The Extent of Rent Sharing along the Wage Distribution

Alessia Matano[1] and Paolo Naticchioni[2]

January 2015

Abstract

The relationship between rent sharing and wages has generally been evaluated on average wages. This paper uses a unique employer-employee panel database to investigate the extent of rent sharing along the wage distribution in Italy. We apply quantile regression techniques and control for national level bargaining, unobserved heterogeneity and endogeneity. Our findings show that the extent of rent-sharing decreases along the wage distribution, suggesting that unskilled workers benefit most from firms' rents. We provide evidence supporting an explanation based on the role of the unions, which are more interested in favoring unskilled workers.

JEL Classification: C33, J31, J41, L25.

Keywords: Rent Sharing, Wage Distribution, Quantile Regressions, Quantile IV fixed effects regressions, Unions.

---

[1] Alessia Matano, Universitat de Barcelona, AQR-IREA and University of Rome "La Sapienza", Dipartimento di Analisi Economiche e Sociali. Email: amatano@ub.edu.

[2] Corresponding Author: Paolo Naticchioni, Roma Tre University and IZA, Via Chiabrera 199 – 00145 Rome (IT), email: p.naticchioni@gmail.com.

We are grateful to the research partnership between ISFOL (Rome) and Dipartimento di Analisi Economiche e Sociali - University "La Sapienza" of Rome, for access to the INPS and AIDA databases. We also thank Roger Koenker for his comments and for making his procedure available, Christian Dahl, Antonio Galvao and Christian Hansen for access to the routines and for suggestions concerning implementation of the quantile procedures, and participants in seminars at University of Barcelona, AIEL 2012 conference (Naples), LEED Workshop 2013 (Lisbon), EALE 2013 (Turin), SIE 2013 (Bologna). Alessia Matano also expresses her gratitude for the support received from Spain's Ministerio de Ciencia e Innovación (VI Plan Nacional de Investigación Científica, Desarrollo e Innovación Tecnológica 2008-2011) through the action ECO2010-16006.

1. Introduction

European countries are usually taken as examples for non-competitive labor markets because of the important role played by labor market institutions. The economic literature has largely investigated how wage setting works in non-competitive labor markets, and how rent sharing can emerge in such markets. Non-competitive theories, such as efficiency wage and bargaining models, can predict a positive relationship between wages and profits. In particular, bargaining models underline that wages result from a bargain between employer and employees which generates a long-run positive relation between wages and profits. In this setting, wages are determined by workers' outside options, by quasi-rent (firm profits evaluated at the opportunity cost of labor) and by the relative bargaining power of the parties involved (Hildreth and Oswald, 1997).

At the empirical level many papers have tested the existence and extent of rent sharing (Abowd and Lemieux, 1993, Van Reenen, 1996, Margolis and Salvanes, 2001, Martins, 2009, Card et al., 2013, etc.). However, these analyses have generally been carried out taking into account average wages. In this way there can be no insight into the distributional consequences of rent sharing, i.e. it is not possible to take into account the difference in the degree of rent sharing for workers located at different points of the distribution.

The aim of this paper is to evaluate the degree of rent sharing along the whole wage distribution in order to achieve a better understanding of the mechanisms behind the relation between profits and wages. Previous empirical investigations have analyzed rent sharing across categories of workers, defined by education and by occupation. The main drawback using this approach is that workers belonging to the same education/occupation level are usually associated to very high within group wage heterogeneity. For instance, according to the 1996 data of the European Community Household Panel, almost 50% of Italian graduates were not employed in the top quartile of the wage distribution, and around 20% had a wage lower than the median, suggesting a substantial heterogeneity within educational levels. A similar argument can be applied when considering blue collar and white collar workers, where especially in the white collar category secretaries coexist with managers, with huge differences in terms of productivities and wages. We make use of quantile regressions methodologies to deal with this heterogeneity, since percentiles of the wage distribution can be more closely associated to the productivity of workers in the labor market.

Furthermore, there could be various different reasons why rent sharing is not uniform along the workers' wage distribution. On the one hand, it might be argued that if bargaining at the firm level was mainly organized by unions, low and median skilled workers might enjoy a higher degree of rent sharing than high skilled workers. On the other hand, if bargaining occurred mainly at the individual level, rent sharing might favor high skilled workers, who can benefit from higher individual bargaining power and from performance pay schemes (Lemieux et al., 2009). Hence, given the ambiguous theoretical predictions, the analysis of rent sharing along the wage distribution is mostly an empirical issue, and to the best of our knowledge this is the first paper that addresses this issue along the whole wage distribution.

In our analysis we make use of a unique employer-employee panel database from 1996 to 2003 for Italy, constructed by merging the INPS (the Italian Social Security Institute) employer-employee panel database with the AIDA database (provided by Bureau Van Dick) which contains detailed information on the balance sheets of the Italian capital-owned firms.

On the econometric side, our empirical analysis takes into account all the issues which have been proved to be relevant when addressing the relationship between rents and wages.

We begin by estimating the impact of quasi-rents on wages using cross-sectional quantile regressions (Koenker and Basset, 1978), controlling for observed worker and firm heterogeneity. In the estimation we use as proxy for the opportunity cost of labor the minimum wage corresponding to the national contract applied to each worker and, within the national contract, to the exact occupation level ('livello di inquadramento') the worker belongs to. We argue that this is a more accurate measure to control for the opportunity cost of labor with respect to the average industrial wage, as generally used in the literature. Another advantage in using individual minimum wages is that they allow controlling in the estimates for the first national collective bargaining level, since minimum wages in Italy are formally bargained at the national level between unions and employer associations. The cross-section estimates show that the impact of rent sharing is positive for all percentiles analyzed and is decreasing along the wage distribution: rent sharing elasticities range from 6.9% at the 10th wage percentile to 4.7% at the 90th wage percentile.

A second step in the analysis is to control for the unobserved worker heterogeneity, which can affect the relationship between profits and wages (Card et al., 2013, Arai and Heyman, 2001, Margolis and Salvanes, 2001, Martins, 2009). By applying quantile fixed effects estimates that explicitly take into account the individual unobserved heterogeneity (Canay, 2011), the impact of rent sharing is significantly reduced along the whole wage distribution and rent sharing elasticities are still decreasing along the wage distribution.

The last step of the empirical analysis investigates the endogeneity issue, which has been proved to be a serious concern in the analysis of rent sharing since endogeneity could cause serious coefficients underestimation (Card et al., 2013, van Reenen, 1996), an underestimation which could even be exacerbated by the introduction of fixed effects in the specification (Nickell, 1981). Therefore, we apply IV quantile fixed effect estimation techniques (Galvao, 2011, Galvao and Montes-Rojas, 2010). For the instrument, we exploit the intuition developed in Card et al. (2013) by using a weighted average of the real sales per employee in other provinces of Italy in the same 3-digit industry. The idea is that real sales per employee in the same industry – which represents national industry demand shocks - affect the profitability of the firms. Further, these sales relate to firms in other provinces of Italy and are therefore assumed to be uncorrelated with local labor market conditions. Consistently with the related literature, by applying an IV methodology estimates increase along the whole wage distribution and by a large extent, thus pointing out that previous fixed effects estimates suffered by a serious degree of underestimation. In particular, the elasticity of wages with respect to rent stands at 7.3% at the 10th percentile, 4.8% at the median and 3.7% at the 90th percentile, confirming that the degree of rent sharing is decreasing along the wage distribution.

