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May 2009

Online at <https://mpra.ub.uni-muenchen.de/15487/>

MPRA Paper No. 15487, posted 03 Jun 2009 00:14 UTC

# **A multilevel analysis on the economic impact of public infrastructure and corruption on Italian regions<sup>\*</sup>**

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This version May 2009

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**Keywords:** corruption; public expenditure; infrastructure, random coefficients; regional public accounts.

**JEL Classification:** H54; O18 ; R11; R58.

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<sup>\*</sup> I thank Isidoro Mazza for helpful comments. The usual disclaimer applies. Financial support of the Department of Development and Social Cohesion (DPS) of the Italian Ministry of Economic Development is acknowledged.

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## **1 Introduction**

This paper illustrates results concerning an empirical analysis dealing with the economic impact of public spending on infrastructure – as recorded in the novel Regional Public Accounts (RPA) database- on GDP augmented with corruption.

The study differs from existing literature for three main aspects.

First, it adopts a random coefficients model (RCM) approach in order to estimate the economic impact of public expenditure on infrastructure across Italian regions<sup>1</sup>.

The rationale for using a RCM is that a drawback common to studies on infrastructures and productivity is that they do not take into account parameters heterogeneity (Romp and de Haan, 2007).

At first glance, within a regional context, differences involving infrastructures' impact on economic performance could make little sense, however in the Italian regional setting characterised by *economic dualism* such kind of analysis might be of some interest in explaining differences in economic performance systematically reported across northern and southern regions.

Second, it uses an “objective” proxy for corruption proposed by Golden and Picci (2005) as explanatory variable. Indeed, many theoretical considerations lead scholars on this field to think that, in general, *social capital* matters (Putnam, 1993). Public infrastructure spending, in particular, “are the classic locus of illegal monetary activities between public officials [...] and businesses” (Golden and Picci, 2005). Moreover, there are reasons to assume that corruption and public investment vary together (Rose-Ackerman, 1999) or, more generally, that corruption modifies patterns of public spending (Mauro, 1998; Coppier, 2005; Shaw, Katsaiti et al., 2007).

Nevertheless, studies already realised use measure of corruption based either on surveys or, on number of crimes against the public administration (regarding the Italian case see Del Monte and Papagni 2001, 2007), although both measure presents some intrinsic weakness: the former being susceptible to be affected by sample selection or to become self-referential and, the latter, to reflect rather than the effective level of corruption the effectiveness of judiciary power in fighting corruption (Golden and Picci, 2005)

Hence, the interest to use an *objective* measure of corruption like the one introduced by Golden and Picci (2005), based on the difference between a measure of the physical quantities of public infrastructure and the cumulative price government pays for public capital stocks<sup>2</sup>.

The third novelty is represented by the use of RPAs. Indeed, under the project of RPAs data on public spending for each region divided by level of government and economic sectors are available giving the possibility for more accurate analysis with respect to previous works both in terms of reference universe and sector detail.

The rest of the paper is organised as follows. A descriptive analysis of data is provided in section 2; section 3 shows results of some preliminary tests regarding typical data-related problems: endogeneity, multicollinearity, and, stationarity. Once concluded this preliminary analysis, section 4 illustrates results achieved by RCMs regression; section 5 deals with the robustness of result regarding the proxy for corruption here utilised. Section 6 develops some concluding remarks.

## **2 A descriptive analysis of data.**

This section a briefly describes data utilised in the empirical analysis developed further.

The number of years covered spans over a 10 year period from 1996 to 2006. The dependent variable is regional GDP per capita taken from ISTAT (various years), while public expenditure's explanatory variables come from the RPAs.

During the sample considered the public sector as a whole spent on average 766.736 millions of Euros with an increase of 49,51% with respect to the initial year considered.

Regarding the expenditure composition by level of government, the central government spent a share around 50-60% with a clear tendency towards its reduction (62% in 1996 and 51% in 2006).

Despite the political discussion on “devolution”, regional administrations spent on average, 95.924 millions of Euros, with a share that, although increasing, is still around 10-13% of total expenditure.

Central public corporations, with an average expenditure of 129.781, account for a share of 15-20% (22% in 2006) confirming a strong role in delivering public goods and services and, in turn, public policy.

Going more into the detail should be noted that although RPAs classify public spending in 30 sectors, in order to deal with a lower number of explanatory variables, expenditure in capital account are grouped into four macro-sectors reflecting as many forms of government intervention, namely: economic infrastructure (EI), human capital (HC), social infrastructure (SI), and, residential building (RB). For the composition of each sector see Volpe (2007).

