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RETIREMENT, PENSION ELIGIBILITY AND HOME PRODUCTION

by Emanuele Ciani*

Abstract

I estimate the effect of retirement on housework by exploiting the discontinuity in pension eligibility generated by the Italian social security rules. Using microdata from the 2007 wave of the Survey on Income and Living Conditions (SILC), I show that women increase their time spent on home production by more than 400 minutes per week. For men, there is on average no evidence of a significant change, which differs from the results of studies in other countries. However, estimates are heterogeneous by marital status, suggesting that married men do not increase time spent on household production because they can rely on their spouses. I also discuss other possible explanations, in particular men dedicating their time to 'semi-leisure chores' that do not fall under the definition of housework used in SILC. Overall, results suggest that retirement does not lead to a more equal distribution of 'core' household chores between genders.

JEL Classification: J22, J26, D1.
Keywords: retirement, house work, regression discontinuity.

Contents

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1 Introduction[1]

The study of the change in time allocation around retirement plays an important role in the analysis of home production. The abrupt decrease in market work can be exploited to study the reallocation of total time between housework and leisure. This is important for understanding the limits of the standard model of home production, the importance of social norms, and the strength of gender differences (Burda et al., 2006).

In this paper I provide new evidence regarding the change in time spent on producing household goods at retirement, using data from the 2007 cross-section of the Italian Survey on Income and Living Conditions (SILC). The focus is on the different behaviour of men and women, for which Italy is an interesting case study, given the strong gender differences in the engagement in housework over the entire life-cycle. The main problem of the empirical analysis is that the cross-sectional comparison between employed and retired individuals at any given age can provide a biased estimate for the quantity of interest. This is because retirees may have different preferences for leisure and housework (Rogerson and Wallenius, 2012). To manage this problem, I use the fuzzy Regression Discontinuity Design (RDD) outlined in Battistin et al. (2009), which exploits the discontinuities in pension eligibility induced by the Social Security rules. While they employed it to estimate the drop in consumption at retirement, I focus on time spent on housework, for which no information was available in their dataset.

To the best of my knowledge, only Stancanelli and van Soest (2012) used (fuzzy) RDD to address this question. They exploited the discontinuity in retirement at age 60 induced by the French system to estimate the causal effect of either partners' retirement on housework in couples. An advantage in studying the Italian setting is that eligibility depends on both age and years of contributions, generating discontinuities in retirement even when keeping one or the other fixed. Furthermore, the system has been subject to several reforms in the last two decades, hence different rules apply to individuals who retired in different years. At the same time, the Italian case is interesting in itself, because comparative international evidence shows that gender differences are stronger than in other countries, with Italian men spending much less time on household production (Burda et al., 2006).

The main results from my RDD estimates show that women increase home production by more than 400 minutes per week on average. In contrast, for men there is no evidence of a significant increase. This gender difference has no parallel in studies from Germany (Schwerdt, 2005; Luhrmann, 2010; Bonsang and van Soest, 2015), France (Stancanelli and

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van Soest, 2012), Spain (Luengo-Prado and Sevilla, 2013) or the US (Aguiar and Hurst, 2005; Szinovacz and Harpster, 1994; Szinovacz, 2000; Hurd and Rohwedder, 2005, 2006). Differently, for Italy the point estimate is very close to zero for married men living with their partner, suggesting that retirement does not lead to a more equal distribution of housework. However, there seems to be an increase for men living without a partner. Another possible reason is that retired men mostly dedicate their increased available time to activities, such as gardening, that are not part of the "core" household chores and are not included in the SILC definition of housework. Other important changes that the literature has found to be associated with retirement, in particular an improvement in health and a decrease in household size, do not explain the result because they do not take place in this case.

The structure of the paper is as follows. Section 2 provides a discussion of the economics of time allocation and retirement, with a focus on the Italian case. Section 3 presents the identification strategy, while Section 4 introduces the dataset. The main results are reported in Section 5, while robustness checks are conducted in Section 6. Section 7 discusses possible explanations for the results. The final section compares them with evidence from other countries and concludes.

2 Time allocation, home production and retirement

The standard Gronau (1977) model of labour supply with home production predicts that individuals who have a lower market wage and higher marginal productivity in home production are more likely to specialize in the latter. This depends also on the degree of substitutability between home produced goods and market purchased ones (Burda et al., 2006). Gender differences can therefore arise because of women earning lower wages on average. Differences in housework productivity are also a possible reason, but this is far from clear given that technological advances in home production have strongly facilitated most of the household chores, increasing the average productivity in housework but at the same time reducing the marginal one (Greenwood et al., 2005; Albanesi and Olivetti, 2007; Burda et al., 2006; Alesina et al., 2011). More importantly, gender differences may arise for those living in a couple due to the stronger bargaining power of men in the division of the full household income (monetary income plus the value of household produced goods, in the spirit of Apps and Rees, 1997). This may be traced back to social and cultural reasons, but also to the fact that men are more likely to be the breadwinner (Alesina et al., 2011). Despite of the gender differences in bargaining power, evidence collected by Burda et al. (2013) from 27 countries shows that the amount of total work (paid work plus housework) is quite similar across genders in rich countries, both in couples and among singles. This may be due to a social norm which results in the coordination of the amount of leisure.

