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Abstract. In view of the lack of area-wide regional price data, Aten and Heston (2005) adopted an econometric approach in estimating regional price level from an international perspective. At a regional level, however, spatial economic theory may contribute to advancement in explaining price level determination. Recently, the market potential approach as an essential element of NEG models has been employed for explaining observed regional land prices (Brakman et al., 2004). The present paper relies on the price mechanisms of the Helpman model (Helpman, 1998) in developing spatial-econometric models for regional price level and its major components. By evaluating a south German sample, first evidence on the empirical significance of NEG-based price level models is provided.

Key words: Regional price level, Helpman model, spatial-econometric models, ESDA

JEL: C21, R13, R31

1. Introduction

In international studies of economic growth and development, real income based on purchasing power parity (PPP) is ordinarily used as an indicator of the level of development and standard of living. Nominal income measures fail to mirror wealth or productivity gradients as cross-country differences in income per capita or labour productivity can be attributed, to a large extent, to price differences and purchasing power parities (see e.g. Jones, 2002). Regional economic theory demonstrates that the law of one price does not hold either across regions (see e.g. McCann, 2001). In fact, from some price surveys it is clear that substantial price differentials do occur at a regional level (Rostin, 1979; Ströhl, 1994).

Newer regional economic theories point out that spatial price differentials are expected to play a crucial role in shaping the economic landscape. The price index effect highlighted in New Economic Geography (NEG) models gives reason for the existence of forward linkages that operate towards agglomeration (Krugman, 1991; Fujta et al., 1999). While low commodity prices may most notably be observed in big cities, prices of non-tradable goods tend to fall with distance from centres (Tabuchi, 2001). Due to congestion effects, land scarcity and other urban costs, prices of non-tradables tend to be high in agglomerated areas. In Helpman's NEG model (Helpman, 1998), the immobile housing sector acts as a force towards dispersion. In contrast to Krugman's fundamental core-periphery model (Krugman, 1991; Fujita et al., 1999), Helpman's model is able to explain regional price level as a result of the tension between opposing centripetal and centrifugal forces.

EU regional and cohesion policy aims to promote regional development, improve the environment and develop transport infrastructure. Development programmes are financed mainly by the Structural Funds and the Cohesion Fund, which account for about 95% of all community regional aid. The largest regional assistance is made for so-called objective 1-areas defined by a GDP per capita of 75% or less of the EU average. Due to a lack of area-wide price data, at present criteria for promotion as well as the control of efficiency are mainly based on nominal regional income. As regional income gradients can be partly or completely compensated for by spatial price differentials, the extent of distortions may seriously matter. The same applies to testing for regional convergence and New Economic Geography (NEG) models. In testing the popular market potential approach, for instance, a wage equation is usually derived under the assumption of equal price levels in all regions (cf. Roos, 2001; Hanson, 2005; Niebuhr, 2004).[1]

National statistical offices do not gather price data area-wide. Hence, price levels are generally not known at a regional level of districts or provinces. In most countries price indices are even missing on the state level. Official statistics confirm that the collection of regionally disaggregated data all over the country would be extremely expensive (Ströhl, 1994; BSMWVT, 2003). Given this current state, econometric techniques have recently been employed for estimating regional price levels. Aten and Heston (2005) estimate regional price levels using spatial-econometric models calibrated with national consumer price indices.

Although the econometric approach seems to be promising, the breakdown of country estimations to a regional level is not easily justified. First, the econometric models built from international studies are primarily demand-orientated and not by grounded in regional economic theory. Second, the calibration with national consumer price index does not

[1] Alternatively, prices can be eliminated by assuming equal real wages (c.f. Roos, 2001; Hanson, 2005; Niebuhr, 2004). Fingleton (2004) determines regional price levels of traded goods by nominal regional wages and a distance-decay function for the purpose of testing New Economic Geography theory vs. urban economics theory in explaining spatial concentration of economic activity.

necessarily imply an adequate explanation of regional price levels. Third, there is not a priori guarantee that responses of explanatory variables and spatial effects at the national and regional level are identical.

Further development of the econometric approach to regional price level estimation may be achieved by addressing the above mentioned aspects. The German Federal Statistical Office conducts comparisons on price levels in larger time intervals across selected cities (Rostin, 1979; Angermann, 1989, Ströhl, 1994). The last price comparison involved 50 cities and took place in 1993 (Ströhl, 1994). Unfortunately, with the exception of the four-city comparison in 1987, the non-traded goods are only represented fragmentarily in consumer price indices. In price level comparisons of 1978 and 1993 housing rents are completely disregarded despite their share of over 20 per cent in the basket of goods and services. Evaluating regional price level models by such consumer price indices would take away an essential centripetal force from the outset.

To the best of our knowledge, in Germany price level data on the district level (*kreisfreie Städte* and *Landkreise*) covering all main groups of the basket of goods and services have only been made available from surveys of the Bavarian State Office for Statistics and Data Processing. The price data were used for comparisons of real purchasing power across Bavarian regions by the Bavarian State Ministry of Economic Affairs, Transport and Technology (BSMWVT, 2003). The regional income study not only provides consumer price indices for different-sized districts (large, moderate, small), but also allows for analyzing sub-price indices of tradables, non-tradables and housing.

The present paper deals with developing regional price level models on the groundwork of NEG theory. Up to now, NEG theory has been drawn only partially for regional price level determination (Brakman et al., 2004). We particularly rely on the Helpman model (Helpman, 1998) in building spatial econometric models for a consumer price index and its major components. Spatial price effects are analyzed by means of exploratory spatial data analysis (ESDA) and spatial Lagrange multiplier (LM) tests (Anselin, 1995; Anselin et al., 1996) as well as by trend surface and spatial expansion models (Anselin, 1992; Jones and Cassetti, 1992; Casetti and Poon, 1995). First empirical evidence on the significance of NEG theory in explaining regional price level is provided from the southern German sample.

The subsequent sections of this paper are organized as follows. In section 2 we discuss economic forces relevant for regional price level determination from the perspective of the Krugman and Helpman model. Section 3 deals with spatial methods and models for analyzing regional price indices. While section 3.1 introduces tools of exploratory spatial data analysis (ESDA), section 3.2 deals with spatial-econometric model building. The regional data set is described in section 4. In section 5, we analyze spatial price effects and evaluate different price level models empirically. Section 6 concludes.