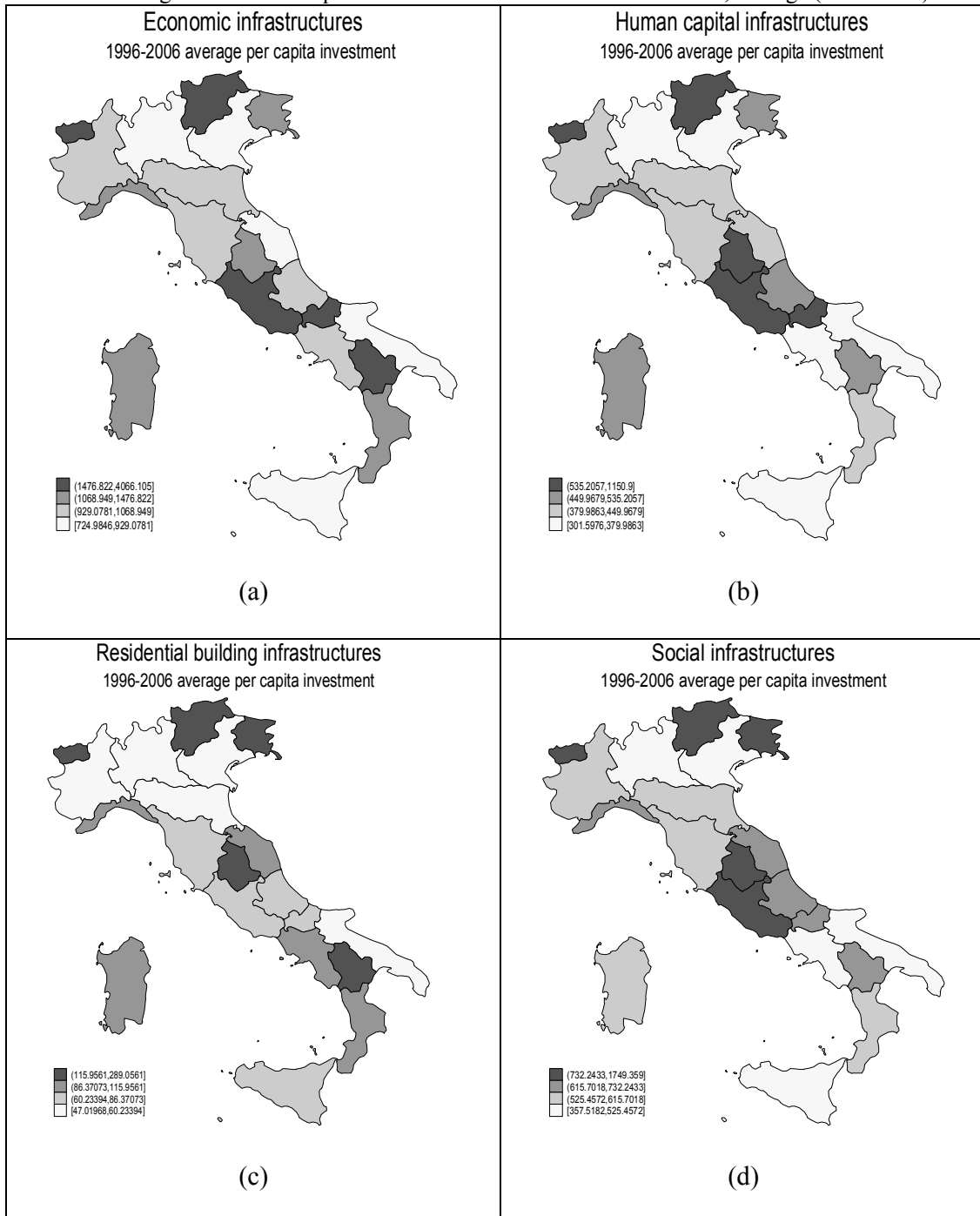
When not stated differently, in what follows variables are always expressed in per capita terms. Moreover, lower-case variables refers to variables expressed in natural logarithm.

Figure 1 (a)-(d) report the average value of per capita expenditure in each macro-sector during the period from 1996 to 2006. The upper part could be interpreted as representing public effort in providing territories with *production-oriented* infrastructure. Namely: economic infrastructures and human capital, while the lower part of the same figure (c)-(d) concerns *social policy* infrastructures: residential building and social infrastructure.

Certainly, a detailed analysis of pattern of public expenditure in infrastructure as emerge from graph reported below, goes further the purpose of present section. Reader

interested in this argument are addressed to a document of the DPS edited by Servizio Progetti Studi e Statistiche (2007).