As possible explanation for this decreasing pattern, one might argue that in Italy bargaining takes place mainly at the national, local and firm level, where the role of the unions is more effective, while individual level bargaining plays a less important role. We test this possible explanation exploiting the regional variability in the union membership rate, considered as a proxy for union power. Focusing on the manufacturing sector, we show that in regions where union power is high rent sharing for unskilled (10th percentile) and medium skilled (50th percentile) workers is higher with respect to regions where union power is low. We also derive evidence suggesting that where union power is high rent sharing for skilled workers (90th percentile) is lower, meaning that individual bargaining for skilled workers is less effective with respect to the case where union power is low.

The structure of the paper is as follows. In Section 2 we review the theoretical and empirical literature on the relationship between profits and wages. In Section 3 we describe the data we use throughout the empirical analyses. Section 4 discusses the empirical specification and presents the main results. Section 5 concludes.

2. Related Literature

Non-competitive theories underline that firms may pay a wage over the level set in the competitive labor market for various different reasons. First, it is possible that firms pay higher wages on the basis of efficiency wage arguments (see Shapiro and Stiglitz, 1984, Krueger and Summers, 1988). Second, according to bargaining theories, profits and wages can move together due to the bargaining over wages between employers and employees. More specifically, in a bargaining framework, wages at the firm level are determined by workers' outside options, by the quasi-rent (firm profits evaluated at the opportunity cost of labor) and by the relative bargaining power of the parties involved (Hildreth and Oswald, 1997).[3]

As for the empirical evidence, many studies explore the existence and the extent of rent sharing in different countries, using various methodologies and various kinds of data. Hildreth and Oswald (1997) make use of firm level data for the UK providing evidence in favor of a significant positive relationship between profits and wages, controlling for observed work heterogeneity and firm characteristics and applying GMM techniques (or using lagged values of profits) to control for the endogeneity of profits. Similar findings are derived by Blanchflower, Oswald and Sanfey (1996) for the US, using industry level data matched with individual data.

Other papers use instrumental variables techniques to control for the endogeneity of profits. Abowd and Lemieux (1993), in the case of Canada, use instruments related to

[3] Note that also within a modified version of the competitive model it is possible to have a positive correlation between wages and profits. In particular, in the presence of short-run frictions, such as those experienced by firms facing an upward sloping labor supply curve, positive demand shocks could lead to a rise in total firm profits and wages (Hildreth and Oswald, 1997). However, in the long-run, wages adjust to the competitive level. Hence, a test for rent sharing cannot rest on the evidence of a short-run correlation between profits and wage.

international performance, namely the industry import and export prices, finding a very large degree of underestimation in the extent of rent sharing when not controlling for the endogeneity between profits and wages. Van Reenen (1996) analyzes the case of the UK using different measures for profits (net profits per head, quasi-rents and Tobin Q), and past innovations as instruments. His findings suggest a substantial amount of rent sharing in the UK, and serious underestimation when not controlling for endogeneity.

More recently, various papers have made use of matched employer-employee panel data in order to control for unobserved worker heterogeneity. Margolis and Salvanes (2001) investigate the case of France and Norway. They apply IV techniques using as instruments sales and operating subsidies, finding relevant rent sharing only in the case of Norway. In the case of France they show that when taking into account the unobserved individual heterogeneity in the IV estimation, rent sharing estimates turn out to be not significant. Similarly, using employer-employee data Arai (2003) analyzes the case of Sweden. He uses time-average of lagged values of profits as instruments and controls for observable firm characteristics to verify the relevance of different theoretical explanations for the relationship between profits and wages (rent sharing, efficiency wages, short-run labor market frictions). He finds robust evidence of rent sharing, in line with bargaining theories, and this effect does not differ across the different worker categories.[4] In another related paper, Arai and Heyman (2001) make use of a larger employer-employee matched dataset and apply instrumental variable techniques. They use different instruments such as lagged values of profits, demand elasticity (based on predicted response in sales due to higher prices) and measures indicating the degree of competition in the product market. Their findings confirm that rent sharing is underestimated when not controlling for endogeneity, and that even greater estimates are provided when demand elasticity is used as instrument.

Also Martins (2009) makes use of matched employer-employee panel data to derive evidence of rent sharing for Portugal in the period 1993-1995. His findings strongly support the need to take into account the role of both the unobserved individual and firm heterogeneity, as well as endogeneity (the interaction between the exchange rate and the share of total exports in sales is used as instrument).

[4] However, note that the results of this analysis could be affected by the very small sample size compared with other studies that use employer-employee data.

Another interesting related paper is Guertzgen (2009), which focuses on how rent sharing is affected by the different levels of bargaining in Germany, using firm-worker level data and GMM techniques. He shows that rent sharing is higher where there is no collective agreement coverage and in the presence of firm-specific contracts. Moreover, he also shows that blue collar workers in uncovered establishments seem to benefit more from the local bargaining power of works councils, i.e. local unions.

Rusinek and Rycx (2013) also analyze the impact of different levels of bargaining (industry and firm level) on the extent of rent sharing, using an employer-employee database for Belgium, a country where the relative importance of industry and firm level agreements (the degree of centralization) differs significantly across industries. Their results show that, after controlling for the endogeneity of profits and heterogeneity among workers and firms, in industries where agreements are more likely to be renegotiated at firm-level ('decentralized industries'), wages and profits are positively correlated regardless of the type of collective wage agreement. On the contrary, where firm-level wage renegotiation is less likely ('centralized industries'), wages are only significantly related to profits for workers covered by a firm-level collective agreement.[5]

As for Italy, empirical evidence on rent sharing is somewhat wanting. One of the few exceptions is the recent paper by Card et al. (2013), which analyzes the degree of rent sharing and tests the hold-up hypothesis in the Italian region of Veneto for the period 1995-2001. By using INPS-AIDA matched employer-employee data, they perform an accurate analysis taking into account all the relevant issues needed to identify the extent of rent sharing (the workers' and firms unobserved heterogeneity and the endogeneity of profits). Their findings show that there is evidence of a substantial degree of rent sharing in Veneto, and that profits are shared with workers after capital costs are fully deducted from profits.[6]

[5] See also Martins (2007) for a survey of the main empirical results and methodologies applied in the rent sharing literature.

[6] Another paper on the Italian case is Pistoresi and Strozzi (2003). Their main findings are that rent sharing in Italy arises only at the centralized level of wage bargaining, while decentralized wage negotiations do not lead to any degree of rent sharing between unions and employers. However, since they use time series techniques and industrial data, they cannot take into account the within-industry heterogeneity (observed and unobserved worker and firm heterogeneity).

