each year that is engaged in *M* activities, which is equal to the total industry employment divided by total employment. The quantity λ_r is equal to province *r's* share *of* the total number of *M* (industry) workers in the economy. Likewise ϕ_r is equal to the proportion in province *r* of the total number of workers in non-industrial sectors in the economy.

The final quantity needed to calculate P_r is the elasticity of substitution of *M* sector varieties σ. Rather than estimate this value, it is assumed to equal 6.25, which is a central value among the range of estimates provided by the literature[9], and equal to the mid-point of the range given by Head and Mayer (2003). Neither of the two main alternative

[9] Head and Ries (2001) estimate values ranging from 7.9 to 11.4, the range is 5 to 10 in Harrigan (1993), 3 to 8.4 in Feenstra (1994), and there are point estimates of 9.28 in Eaton and Kortum (2002) and 6.4 in Baier and Bergstrand (2001).

approaches to obtaining the value of σ is feasible in the current context. One is direct nonlinear estimation, as carried out for example by Mion (2004) and Brakman *et al.* (2006), but this would be difficult to operationalise given the iterative estimation methods used here. The other alternative is the two-step linear estimation approach of Redding and Venables (2004), but this relies on bilateral trade flows which are unavailable for Spain's provinces. Anderson and van Wincoop (2004) summarize various estimates, which are largely within the range 5 to 10.

4. RESULTS

The advantages of panel models are well known, most significantly they allow one to control for individual-specific heterogeneity and the simplest way to do this is by introducing fixed effects (using either dummy variables or equivalently mean deviations, to allow for different intercepts). Alternatively, we introduce spatially autocorrelated random-effects, allowing for inter-province heterogeneity via an individual-specific random component in the disturbance term. We also include fixed year effects in some of our specifications. Our most complex model is equation (7), which for year t is

$$\ln w_t^M = \rho W \ln w_t^M + a_1(\ln P_t - \rho W \ln P_t) + a_0 + a_2 \ln H_t + a_3 \ln K_t + a_4 \ln V_t + \xi_t + (I - \rho W)\omega_t$$

All other models fitted are nested within this specification, as a result of setting parameter ρ and the a parameters to zero[10].

[10] Our models 1 to 4 contain an endogenous spatial lag. For the 'standard' SAR(1) spatial lag model, and associated with the problem of endogeneity, Pinkse and Slade (2010) are concerned about weak identification, most typically due to weak instruments. Kelejian and Prucha(1998) give the rationale for the use of low order exogenous lags as ideal instruments. We use the first order lags. In models 2 to 4. Also for models 5 to 10 we avoid weak instruments. McMillen (2010) is concerned about incorrect functional form and omitted variables, calling for alternatives to the spatial lag model and its usual estimators. For models 1 to 4 this is not an issue since we test a causal relationship within an explicit parametric model, as given by equation (6). However we agree that generally there is insufficient theory for an exact W matrix structure, and results that are robust to alternative specifications are preferable. In our case we obtain estimates from a variety of estimation methods that support the overall thesis of the paper. Additional consideration of issue of endogeneity is also given in Fingleton and Le Gallo(2007, 2008, 2009).

Model 1 in table[11] 1 estimates our closest approximation to the equation (7) specification. The estimates were obtained using the ML approach initially developed by Elhorst (2003) with the likelihood function accommodating the endogeneity of the spatial lag $W \ln w_t^M$. Note also that we have allowed for the parameter restrictions involving ρ, using an iterative routine to ensure that the parameter equalities are satisfied. The iterations commence with $[\ln P_t - \rho_1 W \ln P_t]$ defined by assuming an initial value of ρ denoted by ρ_1, which allows subsequently estimation as the coefficient on $W \ln w_t^M$. In the next iteration this estimated $\rho = \rho_2$ gives an update of vector $[\ln P_t - \rho_2 W \ln P_t]$, leading to another estimate $\rho = \rho_3$, and so on. The iterations terminate at iteration k where $\rho_k - \rho_{k-1} < 0.00001$.