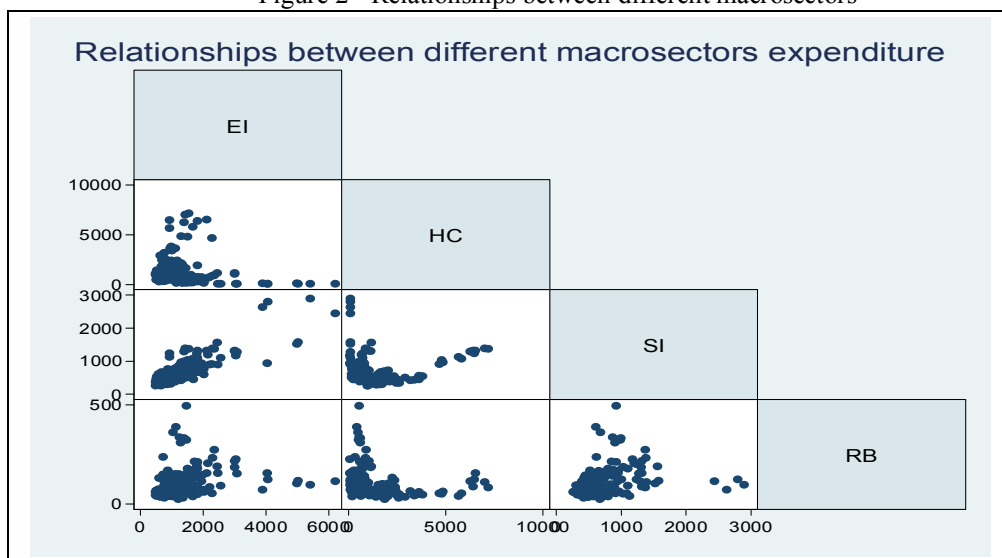
Figure 1- Public Expenditure on infrastructures. Macro-sectors, average (1996-2006).



Nevertheless, as general considerations, can be noted - from figure 1(a) - that Valle D'Aosta, Trentino, Sardinia, and, Basilicata spend more than the national average in the key sector to regional development represented by economic infrastructures. Considering the system as a whole, except – to some extent - Sicily and Puglia, emerges a generalised effort in order to endow with economic infrastructure southern regions. Expenditure on residential building follows more or less a similar pattern. With respect to human capital and social infrastructures it is worth to remark the result concerning Valle D'Aosta, Friuli, and Trentino which spend more than northern regions and generally more than average expenditure relative to the whole northern macro-area. However, as noted in the DPS' document cited above, data regarding social infrastructures are strongly influenced by expenditure in general administration, and once eliminated this sector differences become weaker. Among southern regions Basilicata and Sardinia shows a level expenditure generally higher than others in the same geographical area.

Figure 2 aims to give an intuitive image of patterns registered in public expenditure.

Figure 2 - Relationships between different macrosectors





Generally speaking data show a clear positive relationship between economic infrastructure and social ones. Economic infrastructure is also positively correlated with residential building expenditure, but is negatively correlated with expenditure in human capital which, in turn, is negatively correlated with residential building. Table 1 below reports the whole set of correlations concerning macro-sectors of public spending and their correlation with GDP showing that infrastructure spending is positively correlated with all four categories of infrastructures considered.

**Table 1. Public expenditure on infrastructure correlation matrix**

	<b>GDP</b>	<b>EI</b>	<b>HC</b>	<b>SI</b>	<b>RB</b>
<b>GDP</b>	1				
<b>EI</b>	0,39	1			
<b>HC</b>	0,27	-0,18	1		
<b>SI</b>	0,53	0,83	0,03	1	
<b>RB</b>	0,07	0,27	-0,26	0,30	1

This fact can be seen as a preliminary confirmation of the argument that a region well endowed with infrastructure has a relative advantage in terms of economic performance.

However, the direction of causality has to be tested. Next section deals with public expenditure endogeneity and other data related problems.

### **3 Preliminary tests**

This section deals with some typical data-related problem concerning public infrastructure and economic performance (for a review of data gaps and problems see Infrastructure Canada, 2007, p.44).

The first problem analysed concerns public expenditure (potential) endogeneity. Indeed, public infrastructures exogeneity with respect to (different measure of)

productivity is one of most debated point on empirical analysis concerning the economic impact of infrastructures.

The main argument used in order to question the hypothesis of infrastructure exogeneity is based on the idea that GDP might affect infrastructure in the sense that the higher the level of GDP the higher the demand for infrastructure.

In order to test for endogeneity both the Wu (1973)-Hausman (1978) and Durbin (1954)-Wu (1973)-Hausman (1978) were performed using as instruments the one-period-lagged variables (l1ei, l1hc, l1rb, l1si).

The rationale for using lagged values of public expenditure is supported by theoretical arguments based on incremental budgeting theory (Wildasky, 1975; Dempster and Wildasky, 1982).

From the strict econometric side, chosen instruments do not cast considerable doubts: the Anderson (1951)'s underidentification statistic shows that the model is identified, that is to say that, as expected, instruments are "relevant" in the sense that they are correlated with (assumed) endogenous regressors, and the Sargan (1958)-Hansen (1982)'s J statistic for overidentifying restrictions does not reject the null hypothesis that our instruments are uncorrelated with the error term (and that the excluded instruments are correctly excluded from the estimated equation).