Interestingly, this stylised fact is weaker in (rich) Catholic countries. Indeed, Italy is quite an outlier with respect to the iso-time norm. Empirical evidence from the time use surveys, discussed by Burda et al. (2006), Bloemen et al. (2010) and Addabbo et al. (2012), among others, shows that women do more total work (housework plus paid employment) than men. Furthermore, the gender gap in housework is much stronger with respect to other countries, although it had declined over time (Burda et al., 2006). Clearly, the smaller amount of time spent on paid work is likely to be due to fiscal disincentives due to the presence of family allowances (Alesina et al., 2011; Colonna and Marcassa, 2013) and by the large gender wage gap (Zizza, 2013). The female activity rate was still limited at 50.6 in 2007 (compared to a EU-27 average of 63.2), although women's labour supply had been growing faster than the one for men. Nevertheless, it is striking that Italian women appear to do more total work than men, irrespective of their employment status. If the gender differences are mostly due to the higher marginal value of men's time, it is interesting to understand whether retirement leads to a re-balance in time spent on housework, if not on total work. This may not happen for those living as a couple. Both sources of men's bargaining power are likely to persist after retirement, given that cultural and social norms are hardly affected by it, and pensions are obviously strongly related to the individual career.

3 Identification strategy

3.1 Variability in pension eligibility and retirement

In 2007, people could retire following two alternative paths. The first was to meet the requirement for a seniority pension, based on a combination of age and social contributions (35 years of contributions and 57 years of age for employees, or 58 for the self-employed). This path could be taken without age limits if the worker had at least 39 years of contribution (40 if self-employed). The second path led to an old age pension instead, and was based on a National Retirement Age (NRA) of 65 for men and 60 for women, plus a minimum requirement of social contributions.[2] For the purpose of identification, it is important that the two paths combined create a discontinuity at a certain combination of age and social contributions, which is defined by law and does not correspond to any other administrative rule. It seems quite safe to further assume that, if it was not for retirement itself, the year of eligibility would not correspond to any other significant event that could influence the amount of housework. If these conditions hold, the shift in the fraction of retired individuals at eligibility can be used to identify the change in household production, in a RDD design.

Apart from the discontinuity, eligibility rules have been changed almost every couple of years starting with the 1992 reform. Other major reforms took place in 1995, when the combined age-contributions requirement was introduced, and in 1997, when requirements were

[2]Among individuals who left employment in 2007, 58% of men retired with a seniority pension, exploiting their social contributions, while almost 80% of women did so with an old age pension (source: National Institute for Social Security - INPS).

strengthened again. Further changes were made in 2000, 2004, 2005 and 2007. To clarify the variability, Table 1 shows the rules that applied between 1997-2007. Manacorda and Moretti (2006), Battistin et al. (2009) and Bottazzi et al. (2006) provide evidence that the reforms had the intended effect on retirement behaviour and expectations (although the latter did not fully adjust to the new future entitlement rules). On the one hand, these continuous changes made it quite difficult for the single individual to manipulate his/her eligibility or to predict the timing of retirement exactly. This increases the likelihood that the discontinuity in eligibility can be successfully used to identify the effect of retirement on housework. On the other hand, these changes may have shifted different categories along distance to eligibility. Although these shifts are arguably exogenous, they may have altered the composition of the population around eligibility. In section 6.1 I discuss whether this affects the results.

3.2 Econometric model and issues

In order to provide a more formal discussion of the underlying assumptions, I follow Battistin et al. (2009) and define S_i as time to/from eligibility, $D_i \equiv \mathbf{1}[S_i \geq 0]$ as the dummy for being eligible, R_i as a dummy for being retired from work. Individuals are indexed by $i = 1, ..., N$. Let Y_{1i} be the time spent on home production if i was retired, while Y_{0i} if s/he was still employed. For each single individual, in the cross-sectional data I actually observe only one or the other, so that the observed outcome is (Hahn et al., 2001)

$$Y_i = \delta R_i + \varepsilon_i, \tag{1}$$

$$\varepsilon_i \equiv Y_{0i},\ \delta \equiv Y_{1i} - Y_{0i}. \tag{2}$$

The effect of retirement, δ captures the increase in homework that is associated with retirement. To simplify the discussion, I assume for the moment that this effect is constant across individuals. Theoretically, we expect δ to be positive, due to both the increase in total time availability and to the reduction in the value of market work. However, estimation is complicated by the fact that individuals for which housework is more valuable than market work are more likely to go into retirement earlier. Therefore, in the cross-section, retired individuals may have been doing more housework even if not retired, that is $Cov(R_i, \varepsilon_i) > 0$, leading to an upward bias.