Table 1: Iterative Estimates[1] using ML and 2SLS (T=20, n=47)

REGRESSORS	Parameter Estimates			
	Model 1*	Model 2[#] 2SLS	Model 3[#] 2SLS	Model 4[#] 2SLS
Productivity Spatial Lag [$Wlnw^M$]	0.013 (0.27)	-0. 132 (-0.41)	0172 (0.19)	-0.166 (-0.43)
Market Potential [lnP-$\rho WlnP$]	0.312 (11.10)	0. 301 (5.13)	0.323 (6.37)	0. 292 (5.11)
Public Capital [lnK]	0.123 (4.59)	0. 111 (2.71)	0.124 (4.27)	0. 107 (2.34)
Private Capital [lnV]	-0.049 (-3.65)	-0.047 (-3.24)	-0.050 (-3.33)	-0.045 (-2.52)
Human Capital [lnH]	0.0406 (4.38)	0. 0415 (4.16)	0.0410 (4.10)	0. 041 (4.08)
R squared	0.957	0.853	0.854	0.852
Error variance	0.0018			
Squared Correlation[2]	0.958			
Estimation Method	ML	Instrumental variables	Instrumental variables	Instrumental variables

[11] For simplicity, table 1 omits the estimates of the time and individual fixed effects, focussing on the variables of substantive interest.

Instruments	*lnP_WlnP* *lnK* *lnV* *lnH* *WlnK* *W lnV* *WlnH* *Temporal dummies*	$W \ln w_t^M$ *lnK* *lnV* *lnH* *WlnK* *W lnV* *WlnH* *Temporal dummies*	*lnK* *lnV* *lnH* *WlnK* *W lnV* *WlnH* *Temporal dummies*

Notes:
*t-ratios given in brackets beneath the estimates
z-ratios given in brackets beneath the estimates
[1] Spatial and time period fixed effects included in each model
[2] Between fitted and actual productivity
Time sub indexes have been omitted from the table and from the comments to simplify notation.
Model 2: 2SLS instrumenting endogenous spatial lag
Model 3: 2SLS instrumenting compound market potential variable
Model 4: instrumenting both endogenous spatial lag and compound market potential variable

The construction of P_t introduces two-way causation involving w^M, in addition P_t may possess measurement error for two reasons. One is that the definition of the *M* sector may not be exact, and second we are assuming that the elasticity of substitution σ is equal to 6.25. In all subsequent models, we use instrumental variables for P_t to allow for endogeneity and to satisfy error distribution assumptions. On the other hand we retain an assumption throughout of exogeneity for the other variables. In the case of human capital (*H)* this is because we assume complex determinants of educational attainment levels, so that any direct feedback from wage levels will be relatively weak. We note that in Spain inter-province migration in response to wage differentials is not strong. In the case of Public capital stock (*K)*, the assumption is that this is mainly controlled by government policy and this is not driven by wage levels. Likewise private capital (*V*) is a response to a multiplicity of variables, so it would be unlikely that its spatial distribution is a response to wage level variations. We test these assumptions explicitly in the context of model 6 below. In this the estimates are compared with strictly consistent estimates as a result of instrumenting all six variables, and the outcomes indicate no significant difference between the two sets of estimates, supporting the assumption that the variables are exogenous. The Hausman exogeneity test shows that the test statistic is not an extreme value with reference to the relevant χ^2 distribution under the null, being equal to 5.53, which has a p-value of 0.4783 in the χ^2_6 distribution. The estimates presented in table 1 suggest that there are highly significant and direct (within-province) effects due to market potential, public capital stock and human capital. Interestingly, there is evidence of a significant negative effect due to private capital, although the elasticity is comparatively small, and in our subsequent models that there is no significant private capital effect. The model includes spillover effects in the form of the endogenous lag $W \ln w_t^M$ and the suggestion here is that there are no inter-Provincial spillovers of worker efficiency levels (A) *per se*. We next turn to modelling spillovers by means of exogenous spatial lags, as exemplified by equation (10),

$$\ln w^M = a_1 \ln P + a_0 + a_2 \ln H + a_3 \ln K + a_4 \ln V + a_5 W \ln H + a_6 W \ln K + a_7 W \ln V + \psi \qquad (10)$$

