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Giordano Mion

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# Spatial Externalities and Empirical Analysis: The Case of Italy*

Giordano Mion†

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**Abstract**

This paper aims at assessing the role of market linkages in shaping the spatial distribution of earnings. Using a space-time panel data on Italian provinces, I structurally estimate a NEG model in order to both test the coherence of theory with data, as well as to give a measure of the extent of spatial externalities. Particular attention has been paid to those endogeneity issues that arise when dealing with both structural models and spatial data. Results suggest that final demand linkages influence the location of economic activities and that their spread over space is, contrary to previous findings, not negligible.

**Keywords**: Economic Geography, Spatial Externalities, Market Potential.
**JEL Codes**: F12, R12, R32.

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†Giordano Mion: CORE (Université Catholique de Louvain-la-Neuve, Belgium), and Dipartimento di Scienze Economiche (University of Bari, Italy). Address: CORE-UCL, Voie du Roman Pays 34, 1348 LLN, Belgium. Email: mion@core.ucl.ac.be.

# 1 Introduction

Economic activities are certainly not equally distributed across space. However, despite some interesting early contributions made by Hirschman, Perroux or Myrdal, this issue remained largely unaddressed by mainstream economic theory for a long while. As argued by Krugman [26], this is probably because economists lacked a model embracing both increasing returns and imperfect competition in a general equilibrium framework.

The relatively recent new economic geography literature (NEG) has finally provided a collection of general equilibrium models explicitly dealing with space, and able to account for many salient features of the economic landscape.[1] As Krugman [26] pointed out, there is a strong connection between the NEG and some older fields in economics. To a large extent, what has been done is in fact rediscovering concepts and ideas that did not receive much attention by mainstream economic theory because of their lack of a rigorous formal counterpart. Within this group of overlooked contributions, and of particular interest for this paper, is the literature on *market potential* starting with Harris [19]. This strand of literature argues that a location's attractiveness for firms depends on its access to markets. The quality of this access is often measured by an index of market potential, which is a weighted sum of the purchasing power of all other locations, with weights inversely related to distance. Although this approach has proved to be empirically quite powerful, it totally lacked any microeconomic foundation. At that time, there were in fact no rigorous explanations of why a correlation between market access and firms' location should exists. However, Fujita, Krugman and Venables [15] show that market potential functions can be obtained from spatial general-equilibrium models, thus providing the theoretical background for the use of such an approach.

The main objective of this paper is thus to estimate a market potential function, derived from a NEG model, using data for Italian provinces. The particular framework used is a multi-location extension of Helpman [22] two-location model, originally introduced and estimated by Hanson [17] for the US counties, in order to:

1. Obtain estimates of structural parameters to infer the consistency of Helpman's model with reality.

2. Evaluate the theory-based market potential function in the light of the empirical literature on market potential, in order to investigate the specific contribution of the model in understanding firms' location.

3. Give an idea of the extent of spatial externalities by measuring how far in space a shock in one location affect the others.

I depart from the the existing literature, and in particular from Hanson [17], in several ways. First, a rigorous estimation technique, derived from Spatial Econometrics and Dynamic Panel Data, have been implemented in order to tackle some unaddressed endogeneity issues that naturally arise when dealing with

[1] See Fujita and Thisse [14], Ottaviano and Puga [28], and Fujita, Krugman and Venables [15] for a review of the literature.

structural spatial models. Second, I introduce a new measure of equilibrium local wages that is needed in order to account for those structural differences, like labor mobility, that make international comparisons of agglomeration forces problematic. Finally, I make use of several distance decay functions in order to investigate the sensitivity of my results to the particular assumption made about transportation technology. Interestingly, in my preferred specification, the results indicate that the spatial scope of agglomeration externalities is larger than what emerges from former studies.

This paper contributes to the growing empirical literature on the location of economic activities. There are, however, different approaches in this research field, each relying on a different agglomeration mechanism.[2] Agents may in fact be drawn to regions with pleasant weather or other exogenous amenities.[3] However, both human capital accumulation stories[4] and localized spillovers, like Marshall or Jacobs externalities, may also contribute to geographic concentration.[5] By contrast, here I stress increasing returns and market interactions, as opposed to factor endowments (exogenous amenities) and technological externalities (human capital and technological spillovers), taking the NEG framework as the theoretical basis for my investigations. Other examples of this market-linkages approach can be found in Combes and Lafourcade [8], Head and Mayer [21], and Teixeira [37].

The rest of the paper is organized as follows. In Section 2 I give some insight on the mechanics of Helpman [22] model, and I present the structural equation that I will estimate. Section 3 deals with data issues, while in Section 4 I discuss econometric concerns. Detailed estimation results are presented in Section 5. Finally, in Section 6 I draw some conclusions and suggest directions for further research.

## 2 NEG and Market Potential

The NEG literature offers the possibility to treat agglomeration in a flexible and rigorous way by means of increasing returns (IRS), imperfect competition, and product differentiation. In this Section, I am particularly concerned with Helpman's [22] model, which will be the theoretical ground on which I will construct the econometric analysis.[6]

Helpman's [22] model is actually a two-good, two-factor, two-region model that closely resembles the well-known core-periphery model by Krugman [25]. In both cases, there is an IRS manufacturing sector, producing a differentiated product under monopolistic competition, where the only input is an inter-regionally mobile workforce. Workers/consumers migrate from one region to another according to differences in real wages, while firms look for high profitable locations. However, while in Krugman [25] the other good is homogenous, freely tradable and produced under constant returns to scale (CRS) by a sector specific immobile labor force (farmers), in Helpman [22] it is instead a non tradable good (like housing services) that is produced with an exogenously distributed sector specific capital under CRS. As for the

[2] See Hanson [18] for a survey of the empirical literature on agglomeration economies.
[3] See for example Rosen [33], and Roback [32].
[4] See Lucas [27], and Black and Henderson [6].
[5] As for the impact of localized externalities on productivity and growth see Henderson, Kuncoro, and Turner [24], and Ciccone and Hall [7].
[6] For a detailed exposition of the model, see Helpman [22] and Hanson [17].

distribution of capital ownership, in Helpman [22] this is supposed to be public, i.e. each individual mobile worker/consumer owns an equal share of the total capital/housing stock $\bar{H}$. Equilibrium real wages are equalized across regions unless some areas become empty. Contrary to Krugman [25], this is, however, a very unlikely outcome because it implies that in abandoned regions the price of housing is zero. Therefore, locations where manufacturing activities agglomerate are characterized by high housing prices, and this act as a dispersion force against the tendency for firms to concentrate close to big markets (the so-called market access effect).

Depending on the level of transportation costs ($f$), elasticity of substitution among varieties ($\sigma$) and the share of traded goods in consumers' expenditure ($\mu$), manufacturing activities will be dispersed or agglomerated. In the second case firms will be disproportionately distributed with respect to a region size. In particular, indicating with ($H_i$) the stock of housing of region $i$, those locations with an above (below) average endowment will have a more (less) than proportional share of manufacturing in equilibrium. Both a higher share of tradable goods ($\mu$) or a lower elasticity ($\sigma$) induce more agglomeration. In the case of $\mu$, this is due to the fact that concentration of firms and consumers in the same place allow to avoid transportation costs thus increasing real consumption. The greater is the share of these goods in the consumption of migrating workers, the stronger is this centripetal force. The role of $\sigma$ is instead to counterbalance the usual centrifugal force that works against concentration: price competition. A lower elasticity of substitution $\sigma$ makes in fact varieties more differentiated, relaxing local competition among sellers.