**Table 2. Underidentification and Overidentification tests.**

<b>Underidentification test (Anderson canon. corr. LM statistic)</b>	19,875
Chi-sq(1) p-value	0,000
<b>Sargan statistic (overidentification test of all instruments)</b>	0.000

Results of endogeneity tests are reported in table 3 below.

**Table 3- Tests for endogeneity of hc ei si rb.**

<b>Wu-Hausman F test</b>	F(4,190)	1.395
	p-value	0,237
<b>Durbin -Wu-Hausman Chi-sq test</b>	Chi-sq(4)	5,708
	p-value	0,222

Both tests do not reject the null hypothesis that all categories of public infrastructure are exogenous at a high level of significance not rejecting, in turn, the principal hypothesis that infrastructures are (exogenous variables and) not *accommodating factors*.

An additional potential problem in this field is represented by multicollinearity. Indeed, it is likely that expenditures on public infrastructures are highly correlated: “wealthier” regions tend to spend more on everything, and “poorer” regions spend less on everything. Not surprisingly, it can be difficult to estimate the effect of any particular expenditure category on GDP.

However, the variables under consideration fulfill the threshold of Variance Inflation Factor (VIF) and *tolerance* (1/VIF) in order to individuate excessive or serious multicollinearity (O’Brien, 2007) also when the proxy for corruption is included. Therefore, under this aspect, the four sectors can be used as explanatory variables being confident that results of analyses are quite solid on statistical grounds.

**Table 4. VIF and tolerance of four macro-sectors of public infrastructure**

Variable	VIF	1/VIF
rb	1.27	0,79
corr	1,40	0,71
ei	3,69	0,27
hc	4,53	0,22
si	7,83	0,13

Although the sample considered in this work is “too short” to properly consider the problem of data stationarity, an analysis of stationarity of data concerning *gdp ei si hc* is performed. The test utilised is the Im, Pesaran et al. (2003) (IPS) test for unit roots in panel data because of many desirable properties (Konya, 2001) with respect to other available tests (Levin and Lin, 1992; Maddala and Wu, 1999; Quah 1994). Table 5 shows results concerning the tests under discussion.

**Table 5- Unit roots IPS test**

Variable	Lags	Im, Pesaran e Shin (t-bar)	P-value
gdp	1	-1.896	0.040
ei	1	-1.915	0.033
hc	1	-2.078	0.005
si	1	-1.941	0.026

From table 5 it emerges that for all variables considered the null hypothesis of non-stationarity is rejected at significance level of 5%. Once more, should be noted that this result could be heavily influenced by the length of the sample considered.

Nonetheless, since all tests performed give the expected result the analysis can be performed being confident that its results is not affected by endogeneity, multicollinearity, and, data non-stationarity problems.

#### 4 Estimation results

This section shows result concerning the estimation of the economic impact of infrastructure on gdp across Italian regions augmented with corruption.

Moving from Baltagi, Song et al. (2001) - that estimate a Cobb–Douglas production function examining the productivity of public capital in each state's private output - the estimation utilises a RCM, with the precise purpose to capture heterogeneity across Italian regions.

The interest for this issue comes from the consideration that parameter heterogeneity is not yet well explored in literature leading Romp and de Haan(2007) to assert that

although economists often test the residuals of their regressions for heteroskedasticity and structural change [...]. Unfortunately, none of the studies reviewed [by them] take the issues of model uncertainty and outliers and parameters heterogeneity seriously into account, which casts considerable doubt on their findings ( Romp and de Haan, 2007).

Furthermore, the Italian regional case is particularly interesting due to the persistent economic dualism still present despite the massive amount of capital spent in order to fill the gap. At this regard it is worth noting that a policy objective, concerted with the EU, consists in assigning a share of 45% of total expenditure in capital account to southern regions.

The analysis is augmented with a proxy for corruption proposed by Golden and Picci (2005), based on the difference between a measure of the physical quantities of public infrastructure and the cumulative price government pays for public capital stocks, to empirical test the theoretical argument that corruption is a relevant structural factor in explaining different economic performance across Italian region. For the best of my knowledge this is the first work using this measure since its proposition in 2005.

