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THE CONSEQUENCES OF PUBLIC EMPLOYMENT: EVIDENCE FROM ITALIAN MUNICIPALITIES

by Marta Auricchio*, Emanuele Ciani*, Alberto Dalmazzo** and Guido de Blasio*

Abstract

We investigate the consequences of public employment on local economies. We start by presenting a spatial-equilibrium framework, to highlight that the housing market is an important channel through which a variation in public employment affects private employment. We then provide empirical evidence from Italian municipalities, focusing on the strong contraction in the public sector workforce that occurred between the last two Censuses (2001-2011). We use an IV identification strategy that exploits the fact that variations in local public employment were strongly influenced by central government decisions, with little reference to the economic conditions of the municipalities. Our results suggest that exogenous contractions in public employment lead to an increase of private jobs, and that competition in the housing market seems to be a relevant explanation for this finding.

JEL Classification: J45, J60, R12.
Keywords: local labor markets, public employment.

Contents

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1. Introduction[1]

The interaction between public and private employment is a long-standing issue in the policy debate. The expansion of public jobs may come as a policy reaction to low private employment. At the same time, private employment may react to public employment. On the one hand, the goods and services produced by the public sector may favor a business-friendly environment or better amenities and, further, boost the demand for privately produced goods. On the other hand, public employment may crowd out the private sector employment, as it can raise the costs of production by increasing the local cost of scarce resources, and/or reduce the labor supply available to the private sector. Which one of the two effects dominates is an empirical question, and is essential to understand whether an increase in public jobs, which might also be motivated by unemployment concerns, is effective or not.

Given the relevance of this question, the macroeconomic literature has recently witnessed an expanding number of contributions that model the public sector within a search framework. Burdett (2012) and Gomes (2014) suggest that higher public wages lead to a contraction in private employment. From the empirical point of view, this crowding out effect is confirmed by

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previous studies based on international comparisons (Boeri et al, 2000; Algan et al, 2002; Behar and Mok, 2013).

Although policy decisions are often made at the national level, the expansion of the public sector is usually very differentiated at the local level, both because of historical reasons, in particular past policies, and administrative concerns, for instance the need of a minimum set of services even in low density areas. The analysis at the local (sub-national) level can therefore exploit this source of variation to shed light on the relation between private and public employment and to understand which are the main mechanisms behind it, studying also the effect on local prices. In our perspective, the existing estimates obtained from cross-country comparisons might be plagued by concerns about the causal interpretation. Furthermore, forecasts obtained by macro-econometric models, which are usually based on time series at the national level, rely on strict identification assumptions (Moretti, 2010). Therefore, the estimates that exploit the variability at the local level might provide additional empirical evidence to inform the macro debate.

According to most studies, in Italy public sector wages are higher than in the private sector (Giordano, 2010; Depalo et al, 2015). Thus, an exogenous increase of public employment in one area will likely lead to an upward pressure on salaries, and also increase the demand for locally produced goods, in particular housing. This demand-driven push may have beneficial effects on the local non-tradable sector but, at the same time, if house prices increase, private sector workers are more likely to leave for other areas. On the other hand, the presence of a larger public sector can have impact on the productivity of local firms, if the additional public employment allows for the provision of better services. Employment in the tradable and non-tradable sectors might, therefore, have an additional (supply-driven) push. Nevertheless, this effect could also go the other way round if, for example, a larger public sector generates obstacles to doing business.

Empirically, Faggio and Overman (2014) provide a thorough econometric analysis of the impact of annual variations in public employment on private employment and working age population growth, focusing on British Local Labor Markets and using a shift-and-share instrumental variable approach. They find no crowding out for aggregate private employment, but this is the result of a negative effect on manufacturing (tradable) and a positive impact on services and constructions (non-tradable). Similar analyses using the shift-and-share approach on geographical disaggregated data from other countries were performed by Senftleben-König (2014) and Ranzani and Tuccio (2017). The former finds a sizeable negative effect in German local labor markets between 2003 and 2007, driven by a strong crowding out in the tradable sector. Ranzani and Tuccio estimate an overall negative effect on private employment in three African countries (Ghana, Mali, Mozambique), but driven by agriculture, while other sectors gains from increased public workforce. Jofre-Monseny et al (2016) propose a matching model and, using Spanish data between 1980 and 2001, estimate that an increase in local public employment crowds-in non-tradable jobs and crowds-out tradable ones, with an overall positive effect. Differently from Faggio and Overman (2014), their analysis exploits the status of a provincial city as an instrument for the variation in public employment. They also calibrate their matching model on the Spanish economy and confirm the positive effect, which is associated with an increase in land prices and wages. In a similar fashion, Caponi (2017) proposes a regional search model and calibrates it with Italian data, finding however significant crowding out, driven by the fact that public wages are set nationally and are particularly high in low-productivity regions. Finally, other studies have been looking at relocation episodes. Faggio (2015) studies the effects of a relocation policy for civil service workers in Britain. She finds a positive multiplier effect on the private sector in receiving areas, mostly driven by services. Becker et al (2015) evaluate the impact of public employment on private sector activity, by considering a relocation episode: the move of the

German federal government from Berlin to Bonn in the wake of the Second World War. They basically find no effect on private employment. Faggio et al (2016) exploit the opposite natural experiment, which occurred when the government was brought back to Berlin in 1999. Their evidence supports a different result, that the increase in public employment brought a crowding-in of service sector jobs and no impact on manufacturing.

In this paper we contribute to the literature in two respects. Firstly, we highlight that crowding out may also occur because of competition on the housing market. To this end we outline a spatial-equilibrium model with public employment.[2] In this framework, an increase in local public employment will create an upward pressure on housing prices. This may decrease the relative attractiveness of the area for private workers, as their real wage tends to decrease. Depending on their idiosyncratic preferences for the area, they may move to other places. In the absence of any impact of public employment variations on local productivity and the quality of amenities, the interplay between private and public employment is entirely driven by the competition in the local housing market. However, the actual impact of public employment is ambiguous whenever local productivity or amenities are positively affected by the expansion of the public sector.

Secondly, we provide empirical evidence on the local consequences of public employment using data from Italy at the municipality level between 2001 and 2011. This is an interesting case study because during that decade public employment experienced a significant contraction. Our empirical analysis follows Faggio and Overman (2014), who proposed an IV strategy where public employment growth is instrumented with a predicted change derived by applying national variations to the share of public employment at

[2] Our spatial model is related to the matching model in Jofre-Monseny et al (2016) who, however, explicitly consider three sectors (tradable, non-tradable, and public sector) and agglomeration externalities. There, public employment tends to crowd-in employment in local non-tradables, while it has smaller effects on activity in the local tradable sector. These issues are also quite central in the model presented in Becker et al (2015), where different tradable goods are produced in different areas and transportation across locations involves iceberg costs.

the beginning of the period, in a shift-and-share fashion. In our case, this approach is justified by the fact that, over the period 2001-2011, Italian central authorities cut down local public employment through a halt in turnover, essentially for nationwide budgetary reasons and with little reference to the economic conditions of the municipalities. Although we draw on Faggio and Overman's strategy, we differentiate from them in several respects. First of all, in line with the theoretical analysis, we focus not only on the impact on private employment but also on house prices. Secondly, we conduct a set of additional sensitivity checks intended to probe the appropriateness of our IV approach. By using a shift-and-share instrument for the private sector, we discuss to what extent our IV results are driven by simultaneity. We also show that a reasonable degree of failure on the exclusion restriction does not invalidate our estimates (Oster, forthcoming), while the findings nicely survive from Machine Learning (ML) selection of covariates (Belloni et al, 2014). Our main results highlight clean crowding out: an additional public employee reduces private employment by 0.6/0.8 unit on average. The effect is driven by the tradable sector, while there seems to be no impact on the non-tradable. Furthermore, the result seems to be at least partially driven by competition on the housing market, because house prices rise, as predicted by the theory. Section 2 presents the theoretical model and its main predictions. Section 3 discusses the identification strategy and presents the data. Results are showed in subsection 3.3. Conclusions follow.

2. Theoretical Background

In order to analyze the role of public employment, we outline a Roback's (1982) spatial model with "mobility costs", as in Moretti (2011), Kline and Moretti (2014), and others. The details and the solution of the model sketched here are reported in the Appendix A. The economy is composed of two regions, denoted by $c = \{a, b\}$. Firms are fully mobile across areas and produce a *tradable* good with a Cobb-Douglas, constant-returns-to-scale technology.[3] Production requires only skilled and unskilled labor and sells at a price equal to one across areas. Skilled and unskilled individuals, instead, do have idiosyncratic preferences for locations. A skilled individual living, say, in area *a* has Cobb-Douglas preferences denoted by $U_a^s = \ln\left(A_a{}^s\right) + (1-\gamma)\cdot\ln\left(x_a\right) + \gamma\cdot\ln\left(L_a\right) + \varepsilon_a^s$, an increasing function of the consumption of tradables, x_a, and housing services, L_a. Utility is also increasing in local amenities which appeal to the skilled, denoted by $A_a{}^s$, and in the realization of the preference shock ε_a^s for location *a*. A high realization of ε_a^s implies that the individual may be unwilling to move from place *a* to *b* even when amenities and the wage-rent ratio in location *b* are larger than in location *a*.[4] Hence, preference shocks generate "mobility costs": labor supply is not perfectly elastic across locations, differently from the basic Roback's model. Unskilled workers have similar preferences, given by $U_a^u = \ln\left(A_a{}^u\right) + (1-\gamma)\cdot\ln\left(x_a\right) + \gamma\cdot\ln\left(L_a\right) + \varepsilon_a^u$. The model is closed by the equilibrium condition for the market of local housing services. Individual demands for housing, L, are aggregated across skilled and unskilled individuals employed in the private and public sector. Housing supply is an increasing function of both residential land and local rents.

Public employees can be skilled and unskilled. The size and allocation of public employment across regions is exogenously determined by the central government. We also postulate that the wages for public employees may differ

[3] The models in Faggio and Overman (2014), and Jofre-Monseny et al. (2016) allow for the presence of both local tradable and non-tradable sectors. We also developed a two-sector version of the model, where the non-tradable local sector is subject to "multiplier effects", after Moretti (2010). The two-sector model is available as an additional Appendix in the authors' website (Auricchio et al, 2016).

[4] The idiosyncratic preference shock ε_c^s, for $c = \{a, b\}$, is assumed to follow a Type I Extreme Value distribution (see, e.g., Kline and Moretti, 2014, and Diamond, 2016), with a scale parameter accounting for the size of "mobility costs". The location decision depends on the difference between the idiosyncratic preference shocks of the two areas. See also Ch.2 in Anderson et al (1992).

between skilled and unskilled, but within each category they are equal across regions, and set at the national level. Consistently with evidence from Italy reported in Giordano (2010) and Depalo et al (2015), wages in the public sector are taken to be larger than the corresponding ones in the private sector.[5]

The basic mechanism at work in the present framework hinges on the local housing market. Suppose, at least initially, that local public employment has no impact on local amenities and productivity. Then, the only effect of public employment is that public employees will come to compete on the local housing markets with individuals employed in the private sector. In particular, an increase in the mass of local employees will increase the local demand for housing, and displace individuals who were employed in the local private sector. The effect on local rents is thus driven by such two opposing effects on housing demand.

These implications can be made sharper by introducing some notation. As is clear from the Appendix A, the model is solved by log-linearizing equations and calculating deviations around "symmetry". In other words, we derive our results by assuming that the two areas are initially identical, and then we suppose that public employment in location *b* increases more than public employment in location *a*. In symbols, we denote such an event as $\tilde{\Pi} > 0$. We then ask how *private* employment in area *b* will change, relative to private employment in area *a*, after the public employment shock. By denoting the relative changes in skilled and unskilled private employment respectively by $\tilde{N}$ and $\tilde{n}$, and the change in relative rents by $\tilde{r}$, we obtain the following solutions:

$$\tilde{N} = B_1 \cdot \tilde{\Pi}; \quad \tilde{n} = B_2 \cdot \tilde{\Pi}; \quad \text{and} \quad \tilde{r} = D \cdot \tilde{\Pi}, \qquad (1)$$

[5] As in Faggio and Overman (2014), we abstract from the explicit consideration of the public sector budget constraint by postulating that local public sector wages are financed from *national* taxation: such taxation, indeed, does not alter the *relative* conditions of the two areas considered.

where $\{B_1, B_2, D\}$ are expressions reported in the Appendix A which depend on the parameters of the model, including the size of mobility costs. We can thus give a sum-up of the main implications, starting with changes in private employment and, then, consider rents. Notice that in the real world different areas are hardly symmetric at the beginning of any period. In the empirical application we discuss in detail how this heterogeneity is accounted for in the specification choice and in the selection of control variables.

Private employment changes. Independently of the size of mobility costs, we obtain that the direct effect of local public employment is unambiguously negative, since $\{B_1, B_2\} < 0$. Thus, if it has no indirect effect through local amenities or productivity, an increase in local public employment will always crowd out private employment.

Also, if skilled and unskilled individuals have the same measure of mobility costs, it will hold that $B_1 = B_2$. As a consequence, the impact of public employment on private employees will not change the local skill mix. The size of private employment displacement depends on the size of mobility costs. Displacement is smaller when mobility costs are higher. In other words, the absolute values of $\{B_1, B_2\}$ get smaller the larger mobility costs.

Finally, if the skilled bear mobility costs smaller than the unskilled, a local public employment shock will worsen the skill mix, that is, the skilled in the private sector will decrease faster than the unskilled.