This builds on the fact that most of the literature on the effects of public investment focuses on whether or not infrastructure has productive effects and pays relatively little attention to how public, private or human capital might shift economic activity from one place to another. There are reasons to believe that capital investment in neighbouring provinces might have a positive impact on productivity within a given province. Public capital investment might enhance connections such as roads, railways or airports. On the other hand, negative spillovers might exist perhaps due to the migration of factors to locations with superior infrastructure stocks. Public investment in one region could have a negative effect on other regions that are its closest competitors for labour and mobile capital. It is argued that spillovers might be one of the causes behind the different effects of public investment on productivity and could explain why studies of national time-series data typically find larger output elasticities for public capital than elasticities estimated by papers that analyse regional data. The existence of spillovers could be responsible for the apparently low impact of public investment on regional productivity. The estimates are given in table 2.

Table 2: Fixed effects models with exogenous lags (T=20, n=47)

REGRESSORS	Model 5#	Model 6#
constant	-------	-1.595718 (-3.26)
Market Potential [*lnP*]	0.3179722 (1.48)	0.1728635 (22.54)
Public Capital [*lnK*]	0.2033465 (6.90)	0.2125152 (7.21)
Private Capital [*lnV*]	-0.0552563 (-2.90)	-0.0496475 (-3.45)
Human Capital [*lnH*]	0.039625 (3.00)	0.0322421 (3.60)
Spatial Lag Public Capital [*WlnK*]	-0.3727936 (-6.87)	-0.317303 (-8.95)
Spatial Lag Private Capital [*W lnV*]	0.0342643 (0.32)	0.084841 (3.50)
Spatial Lag Human Capital [*WlnH*]	0.0260104 (0.56)	------
Time dummies	Yes	no
R squared	0.8628	0.857
Wald test statistic	10.94 ($p = 0.0009$)	17.48 ($p < 0.0001$)
Estimation Method	Instrumental variables	Instrumental variables
Instruments (including exogenous regressors)	*lnK* *lnV* *lnH* *WlnK* *W lnV* *WlnH* *Time dummies* *lnV_c*	*lnK* *lnV* *lnH* *WlnK* *W lnV* *Time dummies* *lnV_c*

lnH_c

Notes:
[#] z-ratios given in brackets beneath the estimates

The table 2 estimates are for fixed effects panel models. The model 5 estimates show that the variable market potential (*P*) is correctly signed but is not significant, it does indicate that high levels of market potential correspond to higher wage rates (one-tailed p value is 0.0694). A possible reason for the insignificance of *P* is the high level of correlation between *P* and the time dummies. Since market potential is clearly endogenous, its effect is identified by two instruments, thus over-identifying and allowing a test of over-identifying restrictions (orthogonality conditions). The two instruments are dummy variables *lnH_c* and *lnV_c* , which equal 0 if *lnH* and *lnV* are at most equal to their means, and are equal to 1 otherwise[12]. The first stage (within) regression gives an overall R-squared equal to 0.7871, and the F statistic for the regression is 7861.93, which is an extreme observation in the $F_{27,866}$ reference distribution, thus allowing rejection of the null hypothesis that the instruments are unrelated to the endogenous variable. The individual t ratios are -4.47 for *lnV_c* and 0.95 for *lnH_c* . Restricting these two identifying instruments to zero in the first stage regression gives a joint *F*-statistic equal to 10.79, which has a p-value equal to 2.3518e-005 in the $F_{2,866}$ distribution. This reaffirms that we do not have weak instruments. We test the over-identifying restrictions via the Stata command xtoverid (Schaffer and Stillman, 2010). This gives values of the Sargan-Hansen test statistic equal to 0.708, 0.682 (robust), with p-values in the χ_1^2 distribution equal to 0.4001 for the Sargan test and 0.4090 for the Hansen J statistic