2. Regional price indices and Helpman model

Models of New Economic Geography (NEG) explain the development of agglomerations by means of increasing returns to scale, transportation costs and factor mobility. The Krugman model represents a prototype that consists of a two sectoral core-periphery structure (Krugman, 1991; Fujita et al., 1999, Ch. 4). In the modern sector varieties of a differentiated good are produced with increasing returns to scale by mobile workers across regions. The manufacturing sector is characterized by monopolistic competition. Industrial goods are traded across regions with transport costs increasing with distance.[2] Goods produced in the

[2] In NEG models, transport costs are typically modelled by an "iceberg technology". Only a fraction τ of a variety shipped arrives at the destination. The part $1-\tau$ that "melts" away increases with distance.

traditional sector are homogenous and termed "agricultures". The traditional sector operates with constant returns to scale under perfect competition. For agricultures no transport costs are incurred. Workers in the traditional sector are viewed to be immobile.

Forward linkages between firms and consumers act as centripetal forces that are strongly based on the price index effect (Robert-Nicoud, 2005). When firms move from one region to another, consumers will find a larger range of manufactured goods in the destination region. Because of their preferences for variety, they will benefit from the enlarging number of firms at their place of residence by saving transport costs. As each firm produces only one single variety in equilibrium, consumers' expenditures are spread over a larger number of differentiated goods. As a result, the price index of tradable goods (P_T) will decrease in regions with an increasing number of firms. This effect can be seen from the representation

$$P_{T,r} = \left[\sum_{s=1}^{n} N_s \left(p_s^M \cdot T_{sr} \right)^{1-\sigma} \right]^{1/(1-\sigma)} \quad , \ r=1,2,\ldots,n, \tag{1}$$

of the price index for tradables in region r. p_s^M is the uniform price of manufactures in region s, N_s the number of varieties (= number of manufacturing firms) produced in region s and T_{sr} ($T_{sr}>1$) the transport costs incurred by shipping a variety from s to r. The parameter σ denotes the elasticity of substitution between any two varieties.[3] As all prices of manufactures are equal in equilibrium, the price of a variety produced by a firm that moves from s to r drops from $p_r^M T_{rs}$ to p_r^M. By virtue of the price index effect core-periphery models of Krugman-type predict a lower price level of industrial goods in agglomerations compared to peripheral areas (Krugman, 1991; Fujita et al., 1999, Ch. 4; Fujita and Thisse, 2002, Ch. 9).

The price of the non-traded good ("agriculture"), P_{NT}, is fixed and identical in all regions. Thus, the overall price index of Cobb-Douglas type,

$$P_r = P_{T,r}^{\mu} \cdot P_{NT,r}^{1-\mu}, \tag{2}$$

with μ as the share of expenditures spent on manufactures, does not cover any centrifugal forces. Potential centrifugal forces that might arise from congestion or housing costs are completely removed. Retaining the structure of the modern industrial sector, Helpman (1998) introduces a dispersion force into Krugman's core-periphery model by replacing the agricultural good with housing. As stock of housing, H_r, is fixed in all regions, prices for housing tend to be high in densely populated centres and low in sparsely populated areas. In equilibrium, housing income and housing expenditures must be equalized:

$$P_{NT,r} \cdot H_r = (1-\mu) Y_r . \tag{3}$$

When income Y_r rises with immigration of firms and workers, housing rents $P_{NT,r}$ will increase if housing stock H_r is constant:

$$P_{NT,r} = (1-\mu) Y_r \cdot H_r^{-1} . \tag{4}$$

Thus, overall regional price level depends on the relative strength of centrifugal and centripetal forces. While transport costs act as an agglomeration force, housing costs operate towards dispersion.

The equilibrium relation for the price index of tradables reads

[3] By virtue of normalization, σ also represents the price elasticity of demand of varieties.

$$P_{T,r} = \left[\sum_{s=1}^{n} \lambda_s \cdot \left(w_s \cdot T_{sr} \right)^{1-\sigma} \right]^{1/(1-\sigma)} , \quad (5)$$

where λ_s is the share of region's share of total manufacturing labour force. Transport costs T_{sr} incurred by shipping a unit of manufactures from s to r are measured as a function of distance d_{sr} between s and r:

$$T_{sr} = T\left(d_{sr}\right), \; f\left(d_{sr}\right) > 0 . \quad (6)$$

Using the equilibrium condition that real wages are equalized across regions,

$$\frac{w_r}{P_{T,r}^{\mu} \cdot P_{NT,r}^{1-\mu}} = \frac{w_s}{P_{T,s}^{\mu} \cdot P_{NT,s}^{1-\mu}} \text{ for all } r \neq s, \quad (7)$$

the price index for tradable goods of region r, $P_{T,r}$, can be restated in terms of the fundamental economic variables Y_r, H_r and w_r:

$$P_{T,r} = \gamma \cdot Y_r^{\mu\text{-}1} \cdot H_r^{(1-\mu)/\mu} \cdot w_r^{1/\mu} . \quad (8)$$

In (8) γ is a constant. Let h_r be each individual's share in total housing income and L the labour force. Then region r's total income amounts to

$$Y_r = \lambda_r \cdot L \cdot (w_r + h_r) . \quad (9)$$

Because of housing is equally owned by individuals, h_r is a constant (Roos, 2001). Thus it may be advantageous to solve equation for w_r, in order to eliminate wage in the representation (8) of $P_{T,r}$. If we neglect the share of housing income, h_r, for simplification[4], the price index of tradable goods takes the form

$$P_{T,r} = \phi \cdot Y_r \cdot H_r^{(1-\mu)/\mu} \quad (10)$$

with ϕ as a constant.

Equations (4) and (10) can be viewed as the major components of NEG-based econometric models for compound prices for tradable and non-tradable goods, respectively.[5] Income asserts a positive influence on both sub-price indices, whereas the effect of housing stock is different. While an increasing housing stock entails a fall in housing rents, it is at the same time accompanied by a rise of prices of manufactures. If we combine both price equations (4) and (10) according to the Cobb-Douglas type overall price index (2), the housing variable would cancel out. Higher cost of living due to scarcity of housing stock is completely offset by a fall in prices of manufactures. Housing costs are not such a substantial centrifugal force to be able to outweigh opposite concentration forces. This potential weakness of the Helpman model can be tested in the space of a regional price level model.

[4] Cf. Brakman et al. (2004) who define income from the start only in terms of wages.

[5] The basic idea of combining of some equilibrium relationships in order to obtain testable "price equations" resembles Hanson's derivation (Hanson, 2005) of a NEG-based market potential function in form of the so-called wage equation for empirical testing purposes (see e.g. Mion, 2003; Niebuhr, 2004; Brakman et al., 2004).

