Understanding Italian inequality trends*

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Abstract

This paper develops a unifying framework for analysing the effects of: (i) the changing distribution of individual incomes by main factor sources, (ii) the increasing participation of wives in the labour force and (iii) the changing distribution of family types on the distinctive trends towards inequality in equivalent household income in Italy between 1977 and 2004.

Changes in the distribution of work and pension incomes explain most of the trend. The higher average likelihood of wage-earning wives had an unequalising effect on households on the left tail of the income distribution. Little is explained by the changing distribution of family types.

JEL: D31, D63, C51 **Keywords:** Inequality trends, equivalent income, counterfactual analysis, Italy.

1 Introduction

According to recent comparative studies on OECD countries, the highest level of equivalent income inequality is found in the US, followed by the UK and Italy (Atkinson et al., 1995; Brandolini and Smeeding, 2008). However, while

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the US and UK have presented a roughly increasing trend of equivalent income inequality since the end of the 1970s and for at least twenty years, the Italian trend is unusual. Having begun by gradually decreasing up to the end of 1980s reaching its minimum in 1991, it rose dramatically in the early 1990s. Since then it has remained roughly constant until the mid 2000s.

Several economic hypotheses can be invoked to explain the interesting change in direction of inequality trends in Italy. Since the end of the 1970s, Italy has undergone several socio-demographic changes: the Italian population has grown older, living in smaller families, with fewer children and more working adults. During the same period, there were changes in the macroeconomic and institutional frameworks with likely impacts on the distribution of individual disposable income. However, rather than looking at a whole set of possible causes,[1] this paper focuses mainly on three: (i) the changing distribution of individual incomes, distinguished by main source (dependent employment, self-employment and pension), (ii) the increasing participation of wives in the labour force, (iii) the effect of the changing distribution of family types on trends in equivalent household income inequality. These factors have been chosen for their preeminent importance in inequality analysis and for their important behavioural links.

Available methodologies for inequality decomposition based on the mathematical properties of inequality measures or on the estimation of data generating processes are either unable to jointly analyse the role of socio-demographic changes in the population and of changes in income source distribution changes or are not robust or are too demanding in terms of data requirements.

This paper suggests that an analysis of Italian household income inequality trends should *jointly* test for the above-mentioned three factors and, by applying and extending Daly and Valletta (2006), it proposes a unifying framework by developing counterfactuals, i.e. answering "what if" questions. The methodology used is a combination of the rank-based methodology proposed by Burtless (1999) to address issue (i) and the conditional estimation based on propensity score methods made popular by DiNardo et al. (1996) to address issues (ii) and (iii). This methodology will be applied to the evolution of Italian household income inequality between 1977 and 2004.

By applying this methodology, it is possible to disentangle the main forces determining the distinctive pattern of household income inequality in Italy. In particular, this paper highlights the importance of studying the distribution of self-employment and pension income and not only that of wage income for understanding household income inequality trends. In fact, changes in the distributions of work and pension incomes account for most of the trend, but the changing distribution of pension income had an equalising effect across the whole period while work income distribution shows a widening gap between below- and above-the-median income quantiles. Changes in the self-

[1]Among several contributions addressing the causes of wage and income inequality, see Topel (1997); Gottschalk and Smeeding (1997); Acemoglu (2002).

employment distribution had a similar impact to employment on total household income, although self-employment accounts for less than one third of total work income, so its relative role is larger. In contrast to other countries, the increased probability of earning wives is found to have had an unequalising effect for households in the left tail of the income distribution. As for the upper tail, at first it contributed to reducing inequality but then, after 1991, to increasing it. However, little is explained by the changing distribution of family types. This methodology could also prove useful for the analysis of inequality in other countries.

The structure of the paper is as follows. Section 2 describes the methodology adopted here. Section 3 describes the data set, defines the measure of income used, and presents some figures on trends in Italian inequality. Section 4 presents the results and Section 5 the conclusions.

2 Methodology

The methodology used here, first proposed by Daly and Valletta (2006), combines the Burtless (1999) approach with that of DiNardo et al. (1996). The former is used to account for the changing distribution of individual income sources and the latter for the effect of changing household characteristics on equivalent income distribution. Here, the notation follows closely that of DiNardo et al. (1996) (henceforth, DFL), although some complications arise due to the consideration of household equivalent income instead of individual income.

Let us first view each individual observation of equivalent household income as a vector of equivalent income, y, a set of household characteristics, λ, - comprising family type, π, participation of wives in the labour force, ϕ, and other family characteristics, μ - and time, t, belonging to the joint distribution $F(y, \lambda, t)$. Since equivalent household income is obtained as a transformation of the sum of the total individual incomes, z, of all members of the same household, the distribution of equivalent household income at time t can also be represented as $F(z, \theta, \lambda, t)$, where θ stands for the household composition parameter which combines individual income observations into households and transforms their incomes into equivalent household income. Finally, as total individual income is assumed to come from at most three different sources (employment, z^e, self-employment, z^s, and pension, z^p), the distribution of equivalent household income can be seen as a vector $(z^e, z^s, z^p, \theta, \lambda, t)$ belonging to the distribution $F(z^e, z^s, z^p, \theta, \lambda, t) \equiv F(y, \lambda, t)$. The joint distribution of incomes and household characteristics at one point in time is the conditional distribution $F(z^e, z^s, z^p, \theta, \lambda|t) \equiv F(y, \lambda|t)$. The density of equivalent household income at one point in time, $f_t(y) \equiv f_t(z^e, z^s, z^p, \theta)$, can be written as:

$$f_t(y) \equiv f(y; t_y = t, t_\lambda = t) = \int_{\lambda \in \Omega_\lambda} f(y|\lambda, t_y = t) dF(\lambda | t_\lambda = t) \quad (1)$$
$$= \int_{\lambda \in \Omega_\lambda} f(z^e, z^s, z^p, \theta | \lambda, t_{e,s,p,\theta} = t) dF(\lambda | t_\lambda = t) \equiv f(y; t_{e,s,p,\theta} = t, t_\lambda = t)$$

where Ω_λ is the domain of definition of household characteristics, $t_e \equiv t_{z^e}, t_s \equiv t_{z^s}, t_p \equiv t_{z^p}$. While $f(y; t_y = 04, t_\lambda = 04) \equiv f(y; t_{e,s,p,\theta} = 04, t_\lambda = 04)$ represents the actual density of equivalent incomes in 2004, $f(y; t_y = 04, t_\lambda = 91) \equiv f(y; t_{e,s,p,\theta} = 04, t_\lambda = 91)$ represents the density of equivalent incomes that would have prevailed in 2004 had the distribution of household characteristics remained as it was in 1991.

Assuming that the 2004 structure of incomes, which is represented by the conditional density $f(z^e, z^s, z^p, \theta | \lambda; t_{e,s,p,\theta} = 04)$, does not depend on the distribution of family characteristics, we can write

$$f(y; t_{e,s,p,\theta} = 04, t_\lambda = 91) = \int f(z^e, z^s, z^p, \theta | \lambda, t_{e,s,p,\theta} = 04) dF(\lambda | t_\lambda = 91) \quad (2)$$

where the reweighting function is defined as $\psi_\lambda = dF(\lambda | t_\lambda = 91)/dF(\lambda | t_\lambda = 04)$ and can be estimated with regression models. Assuming further that $\lambda \equiv \{\phi, \pi, \mu\}$, the distribution of household characteristics can be factorised as $F(\lambda | t_\lambda = t) = F(\phi | \pi, \mu, t_{\phi|\pi,\mu} = t) \cdot F(\pi | \mu, t_{\pi|\mu} = t) \cdot F(\mu | t_\mu = t)$, and counterfactual densities holding either ϕ or π or both at a different time period, can be obtained following the methodology outlined by DFL.[2]

Similarly to what DFL highlighted (p. 1011), the density (2) should be called the "density that would have prevailed if household attributes λ had remained at their 1991 level *and* income earners had been paid their incomes according to the employment, self-employment and pension schedule existing in 2004 *and* household composition had remained as in 2004", as the counterfactual density (2) ignores the impact of changes in the distributions of z^e, z^s, z^p on the structure of incomes. Hence, this procedure would allow us to disentangle the effects of changing household characteristics on the change of equivalent income distribution, but would have nothing to say about the role of the changing distribution of individual incomes over the periods considered.