There are basically two reason for which I prefer to use Helpman instead of Krugman model for my empirical investigation of Italian provinces. First of all, Helpman's model seems to be more suitable to describe the kind of forces at work at low-level spatial scale, where congestion costs and the price of land are key localization factors for both firms and consumers. In particular, the fact that in Krugman [25] equilibrium nominal wages are lower in regions where agglomeration takes place is particularly disturbing. Moreover, from an empirical point of view, Helpman [22] is also preferable because of the less extreme nature of its equilibria. Although the production of manufactured tradables is certainly highly agglomerated in space, the full concentration in very few places, that is quite a standard outcome in Krugman [25], is far too extreme.

In order to give a useful interpretation of the kind of investigations I want to deal with, as well as to link them to previous studies, one has to come back to Harris's [19] market-potential concept. Actually, Harris's [19] market-potential relates the potential demand for goods and services produced in a location $i = 1, 2, .., \Phi$ to that location's proximity to consumer's markets, or:

$$MP_i = \sum_{k=1}^{\Phi} Y_k g(d_{ik}) \tag{1}$$

where $MP_i$ is the market potential of location $i$, $Y_k$ is an index of purchasing capacity of location $k$ (usually income), $d_{ik}$is the distance between two generic locations $i$ and $k$ and $g()$ is a decreasing function. The higher is the market potential index of a location, the higher is its attraction power on production activities.

In Helpman's model, a good measure of the attractiveness of location $i$ is given by the equilibrium nominal wages $w_i$. Although firms makes no profits in equilibrium (no matter where they are located), the wage they can afford expresses their capacity to create value once located in a particular region. In fact, if centripetal forces take over, those locations that attract more firms and consumers will also have higher equilibrium nominal wages, thus leading to a positive correlation between agents' concentration and $w_i$. Following Hanson [17], one can combine some equilibrium equations and apply logarithms to simplify things in order to get the following expression for $\ln(w_i)$:

$$\ln(w_i) = \kappa_3 + \sigma^{-1} \ln \left[ \sum_{k=1}^{\Phi} Y_k^{\frac{1-\sigma(1-\mu)}{\mu}} H_k^{\frac{(1-\mu)(\sigma-1)}{\mu}} w_k^{\frac{(\sigma-1)}{\mu}} f\left(d_{i,k}\right)^{(\sigma-1)} \right] \tag{2}$$

where $\kappa_3$ is a function of behavioral parameters ($\mu$, $\sigma$), and $f()$ is, for the moment, a generic decreasing function of distance that I will parametrize explicitly afterwards.

Equation (2) really looks like a market-potential function. It tells us that as long as agglomeration forces are active ($\sigma(1-\mu) < 1$), the nominal wage in location $i$ (and thus local firms' profitability) is an increasing function of the weighted purchasing power coming from surrounding locations ($Y_k$), with weights inversely related to distances $d_{ik}$ trough the transport technology function $f(.)$ (this is the market access component). Crucially, (2) tells us more than the simple market potential equation (1). The distribution of economic activities should in fact depend upon prices, because an increase in other locations' housing stock ($H_k$) or wages($w_k$), causes $w_i$ to increase in the long-run in order to compensate workers for lower housing prices and higher earnings they can enjoy elsewhere. Although quite powerful from an empirical point of view, relations like (1) were not obtained from a theoretical model and, compared to (2), did not control for wages and prices of others locations.

## 3 From Theory to Econometrics: Data Issues

One of the most common problems in using micro-founded economic models for empirical purposes is the choice of good proxies. Estimation requires actual data, and in some circumstances the choice of the statistic that is best suited to approximate a theoretical variable becomes a difficult task. As for the case of equation (2), the variables $H$, $Y$, and $d$ do not raise particular interpretation problems. $H$ is meant to represent those goods and factors that are immobile for consumption or production. Expenditure on housing services actually constitutes a large part of these costs. A good proxy is thus given by the total size of houses available (for both for family and commercial use) in a region measured in square meters. The variable $Y$ should instead represent the demand for goods, and a reasonable solution is to take total household disposable income as a measure of province's purchasing power. Finally, $d$ is the distance between two generic locations. The unavailability of more sophisticated measures of distance has led us to use a physical metric. In particular I adopt the crow flight distance between the centers of each province (as obtain by polygonal approximation) using GIS software.

However, when one thinks about $w$ some complication arise. One natural solution, followed by Hanson

[17], is to consider it as just labor income, thus using county statistics on average earnings of wage and salary workers. Although this solution may be to some extent acceptable for the US, it seems difficult to argue the same for Europe and in particular for Italy. First, it is a wide spread opinion that in Europe conditions of local supply and demand play little role in the determination of wages[7], thus making them unsuited to express re-location incentives. In some countries, and this is the case for Italy, wages are in fact set at the national level for many production sectors. Second, the relatively scarce mobility of people prevents the price system to clear labor markets excess-supply.[8] Agglomeration externalities are thus likely to magnify regional imbalances in both income and unemployment rates rather than shifting massively production activities.

In line with these considerations, US economic activities are more spatially concentrated than in Europe. The 27 EU regions with highest manufacturing employment density account for nearly one half of manufacturing employment in the Union and for 17% of the Union's total surface and 45% of its population. The 14 US States with highest manufacturing employment density also account for nearly one half of the countries manufacturing employment, but with much smaller shares of its total surface (13%) and population (21%). By contrast, in Europe agglomeration is more a matter of income disparities and unemployment. 25% of EU citizens live in so-called Objective 1 regions. These are regions whose Gross Domestic Product per capita is below 75% of the Unions average. By contrast only two US states (Mississippi and West Virginia) have a Gross State Product per capita below 75% of the country's average, and together they account for less than 2% of the US population. Moreover, regional employment imbalances are a special feature of European economic space. The case of Italy is best known, with Campania having a 1996 unemployment rate 4.4 times as high as Valle d'Aosta. However, as shown by Overman and Puga [30], large regional differences exist in all European countries.

These considerations suggest the existence of similar forces shaping the distribution of economic activities in an asymmetric way. The point is that the structural differences between US and EU cause these forces to have a more visible impact on different economic indicators. At this stage, it is probably better to come back to Helpman [22] to look for some guiding insights. In that framework, $w$ is the zero-profit earnings of the only production factor for tradables (labor), and *it turns out to be a measure of the attractiveness of a location for firms.* As long as mobility is limited, the transfer of firms in some locations would produce unemployment in the abandoned ones while pushing the factor market to full employment in the former. However, the fact that basic wages are more or less fixed does not prevent firms to give workers, if they have the means, other form of remunerations in order to attract them. Therefore, one can think to use total labor expenditure per employee as a measure of the shadow wage. However, labor is not the only production factor in real world. In Helpman [22] it stands for the aggregate of mobile factors, as opposed to the immobile ones ($H$), and so mobile capital remunerations should also be included in the construction of a good proxy. Furthermore, profits need also to be accounted for as they are, in principle, precisely those indicators leading firms to relocate. In this light, *it thus seems problematic to associate $w$*

[7] See Bentolila and Dolado [4], and Bentolila and Jimeno [5] for an empirical assessment.