3. ESDA and spatial-econometric price level models

3.1 Tools of exploratory spatial data analysis (ESDA)

Exploratory spatial data analysis (ESDA) aims at identifying spatial properties of data for detecting spatial patterns, formulating hypotheses for geo-referenced variables and assessing spatial models (Haining and Wise, 1997). Instead of hypothesis testing, the toolbox of ESDA primarily consists of descriptive and graphical methods. In recent times much work has been done in developing locational tools for spatial data analysis. These tools focus on detecting spatial instationarities by identifying atypical areal units. In our exploratory spatial data analysis of regional price levels we first introduce a concept for measuring global spatial autocorrelation. Subsequently we give an outline of matching methods for exposing spatial heterogeneity.

Spatial dependence of regional price levels[6], P, can be measured by choosing a special variant of the general cross-product statistic

$$\Gamma = \sum_{r=1}^{n} \sum_{s=1}^{n} w_{rs} \cdot A_{rs} . \qquad (11)$$

Γ reflects a match between locational and attribute similarity. While locational similarity is captured by the (standardized) spatial weight matrix $\mathbf{W} = [w_{rs}]$ of dimension nxn, attribute similarity is here measured by a function A of price levels in region r and s.

Basically the elements w_{rs} of the weight matrix $\mathbf{W}$ can be based on the concepts of contiguity and distance (Anselin, 1988a, pp. 17). As we will work with a sample of non-contiguous counties and cities, only the latter concept proves to be feasible. Geographically distance-based weights rely on the hypothesis that spatial interaction between two regions r and s decreases with greater remoteness.[7] This may be attributed to increasing costs of moving people and goods between areal units. The relevance of geographical nearness finds itself in Tobler's first law of geography stating that "everything is related to everything else, but near things are more related than distant things" (Tobler, 1979).

Cliff and Ord (1981, pp. 17) suggest a general spatial weight matrix built on both distance and common border length (Cliff and Ord, 1981, pp. 17). Usually, however, common border length is viewed as circumstantial with regard to spatial interaction. Thus, the Cliff-Ord weight weights reduce to the Pareto form[8]

$$w_{rs}^{*} = 1/d_{rs}^{-b}, \ b > 0. \qquad (12)$$

The entries w_{rs}^{*} are termed unstandardized spatial weights. The distance decay parameter b controls the degree of downweighting of prices from spatial units with increasing distance to region r. The larger b, the less important are goods prices in far remote areas for region r's own price level. Simple inverse distance weights are given by a distance decay parameter of 1. The gravity model, however, causes the parameter b to be set equal to 2 (see e.g. Zhang and Kristensen, 1995; Bang, 2005). Inverse quadratic distance weights have proved better at

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6 For the ease of exposition, here we use simple the symbol P for any price indices.

7 For a generalization of the geographical distance concept to the notion of an economic distance see e.g. Conley and Ligon (2002). Economic distances between spatial units could be defined by, for instance, trade flows or transport costs.

8 The diagonal elements w_{rr}^{*} are set equal to 0.

adjusting for impedances like traffic congestion and topological barriers that tend to occur more frequently with growing distance (Gimpel et al., 2003).

A row-standardization of the original spatial weight matrix **W*** is preferable for a better interpretation (see e.g. Beck et al., 2004):

$$w_{rs} = w^{*}_{rs} / \sum_{s=1}^{n} w^{*}_{rs} \,. \tag{13}$$

For the row-standardized weight matrix **W** = [w_{rs}] the range of parameter space is no longer dependent on the scale of the distance (Morenoff et al., 2001). Moreover, the spatial lags are measured in the same units as the attribute variable.

Moran's I is most often used as a global measure of spatial autocorrelation. It is based on the general cross-product statistic with the functional choice

$$A_{rs} = (P_r - \overline{P}) \cdot (P_s - \overline{P}) \tag{14}$$

for the attribute variable P:

$$I = \frac{\sum_{r=1}^{n} \sum_{s=1}^{n} w_{rs} (P_r - \overline{P})(P_s - \overline{P})}{\sum_{r=1}^{n} (P_r - \overline{P})^2} \,. \tag{15}$$

Row-standardization of the spatial weight matrix guarantees that the Moran coefficient will approach 1 in case of perfect positive spatial autocorrelation.[9] According to (15) Moran's I turns out to be the ratio of the sum of cross-products of the regional price level P with its spatial lag W_P,

$$W\_P_r = \sum_{s=1}^{n} w_{rs} \cdot P_s \,, \tag{16}$$

to the sum of squared price level deviations from the mean $\overline{P}$. Thus, with a standardized weight matrix W, Moran's I can be interpreted as the slope of a regression of the spatial lag variable W_P on the geo-referenced variable P.[10] For spatially independently distributed attribute variables, I has the expected value E(I) = -1 / (n-1). Values of I above E(I) point to positive spatial autocorrelation, whereas values below E(I) indicate negative spatial autocorrelation.

The prevalence of positive spatial price dependencies can arise by diverse stochastic spatial processes. A global pattern of spatial autocorrelation, however, may mask spatial heterogeneity. Not only the strength but even the direction of spatial dependency can vary significantly across space. Atypical regions may exert considerable influence on the overall picture. Spatial instability in the global pattern can be revealed with the aid of tools of local spatial analysis. For detecting spatial heterogeneity in regional price levels we particularly make use of the local Moran coefficient and the Moran scatterplot (Anselin, 1995).

The local Moran coefficient,

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9 With the unstandardized weight matrix **W*** instead of **W**, a scaling factor has to be introduced (see e.g. Upton and Fingleton, 1985, p. 170). Using **W**, influences from outside regions upon the areal units is normalized.

10 Note that despite this interpretation the standard t-test for regression coefficients is not applicable (see Cliff and Ord, 1981, pp.42).

$$I_r = \frac{(P_r - \overline{P}) \sum_{s=1}^{n} w_{rs}(P_s - \overline{P})}{\sum_{r=1}^{n} (P_r - \overline{P})^2}, \tag{17}$$

is a local indicator of spatial association (LISA) that measures the contribution of the *r*th region to Moran's I as to the attribute variable P. Up to a factor of proportionality, the I_r's average to the global I. In the usual case of contiguous spatial arrays, LISA indicates spatial clustering of similar attribute values. When working with a sample of non-contiguous areal units, local Moran's I provides a somewhat different information. For districts in line with the global pattern of spatial autocorrelation, the I_r's indicate the strength of spatial price dependence with nearby regions.[11] Thus, we can establish which regions contribute above and below average to global Moran's I. Our interest may, however, lie even more in identifying 'pockets of instability' in the form of spatial units that do not share the prevailing orientation. They are formed by outlying regions with respect to the global spatial correlation structure. Given a positive global spatial autocorrelation, I_r values above the expectation -1/(n-1) – particularly positive I_r values – indicate regional conformity with the overall tendency.[12] On the other hand spatial units with I_r values below the expectation -1/(n-1) are potential spatial outliers with respect to orientation.