To investigate the role of the changing distribution of income sources, the distribution of a certain type of income in a year is replaced with the distribution of the same type of income in another year using a modified version of

[2] In theory, the DFL method could be extended by further partitioning the vector of household or individual characteristics and by estimating different weights. However, the main problem with such extensions is that small errors in estimating the conditional probabilities in the denominators of reweighting functions can produce potentially large errors in the weights (for a thorough discussion of this issue, see DiNardo, 2002).

the rank-based Burtless (1999) method (henceforth, the B method).

From a practical point of view, the B method applied for instance to the wage distribution works as follows. Let us assume that the yearly wage income of individual i in year $t = 1991$ is $(z_i^e|t_e = 91) \in p_{j,91}$, where $p_{j,91}$ is the j-th quantile interval (i.e. the interval between the $(j-1)$-th and j-th quantiles) of the wage distribution. The counterfactual wage earned by individual i in year $t = 2004$ is equal to $(z_i^e|t = 91) - m_{i,91}^e + m_{i,04}^e \times m_{91}^e/m_{04}^e$, where $m_{i,t}^e$ is the mean of the wages belonging to the same quantile interval of individual i in the year t's wage distribution and m_t^e is the overall mean wage income in year t, with $t = 1991, 2004$. The m_{91}^e/m_{04}^e normalisation is performed to preserve the overall mean income observed in year $t = 1991$,[3] while distribution by quantiles is used given that income vectors vary in size from year to year as the surveys are on different samples of households. The same procedure can be applied using other sources of income (i.e. self-employment, z^s, and pension income, z^p) instead of wages. By applying the same transformation to all individual wages in the sample at period $t = 1991$, we obtain the counterfactual distribution of wages which, depending on the assumptions adopted for the distribution of different types of income, can then be summed with actual or counterfactual self-employment and pension income vectors received within the same household in the year $t = 1991$, and eventually equivalised to compute the counterfactual equivalent household income vector. Using the B methodology, we turn a blind eye to the process that generated individual employment, self-employment and pension income, due to the fact that available data allows for an analysis of the determinants of wage income only.

It is important to highlight the assumptions underlying this method. In its current application, it is assumed that each person chooses in advance whether to work as an employee, self-employed, or not to work and, in the case of entitlement, to receive a pension income. This choice is irreversible, i.e. once he receives a wage income he will keep receiving a wage income in all counterfactuals developed with the B method. Secondly, conditional on the choice of income type, the employee (self-employed/pensioner) workforce in the x-th quantile of an income distribution is worth that value with no error and the empirical distribution function of a type of income coincides with its true schedule. Thirdly, while distributions of incomes are allowed to differ depending on sources (dependent employment/self-employment/pension), different distributions are not allowed for males and females or for male household heads (as in Daly and Valletta, 2006) or for household heads and their spouses (as in Burtless, 1999).[4]

Assuming that the distribution of dependent employment, self-employment

[3] The same normalisation was also applied by Fournier (2001), who used the B method on Taiwanese data.

[4] In other words, males and females compete in the same market and if, in a given year, a gender bias exists, a female person will belong to a lower quantile of the same market income distribution than a male, *ceteris paribus*

and pension income are mutually independent, "the density that would have prevailed in 2004, if employees had been paid according to the 1991 wage schedule *and* self-employed and pensioners' income schedule, and if household composition and attributes were left unchanged at 2004" is:

$$f(y; t_{s,p,\theta,\lambda} = 04, t_e = 91) \quad = \quad \int_{\lambda \in \Omega_\lambda} dF(z^e, z^s, z^p, \theta, \lambda | t_{s,p,\theta,\lambda} = 04, t_e = 91). \tag{3}$$

Assuming that, instead of dependent employment income (z^e), only self-employment (z^s) or pension income (z^p) remained at the 1991 level, counterfactuals similar to (3) can be computed.

It is then possible to remove the assumption of mutual independence of different sources of income and, conditional on the *ex ante* choice of income type of each individual in the sample, the "the density that would have prevailed in 2004, if employees, self-employed persons and pensioners had been respectively paid according to the 1991 employment, self-employment and pension income schedule *and* household composition and attributes were left unchanged at 2004" can be written as

$$f(y; t_{\theta,\lambda} = 04, t_{e,s,p} = 91) \quad = \quad \int_{\lambda \in \Omega_\lambda} dF(z^e, z^s, z^p, \theta, \lambda | t_{\theta,\lambda} = 04, t_{e,s,p} = 91). \tag{4}$$

The B and DFL decomposition can then be combined. Assuming that the 2004 income structure does not depend on the distribution of family characteristics, θ, the "density that would have prevailed in 2004, if employees, self-employed persons and pensioners had been respectively paid according to the 1991 employment, self-employment and pension income schedule, the distribution of working-spouse households was as in 1991 *and* other household characteristics and composition were left unchanged at 2004" is

$$\begin{aligned} f(y; t_{\theta,\pi,\mu} &= 04, t_{e,s,p,\phi} = 91) = \int\int f(z^e, z^s, z^p | \phi, \pi, \mu, t_\theta = 04, t_{e,s,p} = 91) \\ &\times \psi_{\phi|\pi,\mu} dF(\phi | \pi, \mu, t_{\phi|\pi,\mu} = 04) dF(\pi, \mu, t_{\pi,\mu} = 04), \end{aligned} \tag{5}$$

where $\psi_{\phi|\pi,\mu} = dF(\phi|\pi, \mu, t_{\phi|\pi,\mu} = 91)/dF(\phi|\pi, \mu, t_{\phi|\pi,\mu} = 04)$ and, assuming that ϕ is a binary variable, it can be obtained by using probit or logit regression models. Finally, the "density that would have prevailed in 2004, if employees, the self-employed and pensioners respectively had been paid according to the 1991 employment, self-employment and pension income schedule, with the distribution of working-spouse households and of family type the same as in 1991 *and* other household characteristics and composition also left unchanged at 2004" is

$$f(y; t_{\theta,\mu} = 04, t_{e,s,p,\phi,\pi} = 91) = \int\int\int f(z^e, z^s, z^p|\phi\pi\mu, t_\theta = 04, t_{e,s,p} = 91) \times \psi_{\phi|\pi,\mu} dF(\phi|\pi, \mu, t_{\phi|\pi,\mu} = 04)\psi_{\pi|\mu} dF(\pi|\mu, t_{\pi|\mu} = 04) dF(\mu|t_\mu = 04), \quad (6)$$

where, $\psi_{\pi|\mu} = I(\pi = j) \times \frac{Pr(\pi=j|\mu,t_{\pi|\mu}=91)}{Pr(\pi=j|\mu,t_{\pi|\mu}=04)}$, $j = 1, 2, 3, 4$ is family type (couple with children, couple without children, single with and single without children, respectively), I is an indicator function taking value 1 if its argument is true and value 0 if false.[5]

[Table 3 about here.]

The combination of the B and DFL methodologies used in this paper is an extension of that used by Daly and Valletta (2006) to analyse inequality and poverty in the US. While Daly and Valletta (2006) only applied it to male household heads, keeping the distribution of other household members unchanged, here the B methodology is applied to all recipients of income from work, regardless of their sex and their role in the family. This extension is motivated by the assumption that market income distribution is independent of the role in the household of income recipients and by the fact that the income of male householders has become less important over time in most Western countries and certainly in Italy (see Section 3 below). Secondly, the B methodology is extended to pension income and the effect of work income is divided into dependent employment and self-employment, as they tend to manifest quite different distributions and trends. The different propensities to work or receive a pension are also considered with the number of months that each individual received income kept constant, with the monthly income vector being replaced but not the annual income. The proposed decomposition is not "exact", but it simply aims to explain as much of the change in inequality as possible.[6]

3 Data and inequality trends

The data set used in this paper is the Bank of Italy Historical Database of the Survey of Italian Household Budgets, 1977-2004 (HD-SHIW, Version 3.0),

[5] For quick reference, the decomposition procedure explained in this section has been summarised, step-by-step, in the upper part of Table 3.