Local rent changes. For what it concerns change in local rents, given by $\tilde{r} = D \cdot \tilde{\Pi}$, it holds that D is non-negative. When mobility costs are negligible, increased demand for local housing by public employees will be matched by a decrease in demand due to reductions in private sector employment, and local rents will *not* change in equilibrium. Indeed, when workers are fully mobile across areas, they will be ready to leave whenever local rents tend to increase: thus, people will move away until rents stay the same. By contrast, the presence of mobility costs implies an increase in local rents.

As clear from the Appendix A, the results exposed so far emphasize the crowding-out effect of public employment through the local housing market. However, if local public employment exerts a positive and sizeable effect on the local quality of life, or on local productivity (as, e.g., in Becker et al, 2015), the crowding-out of local private employment gets smaller. At the extreme, public employment may even crowd-in private employment (in this case, however, there will be a larger positive effect on local rents).

3. Empirics

In this section, we provide empirical evidence on the impact of local public employment using data from Italy. We focus on joint adjustments in private employment and local prices by presenting "reduced-form" estimates, meant to deal with causality issues. Then, we discuss how our findings can be interpreted through the lens of the theoretical predictions. We start (Sect. 3.1) by explaining our IV identification strategy. Then we describe the data (Sect. 3.2). The results are shown in Sect. 3.3.

3.1 Identification

3.1.1 Specification. Our main equation relates growth in private sector employment to growth in public sector between the Census waves 2001 and 2011 (we focus only on the most recent wave because we do not have earlier information on house prices at the municipality level). We follow Faggio and Overman (2014) and model the relation as linear in *contributions* to overall employment growth:

$$\frac{N^{priv}_{c,2011} - N^{priv}_{c,2001}}{N_{c,2001}} = \beta_0 + \beta_{pub} \frac{N^{publ}_{c,2011} - N^{publ}_{c,2001}}{N_{c,2001}} + x_{c,2001}\beta_x + \epsilon_{N_{priv},c} \qquad (2)$$

where $c = 1, \dots, M$ are Municipalities. In this specification, the coefficient β_{pub} can be interpreted as the unit change in private employment associated with one unit change in public employment. $x_{c,2001}$ is a $1 \times K$ vector of control variables with reference to the beginning of the period. For house prices per square-meter, we instead estimate

$$\frac{pr_{c,2011} - pr_{c,2003}}{pr_{c,2003}} = \gamma_0 + \gamma_{pub} \frac{N^{publ}_{c,2011} - N^{publ}_{c,2001}}{N_{c,2001}} + x_{c,2001}\gamma_x + \epsilon_{pr,c} \, .$$

(3)

The variation is taken with respect to 2003 because of data constraints (see the data section below). In this case, the coefficient of interest can be read as the percent change in house prices associated with a 1% contribution of public employment to growth (the s.d. of the latter is 5%). Peri and Sparber (2011) suggest that this specification avoids the problem of spurious correlation that affects growth-to-growth or changes-to-changes specifications (for a similar specification, see also Card, 2007).

Our unit of observation is the municipality. That is, the smallest administrative jurisdiction unit in Italy and the ideal geographic reference point for our analysis, which focuses on the local impact of the public-employment variations that are decided at the administrative level (not at the functional one, such as the local labor market). Municipality's boundaries in Italy are defined to cover the entire surface. A map with all boundaries can be found in Appendix B.

3.1.2 Baseline controls. The choice of considering administrative entities, however, comes at some costs. The possibility that what happens in one single municipality spills over neighboring municipalities cannot be excluded: this occurrence would put our identification strategy in danger by invalidating the SUTVA (Stable Unit Treatment Value Assumption). This is why we include in $x_{c,2001}$ some control variables that are likely to differentiate out the potential linkages between the single municipality that experiences a given variation in public employment and its surroundings. Monte et al (2015) suggest that commuting to other areas captures most of cross-border spillovers. Thus, we include the best available proxy for mobility, which is the fraction of the population aged below 64 resident in the municipality that moves daily to other municipalities for work or study.

Although the theoretical model assumes perfectly competitive markets, several areas in Italy are far from full employment. This may lead to bias in our results. Places characterized by a larger fraction of non-employed individuals might display smaller crowding out, since both the local labor market and the local housing market are slack. We thus control also for the unemployment rate (in population aged 15 or more) and for the overall participation rate. By the same token, we control for slackness in the housing market and add an index of housing availability equal to the fraction of vacant housing units over total housing in the municipality. Moreover, given that each municipality might have ties with the surrounding ones, we include the simple average of these four variables across the other municipalities of the same Local Labor Market (LLM), defined by ISTAT as an approximately self-contained area in terms of commuting (on the basis of census data).[6] All these variables, including the mobility index, are defined with reference to the beginning of the period. In Sect. 3.3 we also check whether our results are driven by reallocations of workers and residents from nearby municipalities, by switching the unit of analysis to the LLM level. If all the results are simply driven by reallocations within very short distances, then we should find milder effects – if any – at the LLM level of aggregation.

[6] Some of these indicators for commuting and idle labor are not perfect (for instance, they include age groups that we are not interested in) but, to the best of our knowledge, they are the closest approximation that we can build on the available data.

From an econometric perspective, our equation is basically a difference-in-differences specification, with a continuous regressor, and therefore we must worry about the possible failure of the parallel trend assumption. We therefore include in $x_{c,2001}$ the past trends (between 1991 and 2001) of private and public employment contributions to growth, plus the past growth in working age population (expressed as a contribution over employment).[7] Similarly, we include the past trends for the same variables in the rest of the LLM. As our main dependent and explanatory variables have previous total employment as denominator, we also include the initial (1991) levels of the log of employment, total population and house prices.

3.1.3 Additional controls. We select all the covariates above guided by theory and econometric concerns. However, we have a much wider set of available variables at the municipality level from a recently released database (*ottomilacensus*). These variables are indices calculated from the 1991-2001 population and housing censuses, and include information on demographic structure, housing conditions, self-reported occupational status, commuting and social vulnerability. A complete list is available in the Appendix B, while we refer to the website of *8milacensus* for further details.[8] Given that our estimates might depend on these characteristics, we also include all available variables (with no missing information) measured both in 2001 and 1991.[9] To address concerns about the inclusion of too many additional covariates, in Section 3.3.3 we also discuss whether results are robust to selecting only the most relevant one following a procedure proposed by Belloni et al (2014).

[7] Unfortunately, we do not have house price trends at the municipality level for that period.
[8] These also include the 1991 values for the four indices mentioned above (unemployment rate, labor force participation rate, fraction of unoccupied houses, commuting index).
[9] We use the time series of data with municipality boundaries fixed at 2011. http://ottomilacensus.istat.it/fileadmin/download/Descrizione_degli_indicatori_serie_confini_2011.xlsx (last access: 07/12/2015)

3.1.4 IV approach. In spite of the relevance of the baseline and additional controls, we still have two potential problems of identification:[10]

- *Omitted factors* may influence both private and public employment. For instance, an increase in local productivity or quality of life that is not caused by variations in public employment might still spur both private and public demand for labor.
- *Simultaneity* cannot be excluded, as private sector employment may also influence public employment. For example, the local authority may adjust its public employment target by looking at the growth of private employment.

To tackle these issues we adopt an IV strategy, which builds on Faggio and Overman (2014). Our instrument derives from the well-known Bartik (1991) logic, applied to the public sector and to the specification in contributions to growth. We sum up national growth in each sector *j* of public employment, and we multiply it for the public employment weight in *j* for that municipality in the previous period:

$$inst_{c,2011}^{publ} = \sum_{j\in pub} \frac{N_{j,c,2001}^{publ}}{N_{c,2001}} \times \frac{N_{j,-c,2011}^{publ} - N_{j,-c,2001}^{publ}}{N_{j,-c,2001}^{pub}} \tag{5}$$

$$\frac{N_{c,2011}^{publ} - N_{c,2001}^{publ}}{N_{c,2001}} = \gamma_0 + \gamma_1 inst_{c,2011} + x_{c,2001}\beta_x + \eta_c \,. \tag{6}$$

To be precise, for each municipality *c* the national growth is calculated by omitting the municipality itself. Intuitively, the instrument is the predicted contribution of public employment to overall local employment growth, calculated using national trends, which are strongly influenced by the central government decision to downsize the expenditure in human resources. Following the aggregation algorithm previously discussed, public employment

[10] Measurement error is less likely to be a concern, given that we are using Census data.

is spread across 8 sectors, whose description can be found in Appendix B. The main ones are the two education categories, that together account for around one third of total public employment in 2011, the human health activities, including hospitals, which accounts for approximately one fourth, and the first category (administration of the state and the economic and social policy of the community), accounting for slightly more than one fifth.

In order to use this predicted growth as instrument, we impose $E\left[(\epsilon_{N_{priv},c}, \epsilon_{pr,c})|1, inst_{c,2011}^{publ}, x_{c,2001}\right] = (0,0)$. This assumption is credible as long as policies at the national level set targets for public employment adjustment that are to be followed at the local level. However, it is not necessary that these rules are precisely followed at the local level (which would, by itself, make public employment exogenous). Such deviations from the rule, which are captured by η_c , are essentially what causes endogeneity (as long as $E\left[\eta_c \epsilon_{N_{priv},c}\right] \neq 0$ and $E\left[\eta_c \epsilon_{pr,c}\right] \neq 0$). The instrument is going to be valid as long as predicted growth, which captures the policy target, is not related to specific shocks to the private sector ($\epsilon_{N_{priv},c}$) and to the house prices ($\epsilon_{pr,c}$).

The instrument seems to be appropriate for the Italian case. First, local public sector employment is overwhelmingly financed through transfers from the central government, not local taxation; consequently, the allocation of public employees over the national territories are mostly decided at the central level. When the public budget constraint is local there would be an obvious link between the local private sector and the public one, as richer local economies can afford better public services. In our case this direct link is not there. However, the fact that decisions are centralized does not necessarily imply that public servants allocated to a given area do not reflect the economic fortunes of the place. For instance, lagging areas might get an higher share of centrally decided public workers, in the attempt to counterbalance local unemployment (see: Alesina et al., 2000). In the decade we consider this redistributive motive has been greatly impaired, because of the limitation imposed by EU and national legislations. In particular, several laws (in 2002, 2004, 2006) introduced a total or partial stop to new hires. This led to a dramatic slowdown in turnover, and prevented the replacement of employees entering retirement, especially where local authorities were not meeting budgetary targets.[11] Such stops in turnover can essentially be interpreted as proportional cuts in employment, where the fraction of public employees entering retirement is not replaced by new hires.[12] Second, the nationwide decisions referring to public employment have a sectoral component, as they are bargained with sectoral labor unions (they also depend on the strength of unions vis-à-vis the incumbent government; for instance, school teacher unions, which are traditionally left-wing oriented, usually get better deals with center-left governments).

The instrument used by Faggio and Overman (2014), which is

$$\mathrm{inst}_{\mathrm{c},2011}^{publ,fo} = \frac{\mathrm{N}_{\mathrm{c},2001}^{publ}}{\mathrm{N}_{\mathrm{c},2001}} \times \frac{N_{-c,2011}^{publ} - N_{-c,2001}^{publ}}{N_{-c,2001}^{publ}}, \tag{7}$$

neglects the sectoral composition of public employment and, therefore, uses spatial heterogeneity in the public employment share at the beginning of the period as the only source of variation. However, decisions regarding the size of the public sector are possibly different across different activities. For instance, the cuts imposed on employees in the administration of local

[11] Between 2001 and 2011 the total number of employees in the public sector has decreased by 11 percent; instead in the previous census period (1991-2001) the number of employees in the public sector had been substantially stable (the increase was lower than 3 percent). We also obtain a very similar picture when considering contributions of the public sector to total employment growth. The small national change in public employment in 1991-2001 does not allow us to obtain, for that period, an instrument with sufficient variation. Furthermore, during the previous decade we do not have a similar national intervention on public employment which can justify this empirical strategy. We therefore focus only on 2001-2011.

[12] Clearly, this fraction is not necessarily the same in all municipalities, as it depends on the age structure of public employment. However, the purpose of the instrument is to exclude variations in public employment that may be systematically related to private employment growth. From this perspective, the age structure of the public employment observed in the decade 2001-2011 depends on the hiring decisions made from the 1960s to the 1980s when public sector employment boomed.

authorities may not be the same as those applied to health care centers run by the national health service. Nevertheless, we also checked that using $\mathrm{inst}_{\mathrm{c},2011}^{publ,fo}$ still delivers very similar results (see Section 3.3.4).

3.2 Data.

We exploit municipality-level data on private and public employment from the 2001 and 2011 Italian Industry and Service Census. The Census gathers data on local production units of firms, enterprises, institutions at the 31st of December, the reference date of each census. The subjects of the Census are legal-economic units operating in industrial and service sectors, public institutions, and non-profit institutions.

In the rest of the paper, "private sector" covers all enterprises carrying on economic activities which contribute to gross domestic product at market prices, in the industry, commerce and services sectors. Differently, the "public sector" refers to Public Institutions, defined as "economic entities that are capable of producing non-market goods and assets, intended for the benefit of the community and entirely financed by households, enterprises, nonprofit institutions and other public institutions".[13] The municipal enterprises and other government-controlled enterprises are classified in the Census as units operating in the industrial and service sectors. For this reason, privatizations concerning this kind of enterprises in last decades do not raise reclassification issues for our purposes. Still, it is important to emphasize that our definition of public sector employment excludes those firms that are directly or indirectly owned by central or local governments, as long as they produce market goods or services. In this paper we exclude non-profit enterprises because ISTAT has deeply changed the Census methodology and definition of the sector in the last decades. Furthermore, some changes in legal status have induced transitions between the public and the non-profit sector. This circumstance affects units that are scarcely relevant from the point of view of employment which,

[13] http://siqual.istat.it/SIQual/visualizza.do?id=8888952 (last access: 06/04/2016).

instead, is mostly concentrated in Ministries, Regions and Municipalities (whose legal status has obviously not changed). Unfortunately, ISTAT does not release specific data on such transitions, and we are therefore forced to exclude the non-profit sector from the analysis. In 2011, these organizations included approximately 681 thousand employees, compared to 2.842 million in public institutions and 16.242 in the private sector.