[8] Eichengreen [11] estimates that the elasticity of interregional migration with respect to the ratio of local wages to the national average is 25 times higher in the US than in Britain. The difference with respect to Italy is even larger.

*to wages only, and this criticism apply to the US case too.*

The solution I will adopt tries to address these issues. First, in order to be consistent with Helpman's model, remuneration of immobile capital should not enter in the computation of $w$. The reason is that $w$ should measure the incentive for mobile factors (labor in Helpman [22]) to move towards agglomerated areas. Expenditure in housing services actually represent a large part of those remuneration. Therefore, using statistics on rented house numbers and prices from the Italian National Institute of Statistics (ISTAT), I have constructed a measure of housing spending per province and, after subtracting it from GDP, I have divided by active population (to control for different unemployment rates) to get my $w$.[9] The variable obtained is meant to capture the average remuneration of all mobile factors, as well as profits, and it is only indirectly related to local wages. In Section 5, I will provide further (empirical) evidence to justify my measure of $w$ for the Italian economy.

Table 1 contains summary statistics on $w$, $H$, and $Y$, as well as on provinces' surface and their population. All nominal variables are in 1996 prices and the unit is one million liras. Total housing area $H$ is measured in square meters, while population is in thousands of people and provinces area is expressed in square kilometers. Data are time-averaged and refer to the interval 1991-1998. Statistics on rented-house numbers and prices come from ISTAT. Data on GDP, population, employees, housing stock, and households' disposable income come from SINIT database (Sistema Informativo per gli Investimenti Territoriali). The latter data-set has been collected from the "Dipartimento Politiche di Sviluppo e Coesione - Ministero dell'Economia e Finanze". Finally, distances have been obtained with GIS software and are expressed in kilometers.

# 4 Econometric Specification

## 4.1 Main Concerns

As mentioned in the Introduction, the main goal of this paper is to estimated a structural model and in particular equation (2). However, the data-set I have is a panel covering two dimensions: space and time. Therefore, the actual formulation I use is:

$$\ln(w_{i,t}) = \kappa_3 + \sigma^{-1} \ln \left[ \sum_{k=1}^{\Phi} Y_{k,t}^{\frac{1-\sigma(1-\mu)}{\mu}} H_{k,t}^{\frac{(1-\mu)(\sigma-1)}{\mu}} w_{k,t}^{\frac{(\sigma-1)}{\mu}} f\left(d_{i,k}\right)^{(\sigma-1)} \right] + \varepsilon_{i,t} \tag{3}$$

where indexes $i$ and $t$ corresponds, respectively, to space and time, while $\varepsilon_{i,t}$ is a random term that, for the moment, is just assumed to be serially uncorrelated in the time dimension, that is $Cov(\varepsilon_{i,t}, \varepsilon_{i,s}) = 0, \forall t \neq s$. Later on, I will explicitly test this assumption.

The first choice to make is the geographical reference unit. On one hand, this should not be too large in order to account for the nature of externalities that the model wants to capture. Helpman [22]

[9] Actually, I subtract people looking for their first job from active population before computing $w$. The number of those looking for their first occupation is in fact closely related to factors (like social habitudes), that are both external to the model and vary a lot across Italy, thus introducing a potential source of bias.

is in fact best suited to describe agglomeration forces at low/medium scale spatial level. The tension between easy access to cheap commodities and high costs of non-tradable services like housing is certainly a good metaphor for metropolitan areas, but the more one departs from this example the more other forces are likely to be at work. On the other hand, a too high geographical detail could also misrepresent the tensions at work, as well as to requiring an intractable amount of information. To give an example, if one decides to work on the approximately 8, 100 Italian municipalities, he will need a matrix of distances with $8,100 * (8,100 + 1)/2 = 32,809,050$ free elements to evaluate. My choice is thus a compromise between these two different needs, and consists in taking the 103 Italian provinces as reference units.

Turning to specification issues, I should argue why I choose (2) in order to get the estimates of structural parameters. In principle, this objective would be better achieved using simultaneous equations techniques directly on Helpman [22] equilibrium equations. However, apart from implementation difficulties, it is the unavailability of reliable statistics for manufacturing goods at any interesting geographical level that makes this solution unapplicable. Data on prices can in fact be found at the regional level for Italy. This is probably too much an aggregate unit for my purposes because of the lower inter-regional labor mobility, as well as the limited number of cross-section observations (just 20). Equation (2) is instead a reduced-form, in an algebraic sense, of the model that does not contain these two variables, and for which it is thus possible to find adequate local data.

Another important aspect concerns missing variables like local amenities (nice weather, ports, road hubs, etc.) and localized externalities (especially human capital ones), that possibly influence the distribution of earnings, but are not included in the analysis. As long as these variables are correlated with regressors, and they are indeed likely to be, standard econometric techniques would fail. Anyway, when one thinks about both amenities and externalities it is clear that, if these factors change over time, their variation is very slow. The quality of the work force, as well as the presence of infrastructure and the network of knowledge exchange is thus reasonably constant (for a given location) in a short interval of time. I can thus try to overcome this problem with an appropriate choice of the estimation interval, (that should not be too long) in order to treat them as correlated fixed effects $\mu_i$. The random term would thus become $\varepsilon_{i,t} = \mu_i + u_{i,t}$ and, applying a time-difference on (3), the term $\triangle\varepsilon_{i,t} = \varepsilon_{i,t} - \varepsilon_{i,t-1}$ would then simplify to the time-varying component difference only: $\triangle u_{i,t} = u_{i,t} - u_{i,t-1}$.[10]

To actually implement estimations, one needs to define define distance weights $f(d_{i,k})$ in equation (3). These weights should measure the degree of economic interaction among locations. Actually, Hanson [17] uses the exponential form $f(d_{i,k}) = \exp^{-\tau d_{ik}}$, where $\tau \in [0, \infty)$ is an (inverse) measure of transportation costs to be estimated, and $d_{i,k}$ is distance between $i$ and $k$. However, such a specification gives rise to some odd results. For example, according to Hanson's [17] estimates, travelling two km appears to multiply the price of a good by 50. This quite unrealistic result comes from the fact that an inverse exponential goes to zero very fast (faster than any polynomial function). Therefore, as a robustness check, I will try an alternative specification that is more rooted in trade theory analysis: the power function $f(d_{i,k}) = \theta\, d_{i,k}^{\psi}$.

[10] It is worth noting that localized externalities need a model where production is disaggregated by sector in order to be captured. In fact, Marshall and Jacobs's externalities are usually measure by indexes of (respectively) industrial specialization and diversification. Therefore, it seems problematic to include them in my aggregate model.

Helpman [22] is essentially a trade model, so that a good proxy for economic interaction is given by trade flows. In this respect, the power function has been extensively used in both gravity equation and home bias literature[11]. Comparing results obtained with these two measures, and in particular the associated goodness of fit, will help us to shed some light on the spatial scope of agglomeration externalities.[12] As a further check on the pervasiveness of final-demand linkages, I will also perform some unstructured linear estimations using many distance-band matrices in the same spirit as Rosenthal and Strange [35] and Henderson [23].