[6] In the inequality decomposition literature, a decomposition method is "exact" if it can express inequality as an additive function of some variables without residual. The decomposition proposed here is not exact as it is expected to deliver a residual, which might come from the parameter of household composition (θ), the change in other household characteristics (μ), the approximation errors deriving from simulation B, and the variability of estimated weights obtained using the DFL procedure

which gathers data collected in all issues of the Survey of Household Income and Wealth (SHIW) from 1977 to 2004. The SHIW is a long-standing cross-sectional survey and is representative of Italian households. The HD-SHIW is the only data set that enables the measurement of changes in Italian household income distribution since the end of the 1970s and relates it to individual, household characteristics and income components. Unfortunately the data set only records incomes net of taxes and social contributions, making the direct effects of tax-benefit policies on inequality over time impossible to measure. Under the sampling design, each household is assigned a weight that is inversely proportional to the probability of its inclusion in the sample. The HD-SHIW weights are finally post-stratified to align the structure of the sample with that of the reference population.[7]

In order to minimise the measurement problem, data on capital, building and real estate income, which are often estimated by the Bank of Italy from declared stocks, are not considered here. Total income is defined as the sum of work (e.g. dependent employment and self-employment) and transfer (mainly pension) income, which is about 20% smaller on average than total income when it also includes capital income.[8] As the problem of the robustness of inequality estimates is particularly serious for some inequality indices that consider the entire distribution, the household income variable was censored at the 99th quantile.[9]

The unit of analysis is the individual, whose level of wellbeing is measured by his household equivalent income. A household is defined as the group of people sharing the same dwelling, independently of their kinship.[10] The equivalent income of individual i belonging to household h is equal to $y_i = \sum_{j \in h} z_j / S$, where $\sum_{j \in h} z_j$ is the sum of all individual incomes accruing to household h ($h = 1, 2, ...H$), and S is the equivalence scale. A special case of the equivalence scale is the parametric class proposed by Buhmann et al.

[7]For further details on the data set, see Banca d'Italia (2006), and for a thorough discussion of the quality of the SHIW data and a review of other Italian data sets, see Brandolini (1999).

[8]The definition of total income excluding capital, building and real estate income is sometimes used also by researchers at the Bank of Italy, see for instance D'Alessio and Signorini (2000).

[9]In other words, all values grater than the 99th quantile were replaced with the 99th quantile without removing them. Some authors prefer to truncate the data set beyond a certain threshold. However if we were to perform truncation instead of censoring, we would lose all information (such as labour force participation, household composition and type of earnings) about top income households, which is relevant for the present analysis. Although seemingly arbitrary, it should be noted that various surveys currently used for income analysis, including the Current Population Survey for the US, are top-coded (for instance, see Gottschalk and Smeeding, 1997). Brandolini and D'Alessio (2001) adopted an analogous top-coding procedure by using HD-SHIW.

[10]This leads to a slight overestimation of the average family size, which cannot be corrected at all since the relevant information is missing from the data set. For simplicity, in what follows household and family are often used as synonyms although the data allow us to observe households only.

(1988):

$$S = (N_h)^\epsilon, \tag{7}$$

where N_h is the household size h and ϵ is the equivalence scale parameter, the elasticity of the scale rate with respect to household size.[11] For most of the results reported in this paper, the equivalence scale parameter is set to $\epsilon = 0.5$, which is often referred to as the square root equivalence scale. No sample selection was performed and all households and individuals in the original data set are included in the analysis. SHIW allows interviewees to freely choose the household head, regardless of the role in the household or income received. In the case of a female household head with a cohabiting partner the male partner is recoded as the head and the female as the partner/wife just for the sake of uniformity. All figures are weighted using HD-SHIW sampling weights.

Figure 1 shows the trends in a number of inequality indices, including the Gini and three indices of the Generalised Entropy (GE) class in the top panel, and four quantile ratios in the bottom panel, each normalised to make the 1977 value equal to 100. The three $GE(a)$ indices used have parameter $a = 0, 1, 2$, and are also known respectively as the mean logarithmic deviation or $GE(0)$, the Theil index or $GE(1)$, and half of the square of the coefficient of variation or $GE(2)$. The Gini and GE indices show that equivalent household income inequality decreased from the end of the 1970s through to 1991, although with some fluctuations. After 1991 all inequality measures jumped to levels very close to those at the beginning of the period under consideration and remained rather stable afterwards. All three GE indices moved in very similar way until 1989. Between 1991 and 2002, the mean log deviation was higher than the other GE indices. As GE indices differ in their sensitivity to differences at different points of the distribution (in particular GE indices, with a larger and more positive coefficient a, are more sensitive to income differences at the right tail of the distribution and, vice versa (Cowell, 1995)), the top panel of Figure 1 suggests that we look at the left tail of the distribution to understand the increase in household income inequality since 1991.

[Figure 1 about here.]

All quantile ratios in the bottom panel show a decreasing trend until the end of the 1980s and the beginning of the 1990s. They increased significantly between 1991 and 1993 and remained relatively stable afterwards (or decreased slightly in the case of the 95/5 quantile).[12] During the whole period from 2004-1991, the 95/5 quantile ratio increased by over 40%.

[11] The equivalence scale (7) assumes that all household incomes are gathered and equally distributed across the household and considers the size as an approximation of household needs. Buhmann et al. (1988) show that this scale provides a good approximation to most scales used in empirical studies of income distribution in developed countries, including some scales based on other characteristics in addition to size, such as the age of household members.

[12] Some readers might be sceptical about such a dramatic increase of equivalent income inequality between 1991 and 1993, arguing that aside from the crisis of September 1992,

Figure 2 presents the nonparametric density estimation of equivalent household income in the years 1977, 1991, 2004.[13] Between 1977 and 1991 some households with equivalent income below €5000 improved their relative position and after 1991 the trend partly reversed.

[Figure 2 about here.]

A number of factors certainly played a role in such a distinctive trend in Italian equivalent income inequality. The consumer price annual inflation rate has gradually decreased since peaking at 20.8% in 1980 and has never been higher than 5% since 1992. As low income households are typically more vulnerable to inflation due to their lower ability to hedge against it (Mulligan and Sala-i-Martin, 2000), their limited bargaining power in the political process (Albanesi, 2007) and the frequently regressive effects of bracket creeping (Immervoll, 2005), the decreasing rate of price inflation may have contributed to the downward trend of inequality up to the end of the 1980s, becoming a much less relevant factor afterwards. Following the September 1992 currency and financial crisis, Italy went through its deepest economic crisis since WWII, which led to a series of tough policies, aimed at controlling public finances and limiting the expansion of public debt. They included two thorough reforms of the pension system, a reform of collective bargaining, the tightening of the rules for regional health care finance and management, increase of taxes on building and estate properties, labour and on small business income. These policies reduced the progressivity of taxation and to some degree failed to cope with the expansion of tax evasion, especially among the self-employed, or the slowing rate of growth of the national economy (Ciocca, 2004; Castronovo, 2006).

This paper looks primarily for the microeconomic causes of the distinctive trend towards equivalent income inequality in Italy, focussing in particular on socio-demographic and household income composition effects rather than on economy-wide factors, thereby fitting the paper into the existing literature on inequality in household income. The Italian population has experienced several changes since the end of the 1970s (Table 1). For instance, there has been a great change in household composition: in 1977 over 80% of households were couples (with or without kids) and less than 10% were single-person households; in 2004 the former type of household accounted for less than 63% and the latter for over 25%. There was a clear trend in household size distribution, with larger households (i.e. with 3 or more members) down from nearly two

possible problems in the sampling procedure may have also played a role. Unfortunately there are no alternative data available to properly corroborate the SHIW data. However, even looking at the 1,837 households (about 20% of the total) that were interviewed in both 1991 and 1993, we find a similar jump in income distribution (I would like to thank Andrea Brandolini for suggesting to perform this analysis).