Nevertheless, one of the main determinants of retirement status is the availability of a pension. Eligibility rules dictate that workers can claim it only starting from a specific moment in time. This creates a strong discontinuity at $S = 0$ in the value of retiring with respect to the value of continuing to work. Therefore, we would expect a jump in the proportion of retired individuals:

$$\text{(A1). } R_i = \gamma_D D_i + h_R(S_i) + \xi_i$$
$$\text{with } \gamma_D \neq 0;\ h_R(S_i = s) \text{ continuous at } s = 0;\ E(\xi_i | S_i) = 0.$$

Clearly, eligible and non-eligible individuals may be, overall, quite different with respect to the potential time spent on housework. For instance, older individuals are more likely to be caregivers, both for their own elderly parents (Attias-Donfut et al., 2005, 2008) and for their grandchildren (Hank and Buber, 2009; Dimova and Wolff, 2011; Rupert and Zanella, 2014; Battistin et al., 2014). Furthermore, changes in household composition (Battistin et al., 2009) and health (Coe and Zamarro, 2011) may, over time, alter the amount of time needed to produce household goods. Finally, if wages decrease in the last fraction of working life it might be that more senior workers reduce their overtime work to increase caregiving or other home production activities. More generally, the potential time spent on household production if the individual is not retired ($\varepsilon_i = y_{i0}$) is likely to be a function $h_Y(S_i = s)$ of the distance to eligibility, which depends on both seniority and age. However, it seems safe to assume that this function does not change discontinuously at eligibility:

$$\text{(A2). } E[\varepsilon_i | S_i = s] = h_Y(S_i = s),\ h_Y \text{ continuous at } s = 0$$

so that

$$Y_i = \delta R_i + h_Y(S_i) + \eta_i, \tag{3}$$
$$\eta_i \equiv \varepsilon_i - h_Y(S_i) \tag{4}$$

The validity of this assumption also depends on the ability of individuals to manipulate their eligibility status. In this case, it is difficult for workers to directly change their S at a specific point in time, for instance when they are close to retirement, because it depends on their entire history of contributions as recorded by the National Social Security Institution. More importantly, given that requirements have been subject to several reforms since 1992, individuals were not able to exactly predict the timing of their eligibility in advance.

Under assumptions (A1) and (A2), the average causal effect is equivalent to the ratio of the discontinuities in the reduced forms $E[Y_i|S_i = s]$ and $E[R_i|S_i = s]$ at $s = 0$, because any change in household production at eligibility can be attributed to retirement. However, identification is complicated by the fact that S is not directly observed. Instead, I recovered it using information on current age, age at first job, years spent in paid work, years of social contributions and job description. This introduces two additional problems.

First of all, in SILC I can measure time/to from eligibility only in discrete units (years).

As argued by Lee and Card (2008), this forces us to choose a parametric approximation:

$$Y_i = \alpha_0 + \delta R_i + P_i \alpha + \eta_i^* + \eta_i, \tag{5}$$

$$R_i = \gamma_0 + \gamma_D D_i + P_i \gamma + \xi_i^* + \xi_i, \tag{6}$$

where P_i is a polynomial in S_i possibly interacted with D_i, while $\eta_i^* \equiv h_Y(S_i) - P_i\alpha$ and $\xi_i^* \equiv h_R(S_i) - P_i\gamma$ can be interpreted as the residuals from the Best Linear Projection ($BLP[\cdot]$) of the true functions h_R and h_Y on the vector P_i. With the same polynomial in both eq. (5) and (6), the causal effect δ can be recovered using 2SLS and instrumenting R_i with D_i, as long as:

(A3). $BLP[\eta_i^* | D_i, P_i] = BLP[\eta_i^* | P_i] \equiv 0.$

This implies that the approximation does not introduce any discontinuity in the main equation of interest (5), so that D_i can be excluded from it. Note that the equation for retirement is only a first stage, and therefore we only need it to be the best linear projection.[3] With this caveat in mind, the main estimates will employ a simple 2SLS strategy, choosing the polynomial that provides the best fit in the reduced form for Y.[4]

The second problem is due to the fact that S is discrete because it is rounded in years. Dong (2014) shows that the OLS estimator for the discontinuity in Y at eligibility is biased. Nevertheless, she showed that, under certain conditions, the bias can be recovered if the marginal distribution of the true continuous distance is known. I followed her suggestion to calculate bias-corrected estimates assuming a uniform distribution inside each year interval.

Finally, if there are heterogeneous treatment effects δ_i, then I can still interpret the 2SLS coefficient as a Local Average Treatment Effect for those who retire as soon as eligible. In this case, I also need R_i as a deterministic function of S_i to be monotonic *near* $s = 0$, while δ_i and $R_i(S_i)$ must be jointly independent of S_i (see Hahn et al., 2001). This can be defended using the same arguments advanced for assumptions (A1) and (A2). Despite its local properties, this local average effect is still of interest. First of all, a substantial fraction of individuals leave work as soon as they meet eligibility, or shortly afterwards (as I will show using data from SILC). Secondly, this is the quantity of interest if we want to understand the effect of marginal changes in eligibility rules on housework.

[3]The reason is that, under assumption (A3), the discontinuity in the BLP of Y_i on (D_i, P_i) would be equal to $\delta\gamma_D^*$, where γ_D^* is the coefficient associated to D_i in the BLP of R_i on (D_i, P_i). Therefore 2SLS would still be consistent. Caution should be applied, because if the equation for R_i is only a BLP, then testing for a discontinuity in it is not equivalent to testing the presence of a discontinuity in the true retirement equation. Therefore I may be using a discontinuity in retirement that does not exist, for instance confounding a jump with a kink.

[4]One might prefer to look at the two reduced forms $E[Y_i|S_i]$ and $E[R_i|S_i]$ separately and then estimate δ as the ratio of the two discontinuities. In this parametric setting, however, using 2SLS has the advantage of being clearer, given that it is equivalent to an instrumental variable approach.

