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Rent sharing as a driver of the glass ceiling effect

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ABSTRACT

In this paper we show that rent-sharing plays a role in explaining the glass ceiling effect. We make use of a unique employer–employee panel database for Italy from 1996 to 2003, which allows controlling for observed individual and firm heterogeneity and for collective bargaining. Moreover, by means of IV quantile fixed effects estimates we can cope with unobserved heterogeneity and endogeneity. A discussion of different explanations is provided.

1. Introduction

The glass ceiling effect is one of the stylised facts concerning the gender pay gap. One of the pioneering works is Albrecht et al. (2003) that uses quantile regressions and Swedish data finding an increasing gender pay gap along the wage distribution. Other papers have then extended this finding to most of the OECD countries (Arulampalam et al., 2007). Although the glass ceiling phenomenon is observed in most OECD countries, the understanding of the reasons behind it represents an open field of research (Booth, 2007), with relatively few papers testing explanations from an empirical point of view ((De la Rica et al., 2010; Bertrand and Hallock, 2001), among others).[1]

In this paper we propose a new explanation for the glass ceiling effect, investigating whether men and women differ in their efficacy to extract rents from firms, and whether this difference increases along the wage distribution. We make use of a unique employer–employee panel database for Italy.

The relation between rent-sharing and the gender wage gap has been previously investigated by Nekby (2003), both at the conditional mean and along the wage distribution.[2] Nonetheless, this paper did not properly control for unobserved heterogeneity and endogeneity, issues that have been proved to be crucial in the estimation of rent-sharing (Abowd and Lemieux, 1993; Card et al., 2010). We can cope with these issues by using IV quantile fixed effects estimates.

2. Data description

We make use of a unique panel version of the administrative employer–employee database provided by INPS (Italian Social Security Institute). The sample units are industrial and service dependent workers, both part-time (converted to full-time equivalent) and full-time, in standard labour market contracts (blue collar,

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[1] Bertrand and Hallock (2001) is actually related to a slightly different literature that investigates the behaviour of selected group of workers, such as CEOs and top executives.

[2] Also Plasman et al. (2004) investigate the impact of rent-sharing on the gender pay gap. However, they do not address the glass ceiling phenomenon since their analysis is restricted to the conditional mean.

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white collar and managers), aged between 15 and 64 (when they first enter in the database), with al least two observations in the panel. We merge the INPS dataset with the AIDA database, which includes information on the balance sheet of (capital-owned) firms from 1996 to 2003, in such a way restricting the sample to workers employed in capital owned firms.[3]

Our main independent variable is quasi-rent per worker, i.e. the rents per worker evaluated at the opportunity cost of labour, which is defined as the revenue per worker (operative income –which equals to net profits–plus the wage bill), minus the alternative wage, as in Van Reenen (1996).

An important value added from our matched employer–employee database concerns the way of computing the alternative wage, since it allows controlling accurately for the collective bargaining, i.e. for the part of bargaining that takes place at national/sectoral level and that is not related to individual negotiation. In Italy the collective bargaining is characterised by two levels: a first centralised (national) level where minimum wages for all occupations are set in all industries; a second decentralised level where the employer and employees (individually or at the firm/territorial level) can bargain over wages and other working conditions. To control properly for the first national level of bargaining we introduce in our estimation the minimum wage corresponding to the worker's specific national contract and, within the contract, to the specific occupation (''livello di inquadramento''), as in Card et al. (2010). This turns out to be a more reliable measure of the alternative wage with respect to using average industrial wages. Since in Italy there are more than 200 national contracts, we restrict our sample to the greatest 26 national contracts (80% of the total sample), to have enough variability within each contract-occupation cell.

Table 1 shows the descriptive statistics by gender of the variables of the analysis.

3. Econometric specifications and results

Since our analysis concerns the whole wage distribution, we make use of the quantile regression approach. The baseline specification is the following:

$$\begin{aligned}\ln(w_{i,t}) = {} & \alpha_\theta + \chi_\theta * \ln MW_{c(i,t)} + B'_\theta * \text{I\_Char}_{i,t} \\ & + \beta_\theta * \ln \mathit{Firmsize}_{j(i,t)} + \gamma_{1,\theta} * \ln \mathit{QuasiRents}_{j(i,t)} \\ & + \varphi_{s,\theta} + \lambda_{a,\theta} + \delta_{t,\theta} + \varepsilon_{i,t,\theta}\end{aligned}$$

where θ refers to the percentile, i to individuals, $j(i, t)$ to the firm where the worker i is employed at time t, $c(i, t)$ to the national contract the worker is subject to, and s to industry. The dependent variable in our regressions is the (log) real gross weekly wage. The term $\text{I\_Char}_{i,t}$ is a set of observed individual characteristics (age, age squared, tenure and occupation dummy). $MW_{c(i,t)}$ is the national contract minimum wage that controls for first level bargaining. $\mathit{QuasiRents}_{j(i,t)}$ is quasi-rent per employee. $\mathit{Firmsize}_{i,t}$ is the proxy for firm heterogeneity, while $\phi_s, \lambda_a, \delta_t$ are industry, area (five macro-areas in Italy: Northwest, Northeast, Centre, South and Islands) and time dummies respectively. All the variables of interest are in logarithms and therefore we estimate elasticities.