The general equation is given by<sup>3</sup>

$$\begin{aligned} gdp_i = & \beta_{1_i} + \beta_{2_i}ei + \beta_{3_i}si + \beta_{4_i}hc + \beta_{5_i}corr + \\ & + \beta_6popden + \beta_7gdp96 + \beta_8D2001 + \beta_9DLA + \varepsilon \end{aligned} \quad (1)$$

where *ei*, *si*, and *hc* represent public expenditure in infrastructure as already defined. As already said, the variable *corr* represent the proxy for corruption proposed by Golden and Picci (2005) rearranged in order to render more intuitive the interpretation of its coefficient. The question might be explained as follows. Since, the measure of corruption is inversely scaled – in the sense that it has lower numbers for higher levels of corruption- if empirical analysis should confirm (as it does) that corruption has a negative effect on *gdp*, this fact will result in a positive coefficient for it. An easy way to “correct” this counterintuitive fact consists in taking the inverse of the original measure, so that a negative effect on *gdp* will result in a *negative* coefficient for corruption in the regression analysis. A second, and last, computational transformation adopted consist in multiplying it by 100 in order to avoid negative values of corruption in taking logarithm. Following three variable might be thought as a set of control variables. The variable labelled as *popden* represents the population density trying to capture general development characteristics of each region; *gdp96* represents the level of *gdp* at the beginning of the period considered; *D2001* represents a year’s *dummy* relative to the constitutional reform in a federalist sense realised in 2001<sup>4</sup>; *DLA* is region dummy relative to region Lazio to take into account that all expenses not regionalised are assigned to this last region.

Moving from the basic equation introduced above four alternative regressions have been considered, each one obtained letting the intercept and respectively the coefficient for *ei*, *si*, *hc*, and *corr* vary according to a RCM setting<sup>5</sup>. In other words, for

each alternative regression the intercept is estimated as a fixed part common to all regions and a random part region-specific, thus in symbols we have

$$\beta_{1i} = \beta_1 + \zeta_{1i}$$

where  $\zeta_{0i}$  is a random intercept component and the coefficient relative to variable from time to time considered is determined as

$$\beta_{ji} = \beta_j + \zeta_{ji}$$

where  $\zeta_{ij}$  is a random slope component.

Models of the type just specified have several potential desirable features when used for analysis concerning heterogeneous cross-section units. In particular, they might represent a “good” solution for data clustering given that not only the intercept might vary, but the slopes may be an issue also. Therefore, if slopes and intercepts vary, a random coefficients model may fit better.

Furthermore, we could be interested in analysing variations across observational units *per se*, focusing the analysis on the different pattern of impact of the independent variable(s) rather than to (the prediction of) the dependent variable.

Certainly, the difference between analysis focusing on the prediction of the dependent variable and those focusing on the impact of explanatory is subtle. Nevertheless, the second research question might be of some interest in context like the Italian regional case characterised by a strong north-south dualism despite different *ad hoc* political measure – like the famous “Cassa per il Mezzogiorno” instituted in 1950 or the recent *rule of 45%* cited above - designed to fill the gap<sup>6</sup>.

Table 6 below reports estimation results concerning both the fixed part of coefficients  $-\beta_j$  in terms of the theoretical framework introduced above- and the standard deviation of the random parameters for each variable estimated using the RCM estimation method.

**Table 6. Public Infrastructure impact on gdp. P-value in parenthesis. Standard errors in brackets.**

Gdp		Regression			
	(1)	(2)	(3)	(4)	
corr	<b>-0.424263</b> (0.000) <b>[0,050]</b>	-0.3857345 (0.000) [0,068]	-0.4030503 (0.000) [0,062]	-0.3672979 (0.000) [0,062]	
ei	0.1514033 (0.000) [0,024]	<b>0.1501155</b> (0.000) <b>[0,027]</b>	0.1630842 (0.000) [0,025]	0.1587971 (0.000) [0,025]	
hc	0.1427796 (0.000) [0,044]	0.1370628 (0.000) [0,047]	<b>0.1323167</b> (0.000) <b>[0,052]</b>	0.1608181 (0.000) [0,051]	
si	0.3164714 (0.000) [0,034]	0.3402164 (0.000) [0,035]	0.3382694 (0.000) [0,034]	<b>0.3097269</b> (0.000) <b>[0,045]</b>	
Standard deviation					
random slope	0,162	0,393	0,109	0,085	
random intercept	0,802	0,236	0,656	0,548	

Note: Mixed effects – REML regression with 220 observations and 20 groups (regions). Values reported in bold refer to variable for which from time to time was estimated also the random coefficient.

Estimates synthetically reported in table 6 shows that all variables are statistically significant and show the expected sign.

Indeed, all categories of public infrastructure have a positive impact on *gdp*. Economic infrastructures (ei), in particular, have coefficients essentially comparable to that (0.13) obtained by Marrocu, Paci et al. (2006) using the same dataset of RPA but a different methodology and also with different data and sample.