A final issue is related to endogeneity. First, the presence on the right hand side of a weighted sum over space of the same variable appearing as independent ($w_{i,t}$), is a source of bias. According to spatial econometrics, this sum can be seen as a spatially-lagged endogenous variable and it is well known that, in this case, the least squares method does not work regardless of error properties[13]. Furthermore, in my structural model, $w_{i,t}$ is determined simultaneously with income $Y_{i,t}$. The circularity between factor earnings and income is certainly not debatable in economic theory, and in my framework implies that the explanatory variable $Y_{i,t}$ is correlated with disturbances $u_{i,t}$. Finally, even if the amount of fixed factors $H_{i,t}$ is supposed to be exogenous in Helpman model, it is not difficult to imagine that, for example, pressures on the housing market do not simply lead to price movements, but also encourage construction of new buildings.

The solution proposed by Hanson [17] to account for endogeneity is to use aggregated spatial variables as regressors on the right hand side of equation (3), and then apply non-linear least squares (NLLS). Following his reasoning, $u_{i,t}$ should in fact reflect temporary shocks that influence local business cycles. The finer the geographical unit I use for locations, the smaller is the impact of such shocks on geographically aggregated variables. Furthermore, if these shocks are really local, their spread to other regions should be negligible. Now, if this is true, then there should not be any significant correlation between the shock $u_{i,t}$ of a small county $i$ and (for example) the state-level values of $w$, $Y$, and $H$. Actually, Hanson [17] uses data on $w$ for the 3075 US counties as dependent variables and, for each county $i$, he uses data on $w$, $Y$, $H$ and distances at the state level as independent variables on the right hand side of (3). Formally speaking, the two indexes $i$ and $k$ do not correspond anymore to the same location set, with index $i = 1, 2..., \Phi$ corresponding to US counties, and $k = 1, 2, ..., \Phi^*$ corresponding to US states.[14]

A few considerations are in order. Hanson's idea sounds like instrumentation. He actually employs state level values on the right-hand side precisely because he needs something that is uncorrelated with the disturbances, but still linked with the (real) explanatory variables at county level. Indeed, these are the features of good instrumental variables. Therefore, as an alternative estimation methodology, one can

[11] See Disdiez and Head [10].

[12] In order to make spatial econometrics techniques directly applicable I will estimate $\theta$, while using for $\psi$ values coming from the literature. The distance weight $f\left(d_{i,k}\right)$ in (3) is raised to the power $\sigma - 1$, and so I am actually interested in $\psi\left(\sigma - 1\right)$. Following Disdiez and Head [10], a reasonable estimate for $\psi\left(\sigma - 1\right)$ is $-1$, so that the distance decay I will use is: $f\left(d_{i,k}\right)^{\sigma-1} = \theta d_{i,k}^{-1}$. Furthermore, as standard in spatial econometrics, I will give a zero weight to observations referring to the same location, i.e. $f\left(d_{i,i}\right)^{\sigma-1} = 0$.

[13] See Anselin [1].

[14] In equation (3) for instance he has, for a given year $t$, a sum of $\Phi^* = 49$ terms (the number of US continental states plus the district of Columbia) on the right hand side, for each of the $\Phi = 3075$ equations to fit.

think of keeping county level variables on the right hand side, and use geographically aggregated data directly as instruments for the estimation. Clearly, as long as Hanson's strategy works, the other should work as well. Furthermore, a non-linear instrumental variable approach (NLIV) would be conceptually preferable because it allows to maintain a homogeneous space unit on both sides of (3). In the spatial econometrics literature, it is in fact well known that the level of aggregation matters a lot.[15] However, there is another aspect in favor of instrumental variables: efficiency. By aggregating explanatory variables, Hanson loses a lot of information, ending with a sum of just 49 terms instead of 3075. By contrast, all the information contained in county data would be preserved with instrumental variables as one can keep a fine geographical level also on the right-hand side. I will pay further attention to these two consideration in the Section devoted to estimation.

There is, anyway, something unclear in the crucial identifying assumption on which the two above described procedures rely. Technically speaking, they amount to assume that time-varying residuals $u_{i,t}$ are uncorrelated among themselves as well as with spatially aggregate values of $w$, $Y$, and $H$. The first assumption is quite clear, and can be tested using spatial econometrics tools like the Moran correlation test[16]. The second is, by contrast, quite obscure and needs to be better clarified. For the shock of county $i$ to be uncorrelated with the state-level values of $w$ (which are nothing else that averages of the corresponding $\Phi$ county values $w_i$), one needs $Cov(u_{i,t}, w_{k,t}) = 0 \; \forall \, i \neq k$. In other words, the local shock does not spread over other locations, resulting in a negligible degree of "spatial interaction". The fact that error terms are not spatially correlated limits the degree of spatial interaction in the sense that $u_{k,t}$ has, for example, no impact on $w_{i,t}$ through $u_{i,t}$ because the latter is uncorrelated with $u_{k,t}$. However, $u_{k,t}$ does have an impact on $w_{i,t}$ through $w_{k,t}$ because the latter figures as an explanatory variable in (3), and $w_{k,t}$ is itself a function of $u_{k,t}$. Therefore, as long as estimates obtained with NLLS or NLIV are significant, the correlation between $u_{k,t}$ and $w_{i,t}$ through $w_{k,t}$ could not be negligible and aggregate variables cannot be used as instruments. Put differently, as long as the theoretical model has something interesting to say about local factor earnings, consumers' expenditure, and non-tradable goods, then the aggregation trick does not work.

In the next Subsection I introduce an alternative estimation strategy, that can potentially be used for many spatial structural models applications, in order to properly address endogeneity.

### 4.2 Escaping the Endogeneity Trap

A possible way-out from this endogeneity trap that, at the same time, would allow us to preserve the same space dimension for all variables, could be to better exploit the information coming from the time dimension, using dynamic panel data à la Arellano and Bond [3]. This basically consists of estimating the model in first differences (in order to get rid of fixed effects), while using past levels of endogenous variables

[15] In order to explore the extent of this possible inhomogeneous data bias, I will perform a comparative estimation using the two techniques: a non-linear least squares Hanson type, and a non-linear instrumental variables one.

[16] See Anselin [1], and Anselin and Kelejian [2].

(starting from $t-2$) as instruments. However, for this procedure to work correctly, time-dynamics should also be accounted for.

NEG models are designed mainly to reply to theoretical rather than empirical purposes. Compared to applied macro-economic models, they are in fact represented by systems of equilibrium equations in which almost all variables are endogenous, making the identification task problematic to solve for a given time $t$ (i.e. using only the cross-section dimension). This is precisely the reason for which a panel approach is preferable. Now, since endogeneity comes from the simultaneous nature of these models linking, in equilibrium, the $\Phi$ economies, one can think of using the weak-exogeneity assumption and applying the appropriate GMM estimator directly to (3). However, such an approach rests on the hypothesis that the simultaneity problem is fully contemporaneous, ruling out any dynamic behavior.[17] In the real world, it is unlikely that data do not exhibit a time dynamics so that the impact of a shock $u_{i,t}$ is entirely exhausted at $t$ without spreading over time. For example, frictions in the factors market, like the presence of unobserved sunk costs for migration or unions' power, would cause variables to adjust in a sluggish way toward their equilibrium level, thus justifying the time persistency of a shock. This is why I prefer to resort to dynamic panel data techniques à la Arellano and Bond [3].