3. The Italian institutional setting and Data Description

The institutional issues related to this paper concern the Italian wage setting. Since the beginning of the nineties there has been a two-level wage bargaining system, which is similar to schemes used in other European countries such as Germany. The first level concerns national collective bargaining, which has to preserve the purchasing power of wages at the sector level by incorporating the expected inflation rate in wage increases. This is done concretely by setting minimum wages for all workers covered by the related National collective agreements, which are renewed every 2-4 years. Minimum wages are different in each industry, and within industries different minimum wages are assigned to different workers in different occupation levels (*'livelli di inquadramento'*): this means that minimum wages are settled, at different levels, for blue collar workers, white collar workers and managers.

The second level of bargaining is decentralized, and encourages rent sharing through performance-related pay schemes at the region/firm level.[7] This second level is not compulsory for firms and unions, while it is compulsory respecting the lower bound set by the minimum wage of the first national bargaining level.

As for the data, we use a panel version of the administrative database provided by INPS (Italian Social Security Institute) and elaborated by ISFOL.[8] It is a matched employee-employer dataset, constructed by merging the INPS employee information database for the period 1985-2003 with the INPS employer information database.[9] The database contains individual information such as age, gender, occupation, workplace, date of beginning and end (if any) of the current contract, the kind of national contract and the related minimum wage, the social security contributions, the worker status (part-time or full-time), the real gross yearly wage and the number of weeks worked. We then have some information concerning the firm such as the plant location (province), the number of employees and the sector (NACE Rev.1.1). We focus on male

[7] Apart from the wage setting issue, the second level bargaining may also concern other work dimensions, such as hours worked, working conditions, etc. Furthermore, note that individual bargaining for all workers is always allowed by labor legislation in Italy.

[8] ISFOL stands for "Institute for the Development of Vocational Training". The sample scheme has been set up to follow individuals born on the 10th of March, June, September and December and therefore the proportion of this sample on the Italian employees' population is approximately of 1/90.

[9] For the information on employers we also make use of the ASIA ("Italian Statistical Archive of Operating Firms") database, provided by ISTAT. This database has been used since 1999, because the INPS employer database was no longer available as from 1998. The two databases provide the same set of information (firm size and sector).

and female prime-age workers, aged between 25 and 49 (when they first enter the database), working in the industrial and service sectors, both part-time (converted into full-time equivalent) and full-time, employed in standard labor market contracts: blue collar and white collar workers.[10]

We merge the INPS dataset with the AIDA database, from 1996 to 2003. AIDA is a database on Italian (capital-owned) firms provided by Bureau Van Dijk which contains information on the balance sheet such as value added, profits, sales, production and costs of production.[11]

The two databases are merged by using as key variable the tax code or the VAT number (*codice fiscale* or *partita IVA*) of the company.[12] After the merge, the panel version has been constructed considering only one observation per year for each worker. For those workers who display more than one observation per year we selected the longest available contract in terms of weeks worked. We also eliminated extreme observations below (above) the 1st (99th) percentile of the wage and quasi rent distributions.[13] Further, we dropped those observations for which the growth rate of wages from year to year was higher (lower) than 100%(-50%) and where the growth rate of the quasi rent variable was higher (lower) than 500%(-500%). These thresholds were computed taking into account the growth rate values corresponding approximately to the 1st and 99th percentile of the related growth rates distribution. We also eliminated those observations where the percentage difference in the firm size reported in AIDA and the one reported in INPS exceeds 5% (in this way the correlation between the firm size reported in AIDA and the firm size reported in INPS is equal to 0.96). Finally, we dropped workers for whom data on the minimum wage is not available. In fact, our database does not include minimum wages for the – nearly - 300 national contracts. We have this information for the 39 major contracts, which nonetheless cover more than 75% of the whole sample.

---

[10] The sample also includes managers. However, since they account for a relatively small fraction of workers in the sample (only about 1%, because most of the managers are not covered by the INPS archive) we include this category within the white collars.

[11] The data have been deflated using the valued added deflator for value added, profits, sales, production and costs of production. The value added deflator derives from our elaboration of ISTAT data on regional economic accounts and is defined at the sectoral and regional level. The base year is 2002.

[12] Note that AIDA contains capital-owned firms with total value of production equal to or higher than 950.000 euro, while INPS data cover workers employed in all kinds of companies whatever the legal status and amount of total value of production. Therefore, it is possible to match only the INPS records of firms that are included in the AIDA database.

[13] Note that we eliminate all the observations of workers for which there is at least one outlier.

We end up with an employer-employee panel database constituted by 25,796 workers for 123,178 observations for the period 1996-2003.

4. Econometric Analysis

4.1 Econometric Strategy

In this section we analyze the impact of rents on wages. Since our focus is on the relationship between rents and wages along the wage distribution, we start by performing standard quantile regressions (Koenker and Bassett, 1978). We use the INPS-AIDA employer-employee database from 1996-2003. The baseline specification is quite standard in the rent sharing literature (see for instance Van Reenen, 1996), and it is as follows:

$$\ln(w_{\theta(i,t)}) = \alpha_\theta + \chi_\theta * \ln MW_{c(i,t)} + B'_\theta * I\_Char_{i,t} + \beta_\theta * \ln Firmsize_{j(i,t)} + \\ + \gamma_{1,\theta} * \ln QuasiRents_{j(i,t)} + \phi_{s,\theta} + \lambda_{a,\theta} + \delta_{t,\theta} + \varepsilon_{i,t,\theta}$$

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 (along with its level) the worker is subject to, s to industry. The dependent variable in our regressions is the (log) real gross weekly wage in euro.[14] As main independent variable we use the quasi-rent per worker, $QuasiRents_{j(i,t)}$, which are defined as rents per worker evaluated at the opportunity cost of labor, i.e., the revenue per worker (net profit per worker plus the wage bill per worker) minus the alternative wage, as in Martins (2009) and Card et al. (2013). The term $I\_Char_{i,t}$ is the set of observed individual characteristics, such as age, age squared, tenure (in three categories, 1-2, 3-10, more than 10 years) and occupation dummy (blue collar and white collar). $MW_{c(i,t)}$ is the national contract minimum wage. $Firmsize_{i,t}$ is the proxy for firm heterogeneity, while φ_s, λ_a, δ_t are industry, area (five macro-areas in Italy: Northwest, Northeast, Centre, South and Islands) and year dummies respectively. All the relevant variables are in logarithms and therefore we estimate elasticities. Table 1 shows the descriptive statistics of the variables of the analysis.

[14] Wages have been deflated using as deflator the National Consumer Price Index (FOI index, *Indice dei Prezzi al Consumo per le Famiglie di Operai e Impiegati*, ISTAT). The base year is 2002.