[12] Since we are assuming that *lnH* and *lnV* are exogenous, then it is reasonable to assume that their dummy representations will also be exogenous.

respectively. This affirms that our orthogonality conditions have been satisfied. The determinants of labour efficiency within each province (*lnH*, *lnK, lnV*) are statistically significant, and although the exogenous lags *WlnH* and *WlnV* are insignificant, *WlnK* is significant and negative. A test of the significance of the group of time dummies indicates that they are insignificant, given the other variables in the model. Restricting the time dummy parameters to zero gives a test statistic of 22.41, which when referred to the χ^2_{19} distribution has an excedence probability of 0.2645, which is sufficiently large to allow non-rejection of the null that the parameters are zero. Model 6 is the outcome of nullifying these insignificant variables, but retaining those that turn out to be significant when re-introduced. In order to satisfy the assumptions of the test of the consistency of the equivalent random effects specification as discussed below, we slightly change the instrument set. We eliminate the weaker of the two dummies, retaining *lnV_c*, and add the time dummies thus ensuring overidentification. The first-stage within regression remains highly significant with F = 7213.06 which has a near zero probability in the $F_{25,868}$ distribution. Our test of instrument orthogonality gives test statistics equal to 22.283 and 27.222 (robust), both of which have sufficiently high p-values (0.2704, 0.0996) leading us not to reject the orthogonality null.

However we do not consider model 6 to be our final preferred specification. The negative coefficient on private capital does not seem to be appropriate, and the impact of investment in public capital is counterintuitive in the models considered thus far. Consider an increase in public capital in province *i* in a specific year *t*. While our model implies that productivity will increase in *i* as a result, it also implies reductions in productivity for reasons explained above in those regions which are 'near' to *i* as defined by the matrix W. This will involve several regions and can be complex with both endogenous and exogenous spillovers (see LeSage and Pace (2009), Anselin and Le Gallo (2006), Le Gallo, Ertur and Baumont (2003)). In our simpler case with only exogenous spillover, the impact of a unit change in log public capital in *i* on log productivity in *i* at time *t* is given by

$$\frac{d \ln w_{it}^{M}}{d \ln K_{it}} = a_3$$

The effect of a unit change in log public capital in *j* on log productivity in *i* is given by

$$\frac{d \ln w_{it}^{M}}{d \ln K_{jt}} = a_6 W_{ij}$$

And combining these two impact gives

$$\frac{d \ln w_{it}^{M}}{d \ln K_{jt}} = a_3 \iota + a_6 W_{ij}$$

With $\iota = 1$ when $i = j$ and 0 otherwise. Clearly the vector $d \ln w_t^M$ will contain positive impacts for $a_3 > 0, i = j$, negative impacts $a_6 W_{ij}$ (when $W_{ij} > 0, a_6 < 0, i \neq j$) and zero impacts ($W_{ij} = 0, i \neq j$). In fact the Wald tests reported in table (2) reject the null hypothesis that the coefficient on public capital (a_3) is equal to minus the coefficient on the spatial lag of public capital (a_6). Consider next the implication of the impact of a hypothetical simultaneous unit change in log public capital in all provinces on log productivity in all provinces, which is an *n x 1* column vector equal to

$$\frac{d \ln w_t^{M}}{d \ln K_t} = (I \ldots$$

In which I is an n x n identity matrix, $\tilde{\iota}$ is a n x 1 column vector with elements each equal to a_3 and $\tilde{\iota}$ is a similar vector with elements equal to a_6. This turns out to be a positive if $|a_6| < a_3$ and a vector of zeros when $a_3 = -a_6$, but if, as is the case with our model 6 estimates, $|a_6| > a_3$, then a simultaneous increase in $\ln K_t$ across all provinces causes $\ln w_t^M$ to fall across all provinces. However this is a very hypothetical situation that will not materialize in the real world, as different Provinces will exhibit different public capital stock growth rates, and as a result the effect of public capital stock on each region's productivity in practice will be different. In a Province where public investment is lower than the average of its neighbours' public investment, the net effect will be negative because the positive effect on productivity will be overwhelmed by the negative spillovers effect due to the increase of public capital stock in its neighbours. On the other hand, in a Province with a higher level of public investment than its neighbours, the net effect will be positive. In fact we show that even if this hypothetical situation did occur, and the unlikely event of a simultaneous unit increase in public capital across all Provinces did occur, in our final model (described below) there is no evidence that positive effects will be outweighed by the spillover of a negative effects from neighbours.