In particular, in order to account for the time dynamics, a time-lagged value of $\ln(w_{i,t})$, as well as a complete set of time dummies, will be added to regressors in the estimation of (3). As long as tests on residuals will not detect a significant time correlation, one can be confident that this solution successfully controls for the time dynamics.[18] Then, following Arellano and Bond's [3] idea, I can apply a first difference and use past levels of endogenous variables, starting from $t-2$, as instruments for the estimation. Although, contrary to the usual panel framework, observations are not independent in the cross-section dimension (and this is a peculiarity of spatial data), convergence is achieved, as showed by Anselin and Kelejian [2], as $\Phi$ goes to infinity if error terms are spatially uncorrelated.

Formally speaking, the set of identifying restrictions on which my procedure relies is:

1. $Cov(u_{i,t}, u_{k,t}) = 0 \,\forall\, i \neq k$
2. $Cov(u_{i,t}, u_{i,s}) = 0 \,\forall\, t \neq s$
3. $E[u_{i,t}|x_{i,s}] = 0 \,\forall\, t > s$

where $i, k = 1, 2..., \Phi$ and $s, t = 1, 2..., T$. The first set of restrictions requires absence of spatial correlation, and can be investigated by means of a Moran test. The second calls for absence of residual time-dynamics. The Arellano and Bond [3] GMM estimator is in fact incompatible with disturbances having an AR

[17] Unreported GMM estimations (based on the weak-exogeneity assumption) on a linearized version of equation (3), support the introduction in the model of some dynamic component. In particular, the Sargan test on over-identifying restriction rejects the validity of instruments and, crucially, the tests on residuals detect a significant time correlation thus suggesting the presence of a misspecified time-dynamics.

[18] A slightly more general formulation would consist in using an error autoregressive process: $u_{i,t} = \delta u_{i,t-1} + v_{i,t}$. I choose the other one mainly for computational reasons. They both amount to put some time-dynamics in the data, with the difference being that the in the second the lags of the original regressors should also be introduced in the estimation with their own (restricted) parameters. These restrictions require the implementation of a non-linear recursive procedure. Furthermore, specification tests did not detect any unaccounted source of endogeneity, suggesting that my choice is a good compromise between computational issues and the need to control for time-dynamics.

structure: the dynamics need to be captured into the model, as I am trying to do by adding a time-lagged value of $\ln(w_{i,t})$, as well as a complete set of time dummies, to (3). Tests on the residuals' time correlation will then allow investigation of the validity of such an assumption. Finally, the third type of condition expresses weak exogeneity and, together with the others, makes past values of endogenous variables good instruments. It is important to stress that, contrary to Hanson's procedure, the validity of instruments can be directly assessed here by means of a Sargan test on over-identifying restrictions. Furthermore, my strategy allows us to keep the same geographical dimension for dependent, explanatory, and instrumental variables, thus avoiding a possible inhomogeneous data bias.

However, non-linearity of equation (3) is a computational challenge. It certainly complicates the implementation of simple panel techniques but, more importantly, could cause estimations to be extremely unstable. As known in applied econometrics, the combination of non-linearity, endogeneity, and instrumentation is a dangerous mix that causes criterion functions to have many local minima. The solution I adopt is then to estimate a linearized version of equation (3). This approach is not new for NEG applied models, and has been pioneered by Combes and Lafourcade [8] with promising results. In an unpublished Appendix, available from the author upon request, I formally derive the following linear counterpart of (3) which is given by:

$$\ln(w_{i,t}) = a + \sum_{k=1}^{\Phi}[(B_1\bar{Y}_{k,t} + B_2\bar{H}_{k,t} + B_3\bar{w}_{k,t})\, d_{i,k}^{-1})] + \varepsilon_{i,t} \tag{4}$$

where $B_1 = \theta\frac{1-\sigma(1-\mu)}{\sigma\mu}$, $B_2 = \theta\frac{(1-\mu)(\sigma-1)}{\sigma\mu}$, $B_3 = \theta\frac{\sigma-1}{\sigma\mu}$, and for example $\bar{Y}_{k,t} = \ln(Y_{k,t})\frac{Y_{k,t}}{\sum_{k=1}^{\Phi} Y_{k,t}}$.

Equation (4) is now linear in the 3 parameters $(B_1, B_2, B_3)$ and, after adding time dummies and a time-lag of $\ln(w_{i,t})$ to control for time-dynamics, one gets the final regression equation:

$$\ln(\mathbf{w}_t) = \mathbf{i}\, dum_t + \ln(\mathbf{w}_{t-1})A + \mathbf{W}\,\bar{\mathbf{Y}}_t B_1 + \mathbf{W}\,\bar{\mathbf{H}}_t B_2 + \mathbf{W}\,\bar{\mathbf{w}}_t B_3 + \boldsymbol{\varepsilon}_t \tag{5}$$

where bold variables are column vectors containing observations for the $\Phi$ locations at time $t$, $\mathbf{W}$ is a $\Phi x \Phi$ spatial weighting matrix with generic element $\mathbf{W}_{i,k} = d_{i,k}^{-1}$, $\mathbf{i}$ is a vector of ones, and $dum_t$ is a time-dummy. Equation (5) will be the one I will use for my dynamic panel investigations. With estimates of $B_1$, $B_2$, and $B_3$ in my hands, I can then trace back the implied values of $\mu$, $\sigma$, and $\theta$ and, using the Delta method, make inferences on the latter.

To account for possible structural differences between continental Italy and the two islands of Sicily and Sardinia, I compute estimates using continental provinces only. Further details about spatial aggregation and instruments are given in the Appendix.

# 5 Estimations

## 5.1 Regressions and Instrumentation

In this Section,I will first use data on the 103 Italian provinces to estimate equation (3) using both Hanson non-linear least squares (NLLS) procedure, and a non-linear instrumental variables (NLIV) one. The two methods consist of cross-sections and rest on the same statistical assumptions, with the second being preferable because it does not mix observations referring to different geographical units. These first regressions will allow us to compare directly results with Hanson [17], as well as to shed some light on the bias coming from space-inhomogeneous observations. The two points in time I consider to make time-difference are 1991 and 1998. For the NLIV estimation I have then used, for each province, the change (over the time interval 1991-1998) in the logarithm of the variables $w$, $Y$, and $H$ of the corresponding 11 NUTS (Nomenclature of Territorial Units for Statistics) level 1 regions as instruments. For NLLS, I have instead used directly the values of $w$, $Y$, and $H$, corresponding to the eleven zones, as regressors before taking first differences and applying least squares.[19]

Subsequently, I will go through my preferred specification, the panel estimation of (5), using Arellano and Bond [3] estimator. At the cost of linearization, this method should allow us to address properly the endogeneity issue.[20] Crucially, a test on the validity of instruments can be actually performed in this framework. The database used in this case will consist of yearly data from the entire period 1991-1998. Panel estimates are two-stage GMM and have been obtained with DPD 98 for Gauss. The model is estimated in first differences, using past levels of $\ln(w_{i,t})$, $\mathbf{W}_i\bar{\mathbf{w}}_t$, $\mathbf{W}_i\bar{\mathbf{Y}}_t$, and $\mathbf{W}_i\bar{\mathbf{H}}_t$ (where $\mathbf{W}_i$ refers the generic row $i$ of matrix $\mathbf{W}$) as instruments starting from $t-2$. Table 2 contains their (total) contemporaneous serial correlation matrix. Coherently with the requirement of good instruments, Table 2 shows that they are quite uncorrelated among themselves, while coefficients of the regressions of instruments on explanatory variables are always significant with $R^2$ ranging from 0.22 to 0.51.

