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Spatial dynamics of public pharmaceutical expenditure

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A regression model for per capita public pharmaceutical expenditure is analyzed. The necessity of simultaneously controlling for dynamic patterns and spatial spillover in such analyses is demonstrated. In contrast to previous studies of impact of small-area variation, the present study exploits important aspects related to spatial dynamics as the effects of spatial spillover are analyzed and interpreted within a framework of spatial dynamics and spatial error-correction. It is shown that such dynamics bear important implications related to spatial convergence of a pharmaceutical market. The paper is accessible to an audience experienced with linear regression; basic exposure to spatial statistics is helpful but not strictly necessary.

Introduction

During the last decade, public pharmaceutical expenditures in Spain have grown at a rate superior to the total public health care expenditure (DARBÁ, 2003a, DARBÁ, 2003b). Thus, public pharmaceutical expenditures make up an increasing proportion of total public health care expenditures. Indeed, pharmaceutical expenditures made up 16.8% in 1991 and had in 2002 increased to make up 23% of total health care expenditures (LOPEZ-CASASNOVAS *et al.*, 2005). This growth is found not only in Spain, but is a general feature of European Union countries (ESS *et al.*, 2003); however, Spanish pharmaceutical expenditures as a share of public health care expenditures exceed EU averages (LOPEZ-CASASNOVAS, 2005). It is thus crucial to analyze the causes of this growth differential in order to focus on a rational use of medicine.

The regulation of the pharmaceutical market in Spain is shared between national regulatory bodies and regional authorities. There are notable differences in health resources supply and health care expenditure across regions (LOPEZ-CASASNOVAS et al., 2005) and there is evidence of regional variation in prescription rates and expenditure per prescription resulting in regional heterogeneity in pharmaceutical expenditures, and in pharmaceutical expenditures as a share of total regional health care expenditures (COSTA-FONT and PUIG-JUNOY, 2004).

Studies on pharmaceutical expenditures from a regional perspective are very scant, although it is possible to find a few articles dealing with the analysis of regional health care expenditures (see e.g. KITCHENER et al., 2003,

LEVAGGI and ZANOLA, 2003, LOPEZ-CASASNOVAS and SAEZ, 2001, MOSCONE and KNAPP, 2005). Despite the ample body of evidence of geographical differences in the use of health-care procedures in the literature on small-area variation in health care (FOLLAND et al., 2003, HAM, 1988, JOINES et al., 2003, WENNBERG and GITTELSOHN, 1973, WESTERT et al., 2004), few studies have examined the geographical variability in the use of pharmaceuticals (see e.g. DUBOIS et al., 2002, METGE et al., 1999, MORGAN, 2005). The causes of variation discussed in the literature are the prevalence of diseases, mixed opinions about the effectiveness of surgery, practice style, health supply resources and differing patient preferences.

Only a few studies of small-area variations have considered spatial variation in medical practice. WESTERT et al. (2004) studied spatial disparities in hospital discharges (as measured by coefficients of variations) and found these disparities to be approximately unchanged during the 1980s and 1990s. JOINES et al. (2003) found that hospitalization rates for low back problems varied significantly across the counties of North Carolina. They further found that counties with similar rates clustered geographically and concluded that spatial effects are important and should be considered in small area studies. MOSCONE and KNAPP (2005) explored the spatial patterns of mental health expenditure and established – as in Joines et al. – the importance of controlling for spatial spillover. Moscone and Knapp's study found a positive significant spatial effect suggesting that adjacent local authorities mimic the behavior of their neighbours and tend to have similar mental health expenditure. In contrast to the present study, however, none of the two latter studies incorporated dynamic properties as a part of their analysis.

The present study will focus on the regional variations in the models of public pharmaceutical expenditures in Spain, and will thus contribute to the literature on small-area variation and determinants of health care expenditure. The aim of our study is to analyze the determinants of provincial-level pharmaceutical expenditures in Spain while controlling for spatial effects. However, in contrast to the above mentioned small-area studies, we go deeper into the nature of spatial spillover and show that this spillover contains valuable information regarding the conditional spatial convergence of expenditure models.

In the case of Spain, large provincial variations in models of public pharmaceutical expenditures are to be expected. It is thus important to analyze not only the effects of change over time in the determinants on the expenditures, but also the effects of inter-provincial differences in determinants on inter-provincial differences in expenditures. Both targets are achieved by using pooled cross-sectional data in the analyses to follow.

In order to obtain efficient results, a Seemingly Unrelated Regression (SUR) framework is advocated. The SUR is specifically designed to capture inter-temporal residual correlation and time-varying residual variances, and is favored over the simpler but less efficient OLS framework. We show that the SUR framework leads to a global (i.e., non-spatial) specification. Potential spatial spillover effects are modeled to provide a local (i.e., adjusted for spatial dynamics) specification. Deeper analyses of the spatial dynamics are obtained by applying a spatial speed of adjustment re-specification and spatial error-correction specifications. It is concluded that spatial convergence of expenditure behavior is present, but that the speed of this convergence is relatively slow, due to a large endogenous spillover expressing rigidities which are caused by supra-provincial forces such as regulation and/or market imperfection.