[13] Income was adjusted to take account of inflation using the consumer price index.

thirds to less than half of the total.[14] The participation in the labour market of Italian wives, which has traditionally been low compared to that of other Western countries, showed an upward trend. In 1977, six wives out of ten had zero income, by 2004 the proportion was nearly reversed.

[Table 1 about here.]

Figure 3 presents inequality measures of individual income according to main source, namely dependent employment, self-employment, and pension income and is to be read in columns similarly to Figure 1.[15] As long as they receive an income, all individuals in the sample are considered, regardless of their role in the household. Although individual inequality trends by income type also show at first a decreasing trend and then a significant increase after 1991, they each present their own distinguishing features. For instance, employment income inequality, as measured by the Gini or GE indices, hit its the minimum in 1989, while the ratio of rich and median households remains relatively stable, as measured by the 90/50 quantile ratio. Wage inequality dynamics appear to be driven by a relative impoverishment of poor households,[16] as measured by the 50/10. Similarly, self-employment income inequality shows a large increase of the 50/10, with a relatively stable 90/50 quantile ratio. The increased dispersion of wages has been convincingly explained as an effect of the complete abolition in 1992 of the automatic indexation of wages to the consumer price index (the so-called "*scala mobile*"), a gradual process that began in the mid 1980s (Erickson and Ichino, 1995; Manacorda, 2004; Devicienti, 2003). However this institutional change did not have a direct effect on the earnings of the self-employed and its indirect effects (e.g. due to cross-migration between occupations) are unlikely to have been large.

[Figure 3 about here.] [Table 2 about here.]

The trend in inequality for pension income is however rather different: it started to rise after 1987 and stopped only in mid 1990s. In particular, median pension income rose relative to both the 10th quantile (40% in two decades) and the 90th quantile (the 90/50 quantile ratio reduced by over 30%). Here it should be noticed that, starting from the mid-1980s, pensioners increasingly consist in people who worked for long and stable periods in large firms and who have a full history of national insurance contributions. In contrast, those who retired earlier had on average a short history of contributions due to the war and had worked irregularly for small firms, which frequently evaded payment of social contributions. Moreover, the relative increase in the median pension

[14] Decreasing household size is the most likely reason why fewer households have three or more income earners, even though more wives participate in the labour market

[15] For a table with the whole set of inequality figures and its standard errors, see Table 2.

[16] Here households are called poor if their income is below the 10th quantile, with no reference to a well-defined poverty line. Analogously, rich households have an income at or above the 90th quantile.

may be due to composition effects as the share of pensioners who worked in the public sector and enjoyed relatively better pensions than other sectors (at least until the two mid 1990s pension reforms), nearly doubled between 1981 and 2000, while the share of those who worked in the private sector remained largely constant over the same period.

4 Results

The methodology outlined in Section 2 has been applied focussing on a few relevant years, in particular the very first and last years of the period considered and the year in which equivalent household income inequality reached its minimum, i.e. 1977, 2004 and 1991. The rate of growth of Italy's GDP in all the three years was similarly close to 2% and all three years were at least one year away from the economic cycle's peaks and lows. Results are reported, along with their bootstrap standard errors, in two tables, which compare equivalent household incomes in 1991 with those of 1977, and those of 2004 with 1991, using different assumptions about counterfactual incomes.

The bottom part of Table 3 summarises the primary order decompositions performed, with reference to the equations introduced in Section 2.[17] The variable ϕ defines participation of wives in the labour force and takes a value equal to 1 for households where the male head is part of a couple and his partner is younger than 65 and participating in the labour market and equal to 0 if she is not in the labour market; π defines the family type and takes one of four discrete values (couple with kids, couple with no kids and single with kids and single without kids); μ is the vector of other family characteristics.[18] The models for DFL weight estimation were tested for robustness and more parsimonious models were preferred given that the predictive ability of a model, which is at the core of the DFL method, is well known to decrease with the model's complexity.

[Table 3 about here.]

The analysis here is presented only using some common dispersion indices and no non-parametric density estimate given that there was no notable change in the profile of the income distribution for the the nature of changes that occurred in Italy over the period considered. This was different, for instance, from the US, as documented by DFL. The set of inequality measures used is the same as those used in figures 1 and 3.

[17]In its empirical application, 500-quantile distributions were used in the B methodology, although results do not change significantly even when using a larger number of quantiles.

[18]This includes geographical area of residence, town/city of residence size, sex, main occupation, sector of occupation and income of householder, age and years of education of the householder and their squares, average age of the household and its square, number of kids in the household and its square, and various interactions.

Table 4 presents the first set of results: the primary order decomposition of household equivalent income inequality between 1977 and 1991. Column E of Table 4 shows that the changed distribution of income sources accounts for a very large proportion of the reduction in equivalent household income inequality. In fact, keeping the 1977 joint distribution of individual dependent employment, self-employment and pension income as it was in 1991, conditional on the choice of income type, would have roughly cancelled the difference in the 90/10 and 95/5 quantile ratios. The changed dispersion of individual incomes accounts for roughly 80% of the total change of the Gini and GE indices and largely over-explains the reduction of the 50/10 ratio. Assuming the mutual independence of different sources of income, and looking at columns B, C and D, we can see that while the distribution of pension income accounts for less than 10% of the reduced inequality (with the exception of the 50/10 ratio, of which it accounts for 22%), the distribution of work income and in particular of self-employment income accounts for the largest proportion of the reduced inequality. Conversely, as columns F and G show, a much smaller part of inequality dynamics is attributable to the demographic changes in households and is actually negligible in the case of household types. The increased female participation had an unequalising effect in the left tail and an equalising effect for households in the right tail of the income distribution, as emerges from the indices that focus on the left tail of the income distribution (i.e. the mean log deviation and the 50/10 quantile ratio). Overall, this decomposition nearly explains the whole change in the 90/50 quantile ratio, and a rather large share (over 80%) of the total change for most of the other indices.

[Table 4 about here.]

Table 5 presents the primary order decomposition of household equivalent income inequality changes between 2004 and 1991. It shows that the changing dispersion of income sources accounts for a large proportion of the increased inequality, although less than for the 1991-1977 period. In particular, while pension income, when significant, had an equalising effect, the changing work income distribution had an unequalising effect. The changing distribution of employment income had a more pronounced disequalising effect on the right tail, which is consistent with the progressive loss of trade union bargaining power and with past results showing that the gradual abolition of the automatic wage indexation ("*scala mobile*") between the late 1980s and 1991 unleashed the decompression of wages seen elsewhere during the 1980s and already mentioned above with reference to Figure 3. The changing distribution of self-employment income had by contrast a larger effect on the left tail of the equivalent household income distribution, possibly due to the recession of the early 1990s and the introduction of a number of new taxes on small business income after 1992. In contrast to the previous period considered, changing female participation in the labour force had an unequalising effect regardless of the inequality index used (accounting for an additional 6-15% of total change).[19]

[Table 5 about here.]

The result of the inequality enhancing effects of participation in the labour force by wives is in contrast with what was found, for instance, by Daly and Valletta (2006) for the US and by Biewen (2001) for East Germany, but not with what was found by others (see for instance Shorrocks (1983); Karoly and Burtless (1995) or Cancian and Reed (1998)). As for Italy, a simple logit regression of the probability of participation in the labour force by wives conditional on their own personal characteristics and their partner's income, age and years of education shows that, while in 1977 women coupled with well-educated partners were, *ceteris paribus*, less likely to work, the reverse was true in 1991 and 2004. As better educated people have a higher average market income, the increase in participation of wives in the labour market was mostly concentrated in relatively well-off families, increasing differences between households with above median income and households with below median income. Further insights can be gained by looking at the unconditional probability of female labour force participation at different levels of income over the whole period. Table 6 shows that the participation of wives did not increase uniformly across all household income vigintiles, but that it was positively correlated with household income, at least in the 1991-2004 period. Something different occurred in the period 1977-1991. Female labour participation increased among households with around the median income. This had the effect of augmenting the gap in relative incomes with poorer and less active households, thus exerting a strong inequality enhancing effect at the lower half of the income distribution (cf. the GE(0) index and the 50/10 percentile ratio in Table 4). Conversely, the increased participation of median income households reduced the gap in relative incomes with top-income households, thereby reducing inequality (cf. the GE(2) index and the 90/50 percentile ratio). These two effects combined produced inconclusive results in the 1977-1991 period as they caused a clear increase in inequality using the 90/10 and 95/5 percentile ratios and a decrease in inequality using the Gini and Theil indices.