## 5.2 Discussion of Results

Tables 3 and 4 show respectively NLIV and NLLS estimates of the non-linear market potential function (3). In order to facilitate the comparison with Hanson [17], in these estimations I have used the inverse exponential as a distance decay. The first column of each table refers to results on all provinces while the second to data on continental provinces only. However, in all specifications, the two set of estimations do not differ significantly, so that I will refer directly to estimates on all provinces. First of all, one can notice that estimates from table 4, which are obtained with the same methodology proposed by Hanson [17], look

[19] Further details about spatial aggregation and instruments are given in the Appendix.

[20] Linearization could in principle lead to an estimation bias. However, my results suggest that this bias is reasonably small. As both NLLS and NLIV estimations turn out to be quite unstable (because of the many local minima of the criterion function), I try to get some help from the linearized specification of the model. Starting from equation (4), I have basically followed the two procedures behind non-linear estimations in order get simple linear criterion functions from which I got the correspondingly uniquely identified coefficients. I have then used these coefficients as starting values in the non-linear procedures, obtaining more reliable estimates (the corresponding value of the criterion functions seems to be the global minima) that are very close to the initial linear ones (within a range of $\pm 15\%$).

very similar to his findings. Although precaution is needed, because the limited data set dimension causes standard errors to be quite high, this suggest that *the different proxy of $w$ I use for Italy is a good choice.* I am in fact able, replicating his technique, to get something that is perfectly consistent with the results Hanson got using local wages for US.

However, a closer comparison of Tables 3 and 4, reveals immediately two important things. Although both procedures rest on the same statistical hypothesis, NLIV estimates are more precise and, with particular reference to $\sigma$, quite different from NLLS ones. As argued in the above Section, precision is a consequence of the more efficient way in which NLIV treats the information. Moreover, the fact that Hanson's procedure actually mixes county with state data in the same regression equations could lead to an aggregation bias. Coherently with my NLLS results, in Hanson [17] values of $\sigma$ lies between 6 and 11. By contrast, NLIV here indicates something around 2, suggesting that the magnitude of the aggregation bias is important. In both cases the the Moran statistic does not detect a significant spatial correlation in residuals[21]. However, as argued in the previous Section, this does not suffice to rule out endogeneity problems. It is in fact the significance of the estimates themself that suggests that the aggregation trick does not work. As for the scope of spatial externalities, unreported estimations obtained using the polynomial function as distance function indicate that, while estimates of structural parameters are almost unchanged, the goodness of fit increases significantly. The generalized $R^2$ passes in fact from 0.34 (0.23) to 0.43 (0.35) in the NLIV (NLLS) specification for all provinces, suggesting that the underlying degree of spatial interaction is better captured by the slower declining power function.

Table 5, which is the most important for us, shows my panel results obtained using the power function. One could first note that the implied values of $\sigma$, $\mu$, and $\theta$ are all very precisely estimated, with values lying in the corresponding interval predicted by theory. As for $\mu$, its estimate is in fact between 0 and 1 and in line with more reasonable values of the expenditure on traded goods than Hanson's estimates. Actually, in Helpman's stylized model, product $M$ is probably best seen as the aggregate of traded goods, as opposed to the non-traded ones ($H$), like housing and non-traded services. In Italy, the share of expenditure on housing is around 0.2 (for US it is almost the same), implying that estimated $\mu$ cannot be smaller than 0.8. However in Hanson [17], as well as in my NLLS and NLIV estimates, $\mu$ is always too high with values around 0.9 or even bigger.

For the elasticity of substitution, I got estimates between 3 and 4 that are significantly different from Hanson's findings. Although recent empirical studies indicate, using sectoral data, values of the elasticity of substitution between 4 and 9[22], I do not believe that these values are coherent with my underlying framework. Helpman [22] is in fact a very aggregated vision of the economy with just two sectors: traded goods ($M$), and non traded ones ($H$). Consequently, the aggregate $M$ contains goods that are actually very different from consumers' point of view (like cars and shoes), and one cannot certainly expect to find high values for their elasticity of substitution.

[21] The null hypothesis of the test is the absence of spatial autocorrelation. The test statistic can be corrected, as I actually do here, to account for both endogeneity in regressors and instrumentation, and is asymptotically distributed as a standardized normal. See Anselin [1], and Anselin and Kelejian [2] for further details.

[22] See Feenstra [12], and Head and Ries [20].

As earlier mentioned, a crucial difference between the theory-based market potential (2) and the Harris type (1), is that the second does not control for wages and prices at other locations. In Helpman [22], an increase in other locations' housing stocks ($H_k$) or wages($w_k$), cause $w_i$ to increase in the long-run in order to compensate workers for lower housing prices and higher earnings they can enjoy elsewhere. Estimations suggest that both variables actually play a significant role, as explicitly measured by the significance of $B_2$ and $B_3$, in understanding the forces at work in a spatial economy.

Turning to endogeneity and correlation issues, one can notice that all specification tests support my panel estimation. The Sargan test on over-identifying restrictions does not in fact reject the validity of instruments. Furthermore, the two tests on time autocorrelation behave in the correct way. If the $u_{i,t}$ are not correlated over time, then one should detect a significant (negative) first order correlations in differenced residuals $\triangle\hat{u}_{i,t}$, and an absence of "pure" second order correlation[23]. As one can see, this is actually what I found. This suggests that the inclusion of the dynamic term $\ln(\mathbf{w}_{t-1})$ in the equation, which turns out to be strongly significant, has probably allowed us to properly "capture" the time-dynamics (that one needs to control for) in the model. Finally, to exclude the presence of residual spatial correlation an adequate test is needed. Anyway, as far as I know, there is still no test procedure that exploits both the time and cross-section information, that at the same time accounts for endogeneity and instrumentation. However, one can certainly test year by year, and this is what I have actually done in Table 6 where the Moran statistic has been calculated for those years in which a sufficient number of instruments were available. As one can see, I did not find evidence of a significant spatial correlation. Finally, as in the case of NLLS and NLIV, changing the decay matrix does not alter estimates dramatically but the $R^2$ of the specification with the inverse exponential decay is lower (0.36 for all provinces).

In order to have a better idea of the spatial extent of agglomeration forces, I have simulated the effect on $w$ caused by an exogenous temporary shock on income, as measured by equation (5). Using panel estimates from Table 5 (first column) I have first evaluated equilibrium wages by means of (5), using actual data on $\ln(w_{i,t-1})$, $\bar{w}_{k,t}$, $\bar{Y}_{k,t}$, and $\bar{H}_{k,t}$ for $t$ =1992. Then, I have decreased the 1992 income of the 5 Latium provinces (Roma, Latina, Frosinone, Viterbo and Rieti) by 10% before re-computing $\ln(w_{i,t})$. Finally, as (5) contains a dynamic term linking $\ln(w_{i,t})$ with its past values, I have computed the sum of yearly changes on $\ln(w_{i,t})$ induced by this shock on income, occurring in 1992, over the entire period 1992-1998.