4 Data and descriptives

4.1 Definition of housework, retirement and eligibility

The Italian component of the European Union Survey of Income and Living Conditions is a stratified sample of household population conducted by the Italian National Statistical Office (ISTAT) every year since 2004.

In this paper, I use the 2007 wave, because respondents were asked *"On average, how much time per week do you spend on domestic and family-related work (household chores, shopping, caregiving of other members), in hours and minutes?"*.[5] From the answer to this question I define the main variable of interest, Y_i, measured in minutes per week. The survey collects only the aggregate housework time, without distinguishing between different components. This is a clear limitation with respect to the Italian Time Use Survey 2008-2009 (TUS). Unfortunately, the TUS does not collect information on years of contribution, so that it is not possible to replicate the RDD.

In the definition of the main variables (R and S) I follow Battistin et al. (2009) as close as possible. I defined workers as individuals with "employed" as self-reported employment status, excluding those who did not work in the week prior to the interview because of being temporarily unemployed or under a temporary public layoff scheme called *cassa integrazione.* In line with the identification strategy, based on eligibility for a pension, I identified retired individuals as those who reported not to be working in the week prior to the interview because they were "*in pensione da lavoro"*, literally "in work-related pension". The chosen definition of retirement implies that it corresponds to zero hours of paid work.

The calculation of distance to eligibility requires information on the individual working history. One crucial variable is number of years the individual has contributed towards social security, which is directly asked to all respondents who have some working experience. It also includes years that the individuals have covered ex-post by paying a fee. This could be done for the time spent in military service or in higher education. The latter is particularly problematic in case of individuals sorting around the threshold, and will be discussed in Section 6.1. Although the SILC variable is recorded in years, the actual institutional unit of measurement is shorter (weeks).

Another crucial variable for retired individuals is the year of retirement, which is unfortunately not directly reported in SILC. Instead, I calculate the age at retirement as the age in which the respondent began the first regular job plus the number of years spent in a paid job, plus one. The final correction is taken to account for rounding (see Appendix A for a

[5] This information in SILC was already used by Addabbo et al. (2012) to study the interaction between working age partners in paid employment and (unpaid) housework. The question on housework was also asked in the following wave (SILC 2008), but unfortunately it contains a large number of missing values (18.05%), which casts doubts on its validity. The 2010 wave also contains a similar question, but (i) it is posed only to a fraction of the sample (ii) the years of contributions variable has not been released with the microdata, so that the RDD design cannot be implemented.

more detailed discussion of this point).

Distance to/from eligibility S is calculated as age at interview minus age at eligibility. Because of the different reforms that took place between 1992 and 2007, to calculate the latter I first need to understand which rules to apply to each individual. To this purpose, for retired individuals I use the requirements that applied in the year in which s/he went into retirement, while for workers those applying in the year of the interview. Essentially, the requirements used in the calculations are those from Table 1, focusing on the year of retirement for retirees and on 2007 for workers (requirements for the years before 1997 are not reported in the Table, but are available on request). The employment category refers to the last job for retirees and to the current one for workers (see Appendix A for more details).[6]

Using the rules applying to each individual, I then calculate the age at which s/he became or will become eligible for a pension. For the seniority path, the rule is a combination of contributions and age. For retirees I assume that years of contributions grew 1:1 with respect to age in the years before retirement, and would have grown at a similar pace in the following years if the individual had not retired. Similarly, for workers I assume a 1:1 increase of contributions with age around 2007.[7] For the old-age path, age at eligibility is basically already defined by the requirements, apart from some old individuals who still do not meet the minimum contribution requirement at the time of interview. Finally, to identify age at eligibility I take the minimum between the age at eligibility for the two paths.

The definition of eligibility used in this paper has some limits, that I discuss in Section 6.2, where I also show estimates based on an alternative strategy. More in general, the process of recovering S from other survey information clearly introduces measurement error, which smooths the discontinuity in R at $s = 0$. In particular, if S was correctly measured I should not observe anyone in retirement status before being eligible, that is when $S < 0$. As argued by Battistin et al. (2009), 2SLS is consistent as long as the measurement error process is statistically independent from (Y, R) conditional on the true value of distance to/from eligibility. One concern is that S is necessarily calculated differently for workers and retirees. In particular, the need to determine the year in which the individual has gone into retirement introduces an additional source of measurement error that has no counterpart for workers. For women, whose retirement behaviour is influenced more by the National Retirement Age (NRA), in the Appendix I also estimated the effect on household production

[6] There were some differences for blue-collars and for those who started their career before 1992. I took them into account. In the public sector the rule for the old-age path differed between specific sub-categories (e.g. central administration, military, local authorities). I always apply those for the central administration, as I am unable to identify the single groups in the data. Anyway, in Section 7.1 I show that results are not driven by the public sector category.

[7] The assumption can also be problematic in case of a job-loss, as individuals are covered by unemployment benefits and/or job supplement schemes lasting only for a limited amount of time. Nevertheless, it would make sense to apply a different coefficient only if we knew with certainty that some workers will interrupt contributing for some years. Applying uniformly a progression coefficient smaller than one would simply move workers proportionally away from the threshold.

using only age as running variable. Unfortunately, this strategy does not work for men, whose retirement pattern is smooth with respect to age.