Table 1: Descriptive Statistics of the Variables of the Analysis				
Variable	**Mean**	**Std. Dev.**	**Min**	**Max**
Log Real Weekly Wage	5.98	0.28	4.39	8.86
Log Real Weekly Minimum Wage	5.69	0.12	5.37	6.80
Female	0.31	0.46	0.00	1.00
Age	37.72	9.72	25	56
Age Squared	1,477.83	755.89	625	3,136
Blue Collars	0.63	0.48	0	1
White Collars and Manager	0.37	0.48	0	1
Log Firm Size	4.65	1.51	0	10.69
Log Quasi-Rent per Employee	3.01	0.96	-6.14	5.01
Log Real Sales per Employee other provinces (instrument)	5.23	0.50	2.63	7.16
Tenure 1-2	0.33	0.48	0	1
Tenure 3-10	0.48	0.50	0	1
Tenure >10	0.19	0.37	0	1
North East	0.30	0.46	0	1
North West	0.42	0.49	0	1
Centre	0.16	0.37	0	1
South	0.09	0.28	0	1
Island	0.03	0.17	0	1
Number of Observations	123,178			
Number of Workers	25,796			

Source: Panel ISFOL on INPS-AIDA data. Note: Sectoral dummies are defined according to Nace Rev 1.1, and the related descriptive statistics are omitted from the table for the sake of space. As for the main aggregates, the industry accounts for around 58% of the observations, while the service sector for 42%.

In the first specification, as benchmark estimates, we perform cross-sectional quantile estimates where, as already pointed out, we use as alternative wage the minimum wage which captures the extent of the first (national) level of bargaining. It is worth noting that the minimum wage turns out to be a very accurate measure to control for first level bargaining at the national level. At the same time it is the best available measure of the opportunity cost of labor. We believe this measure represents a valuable improvement with respect to the average industrial wages generally used in the literature, mainly because it is related to the specific contract (and within the contract to the specific level) the worker belongs to.

Since an important concern in our analysis is to tackle the issue of the unobserved individual heterogeneity that can bias the cross sectional estimates, we then carry out quantile fixed effects estimates (Canay, 2011). In fact, in the literature unobserved worker heterogeneity has been proved to be very important in affecting the relationship between rents and wages since high-skilled workers may sort into highly profitable firms (Card et al., 2013, Martins, 2009, Arai and Heyman, 2001, Margolis and Salvanes, 2001).

Finally, in order to control also for the issue of the endogeneity between profits and wages (due to simultaneous determination and to possible measurements error) we also apply an IV strategy. The literature has stressed that in case of endogeneity the (attenuation) bias in the cross-sectional estimates can be severe, and may also be aggravated by a fixed effects strategy (Card et al., 2013).

Therefore, we use a very recently developed estimation strategy of IV quantile fixed effects estimates (Galvao, 2011, and Galvao and Montes-Rojas, 2010), which is an extension of the IV quantile procedure of Chernozukov and Hansen (2008) that allows for the inclusion of fixed effects as introduced in Koenker (2004).[15] As instrument we exploit the idea developed in Card et al. (2013) by using a weighted average of the firm sales per employee in other provinces of Italy but in the same three-digit industry of the firm considered. The weights are the inverse of the distance between provinces. The idea is that industry sales, which represent industry demand shocks, affect the profitability of the firms while, at the same time, they are not correlated with local labor market conditions since they concern firms in other provinces of Italy.

4.2 Results

Table 2 shows the cross-sectional quantile estimates of the impact of profits per employee on workers' wages, by using the minimum wage as a measure for the opportunity cost of labor.

[15] For a detailed description of the procedures applied see the methodological annex and Canay (2011), Galvao (2011) and Galvao and Montes-Rojas (2010).

Table 2: Cross Sectional Quantile Regressions of Wages on Quasi Rents, with Control on First Level of Bargaining.

	q10	q25	q50	q75	q90
Ln Quasi Rent	0.069***	0.059***	0.057***	0.054***	0.047***
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]
Ln Minimum Wage	1.449***	1.539***	1.659***	1.755***	1.775***
	[0.009]	[0.005]	[0.007]	[0.009]	[0.018]
Female	-0.086***	-0.076***	-0.091***	-0.120***	-0.150***
	[0.002]	[0.001]	[0.001]	[0.001]	[0.002]
Age	0.009***	0.010***	0.011***	0.012***	0.016***
	[0.001]	[0.001]	[0.001]	[0.001]	[0.002]
Age Squared	-0.000***	-0.000***	-0.000***	-0.000***	-0.000***
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Tenure 3-10	0.057***	0.041***	0.025***	0.015***	0.009***
	[0.002]	[0.001]	[0.001]	[0.002]	[0.002]
Tenure >10	0.089***	0.065***	0.042***	0.025***	0.015***
	[0.003]	[0.002]	[0.002]	[0.003]	[0.004]
White Collar and Manager	0.061***	0.052***	0.061***	0.086***	0.131***
	[0.002]	[0.001]	[0.002]	[0.002]	[0.004]
ln Firm Size	0.014***	0.015***	0.015***	0.012***	0.009***
	[0.000]	[0.000]	[0.000]	[0.000]	[0.001]
Constant	-3.128***	-3.445***	-4.000***	-4.363***	-4.354***
	[0.061]	[0.040]	[0.043]	[0.058]	[0.103]
Area, Time and Sector dummies	yes	yes	yes	yes	yes
N. Observations	123,178	123,178	123,178	123,178	123,178
N. Individuals	25,796	25,796	25,796	25,796	25,796
R squared	0.35	0.38	0.41	0.43	0.44

Notes: ***,** and * denote significance at 1%, 5% and 10% respectevely.

The main relevant variable, rent sharing, displays a non-uniform impact along the wage distribution. In particular, elasticity estimates turn out to stand at 6.9% at the 10th percentile, 5.7% at the median and 4.7% at the 90th percentile.[16] Moreover, since these elasticities have been computed by controlling for the importance of the first (national) level of bargaining, they suggest that there is a non-negligible rent sharing that

[16] In this literature it is quite standard to provide a measure of the "Lester" range. The "Lester" range is defined as the elasticity of wages with respect to quasi-rent multiplied by four times the ratio between the standard deviation of quasi-rent and mean quasi-rent (Lester, 1952). It provides a measure of how much the wage of a worker increases moving from a firm at the bottom of the profit distribution (two standard deviations below the mean) to a firm at the top of the profit distribution (two standard deviations above the mean). In this paper we are unable to provide measures for the "Lester" range, since we are working with quantiles and not with average wages. Nonetheless, we provide a computation of the Lester range, which amounts to 19%, based on the OLS estimates included in table A1 in the appendix.

essentially takes place at the individual, local or firm level (consistently with Van Reenen, 1996).[17]

The cross sectional standard quantile regressions are likely to be biased since they do not take into account the workers' unobserved heterogeneity. Therefore we run quantile fixed effects estimates (Canay, 2011), enabling the introduction of fixed effects in the estimation, in such a way as to capture time invariant worker characteristics such as ability and education. Table 3 shows the results. The estimates change significantly: the coefficients are much reduced in magnitude (around 60%) and are still slightly decreasing along the wage distribution.

These results are consistent with previous empirical evidence showing that taking into account the unobserved worker heterogeneity entails a sharp reduction in the estimated degree of rent sharing (see for instance Card et al., 2013, Martins 2009).