The Moran scatterplot is a flexible exploratory tool that supplements information provided by local Moran's I. It consists of a bivariate plot of spatially lagged (sub-)price levels against the consumer (sub-)price index after standardization. By the quadrant scheme both groups of areas with local negative and positive spatial autocorrelation can be divided according to the kind of similarity or dissimilarity with nearby regions. In Table 1 four types of local spatial association are distinguished according to possible positions of the pairs (P_r, W_P_r) in the Moran scatterplot.

Table 1: Types of local spatial association

		Spatially lagged price level (W_P)	
		high	low
Regional price level (P)	high	Quadrant I: HH	Quadrant IV: HL
	low	Quadrant II: LH	Quadrant III: LL

With the aid of the Moran scatterplot areas with a coinciding local and global autocorrelation pattern can classified in HH or LL type regions. Spatial units lying in quadrant II (LH type) and IV (HL type) have a low (high) own price level, but they are surrounded by nearby regions with a high (low) price level.

According to the match of Moran's I with the slope of a spatial regression, various diagnostics for identifying outliers used in standard regression analysis can be employed. We make particular use of the "hat" values and Cook's distance (Cook and Weisberg, 1982) for identifying leverage points and influential values with respect to spatial association.

[11] Nearby regions are here defined by distance. They need not necessarily be neighbours.

[12] Global negative spatial autocorrelation is rarely observed in reality. It describes patterns where neighbouring or near-distant areas are unlike.

3.2 NEG-based spatial-econometric price level models

The price index for tradable goods (10) in the Helpman model is shown to be a non-linear function of income and housing stock. Nonlinearity is as well present in the determination of prices for housing (4) by virtue of the interaction of income and housing stock. Both equations can, however, be linearized by taking the logarithms on both sides. Summarizing the effects of all ignored influences by error terms ν and η, we obtain the log-log models

$$\ln P_{T,r} = \ln \phi + \ln Y_r + \frac{1-\mu}{\mu} \cdot \ln H_r + \nu_r \tag{10'}$$

and

$$\ln P_{NT,r} = \ln(1-\mu) + \ln Y_r - \ln H_r + \eta_r \tag{4'}$$

for sub-price indices of tradables and non-tradables. The models (10') and (4') represent the core of NEG-based econometric regional price level models. Additional variables may enter the model as control variables. In international price comparisons, it is particular controlled for geography, while income, openness and human capital are used as economic influence variables in a largely ad hoc manner (Aten and Heston, 2005).

According to equation (2) consumer price index P is determined by all model variables affecting P_T and P_{NT}. Let α_0 be the intercept, β_j the regression coefficients of the NEG-variables and γ_k the regression coefficients of the control variables X_k. Then the NEG-based econometric model for the consumer price level is of the form

$$\ln P_r = \alpha_r + \beta_1 \cdot \ln Y_r + \beta_2 \cdot \ln H_r + \sum_{k=1}^{m} \gamma_k \cdot X_{kr} + \varepsilon_r \tag{18}$$

with ε as the error term. The Helpman model imposes the following parameter restrictions:

for $\ln P_{T,r}$: $\beta_1 > 0$, $\beta_2 > 0$, (19)

for $\ln P_{NT,r}$: $\beta_1 > 0$, $\beta_2 < 0$, (20)

for $\ln P_r$: $\beta_1 > 0$, $\beta_2 = 0$. (21)

Instead of using individual geographic variables like climate, height, precipitation, water access, we capture spatial heterogeneity by latitude and longitude. They exert direct effects on regional prices, but as well captured influence of individual geographic variables not explicitly considered (Aten and Heston, 2005). Both geographic variables may interact with economic forces like income.

If spatial effects are ignored, the disturbances ε_r will not be an independently identically distributed random variable. They can manifest in form of spatial heterogeneity and/or spatial dependence (Anselin, 1988b, pp. 8; LeSage, 1999, pp. 3). The former type of spatial effect is strongly linked to nonstationarities of price level itself or in its relationship to other variables across space. Spatial autocorrelation refers to the fact that near phenomena are often more strongly related than distant matters of fact. Significance tests may become invalid and parameter estimates biased if spatial effects are ignored.

Spatial heterogeneity of regional prices may be present in form of south-north and/or west-east price gradients. In this case it can be captured by a trend surface model which consists of a polynomial equation in the coordinates x_r and y_r of a representative point in region r (Anselin, 1992). A second order trend surface model, for instance, takes the form

$$z_r = a + b_1 \cdot x_r + b_2 \cdot y_r + b_3 \cdot x_r^2 + b_4^2 \cdot y_r^2 . \tag{22}$$

Rural districts are usually represented by the latitude and longitude coordinates of the district town. When the district town is itself an urban district, it is replaced by the next largest city in the district. The polynomial trend surface (21) is entered into the price level model (20) omitting the nonsignificant effects.

In case of spatial heterogeneity caused by space interacting with economic price determinants, the global price level model (20) may be expanded over space by allowing regression coefficients to be functions of the coordinates x_r and y_r. A simple linear spatial expansion of the income parameter β_1 reads

$$ß_{1r} = \beta_0 + \beta_{1x} \cdot x_r + \beta_{1y} \cdot y_r. \tag{23}$$

Replacing $ß_{1r}$ with β_1 in equation (20) and separating the direct income effect from the interactions effects gives the spatial expansion model (Casetti, 1997; Jones and Casetti, 1992). By means of spatial expansion, trends in the regression coefficients over space can be measured.

Substantive spatial price dependence can be accounted for by introducing a spatial price lag (16) with ln P_r instead of P_r,

$$W\_\ln P_r = \sum_{s=1}^{n} w_{rs} \cdot \ln P_s \,, \tag{24}$$

into the non-spatial model (18):

$$\ln P_r = \alpha_r + \rho \sum_{s=1}^{n} w_{rs} \cdot \ln P_s + \beta_1 \cdot \ln Y_r + \beta_2 \cdot \ln H_r + \sum_{k=1}^{m} \gamma_k \cdot X_{kr} + \varepsilon_r \,. \tag{25}$$

Equation (25) takes the structure of a mixed regressive spatial autoregressive model (Anselin, 1988a, pp 34; LeSage, 1999, pp. 45). The spatial autoregressive parameter ρ reflects the impact of the distance-weighted price level in nearby regions on region r's price level.

If nuisance causes spatially autocorrelated errors, we will employ the price level model (18) with a spatially autocorrelated error process

$$\varepsilon_r = \sum_{s=1}^{n} \lambda \cdot w_{rs} \cdot \varepsilon_s + \nu_r \,, \tag{26}$$

where ν is an independently identically distributed random variable. In case of normally distributed disturbances, consistent and efficient estimators of the regression coefficients are obtained in both spatial dependence models by maximum likelihood (Anselin, 1992).