Focusing on the elasticity of the proxy for corruption emerges that it is higher than relative to public infrastructures and, in particular, of magnitude comparable - in absolute value - to that relative to social infrastructure. At this regard it is worth to note that also in Del Monte and Papagni (2001), although with a different estimation method, the coefficient for corruption – in their work measured as the number of crime against the public administration- is comparable in absolute value to their estimate relative to the share of the real public investment (considered as a whole) in the real GDP. The two coefficients being respectively -.138 and .156. Therefore, results obtained in this work confirms that corruption plays a significant role in affecting the economic performance of Italian regions.

In order to test the effective advantage of using a RCM approach compared to OLS a likelihood-ratio (LR) test comparing the fitted mixed model to standard regression with no group-level random effects was performed for each regression. The null hypothesis of the test is that the random part of the parameter -  $\zeta_{ji}$  in the notation adopted here - is zero. Thus, rejection of the null hypothesis can be interpreted in the sense that a RCM is preferable.

Preliminarily should be noted that due to technical complications<sup>7</sup>, while we can compute the exact distribution of the test in the one random parameter case (Self and Liang, 1987), an appropriate and sufficiently general computation methods for the more-than-one-parameter case have yet to be developed. Nonetheless, theory (e.g. Stram and Lee, 1994) and empirical studies (e.g. McLachlan and Basford, 1988) have demonstrated that, whatever the distribution of the LR test statistic, its tail probabilities are bounded above by those of the chi-squared distribution with degrees of freedom equal to the full number of restricted parameters. Therefore, using as reference

distribution the chi-squared with full degrees of freedom a conservative test is produced. The reported significance level for the LR test is an upper bound on the actual significance level. As such, rejection of the null hypothesis based on the reported level would imply rejection based on the actual level.

Once aware of these technical problems we can pass to results. Table 7 below shows result for each regression. The first column should be read in the sense that result refers to a regression in which the random intercept *and* the coefficient relative to the variable reported in it have been treated as “random”.

**Table 7. LR test vs. linear regression**

Variable	Lr-test Chi-sq (2)	p-value
corr	96.39	0.000
ei	90.59	0.000
hc	93.30	0.000
si	92.04	0.000

All tests performed shows that the null hypothesis of inexistence of difference in parameters across region has to be rejected, reinforcing theoretical consideration leading to the use of a RCM.

Hence, it is worth focusing the attention to the random part of coefficients<sup>8</sup> -  $\zeta_{ji}$  in terms of the theoretical framework introduced above - and considering the correlation between coefficients predicted in the four regression considered we obtain the following correlation matrix in which  $r_{.1}$  represents the random slope for each variable, while  $r_{.0}$  represents the random intercept predicted in the regression in which the relative variable was allowed to have a random slope. To be clearer,  $r_{corr1}$  – for

example - represents the random slope(s) for the proxy for corruption (*corr*) predicted after regression (1), while  $r\_corr0$  represents the random intercept(s) predicted after the same regression (1).

**Table 8. Random coefficient (BLUPs) correlation matrix.**

	r_ei1	r_ei0	r_corr1	r_corr0	r_hc1	r_hc0	r_si1	r_si0
r_ei1	1							
r_ei0	-0,9347*	1						
r_corr1	-0,2505*	0,1421*	1					
r_corr0	0,2505*	-0,1421*	-1*	1				
r_hc1	0,5751*	-0,7101*	-0,0820	0,0820*	1			
r_hc0	-0,4634*	0,6380*	0,0393	-0,0393*	-0,9892*	1		
r_si1	0,4331*	-0,5765*	-0,1556*	-0,1556*	0,8681*	-0,8704*	1	
r_si0	-0,2825*	0,4732*	0,0844	-0,0844	-0,8313*	0,8604*	-0,9827*	1

\*5% level of significance

Table 8 gives rise to some interesting observations. At first glance, it shows that random slopes and random intercept are negatively correlated for all variables considered. Therefore, the higher the intercept the lower the slope and *vice versa*. This fact could be interpreted – to some extent - as evidence of “convergence” in the impact of public expenditure, in the sense that, *ceteris paribus*, higher intercept are associated with flatter curves.

Furthermore, focusing the attention to the random slopes can be observed that random slopes regarding infrastructures are significantly positively correlated each other. The interpretation in terms of policy that might be done to this last result is in favour to a policy of *balanced growth* – as opposed to *unbalanced growth* - whereby many interdependent public investment projects are started simultaneously based on the principal justification of external economies (Hansen, 1965).

Enlarging the analysis to corruption, the same table shows that random slopes relative to all categories of infrastructure considered are negatively correlate to random coefficient for corruption meaning that the lower<sup>9</sup> the “correction” for corruption the higher the “correction” relative to infrastructure (let say economic infrastructure). This fact can be generally interpreted in terms of the theoretical argument that the negative effect of corruption could prevail when, due to infrastructures scarcity, the productivity of public spending is high, so that corruption, stealing public resource to high-productivity-infrastructures, exerts its negative effect in a more incisive way (Del Monte and Papagni, 2001).