Finally, we present the IV estimates to tackle the endogeneity between rents and wages; in fact, endogeneity can cause serious underestimation of the degree of rent sharing, which can also be worsened by a fixed effects strategy (Card et al., 2013). The estimation was carried out simultaneously on three percentiles (10^{th}, 50^{th}, 90^{th}) for computational reasons. Moreover, since it is not possible to test the weakness of the instrument in this procedure, we carried out a standard IV fixed effects estimation on average wages (see table A1 in the appendix), checking the first stage F-statistics. The F-value for the instrument in the first stage is significant and higher than the threshold value of 10, confirming that the instrument chosen is not weak.

When endogeneity is taken into account, the results change significantly (Table 4). In fact, the elasticities of rents with respect to wages are now greater, and the highest increases are to be seen in the lower tail of the wage distribution. In particular, rents show a decreasing impact along the wage distribution with elasticities ranging from 7.3% at the 10th percentile to 4.8% at the median and to a 3.7% at the 90^{th} percentile. These estimates are consistent with those of Card et al. (2013), who find an elasticity of (average) wages with respect to rents of 4.5% for Veneto in Italy.

[17] As for the control variable in the estimation, the results are as follow: the impact of minimum wage is positive and increasing along all the wage distribution and its elasticity is higher than 1, meaning that an increase in the minimum wage implies a more than proportional increase in the corresponding worker's wage; the age coefficients show a concave pattern, which is increasing along the wage distribution; the gender wage gap is higher at the highest percentiles; the return to tenure is positive and decreasing along the wage distribution and the occupation dummy is positive and increasing, highlighting higher wages for higher occupation categories; the firm size has only a slightly decreasing impact along the wage distribution.

Table 3: Quantile Fixed Effects Regressions of Wages on Quasi Rents.					
	q10	**q25**	**q50**	**q75**	**q90**
Ln Quasi Rent	0.027***	0.021***	0.020***	0.020***	0.019***
	[0.001]	[0.000]	[0.000]	[0.000]	[0.001]
Ln Minimum Wage	0.855***	0.864***	0.880***	0.899***	0.914***
	[0.006]	[0.003]	[0.003]	[0.004]	[0.006]
Age	0.040***	0.037***	0.035***	0.033***	0.030***
	[0.001]	[0.000]	[0.000]	[0.000]	[0.001]
Age Squared	-0.000***	-0.000***	-0.000***	-0.000***	-0.000***
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Tenure 3-10	0.048***	0.028***	0.012***	0.002***	-0.011***
	[0.002]	[0.001]	[0.001]	[0.001]	[0.001]
Tenure >10	0.054***	0.027***	0.009***	-0.005***	-0.026***
	[0.002]	[0.001]	[0.001]	[0.001]	[0.002]
White Collar and Manager	0.047***	0.049***	0.048***	0.049***	0.056***
	[0.002]	[0.001]	[0.001]	[0.001]	[0.002]
ln Firm Size	0.009***	0.009***	0.010***	0.011***	0.011***
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Constant	-0.097***	0.004	0.025	0.009	0.060*
	[0.033]	[0.016]	[0.016]	[0.018]	[0.036]
Area, Time and Sector dummies	yes	yes	yes	yes	yes
N. Observations	123,178	123,178	123,178	123,178	123,178
N. Individuals	25,796	25,796	25,796	25,796	25,796
R squared	0.42	0.49	0.53	0.52	0.48

Notes: ***,** and * denote significance at 1%, 5% and 10% respectevely.

This evidence suggests that once having controlled for the national centralized level of bargaining, rent sharing in Italy is such as to favor unskilled workers.[18] This finding is consistent with the idea that in Italy the unions are relevant not only at the national level, but also at the local/firm level. Moreover, this result is also in line with Bagger et al. (2013) who, using a structural matching model, have shown that the workers' bargaining power decreases slightly with the education level. Similar findings are

[18] By dividing firms with respect to the quartiles of the profit distribution and workers with respect to the quartiles of the wage distribution, we find evidence that high paid workers are mostly employed in high profits firms, while low paid workers are mostly employed in low profits firms. It is also interesting to note that the rate of growth of profits is not the same in the four quartiles of the profits distribution. Our descriptive analysis (available upon request) shows that the firms that enjoy higher growth rates in profits are those in the top quartile of profits (on average 8% per year). Combined with previous results, this evidence suggests that low-skilled workers are characterized by a higher degree of rent sharing than high-skilled workers, but at the same time they are employed in firms which experience relatively lower growth rates, thus partially balancing out (in cumulative terms) their greater rent sharing elasticities.

derived by Guertzen (2009) and Kohn and Lembcke (2007), who find that rent sharing is greater for blue collar workers.

Table 4: IV Quantile Fixed Effects Regressions of Wages on Quasi Rents.			
	q10	**q50**	**q90**
ln Quasi Rent	0.073***	0.048***	0.037***
	[0.002]	[0.001]	[0.002]
ln Minimum Wage	0.847***	0.889***	0.963***
	[0.004]	[0.002]	[0.005]
Age	0.037***	0.030***	0.027***
	[0.001]	[0.000]	[0.001]
Age Squared	-0.000***	-0.000***	-0.000***
	[0.000]	[0.000]	[0.000]
Tenure 3-10	0.053***	0.010***	-0.015***
	[0.001]	[0.000]	[0.001]
Tenure >10	0.057***	0.008***	-0.029***
	[0.001]	[0.000]	[0.001]
White Collar and Manager	0.034***	0.033***	0.040***
	[0.001]	[0.000]	[0.001]
ln Firm Size	0.011***	0.013***	0.014***
	[0.000]	[0.000]	[0.000]
Constant	16.109***	16.375***	16.205***
	[0.041]	[0.015]	[0.027]
Area, Time and Sector dummies	yes	yes	yes
N. Observations	123,178	123,178	123,178
N. Individuals	25,796	25,796	25,796

Notes: ***,** and * denote significance at 1%, 5% and 10% respectevely. The instruments are the linear projections of other provinces average sales per employee on the endogeneous variables.

4.3 The role of unions

In order to provide evidence supporting the intuition that unions contribute to have a decreasing rent sharing impact along the wage distribution it is not possible to use individual information on unions membership since in many European countries, including Italy, this is a sensitive data information. For such a reason, we resort to regional data, splitting the sample according to information on unions membership at the regional level, which can be considered as a proxy of unions regional strength. We also focus on manufacturing, since the literature has shown that unions in this sector are supposed to be more organized and with greater power (see for instance Booth, 1995, Disney, 1990).[19] We split the sample according to the median of unions membership rate computed for all (20) Italian regions.[20]

Estimates for the 10th, 50th, and 90th percentiles distinguished by regions characterized by high union power (above the median) and low union power (below the median) are shown in Table 5. As first remark, it is worth noting that estimates are always greater than those derived for the whole sample, suggesting that rent-sharing is as expected greater in manufacturing, where unions are supposed to have greater power. When comparing the two samples, above and below the median of the regional union power, it emerges that when union power is low estimates along the distribution are only slightly decreasing, while the ones in regions with high union power are strongly decreasing. Further, where union power is high, estimates are greater at the 10th and the 50th percentiles with respect to estimates in regions with low union power. In particular, differences are equal to 2 percentage points at the 10th percentile (12.8 vs 10.8) and to 0.7 (10 vs 9.3) at the 50th percentile, suggesting that the role of unions is stronger for unskilled than for medium skilled workers.[21] Interestingly, the coefficient estimate at the 90th percentile is much lower in regions with high union power than in those with low union power (4.6 vs 9.2), i.e., skilled workers receive a higher amount of rent-sharing when unions have low power. This might be explained by the fact that when unions have low power skilled individuals can capture higher rents through individual bargaining, without the mediation of unions, while when unions have high power skilled workers either have lower power in individual bargaining or the amount of their rent sharing is still bargained, at least partially, by unions, which favor unskilled workers.