4. Data

Regional price level models are estimated and tested for a southern German sample of districts. The data on

- the consumer price index,

and three sub-price indices on

- tradable goods, non-tradable goods and housing are available from a comparison of real purchasing power in 2002 across 21 Bavarian municipalities.[13] The municipalities are divided in three size groups (Table 2). While G and K communities are urban districts (*kreisfreie Städte*), A communities belong to rural districts (*Landkreise*). Because of data availability and comparability with regard to explanatory variables, we relate our data analysis on NUTS3 regions, which means that price indices of A municipalities are considered as representative for the price level in the corresponding rural district.

The number of price representatives the consumer price index is based on rises with the size group. The comparison covers price representatives from the products categories housing, food, furniture and home appliances, energy, traffic, leisure, entertainment, culture, other goods and services.

Table 2: Size classes of municipalities

Size class	Municipality	Number of price representatives
A (small municipality)	Deggendorf, Neuburg a. d. Donau, Lohr a. Main, Bad Reichenhall, Neustadt b. Coburg, Cham, Dinkelsbühl	109
G (medium-sized municipality)	Regensburg, Würzburg, Bamberg, Bayreuth, Schweinfurt, Landshut, Passau, Weiden i. d. Opf., Ansbach, Rosenheim, Lindau	204
K (large municipality)	Munich, Nuremberg, Augsburg	646

Source: BSMWVT (2003).

We use disposable income of households as an indicator of regional income. District data for this variable are available from the "National Accounts from the states" ("Volkswirtschaftliche Gesamtrechnung der Länder") provided by the Statistical State Office Baden Württemberg. Housing stock is measured by the number of dwellings in residential and non-residential buildings. On the district level, data on qualifications are available for the employees subject to the social security system. Using this data, an indicator of human capital is obtained by the proportion of employees with a university degree or a degree at an advanced technical college. Data on dwellings and employees is from the CD "Statistik regional" published by the Federal Statistical Office Germany. Geographic coordinates, longitudes and latitudes, of southern German municipalities are compiled from the web catalogue and information portal Informationsarchiv.com.[14]

5. Exploratory spatial data analysis of regional price indices

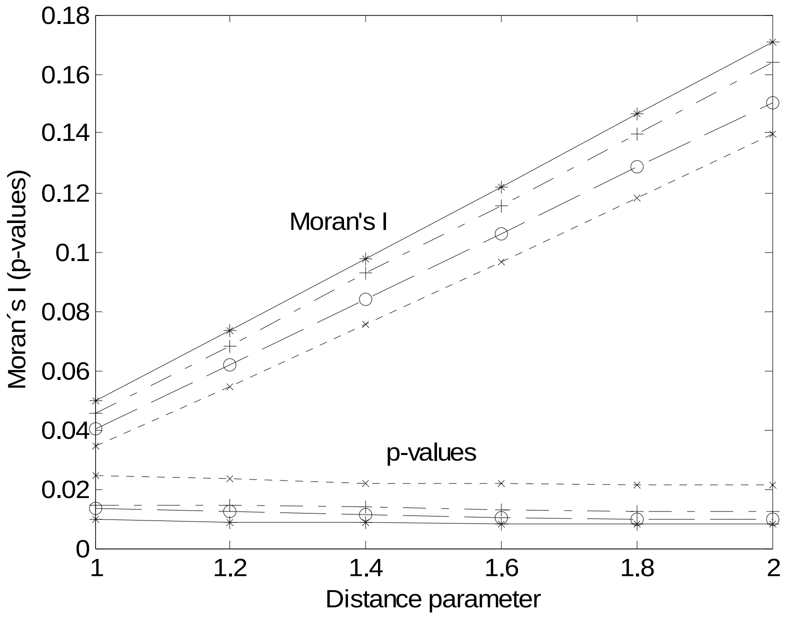
In order to discover spatial price dependence and local nonstationarity across southern German regions, tools of spatial exploratory data analysis (ESDA) are employed to price indices. Global spatial price autocorrelation across Bavarian regions is explored using Moran's I. Figure 1 displays the Moran coefficients for the range [1; 2] of the decay

[13] BSMWVT (2003). The comparison is an updating of a study of the GFK Group, Nuremberg, on behalf of the Bavarian Ministry of Economic Affairs, Transport and Technology.

[14] URL: http://www.informationsarchiv.com/regionalseiten/.

parameter b along with the empirical significance levels (p-values) for the consumer price index (P) and for the three sub-price indices for tradables (P_T), non-tradables (P_{NT}) and housing (P_H). For all four indices Moran's I grows virtually linearly with larger distance decay. The Moran coefficient of the overall price level rises from 0.05 with inverse distance weights to 0.17 with inverse quadratic distance weights. Spatial autocorrelation is somewhat lower pronounced for the sub-price indices. This particularly holds for P_T and P_{NT}, where I rises from 0.04 (b = 1) to 0.13 (b = 2) and 0.034 (b = 1) to 0.10 (b = 2), respectively. Despite the varying strength of spatial association, the p-values show only a relative small decrease over the range $1 \le b \le 2$. For all parameter values b, the Moran coefficients are highly significant ($p < 0.01$) for P and P_M and significant ($p < 0.05$) for P_T and P_{NT}.[15] Thus, price levels do not appear to vary purely randomly across southern German regions. Areas with a relatively high (low) price level tend to be located near areas with price levels above average more than by chance. According to the Moran tests, spatial dependence of regional price indices in Bavaria can be inferred from a broad range of distance-based spatial weight matrices.

Figure 1: Global Moran tests for regional price indices



Notes:
*: Consumer price index, o: Price index of tradables, x: Price index of non-tradables, +: Price index of housing
Moran test [E(I) = -0.05]: Permutation approach: 10000 permutations

An increase of global Moran's I with growing distance decay can certainly be expected as more weight is given to nearby regions. For country income, Aten and Heston (2003) established a rise of Moran's I from 0.35 to 0.73 with replacing inverse distance weights by inverse squared distance weights. Although not compulsory, the latter specification is more "natural" when the gravity model is adopted (Isard et al., 1998, pp. 243; Sen and Smith, 1995). Moreover, in a comparative study Aten and Heston (2003) have shown that the quadratic distance specification matches much better with the contiguity approach over a

[15] Significance tests are based on the permutation approach (Anselin, 1995). For large n Moran's I is asymptotically normally distributed (Cliff and Ord, 1981). Because of the medium-sized sample southern German sample (n=21), the normal approximation would be doubtful.

broad band of distances. Using contiguity matrices with neighbouring regions lying inside a circle of radius of 100 up to 1000 miles, the Moran coefficient decreases from 0.89 to 0.74. This supports the view of Gimpel and Schuknecht (2003) that the squared distance function seems to capture the occurrence of impedance more realistically than the simple inverse distance function. In the following, we therefore confine spatial analysis to the use of a squared distance-based spatial weight matrix.