All our alternative specifications are all random effects models. One advantage of random effects estimation is that it takes account of permanent cross-sectional or between-variation, therefore picking up long-run effects. In contrast, within-unit fixed effect estimation focuses on short-run variation (Partridge, 2005, Baltagi, 2005, Elhorst, 2010). This raises the issue of whether random effects models can be considered to be consistent, satisfying an assumption of lack of correlation between the unobserved effects and the observed variables. The Hausman consistency test compares random and fixed effects estimates, using the test statistic $H = (\beta_r - \beta_f)'(\Sigma_r - \Sigma_f)^{-1}(\beta_r - \beta_f)$, in which the Σs are the respective covariance matrices from the fixed and random effects models. We find that there exists a positive-definite differenced covariance matrix. The $\hat{\beta}$ vectors comprise the parameter estimates that can be compared, and show that *H* is not an extreme value with reference to the relevant χ^2 distribution under the null, since the test statistic *H* is equal to 5.64, which has a p-value of 0.4641 in the χ_6^2 distribution. We thus conclude for model 6, that the random effects estimates are consistent. For the special case of panels with spatial dependence considered below there is a very limited literature on testing fixed versus random effects specifications. Mutl and Pfaffermayr (2008) consider this problem and suggest a solution, but as currently configured this is prohibitively complex in relation to our model set-up.

Our remaining random effects models are a variant on that of Kapoor *et al.* (2007) which introduces spatial dependence in the disturbances[13]. Province heterogeneity is absorbed as a component of the disturbance term, so that there are thus two error components, one component allowing for individual Province heterogeneity, the other component being a transient random effect specific to each province and time. The models also assume that an autoregressive error process describes cross-sectional dependence in the disturbances. Given the two error components, we assume that combined they form a spatial autoregressive process in which the disturbance in one province is simultaneously affected by, and affects, disturbances in 'nearby' provinces.

The spatially autocorrelated process for the disturbances are specified by ψ. Given that I_T is a *T* x *T* diagonal matrix (T is the number of time periods) with *1s* on the main

[13] It is also robust to distributional assumptions for the errors.

diagonal and zeros elsewhere, and I_n is a similar *n x n* diagonal matrix, then $I_{Tn} = I_T \otimes I_n$ is a *Tn* x *Tn* diagonal matrix with *1s* on the main diagonal and zeros elsewhere. These create the *nT* x 1 vector ψ [14]

$$\psi = (I_{Tn} - \gamma I_T \otimes W)^{-1} \xi \tag{11}$$

In which ξ is an *nT x 1* innovations vector. Time dependency is present in ξ due to the permanent error component α , hence

$$\begin{aligned} &\alpha \sim iid(0, \sigma_\alpha^2) \\ &\nu \sim iid(0, \sigma_\nu^2) \end{aligned} \tag{12}$$

$$\xi = (\iota_T \otimes I_N)\alpha + \nu \tag{13}$$

In which the province-specific component is the *n x 1* vector α , the *nT x 1* vector ν is the transient component, ι_T is a *T x 1* matrix with *1s* , and $\iota_T \otimes I_N$ is a *Tn* x *n* matrix of *T* stacked I_N matrices. The resulting *Tn* x *Tn* innovations variance-covariance matrix Ω_ξ is nonspherical. Also $\sigma_1^2 = \sigma_\nu^2 + T\sigma_\alpha^2$.