[Table 6 about here.]

Overall, the 2004-1991 decomposition of equivalent household income leaves a sizable proportion of the total change unexplained, from a minimum of 33% of half the squared coefficient of variation to a maximum of 62% of the 50/10

[19] Results have also been checked for robustness by reversing the decomposition sequence and by changing the equivalence scale as defined in (7), choosing $\epsilon = 0.25, 0.75$, yielding results that are similar to those reported here. This robustness analysis can be found in (Fiorio, 2008) and was omitted here for reasons of space.

quantile ratio. A large portion of unexplained inequality is likely to derive from changes in personal and small business income taxes, an area that saw a whole series of reforms during the 1990s, and cannot be accounted for with available data, or by differences in assortative mating, which is not studied here. However, the amount of inequality that has been explained is much grater than what could and has been explained using other decomposition methods.[20] Moreover, a large residual is not unusual in this type of analysis.[21]

5 Conclusions

This paper applies and extends the methodology of Daly and Valletta (2006), providing a unifying framework with which to analyse the relative importance of socio-demographic changes and changing dispersion of sources of income in explaining equivalent income inequality trends in Italy since the end of the 1970s, overcoming some of the limitations of traditional methodologies.

The main empirical findings are as follows: (a) In the period under consideration individual income dispersion changes played a greater role in accounting for household inequality trends than socio-demographic changes. (b) The changing distributions between dependent employment and self-employment had a similar effect in size and direction, explaining the reduction of inequality before 1991 and its increase afterwards. However, as total self-employment income in absolute terms accounted for less than one third of total work income, its relative importance for decomposing inequality was greater. (c) Pension income distribution changes had an equalising effect throughout the period. (d) The increasing participation of wives in the labour force had particular effects: it was inequality-increasing across the whole period for equivalent incomes below the median, while it was inequality-reducing before 1991 and inequality-increasing after 1991 for incomes above the median.

Understanding the differential effect of different types of income and female participation in the labour force on equivalent household income distribution

[20] Brandolini and D'Alessio (2001) decomposed the mean logarithmic deviation index by population subgroups showing that the demographic characteristic with the largest impact was the classification by sex of the head of household, which however accounted for only 3.3% of inequality change. D'Alessio and Signorini (2000), by using a subgroup decomposition of the Gini index, showed that the increased participation of the female labour force and the higher number of pensioners had a role in the change in household inequality, although they could not explain more than 10% of the total change.

[21] For instance, a simulation decomposition of inequality changes in the US (with a methodology analogous to the one adopted here) can leave unexplained more than 40% of the total change in some inequality indices (Daly and Valletta, 2006, Table 2). Single-equation regression-based decomposition of inequality indices applied to Chinese data found a residual between 40% and 90% of the total change depending on the inequality index used (Morduch and Sicular, 2002, Table 2), while structural equation equivalent income inequality decomposition of Taiwan equivalent income inequality leaves about 25% of mean change unexplained (Bourguignon et al., 2001, Table 4).

has important policy implications. The increase in inequality after 1991 possibly suggests more redistribution, considering that among the EU countries during the 1990s Italy was the one with the lowest redistribution from top to bottom (G, 2004) and the lowest level of unemployment benefit coverage (Boeri and Brandolini, 2004). However, the differing effects of changes in the distribution of dependent employment, self-employment and pension income suggest the need to avoid a uniform approach for all types of income. As the number of wives in the labour force increased more for high-income households, contributing to increased inequality especially after 1991, Italian policy-makers should consider targeted policies for increasing female participation at lower income levels.

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Figures and Tables

Table 1: Some socio-demographic statistics.

	Year			
Main occupation of income receivers	**1977**	**1991**	**1993**	**2004**
employees	47.96	45.57	43.76	44.68
self-employed	20.33	17.89	17.03	16.61
pensioners	31.72	36.54	39.21	38.71
Total	*100.00*	*100.00*	*100.00*	*100.00*
Household type	**1977**	**1991**	**1993**	**2004**
couple with kids	58.47	52.68	51.22	40.73
couple no kids	21.99	19.22	19.25	22.08
single with kids	5.64	7.17	8.15	7.37
single	9.88	18.21	17.53	25.65
other	4.02	2.72	3.85	4.17
Total	*100.00*	*100.00*	*100.00*	*100.00*
Household size	**1977**	**1991**	**1993**	**2004**
1 member	9.88	18.21	17.53	25.65
2 members	24.67	23.70	24.64	28.28
3+ members	65.45	58.08	57.83	46.06
Total	*100.00*	*100.00*	*100.00*	100.00
Number of income receivers	**1977**	**1991**	**1993**	**2004**
1	50.05	45.47	44.99	51.13
2	36.22	39.74	41.78	38.43
3+	13.74	14.79	13.23	10.44
Total	*100.00*	*100.00*	*100.00*	*100.00*
Working status female spouse[a] **(if present)**	**1977**	**1991**	**1993**	**2004**
female spouse not in labour force	62.49	45.29	43.67	43.14
female spouse in labour force	37.51	54.71	56.33	56.86
Total	*100.00*	*100.00*	*100.00*	*100.00*

Source: author's calculations on SHIW-HD. Weighted estimates.
Note: Only statistics for some relevant years are provided but the whole series can be obtained from the author upon request. (a) A female spouse is the female partner of a male household head.

Figure 1: Household equivalent income inequality, with square-root equivalence scale ($\epsilon = .5$).