The Spanish pharmaceutical market

In Spain, the prices of publicly financed pharmaceuticals are fully or partially controlled, and the price index of medicine has practically not risen in the last decade. Nevertheless, this does not preclude the fact that new products entering the market are introduced at a sales price higher than that of the already existing ones. Several studies have shown that the replacement of older drugs by newer, more expensive drugs is the single most important reason for the increase in pharmaceutical expenditures (see e.g. DUBOIS *et al.*, 2000, GERDTHAM and LUNDIN, 2004, MORGAN, 2005), whereas the real price index of existing drugs is decreasing. The second most important reason is that a larger quantity is consumed because of increases in the intensity of medication in terms of defined daily doses per patient. Similar results are found in analyses of the increase in pharmaceutical expenditures in Spain (DARBÁ, 2003b, ROVIRA *et al.*, 2001).

The Spanish national health system is a decentralized system in which the regulation of the pharmaceutical market is shared between national regulatory bodies and the regional authorities – called Autonomous Communities (AC) - however, most of the key regulatory bodies are run centrally at the national level

to reduce diversity and maintain overall control (COSTA-FONT and PUIG-JUNOY, 2004); see Figure 1 for a map of provinces by AC.

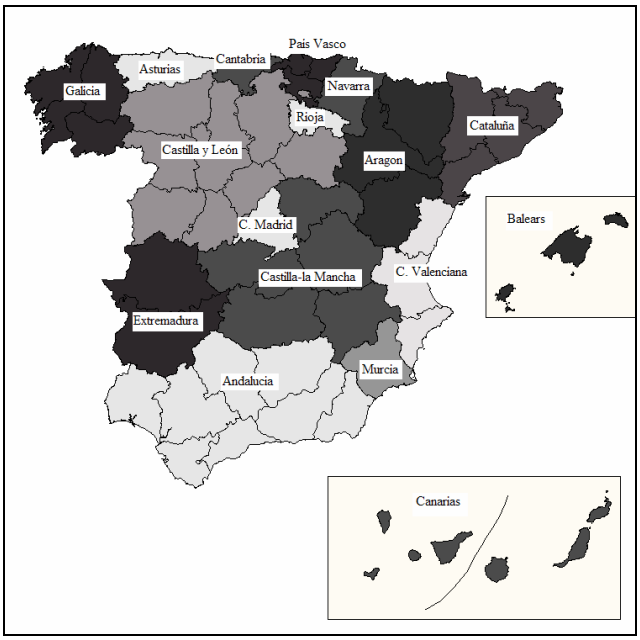

Figure 1. Provinces by Autonomous Communities

Even though cost containment has been a major priority for publicly financed pharmaceuticals this has not resulted in significant savings in public expenditures (COSTA-FONT and PUIG-JUNOY, 2004, DARBÁ, 2003a, DARBÁ, 2003b). The average price for pharmaceuticals is below EU averages with older drugs priced significantly below the EU average (PUIG-JUNOY, 2004). The market for generic drugs was small compared to the EU average, accounting for 3% of total drug sales in 2000 and had increased to 6.4% of total drug sales by 2003 (COSTA-FONT and PUIG-JUNOY, 2004). There seems to be significant regional heterogeneity in the use of generics (COSTA-FONT and PUIG-JUNOY, 2004). New drugs are not priced significantly below the EU average and these drugs account for the largest market share (COSTA-FONT and PUIG-JUNOY, 2004, DARBÁ, 2003a, DARBÁ, 2003b).

Different cost containment policies such as negative lists of excluded drugs, the regulation of profits, repayments from pharmaceutical companies, the reference pricing system and the promotion of the use of generic drugs have had little effect on the overall increase in pharmaceutical expenditures. Some of these policies are under the devolved responsibility of the 17 regional health systems. The ACs have gradually become significant actors in the pharmaceutical policy along with the decentralization process starting in the early 1980s until the completion of the devolution process in 2002.

Funding is mainly centrally collected and distributed to the ACs. Until 2001, regional health care financing was decided in a separate negotiation between the Minister of Health and the corresponding Regional Ministers in the 17 ACs, mainly allocating funds as block grants following the lines of an unadjusted capitation formula (LOPEZ-CASASNOVAS et al., 2005). Since 2002 the health care expenditures have been allocated as part of general financing using a capitation formula with some demographic adjustments. Health care expenditures account for around 40% of the ACs' total funding. The ACs have some possibilities of raising funding by levying higher taxes; however, various central funds strive to maintain territorial equity.