4.2 Descriptive analysis

To better understand the content of the SILC information on housework, Table 2 compares it with the TUS survey, where "family related" work consists of cooking, doing the dishes, cleaning the house, doing the laundry, sewing, knitting, shopping, general administrative work, and caring for adults or children (inside or outside the household). It also includes gardening, taking care of pets, maintenance of the house and vehicles. To focus the analysis on a sample relevant for the analysis, I focus on retirees and workers (according to the self-defined economic status, which is comparable across the two surveys) between 50 and 70 years of age. On average, time spent on housework is lower in SILC with respect to "family related" work from TUS (column TUS (A)). The difference is proportionally larger for men, in particular when retired. The difference is likely to be related to the fact that the general question posed in SILC seems to exclude some activities. While caring and shopping are explicitly mentioned, "household chores" are likely to be associated with cooking and "core" household work, as defined by Stancanelli and van Soest (2012, pg. 7): "cleaning, doing the laundry, ironing, cleaning the dishes, setting the table, and doing administrative paper work for the household". Differently, it might exclude "semi-leisure" chores, such as gardening. In column TUS (B) I redefined the variable in TUS, keeping only shopping, cooking, caregiving and "core" household work. As expected, the averages for men drop down, in particular for retirees, getting closer to those from SILC. Finally, in SILC, the caregiving part of the question seems to explicitly refer to other members of the household, although respondents may have considered other relatives as well because the general question mentions both "domestic" and "family" work. If I exclude care outside the household from TUS (column TUS (C)) leads to averages that are closer to SILC.

Focusing on the overall SILC sample, with no age restrictions, Table 3 shows predicted values of minutes/week spent on housework, market (paid) work and total work from a regression of these variables on socio-economic characteristics (see Addabbo et al., 2012, for similar descriptives on housework in SILC). For both men and women, holding other variables constant, workers and students are those spending less time on housework. Although employed women do less market work than men, they dedicate much more time on housework too, hence their total work is larger. Retirees spend a considerably larger amount of time on household production. Retired women are actually close to housewives, although still distant by almost 200 minutes/week.

For women, the differences based on other characteristics seem to follow the pattern of market wage: where this opportunity cost is larger (the North, densely populated areas, more educated individuals), the time spent on housework is smaller. Nevertheless, this is not necessarily the main explanation, because the pattern of market work is less related to these characteristics. For men, those living in the North are more involved in both home production (consistently with results from Bloemen et al., 2010) and market work, while heterogeneity by education and population density is less relevant.

Finally, while married men living with their wife spend less time than those not living with a partner, the opposite is true for women. The distribution of housework is more balanced across genders in cohabiting couples (who represent only 5 percent of total couples in the sample). This is in line with previous literature from Germany (Barg and Beblo, 2012), which suggests that this can be explained by sorting into marriage of partners who agree on a "traditional" division of work, and by a specialisation-reinforcing effect of marriage.

Figure 1 restricts the attention to the time spent on household production by workers and retirees aged 50-70. Two main stylised facts can be drawn from this. The first is that, at any age, the average Y is larger for retirees than for workers. Secondly, for females the difference between retirees and workers is almost double that for males, which suggests that retirement leads to an increase in gender-gap. Nevertheless, the cross-sectional comparison between individuals in the two employment status may be misleading, because at any age those who are already retired may have different preferences for house and market work. To address this endogeneity concern, in the rest of the paper I exploit the increase in the proportion of retirees at pension eligibility, in an RDD framework.

4.3 Sample selection for the RDD

A full table reporting sample selection is available in Appendix A. For the RDD I focus only on workers or retirees. I also exclude all proxy interviews, which is the case when another household member provides the information about an individual who is not available at the time of interview. The reason is that they are likely to increase measurement error and not be particularly reliable for Y. There are few missing values for housework.

As in Battistin et al. (2009), I keep only the window $S \in [-10, 10]$, in order to limit the influence of observations far from the eligibility threshold, and I exclude observations with $S_i = 0$. The fact that contributions, age at first job and time spent on paid work are measured in years implies that the observed S is obtained by rounding either up or down, so that $S_i = 0$ includes both cases at the left and at the right of eligibility. One simple solution, suggested by Dong (2014), is to discard observations with $S_i = 0$.

I do not use sample weights. Nevertheless, in section 5.3 I discuss what happens when I include stratification variables in the regression or I employ sample weights.

The final sample includes 3970 observations for men and 2701 for women. Full descriptive tables are reported in Appendix A due to space constraints. The average age is similar across genders (56.9 for men and 56.5 for women) but, as expected, men spent on average more years in paid work (32.9 vs 28.8) and therefore contributed longer towards social protection. Women are more likely to be public employees (36 percent as opposed to 19 for men) and less likely to be self-employed (24 vs 28). They are also less likely to be married (71 percent as opposed to 82 for men), mostly because a larger fraction of them is widowed (10 vs 2). The level of education is quite similar across genders, with slightly more than 10 percent holding a university degree, around 37 percent with a high school diploma and less with lower qualifications. Also, their health status and the distribution across areas are similar. Clearly, in the presence of heterogeneous effects with respect to these characteristics, differences between genders may cause different estimates for δ. In section 7.1 I explicitly look at heterogeneity. I do not discuss here differences by eligibility status, because what matters for the RDD are those around $S = 0$, which are examined in Section 6.1.