In both private and public sectors, employment includes employees with fixed-term or permanent employment contracts, and the self-employed.[14] The number of contractors (essentially collaborators with non-standard contracts) and the number of temporary workers (apart from those on standard fixed-term contracts) are not taken into account in our definition of employment.[15] The information on these two categories of workers is available only for the Public sector both in 2001 and 2011, while it is not available at the local level for the Private sector in 2011. In Sect. 3.3 we nevertheless analyze what happens if we include both categories in the definition of public employment.

The Census data provide information on private and public employment in local units of firms and public institutions. Data are disaggregated at the industry level, with the ATECO 5-digit classification, corresponding to the NACE classification used by Eurostat. However the ISTAT releases for 1991 and 2001 Census data have implemented the ATECO 5-Digit 1991 ISTAT classification, while the release for the 2011 Census has implemented the ATECO 5-Digit 2007 ISTAT classification. We build an algorithm to solve this reclassification issue that properly aggregates the entries at 3-Digit level using the 2002 ISTAT classification. The algorithm aggregates the entries for the three different classifications in order to guarantee that each final (re-aggregated) entry in 1991 is assigned to only one

[14] This entry in particular can only concern the private sector employment.

[15] The relative weight of the number of outworkers and the number of temporary workers has grown from 3.4 to 4.3 percent in the public sector between 2001 and 2011 (http://www.istat.it/it/files/2013/07/06-Scheda-Istituzioni-pubbliche-DEF.pdf, last access: 06/04/2016).

(re-aggregated) entry in the 2007 classification (see the Appendix B for more details). Unfortunately, Census data about employment in local units do not collect information on the skill level of the workforce.

In order to include a set of additional variables and controls, we exploited the information on population and housing Censuses available on *8milacensus* (http://ottomilacensus.istat.it/).[16] The time series on the house price per square meter is, instead, built by using a Bank of Italy Index on the OMI house prices database (see the Appendix B). However since the available OMI's time series starts in 2003, we used prices in 2003 as a proxy for the prices in 2001. As mentioned already, Census data do not provide any information on local wages. To the best of our knowledge there is no available data source on average wages at the municipality level between 2001 and 2011.

In order to exploit an homogeneous set of observations, we selected only those municipalities that exist in all the censuses considered at each specification.

Since we want to avoid the possibility that our results might be heavily influenced by spurious outliers we winsorized the outcome variables at 5% and 95% levels; we censored all the observations below the 5th percentile to the 5th percentile, and all those above the 95th percentile to the 95th percentile. The instruments have been winsorized only at the 1st and 99th percentiles to avoid losing variability, given that there are few outliers.

Descriptive statistics are reported in Table 1. Between 2001 and 2011, employment grew by 0.8% on average across all municipalities (unweighted), but with significant heterogeneity. The private sector contribution has been overall positive (Figure 1), with an average 3.5% increase, while employees in public institutions decreased (-2,6%). The distribution of the public sector contribution confirms that almost all municipalities experienced a contraction during the decade, suggesting that the repeated halts in turnover were effective

[16] We use the time series of data with municipality boundaries fixed at 2001.

in reducing the public sector workforce. The average national changes were negative in all sectors of public employment, therefore it is not surprising that our instrument, $inst_{c,2011}^{publ}$, is negative everywhere. Its standard deviation is significantly smaller than the actual contribution of public employees to employment growth. This is because it is recovered using the (leave-one-out) national growth in each sector, which compensates what happens in the different areas. Figure 2 shows the geographical distribution of $inst_{c,2011}^{publ}$. Larger negative changes take place, as expected, in the South of Italy (which includes the two main islands, Sardinia and Sicily). However, there is variation also within the Centre-North. To address potential issues related to these regional divides, we check our results by splitting the two main areas and, also, by introducing regional fixed effects (see Section 3.3.4).

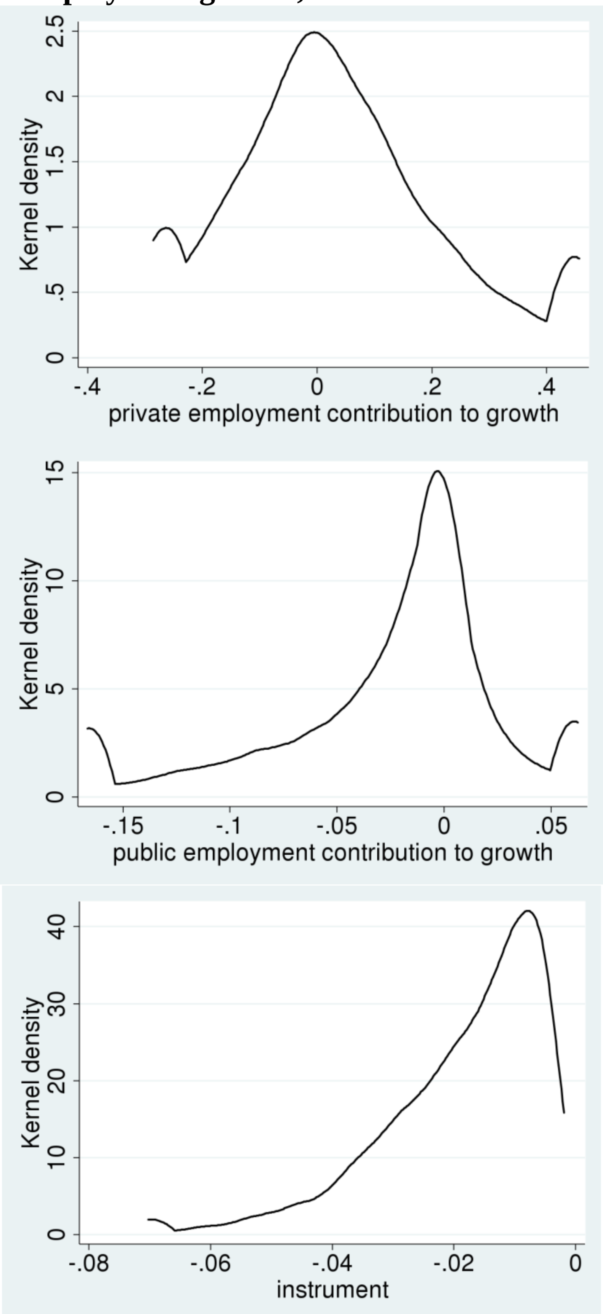
The first panel of Figure 3 shows a positive correlation between variations in private and public employment, although the slope is not very steep. The second panel focuses on the relation between the actual variation in public employment and the predicted one ($inst_{c,t}^{publ}$), which captures the policy rule. The association between averages is quite strong, with a slope of 2.22 (s.e. 0.07, using clustered s.e. at the LLM level), although for each predicted change there is significant dispersion across different municipalities. Given that the standard deviation of the instrument is quite small with respect to the actual variation in public employment, it is useful to standardize the slope to understand the magnitude. An increase by one s.d. in the instrument leads to an increase in the contribution of public employment by 0.55 s.d., which appears to be a reasonable relation.

Finally, the last panel describes the association of the predicted change (the instrument) with the actual variation in private employment. If the instrument has an effect only through its impact on public employment, as argued in Sect. 3.1, then the correlation between the two should reveal the impact of public employment on the private sector. The picture displays a significant and non-negligible negative relation. In terms of s.d., the slope implies that an increase by one s.d. in the predicted public employment contribution (the instrument) leads to a decrease in the contribution of private employment by 0.12 s.d.

Table 1. Descriptive statistics

	mean	p50	sd	min	max	count
private empl contribution to empl growth	0.035	0.016	0.188	-0.286	0.457	8085
public empl contribution to empl growth	-0.026	-0.012	0.055	-0.167	0.063	8085
growth in house prices X sqm	0.284	0.211	0.280	-0.098	0.935	8085
Predicted public empl contrib (instrument)	-0.019	-0.015	0.014	-0.070	-0.002	8085
service contribution to empl growth	-0.050	-0.037	0.095	-0.255	0.126	8085
manufacture contribution to empl growth	0.085	0.069	0.141	-0.161	0.403	8085
population variation (15-64) wrt empl at t-1	0.018	0.027	0.371	-0.796	0.784	8085
private employment (total 2011)	2031	461	15508	1	949956	8085
public employment (total 2011)	351	52	3050	0	203607	8085
house price X sqm (total 2011)	1086	951	604	0	11275	8085
population 15-64 (total 2011)	4787	1581	25589	20	1692869	8085
manufacture private employment (total 2011)	500	108	1698	0	71677	8085
service private employment (total 2011)	1531	303	14067	1	886909	8085
employment (total 2011)	2382	527	18377	4	1153563	8085
unemployment rate (2001)	0.101	0.059	0.088	0.000	0.513	8085
labor force participation rate (2001)	0.474	0.478	0.069	0.167	0.714	8085
fraction of unoccupied houses over total housing in urban areas (2001)	0.254	0.198	0.195	0.000	0.963	8085
mobility index (2001)	0.310	0.317	0.119	0.000	0.639	8085

Figure 1. Density of private, public and predicted public (instrument) contributions to employment growth, 2001-2011

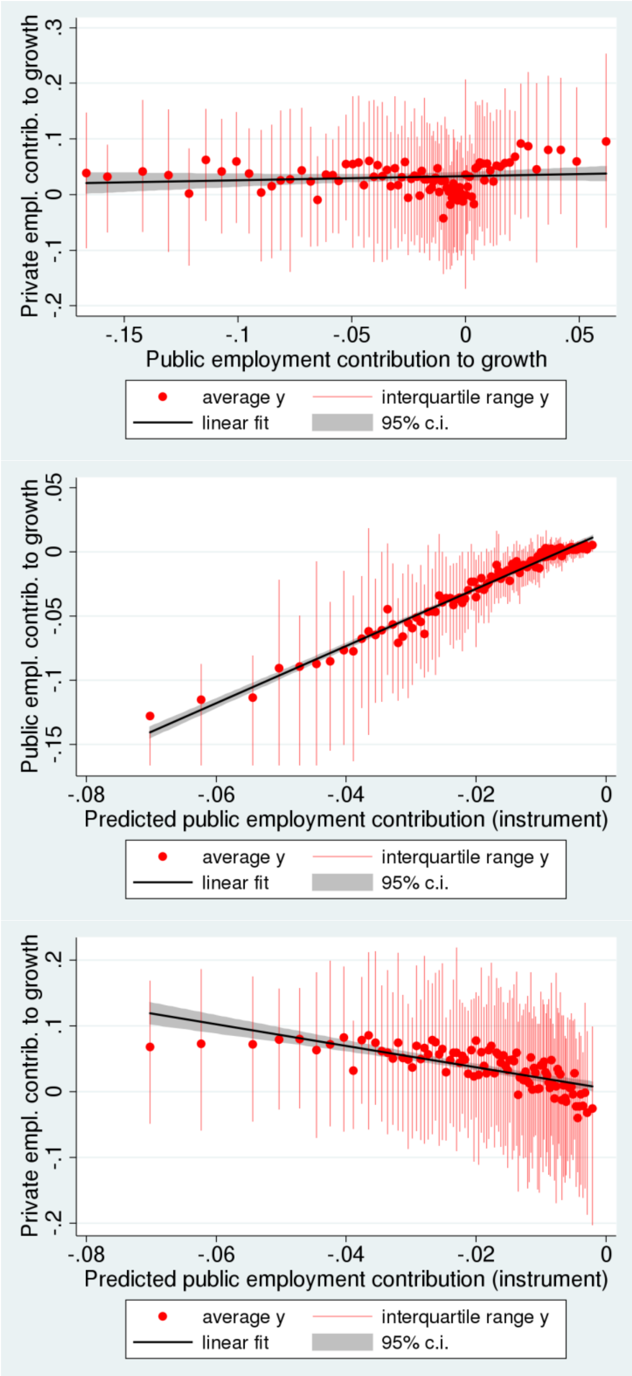


Note: The graphs show densities estimated with a kernel density estimator (Epanechnikov kernel, Silverman's rule-of-thumb bandwidth). See Section 3.1.1 for the definition of the variables.

Figure 2. Distribution of the predicted public contributions to employment growth (instrument) across municipalities, 2001-2011

Note: In the analysis we kept only municipalities that exist in both years. “No data” refers to the excluded ones.

Figure 3. Private and public contributions to employment growth, Italian municipalities, 2001-11



Note: each graph is obtained by splitting the distribution of the x-variable in 100 percentiles and then showing the relation between the average y and the average x in each percentile group (with fitted lines); for each percentile it also shows the interquartile range of the y variable.

3.3 Results.

3.3.1 The impact on private employment. Table 2 reports our baseline estimates for a sample of 8,085 Municipalities, over the period 2001-2011. We start by including only the baseline controls. OLS estimates from panel A (Column 1) suggest that public sector employment contribution has a positive impact on private sector employment. The result is similar when we introduce baseline controls (Column 2), which account for pre-determined difference in activity rate, unemployment, housing supply and commuting, as well as initial levels and past-trends. In column (3) we include all other variables available at the municipality level from the *ottomilacensus* database. These are indices calculated from the two previous waves (1991 and 2001) of the population and housing censuses, and they broadly refer to the demographic structure (age, education, household compositions), housing conditions (housing availability, housing density, buildings age), self-reported occupational status (also distinguished by main sector of activity), commuting and social vulnerability (defined on the basis of household members characteristics and employment status). See the Appendix B for a full list of included variables. When we control for the whole set of controls, the estimate shrinks and becomes less precise, indicating that there may be some positive bias. As we argued, in this context OLS estimates are hardly convincing.