There are some inter-regional inequalities in health expenditure per capita but the coefficient of variation in regional health care expenditure per capita is one of the lowest among health care systems for which territorial health care expenditures may be identified (see LOPEZ-CASASNOVAS and SAEZ, 2001). There seems to be significant differences in hospital specialization, physician density and technology and it has been suggested that this diversity can be partly explained by differences in particular GDP and population structures (LOPEZ-CASASNOVAS et al., 2005). The regional inequality in health expenditure is however not correlated with inequality in health outcomes (LOPEZ-CASASNOVAS et al., 2005).

Modeling spatial dynamics

The point of departure is a linear regression model. Assuming initially one cross-section of N observations, this reads as

$$y_t = \alpha_0 i_N + X_t \beta_0 + \upsilon_t, \upsilon_t \sim N(0, \sigma_{tt}^2 I) \qquad (1)$$

where X_t is an N by K dimensional matrix of explanatory variables, y_t an N dimensional vector of endogenous observations, β_0 a K dimensional coefficient vector, α_0 a constant term, i_N a column vector of N ones, and υ_t a residual with variance σ_{tt}^2. Applying pooled data for T periods leads to T equations, one for each time period. The residuals of the T equations are correlated, and the variances for the cross-sections vary over time. Between any two time periods, the residual covariance reads as

$$E(\upsilon_t \upsilon_s^{'}) = \sigma_{ts}^2 I, \quad t, s = 1,..,T. \qquad (2)$$

The model defined by (1)-(2) is estimated efficiently by applying Feasible Generalised Least Squares (F-GLS) estimation to obtain the ZELLNER (1962) Seemingly Unrelated Regression (SUR) estimate for β_0, denoted β_0^{SUR}. Further, β_0 is allowed to vary with time by extending X_t of formula (1) with interactions between X_t and a time trend.

In terms of spatial econometrics, (1) expresses the relationship as a global equilibrium between non-spatial variables, i.e. β measures the effects of the exogenous variables (*X*) on public pharmaceutical expenditures (*y*) assuming a global equilibrium to be present, so that variations in expenditure across regions are only caused by variations in exogenous determinants. For the case of conditional convergence, regional adjustment towards the global equilibrium is expected to take place. In contrast, when convergence is missing, local discrepancies will tend to persist. Thus, we have to investigate what kind of spatial dynamics is driving the economy.

Local developments may spread out at first in the neighboring provinces before diffusing over the entire economy. Observed spatial spillover may confirm a marked spatial dimension of the provincial adjustment processes. It may be attributed to rigidity barriers such as governmental or provincial regulations, supplier and demander mimicking behavior or imperfections of the pharmaceutical market. Not only the spatially lagged values of the dependent variable, but also those of the exogenous variables generally have to be taken into account during the transition. Thus, a local model can be established of the SAR-SDL (Spatial Autoregressive-Spatially Distributed Lag; ANSELIN, 1988; FLORAX, 1992) form (omitting throughout the t subscript for ease of notation);

$$y = \alpha_0 i_N + \alpha_1 Wy + X\beta_0 + WX\beta_1 + \nu, \qquad (3)$$

where β_0 is the first-order direct (intra-provincial) effect of X, while β_1 measures the first-order exogenous spatial spillover caused by spatial lags of X. The coefficient α_1 measures the first-order endogenous spatial spillover, while *W* is an N by N matrix defined so that w_{ij} equals 1 if provinces *i* and *j* are neighbors ($i \neq j$) and 0 otherwise. By dividing each element in W by the number of non-zero elements in the row it belongs to, the product *Wy* is conveniently understood as a variable which, for each province, holds the average of y in the neighboring provinces. It is important to understand that β_0, β_1 and α_1 measure only the first-order effects caused by X, *WX* and y respectively, but not the full aggregated effects. Calculations of the latter are considerably more involved because of the presence of contemporaneous spillover. Informally spoken, region i spills over into region j, which in turn spills over into region i, including a part of the spillover received from there. A formal treatment of this contemporaneous aggregate spillover and the way it forms the aggregate effects is technically involved and not strictly needed to understand the presentation; we refer to the Appendix for a detailed treatment.

An alternative view of the distinction between the global model (1) and the local model (3) is easily obtained by rewriting the latter as a 'speed of spatial adjustment' specification (KOSFELD and LAURIDSEN, 2004):

$$y = \kappa_0 i_N + X\kappa_1 + \theta_0 \Delta y + \Delta X\theta_1 + \varepsilon, \qquad (4)$$

with $\kappa_0 = \frac{\alpha_0}{1-\alpha_1}$, $\kappa_1 = \frac{\beta_0 + \beta_1}{1-\alpha_1}$, $\theta_0 = -\frac{\alpha_1}{1-\alpha_1}$, $\theta_1 = -\frac{\beta_1}{1-\alpha_1}$, and $\varepsilon = \frac{\nu}{1-\alpha_1}$, where $\Delta = I_N - W$ is a spatial difference operator, so that the product Δy returns a new variable measuring the discrepancy of y between a region and the surrounding regions. Comparison of (3) to (1) thus shows that the larger the values of β_1 and α_1 are, the larger the discrepancy will be between the local and the global model. Furthermore, Equation (4) shows that if α_1 equals 1, then regional discrepancies will persist and not fade out throughout the spatial structure. Thus, α_1 can be interpreted as a spatial speed of adjustment parameter: the larger α_1 is, the slower the adjustment, and vice versa for a small α_1. Model (4) is equivalent to the spatial Bewley (SBE) specification established by LAURIDSEN (2006).