5 The change in housework at retirement

5.1 Graphical analysis

Figure 2 shows the pattern of retirement and household production with respect to the distance S to/from eligibility from a pension. For both genders I observe a small proportion of individuals who retired before meeting the eligibility criteria. Between $S = -1$ and $S = 1$ there is a large step-up in the fraction of retirees, which continues at a declining rate until reaching 90% or more at $S = 10$.

Time spent on housework increases slightly before eligibility is met. After it increases progressively for men, but there is no clear evidence of discontinuity. I observe an increase at $S = 0$ around 50 minutes/week, but it is followed by alternate drops and rises. For women, time spent on home production is quite constant before eligibility. I then observe a jump at $S = 0$ by nearly 150 minutes/week, followed by an increase. A linear polynomial predicts a discontinuity. A quadratic does not, but it is important to note that it seems to overfit the mean for Y at $S = 0$, predicting a lower value.[8]

5.2 Estimates of the jump in retirement at eligibility

To test for the presence of a discontinuity in retirement at eligibility, Table 4 shows regressions of R_i on the eligibility dummy D_i, a polynomial in S_i and their interactions. For model selection, I focus on minimizing the Akaike (AIC) and Bayesian (BIC) information criteria (Lee and Lemieux, 2010). Lastly, I test whether the constraints imposed by the polynomial specification are rejected, using Lee and Card (2008) G statistic. It compares the regression with an unrestricted one including a dummy for each value of S.

[8]The comparison of predicted values with the sample average at eligibility is useful in evaluating the polynomial fit, because I am not using observations with $S_i = 0$ in estimating the regressions.

For men (columns (1)-(3) in Table 4), there is clear evidence of a discontinuity in retirement behaviour. This is in line with previous findings (Battistin et al., 2009) and with theoretical expectations. For women (columns (4)-(6) in Table 4), the estimated discontinuity in R at $S = 0$ is small and not statistically significant using the 3rd order polynomial. However, with a quadratic it is around 24 percentage points and statistically significant with either robust or clustered standard errors. The statistical tests do not give clear indications. The G test is passed at the 5% level both with the cubic and the quadratic, although the latter has a smaller p-value. The Akaike information criteria lead us to choose the cubic regression, but the Bayesian is minimized for the second order, and it should be noted that the R^2 changes only at the third decimal place between the two models. Given the strong institutional reasons for expecting a jump at eligibility, I find it reasonable to focus on the quadratic specification and take it as supporting evidence in favour of the presence of a discontinuity. Dong's correction suggests a smaller jump (0.183), but still statistically significant at conventional levels.

5.3 The effect of retirement on housework

Given the evidence of a jump in retirement at eligibility, I expect that, in the presence of an effect on household production, I should also observe a discontinuity in Y around $S = 0$. In Table 5 I show regressions of Y on D on linear polynomials in S, interacted or not with S. I do not consider higher orders, given that information criteria invariably lead us to prefer the simplest specification and graphical analysis did not show large differences (extended tables including quadratic polynomials can be found in Appendix A).

Despite the strong evidence of a jump in retirement at eligibility for men, none of the estimated models show a discontinuity in the average time spent on home production (Table 5, upper panel, columns (1)-(2)). Regression analysis is therefore in line with the intuition resulting from graphical inspection. To recover the causal effect δ of retirement on housework, I use 2SLS, instrumenting R with D. The highest estimate (Table 5, lower panel, column (2)) is 73 minutes/week, obtained when only S is included. It is only 25 percent of the relative OLS estimate (see the last row of the Table), and it is not statistically significant at conventional levels. The change is more significant if compared to the counterfactual at eligibility. Without covariates, and with $S = 0$ corresponding to eligibility, this is captured by the constant (436 minutes/week), so that the increase is around 17 percent. However, it is still far from what would lead to a more equilibrated distribution of housework across genders. In terms of magnitude, consider also that the equivalent 2SLS regression using time spent on market work as a dependent variable predicts a drop of 2468 minutes/week associated with retirement. Therefore, of this increase in time, only around 3 percent is spent on housework.

Differently, for women, columns (3)-(4) of Table 5, upper panel, show a discontinuity in Y at $S = 0$, around 219 minutes/week using a linear polynomial without interactions. For both specifications the G test fails to reject the null of correct specification.[9] Dong's correction does not lead to different conclusions.

Using the simplest linear specification, and instrumenting R with D, the 2SLS estimate for δ (Table 5, lower panel, column (4)) is 430 minutes/week, statistically significant at the 5% level. If expressed as an elasticity with respect to the counterfactual at eligibility (1579 minutes/week), it is nearly 28 percent. It also represents a substantial fraction (22 percent) of the increase in available time given by the drop in paid work estimated by the corresponding 2SLS regression (1961 minutes/week). If I use a linear polynomial with interactions (Table 5, upper panel, column (3)), the estimated effect is quite similar.