In Panel B we show how our instrument is related to the public sector contribution to employment growth (first stage). The relation is positive and quite strong, as expected given that the national rules led to a significant halt in turnover. The magnitude of the coefficient appears to be large, around 2, but it is hard to interpret it because the scale of the instrument depends on the aggregate variation, which may differ from the average one as it compensates the size of the different municipalities. If we look at the standardized coefficient, the relation appears reasonable: one s.d. increase in predicted contribution of public employment to employment growth ($\mathrm{inst}_{c,2011}^{\mathrm{publ}}$) leads to around half a s.d. increase in the actual contribution. The results are not significantly affected by the introduction of covariates, suggesting that the variation induced by our instrument is not related to other sources of unobserved heterogeneity.

Table 2. The impact of local public sector employment on private employment

	(1) No controls	(2) Baseline controls	(3) Adding all available controls
	Panel A		
	Dependent variable: Private employment (OLS)		
Public empl	0.129***	0.132***	0.082*
	(0.047)	(0.047)	(0.046)
	Panel B		
	Dependent variable: Public employment (OLS - first stage)		
$\mathrm{inst}_{c,2011}^{\mathrm{publ}}$	2.218***	1.999***	1.997***
	(0.069)	(0.099)	(0.094)
Standardized coeff.	0.550	0.493	0.495
F	1046	401	449
	Panel C		
	Dependent variable: Private employment (2SLS - second stage)		
Public empl	-0.748***	-0.625***	-0.765***
	(0.095)	(0.124)	(0.133)
# controls	0	18	125
Observations	8,085	8,085	8,085

Notes: * p-val<0.01, ** p-val<0.05, *** p-val<0.01. The unit of observation is the municipality across 2001-2011. We kept only municipalities that exist in both years. Both public and private employment are expressed as contributions to overall (public+private) employment growth. The standard errors, in brackets, are clustered at the LLM level (2001 definition). We censored the contribution to growth at the 5th and 95th percentiles, while the instrument is censored at the 1st and 99th. The instrument is $\mathrm{inst}_{c,2011}^{\mathrm{publ}}$. For the list of covariates in column (2), see Section 3.1; coefficients are reported in the Appendix B. Column (3) includes also all the other available controls at the municipality level (with no missing values) released by ISTAT in the *ottomilacensus* database, both in year 2001 and 1991 (see the Appendix B). Estimates are produced using the command ivreg2 by Baum et al (2010).

The 2SLS results without including covariates are presented in Panel C (column 1). They suggest that an exogenous increase in public employment brings a substantial displacement of private workers. During the decade that we examine, the actual change involved a contraction of public employment, which therefore caused an increased in private employment. Comparing the results with the OLS estimates, the issues of omitted variables and reverse causality seem to have biased upward the least-square estimates. The most likely explanation is that there are some unobserved shocks which, at the same time, stimulated both private employment and the demand for public services. Simultaneity may have worsened the bias. Results get smaller when including the baseline controls, but become again closer to the one without covariates when we include all available information.

3.3.2 Substantiating the IV approach. One way to assess the plausibility of our estimates is to look at the inverse relation, i.e. how changes in private employment lead to variations in the public workforce. Using a shift-and-share reasoning, we can build an instrument for the private sector contribution to employment growth as:

$$inst_{c,2011}^{priv} = \sum_{j \in priv} \frac{N_{j,c,2001}^{priv}}{N_{c,2001}} \times \frac{N_{j,-c,2011}^{priv} - N_{j,-c,2001}^{priv}}{N_{j,-c,2001}^{priv}}. \qquad (8)$$

Table 3, column (1), shows that this instrument is a good predictor of the contribution of private sector to employment growth. The relation is less strong than in the public sector, but this is in line with previous findings for the local multiplier of the tradable sector reported in de Blasio and Menon (2011). The impact on public employment is estimated using 2SLS in column (2). The coefficient is a precisely estimated zero. On the one hand this is reassuring, because if the shift-and-share approach was inducing a spurious negative correlation between public and private sectors, then we would expect this issue to show up in this regression as well. On the other hand, this suggests that simultaneity is not the main cause of bias in OLS, which is more likely to be driven by unobserved heterogeneity.

Table 3. The impact of local private sector employment on public employment

	(1) Private employment (OLS – First stage)	(2) Public employment (2SLS – second stage)
$\text{inst}^{\text{priv}}_{\text{c,2011}}$	0.443***	
	(0.043)	
Private empl		-0.025
		(0.022)
F	107	
Observations	8,085	8,085

Notes: * p-val<0.01, ** p-val<0.05, *** p-val<0.01. The standard errors, in brackets, are clustered at the LLM level (2001 definition). The instrument, $\text{inst}^{\text{priv}}_{\text{c,2011}}$, is censored at the 1^{st} and 99^{th}. Regressions include all controls as in column (3), Table 2. See Table 2 for other comments.

Another concern is that, as in most settings using instruments derived using a shift-and-share approach, $\text{inst}^{\text{public}}_{\text{c,2011}}$ might be only "plausibly" exogenous (Conley et al, 2012), i.e. correlated, to some extent, with the unobserved heterogeneity. One possible sensitivity check is to include an additional control which is good at predicting the outcome and therefore potentially correlated with the error term, and see whether the coefficient of interest changes.[17] A good candidate is $\text{inst}^{\text{priv}}_{\text{c,2011}}$. Although both $\text{inst}^{\text{publ}}_{\text{c,2011}}$ and $\text{inst}^{\text{priv}}_{\text{c,2011}}$depend on the initial share of public employment, they also vary according to the sectoral composition within public and private employment, hence we can include them together in the regressions.

Considering that the 2SLS estimate is the ratio of the coefficients on $\text{inst}^{\text{publ}}_{\text{c,2011}}$ in the reduced form and in the first stage, in Table 4 we discuss how these two coefficients change separately. This helps to understand what drives

[17] An alternative sensitivity check is the one suggested by Conley et al (2012). They propose to estimate how strong the violation of the exclusion restriction should be for the 2SLS to be entirely driven by it. The violation is captured by the coefficient on $\text{inst}^{\text{publ}}_{\text{c,2011}}$in the main equation, because this coefficients is assumed to be 0 if the exclusion restriction holds. In order to explain the 2SLS estimate, this coefficient should be three times the coefficient on $\text{inst}^{\text{priv}}_{\text{c,2011}}$ from Table 3. If we think in terms of standardized coefficients, it should still be at least 60% of it.

the changes and it allows us to assess the magnitude of the potential bias by exploiting some suggestions from Oster (forthcoming). In columns (1) and (2) we show that $\mathrm{inst}_{c,2011}^{priv}$ does not actually enter into the first stage, as its coefficient is zero. When we add it in the reduced form (column 4), it is instead positive and significant (as in Table 3), as it should be.[18] The coefficient on $\mathrm{inst}_{c,2011}^{publ}$ decreases when both variables are included. The corresponding 2SLS estimate is thus a bit smaller (-0.687) with respect to the main result from Table 2, but not much different.

As argued by Altonji et al (2005) and Oster (forthcoming), looking only at the changes in the coefficient of interest when introducing additional controls is not sufficient to evaluate the potential omitted variable bias. Given that the additional control may be only a proxy of unobserved heterogeneity, we should also look at changes in R2, to understand how much the additional covariate helps in predicting the outcome. In this case the R2 increases by approximately 8%. To assess the sensitivity of the results, Oster (forthcoming) proposes to assume that the maximum R2 that we would obtain if we were to include all the relevant unobserved heterogeneity is 1.3 times the one that we observe including all controls.[19] She then proposes two alternative statistics. One is to estimate a bound for the coefficient on $\mathrm{inst}_{c,2011}^{publ}$ by further assuming that that selection on unobservables is proportional to the selection on the observables (in this case $\mathrm{inst}_{c,2011}^{priv}$only, as we assume that all the other controls are nuisance and therefore we partial them out), with a degree of selectivity equal to one (we refer to her paper for full details). In this case the estimated bound is -0.732, which, if rescaled by the first stage, would give us a value for the 2SLS estimate equal to -0.366. The other statistic is the degree of

[18] In terms of standardized coefficients it is also larger than the one associated to $\mathrm{inst}_{c,2011}^{publ}$.

[19] One could assume that the maximum R2 is equal to one. However, as Oster suggests, this may be overly conservative, as in many settings it is unlikely that one would ever estimate large R2, because idiosyncratic but independent shocks are likely to play a role. This is true also in our setting, where the main equation is already in first differences. For the calculations we used the psacalc ado file written by Oster.

selection on unobservables with respect to the selection on observables that would explain the whole result. If the value is above 1 (equal selection) than the estimate can be considered more robust. In our case the estimated degree of selection is 2.08. These results suggest that, although the exclusion restriction may not be fully respected, our conclusions about a negative impact of public employment are robust to a substantial degree of violation of the assumption.

Table 4. Changes in first stage and reduced form when controlling for $\mathbf{inst}_{\mathbf{c,2011}}^{\mathbf{priv}}$

	(1)	(2)	(3)	(4)
	Public empl (first stage)		Private empl (reduced form)	
	All controls	+ $\mathrm{inst}_{c,2011}^{priv}$	All controls	+ $\mathrm{inst}_{c,2011}^{priv}$
$\mathrm{inst}_{c,2011}^{publ}$	1.997***	2.000***	-1.528***	-1.374***
	(0.094)	(0.094)	(0.261)	(0.263)
- *Std. coeff*	*0.495*	*0.496*	*-0.111*	*-0.099*
$\mathrm{inst}_{c,2011}^{priv}$		0.006		0.431***
		(0.008)		(0.043)
- *Std. coeff*		*0.496*		*0.170*
R2	0.425	0.425	0.182	0.196
Observations	8,085	8,085	8,085	8,085

Notes: * p-val<0.01, ** p-val<0.05, *** p-val<0.01. The standard errors, in brackets, are clustered at the LLM level (2001 definition). Regressions include all controls as in column (3), Table 2. See Table 2 for other comments. "Std. Coeff" is the standardized coefficient relative to the two variables.

3.3.3 Housing prices and other outcomes. Overall, our results indicate that a decrease by one unit of public employment lead to 0.6-0.8 more workers in the private sector. A conservative conclusion, taking into account other possible sources of unobserved heterogeneity, would still confirm that the relation is negative. Interestingly, our conclusion is in line with macroeconomic estimates, although the magnitude of the crowding out is estimated to be lower. Analyzing a panel of OECD countries, Algan et al (2002) suggest a 1.5

displacement effect; Behar and Mok’s (2013) estimates from a panel of 194 countries are around 1.

Our results are, instead, different from studies exploiting variations at the local level, with the exception of the working paper from Senftleben-König (2014) who also find a negative effect for Germany.[20]

One possible explanation for the different results for Italy is the role of the housing market. The theoretical model suggests that, in the absence of an effect of public employment on amenities and productivity, the increase in housing demand leads to a displacement of private workers. Nevertheless, the variation of house prices in equilibrium depends on how many private workers will leave the area to move somewhere else, which is ultimately related to their mobility costs. As predicted by the theory, the effect may be zero or positive. In this respect, Jofre-Monseny et al (2016) also point out that, in their model, crowding out occurs when the land (and housing) supply is very inelastic. This may justify the fact that they find a positive effect for Spain, given that the cities driving their results “are relatively small provincial capitals which can be considered to be cities with a rather elastic land supply” (idem, pg. 29). Differently, Faggio and Overman (2014, p.103) find basically no overall impact in Britain, where strong regulations in the housing market are likely to hamper the multiplier effect associated with an increase in public employment (pg. 103). For Italy, there is evidence that the strong rigidity in the housing market hampers the ability of local economies to adjust to external shocks (see Ciani et al, 2017, and Accetturo et al, 2017). Unfortunately, none of the studies, including those from relocation episodes, provide estimates about the effect on local house prices to provide a full comparison.

The OLS regression for house prices without controls, in Table 5, panel A, column (1), actually displays a negative relation. This, however, seems to be driven by other differences. Including controls (columns 2 and 3),

[20] The paper by Ranzani and Tuccio (2017) also finds a negative impact in Ghana, Mali and Mozambique, but it is entirely driven by a contraction in agricultural workers, which is much less relevant in Italy.

the coefficient turns out to be positive, although small and not statistically significant at standard levels. It is indeed likely that the trends in house prices are affected by different amenities and past trends. Furthermore, it must be pointed out that our results are average values at the city level as provided by the *Osservatorio del mercato immobiliare*. Although the averages account for some broader categories and status of the building, they are not derived from an hedonic price regressions on microdata. Therefore results are more likely to depend on how well we control for potential confounders.

In line with the OLS results, the IV regression without controls is negative. When including all available covariates it turns out to be positive and statistically significant. This suggests that public employment indeed creates pressure on the local housing market, therefore contributing to the crowding out. In the decade that we study, the contraction of public employment has led to the opposite situation, reducing local prices and, therefore, stimulating local labor supply to the private sector.

Table 5. The impact of local public sector employment on house prices

	(1) No controls	(2) Baseline controls	(3) Adding all available controls
Panel A. Dependent variable: House prices (OLS)			
Public employment	-0.364***	0.062	0.075
	(0.101)	(0.086)	(0.068)
Panel B. Dependent variable: House prices (2SLS - second stage)			
Public employment	-1.429***	0.003	0.505**
	(0.253)	(0.281)	(0.205)
# controls	0	18	125
Observations	8,085	8,085	8,085

Notes: * p-val<0.01, ** p-val<0.05, *** p-val<0.01. The dependent variable for house prices is a growth rate ($[pr_{2011} - pr_{2003}]/pr_{2003}$), and it is censored at the 5th and 95th percentiles. See Table 2 for other comments

Coming back to local employment, most of the papers find that public employment has a positive impact on private employment in the non-tradable sector, but a negative impact on employment in manufacturing. In Table 6, columns (1) and (2) we split private employment into different components. Although the model we sketch here does not distinguish between tradable and non-tradable sectors, this distinction could be relevant.[21] In particular, as in Jofre-Monseny et al (2016), we might consider a basket of non-tradables (which encompasses housing services) which is produced using also a local production factor (e.g. land). Then, the increased demand due to the expansion of the public sector would not be limited to the housing market but also to other non-tradables.