The spatial adjustment may be instructively interpreted in terms of spatial error-correction. Some easy manipulations of (3) provide alternative representations (LAURIDSEN, 2006):

$$\Delta y = \alpha_0 i_N + (\alpha_1 - 1)Wy + \Delta X\beta_0 + WX(\beta_0 + \beta_1) + \nu, \qquad (5)$$

$$\Delta y = \alpha_0 i_N + (\alpha_1 - 1)Wy^* + \Delta X\beta_0 + WX(\beta_0 + \beta_1 + (\alpha_1 - 1)i_K) + \nu, \qquad (6a)$$

$$\Delta y = \alpha_0 i_N + (\alpha_1 - 1)(Wy - WX\kappa_1) + \Delta X\beta_0 + \nu, \qquad (6b)$$

where $y^* = y - Xi_K$.

Forms (5) and (6) are algebraically equivalent to (3) and (4), but provide different interpretations. Equation (5) is a spatial Baardsen specification (SBA), while models (6a) and (6b) are variants of a spatial error-correction (SEC) model (LAURIDSEN, 2006). In contrast to the

SAR-SDL, the SBA and the SEC describe the formation of expected local differences in y as depending on local differences in X and locally lagged values of X. They are distinctive in that the SBA introduces locally lagged levels in y, whereas the SEC introduces the locally lagged discrepancy between y and X. Thus, in the SEC the term $(\alpha_1 - 1)$ represents the local adjustment to any discrepancy.

To improve efficiency, and to adjust for temporal dynamics, data from several time periods should be applied to these spatial specifications if available. This is done by simply implementing (2) to the residuals of (3)-(6) so that SUR versions of these are obtained. As none of the forms (3)-(6) can be estimated consistently using Zellner's F-GLS method because of the contemporaneous y values of Wy, we apply a maximum likelihood procedure to obtain the coefficients of the SAR-SDL-SUR. From these, the parameters of (4)-(6) are straightforwardly calculated, and their standard errors obtained by applying the Delta method (GREENE, 2003) to the covariances of the SAR-SDL-SUR.

Data

Data for 50 Spanish provinces (excluding the autonomous cities of Ceuta and Melilla) were collected and are provided with this paper. These provinces correspond with the NUT-3 level of aggregation according to EUROSTAT. The provinces are assembled in 17 Autonomous Communities (AC). The ACs correspond with the NUT-2 level of aggregation according to EUROSTAT and they present a higher degree of heterogeneity than the provinces. Regarding the decentralisation process, 7 of the ACs got independent responsibilities during the 1980s and 1990s (Cataluña 1981, Andalucía 1984, Comunidad Valenciana y País Vasco 1987, Galicia y Navarra 1991, and Canarias 1994), while the last 10 got responsibility for health care regulation in 2002. Until then these 10 ACs were centrally regulated.

The data were collected annually from 1996 to 2003 from two sources, the National Statistical Institute (INE) and the Ministry of Health and Consumption (MSC). The dependent variable is Public Pharmaceutical Expenditure (EXP) per capita. This variable includes expenditures on extra-hospital drugs managed by the administration, but does not take private purchases into account. To capture the influence of wealth, Gross Domestic Product per capita (GDP) is included as an explanatory variable. Further, to capture the influence of the health care system, variables on the number of pharmacists per 1000 inhabitants (PHARM), the number of hospital beds per 1000 inhabitants (BEDS), and the number of medical doctors per 1000 inhabitants (MED) are included. Finally, to capture the influence of the population structure, the proportions of females (FEM), foreigners (FOREIGN), people over 65 years (OLD), and 0-4 year old children (CHILD) in the population are included. Table 1 presents the data used in this paper, including means and standard deviations (averaged over eight years).

The variables describing the population control for socio-demographic risk factors and are considered to be proxies for need whereas GDP controls for the ability to pay. The variables describing the health care system do not solely reflect supply factors but are a result of interactions between demand and supply factors. Some health care system variables may be considered to be substitutes to the utilization of pharmaceuticals while others are complementary. A priori, one would expect the number of pharmacists to be complementary, whereas we have no unambiguous a priori hypothesis for hospital beds and medical doctors.