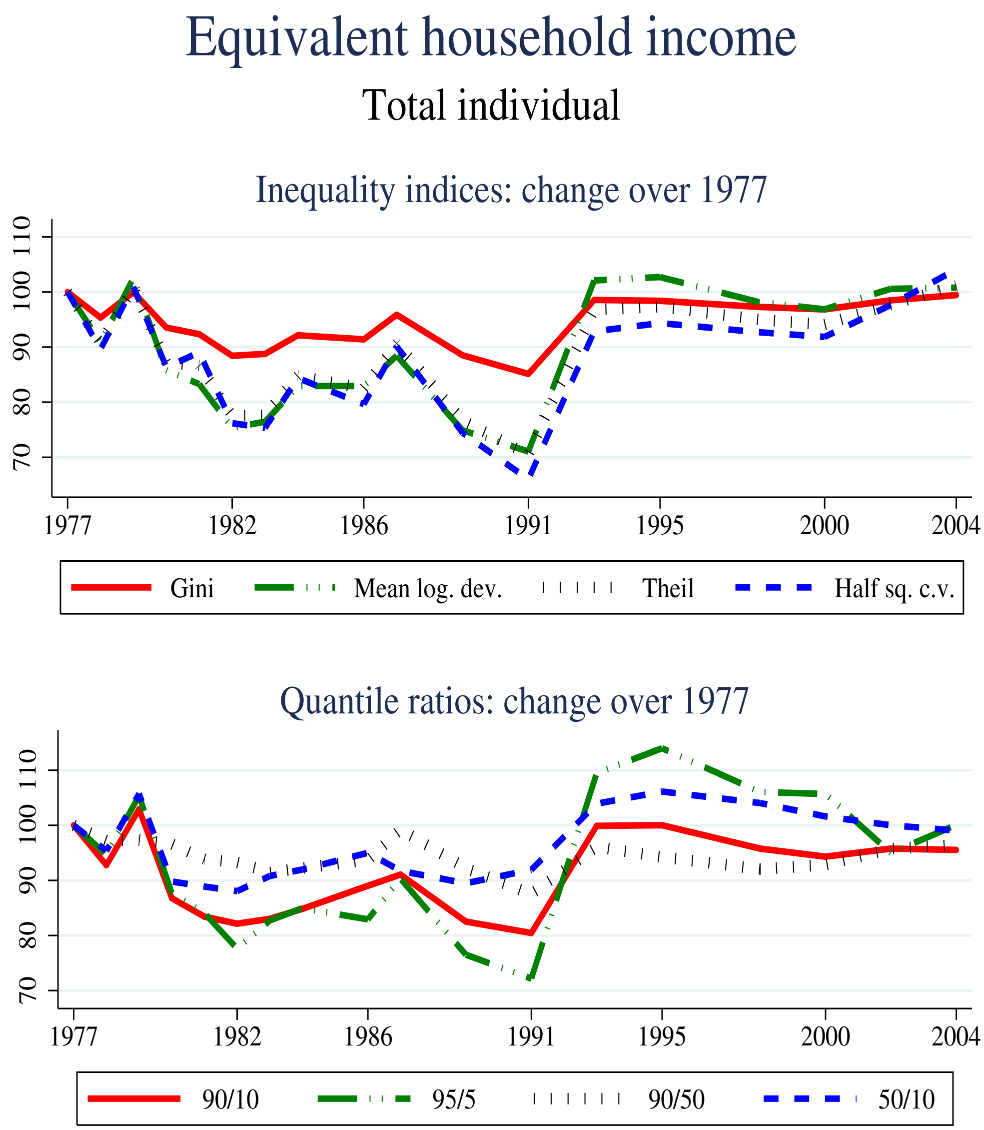
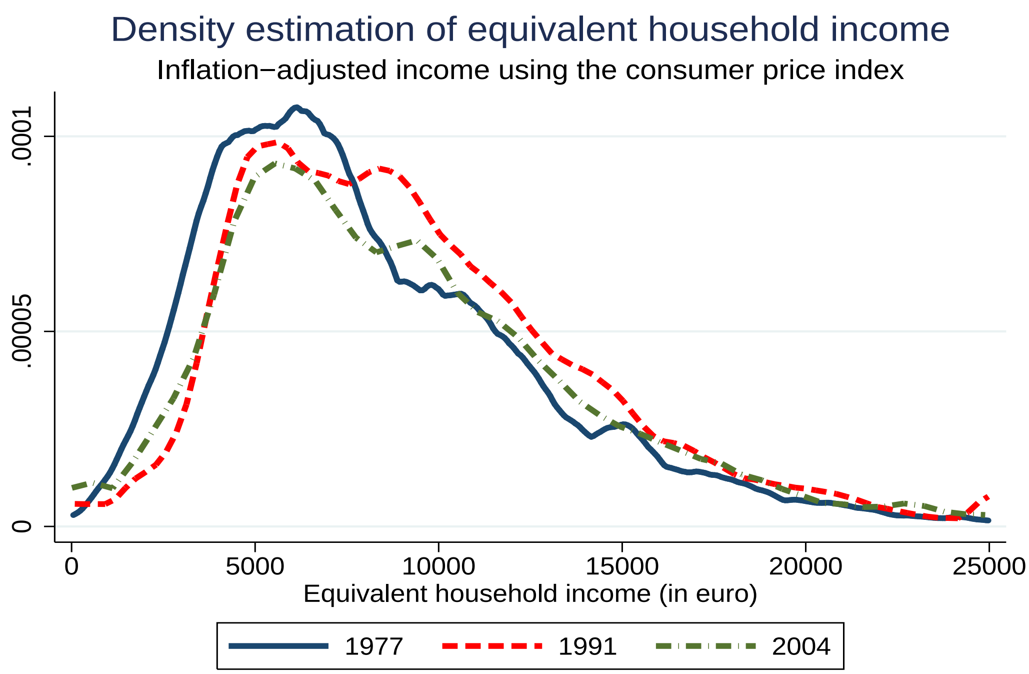


Figure 2: Kernel density estimation of CPI-adjusted equivalent household income (square-root equivalence scale) for three selected years.



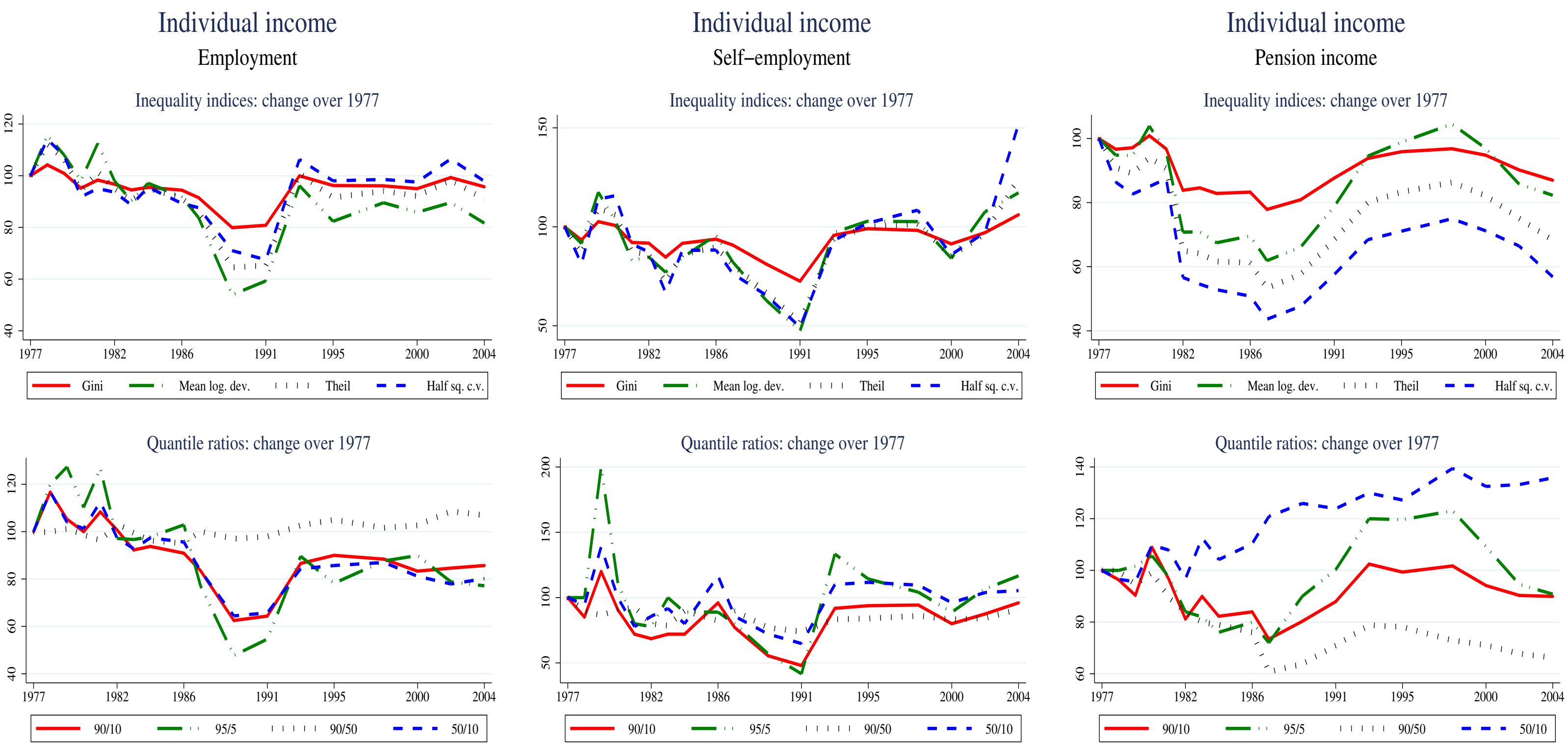


Figure 3: Individual income inequality by type of income.

Table 2: Inequality measures for some relevant years.