Although most areas are in line with the global tendency of positive spatial autocorrelation, a number of regions deviate from that pattern with each of the four price indices. They form pockets of instationarity as their own and surrounding price level depart from one another. Such pockets of instationarity are particular the regions of Nuremberg, Würzburg, Bamberg and Neuburg-Schrobenhausen, each of which lying at least twice in the group of areas with the three most negative I_r's (see Table 3). Other Bavarian cities share the positive spatial association with the bulk of the data, but to a much greater extent. The positive local autocorrelation of price level is for the city of Rosenheim so strongly pronounced that it clearly turns out to be an outlier according to the two-sigma rule.[16] The Bavarian capital of Munich, which appears for all price indices in the group of regions with the three largest positive I_r's, proves to be an outlier with respect to P_T. Both cities along with that of Bad Reichenhall render the largest contribution to global spatial price autocorrelation across Bavarian regions.

Table 3: Local Moran coefficients for regional price indices

Local Moran coefficients	Consumer price index (P)	Price index of tradables (P_T)	Price index of non-tradables (P_{NT})	Price index of housing (P_H)
# of positive I_r's	17	16	15	16
Largest positive I_r	I_{RO} = 1.239 (O)	I_{RO} = 0.982 (O)	I_{RO} = 1.172 (O)	I_{RO} = 1.354 (O)
	I_{MU} = 0.707	I_{MU} = 0.919 (O)	I_{BR} = 0.734	I_{BR} = 0.824
	I_{BR} = 0.683	I_{AU} = 0.138	I_{MU} = 0.483	I_{MU} = 0.719
# of negative I_r's	4	5	6	5
Most negative I_r	I_{NU} = -0.226	I_{RE} = -0.134	I_{NU} = -0.362	I_{NU} = -0.296
	I_{WU} = -0.162	I_{BA} = -0.112	I_{WU} = -0.200	I_{WU} = 0.227
	I_{NS} = -0.063	I_{NS} = -0.109	I_{19} = -0.035	I_{BA} = -0.074

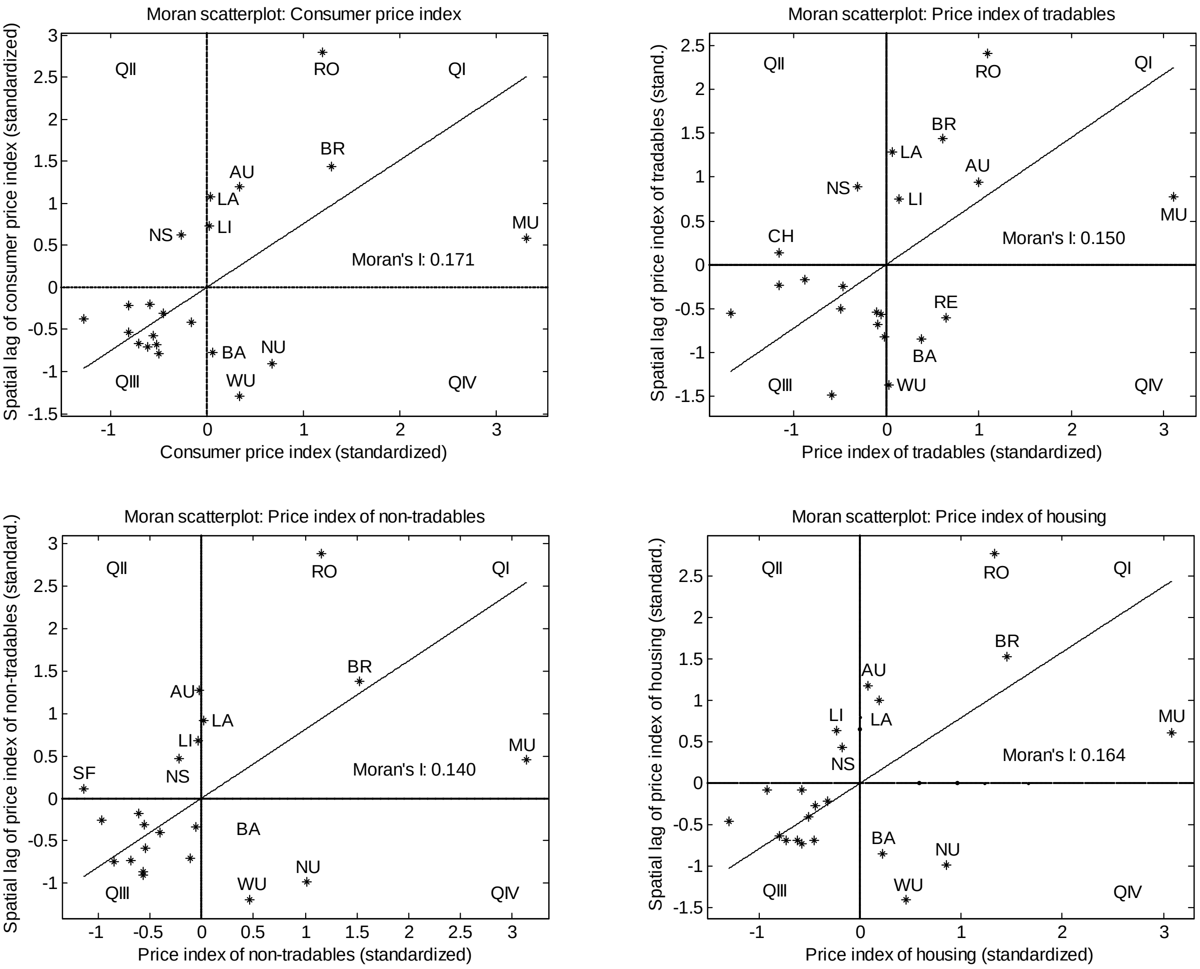
Notes:
MU: Munich, BR: Bad Reichenhall, RO: Rosenheim, NU: Nuremberg, AU: Augsburg, WU: Würzburg, BA: Bamberg, RE: Regensburg, NS: Neuburg-Schrobenhausen, SF: Schweinfurt
O: Outlier according to the two-sigma rule

More insight into the type of local spatial price dependence can be obtained with the aid of Moran scatterplots (Figure 2). In accordance with the local Moran coefficients, most regions are located in quadrants I and III where areal units with positive local spatial autocorrelation are located. The three cities with the three largest I_r's, Rosenheim, Munich and Bad Reichenhall, lie in quadrant I, which is characterized by regions with their own and surrounding high price level. Although its normed residual is not conspicuous, Munich proves to be a high leverage and influential data point for all price indices. Munich's Cooks's

[16] Note that the local Moran coefficients are not bounded to the interval [-1; 1].

distance (CD) exceeds thrice the threshold of one. In contrast to Munich, Rosenheim is already identified to be an outlier by its residual values. Even though it is not a high leverage point, Rosenheim marks an influential data point as its CD value exceeds that of the next ranked region by factors between 3 ½ and 6 ½.