The first step of the analysis is to carry out cross sectional quantile estimates at the 10th, 25th, 50th, 75th and 90th percentiles, separately for men and women, controlling for observed heterogeneity of workers and firms. From Table 2 it is possible to note that rent-sharing estimates are greater for men, and for both men and women they are slightly increasing along the wage distribution. However, using cross sectional regressions we cannot control for the unobserved heterogeneity of workers. Hence, we implement quantile fixed effect estimates, as proposed by Koenker (2004). Estimates in Table 3 prove that, as expected, when controlling for the individual unobserved heterogeneity rent-sharing estimates strongly dampen, consistently with the sorting literature (Mion and Naticchioni, 2009) and with the rent-sharing literature (Card et al., 2010). Moreover, in fixed effects the estimates are basically flat both for men and women and still higher for men than for women.

The last step in order to derive unbiased estimates is to address the endogeneity due to the likely simultaneous determination of wages and profits and to measurement errors, endogeneity that can cause a severe underestimation of rent-sharing (Van Reenen, 1996; Card et al., 2010). Since we are working in a quantile framework we apply a very recent methodology developed by Galvao and Montes-Rojas (2009), Galvao (2011) and Harding and Lamarche (2009). This procedure is an extension of the IV quantile procedure of Chernozhukov and Hansen (2008) that allows for the inclusion of fixed effects as introduced in Koenker (2004).

As an instrument for firm profits we exploit the idea developed in Card et al. (2010) by using for each firm in a given province the average of current total real sales per employee of firms located in all other Italian provinces but operating in the same three-digit sector.[4] The identifying assumption is that national industry demand shocks affect firm level profitability but have no direct effect on local labour conditions.[5]

Results are shown in Table 4. As expected, the degree of underestimation of the fixed effects estimates is substantial. As for men, rent elasticities are quite stable along the wage distribution, ranging from 5.9% at the 10th percentile to 4.4% at the 90th percentile. For women, the extent of rent-sharing is again lower than for men. Further, it is basically stable from the 10th percentile to the median (3.8% and 3% respectively), while it falls substantially at the 90th percentile (1.6%). These findings strongly suggest that, once controlling for unobserved heterogeneity and endogeneity, the rent-sharing impact is such as to increase the gender wage gap along the wage distribution, contributing to generate a glass ceiling effect.[6]

To characterise further our results we analyse other factors that might play a role in explaining the differences in the rent-sharing effect on gender pay gap. First, we consider a possible sorting of women into less profitable firms, pointing out that in our

[3] Data on profits are deflated using the value added deflator (base year, 2002). Further, we clean our data as in the following: we drop observations for which the difference in absolute value between the firm size reported in AIDA and the firm size reported in INPS was higher than 200 (in so doing the correlation between firm size in the two databases is 99.95), as well as extreme observations below (above) the 1st (99th) percentile of wages and profits per employee; we also drop outliers with respect to the yearly growth rate of wages and profits per employee.

[4] To compute these averages we use weights equal to the inverse of distances between provinces: more weight is given to closer provinces. The weighing procedure increases the explicative power of the instrument. Nonetheless, similar results apply even when weights do not change with distance.

[5] The estimation is carried out simultaneously on three percentiles (10th, 50th, 90th) for computational reasons. Further, this estimation technique does not allow testing the weakness of instruments. The only possible check is to implement a standard IV fixed effects estimation and look at the F-statistic of the first stage. In our case the F-statistics are statistically significant and higher than the threshold value of 10 for both gender categories.

[6] As a robustness check, instead of using the individual minimum wage as a measure of the alternative wage, we make use of a more standard measure used in the literature, the industrial wage (as in (Van Reenen, 1996)). This is computed as the average of individual minimum wages at the national contract level (each national contract roughly corresponds to a different industry). Results are similar from a qualitative point of view. In particular, estimates are still decreasing along the wage distribution both for men and women, with a more substantial drop for women at the top of the distribution (from 0.03 at the 10th percentile to 0.01 at the 90th percentile) than for men (from 0.08 to 0.07), confirming a widening of the gender wage gap at the 90th percentile.