This evidence confirms the intuition that labor market institutions, and in particular unions, contribute, at least in part, to the heterogeneous extent of rent sharing along the wage distribution, suggesting that unions favor the rent sharing of low and medium skilled workers.

---

[19] The literature has also focused on the decline in manufacturing as a possible determinant of the fall in union density in the last decades (for the UK case Disney, 1990, among others).

[20] Data have been derived online from the unions' website and relate to the percentage of members in the main Italian unions (CGIL and CISL) over total dependent workers (data from Italian regional accounts) in each Italian region in 2003. To compute the shares for the manufacturing sector we have relied on the share system computed by Visser ed Ebbinghaus (1999).

[21] Note also that the difference in coefficients at the 10th percentile is statistically different from zero at 10% level, while this is not the case for the difference at the median.

Table 5: Quantile IV Fixed Effects Regressions of Wages on Quasi Rents. Manufacturing by degree of unionization.

	High Union Power			Low Union Power		
	q10	**q50**	**q90**	**q10**	**q50**	**q90**
Ln Quasi Rent	0.128***	0.100***	0.046***	0.108***	0.093***	0.092***
	[0.007]	[0.007]	[0.012]	[0.008]	[0.012]	[0.008]
Ln Minimum Wage	0.945***	0.961***	1.035***	0.925***	0.969***	1.030***
	[0.010]	[0.006]	[0.013]	[0.013]	[0.007]	[0.008]
Age	0.029***	0.024***	0.022***	0.034***	0.026***	0.022***
	[0.001]	[0.001]	[0.001]	[0.002]	[0.001]	[0.001]
Age Squared	-0.000***	-0.000***	-0.000***	-0.000***	-0.000***	-0.000***
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Tenure 3-10	0.041***	0.008***	-0.008***	0.040***	0.008***	-0.016***
	[0.002]	[0.001]	[0.002]	[0.003]	[0.001]	[0.002]
Tenure >10	0.051***	0.010***	-0.014***	0.040***	-0.000	-0.032***
	[0.002]	[0.001]	[0.002]	[0.004]	[0.002]	[0.002]
White Collar and Manager	0.018***	0.021***	0.032***	0.075***	0.074***	0.085***
	[0.003]	[0.001]	[0.002]	[0.004]	[0.001]	[0.002]
ln Firm Size	0.017***	0.021***	0.023***	0.017***	0.018***	0.018***
	[0.001]	[0.001]	[0.001]	[0.002]	[0.001]	[0.001]
Const	17.158***	17.348***	17.241**	5.636***	5.747***	5.620***
	[0.062]	[0.033]	[0.059]	[0.093]	[0.041]	[0.044]
Area, Time and Sector dummies	yes	yes	yes	yes	yes	yes
N. Observations	26,016	26,016	26,016	45,037	45,037	45,037
N. Individuals	5,756	5,756	5,756	8,928	8,928	8,928

Notes: ***,** and * denote significance at 1%, 5% and 10% respectevely. The instruments are the linear projections of other provinces average sales per employee on the endogeneous variables.

5. Conclusions

The innovative contribution of this paper is to analyze the degree of rent sharing along the wage distribution. Previous empirical analyses focused only on average wages. In some cases attention have been paid to average wages of workers' groups defined using education and/or occupation categories, which however does not allow taking into account the substantial heterogeneity within workers' groups. In this paper we address this issue by using quantile regressions, since percentiles of the wage distribution can be more closely associated to the productivity of workers in the labor market.

We make use of a unique employer-employee database for Italy, which merges administrative records for workers (INPS) and balance sheet data for firms (AIDA). Our findings show that the rent sharing impact is not uniform along the wage distribution. In particular, taking into account the first national level of bargaining, unobserved heterogeneity and endogeneity, we find a decreasing pattern of rent sharing along the wage distribution, with elasticities of wages with respect to quasi-rents ranging from 7.3% at the 10^{th} percentile to 3.7% at the 90^{th} percentile of the distribution. One of the possible explanations for this finding refers to the role of the unions in protecting the lowest paid worker categories, since in Italy, as in other European countries, the unions play a crucial role in the bargaining process between employers and employees, while individual bargaining is less important. We provide evidence in favor of this possible explanation exploiting the regional variability in union membership, which can be considered as a proxy for union power. We show that where union power is high the extent of rent sharing is higher for unskilled and medium skilled workers, while it is lower for skilled workers.

References

Abowd J., Lemieux T. (1993), 'The effects of Product Market Competition on Collective Bargaining Agreements: The case of Foreign Competition in Canada', *Quarterly Journal of Economics*, 108, 983-1014.

Arai M. (2003), 'Wages, Profits and Capital Intensity: Evidence from Matched Worker-Firm Data', *Journal of Labor Economics*, 21, 593-618.

Arai M., Heyman F. (2001), 'Wages, profits and individual unemployment risk: evidence from matched firm-worker data', *FIEF Working Paper* series, n.172.

Bagger J., Fontaine F., Postel-Vinay F., Robin J.M., (2013) 'Tenure, Experience, Human Capital and Wages: A Tractable Search Equilibrium Model of Wage Dynamics', available at https://sites.google.com/site/jmarcrobin, forthcoming in *American Economic Review*.

Blanchflower D., Oswald A.., Sanfey P. (1996), 'Wages, Profits and Rent Sharing', *Quarterly Journal of Economics*, 111, 227-252.

Booth, A. (1995), 'The Economics of Trade Unions', Cambridge University Press.

Canay I., (2011), 'A Simple Approach to Quantile Regression for Panel Data', *The Econometrics Journal*, 14(3), 368-386.

Card D., Devicienti F., Maida A. (2013), 'Rent Sharing, Hold Up and Wages: Evidence of Matched Panel Data', forthcoming in *The Review of Economic Studies*. DOI: 10.1093/restud/rdt030.

Chernozukov V., Hansen C., (2008), 'Instrumental Variable Quantile Regression: A Robust Inference Approach', *Journal of Econometrics*, 142, 379-398.

Disney, R. (1990), 'Explanations of the Decline in Trade Union Density In Britain: an Appraisal', *British Journal of Industrial Relations*, 28(2), 165–177.