Although the RDD design does not require including covariates, one may want to understand whether including some standard socio-demographic characteristics leads to different results. To condition on other variables, one solution is to adopt a parametric framework, where the counterfactual $\varepsilon_i \equiv Y_{0i}$ depends linearly on these additional variables, which, therefore, enter all regressions as a vector of covariates X (see Frölich, 2007, for a non-parametric alternative). I tried by including dummies for geographic area, population density, education and employment category. For both genders the estimates for the discontinuity in retirement are basically unchanged, while they are slightly larger for Y. For men, the highest estimate for δ is 100 minutes/week (Table 6), but not statistically significant at conventional levels and still far from the OLS results. The estimates with covariates are also bigger for women: using a linear polynomial with interactions, the result is 493 minutes/week, while it is 466 when only S is included. Overall, the differences with the main estimates are not particularly large and they still lead to the same conclusions. Given that the covariates include the stratification variables, these regressions are also useful to assess whether the regressions are influenced by the choice of not using sample weights. I also tried using them, but results (available on request) are quite similar to those presented in the main tables.

6 Robustness checks

6.1 Discontinuities in baseline covariates

One way to check the plausibility of the continuity assumption (A2) is to inspect whether some baseline characteristics exhibit discontinuities at eligibility. I focus on three sets of variables: (i) geographical dummies for area of residence and population density (which

[9] Although a second order polynomial shows no discontinuity (see Appendix A), the information criteria indicate a preference for the simpler polynomials. A very similar estimate for the discontinuity (207 minutes/week, p-value 0.021) is obtained by a regression of Y on D, S and S^2, with no interactions as in Battistin et al. (2009).

were used for sample stratification), (ii) dummies for the highest educational achievement, (iii) employment categories.

Geographical dummies are relatively smooth (Table 7).[10] We observe an increase at eligibility in the proportion of men residing in the Centre and in the proportion of women in densely populated areas. However, a test for joint significance of all the discontinuities in geographical dummies fails to reject the null at conventional levels for both genders.

Educational dummies do not show discontinuities for women, while for men we observe an increase in the proportion of high school graduates at eligibility and a decrease in those who only completed the middle school degree. This discontinuity is a problem if it is evidence of endogenous sorting of individuals. In the present context, one possibility is that they were able to exploit rules related to their educational level: in Italy university graduates are allowed to pay-back social contributions to cover the years of higher education and become eligible for a pension earlier. But in this case I should have also found an increase in university graduates at eligibility, while I found no evidence of such a discontinuity. Another problem could be the 1963 educational reform, which had an effect on cohorts from 1949 (see Brunello et al., 2012, p. 19). It seems that this is a minor issue in this context. By construction $S = 0$ does not include a single cohort: the proportion of cohorts from 1949 at $S = -1$ is 0.725, quite close to the proportion at $S = 1$ (0.621). Moreover, if this was the problem, we should expect a decrease in the educational level at eligibility, given that those at $S \geq 0$ are older individuals. To further inspect the change in overall educational level, I calculated the total years of schooling by attributing the official length to each degree (this also accounts for some shorter vocational degrees included in the "high school" dummy). As shown in Table 7, there is no evidence of a discontinuity for both genders. I also calculated the difference between age 6 and the age at which the individual completed his/her highest degree. This is larger than years of education, both because of grade retention and individuals taking degrees later in life. This variable seems to show a drop in the "age at highest degree - 6" variable, not necessarily in line with an increase in the educational level. It is not statistically significant, although the joint test for the discontinuities in both additional educational variables has p-value 0.046.

Employment categories are relevant in the definition of S, as rules are quite different between employees and self-employed, and between the public and private sector. For women there is no evidence of a significant discontinuity in these categories. In contrast, for men we observe a decrease at eligibility in the proportion of public employees, compensated for by an increase in the self-employed. Although individuals move across different occupations, it is unlikely that they moved in order to gain from different eligibility rules, also because in this case we would expect an increase in public sector employees (with more generous

[10] I present regressions including both $(1-D) \times S$ and $D \times S$ because for some variables this specification is preferred to the one using only S according to the G test (in particular for those variables used in building S). However, estimates without interactions (available on request) lead to the same conclusions.

rules). One alternative explanation is that the retirement reforms created some discontinuities across workers with different employment histories. The source of these differential treatments does not seem to be precisely manipulable by the single individual, given that the repeated changes in the rules between 1992 and 2007 were hardly predictable at the time s/he started his/her career. However, the resulting discontinuities make individuals across eligibility not completely comparable. In Section 7.1 I show that results for men are not driven by a specific employment category.

A discontinuity in the density function at eligibility might be a sign of individuals sorting around the threshold (McCrary, 2008), even if a continuous density function is neither a sufficient, nor a necessary condition for identification. In this case, I cannot directly use McCrary's test because the running variable is discrete, while the test is designed for the continuous case. Nevertheless, in Appendix A I show plots with the fraction of observations for each value of S. There is no change in the density at $S = 0$ for men. For women there is a drop of around 1 percentage point, if estimated with a linear fit. However, if individuals were able to manipulate their distance to/from eligibility, there would be no reason to expect them to misreport it in order to become ineligible. Given that retirement is not generally compulsory at $S = 0$ according to the Italian rules, most individuals have an incentive to manipulate S_i in the opposite direction, so that I should find an increase in density at eligibility. I therefore do not take the observed drop as evidence of sorting.