Table 6. The impact of local public sector employment on other outcomes (2SLS regressions)

	(1) Manufacture private employment	(2) Service and construction private employment	(3) Population
Public employment	-0.586***	-0.169	0.903***
	(0.064)	(0.103)	(0.193)
Observations	8,085	8,085	8,085

Notes: *** p-val<0.01. All the outcomes are expressed as contribution with respect to the overall (public+private) employment growth and they are censored at the 5th and 95th percentiles. Standard errors, in brackets, are clustered at the LLM level (2001 definition). The manufacture sector includes the extraction of natural resources. The service sector includes construction, as in Faggio and Overman (2014). Population refers only to the working age population (aged 15-64). The instrument is $inst_{c,2011}$; see Table 2 for more details and for the first stage statistics. The regressions include all controls as in column (3) from Table 2.

In the empirical estimates by sector, the dependent variable is always the contribution of the variation in that specific sector to the overall employment growth (the variation between 2001 and 2011 divided by employment in 2001). The effect on local manufacture (Column 1) is negative and significant. The impact on service and construction (Column 2), is not statistically significant and quite smaller, though still negative. This is in line

[21] In the additional Appendix available on our websites (Auricchio et al, 2016) the interested reader can find an extension of our model which considers both tradables and non-tradables.

with previous empirical work showing that the impact of exogenous increases in local activity percolates mostly on non-tradables (see, for instance, Moretti, 2010). The key difference with Faggio and Overman (2014) and Jofre-Monsey (2016) is that we do not find any evidence of a positive effect on non-tradables. One possible reason is that private employment, in Italy, tends to display small multiplier effects. De Blasio and Menon (2011) show that the multiplier effects of the tradable sector on non-tradables is very small in Italy.

The difference between our results and other studies can also be due to the presence of non-linearities, affecting the impact of experiments which are quite varied. For instance, in Jofre-Monseny et al, public employment grew by 133 per cent in the period that they considered, while in Faggio and Overman the average public contribution to employment growth was slightly less than 6 per cent. The studies based on relocation episodes (Faggio,2015; Becker et al, 2015; Faggio et al, 2016) focus only on the receiving area, where public employment grew, while they do not study what happened in the areas from where they were previously located. Only Senftleben-König (2014) estimates are based on a period when, on average, public employment contracted. It might be possible that contractions lead to different results. With respect to the housing market channel, as discussed by Notowidigdo (2011), housing supply is likely to be rigid at the current level of housing stock, as it cannot be destroyed. A negative shock would therefore have a stronger impact on house prices, favoring an increase in private employment. With respect to the absence of a multiplier effect in the non-tradable sector, one possibility is that when public employment contracts, local enterprises that prospered from this source of demand are relatively slow to exit the market.

There is also the possibility that, in the expectation of new public openings, individuals wait for a public job, instead of applying for a private sector vacancy. In our model, the crowding out of the local private sector is generated by individuals leaving the local area. Given that the negative impact on local private employment is around 0.7 units, we expect a positive net impact on local working-age population. One additional public employee should increase population by around 0.3 individuals. In column (3) of Table 6 we focus on the changes in working-age population. Despite of the displacement effect, an increase in public employment seems to increase working-age population by more than we would expect. This implies that the non-employed (working age) population increases, suggesting that mechanisms other than migration may be contributing to our findings. As shown in Giordano (2010) and Depalo et al (2015), wages in the public sector are larger. Furthermore, jobs in the public sector are generally perceived as more stable. This may induce an increase in the population given that earnings are used to support unemployed family members (see also Boeri et al, 2014). Finally, an increase in public employment may increase the propensity of individuals to stay unemployed longer without leaving the area, as suggested by the aforementioned research, and by Burdett (2012) and Gomes (2014) in a search framework. Calibrating a macro-model with two regions and a public sector, Caponi (2017) argues that higher public employment is used to prevent out-migration. Indeed, our results indicate that, in the decade that we analyze, the contraction in public employment led to an outflow of population from those places were the decrease was larger.

Table 7. The impact of local public sector employment on private employment and house prices, including also the variation in institutional quality indices between 2004 and 2011 (2SLS regressions)

	(1) Private employment	(2) House price
Public employment	-0.780***	0.422**
	(0.134)	(0.201)
Obs	8,085	8,085
First stage F	457	457

Note: * p-val<0.10, *** p-val<0.01. Standard errors, in brackets, are clustered at the LLM level (2001 definition). The instrument is $inst_{c,2011}$. The regressions include all controls as in panel C from Table 2, plus the variation at the provincial level in three institutional quality indices proposed by Nifo and Vecchione (2014): Government effectiveness, Rule of law and Corruption. We defer to their paper for a detailed discussion of the variables.

Finally, our results may be driven by changes in local amenities or productivity induced by changes in public employment. For instance, an increase in public employment may enhance the services available to both households and firms, thus limiting the negative impact on the private sector. By contrast, if the increase is designed only as a redistributive measure as suggested by Alesina et al. (2000), it may even reduce overall efficiency by raising the amount of red-tape. The estimates presented so far capture both the impact of rising costs of local non-tradables, and the impact from potential changes in local productivity and amenities. Unfortunately, we do not have good proxies for such features at the municipality level. Nevertheless, Nifo and Vecchione (2014) propose an institutional quality index at the provincial level for each year from 2004 to 2012. Three of its elements are likely to capture possible changes in amenities and productivity: (i) Government effectiveness, which accounts for the endowment of economic and social facilities; (ii) Rule of law, which captures the level of crime against property and justice effectiveness; (iii) Corruption. Even if the variables are defined at the provincial level, we expect that their variation between 2004 and 2011 is able to capture significant changes in the local economic environment that have been taking place contemporaneously with our changes in public employment. If the impact on private employment and house prices is partially due to such changes, then including them together with the other controls may lead to different results. Table 7 shows that the estimates of interest are only marginally affected. Our results do not, therefore, seem to be driven by changes in local amenities or productivity induced by changes in public employment.

3.3.4 Other robustness checks. As suggested by the model, results strongly depend on the mobility of the population. Centre-North vs South differences may play a crucial role.[22] Table 8 provides the results obtained by splitting the sample between Centre-North and South. We find that displacement affects both areas, although the effect is larger in the South (see also Alesina et al 2001). The effect on house prices is larger in the South and closer to zero in the other areas. In a heavily subsidized area like the South of Italy, the contraction of the public sector releases more resources for the private initiative; at the same time, the negative impact on the local housing market is deeper.

Table 8. The impact of local public sector employment on private employment and house prices, by area (2SLS regressions)

	Centre-North		South	
	(1)	(2)	(3)	(4)
	Private employment	House price	Private employment	House price
Public employment	-0.429**	0.100	-1.063***	0.643***
	(0.172)	(0.273)	(0.173)	(0.249)
obs	5528	5528	2557	2557
First stage F	251	251	271	271

Note: * p-val<0.10, *** p-val<0.01. Standard errors, in brackets, are clustered at the LLM level (2001 definition). Censoring of dependent variables and instruments is done at the national level. The instrument $inst_{c,2011}$; see Table 2 for more details and for the first stage statistics. The regressions include all controls as in panel C from Table 2. The South of Italy includes the two main islands (Sicily and Sardinia).

To address concerns about the inclusion of "too many" additional covariates, which may lead to imprecise estimates, in Table 9, Panel A, we select only the most relevant ones by following the "double selection" procedure proposed by Belloni et al (2014). In detail, we use a LASSO algorithm to select those additional covariates (except for the baseline ones) that help explaining the variability of the outcomes, the endogenous variable (contribution of public sector to employment growth) and the instrument. For

[22] Furthermore, Southern regions were less likely to meet the budgetary targets set by the central government and were therefore strongly required to enforce the stop in turnover. Indeed the power of our instrument is higher in southern municipalities.

each outcome, the set of selected variables is the union of the different selections, and therefore the set may be different according to the outcome. In practice, we assume that, among the whole set of covariates, only some have non-zero coefficient and use a penalized criterion (LASSO) to select them, where the penalization is based on the number of selected items. We start from the entire list of additional variables (standardized to have zero mean and unitary variance), and apply the algorithm proposed by Belloni et al (2014).[23] The final estimate for the effect of interest is obtained by running standard 2SLS, including only the selected covariates. The list of selected covariates, which is different for the two outcomes, is reported in the Appendix B. Estimates are similar and in line with the conclusions obtained when including the entire set of covariates.

In Panel B we reconsider our choice of censoring the outcomes and the contribution of public employment at the 5th and 95th percentiles, by changing it to the 1st and 99th percentiles. Results are in line with our main estimates, even if slightly smaller.

Panel C instead illustrates the results obtained by using the Faggio and Overman's (2014) instrument. The results are very similar to those depicted so far. Note also that the first-stage power of the original instrument is actually stronger. If the differences between the sectors are not so relevant, then their instrument may actually be more precise, as it avoids the measurement error introduced by first calculating the prediction at the sectoral level and then aggregating for the overall instrument.

A different issue concerns the geographical distribution of the predicted cuts in public employment, which are more concentrated in Southern regions. In Panel D we include regional fixed effects, which are aimed at capturing regional trends. Results are, again, quite similar to the main ones.

[23] We use the Stata code made available by the authors. The controls suggested by the theory are included as non-delectable controls and have been standardized as well.

Table 9. The impact of local public sector employment on private employment and house prices; robustness checks

	(1) Private employment	(2) House price
Panel A: "Double selection" of the additional controls		
Public employment	-0.626***	0.584***
	(0.134)	(0.210)
First stage F	401	428
Obs	8,085	8,085
Panel B: Censoring only at the 1th and 99th percentile		
Public employment	-0.645***	0.433***
	(0.124)	(0.164)
First stage F	360	360
Obs	8,085	8,085
Panel C: Using Faggio and Overman's instrument		
Public employment	-0.843***	0.542***
	(0.119)	(0.204)
First stage F	627	627
Obs	8,085	8,085
Panel D: Regional fixed effects		
Public employment	-0.776***	0.517***
	(0.128)	(0.169)
First stage F	533	533
Obs	8,085	8,085
Panel E: Estimates at the LLM level		
Public employment	-0.966***	1.440
	(0.445)	(1.123)
First stage F	34	34
Obs	686	686

Note: ** p-val<0.05, *** p-val<0.01. Standard errors, in brackets, are clustered at the LLM level (2001 definition). The estimates in panel A are obtained by 2SLS, including all covariates suggested by the theory (column 2) plus an additional set of covariates from the *ottomilacensus* database, selected following the procedure suggested by Belloni et al (2014) and the ado program written by them. All covariates have been standardized before running the selection (see the Appendix B for a list of selected covariates). The algorithm converges in few iterations. The regressions in the other panels include all controls as in column (3) from Table 2. In the estimates at the LLM level, the average of controls is across municipalities and it is weighted by population size. See Table 2 for other details.

Panel E shifts the unit of analysis from the municipality to the LLM. We focus on the specification with all controls, though we do not include the average commuting index, as is not relevant for this analysis. Displacement is still confirmed, with a larger coefficient, but anyway close to the one estimated at the municipality level. The effect on house prices is still positive, but now largely imprecise. The noisy estimate may be due to the fact that we had to aggregate house prices at the LLM using population as weight, which may lead to a poor proxy of the actual variable. These results confirm qualitatively the main conclusions. Most importantly, they do not lend support to the possibility that all the effects at the municipality level are simply driven by relocations to/from nearby municipalities. We also tried a different strategy, randomly selecting only one municipality from each LLM. This should limit the possibility of localized spill-overs. We repeated the estimation 999 times and we averaged the estimated coefficients. Results are in line with the regressions at the LLM level, with an average displacement effect of –.959 and an impact on house prices of 0.527.

One additional issue concerns the fact that, during the last decades, several activities run by public services have gone through a process of privatization, in particular, utilities, transportation, telecommunication, postal services, and waste management. Our definition of public employment does not include these sectors. As discussed in Section 3.1, firms in these sectors are considered private as long as they produce market goods or services. Nevertheless, employment in heavily regulated sectors may depend quite tightly on public employment decisions. In column (1), Table 10, we redefine private sectors to exclude sectors that, in Italy, are strongly regulated or are characterized by a strong presence of state controlled firms and investments. In particular, we exclude recycling, utilities and electricity production, transportation and related services, telecommunication, postal services, monetary intermediation, education, health care, waste disposal, private associations, cultural services, family services. Results are quite similar to the main ones, which indicates that the effect is driven by the less regulated sector. When we focus on the heavily regulated one (column 2), there is no crowding-out. In our definition of public employment, the sector where concerns about liberalization may anyway show up is the one relative to health care, residential care and the residual category, which includes a bunch of minor activities that are not the core business of Public administration. In Column (3) we unbundle public employment to single out these sectors. We define two separate instruments based on the two different components. The first stage is still strong for both variables. Although it is true that stronger crowding-out is found in the component where we expect results to be partially driven by a contraction in public services and an increase in private intervention, the estimate for the other is quite close to our main results.