Obviously analysis of correlation between complete coefficients ( $\beta_{ji} = \beta_j + \zeta_{ji}$ ) will result in identical pattern and is not reported.

Focusing the attention to the geographical distribution of random effects a test for spatial autocorrelation of random coefficient and random slopes has been performed. Table 9 below shows that the hypothesis of spatial independence of random effects of corruption should be rejected.

**Table 9. Moran’s I Spatial correlogram for Random slope of corruption (BLUP).**

Distance bands	I	E(I)	sd(I)	z	p-value*
(1-2]	-0,099	-0,005	0,018	-5,167	0,000
(1-3]	-0,113	-0,005	0,012	-9,205	0,000
(1-4]	-0,201	-0,005	0,008	-24,043	0,000
(1-5]	-0,049	-0,005	0,007	-6,669	0,000

\*2-tail test

Furthermore, the hypothesis of spatial independence has to rejected for all random coefficients. Probably, the simplest interpretation of this last result is in term of spill over effects between neighbour regions.

A test for spatial autocorrelation of residuals relative to each regression was also run in order to individuate an eventual systematic bias in capturing the effect of variables considered based on geographical ground.

The most commonly used specification test for spatial autocorrelation is derived from a statistic, already used, developed by Moran (1948, 1950a, 1950b). Following Anselin (1999) the Moran's I statistic, in matrix notation, can be expressed as follows

$$I = \frac{N}{S_0} \frac{\varepsilon' W \varepsilon}{\varepsilon' \varepsilon}$$

where  $N$  is the number of geographical units considered,  $S_0 = \sum_i \sum_j w_{ij}$  is a standardization factor that corresponds to the sum of the weights for the non-zero cross-products,  $\varepsilon$  indexed the vector of residuals, and  $W$  is a spatial weights matrix.

Moran's  $I$  tests have been computed both in the cumulative and in the consecutive case for five different distance bands for each regression and none of tests performed reject the null hypothesis of spatial independence of residuals. This fact can be interpreted as evidence that our RCMs do "a good job" in capturing systematic differences (especially) between northern and southern regions.

## **5 The robustness of result concerning corruption.**

In order to test the validity of result reported above, in this section, I compare results achieved with the measure of corruption utilised in regressions presented above with those obtained using, everything else be equal, the Putnam's index<sup>10</sup> (putnamindex) and geographical latitude (LAT)<sup>11</sup> as additional controls.

The rationale for using these two control is twofold.

First, with regard to corruption appendix A (available upon request) shows that the measure of corruption here utilised is significantly spatial correlated. Therefore, one could doubt that - once considered as control - it rather than capturing the effect of corruption acts a sort of “geographical control”. If so, introducing an explicit measure of geographical latitude should result in a statistical insignificance of the measure of corruption because of a better explicative power of a precise geographical control with respect to a variable that acts as its proxy.

Second, the use of Putnam’s index makes sense because it represents a direct – as opposed to the indirect measure given by corruption - proxy of *social capital* for which regional-level data are available and about which a relatively more secure knowledge can be assumed. Even if the measure are strictly correlated (Golden and Picci, 2005) introducing this direct measure in our regression could be interpreted as a test for the explicative power of the indirect measure of social capital represented by corruption.

Nevertheless, neither the measure for geographical latitude nor the Putnam’s index are statistical significant in the regressions where also the measure of corruption is introduced both in the case in which they are considered together (i.e. corr-putnamindex-LAT) and in the case in which they are considered one at time (i.e. corr-putnamindex and corr-LAT)<sup>12</sup>.

Nonetheless, at margin should be noted that the Putnam index became statistically significant when considered separately and, in particular, it has the expected (positive) sign when considered separately.

The fact that LAT is statistically insignificant implies that a difference in the economic impact of infrastructure based merely on geographical ground should be rejected.

Combining these empirical evidences can be drawn the conclusion that although *social capital* - systematically higher in northern regions than in the southern ones - matters, once considered the proxy for corruption, results became more robust in the sense that corruption is statistically significant and with the expected sign in a set of specifications broader than the one characterising the two alternative variable here considered.

## **6 Concluding Remarks**

In this paper I empirically analysed the economic impact of economic of public spending on infrastructure combined with the analysis concerning the role of corruption in this field by means of a set of random coefficient models.

Random coefficients have been estimated for three different categories of public infrastructures: economic infrastructures, human capital infrastructures, and, social infrastructures. Moreover, a different effect of corruption across regions was estimated. Tests based on likelihood-ratio show that the null hypothesis of inexistence of difference in parameters across regions has to be rejected for all variables cited above.

A Moran's test for spatial autocorrelation of random effects predicted for each variable considered was performed showing that, in general, the impact of public spending on infrastructures is spatial correlated.