Figure 1 shows the implied total percentage decrease in the values of $w_{i,t}$ consequent to this simulated shock. Although I am actually under-evaluating the effect of such shock because I do not account for all interdependencies of the model, Figure 1 points out clearly that the impact is certainly not negligible and, contrary to Hanson [17], it is not so geographically bounded. Furthermore, one may note that the shock seems to be "asymmetric", in the sense that southern provinces are more affected than northern ones. This is certainly not surprising in the light of Italian economic geography. Everything else equal, the purchasing power of Latium is in fact more important for the south where local demand, as measured by households' disposable income, is lower than in the richer north.

[23] See Arellano and Bond [3].

As for the robustness of results to the choice of the spatial weights I find that, when performing the simulation exercise with an inverse exponential decay, the spread of the shock is in line with Hanson's findings.[24] Nevertheless, as earlier mentioned, the inverse exponential has quite unrealistic implications, and has proven to be less capable of capturing the degree of spatial interaction in the data. As a further evidence on the pervasiveness of final-demand linkages, I have expanded the number of linear regressors in (5) pre-multiplying $\bar{w}_{k,t}$, $\bar{Y}_{k,t}$, and $\bar{H}_{k,t}$ by the sum of many spatial matrices ($\mathbf{W}_0^{100} + \mathbf{W}_{100}^{200} + ..$), each corresponding to observations for provinces within subsequent rings of 100 km around the centroid of a province (0-100 km, 100-200 km.), up to 400 km.[25] By evaluating the joint significance of the 3 parameters corresponding to each distance-band matrix, one can have another feeling of how agglomeration externalities attenuate with distance. In estimations, variables up to 200 km are still significant, again suggesting that the scope of market-proximity externalities is not so limited. This is actually in line with Rosenthal and Strange [34], who found that reliance on factors sensitive to shipping costs (manufactured inputs, natural resource inputs, and perishability of products) influences agglomeration for the US at the state level. Although my results seems to be at odds with the Rosenthal and Strange [35] and Henderson [23], who found a little role for spatial interaction, it is worth noting that these latter studies are concerned with localized externalities essentially stemming from knowledge flows. In this respect, I agree with Rosenthal and Strange [36], who argue that different externalities are likely to differ substantially in their geographical scope.

# 6 Conclusions and Directions for Further Research

The NEG literature has provided a series of fully-specified general equilibrium models capable of addressing rigorously the agglomeration phenomenon. The combination of increasing returns, market imperfections, and trade costs creates forces that, together with factor endowments, determine the distribution of economic activities.

Following the approach of Hanson [17], I have first derived a theory-based market potential function (obtained from a multi-location extension of Helpman [22] model), relating the attractiveness of a location to the spatial distribution of factor earnings, consumers' expenditure, and non-tradable goods. Using a time-space panel data on Italian provinces, I have then estimated a linearized version of this equation by means of an innovative estimation technique, based on Arellano and Bond [3] and Anselin and Kelejian [2], that is needed in order to effectively address those endogeneity issues that arise when dealing with structural models and spatial data. I also provide evidence that the spatial aggregation approach used by Hanson [17] may suffer from a serious bias problem.

[24] In such a simulation, we use estimates of $\sigma$ and $\mu$ coming from panel regressions, while the value of $\tau$ comes from NLIV estimation. The rest of the exercise works as before.

[25] The size of Italian provinces is such that rings of less than 100 km would lead to some isolated locations. Arellano and Bond (1991) estimator is used in regression, i.e. the model is estimated in first differences with lagged values of $\ln(w_{i,t})$, $\mathbf{W}_0^{100}\bar{\mathbf{w}}_t$, $\mathbf{W}_0^{100}\bar{\mathbf{Y}}_t$, $\mathbf{W}_0^{100}\bar{\mathbf{H}}_t$, $\mathbf{W}_{100}^{200}\bar{\mathbf{w}}_t$, $\mathbf{W}_{100}^{200}\bar{\mathbf{Y}}_t$, $\mathbf{W}_{100}^{200}\bar{\mathbf{H}}_t$,... as instruments (starting from $t-2$).

My results are consistent with the hypothesis that final-demand linkages actually influence the distribution of earnings. Furthermore, my simulations suggest that, contrary to Hanson [17], the scope of such spatial externalities is not so limited. This latter result is mainly due to the different choice of the spatial decay matrix. Nevertheless, I provide evidence in favor of the power function specification. As a further check, I also perform some unstructured linear estimations using many distance-bands matrices in the same spirit as Rosenthal and Strange [35] and Henderson [23]. Variables up to 200 km are found to be still significant, further suggesting that attenuation of market-proximity externalities is less rapid compared to other agglomeration forces.

There are several possible directions for further research. One natural extension of my framework would be to obtain estimates using European data. As shown by Overman and Puga [30], national borders are in fact less and less important in Europe, while regions are becoming the best unit of analysis. A second issue is related to the simplifying assumptions that lead Helpman [22] to be cumbersome for empirical interpretation. The fact that $\sigma$ is at the same time a measure of different things in these kind of models is annoying. A promising alternative approach is the one proposed by Ottaviano, Tabuchi, and Thisse [29]. Using a more elaborated demand structure and transportation technology, this model allows in fact a clear separation (by means of different parameters) of elasticity of demand, elasticity of substitution and increasing returns, as well as firms' pricing policies. Finally, a deeper understanding of the functional form that is more suited to describe the degree of spatial interaction among data is needed. In this respect, the pioneering work of Pinkse, Slade and Brett [31] that tries to endogenize the spatial correlation matrix by means of polynomial approximation may turn out to be a useful tool.

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## Appendix: Details of Estimations

To construct instruments for NLIV and regressors for NLLS, I have adopted the following procedure. I first divide Italy into 11 zones using NUTS-1 regions. After having transformed (3) with a time difference, for NLIV estimation I have then used, for each province, the change (over the time interval 1991-1998) in the logarithm of the variables $w$, $Y$, and $H$ of the corresponding zone (reconstructed aggregating provinces data) as instruments. I thus have a set of exactly 3 instruments for the 3 parameters to estimate in (3), and so there is no need of an optimal weighting matrix. For NLLS, I have instead used directly levels of $w$, $Y$, and $H$, corresponding to the eleven zones, as regressors before making first difference and applying least squares. In both cases, I have also neutralized,as in Hanson [17], the specific contribution of each province in the formation of the corresponding zone aggregate variable. As a remedy for spatial heterogeneity, I have used White heteroscedasticity-consistent standard errors. For the Moran test, I used the pseudo-regressors as explanatory variables. Finally, all estimations have been performed with Gauss for Windows 3.2.38.

Panel estimates are two-step GMM ones and have been obtained with DPD 98 for Gauss. The model is estimated in first differences, using past levels of all explanatory variables, from $t-2$ and later, as instruments. The reason why I treat all variables as endogenous is that, in unreported estimations, I actually found evidence that also the housing stock process suffers from simultaneity. Estimations includes time-dummies, while standard errors and tests are all heteroscedasticity consistent.