	Equivalent household income			
	1977	**1991**	**1993**	**2004**
Gini	0.318	0.271	0.314	0.316
s.e.	(0.002)	(0.002)	(0.002)	(0.002)
GE(0)[a]	0.176	0.125	0.180	0.178
s.e.	(0.002)	(0.002)	(0.002)	(0.003)
GE(1)[a]	0.166	0.117	0.161	0.168
s.e.	(0.002)	(0.001)	(0.002)	(0.003)
GE(2)[a]	0.186	0.123	0.172	0.193
s.e.	(0.003)	(0.002)	(0.003)	(0.005)
p90p10[b]	4.486	3.609	4.482	4.286
s.e.	(0.053)	(0.045)	(0.076)	(0.048)
p95p5[b]	7.085	5.114	7.758	7.091
s.e.	(0.090)	(0.078)	(0.184)	(0.111)
p90p50[b]	2.072	1.813	1.992	1.999
s.e.	(0.014)	(0.017)	(0.018)	(0.016)
p50p10[b]	2.165	1.991	2.250	2.145
s.e.	(0.025)	(0.019)	(0.033)	(0.019)

	Individual employment income				Individual self-employment income				Individual pension income			
	1977	**1991**	**1993**	**2004**	**1977**	**1991**	**1993**	**2004**	**1977**	**1991**	**1993**	**2004**
Gini	0.259	0.209	0.259	0.248	0.417	0.303	0.399	0.443	0.325	0.285	0.305	0.283
s.e.	(0.002)	(0.002)	(0.002)	(0.002)	(0.005)	(0.004)	(0.005)	(0.008)	(0.004)	(0.002)	(0.002)	(0.002)
GE(0)[a]	0.150	0.089	0.144	0.123	0.319	0.152	0.310	0.374	0.170	0.134	0.161	0.140
s.e.	(0.003)	(0.002)	(0.003)	(0.002)	(0.007)	(0.004)	(0.008)	(0.014)	(0.004)	(0.002)	(0.002)	(0.003)
GE(1)[a]	0.118	0.077	0.119	0.107	0.293	0.152	0.274	0.365	0.191	0.131	0.153	0.130
s.e.	(0.002)	(0.001)	(0.002)	(0.002)	(0.008)	(0.005)	(0.007)	(0.020)	(0.005)	(0.002)	(0.002)	(0.003)
GE(2)[a]	0.115	0.078	0.122	0.113	0.359	0.177	0.334	0.546	0.248	0.143	0.170	0.141
s.e.	(0.003)	(0.002)	(0.003)	(0.002)	(0.016)	(0.007)	(0.012)	(0.056)	(0.009)	(0.003)	(0.003)	(0.004)
p90p10[b]	4.000	2.575	3.464	3.429	8.333	4.000	7.648	8.000	3.891	3.421	3.986	3.500
s.e.	(0.019)	(0.048)	(0.131)	(0.047)	(0.172)	(0.006)	(0.516)	(0.105)	(0.093)	(0.075)	(0.143)	(0.020)
p95p5[b]	7.778	4.250	6.970	6.000	15.000	6.250	20.000	17.500	5.556	5.571	6.667	5.045
s.e.	(0.315)	(0.117)	(0.403)	(0.218)	(0.403)	(0.220)	(1.061)	(1.596)	(0.168)	(0.110)	(0.039)	(0.130)
p90p50[b]	1.558	1.526	1.600	1.667	2.778	2.057	2.317	2.532	2.862	2.031	2.256	1.896
s.e.	(0.015)	(0.016)	(0.019)	(0.021)	(0.030)	(0.040)	(0.058)	(0.070)	(0.060)	(0.022)	(0.027)	(0.019)
p50p10[b]	2.567	1.687	2.165	2.057	3.000	1.944	3.300	3.160	1.359	1.684	1.766	1.846
s.e.	(0.026)	(0.027)	(0.078)	(0.014)	(0.060)	(0.038)	(0.220)	(0.087)	(0.021)	(0.036)	(0.059)	(0.018)

Source: author's calculations on SHIW historic archive. Weighted samples.

Note: In parenthesis the bootstrap standard errors with 399 bootstrap replications. (a) The generalised entropy index GE(0) is also known as the mean logarithmic deviation, the $GE(1)$ as t Theil index and $GE(2)$ as the half the square of the coefficient of variation. (b) p90p10 stands for the ratio between the 90th and the 10th quantiles, the following percentile ratios measur are defined similarly.

Table 3: Primary order decomposition.

line number	density	simulation method	equation reference[a]
1	$f(y; t_{e,s,p,\theta} = i, t_\lambda = i)$		(1)
2	$f(y; t_{e,s,p,\theta} = j, t_\lambda = j)$		(1)
3	$f(y; t_{s,p,\theta,\lambda} = i, t_e = j)$	B	(3)
4	$f(y; t_{e,p,\theta,\lambda} = i, t_s = j)$	B	(3)
5	$f(y; t_{e,s,\theta,\lambda} = i, t_p = j)$	B	(3)
6	$f(y; t_{\theta,\lambda} = i, t_{e,s,p} = j)$	B	(4)
7	$f(y; t_{\theta,\pi,\mu} = i, t_{e,s,p,\phi} = j)$	B&DFL	(5)
8	$f(y; t_{\theta,\mu} = i, t_{e,s,p,\phi,\pi} = j)$	B&DFL	(6)
	labels in tables of results	total change attributable to each factor[b]	proportion of total change attributable to each factor[b,f]
A	**Total actual change**	$(1-2)$	
B	**empl.**[c]	$(1-2)-(1-3)$	$(3-2)/(1-2)$
C	**self-empl.**[c]	$(1-2)-(1-4)$	$(4-2)/(1-2)$
D	**pension**[c]	$(1-2)-(1-5)$	$(5-2)/(1-2)$
E	**income**[d]	$(1-2)-(1-6)$	$(6-2)/(1-2)$
F	**female FP**[e]	$(1-2)-(1-7)$	$(7-2)/(1-2)$
G	**family type**[e]	$(1-2)-(1-8)$	$(8-2)/(1-2)$
H	**residual factor**	$(1-8)$	$(1-8)/(1-2)$

Notes: $i, j, i \neq j$ can be any couple of years.
(a) This column provides the relevant equations' reference numbers (see Section 2).
(b) Numbers reported in this column are correspond to line numbers defined in the top panel of this table.
(c) total change that can be explained by changes in employment (self-employment/pension) marginal distribution.
(d) total change that can be explained by changes in joint distribution of employment, self-employment and pension incomes.
(e) total change that can be explained by changes in joint distribution of employment, self-employment and pension incomes, and some family characteristics.
(f) This column shows how the total change and the proportion of the total change attributable to each factor were computed.

Table 4: Decomposition of changes in the distribution of household equivalent income, 1991-1977.