Figure 2: Moran scatterplots of price indices



Notes:
MU: Munich, BR: Bad Reichenhall, RO: Rosenheim, NU: Nuremberg, AU: Augsburg, WU: Würzburg, BA: Bamberg, RE: Regensburg, NS: Neuburg-Schrobenhausen, SF: Schweinfurt, LA: Landau, LI: Lindau, CH: Cham

While observations in quadrant I consist of medium-sized and larger cities, most data points of quadrant III represent rural areas characterized by the lowness of their own and surrounding price level. Particularly the city of Regensburg, however, departs from this classification. The cities of Nuremberg, Würzburg and Bamberg turn out to form pockets of instationarity with their own high price level but surrounded by regions with low price levels (quadrant IV). Areas of spatial instationarity are also represented by the districts of Neuburg-Schrobenhausen, Cham and Lindau, with an unexpectedly low price level of their own, while that of nearby regions is above average (quadrant II).

6. Spatial price level models: estimation and testing results

Exploratory spatial data analysis has highlighted both spatial dependence and spatial heterogeneity of price levels across southern German regions. Thus, we have to account for spatial effects in explaining regional (sub-)price indices of southern districts by NEG-based price level models. For all price indices, model specification and estimation is accomplished in two steps. In the first step, we only allow for spatial heterogeneity by adopting a trend surface and/or spatial expansion model. By means of Lagrange multiplier (LM) tests for spatial dependence we will establish whether spatial autocorrelation is still present in a price level model after accounting for spatial heterogeneity. In case of rejection of the null of independence, the kind of spatial dependence, spatial error or lag dependence, is exposed by the LM statistics.[17] In the second step, the final spatial interaction model for the respective (sub-)price index is estimated and evaluated.

First of all, some general issues should be addressed. According to the price equations inferred from the Helpman model income and housing stock are viewed to be the main influence factors for regional sub-price indices. Although housing stock formally cancels out when combining the sub-price indices of tradables and non-tradables according to the Cobb-Douglas form (2), we use both explanatory variables in the spatial model of the consumer price index. This allows us to test the relevance of a centrifugal force in overall price level determination. The price effects of both variables are measured after controlling for human capital and geography. In explaining P and P_T, we found distance-weighted income to be a more powerful proxy for purchasing power than the income of a region alone.

As to geography, all regressions point to purely random effects of longitude, while latitude always appears to be significant. This finding is in line with that of international studies, where latitude is interpreted as a proxy for a broad set of geographic variables (Aten, 2001; Aten and Heston, 2005). For Western Germany a north-south gradient is well-known with regard to unemployment, while a south-north trend holds for income per capita and employment. In the space of Bavaria, northern administrative districts of Upper Palatinate and Upper Franconia are known to experience noticeably higher unemployment than areas in Upper Bavaria. Thus, a clear-cut south-north price gradient seems to reflect differences in economic performance.

No evidence for a substantial interaction effect between income and latitude on the overall price level is revealed. This suggests that spatial heterogeneity in consumer price index will be captured by a trend surface. Table 4 exhibits the estimates of the extended trend surface model for the consumer price index. About 76 per cent of regional variation in the overall price level can be explained by the above discussed economic factors along with a north-south trend. As expected, regional price level rises with growing income, while dwelling capacity and latitude act in the opposite direction. Human capital seems to exert at best a weak positive influence on regional prices. When ignoring spatial dependence, the significance of that explanatory variable fails to be proved. As both robust LM statistics for spatial dependence are significant, but the exceeding probability with LM(lag) is lower than that of LM(error), we have most notably been concerned with spatial lag dependence (Anselin, 1992; Anselin et al., 1996). There is no evidence for heteroscedastic and non-normal distributed errors.

[17] See Anselin and Florax (1995). In OLS regressions we draw inferences on robust LM spatial lag and error tests because of their higher power to discriminate between both kinds of spatial dependence (Bera/Yoon, 1993; Anselin et al., 1996). With spatial lag or spatial error models only non-robust Lagrange multiplier (LM) or Likelihood ration (LR) tests for spatial dependence are available (see Anselin, 1988a, pp.65 and pp. 103).

Table 4: Estimation und testing results: spatial models for consumer price index

	Extended trend surface model (OLS)		Spatial lag model with trend surface (ML)	
	Coefficient	p-value	Coefficient	p-value
Constant	6.363	0.000	3.725	0.022
Spatial price lag			0.463	0.082
Income	0.039	0.016	0.041	0.000
Dwelling capacity	-0.080	0.002	-0.070	0.042
Human capital	0.036	0.142	0.040	0.072
Latitude	-0.043	0.004	-0.029	0.029
	Goodness of fit and diagnostics			
R^2	0.761		0.770	
SER	0.0022		0.0015	
JB	0.183 (0.913)			
BP / Spatial BP	6.107 (0.191)		6.130 (0.190)	
LM(lag) / LR(lag)	5.521 (0.019)		1.648 (0.199)	
LM(error)	4.273 (0.039)		2.435 (0.119)	

Notes:
R^2: Coefficient of determination, SER: Standard error of regression, JB: Jarque-Bera test, BP: Breusch-Pagan LM test for heteroscedasticity (for the spatial variant see Anselin, 1988b), LM(lag): Robust LM spatial lag test, LR(lag): Likelihood ratio test for spatial lag model, LM(error): (Robust) LM spatial error test

Diagnostics for spatial dependence point to a spatial lag model as the final empirical model for the overall price index. Within this model spatial heterogeneity of prices is captured by a trend surface. Because the Jarque-Bera test does not question the assumption of normal distributed errors, efficient estimation can be accomplished by the method of maximum likelihood (ML). Although the degree of overall explanation has hardly been improved, the influences of some explanatory variables are more clearly exposed. Particularly the strong effect of income on regional prices is highlighted. In addition, human capital seems to exert at least a weakly significant effect on regional price level. The negative influence longitude is clearly confirmed. Moreover, spatially lagged prices seem to capture remaining spatial effects after controlling for spatial heterogeneity. No indication of additional spatial error dependence is revealed. Significant negative influence of housing stock on overall price level points to some weakness of implementing the non-tradable goods sector in Helpman’s NEG model. Scarcity of housing stock in centres may increase overall cost of living thereby potentially establishing a strong centrifugal force.