## Table 1: Summary Statistics

| | Mean | Standard Dev. | Min | Max |
|---|---|---|---|---|
| $w$ | 65.1 | 12.98 | 42.63 | 100.52 |
| $Y$ | 13,055,315 | 16,166,672 | 1,772,886 | 115,120,309 |
| $H$ | 19,247,567 | 19,714,984 | 3,271,507 | 130,723,464 |
| Area | 2,925.64 | 1,750.38 | 211.82 | 7,519.93 |
| Population | 557.87 | 615.56 | 92.15 | 3,781.79 |

All nominal variables are in 1996 prices and the unit is one million liras. Housing $H$ is measured in squared meters, while population is in thousand of people and provinces area is expressed in squared kilometers. Data are time-averaged and refers to the interval 1991-1998

## Table 2 : Correlation Matrix of Panel Instruments

| | $\ln(w_{i,t})$ | $\mathbf{W}_i\bar{\mathbf{w}}_t$ | $\mathbf{W}_i\bar{\mathbf{Y}}_t$ | $\mathbf{W}_i\bar{\mathbf{H}}_t$ |
|---|---|---|---|---|
| $\ln(w_{i,t})$ | +1 | +0.54 | +0.36 | +0.32 |
| $\mathbf{W}_i\bar{\mathbf{w}}_t$ | +0.54 | +1 | +0.74 | +0.81 |
| $\mathbf{W}_i\bar{\mathbf{Y}}_t$ | +0.36 | +0.74 | +1 | +0.70 |
| $\mathbf{W}_i\bar{\mathbf{H}}_t$ | +0.32 | +0.81 | +0.70 | +1 |

Variables are in levels, and the entire sample period (1991-1998) have been used to compute time-averaged variances and covariances

## Table 3: NLIV estimates for Helpman Model

| | | |
|---|---|---|
| $\mu$ | 0.8741** | 0.8687** |
| (stand. error) | (0.1939) | (0.1726) |
| $\sigma$ | 1.9196** | 2.0219** |
| (stand. error) | (0.4876) | (0.5327) |
| $\tau$ | 0.1895** | 0.1698** |
| (stand. error) | (0.0523) | (0.0491) |
| $\sigma(1-\mu)$ | 0.2417 | 0.2250* |
| (stand. error) | (0.1314) | (0.1126) |
| $\sigma/(\sigma-1)$ | 2.0874** | 1.9786** |
| (stand. error) | (0.3071) | (0.3201) |
| Wald test joint sign. | 63.231 | 68.818 |
| (degrees of freed, impl. prob) | (df=3, p=0.000) | (df=3, p=0.000) |
| Moran test spat. correl. | 1.212 | 0.961 |
| (implied prob) | (0.2255) | (0.3365) |
| Adjusted $R^2$ | 0.4201 | 0.5136 |
| General. $R^2$ | 0.3392 | 0.3448 |
| Provinces | All | Continental |
| N° of observ | 103 | 90 |

** Indicates estimates significant at 1% level, while * indicates estimates significant at 5%.

## Table 4: NLLS estimates for Helpman Model

| | | |
|---|---|---|
| $\mu$ (stand. error) | 0.9106* (0.4561) | 0.9394 (0.5652) |
| $\sigma$ (stand. error) | 5.9128 (3.9692) | 6.7531* (3.3469) |
| $\tau$ (stand. error) | 0.9351 (2.0882) | 0.7495 (1.1421) |
| $\sigma(1-\mu)$ (stand. error) | 0.5877 (1.7527) | 0.2880 (1.0182) |
| $\sigma/(\sigma-1)$ (stand. error) | 1.2035* (0.6077) | 1.2664* (0.5872) |
| Wald test joint sign. (degrees of freed, impl. prob) | 14.228 (df=3, p=0.0026) | 15.124 (df=3, p=0.0017) |
| Moran test spat. correl. (implied prob) | 0.522 (0.6016) | 0.991 (0.3216) |
| Adjusted $R^2$ | 0.2521 | 0.1987 |
| General. $R^2$ | 0.2346 | 0.2893 |
| Provinces | All | Continental |
| N° of observ. | 103 | 90 |

** Indicates estimates significant at 1% level, while * indicates estimates significant at 5%.

## Table 5: Panel estimates for Helpman Model

| | | |
|---|---|---|
| $A$ (stand. error) | 0.3004 (0.0355) | 0.2731 (0.0348) |
| $B_1$ (stand. error) | 11.1843 (1.5507) | 10.9123 (1.7516) |
| $B_2$ (stand. error) | 27.7367 (4.2910) | 28.3625 (4.5608) |
| $B_3$ (stand. error) | 122.4511 ( 9.2635) | 117.8225 (8.9225) |
| $\mu$ (stand. error) | 0.7735 (0.0316) | 0.759278 (0.0302) |
| $\sigma$ (stand. error) | 3.4335 (0.6796) | 3.2778 (0.7345) |
| $\theta$ (stand. error) | 133.6351 (10.3401) | 128.7353 (10.8708) |
| Wald test joint sign. (degrees of freed, impl. prob) | 333.1169 (df=4, p=0.000) | 322.6295 (df=4, p=0.000) |
| Sargan test (degrees of freed, impl. prob) | 94.1325 (df=80, p=0.3621) | 101.2946 (df=80, p=0.1953) |
| $1^{st}$order time corr. (implied prob) | -3.643 0.0003 | -3.982 0.0000 |
| $2^{nd}$order time corr. (implied prob) | 0.389 0.6972 | -0.104 0.9172 |
| Adjusted $R^2$ | 0.4711 | 0.4568 |
| Provinces | All | Continental |
| N° of sample observ. | 618 | 540 |

## Table 6: Moran Test for panel estimations

| | 1993-1994 | 1994-1995 | 1995-1996 | 1996-1997 | 1997-1998 |
|---|---|---|---|---|---|
| Moran test spat. correl. all provinces | +0.2014 | +0.1301 | -0.7910 | -0.6370 | -0.4814 |
| (implied prob) | (0.8404) | (0.8965) | (0.4289) | (0.5241) | (0.6302) |
| Moran test spat. correl. cont. provinces | +0.4114 | -0.5632 | -0.3120 | +0.1425 | -0.7123 |
| (implied prob) | (0.6808) | (0.5733) | (0.7550) | (0.8867) | (0.4763) |

Test statistics have been computed with residuals of the model estimated in first differences.

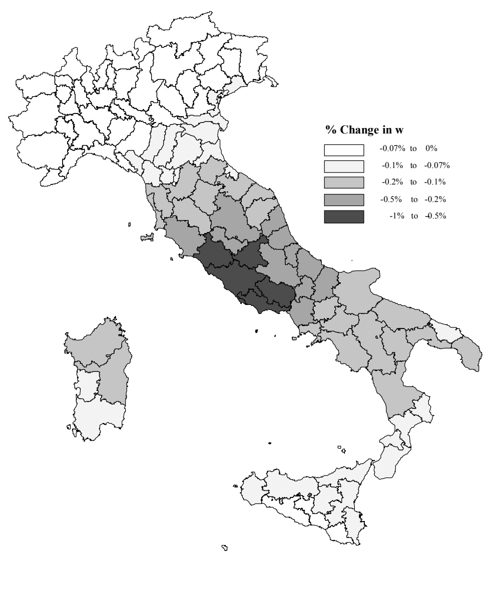


Figure 1: Simulated $w$ changes from income shock to Latium