The main difference between the present model and that of Kapoor *et al.* (2007) is that we include an endogenous right hand side variable *lnP*, following Fingleton (2008). This model is described in detail in the cited references and we therefore do not include technical details of the two-stage least squares/GMM estimation method here. To summarize, the consistent residuals from the first stage of the three stage estimation procedure are two-stage least squared residuals. The consistent 2sls residuals are used in the second stage GMM estimation of the autoregressive error process parameter γ and the error component variances σ_ν^2 and σ_α^2. The estimate $\hat{\gamma}$ then allows the variables to

[14] In this the definition of nearby is given by the *W* matrix which is the row standardised contiguity matrix used previously for the autoregressive variables.

be purged of spatial dependence via a Cochrane-Orcutt transformation, with inference in the third stage based on a comparatively robust approach for IV estimation with nonspherical disturbances (Bowden and Turkington (1984), Greene (2003)). Tables 3 and 4 show the resulting estimates.

Table 3: Random effects models with spatially correlated error components (T=20, n=47)

REGRESSORS	Parameter Estimates Model 7#	Model 8#
constant	-0.119548 (-0.0854058)	0.940426 (2.38336)
Market Potential [lnP]	0.157773 (13.2558)	0.137727 (14.6282)
Public Capital [lnK]	0.183698 (2.76451)	0.0513671 (2.28948)
Private Capital [lnV]	-0.0845292 (-1.83307)	------
Human Capital [lnH]	0.0499766 (0.898652)	-----
Spatial Lag Public Capital [$WlnK$]	-0.12506 (-1.29839)	------
Spatial Lag Private Capital [$W\ lnV$]	0.120654 (1.88765)	------
Spatial Lag Human Capital [$WlnH$]	-0.117273 (-1.67007)	------
γ	0.310715	0.397892
σ_v^2	0.0020129	0.00199911
σ_α^2	0.0175	0.0205
Residual sum of squares	22.0546	24.395
R^{2*}	0.448175	0.396652
Estimation Method	2sls/GMM	2sls/GMM
Instruments	lnV, lnK, lnH, $WlnK$, $WlnH$, $W\ lnV$	lnK, *Time dummies*, lnV_c, lnH_c

	Time dummies
	lnV_c
	lnH_c

Notes: R^{2*} is square of correlation between actual and fitted values.

Table 4: Random effects models with spatially correlated error components (T=20, n=47)

REGRESSORS	Parameter Estimates Model 9#	Model 10#
constant	0.533372 (0.779373)	0.0146005 (0.0257145)
Market Potential [*lnP*]	0.158536 (12.1016)	0.158546 (12.0589)
Public Capital [*lnK*]	0.0851076 (3.26201)	0.0750996 (3.05246)
Spatial Lag Human Capital [*WlnH*]	-0.0634812 (-1.23478)	-------
Spatial Lag Public Capital [*WlnK*]	-0.00789638 (-0.121829)	-0.0774274 (-2.25444)
γ	0.326855	0.392173
σ_v^2	0.0024878	0.00201282
σ_α^2	0.0205	0.0206
Residual sum of squares	24.0769	24.5003
R^{2*}	0.397133	0.386455
Estimation Method	2sls/GMM	2sls/GMM
Instruments	*lnK* *WlnK* *WlnH* *Time dummies* *lnV_c* *lnH_c*	*lnK* *WlnK* *Time dummies* *lnV_c* *lnH_c*

Notes: R^{2*} is square of correlation between actual and fitted values.