Model specification and estimation for the sub-price indices P_T, P_{NT} and P_H is carried out in the same manner as for the consumer price index. For the sake of space saving, however, we only report detailed estimation results of the final spatial models (Table 5). Note that the spatial models for the sub-price indices for non-tradables and housing are of the same structure. The spatial-econometric models for the consumer price index and the sub-price index of tradables differ structurally with respect to the interaction of income and geography.

Table 5: Estimation und testing results: spatial models for P_T, P_{NT} and P_M

	Spatial Lag model with spatial trend and expansion (ML) for P_T		Extended trend surface (OLS) for P_{NT}		Extended trend surface (OLS) for P_H	
	Coefficient	p-value	Coefficient	p-value	Coefficient	p-value
Constant	-2.857	0.356	6.692	0.000	7.257	0.000
Spatial price lag	0.499	0.044				
Income	3.003	0.024	0.063	0.001	0.065	0.009
Income x latitude	-0.061	0.026				
Dwelling Capacity	-0.039	0.122	-0.115	0.015	-0.173	0.010
Human capital	0.065	0.000				
Latitude	0.106	0.071	-0.054	0.023	-0.069	0.013
	Goodness of fit and diagnostics					
R^2	0.818		0.693		0.667	
SSE	0.0007		0.0047		0.0080	
Jarque-Bera			0.297 (0.862)		0.976 (0.614)	
Spatial BP / BP	4.392 (0.494)		2.771 (0.428)		2.997 (0.392)	
LR(lag) / LM(lag)	2.417 (0.120)		2.306 (0.129)		3.097 (0.078)	
LM (error)	2.288 (0.130)		2.072 (0.150)		2.524 (0.112)	

Notes:
R^2: Coefficient of determination, SER: Standard error of regression, JB: Jarque-Bera test, BP: Breusch-Pagan LM test for heteroscedasticity (for the spatial variant see Anselin, 1988b), LM(lag): Robust LM spatial lag test, LR(lag): Likelihood ratio test for spatial lag model, LM(error): (Robust) LM spatial error test

In contrast to overall price level, a spatial lag model with a trend surface does not capture spatial effects with P_T adequately. The LM(error) test still indicates the presence of spatial error dependence. However, because of the high p-values of the regression coefficients for income (p=0.18) and dwelling capacity (p=0.11) one may doubt that the significance of the LM(error) statistic can actually be ascribed to disregarding a spatial error process. As ignoring of spatial error dependence will not entail unbiasedness (Anselin, 1988a), the defect may be due to a more complex form of spatial nonstationarity. Indeed, it turns out that the rejection of the null of spatially independent errors can be attributed to 'hidden' spatial heterogeneity that is due to an interaction between income and geography. If we expand income over geographic space a clear model structure emerges. Income exerts a significant impact on prices of tradables, but the effects abate as one goes north. The influence of the interaction between income and latitude clearly outperforms latitude itself, which operates much more weakly in the opposite direction. Human capital is again highly significant.

We also observe a decrease of the empirical significance level of the spatial price lag when passing from P to P_T. This effect can be expected if spatial interaction is mediated by trade. For this interpretation to hold, lag dependence has to be weak or absent with sub-price indices for non-tradables and housing. This is what is evidenced by the testing results. No further spatial effects occur in the space of extended trend surface models for both sub-price indices. The weak significance of the LM(lag) statistic in the price level model for P_H has no consequences as a spatial price lags prove to be insignificant (p = 0.26). For the most part, spatial dependence with P_{NT} and P_H seems to be attributed to price trends across space.

According to theory, housing stock exerts a strong negative influence on P_{NT} and P_H. Its effect on P_T turns out to be negative, too, albeit not significant. This finding again challenges the price formation mechanism of the Helpman model. However, the results for the sub-price indices match well with the finding for overall regional price level. The lack of influence of human capital on both sub-price indices may perhaps be due to its stronger demand linkages to high tech goods. However, the role of human capital in determination of price level seems to be not at all clear (see Heston and Aten, 2005).

Although we have allowed for spatial effects in all price level regressions, the estimates may suffer from a simultaneity bias due to endogeneity of income. We have met this problem by treating income as endogenous and employing the method of instrument variables (IV-2SLS). Several regressions are run using contemporaneous exogenous variables and lagged income as instruments. By and large, the IV estimators look very alike independently of income treated as exogenous or endogenous. On the other hand, differences between IV and ML estimation turn out to be large. Most conspicuously, estimators of the spatial autoregressive parameter, ρ, are greater than 1 for all spatial models. Thus, a possible simultaneity bias seems to be minor compared to the loss of accuracy of IV-2SLS estimation instead of ML estimation (Das et al., 2003).

7. Conclusion

In international income comparisons, purchasing power (PPP) adjusted measures are generally based on real income. Urban and regional economics theory shows that the law of one price is not expected to hold across regions due to, for example, transport costs, land rents and commuting costs (see e.g. McCann, 2001). Because of a lack of area-wide price data, regional policy and empirical research ordinarily have to rely on nominal income. As high income can be partially or even completely compensated for by high costs of living, regional standard of living as well as catching-up of poorer regions cannot be reliably appraised by nominal indicators.

Therefore, Aten and Heston (2005) aimed at estimating regional price levels from international comparisons of income and prices. Since regional price level estimates are based on national price level models, a lot of issues as to the theoretical foundation as well as estimation properties arise. According to the international orientation of the studies, the recently developed price level models are not drawn to regional economic theory. Moreover, a calibration of the models with national data does not a priori ensure an adequate regional match.

The present paper takes up these issues by investigating the contribution of NEG theory in explaining observed regional price indices. The spatial-econometric price level models are based on price equations derived from the Helpman model. Spatial effects of regional (sub-) price indices are revealed with the aid of spatial exploratory data analysis (ESDA) and spatial LM tests. While spatial dependence and spatial heterogeneity matter for explaining the overall price level and the sub-price index for tradables, spatial interaction in the sub-price indices of non-tradables and housing is captured quite well by trend surface models.

According to Helpman's NEG model, the latter sub-price indices are strongly linked with income and housing. After controlling for human capital, overall regional price level and price index of tradables are as well to a great deal determined by income. In the latter regional price level models, spatial price dependence does not vanish when controlling for spatial heterogeneity. Such behaviour is expected if spatial interaction is mediated by trade. In contrast to the prediction of the Helpman model, housing stock seems to be not neutral with respect to overall price level. Significant negative influence of housing stock on overall price level points to some weakness of modelling the non-tradable goods sector. Cost of living may in fact be negatively affected by housing scarcity in centres. Thus, the strength of the centripetal force represented by housing seems to be underrated in the Helpman model.

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