Figure 2 shows the distribution of variables (averaged over eight years) by provinces. Spatial patterns are predominant, though not of a unique nature. For expenditures, a clear indication of spatial spillover is seen. Comparing the maps in Figure 2 to the map of ACs in Figure 1, this spillover seems to be of an intra- as well as a supra-AC nature. Furthermore, there seem to

Table 1: Data used in the study

Variable	Description	Source	Mean	Std. D.
EXP	Pharmaceutical Expenditure per capita	MSC, Inst. of Sanitary Information	164.899	31.710
GDP	GDP per capita	INE, National Statistical Inst.	9241.57	1766.14
PHARM	Pharmacists per 1000 inhabitant	INE, Social Indicators, 2004	1.206	0.225
BEDS	Hospital beds per 1000 inhabitants	MSC, National Hospital Catalogue	0.004	0.001
MED	Medical doctors per 1000 inhabitants	INE, Social Indicators, 2004	4.183	0.739
FEM	Population proportion of females	INE, National Statistical Inst.	0.506	0.006
FOREIGN	Population proportion of foreigners	INE, National Statistical Inst.	0.018	0.019
OLD	Population proportion over 65 years	INE, National Statistical Inst.	0.185	0.042
CHILD	Population proportion from 0 to 4 years	INE, National Statistical Inst.	0.090	0.016

be some tendencies to North/West-South/East contrasts. With respect to GDP, medical doctors, hospital beds and, to some extent, pharmacists a clear North-South contrast is evident. This is also the case for some of the population characteristics, especially proportions of elderly, children and to some extent females, while foreigners seem to cluster especially in the East coast provinces.

Results

The model assumed for pharmaceutical expenditures employs a multiplicative Cobb-Douglas specification, which is linearized by applying logarithmic transforms to the variables. BECH et al. (2006) estimated a similar model applying the SUR framework. They found that a model with common coefficients across years was adequate, as the variation over time of the coefficients could be captured by adding interaction variables between a time trend and the explanatory variables. Table 2 reports the global SUR model (i.e., without spatial spillover) and a local SAR-SDL-SUR model (adjusted for endogenous and exogenous spillover).

It is first clear that the LR test for the global versus the local model strongly rejects the former in favor of the the local model, the large coefficient of the endogenous spatial lag (*W*EXP*) clearly supports the assumption that provincial pharmaceutical expenditures are not only