Table 10. The impact of local public sector employment on private employment, 2SLS

	(1) Private empl in less regulated sectors	(2) Private empl in more regulated sectors	(3) Private empl
Public employment	-0.721***	-0.025	
	(0.112)	(0.038)	
Public empl without health, residential care and residual cat.			-0.778***
			(0.126)
Public empl in health, residential care and residual cat.			-1.281***
			(0.337)
H0: difference between two sectors=0 (p-val)			0.093
First stage F	449	449	253
Observations	8,085	8,085	8,085

Notes: * p-val<0.01, ** p-val<0.05, *** p-val<0.01. The standard errors, in brackets, are clustered at the LLM level (2001 definition). The regressions in the other panels include all controls as in column (3) from Table 2. The less regulated private sector in column (1) excludes recycling, utilities and electricity production, transportation and related services, telecommunication, postal services, monetary intermediation, education, health care, waste disposal, private associations, cultural services, family services. In column (3) the employment contribution to growth of the two sectors has been instrumented by two instruments designed in the same way as $inst_{c,2011}$ but limited to the specific sector. The relative First stage F is the “Kleibergen-Paap rk Wald F statistic”.

Our empirical strategy is justified by the way public sector was reduced over the period 2001-11 through national interventions that reduced turnover. Furthermore, we have information on house prices only for the most recent years. Both reasons led us to focus only on that decade. Nevertheless, we can make use of the previous decade to understand whether our findings are driven only by the particular choice of the instrument, by focusing on an alternative identification strategy where municipality fixed effects are included in order to address time-invariant unobserved heterogeneity. In Table 11 we use the whole 1991-2011 span. We do not add controls because some of them are debatable in a longitudinal regression (initial levels in particular) and, furthermore, not all of them are available in both decades (in particular the double lagged ones). Results using pooled OLS are positive and coherent with what we estimate in the last decade (Table 2). When we introduce municipality fixed effects the relation between public and private employment turns out to be negative and statistically significant. Although the magnitude is smaller than the relative IV estimates, this confirms that the bias in OLS is positive and supports our conclusion of crowding-out.

Table 11. The impact of local public sector employment on private employment; 1991-2011

Dependent variable: Private employment	(1) Pooled OLS	(2) Municipality FE
Public empl	0.278***	-0.157***
	(0.041)	(0.042)
Observations	16,170	16,170

Notes: * p-val<0.01, ** p-val<0.05, *** p-val<0.01. The standard errors, in brackets, are clustered at the LLM level (2001 definition). The regressions include no covariates.

We also checked whether results are different if we include collaborators and temporary workers (other than those on standard fixed-term contracts, who are already included) in the definition of public employment. Point estimates, available on request, are only slightly affected. Finally, we tried to exclude municipalities with more than 100 thousand inhabitants, like

Rome, which has a large share of total public employment. Results are basically unaffected.[24]

4. Conclusions

We proposed a spatial equilibrium model to discuss how changes in public employment affect private employment and house prices. In the absence of any effect of public employment on amenities or factors productivity, the interaction between private and public employees is mainly due to competition in the local market for housing. An increase in local public employment exerts pressure on the local demand for housing. Depending on their preferences for the location, some private sector workers may decide to leave when facing a higher local cost of living, while the others will stay despite the drop in their real income.

Our empirical analysis of decadal changes in public employment in Italian municipalities confirms the importance of this channel. We find a marked crowding out of private employment and a positive impact on house prices. The increase in house prices suggests that mobility costs (or idiosyncratic location preferences) we embedded in the theoretical model are important even in the medium run, as captured by decadal changes. Finally, our estimates do not seem to be driven by changes in the local economic environment.

These results are particularly useful in assessing the economic consequences of employment policies in the public sector. Local contractions in the public workforce, as in the case of the recent policies that led (and are still leading) to a decrease in turnover, seem to have induced an increase in private sector jobs, although they led to an outflow of population. As long as the effects are symmetrical for cuts and increases in public employment, our

[24] To address the potential concentration in some areas, we also tried to drop the bottom and top deciles according to the population in 2001. Results are similar, although the one on house prices loses statistical significance.

estimates can also be used to evaluate the opposite situation. An expansion in public employment may be used to redistribute resources to laggard regions, so to limit the outflow of local population, but the beneficial effect on local economies is eroded by the negative impact on the private sector and by the increase in rents. This is even more important if we consider that private firms play a crucial role for the sustainability of growth in the longer run, in particular those in the tradable sector, for which crowding out appears to be even more consistent.

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Appendices

Appendix A: A Spatial Model with Public Employment

As mentioned in the main text, the model builds on Roback (1982) and exploits the notion of "mobility costs" (see, e.g., Moretti 2011). The economy is divided into two regions, $\{a,b\}$, possibly characterized by different amenities. All firms use skilled and unskilled labor to produce a *tradable* good. While firms are assumed to be fully mobile across regions, workers are subject to idiosyncratic preference shocks for each location. Such shocks generate "mobility costs" across areas which, in contrast with Roback's original framework, make the local labor supply not perfectly elastic to local real wages. Residential supply in each area may depend on local rent levels, and landowners are absentee.

We now come to the central theme, local public employment. Public employees can be skilled and unskilled. Skilled public employment in regions *a* and *b* is equal, respectively, to $(\hat{N}_a, \hat{N}_b)$, with $\hat{N} \equiv \hat{N}_a + \hat{N}_b$. Similarly, unskilled public employment in regions *a* and *b* is equal, respectively, to $(\hat{n}_a, \hat{n}_b)$, with $\hat{n} \equiv \hat{n}_a + \hat{n}_b$ The size and allocation of public employment across regions is exogenously determined by the public administration. We also postulate that the wages for (skilled and unskilled) public employees are equal across regions, $(\overline{w}^s, \overline{w}^u)$ and set at the national level. We also assume that such wages are not smaller than the corresponding levels in the private sector.

We now describe the fundamentals of the model, starting with individual preferences.

Preferences.

A *skilled* worker in area $c = \{a,b\}$ has Cobb-Douglas utility (see, e.g., Diamond, 2016)

$$U_c^s = \ln(A_c^s) + (1-\gamma)\cdot\ln(x_c) + \gamma\cdot\ln(L_c) + \varepsilon_c^s \qquad (1)$$

with $\gamma \in (0,1)$, which is maximized under the budget constraint $w_c^s = x_c + r_c \cdot L_c$. The term A_c^s denotes local amenities that are particularly attractive to educated individuals, while $\{x_c, L_c\}$ denote, respectively, the consumption of the tradable good (of price equal to one, the *numeraire*), and the consumption of housing services of price equal to r_c. The preference shock for location c

is denoted by ε_c^s. We assume that each ε_c^s, $c = \{a, b\}$, is an i.i.d. Type I Extreme Value (Gumbel) with scale parameter equal to $\rho^s \geq 0$ (see, e.g., Moretti and Kline, 2014). The parameter ρ^s drives the strength of individual preferences for location and, thus, the degree of labor mobility across areas. The closer ρ^s to zero, the more sensitive are skilled workers to differences in local prices and amenities.

Utility maximization by skilled individuals delivers the following indirect utility function:

$$V_c^s = \ln\eta + \ln\left(A_c{}^s\right) + \ln\left(w_c^s\right) - \gamma \cdot \ln\left(r_c\right) + \varepsilon_c^s \equiv \ln\eta + v_c^s + \varepsilon_c^s, \qquad c = \{a, b\} \qquad (2)$$

where η is a positive constant, and $v_c^s \equiv \ln\left(A_c{}^s\right) + \ln\left(w_c^s\right) - \gamma \cdot \ln\left(r_c\right)$.

Location a will be preferred to location b when it holds that $v_a^s + \varepsilon_a^s \geq v_b^s + \varepsilon_b^s$, that is, when inequality $\varepsilon_a^s - \varepsilon_b^s \geq v_b^s - v_a^s$ holds true. Since the difference, $\varepsilon_a^s - \varepsilon_b^s$, between two independent Gumbel-distributed random variables has a Logistic distribution with zero mean and CDF given by[a] $\Lambda(x) = \dfrac{\exp\left\{\dfrac{x}{\rho^s}\right\}}{1 + \exp\left\{\dfrac{x}{\rho^s}\right\}}$, the fraction of people living in location b, defined as $\dfrac{N_b}{N_a + N_b}$, will be given by $\Lambda\left(v_b^s - v_a^s\right)$. It follows that $\dfrac{N_b}{N_a + N_b} = \dfrac{\exp\left\{\dfrac{v_b^s - v_a^s}{\rho^s}\right\}}{1 + \exp\left\{\dfrac{v_b^s - v_a^s}{\rho^s}\right\}}$, which can be re-written as $\dfrac{N_b}{N_a} = \exp\left\{\dfrac{v_b^s - v_a^s}{\rho^s}\right\}$. By taking logs and expressing $v_b^s - v_a^s$ in terms of local prices and amenities, it holds that:

$$\rho^s \cdot \ln\left(\frac{N_b}{N_a}\right) = \ln\left(\frac{A_b^s}{A_a^s}\right) + \ln\left(\frac{w_b^s}{w_a^s}\right) - \gamma \cdot \ln\left(\frac{r_b}{r_a}\right), \qquad (3)$$

As in Moretti (2011) and Moretti and Kline (2014), equation (3) represents skilled labor supply in area *b*'s private sector relative to area *a*, an increasing function of skilled wages in area *b*,

[a] See, e.g., Anderson et al (p.60, 1992).

relative to skilled wages in area *a*. The elasticity of local employment to relative local wages is given by $1/\rho^s$.

Similar expressions hold for *unskilled* workers, who maximize utility $U_c^u = \ln\left(A_c^u\right) + (1-\gamma)\cdot\ln\left(x_c\right) + \gamma\cdot\ln\left(L_c\right) + \varepsilon_c^u$ subject to the budget constraint $w_c^u = x_c + r_c \cdot L_c$. Again, we assume that each ε_c^u , with $c = \{a, b\}$, is an i.i.d. Gumbel with scale parameter $\rho^u \geq 0$. Recall that, when it holds that $\rho^s = \rho^u = 0$, workers are fully mobile across areas, and the local labor supply becomes infinitely elastic to local real wages. Following the same procedure adopted above, we obtain the relative labor supply for unskilled individuals:

$$\rho^u \cdot \ln\left(\frac{n_b}{n_a}\right) = \ln\left(\frac{A_b^u}{A_a^u}\right) + \ln\left(\frac{w_b^u}{w_a^u}\right) - \gamma \cdot \ln\left(\frac{r_b}{r_a}\right), \qquad (4)$$

Technology.

In each area, there are competitive firms which produce a tradable good under constant returns to scale by using skilled and unskilled labor. Local producers of the tradable good use the following technology:

$$X_c = Q_c \cdot \left(N_c\right)^{\theta} \cdot \left(n_c\right)^{1-\theta} \qquad (5)$$

where $\theta \in (0,1)$; Q_c is a local productivity shifter, and (N_c, n_c) are, respectively, the skilled and unskilled labor inputs. Recalling that the economy-wide price for tradables is one, the set of the first-order conditions for profit maximization in the tradable sector can be written as:

$$\theta \cdot X_c = w_c^s \cdot N_c; \quad (1-\theta)\cdot X_c = w_c^u \cdot n_c \qquad (6)$$

Housing.

The model is closed by the equilibrium condition for the local housing market. The local supply of housing services is equal to local demand.[b] Since demands for housing by skilled and

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[b] This is *not* necessarily the case on the local market for tradables, which are supplied in any amount at the economy-wide price of one. Thus, the local "trade balance" need not be zero. Such issues are very common in the open economies literature, where economic policies have different impact across sectors, or where fiscal policies aimed at redistribution reduce country's competitiveness: see Alesina and Perotti (1997).

unskilled individuals are, respectively, equal to $L = \gamma \cdot \frac{w_c^s}{r_c}$ and $L = \gamma \cdot \frac{w_c^u}{r_c}$, the local market clearing condition for housing services can be written as:

$$\bar{\ell}_c \cdot (r_c)^{\delta} = \gamma \cdot \left[\frac{w_c^s}{r_c} \cdot N_c + \frac{w_c^u}{r_c} \cdot n_c + \frac{\overline{w}^s}{r_c} \cdot \hat{N}_c + \frac{\overline{w}^s}{r_c} \cdot \hat{n}_c \right]. \qquad (7)$$

The left-hand side of (7) represents local housing supply, postulated as an increasing function of residential land availability, denoted by $\bar{\ell}_c$, and local rents, with $\delta \geq 0$. In other words, we assume that local housing supply increases with the level of rents whenever δ is strictly larger than zero. On the right-hand side, $\hat{N}_c$ and $\hat{n}_c$ represent, respectively, the units of skilled and unskilled employees in the local public sector, while N_c and n_c denote the units of skilled and unskilled employees in the local private sector.