In model 7, both market potential and public capital are significant, but private capital, human capital, and the spatial lags of all three capital variables are insignificant. We therefore reduce this model, retaining only the two significant variables (model 8) and then reintroduce each of the nullified variables back into model 8 to give five alternative models, each one being model 8 plus one additional variable. Only the spatial lags of public capital and human capital have any explanatory power in the presence of market potential and public capital. Model 9 in Table 4 shows the estimates of a model simultaneously retaining both variables, and the outcome is that neither is significant, probably because they are fairly strongly correlated (r = 0.6126). Model 10 shows the effect of dropping lagged human capital, with the lag of public capital now becoming significant. Likewise, dropping the lag of public capital causes the lag of human capital to become significant. It is therefore difficult to disentangle these two lag effects, although the lack of a significant direct human capital effect, and a clear rationale for the existence of a negative public capital spillover, suggests that we should focus on public capital. Model 10 shows that market potential[15] and public capital have significant positive effects, and although the negative spillover parameter estimate is larger than the positive public capital parameter, the difference is miniscule. The estimated elasticity for public capital stock is 0.075, broadly within the range of variation of the elasticities obtained in other papers. Goerlich and Mas (2001) reported an elasticity of 0.02 and Boscá *et al.* (1999) obtained an output elasticity of 0.026 for public infrastructures (0.035 in the long run). In Mas *et al.* (1994, 1996), the elasticity associated with productive public infrastructures is 0.23 and 0.08, respectively, so the second of their papers gives an outcome close to what we find. Also, the point estimate for the elasticity of substitution is 6.32 (= 1/0.15824) which is very close to the value of 6.25 assumed a priori in the

[15] The joint *F*-statistic on the two identifying instruments in the first stage, lnH_c and lnV_c, is equal to 37.33 with has a p-value equal to 2.6012e-016 in the $F_{2,916}$ distribution.

construction of market potential (the 95% confidence interval is 6.0282 to 6.6856). While the true elasticity of substitution may not be exactly 6.25 as assumed, using instruments allows for the potential measurement error introduced by our assumption.

Compared with our conclusion that there exists a negative effect, the sign on spatial infrastructure spillovers is less conclusive in the literature. Holtz-Eakin and Schwartz (1995), for the US, provide no evidence of spatial infrastructure spillovers. In contrast, Pereira and Roca-Sagalés (2003), Cohen and Morrison (2004) and Bronzini *et al.* (2009) find significant positive spatial spillovers for Spain, the US and Italy, respectively. Most interestingly, Boarnet (1998), using data for California's counties, found that the output of counties is negatively affected by neighbouring counties' infrastructure. Sloboda and Yao (2008) for the US, Delgado and Alvarez (2007) for the Spanish economy, and Pereira and Andraz (2006) for Portuguese regions, argue that public capital provided in a particular region raises the comparative advantage of that region compared with others, and could therefore attract production factors from other locations where output or productivity might consequently decrease. The development of new network infrastructure may alter the location decisions of firms, increasing investments and outputs in some provinces while causing disinvestments and possible job losses for others. Therefore these results suggest that a spatially neutral approach of uniformly increasing infrastructure on a national basis would be ineffective in Spain.

5. CONCLUSIONS

The relationship between public infrastructure and productivity has been analysed using a model based on New Economic Geography theory, with a version of the so-called wage equation estimated using a spatial panel, controlling for heterogeneity across provinces, spatial public investment spillovers between contiguous Spanish provinces and spatial dependence in the disturbances. The substantive conclusion we come to as a result of this analysis is that there is evidence in support of the wage equation involving market potential and public capital, with significant positive effects on the wage level, which we adopt as our proxy for the level of productivity.

The wage equation has provided us with a satisfactory theory on which to base our analysis of the impact of public capital, which we show to be relevant to our understanding of productivity variations, finding that a 1% increase in public capital leads to approximately 0.08% increase in productivity. The statistical significance of public capital is maintained under different model specifications, pointing to the robustness of our estimates. The estimated elasticity is broadly within the range of variation of the elasticities obtained in other papers that find a positive effect of public investment on productivity. In contrast the elasticity associated with human capital is much smaller, and in the final analysis is not significant.

Additionally, an interesting outcome of our analysis is that there is a significant spillover effect involving public capital in 'nearby' provinces, which turns out to be negative in sign. What we find is that if the level of public capital in a neighbouring province increases, then a province's wage or productivity level is reduced. We can interpret this as an effect of competition from neighbours. Public capital provided in one place is thought to enhance the comparative advantage of that place relative to others not receiving the capital investment. As argued by Boarnet (1998), negative output spillovers can result when mobile factors of production migrate to competing locations with the best infrastructure stocks. The negative spillovers of public investment might explain the results obtained in some previous papers which found a non-significant effect of public investment at a national level.

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