Total actual change (A)	empl.[b] (B)	self-empl.[b] (C)	pension[b] (D)	Effect of:[a] income joint distrib.[c] (E)	+ female LFP[d] (F)	+ family type[e] (G)	residual factor[f] (H)
				Gini			
-0.047	-0.011	-0.023	-0.004	-0.038	-0.039	-0.040	-0.008
(0.003)	(0.001)	(0.002)	(0.001)	(0.002)	(0.004)	(0.004)	(0.004)
	[23.1%]	[48.8%]	[7.6%]	[79.6%]	[83.2%]	[84.0%]	[16.0%]
				Mean Log Deviation (GE0)			
-0.051	-0.013	-0.027	-0.004	-0.043	-0.040	-0.040	-0.011
(0.003)	(0.001)	(0.002)	(0.001)	(0.002)	(0.004)	(0.004)	(0.004)
	[26.1%]	[52.6%]	[7.8%]	[84.8%]	[77.7%]	[79.3%]	[20.7%]
				Theil (GE(1))			
-0.049	-0.011	-0.025	-0.004	-0.040	-0.041	-0.041	-0.008
(0.003)	(0.001)	(0.002)	(0.001)	(0.002)	(0.004)	(0.004)	(0.003)
	[23.3%]	[51.2%]	[7.8%]	[80.8%]	[83.1%]	[83.8%]	[16.2%]
				Half squared coefficient of variation (GE(2))			
-0.063	-0.013	-0.033	-0.005	-0.050	-0.055	-0.055	-0.009
(0.004)	(0.002)	(0.003)	(0.001)	(0.004)	(0.005)	(0.005)	(0.004)
	[21.3%]	[51.5%]	[7.6%]	[78.6%]	[86.4%]	[86.5%]	[13.5%]
				Percentile ratio: 90/10			
-0.822	-0.304	-0.402	-0.053	-0.855	-0.743	-0.752	-0.070
(0.063)	(0.049)	(0.052)	(0.039)	(0.053)	(0.092)	(0.089)	(0.089)
	[37.0%]	[48.9%]	[6.4%]	[104.1%]	[90.5%]	[91.5%]	[8.5%]
				Percentile ratio: 95/5			
-1.767	-0.616	-0.673	-0.159	-1.742	-1.200	-1.354	-0.412
(0.132)	(0.091)	(0.119)	(0.064)	(0.123)	(0.306)	(0.260)	(0.264)
	[34.9%]	[38.1%]	[9.0%]	[98.6%]	[67.9%]	[76.7%]	[23.3%]
				Percentile ratio: 90/50			
-0.260	-0.035	-0.053	0.005	-0.154	-0.261	-0.258	-0.001
(0.020)	(0.017)	(0.015)	(0.008)	(0.017)	(0.042)	(0.044)	(0.044)
	[13.4%]	[20.6%]	[-1.9%]	[59.3%]	[100.5%]	[99.5%]	[0.5%]
				Percentile ratio: 50/10			
-0.142	-0.112	-0.141	-0.031	-0.271	-0.098	-0.105	-0.037
(0.028)	(0.018)	(0.021)	(0.017)	(0.022)	(0.036)	(0.035)	(0.038
	[78.7%]	[99.4%]	[21.5%]	[190.6%]	[68.7%]	[73.9%]	[26.1%]

Source: author's calculations using SHIW archive data.

Notes: The table reads left to right. For instance, the change of the Gini index between 1977 and 1991 was -0.047. The changed distribution of employment income caused the Gini to change by -0.011 (the 23.1% of total change), the changed distribution of self-employment income caused the Gini to change by -0.023 (48.8% of total change), etc. Any specific factor can over- or under-explain the total change.
(a) Numbers in parenthesis show bootstrap standard errors. Numbers in brackets show the percentage share of the explained change in the total change.
(b) holding only employment (or self-employment or pension) marginal distribution constant.
(c) holding marginal distributions of employment, self-employment and pension jointly constant.
(d) as in footnote (c) and additionally applying the DFL re-weighting procedure for female labour force participation.
(e) as in footnote (d) and additionally applying the DFL re-weighting procedure for family type.
(f) the residual factor is what is left unexplained after changing income joint distribution, female labour force participation and family types have been accounted for (column G).

Table 5: Decomposition of changes in the distribution of household equivalent income, 2004-1991.

	Effect of:[a]						
Total actual change (A)	**empl.[b] (B)**	**self-empl.[b] (C)**	**pension[b] (D)**	**income joint distrib.[c] (E)**	**+ female LFP[d] (F)**	**+ family type[e] (G)**	**residual factor[e] (H)**
				Gini			
0.046	0.010	0.012	-0.003	0.019	0.026	0.027	0.019
(0.003)	(0.001)	(0.002)	(0.001)	(0.002)	(0.003)	(0.003)	(0.003)
	[22.8%]	[26.5%]	[-6.1%]	[42.7%]	[56.1%]	[58.8%]	[41.2%]
				Mean Log Deviation (GE0)			
0.052	0.012	0.014	-0.003	0.022	0.028	0.029	0.023
(0.003)	(0.001)	(0.002)	(0.001)	(0.003)	(0.003)	(0.003)	(0.003)
	[22.6%]	[26.3%]	[-5.6%]	[42.5%]	[53.9%]	[56.1%]	[43.9%]
				Theil (GE(1))			
0.051	0.011	0.017	-0.003	0.025	0.030	0.032	0.019
(0.003)	(0.001)	(0.002)	(0.001)	(0.003)	(0.003)	(0.003)	(0.003)
	[22.2%]	[32.6%]	[-5.5%]	[48.5%]	[59.9%]	[62.3%]	[37.7%]
				Half squared coefficient of variation (GE(2))			
0.071	0.015	0.028	-0.004	0.038	0.045	0.047	0.024
(0.005)	(0.002)	(0.004)	(0.001)	(0.005)	(0.005)	(0.005)	(0.004)
	[21.3%]	[38.9%]	[-5.0%]	[54.2%]	[64.2%]	[66.4%]	[33.6%]
				Percentile ratio: 90/10			
0.647	0.163	0.111	-0.100	0.158	0.255	0.308	0.339
(0.075)	(0.048)	(0.071)	(0.042)	(0.071)	(0.079)	(0.079)	(0.085)
	[25.3%]	[17.1%]	[-15.5%]	[24.4%]	[39.4%]	[47.6%]	[52.4%]
				Percentile ratio: 95/5			
2.627	0.973	0.332	0.039	0.761	1.151	1.151	1.475
(0.181)	(0.136)	(0.219)	(0.162)	(0.166)	(0.192)	(0.2)	(0.214)
	[37.0%]	[12.6%]	[1.5%]	[29.0%]	[43.8%]	[43.8%]	[56.2%]
				Percentile ratio: 90/50			
0.183	0.075	0.012	-0.017	0.057	0.087	0.096	0.086
(0.024)	(0.013)	(0.014)	(0.013)	(0.018)	(0.024)	(0.023)	(0.026)
	[41.3%]	[6.7%]	[-9.6%]	[31.0%]	[47.6%]	[52.8%]	[47.2%]
				Percentile ratio: 50/10			
0.139	0.000	0.042	-0.031	0.018	0.035	0.052	0.086
(0.031)	(0.023)	(0.033)	(0.021)	(0.03)	(0.03)	(0.03)	(0.036)
	[0.0%]	[30.6%]	[-22.3%]	[13.0%]	[25.2%]	[37.8%]	[62.2%]

Source: author's calculations using SHIW archive data.

Notes: The table reads left to right. For instance, the change of the Gini index between 1991 and 2004 was 0.046. The changed distribution of employment income caused the Gini to change by 0.010 (the 22.8% of total change), the changed distribution of self-employment income caused the Gini to change by 0.012 (26.5% of total change), etc. Any specific factor can over- or under-explain the total change.

Notes: (a) Numbers in parenthesis show bootstrap standard errors. Numbers in brackets show the percentage share of the explained change in the total change.

(b) holding only employment (or self-employment or pension) marginal distribution constant.

(c) holding marginal distributions of employment, self-employment and pension jointly constant.

(d) as in footnote (c) and additionally applying the DFL re-weighting procedure for female labour force participation.

(e) as in footnote (d) and additionally applying the DFL re-weighting procedure for family type.

(f) the residual factor is what is left unexplained after changing income joint distribution, female labour force participation and family types have been accounted for (column G).

Table 6: The average proportion of wives out of the labour force by some relevant equivalent household income vigintiles.

		Wives out of the labour force				
					Change	
vigintile	vigintile interval	**1977**	**1991**	**2004**	**1991-1977**	**2004-1991**
1	0-5	86.53	90.59	88.27	4.7%	-2.6%
2	5-10	64.11	82.44	85.43	28.6%	3.6%
3	10-15	63.41	87.27	79.16	37.6%	-9.3%
9	40-45	64.35	63.78	53.07	-0.9%	-16.8%
10	45-50	81.25	58.04	42.09	-28.6%	-27.5%
11	50-55	80.95	52.63	40.45	-35.0%	-23.1%
18	85-90	38.16	24.98	22.90	-34.5%	-8.3%
19	90-95	31.32	20.61	13.97	-34.2%	-32.2%
20	95-100	48.54	21.40	19.19	-55.9%	-10.3%

Source: author's calculations on SHIW historic archive.

Notes: The j-th income vigintile interval encompasses household equivalent incomes between the j-th and the (j-1)-th vigintile. The proportion of female spouses in the labour force is the complement to 100. A vigintile interval encompasses (income) values between two consecutive vigintiles, which are quantiles dividing the set of data into 20 equal proportions.