Ebbinghaus, B. e Visser, J. (1999), 'Trade Unions in Western Europe since 1945', London: MacMillan/Grove.

Galvao A., (2011), 'Quantile Regression for Dynamic Panel Data with Fixed Effects', *Journal of Econometrics*, 164, 142-157.

Galvao A., Montes-Rojas G., (2010), 'Penalized Quantile Regression for Dynamic Panel Data', *Journal of Statistical Planning and Inference,* 140, 3476-3497.

Guertzgen N. (2009), 'Rent-Sharing and Collective Bargaining Coverage – Evidence from Linked Employer-Employee Data', *Scandinavian Journal of Economics,* 111(2), 323-349.

Guiso L., Pistaferri L., Schivardi F., (2005) 'Insurance within the Firm' *Journal of Political Economy*, University of Chicago Press, 113(5), 1054-1087.

Hildreth A., Oswald A. (1997), 'Rent Sharing and Wages: Evidence from Company and Establishment Panels', *Journal of Labor Economics*, 15, 318-337.

Koenker R., Bassett G., (1978), 'Regression Quantiles', *Econometrica*, 46(1), 33-50.

Koenker R., (2004), 'Quantile Regression for Longitudinal Data', *Journal of Multivariate Analysis*, 91, 74-89.

Kohn, K. and Lembcke, A. (2007), 'Wage Distributions by Bargaining Regime. Linked Employer–Employee Data Evidence from Germany', *AStA Wirtschafts- und Sozialstatistisches Archiv* 1, 247–261.

Krueger A.B., Summers L.H. (1988), 'Efficiency Wages and the Inter-Industry Wage Structure', *Econometrica*, 56(2), 259-293.

Lemieux T., MacLeod B., Parent D., (2009), 'Performance Pay and Wage Inequality', *Quarterly Journal of Economics*, 124(1), 1-49.

Lester R.A., (1952), 'A Range Theory of Wage Differentials', *Industrial and Labor Relations Review,* 5, 483-500.

Margolis D., Salvanes K. (2001), 'Do Firms Really Share Rents with Their Workers?', *IZA Discussion Paper* 330.

Martins P.S. (2007), 'Rent Sharing and Wages', *Reflets et Perspectives de la Vie Economique*, 46(2/3), 23-31.

Martins P.S. (2009), 'Rent Sharing Before and After the Wage Bill', *Applied Economics*, 41(17), 2133-2151.

Matano A., Naticchioni P., (2012), 'Wage Distribution and the Spatial Sorting of Workers', *Journal of Economic Geography*, 12(2), 379-408.

Nickell S.J., (1981), 'Biases in Dynamic Models with Fixed Effects', *Econometrica*, 49, 1417-1426.

Pencavel J., Pistaferri L., Schivardi F. (2006), 'Wages, Employment and Capital in Capitalist and Worker Owned Firms', *ILR Review*, 60(1), 23-44.

Pistoresi B., Strozzi C. (2001), 'Rent Sharing in Wage Determination. Evidence from Italy' *CEPR Discussion Paper* 2939.

Rusinek M., Rycx F. (2013), 'Rent-Sharing under Different Bargaining Regimes: Evidence from Linked Employer-Employee Data', *British Journal of Industrial Relations*, 51(1), pp. 28-58.

Shapiro C., Stiglitz J. (1984), 'Equilibrium Unemployment as a Worker Discipline Device' *American Economic Review*, 74, 433-444.

Van Reenen J. (1996), 'The Creation and Capture of Rents: Wages and Innovation in a Panel of U.K. Companies', *Quarterly Journal of Economics*, 111, 195-226.

Appendix

Table A1: OLS Regressions of Avearge Wages on Quasi Rents			
	(1)	(2)	(3)
	Cross Section	Fixed Effects	IV-Fixed Effects
Ln Quasi Rent	0.063***	0.022***	0.056***
	[0.007]	[0.010]	[0.011]
Ln Minimum Wage	1.575***	0.885***	0.905***
	[0.007]	[0.010]	[0.011]
Female	-0.114***		
	[0.001]		
Age	0.013***	0.035***	0.034***
	[0.001]	[0.001]	[0.001]
Age Squared	-0.000***	-0.000***	-0.000***
	[0.000]	[0.000]	[0.000]
Tenure 3-10	0.033***	0.016***	0.016***
	[0.001]	[0.001]	[0.001]
Tenure >10	0.050***	0.011***	0.012***
	[0.002]	[0.002]	[0.002]
White Collar and Manager	0.098***	0.049***	0.048***
	[0.002]	[0.003]	[0.003]
ln Firm Size	0.013***	0.010***	0.013***
	[0.000]	[0.001]	[0.001]
Constant	-3.658***	-0.028	-0.198***
	[0.046]	[0.057]	[0.064]
Area, Time and Sector dummies	yes	yes	yes
N. Observations	123,178	123,178	123,178
N. Individuals	25,796	25,796	25,796
R squared	0.63	0.19	0.17
F Test Instrument First Stage			2,108.46

Notes: ***,** and * denote significance at 1%, 5% and 10% respectevely.

METHODOLOGICAL ANNEX

The quantile regression methodologies

Standard quantile regressions (Koenker and Bassett, 1978), can be expressed as follows:

$$(1) \quad \ln(w_i) = X_i^{'}\beta(\theta) + u_{i,\theta}$$

where $i=1,\ldots n$ is the observation, θ is the quantile analyzed, $u_{i,\theta}$ is an idiosyncratic error term, $ln(w_i)$ is the dependent variable (logarithm of wages) and X represents the set of explanatory variables. As is the standard practice in this literature, quantile regression coefficients can be estimated by means of the approach proposed by Koenker and Basset (1978), where $\beta(\theta)$ solves the following minimization problem:

$$(2) \quad \min_{\beta}\left[\left(\sum_{i=1}^{n}\rho_{\theta}(\ln(w_i)-X_i^{'}\beta(\theta))\right)\right]$$

$$\text{where } \begin{cases}\rho_{\theta}(u)=\theta u \text{ if } u>0\\ \rho_{\theta}(u)=(\theta-1)u \text{ if } u<0.\end{cases}$$

To take unobserved heterogeneity issues into account quantile fixed effects estimates can be performed by means of two different procedures.