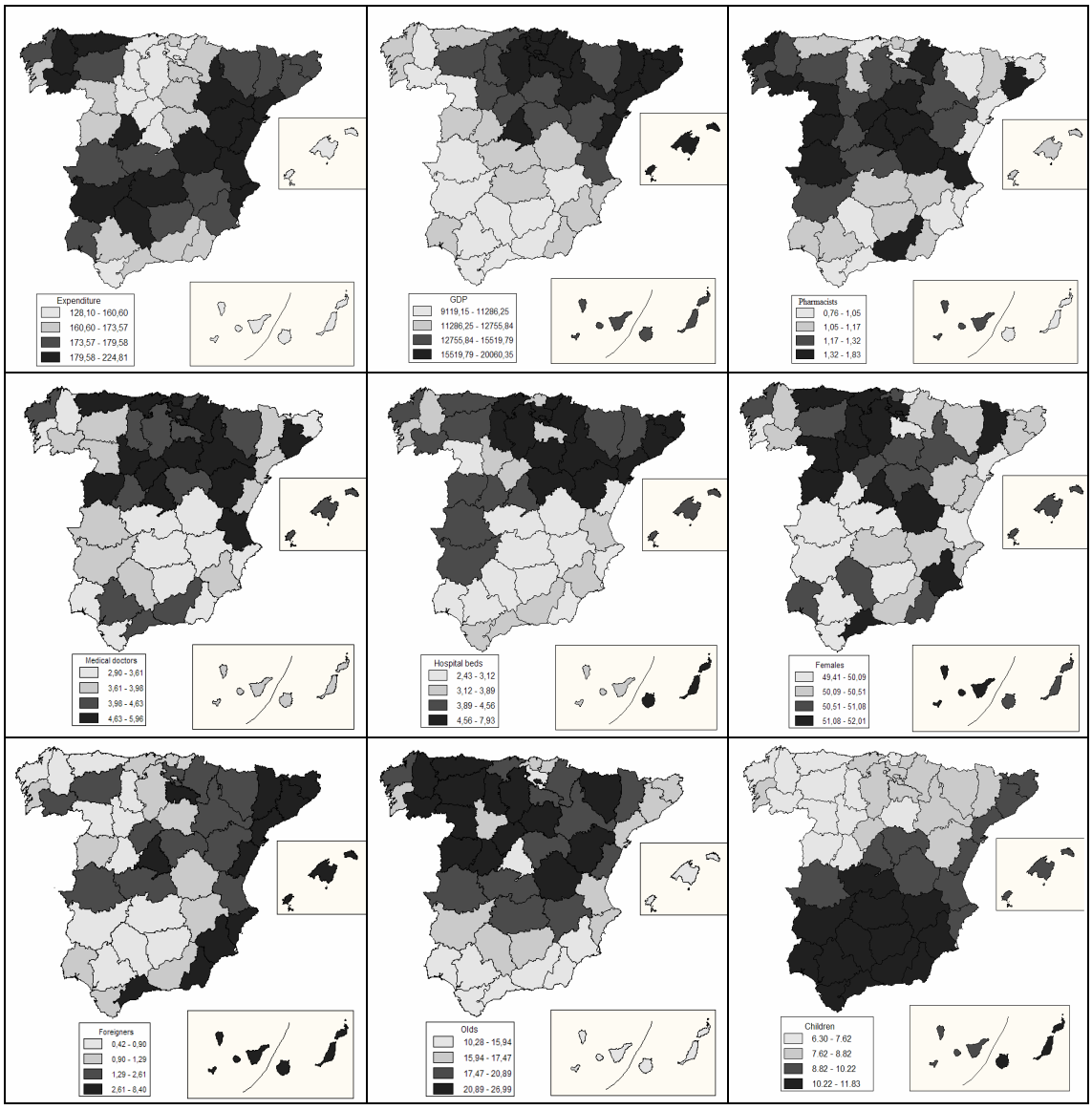

Figure 2. Variables (averaged over eight years) by province

Table 2. Global and local SUR models (dependent variable: EXP)

SUR (global)		SAR-SDL-SUR (local)				SBE-SUR (speed of spatial adjustment)			
Constant	7.645***	*Constant*	3.711***	*W*EXP*	0.659***	*Constant*	10.90***	Δ *EXP*	-1.937***
	(0.570)		(0.079)		(0.003)		(2.875)		(0.022)
GDP	0.033	*GDP*	0.092**	*W*GDP*	0.006	*GDP*	0.289*	Δ *GDP*	-0.018
	(0.034)		(0.036)		(0.063)		(0.170)		(0.185)
PHARM	0.015	*PHARM*	0.046	*W*PHARM*	0.084	*PHARM*	0.384*	Δ *PHARM*	-0.247
	(0.034)		(0.030)		(0.067)		(0.222)		(0.199)
BED	0.005	*BED*	-0.016	*W*BED*	-0.014	*BED*	-0.090	Δ *BED*	0.042
	(0.016)		(0.013)		(0.027)		(0.087)		(0.079)
MED	0.082**	*MED*	0.096***	*W*MED*	-0.139**	*MED*	-0.126	Δ *MED*	0.410**
	(0.037)		(0.035)		(0.070)		(0.197)		(0.206)
FEM	2.212***	*FEM*	-0.209	*W*FEM*	2.883***	*FEM*	7.855***	Δ *FEM*	-8.469***
	(0.628)		(0.618)		(1.014)		(2.910)		(2.981)
FOR	0.013**	*FOR*	0.005	*W*FOR*	-0.006	*FOR*	-0.001	Δ *FOR*	0.018
	(0.006)		(0.005)		(0.010)		(0.030)		(0.031)
OLD	0.231***	*OLD*	0.138***	*W*OLD*	0.020	*OLD*	0.465*	Δ *OLD*	-0.058
	(0.054)		(0.052)		(0.087)		(0.257)		(0.256)
CHILD	0.203***	*CHILD*	0.054	*W*CHILD*	0.119	*CHILD*	0.509**	Δ *CHILD*	-0.350
	(0.043)		(0.050)		(0.082)		(0.221)		(0.243)
*T*GDP*	-0.017***	*T*GDP*	-0.025***	*W*T*GDP*	0.008	*T*GDP*	-0.051***	Δ *T*GDP*	-0.024
	(0.004)		(0.005)		(0.007)		(0.017)		(0.022)
*T*PHARM*	0.008*	*T*PHARM*	0.001	*W*T*PHARM*	-0.006	*T*PHARM*	-0.017	Δ *T*PHARM*	0.018
	(0.005)		(0.003)		(0.009)		(0.030)		(0.027)
*T*BED*	0.001	*T*BED*	0.005**	*W*T*BED*	-0.004	*T*BED*	0.003	Δ *T*BED*	0.012
	(0.003)		(0.002)		(0.004)		(0.014)		(0.013)
*T*MED*	-0.018***	*T*MED*	-0.016***	*W*T*MED*	0.017*	*T*MED*	0.002	Δ *T*MED*	-0.050*
	(0.005)		(0.005)		(0.009)		(0.025)		(0.028)
*T*FEM*	-0.142***	*T*FEM*	0.058	*W*T*FEM*	-0.284***	*T*FEM*	-0.663***	Δ *T*FEM*	0.836***
	(0.053)		(0.058)		(0.096)		(0.256)		(0.284)
*T*FOR*	0.001	*T*FOR*	-0.001	*W*T*FOR*	0.004**	*T*FOR*	0.006*	Δ *T*FOR*	-0.010**
	(0.001)		(0.001)		(0.001)		(0.003)		(0.004)
*T*OLD*	-0.002	*T*OLD*	0.003	*W*T*OLD*	0.008	*T*OLD*	0.032	Δ *T*OLD*	-0.022
	(0.005)		(0.006)		(0.009)		(0.026)		(0.027)
*T*CHILD*	-0.039***	*T*CHILD*	-0.013	*W*T*CHILD*	-0.001	*T*CHILD*	-0.041*	Δ *T*CHILD*	0.001
	(0.007)		(0.009)		(0.012)		(0.022)		(0.036)
LogL = 1367.01		LogL = 1432.01							
AIC = -2626.03		AIC = -2724.02							
LR test of global SUR versus local SAR-SDL-SUR = 130.00***									

Note. Standard errors in parentheses. Significance indicated at 1% (***), 5% (**) and 10% (*) levels.