In what follows, the expression $\Pi_c \equiv \overline{w}^s \cdot \hat{N}_c + \overline{w}^u \cdot \hat{n}_c$, with $c = \{a, b\}$, will denote the local wage-bill for public employment. Moreover, by using expressions (6), we can rewrite (7) as

$$\bar{\ell}_c \cdot (r_c)^{1+\delta} - \gamma \cdot \Pi_c = \gamma \cdot X_c. \qquad (8)$$

By taking logs of (8) and subtracting the expression relative to location *a* from the corresponding one for location *b*, we obtain:

$$\ln\left(\frac{\bar{\ell}_b \cdot (r_b)^{1+\delta} - \gamma \cdot \Pi_b}{\bar{\ell}_a \cdot (r_a)^{1+\delta} - \gamma \cdot \Pi_a} \right) = \ln\left(\frac{X_b}{X_a} \right) = \ln\left(\frac{Q_b}{Q_a} \right) + \theta \cdot \ln\left(\frac{N_b}{N_a} \right) + (1-\theta) \cdot \ln\left(\frac{n_b}{n_a} \right). \qquad (9)$$

Profit maximization in the tradables sector implies that:

$$\ln\left(\frac{w_b^s}{w_a^s} \right) = \ln\left(\frac{Q_b}{Q_a} \right) - (1-\theta) \cdot \ln\left(\frac{N_b}{N_a} \right) + (1-\theta) \cdot \ln\left(\frac{n_b}{n_a} \right). \qquad (10)$$

and

$$\ln\left(\frac{w_b^u}{w_a^u} \right) = \ln\left(\frac{Q_b}{Q_a} \right) + \theta \cdot \ln\left(\frac{N_b}{N_a} \right) - \theta \cdot \ln\left(\frac{n_b}{n_a} \right). \qquad (11)$$

Thus, equation (9) can be re-written as:

$$\ln\left(\frac{\bar{\ell}_b \cdot (r_b)^{1+\delta} - \gamma \cdot \Pi_b}{\bar{\ell}_a \cdot (r_a)^{1+\delta} - \gamma \cdot \Pi_a}\right) = \ln\left(\frac{w_b^u}{w_a^u}\right) + \ln\left(\frac{n_b}{n_a}\right) = \ln\left(\frac{w_b^s}{w_a^s}\right) + \ln\left(\frac{N_b}{N_a}\right). \tag{12}$$

Expressions (3), (4), (10), (11) and (12) constitute a system of five equations which can be differentiated and evaluated around *symmetry*, that is, evaluated for the case when the two locations are initially identical (such that it holds that $x_a = x_b = x$ for every variable in the system). In what follows, we will denote the *average* number of skilled and unskilled workers across locations respectively as $\{N, n\}$, so that $2N = N_a + N_b$, and $2n = n_a + n_b$. Total differentiation yields:

$$\rho^s \cdot \left[\frac{dN_b - dN_a}{N}\right] = \left(\frac{dA_b^s - dA_a^s}{A^s}\right) + \left(\frac{dw_b^s - dw_a^s}{w^s}\right) - \gamma \cdot \left(\frac{dr_b - dr_a}{r}\right), \tag{13}$$

$$\rho^u \cdot \left[\frac{dn_b - dn_a}{n}\right] = \left(\frac{dA_b^u - dA_a^u}{A^u}\right) + \left(\frac{dw_b^u - dw_a^u}{w^u}\right) - \gamma \cdot \left(\frac{dr_b - dr_a}{r}\right), \tag{14}$$

$$\left(\frac{dw_b^s - dw_a^s}{w^s}\right) = \left(\frac{dQ_b - dQ_a}{Q}\right) - (1-\theta) \cdot \left(\frac{dN_b - dN_a}{N}\right) + (1-\theta) \cdot \left(\frac{dn_b - dn_a}{n}\right). \tag{15}$$

$$\left(\frac{dw_b^u - dw_a^u}{w^u}\right) = \left(\frac{dQ_b - dQ_a}{Q}\right) + \theta \cdot \left(\frac{dN_b - dN_a}{N}\right) - \theta \cdot \left(\frac{dn_b - dn_a}{n}\right). \tag{16}$$

$$\begin{aligned} \left(\frac{dr_b - dr_a}{r}\right) &= H \cdot \left[\left(\frac{dw_b^s - dw_a^s}{w^s}\right) + \left(\frac{dN_b - dN_a}{N}\right)\right] + H' \cdot \left(\frac{d\Pi_b - d\Pi_a}{\Pi}\right) = \\ &= H \cdot \left[\left(\frac{dw_b^u - dw_a^u}{w^u}\right) + \left(\frac{dn_b - dn_a}{n}\right)\right] + H' \cdot \left(\frac{d\Pi_b - d\Pi_a}{\Pi}\right) \end{aligned} \tag{17}$$

where $H \equiv \dfrac{\bar{\ell} \cdot (r)^{1+\delta} - \gamma \cdot \Pi}{(1+\delta) \cdot \bar{\ell} \cdot (r)^{1+\delta}} \in (0{,}1)$, and $H' \equiv \dfrac{\gamma \cdot \Pi}{(1+\delta) \cdot \bar{\ell} \cdot (r)^{1+\delta}} \in (0{,}1)$.

By using the notation $\tilde{x} \equiv \dfrac{dx_b - dx_a}{x}$, such that $\tilde{x}$ denotes the difference in percentage change between x in area *b* and x in area *a*, the system from (13) to (17) can be written as follows:

$$\rho^s \cdot \tilde{N} = \tilde{A}^s + \tilde{w}^s - \gamma \cdot \tilde{r} \tag{18}$$

$$\rho^u \cdot \tilde{n} = \tilde{A}^u + \tilde{w}^u - \gamma \cdot \tilde{r} \qquad (19)$$

$$\tilde{w}^s = \tilde{Q} - (1-\theta) \cdot \tilde{N} + (1-\theta) \cdot \tilde{n} \qquad (20)$$

$$\tilde{w}^u = \tilde{Q} + \theta \cdot \tilde{N} - \theta \cdot \tilde{n} \qquad (21)$$

$$\tilde{r} = H \cdot \left(\tilde{w}^s + \tilde{N}\right) + H' \cdot \tilde{\Pi} = H \cdot \left(\tilde{w}^u + \tilde{n}\right) + H' \cdot \tilde{\Pi}. \qquad (22)$$

Equations (18)-(19)-(20)-(21)-(22) constitute a system of five equations in $\left\{\tilde{w}^s, \tilde{w}^u, \tilde{N}, \tilde{n}, \tilde{r}\right\}$. One can use (22) to substitute $\tilde{r}$ away from (18)-(19), and combine the ensuing expressions together with (20)-(21). This procedure yields the following equilibrium solutions for private employment changes:

$$\tilde{N} = \frac{\left[\theta + (1-\theta) \cdot \gamma \cdot H + \rho^u\right] \cdot \tilde{A}^s + (1-\gamma \cdot H)\left(1+\rho^u\right) \cdot \tilde{Q} + (1-\theta)(1-\gamma \cdot H)\tilde{A}^u - \gamma \cdot H' \cdot \left(1+\rho^u\right) \cdot \tilde{\Pi}}{\gamma \cdot H + \rho^u\left[1-\theta(1-\gamma \cdot H)\right] + \rho^s\left[\theta + (1-\theta) \cdot \gamma \cdot H + \rho^u\right]} \qquad (23)$$

and

$$\tilde{n} = \frac{\theta(1-\gamma \cdot H) \cdot \tilde{A}^s + (1-\gamma \cdot H)\left(1+\rho^s\right) \cdot \tilde{Q} + \left[1-\theta(1-\gamma \cdot H) + \rho^s\right] \cdot \tilde{A}^u - \gamma \cdot H' \cdot \left(1+\rho^s\right) \cdot \tilde{\Pi}}{\gamma \cdot H + \rho^u\left[1-\theta(1-\gamma \cdot H)\right] + \rho^s\left[\theta + (1-\theta) \cdot \gamma \cdot H + \rho^u\right]} \qquad (24)$$

From (23) and (24) one can immediately notice that skilled and unskilled private employment are increasing in areas which exhibit a relative advantage in terms of amenities (that is, $\left\{\tilde{A}^s, \tilde{A}^u\right\} > 0$) or productivity ($\tilde{Q} > 0$). By contrast, the direct impact of an increase of local public employment (that is, $\tilde{\Pi} > 0$) on private employment is negative.

The equilibrium expression for $\tilde{r}$ can be obtained from (22) by using (23)-(24), together with (20)-(21). In order to concentrate on the impact of public employment on the local economy, we set productivity and amenity terms equal to zero, that is, $\left\{\tilde{A}^s, \tilde{A}^u, \tilde{Q}\right\} = 0$. By doing so, we implicitly assume that local public employment has no impact on local amenities or productivity. Thus, expressions (23) and (24) reduce to:

$$\tilde{N} = \frac{-\gamma \cdot H' \cdot \left(1+\rho^u\right) \cdot \tilde{\Pi}}{\gamma \cdot H + \rho^u\left[1-\theta(1-\gamma \cdot H)\right] + \rho^s\left[\theta + (1-\theta) \cdot \gamma \cdot H + \rho^u\right]} \equiv B_1 \cdot \tilde{\Pi} \qquad (25)$$

and

$$\tilde{n} = \frac{-\gamma \cdot H' \cdot \left(1+\rho^{s}\right) \cdot \tilde{\Pi}}{\gamma \cdot H + \rho^{u}\left[1-\theta(1-\gamma \cdot H)\right] + \rho^{s}\left[\theta + (1-\theta) \cdot \gamma \cdot H + \rho^{u}\right]} \equiv B_2 \cdot \tilde{\Pi} \qquad (26)$$

The relative change in local rents induced by an increase in local public employment is given by:

$$\tilde{r} = \left\{1 - \frac{\left(1+\theta \cdot \rho^{s} + (1-\theta) \cdot \rho^{u}\right) \cdot \gamma \cdot H}{\gamma \cdot H + \rho^{u}\left[1-\theta(1-\gamma \cdot H)\right] + \rho^{s}\left[\theta + (1-\theta) \cdot \gamma \cdot H + \rho^{u}\right]}\right\} \cdot H' \cdot \tilde{\Pi} \equiv D \cdot \tilde{\Pi} \quad (27)$$

While (25) and (26) show that an increase in local public employment unambiguously reduces skilled and unskilled local private employment, the impact on rents is subject to two opposing forces. On the one side, public employment raises the local demand for housing but, on the other side, it displaces private employees. However, it is quite straightforward to show that the net effect on the local housing market, as summarized by D in (27), is non-negative.

In what follows, we analyze three simple cases based on different assumptions about the size of "mobility cost" measures, $\left\{\rho^{s}, \rho^{u}\right\}$.

Case 1. No mobility costs: $\rho^{s} = \rho^{u} = 0$.

This is the standard case from Roback (1982) onwards. It is immediate to notice from (27) that, absent mobility costs, a relative increase in local public employment ($\tilde{\Pi} > 0$) has *no net impact* on local rents! Thus, when workers are perfectly mobile, the demand for housing generated by public employees is exactly compensated by reductions in private employment. The expression for private employment displacement is given by:

$$\tilde{N} = \tilde{n} = -\frac{H'}{H} \cdot \tilde{\Pi} = -\left(\frac{\gamma \cdot \Pi}{\bar{\ell} \cdot r^{1+\delta} - \gamma \cdot \Pi}\right) \cdot \tilde{\Pi} \qquad (28)$$

The size of impact, given by $\left(\frac{\gamma \cdot \Pi}{\bar{\ell} \cdot r^{1+\delta} - \gamma \cdot \Pi}\right)$, can be either larger or smaller than one.

Case 2. Only the unskilled bear mobility costs: $\rho^{s} = 0, \rho^{u} > 0$.

In this case, the rent expression (27) reduces to:

$$\tilde{r} = \left\{ 1 - \frac{\left(1 + (1-\theta)\cdot\rho^u\right)\cdot\gamma\cdot H}{\gamma\cdot H + \rho^u\left[1-\theta(1-\gamma\cdot H)\right]} \right\}\cdot H'\cdot\tilde{\Pi} \qquad (29)$$

The quantity in braces from expression (29) is now strictly positive: thus, an increase in local public employment will exert a positive impact on local rents.

The expressions for private employment are as follows:

$$\tilde{N} = \frac{-\gamma\cdot H'\cdot\left(1+\rho^u\right)\cdot\tilde{\Pi}}{\gamma\cdot H + \rho^u\left[1-\theta(1-\gamma\cdot H)\right]} \qquad (30)$$

and

$$\tilde{n} = \frac{-\gamma\cdot H'\cdot\tilde{\Pi}}{\gamma\cdot H + \rho^u\left[1-\theta(1-\gamma\cdot H)\right]} \qquad (31)$$

Public employment still displaces skilled and unskilled private employment, but to a lesser extent: the size of the impact is decreasing in ρ^u in both (30) and (31). However, since the size of displacement is relatively larger for skilled workers, who are perfectly mobile in this case, more local public employment will *worsen* the local skill mix, measured by $\tilde{N} - \tilde{n}$.

Case 3. The skilled and the unskilled bear the same mobility cost: $\rho^s = \rho^u = \rho > 0$.

Now, the rent expression (27) reduces to:

$$\tilde{r} = \left\{ 1 - \frac{(1+\rho)\cdot\gamma\cdot H}{\gamma\cdot H + \rho[1+\gamma\cdot H+\rho]} \right\}\cdot H'\cdot\tilde{\Pi} \qquad (32)$$

Again, an increase in local public employment induces an increase in local rents. Private employment displacement is given by:

$$\tilde{N} = \tilde{n} = \frac{-\gamma\cdot H'\cdot(1+\rho)\cdot\tilde{\Pi}}{\gamma\cdot H + \rho[1+\gamma\cdot H+\rho]} \qquad (33)$$

Again, mobility costs reduce the impact of public employment on private employment, that is, the size of the impact is decreasing in ρ. However, since the skilled and the unskilled are assumed to have the same measure of mobility costs, the local skill mix is unaffected.