This evidence combined with the argument that "casual observation [supported with appropriate tests] suggests that, in Italy, infrastructure construction costs should be

relatively uniform in different parts of the country” (Golden and Picci, 2005) could be interpreted as a confirmation of the existence of differences based not only on different amounts of infrastructure spending but also based on their effects.

Focusing the attention to the effects of corruption it emerges that the *objective* proxy for it utilised in this work negatively affects the economic performance of Italian regions. Moreover its negative effect, like the effect of infrastructure spending, is spatially correlated.

Both phenomena could be interpreted as consistent with the theoretical arguments developed in this field (Del Monte and Papagni, 2001). Indeed, the negative effect of corruption is higher when productivity of public spending is high. In other words, corruption exerts its negative effect in a way more incisive when stealing public resource to high-productivity-infrastructures.

In order to test the robustness of result reported above I compared results achieved with the measure of corruption utilised with those obtained using, everything else be equal, the Putnam’s index and geographical latitude as additional controls representing a direct measure of social capital and a direct geographical control.

The robustness analysis shows that differences based merely on geographical bases should be rejected. Regarding the *social capital*, even though it matters, once considered the proxy for corruption results became more robust.



## Notes

<sup>1</sup> All multilevel analyses described in this paper were implemented in Stata© 10.0 using the "xtmixed" command, "cmdlog" file available upon request. Two details about model specification are worth noting here. First, no assumptions were made about the structure of the covariance matrix. Rather, all variances and covariances were distinctly estimated. Second, all models were estimated using restricted maximum likelihood (REML) over maximum likelihood (ML) since the latter is more sensitive to loss of degrees of freedom when dealing with a small number of groups (Snijders and Boskers, 1999).

<sup>2</sup> The issue of the interpretation of this measure as *corruption* rather than as a measure of *efficiency* is developed in length by the Authors, therefore readers interested in this peculiar aspect are addressed to Golden and Picci (2005).

<sup>3</sup> The variable related to residential building expenditure (RB) is not statistically significant according to many different specifications.

<sup>4</sup> Cost. L. 2001, n. 3 "Modifiche al titolo V della parte seconda della Costituzione". Gazzetta Ufficiale n. 248, 24th October 2001.

<sup>5</sup> Due to computational difficulties estimating all coefficients as random is not possible.

<sup>6</sup> For a review of Italian regional development measures see ch.15 of Cellini, R., Ed. (2004). *Politica economica. Introduzione ai modelli fondamentali* Milano, McGraw-Hill.

<sup>7</sup> The question can be summarised as follows. As said the LR test assesses whether all random-effects parameters of the mixed model are simultaneously zero. In the one random-effects parameter case, this parameter, a *variance component* in Stata™ terms, is restricted to be greater than zero. Since the null hypothesis is that this parameter is indeed zero, which is on the boundary of the parameter space, the distribution of the LR test statistic is a 50:50 mixture of a chi2(0) (point mass at zero) and a chi2(1) (point mass at one) distribution. Therefore, significance levels in the one-parameter case can be adjusted accordingly. However, when we have more than one random-effects parameter to be tested, the distribution to be considered becomes unclear. In a model where we have two random coefficients with unstructured covariance matrix of random parameter  $\zeta_{ji}$  (both relative to intercept and slope)

$$S = \begin{bmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{21} & \sigma_{22} \end{bmatrix}$$

for example, since the "random" component of the mixed model comprises three parameters ( $\sigma_{11}, \sigma_{21}, \sigma_{22}$ ), it would appear that the LR comparison test would be distributed as chi2(3). However, there are two complications that need to be considered. First, the variances  $\sigma_{11}$  and  $\sigma_{22}$  are restricted to be positive, and testing them against zero presents the same boundary condition described above. Second, constraints such as  $\sigma_{11} = 0$  implicitly restrict the covariance  $\sigma_{22}$  to be zero as well, and from a technical standpoint it is unclear how many parameters need to be restricted to reduce the model to one with no group-level random effects.

<sup>8</sup> Predicted by Best Linear Unbiased Linear Predictors (BLUPs).

<sup>9</sup> Note that corruption has negative coefficients and note also that random slope and random intercept relative to corruption are perfectly (negatively) correlated.

<sup>10</sup> Data on Putnam index courtesy of Robert Putnam. Note also that even if there is an obvious temporal gap between Putnam's measure, constructed using data from the period 1978–1985, and data here considered a comparison between the original measure and a more recent (1990–1994) study due to Simoni(1997), updating Putnam's index, shows that it is relatively unchanged since the period of original data collection and can be considered as characterized by a strong persistence.

<sup>11</sup> Both for Moran'I linear geographic coordinates and for the geographical latitude as control I utilised data relative to Italian waypoint available at <http://xoomer.alice.it/ntpai/GPS/ISTAT/links.html>

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