Table 3. Spatial error-correction representations of the local model (dependent variable: ΔEXP)

	SBA-SUR			SEC-SUR (formula 6a)		
Constant	3.711***	W*EXP	-0.340***	3.711***	W*EXP*	-0.340***
	(0.979)		(0.002)	(0.979)		(0.002)
Δ*GDP*	0.092**	W*GDP	0.098*	0.092**	W*GDP	-0.248***
	(0.036)		(0.058)	(0.036)		(0.058)
Δ*PHARM*	0.046	W*PHARM	0.130*	0.046	W*PHARM	-0.209***
	(0.030)		(0.075)	(0.030)		(0.075)
Δ*BED*	-0.016	W*BED	-0.030	-0.016	W*BED	-0.371***
	(0.013)		(0.029)	(0.013)		(0.029)
Δ*MED*	0.096***	W*MED	-0.043	0.096***	W*MED	-0.383***
	(0.035)		(0.067)	(0.035)		(0.067)
Δ*FEM*	-0.209	W*FEM	2.674***	-0.209	W*FEM	2.333**
	(0.618)		(0.990)	(0.618)		(0.990)
Δ*FOR*	0.005	W*FOR	-0.001	0.005	W*FOR	-0.340***
	(0.005)		(0.010)	(0.005)		(0.010)
Δ*OLD*	0.138***	W*OLD	0.158*	0.138***	W*OLD	-0.182**
	(0.052)		(0.087)	(0.052)		(0.087)
Δ*CHILD*	0.054	W*CHILD	0.173**	0.054	W*CHILD	-0.166**
	(0.050)		(0.075)	(0.050)		(0.075)
Δ*T*GDP*	-0.025***	W*T*GDP	-0.017***	-0.025***	W*T*GDP	-0.358***
	(0.005)		(0.005)	(0.005)		(0.006)
Δ*T*PHARM*	0.001	W*T*PHARM	-0.005	0.001	W*T*PHARM	-0.346***
	(0.003)		(0.010)	(0.003)		(0.010)
Δ*T*BED*	0.005**	W*T*BED	0.001	0.005**	W*T*BED	-0.339***
	(0.002)		(0.005)	(0.002)		(0.005)
Δ*T*MED*	-0.016***	W*T*MED	0.001	-0.016***	W*T*MED	-0.339***
	(0.005)		(0.008)	(0.005)		(0.009)
Δ*T*FEM*	0.058	W*T*FEM	-0.226***	0.058	W*T*FEM	-0.566***
	(0.058)		(0.087)	(0.058)		(0.087)
Δ*T*FOR*	-0.001	W*T*FOR	0.002*	-0.001	W*T*FOR	-0.338***
	(0.001)		(0.001)	(0.001)		(0.002)
Δ*T*OLD*	0.003	W*T*OLD	0.011	0.003	W*T*OLD	-0.329***
	(0.006)		(0.008)	(0.006)		(0.009)
Δ*T*CHILD*	-0.013	W*T*CHILD	-0.014*	-0.013	W*T*CHILD	-0.354***
	(0.009)		(0.007)	(0.009)		(0.008)

Note. Standard errors in parentheses. Significance indicated at 1% (***), 5% (**) and 10% (*) level

determined within provinces, but that important supra-provincial forces are in play. These could be attributed to governmental and provincial regulation or to the supra-latter. Looking at the coefficients of the spatial lags of provincial nature of the pharmaceutical market caused by the presence of pharmaceutical companies of such a size that they can partly outperform the regional self-determination of expenditure. Further, several spatial lags of exogenous determinants are significant and thus indicate that supra-provincial actors react to supra-provincial levels of determinants, rather than to intra-provincial levels exclusively. Further, a comparison of the global and the local models clearly shows that the direct (i.e., non-spatial) effects of population characteristics are overstated by the global model, while the effect of GDP seems to be understated. On the other hand, the effects of health system characteristics seem to be properly estimated by the global model.