Appendix B: Data Appendix and Additional Tables

B.1 Census data and the economic activity classification

In 1991 the ISTAT standard for economic activity classification was set to ATECO 1991, than it was changed to get to ATECO 2002. ISTAT, following Eurostat requirement, in 2007 released the new ATECO 2007 that implements a quite radical change with respect to ATECO 2002. The ISTAT release for the 1991 and 2001 Census data are classified using the ATECO 5-Digit 1991, while the 2011 Census data is distributed with the ATECO 5-Digit 2007 classification. There is no official transition matrix from the ATECO 1991 classification to the ATECO 2007. There are, however, two different transition matrices, one from ATECO 1991 to ATECO 2002 and the other from ATECO 2002 to ATECO 2007 (both available on the ISTAT web site). Since for our purpose we can work with a less detailed classification, we approximated the 3-digit level classification in both matrices. The main problem is that, even at this level, each ATECO 2007 may correspond to multiple ATECO 2002 (and similarly for ATECO 1991). To solve this issue we use an aggregative mechanism method to build an univocal relationship. We started with the second and more critical transition matrix. We first removed those multiple correspondences that, at a close inspection, resulted to be less relevant. We then aggregated the 3-digit ATECO 2002 codes so that each ATECO 2007 was mapped into only one ATECO 2002 code. We then applied the same aggregation of the 3-digit ATECO 2002 codes to the 1991-2002 matrix. In very few cases this was not sufficient to have each ATECO 1991 code to be mapped to a single ATECO 2002 (re-aggregated) code. After careful inspection, these were marginal cases that we corrected by choosing the most relevant mapping. The do-file aggregating the codes is available with the replication material. However, since that public employment is concentrated in a few specific ATECO codes, and those are only marginally affected by our aggregation method, the reclassification process has no effect on our results. The eight categories obtained after the re-aggregation are:

- Administration of the State and the economic and social policy of the community
- Provision of services to the community as a whole
- Compulsory social security activities
- Pre-primary and primary education
- Information service activities, secondary and tertiary education, other education activities, Arts, Entertainment and Recreation
- Human health activities

- Residential care activities, social work activities without accommodation
- Residual category

B.2 House prices from *Osservatorio del mercato immobiliare*

The time series of housing prices at local level is based on the data released by the "Osservatorio del mercato immobiliare" (OMI) from 2003 onwards. The OMI data base contains semi-annual reports from approximately 8,100 Italian municipalities, in turn divided into about 31,000 homogeneous zones (whose identification is based on socio-economic and urban characteristics). The main sources are private real estate agencies, with a specific collaboration agreements; residually also administrative data on the transactions are considered. For each area and type of building (flats, villas and cottages) a minimum and maximum price are given. First the average is taken as the mid-point and then the price is further averaged across different buildings (with weights that do not vary across different municipalities). Finally the average price at the municipality level is calculated by weighting the different areas (center, semi center and periphery), with municipality-specific weights calculated by Cannari and Faiella (2008) trough information collected in the Bank of Italy surveys of Household Income and Wealth of Italian families (SHIW).

Figure B1. Municipality boundaries

Table B1. The impact of local public sector employment on private employment and house prices (2SLS)

	Columns (2) from Tables 2and 5		*Columns (3) from Tables 2and 5*	
	Private employment contribution to employment growth	House price growth	Private employment contribution to employment growth	House price growth
Public employment contribution to growth	-0.625***	0.003	-0.765***	0.505**
	(0.124)	(0.281)	(0.133)	(0.205)
Unemployment rate (2001)	-0.047	0.041	-0.067	0.017
	(0.064)	(0.063)	(0.105)	(0.140)
Unemployment rate in the rest of the LLM (2001)	0.186**	0.348*	-0.096	-0.171
	(0.078)	(0.191)	(0.103)	(0.196)
Labor force participation rate (2001)	0.365***	0.031	-0.384*	0.382
	(0.072)	(0.107)	(0.198)	(0.287)
Labor force participation rate in the rest of the LLM (2001)	0.222**	-0.136	0.134	0.523**
	(0.100)	(0.283)	(0.102)	(0.242)
Fraction of unoccupied houses (2001)	0.021	0.249***	-0.025	0.004
	(0.020)	(0.032)	(0.038)	(0.055)
Fraction of unoccupied houses in the rest of LLM (2001)	0.072**	0.156*	0.063**	0.055
	(0.030)	(0.081)	(0.031)	(0.066)
Commuting index (2001)	-0.147***	-0.070	-0.235***	0.196*
	(0.043)	(0.074)	(0.080)	(0.113)
Commuting index in the rest of the LLM (2001)	-0.057	0.072	0.069	-0.213
	(0.066)	(0.196)	(0.063)	(0.153)
Log population (2001)	0.111***	0.079***	0.098***	0.024*
	(0.008)	(0.017)	(0.008)	(0.013)
Log total employment (2001)	-0.107***	-0.022	-0.098***	0.006
	(0.007)	(0.015)	(0.008)	(0.012)
Log house prices (2003)	0.058***	-0.134***	0.039***	-0.347***
	(0.011)	(0.036)	(0.011)	(0.033)
Private empl contrib to growth (1991-2001)	0.033**	0.034	0.004	0.025
	(0.014)	(0.023)	(0.015)	(0.019)
Public empl contrib to growth (1991-2001)	-0.121*	0.019	-0.227***	0.210**
	(0.065)	(0.127)	(0.071)	(0.102)
Pop 15-64 growth w.r.t. to empl t-1 (1991-2001)	0.096***	0.120***	0.066***	0.041**
	(0.009)	(0.020)	(0.015)	(0.019)
Priv empl contrib (1991-2001) in the rest of LLM	0.072*	0.031	0.040	0.019
	(0.041)	(0.111)	(0.037)	(0.078)
Pub empl contrib (1991-2001) in the rest of LLM	0.064	-0.181	0.016	-0.056
	(0.081)	(0.206)	(0.083)	(0.157)
Pop 15-64 growth w.r.t. to empl t-1 (1991-2001) in the rest of LLM	0.027*	0.025	0.031**	0.000
	(0.016)	(0.061)	(0.016)	(0.043)
Constant	-0.817***	0.614**	-0.209	1.959***
	(0.083)	(0.291)	(0.132)	(0.265)
Additional variables from *ottomilacensus*			X	X
Obs	8,085	8,085	8,085	8,085

Note: *** p-val<0.01; ** p-val<0.05; * p-val<0.10. Standard errors, in brackets, are clustered at the LLM level (2001 definition). See Table 2 for other comments. The additional variables from ottomilacensus include the 1991 values for the unemployment rate, labor force participation rate, fraction of unoccupied houses and commuting index, plus all variables reported in Table B2.

Table B2: *Ottomilacensus* database, list of variables(1), with name and description; the last two columns indicate whether the variable is selected by the "double selection" procedure in the two main equations.

Code	Name	Description	Included in columns (3) from Tables 2 and 5 with values for		Selected for private employment eq. (Table 9 panel A)	Selected(2) for house prices eq. (Table 9 panel A)
			2001	1991		
P2	Annual population variation between census waves	Geometric average of annual population variation between census waves	X			
P5	Surface covered by urban centers and settlements	Percent of total surface covered by urban centers and settlements	X			
P6	Population resident in settlements and rarely populated areas	Percent of total population resident in settlements and rarely dense areas	X			
P7	Demographic density	Total population over surface in kmq	X			
P8	Male/Female ratio	Ratio (percent) of males to females	X	X		
P9	Fraction of population aged less than 6	Percent of population aged less than 6	X	X	1991	
P10	Fraction of population aged 75 or more	Percent of population aged 75 or more	X	X	2001	
P12	Young dependency index	Ratio (percent) of population aged up to 14 to population aged 15-64	X	X		
P14	Divorce index	Percent of population aged 18 or more that is divorced or legally separated	X	X	1991	
S1	Fraction of foreigners	Foreign-citizen residents per 1000 Italian residents	X	X		
F1	Average household size	Ratio of total population resident in households to number of households	X			
F2	Fraction of households without sub-units	Percent of households without sub-units	X	X		
F3	Fraction of households with two or more sub-units	Percent of households with two or more sub-units	X	X		2001
F4	Fraction of young individuals living alone	Ratio (percent) of households with a single young (less than 35) component to total population aged 15-34	X	X		
F5	Fraction of young one-parent households	Percent of young one-parent households (mother/father aged less than 35 with and without other co-residents) among households composed by a single unit	X	X		
F6	Fraction of young couples without children	Ratio of childless young couples (women aged less than 35) to households composed by a single unit (with or without other co-residents)	X	X		
F7	Fraction of young couples with children	Percent of young couples with children (women aged less than 35) among households composed by a single unit (with or without other co-residents)	X	X		
F8	Fraction of elderly living alone	Ratio (percent) of single-person households aged 65 or more to total population aged 65 or more	X	X		
F9	Fraction of old one-parent households	Percent of old one-parent households (mother/father aged 65 or more with and without other co-residents) among households composed by a single unit	X	X		2001

F10	Incidence of old couples without children	Ratio (percent) of old couples without children (women aged 65 or more) to households composed by a single unit (with or without other co-residents)	X	X		
A1	Fraction of own-housing	Percent of houses owned by the residents	X	X		2001
A3	Potential use of buildings	Ratio (percent) of unused buildings to total buildings	X			2001
A7	Services availability	Average of five ratios, each one calculated as the number of inhabited houses with a specific service available over total inhabited houses (services: drinkable water, bathroom, shower, heating, hot water)	X	X		
A8	Fraction of buildings in good condition	Percent of residential inhabited buildings in good or perfect condition	X			
A9	Fraction of buildings in bad condition	Percent of residential inhabited buildings in poor condition	X			2001
A10	Fraction of historical buildings	Percent of inhabited houses built before 1919	X	X		
A11	Index of housing expansion	Percent of inhabited houses, located in urban centers or clusters, that were built in the last decade	X	X	2001	1991
A12	Surface per inhabitant	Average house surface (sqm) per inhabitant	X	X	2001	1991,2001
A13	Houses under exploitation index	Percent of inhabited houses that meet tone of the following criteria: >80 sqm and 1 inhabitant; >100 sqm and <3 inhabitants; >120 sqm and <4 inhabitants	X	X		1991,2001
A15	Residential mobility	Percent of resident population that changed house in the last year	X			
I1	Gender gap in higher education	Ratio (percent) of percent of male with high-school diploma to percent of females with high-school diploma (in pop aged 6+)	X			
I2	Adults in life-long training	Percent of 25-64 population enrolled in education	X	X		
I4	Illiteracy index	Percent of 6+ population that is illiterate	X	X	1991	1991
I5	Early leave index	Percent of 15-24 population with middle-school diploma (8th grade) that is not enrolled in education	X	X		
I6	Fraction of adults with high-school diploma or university degree	Percent of 25-64 adults with high school diploma or university degree	X	X		2001
I9	Fraction of adults with middle-school diploma	Percent of 25-64 adults with middle-school diploma (8th grade)	X	X		2001
L2	Female labor force participation rate	Female labor force participation rate (pop aged 15+)	X	X		2001
L9	Young unemployment rate	Young (15-24) unemployment rate		X		1991
L14	Young employment rate	Young employment rate (pop aged 15-29)	X	X		1991,2001
L16	Fraction of employment in manufacturing		X	X	2001	2001
L17	Fraction of employment in non-trade services		X	X	1991, 2001	1991,2001
L18	Fraction of employment in trade-services		X	X	1991, 2001	2001

L19	Fraction of employment in medium-high specialization	Percent of employed individuals in types 1,2,3 of occupations (legislators, directors, intellectual scientific occupations with high specialization, technical occupations)	X	X		
L20	Fraction of employment in craftsmen, blue-collar or agricultural positions	Percent of employed individuals in types 6,7 (Craftsmen, blue-collars and agricultural workers)	X	X		
L21	Fraction of employment in low-skilled positions	Percent of employed individuals in type 8 (non-qualified occupations)	X	X	1991	1991, 2001
M1	Daily mobility for work or study	Percent of <65 population moving daily to go to work or study	X	X	2001	2001
M5	Mobility with private transportation means	Percent of the population moving daily that uses private means of transportations	X	X		
M6	Mobility with public transportation means	Percent of the population moving daily that uses public means of transportations	X	X		1991
M7	Slow mobility	Percent of the population moving daily by foot or bicycle	X	X	1991	1991
M8	Short mobility	Percent of the population moving daily that commutes for less than 30 minutes	X	X		
M9	Long mobility	Percent of the population moving daily that commutes for more than 60 minutes	X	X		2001
V4	Fraction of irregular accommodations	Percent of accommodations that cannot be classified as houses	X	X		
V5	Fraction of large households	Percent of households with 6 or more members	X	X		
V6	Percent of households with potential economic disadvantage	Percent of households where there are children, the reference person has less than 65 years and all components are neither employed nor retired.	X	X		
V7	Population living in over-exploited houses	Percent of population living in houses that meet tone of the following criteria: <80 sqm and >4 inhabitants; [40,59] sqm and >5 inhabitants; [60-79] sqm and >6 inhabitants	X	X		
V8	Percent of young 15-29 residents neither active on the labor market nor in training	Percent of young 15-29 residents neither active on the labor market nor in training	X	X		2001
V9	Percent of households in need of assistance	Percent of households with at least two members, with no other sub-units, with all members aged 65+ and at least one aged 80+	X	X		1991,2001

(1) Apart from the variables reported in the list, the controls included in column (3) from Tables 2 and 5 include also the 1991 values for the unemployment rate, labor force participation rate, fraction of unoccupied houses and commuting index. Variables that are not available for all municipalities in a specific wave have not been included. Some variables were excluded a priori from the original *ottomilacensus* database because they were, by construction, highly correlated with other variables, or already included in the baseline covariates (P1 resident population, P4 annual population variation limited to those 15+, P11 Elderly dependency index, A2 average surface of houses, A14 houses overexploitation index, F11 Incidence of old couples with children, I3 Ratio of adults with high-school or university degree to those with middle school degree, L1 male labor force participation rate, L4 fraction of young individuals (15-29) neither in employment nor training, L6 male unemployment rate, L7 female unemployment rate, L10 male employment rate, L11 female employment rate, L12 employment rate, L15 fraction of employment in agriculture, V1 vulnerability index, V2 ranking in the vulnerability index). F1 Average household size, P7 demographic density, P5 Surface covered by urban centers and settlements, P6 Population resident in settlements and rarely populated areas, have been included only for 2001 because they are highly correlated over time.

(2) In the house price equation, the “double selection” procedure also selects the fraction of unoccupied houses for 1991.

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