The first developed procedure is the technique elaborated by Koenker (2004), who estimates quantile regressions adding individuals' dummies in the estimation. Moreover, Koenker (2004) adds to the minimization algorithm a penalty term that takes into account the computational problem arising when estimating such a large number of parameters.[22] This technique minimizes the following expression:

$$(3) \quad \min_{a,\beta}\sum_{k=1}^{q}\sum_{i=1}^{n}\sum_{j=1}^{t_i}\xi_k\rho_{\theta_k}(\ln(w_{ij})-a_i-X_{ij}^{'}\beta(\theta_k))+\lambda\sum_{i=1}^{n}|a_i|$$

where k is the index for the chosen quantiles, i is the index for the (n) individuals, j is the index for the observations per individual (from 1 to t_i), and $\rho_{\theta k}(u)$ is defined as in equation (2). This technique requires the simultaneous estimation of the chosen quantiles, since individuals' fixed effects are assumed to be constant across quantiles to reduce the number of parameters estimated. The weights ξ_k control for the relative influence of the k quantiles on the estimation of the a_i parameters. The last term in the

[22] Indeed, Koenker (2004) claims that the use of the penalty term is necessary since the large number of individual fixed effects can increase the variability of the estimates of the covariates.

above expression represents the penalty term, where λ describes the importance of the penalty term in the minimization formula.[23]

The second quantile fixed effects procedure is developed by Canay (2011). The starting point of the Canay (2011) procedure is the following conditional mean equation:

$$(4) \qquad \ln(w_{ij}) = X_{it}^{'}\beta(\theta_{\mu}) + a_i + u_{ij}$$

where $E(u_{ij} \mid X_i, a_i) = 0$ and i=1….n, j=1…t_i.

Equation (4) implies that the individual fixed effect a_i is present in the conditional mean of $ln(w_{ij})$. Therefore, from eq.(4) it is possible to compute a $\sqrt{T}$-consistent estimator of a_i given a $\sqrt{nT}$ –consistent estimator of $\beta(\theta_{\mu})$. In such a framework, Canay (2011) proposes a two-step estimator. The first step consists in defining the individual fixed effect $\hat{a}_i$ as: $\hat{a}_i \equiv E_T(\ln(w_{it}) - X_{it}^{'}\hat{\beta}(\theta_{\mu}))$ where $\hat{\beta}(\theta_{\mu})$ is a $\sqrt{nT}$-consistent estimator of $\beta(\theta_{\mu})$ (for instance obtained by a standard fixed effects estimation). In the second step it is then possible to define a new dependent variable $\ln(\hat{w}_{it}) = \ln(w_{it}) - \hat{a}_i$, and the two-step estimator $\hat{\beta}(\theta)$ is the one that solves the following minimization problem:

$$\min_{\beta} E_{nT} \left\lfloor \rho_{\theta_{\tau}}(\ln(\hat{w}_{ij}) - X_{ij}^{'}\beta(\theta) \right\rfloor$$

Canay (2011) shows that this estimator is consistent and asymptotically normal under some regularity conditions (see Canay, 2011, for further details).

These two estimations methodologies usually provide very similar results.

To control for the endogeneity bias that can arise by simultaneity in the individual choices regarding locations and wages, we can make use of IV quantile fixed effects estimation. 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). The methodology has been presented in Galvao and Montes-Rojas (2009, 2010), Galvao (2011), and Harding and Lamarche (2009). In particular, we follow Galvao and Montes-Rojas (2009, 2010), who extend the framework allowing the fixed effects to be the same across quantiles. The model we consider is thus the following:

$$(5) \qquad ln(w_{ij}) = X_{ij}^{'}\beta(\theta_k)) + d_{ij}^{'}\delta(\theta_k) + a_i + u_{ij,\theta_k}$$

[23] It is worth noting that if λ is equal to zero a generic quantile fixed effects estimator is derived (the penalty term disappears), while if λ tends to infinity the a_i goes to zero for all i, ending up with an estimate of the model with no fixed effects. Koenker (2004) shows the consistency of this estimation technique, while standard errors can be computed by bootstrap estimations (see Koenker, 2004, for further details). Moreover, because of the longitudinal dimension of the data it is necessary to use bootstrapping over random samples (with replacement) of individuals instead of over random samples of observations, as also done in Abrevaya and Dahl (2008) and Bache et al. (2013).

where $$d_{ij} = f(X_{ij}, g_{ij}, v_{ij})$$

and i=1….n, j=1…t_i.

The first expression in (5) shows that the dependent variable is a function of the exogenous variables X_{ij}, the endogenous variables d_{ij}, a vector of fixed effects a_i and an error term $u_{ij,\theta k}$. The second expression in (5) shows that the vector of endogenous variables d_{ij} is a function of the exogenous variables X_{ij}, a vector of instrumental variables g_{ij} uncorrelated with the error term $u_{ij,\theta k}$, and an error term v_{ij} stochastically dependent on $u_{ij,\theta k}$. In this framework the objective function of the model for a given quantile k is:

$$(6) \qquad \sum_{i=1}^{n}\sum_{j=1}^{t_i} \rho_{\theta_k}(\ln(w_{ij}) - d_{ij}^{'}\delta(\theta_k) - X_{ij}^{'}\beta(\theta_k) - \alpha_i - \hat{g}_{ij}^{'}\gamma(\theta_k))$$

where $\hat{g}_{ij}$ is the least square projection of the endogenous variables d_{ij} on the instruments g_{ij} and the X_{ij} (as suggested in Chernuzhukov and Hansen, 2008, Galvao and Montes-Rojas, 2009, 2010, Galvao, 2011, and Harding and Lamarche, 2009), and the other variables are expressed as in (3). The idea underlying the model is that, in order for $\hat{g}$ to be a good instrument it should be uncorrelated with the error term and therefore it should have a zero coefficient in (6). Thus, for given parameters of the endogenous variables (δ), the quantile fixed effects regression of $(ln(w_{ij})\text{-}d_{ij}\,\delta)$ on $(x_{ij}, a_i, \hat{g}_{ij})$ should generate a zero coefficient (γ) for the variable $\hat{g}$.

From a practical point of view, minimization proceeds in two steps: first, for a given set of δ, equation (6) is minimized with respect to (β, a, γ), deriving estimates of the parameters as function of δ, i.e. $\beta(\delta)$, $a(\delta)$, $\gamma(\delta)$. A consistent estimate for the coefficient of the endogenous variable is then obtained by selecting the value of δ that minimizes a weighted distance function defined on γ:

$$(7) \qquad \hat{\delta} = \min_{\delta} \hat{\gamma(\delta)}' A \hat{\gamma(\delta)}$$

for a given positive definite matrix A. This estimator has been proved to be asymptotically normal and, as mentioned, the estimation can be performed for more quantiles simultaneously.[24]

[24] Note that standard errors are derived from the estimation of a heteroskedasticity consistent variance-covariance matrix. See Galvao and Monte-Rojas (2009, 2010), Galvao (2011), Chernozhukov and Hansen (2008), for further details on the estimation technique and its properties.

Additional References for the methodological annex

Abrevaya, J., Dahl, C. (2008), The effects of Birth Inputs on Birth-weight: Evidence from Quantile Estimation on Panel Data, *Journal of Business and Economic Statistic*, 26(4): 379-97.

Bache, S.H., Dahl C. and Tang Kristensen J., (2013), Headlights on Tobacco Road to Low Birthweight Outcomes, *Empirical Economics*, 44(3), 1593-1633.

Galvao, A., Montes-Rojas, G. (2009), Instrumental Variables Quantile Regression for Panel Data with Measurement Errors, Department of Economics Discussion Paper Series 09/06, City University London.

Harding, M., Lamarche, C. (2009), A Quantile Regression Approach for Estimating Panel Data Models Using Instrumental Variables, *Economic Letters*, 104: 133-135.