A closer look at the speed of spatial adjustment (SBE-SUR) transform of the local model provides a deeper insight into the nature of the spatial adjustment process. The large negative coefficient for the local discrepancy of expenditure (ΔEXP) clearly shows that a local excess expenditure of a province (i.e., a positive ΔEXP) causes a downward adjustment, so that a local convergence of expenditure is in play. Likewise, for the exogenous determinants, the inversely-signed nature of the direct effects and the discrepancy effects indicate the presence of local convergence. These exogenous discrepancy effects are, however, relatively small and weakly significant. To illustrate, the direct effect of GDP is 0.289. If, however, the GDP of a province exceeds the average GDP of the surrounding provinces, then the effect is downward adjusted by 0.018 times the difference ΔGDP. A somewhat stronger effect is in play for the number of pharmacists and the number of medical doctors. For the former, the direct effect is 0.384, while the downward adjustment effect of a positive local difference is 0.247 times the discrepancy in the number of pharmacists. For the latter, the negative direct effect of 0.126 is modified in the case of a positive local difference by a positive adjustment effect of 0.410 times the difference in the number of medical doctors.

Table 3 summarizes the spatial error-correction specification, which provides further insight into the local dynamics. It is seen that local differences in GDP cause local differences in pharmaceutical expenditures. Regarding health care system characteristics, it is obvious that local differences in the number of medical doctors cause local expenditure differences. Turning to population characteristics, it is apparent that local differences in the proportions of old people cause differences in expenditures. In both models in Table 3 the coefficients are the same (-0.340) for the endogenous spatial lag. While it returns simply the spatial lag of the dependent variable in the SBA-SUR model, in the SEC-SUR it embodies spatial error-correction mechanisms. The adjustment coefficient (W*EXP*) takes a negative sign, which generally indicates the working of the error-correction mechanism. As an effective spatial error-correction allows the economy to adjust towards equilibrium by fast absorption of a spatial spillover, it is straightforwardly linked with the existence of conditional convergence. However, due to the large endogenous spillover (indicated with $\alpha_1 = 0.66$) the error-correction (indicated by $\alpha_1 - 1 = -0.34$, which is considerably larger than -1), and thus the speed of absorption of the spatial spillover, is relatively slow. Thus, important rigidities are demonstrated which slow down the opportunities for convergence.

Conclusions

The present study analyzes models of public pharmaceutical expenditures and adds to previous knowledge regarding not only the necessity but also the involved implications of adjusting for spatial spillover effects. The importance of interpreting spatial spillovers within a spatially dynamic error-correction framework is demonstrated. Specifically, it is found that opportunities for convergence are slowed down, as the speed of absorption of spatial spillovers is considerably slow. This in turn is especially due to the presence of a large endogenous spatial spillover, presumably caused by the complex supra-provincial nature of the pharmaceutical market. Thus, the complexity of spatial spillover is clearly illustrated, and the need for further evidence on the implications and the nature of spatial spillover is demonstrated.

SAS code and data sets are available on the CS-BIGS web site.

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Appendix[1]

The distinction between first-order and aggregate spatial spillover was informally touched upon in the treatment of the SAR-SDL model (3). The following is devoted to a formal treatment of this distinction. Specifically, with K explanatory variables, the data generating process for the SAR-SDL can be written as

$$y = \sum_{k=1}^{K} S_k(W)x_k + V(W)i_N + V(W)\varepsilon$$

where

$$S_k(W) = V(W)(I_N\beta_{0(k)} + W\beta_{1(k)})$$

and

$$V(W) = (I_N - \alpha_1 W)^{-1} = I_N + \alpha_1 W + \alpha_1^2 W^2 + \alpha_1^3 W^3 + \ldots$$

as k indexes variable k in the model, so that x_k is the kth explanatory variable, i.e., the kth column of X.

To illustrate the role of $S_k(W)$, consider the expansion of the data generating process equation as:

$$\begin{pmatrix} y_1 \\ y_2 \\ . \\ . \\ y_N \end{pmatrix} = \sum_{k=1}^{K} \begin{pmatrix} S_k(W)_{11} & S_k(W)_{12} & . & . & S_k(W)_{1N} \\ S_k(W)_{21} & S_k(W)_{22} & . & . & . \\ . & . & . & . & . \\ . & . & . & . & . \\ S_k(W)_{N1} & S_k(W)_{N2} & . & . & S_k(W)_{NN} \end{pmatrix} \begin{pmatrix} x_{1k} \\ x_{2k} \\ . \\ . \\ x_{Nk} \end{pmatrix} \cdot$$

$$+ V(W)i_N\alpha_1 + V(W)\varepsilon$$

from which it is clear that the aggregate effect of x_{jk} on y_i is given by

$$\frac{\partial y_i}{\partial x_{jk}} = S_k(W)_{ij} \cdot$$

Especially, it follows further that the aggregate effect of x_{ik} on y_i is given by

$$\frac{\partial y_i}{\partial x_{ik}} = S_k(W)_{ii} \cdot$$

Thus, the aggregate spatial spillover effect as well as the aggregate non-spatial effect of x_k is a complex function, capturing the first-order effects as given by β_0 and β_1 as well as connectivity/feed back loop influences. Feed back loop influences arise from impacts passing through neighbors, and back to the observation itself. These will vary by the location of the observation and the connectivity among observations governed by the spatial weight matrix W.

[1] The authors are indebted to an anonymous referee, who provided a substantial part of the derivations in this Appendix.