The selectivity of women in and out the labour force may create a problem for the RDD design if those who leave work at an early stage of their career, after having accumulated some years of contributions, start reporting to be "retired from work" when they reach the NRA. This would imply that they enter the sample only at eligibility. This does not seem to be particularly relevant in the sample, given that there is no evidence of an increase in the density at $S = 0$. Furthermore, we would expect a decrease in age at retirement and year of retirement at eligibility, as they are built by summing the age at first job and the years spent in a paid job. I estimated the discontinuity in these two variables at $S = 0$, as in Table 7, and in both cases there is no evidence of a significant change.[11]

6.2 Sensitivity analysis

Table 8 discusses additional robustness checks. I show results for the simplest specification, including only S as additional covariates, but results including interactions with D and/or covariates are similar (available on request). Column (1) shows that not adding one to the age at retirement basically leaves the results unchanged. Also keeping observations with $S = 0$ (column 2) does not lead to sensibly different results.

One crucial issue in the definition of S is the choice of using the rules applying in the

[11] I also run the 2SLS regressions by selecting only women with more years of contributions (I tried both with ≥ 20 and ≥ 30), or whose year of retirement is closer to 2007 (≥ 1995 or ≥ 2000). Results are in line with the main conclusion, with a statistically and economically significant coefficient on R.

year of retirement for retirees and in 2007 for workers. An alternative is to identify the *first* year in which an individual could have gone into retirement. For instance, take an individual aged 55 in 2000 (with 35 years of contributions). According to the rules, s/he could have gone into retirement in that year. Instead, suppose s/he decided to work until 2002 and s/he went into retirement with the new rules (57 years of age and 37 of contributions). For this individual, the method used in this paper defines 57 as age at eligibility. However, his/her *first* age of eligibility is 55. I prefer the former definition for two main reasons. First of all, the change in the rules was not always as smooth as in this example, in particular for public sector employees during the nineties. Secondly, calculating the *first* year of eligibility requires an additional level of complication that may introduce further measurement error. Note that, similarly to what happens for workers, the chosen method may bring retirees closer to the discontinuity. However, histograms for the distribution of S (see the Appendix) are not hump-shaped, nor they show a peak around $S = 0$.

Nevertheless, I also run the regressions based on the *first* age of eligibility. For retirees I still need to assume that, up to the year of retirement, contributions had grown 1:1 with age, and that they would have kept growing at the same pace had the individual not stopped working. Similarly, for workers I assume a 1:1 progression of contributions with respect to age around year 2007. I also apply, after 2007, the increase in requirements according to the last reform, Law 243/2007, passed in July 2007, which is the relevant one given that interviews took mostly place in the last quarter. Using this alternative definition of eligibility, the discontinuity in retirement is weaker for both genders. The jump in housework is similar to the main estimates, with a small and not significant discontinuity for men and an increase around 200 minutes/week for women. The resulting 2SLS estimate (column 3) is in line with the main conclusions, with a large increase for women and a smaller one for men.

For retirement status, Battistin et al. (2009) used a similar definition as the one employed in this paper, but they directly controlled for whether individuals were actually recipients of a pension. This cannot be done using SILC cross-sectional data, because income information refers to the calendar year previous to that of the interview. In the selected sample used for estimates, 6% of those self-defined as workers report having received a work-related pension in the previous year, while among those classified as retired 13% report that they have not received a pension. I did not correct their status because this is likely to correct measurement error in R only at some positive distance from eligibility, while leaving the same situation at $S = 0$. The reason is that for someone who retired in the current year we do not know whether s/he is in receipt of the pension or not. If I drop from the sample the workers who receive a pension and the retirees who did not (column 4), estimates of the effect of retirement on household production are smaller for women, but still close to 400 minutes/week and statistically significant. For men they are larger (around 120 minutes/week) but not statistically significant at conventional levels.

The chosen definition of retirement implies that retirees have zero hours of work. Stancanelli and van Soest (2012), instead, used the self-defined economic status, so that some of the retirees may be working for some time during the week. However, in SILC 2007, even among those whose self-reported status is "retired from work", only 1.2% of men and 0.6% of women worked at least one hour in the week previous to the interview. I also estimated the main regressions setting $R=1$ if the individual's self-defined occupational status is "retired from work" (column 5). Results for men are almost unchanged. The estimated effect for women is smaller than that in the main estimates, but still larger than men and statistically significant. The smaller result may actually be explained by some degree of misclassification. Among those reporting to be "in work-related pension" (which corresponds to the definition used in the main results), 6% of women have "housewives" as occupational status, so that it is possible that individuals receive a pension, but do not report to be "retired from work".

Results might be driven by the choice of window size. I checked how they change when this is decreased, using 2SLS regressions including $(1-D)\times S$, $D\times S$ as covariates, and using D as an instrument for R (see graphs in Appendix A for the results). The estimates for men oscillate around zero and they are never statistically significant at the 10% level. For women, $\widehat{\delta}$ is quite stable for $|S|\geq 5$. At size 4, the estimate is almost zero, while for size 3 and 2 the first stage F is very small. One reason is that four points are probably not enough to obtain precise estimates of the linear fit with interactions. Another is that, given that measurement error smooths down the discontinuity in retirement, I need other points away from $S=0$ to partially correct for it. Nevertheless, even using only $S\in\{-1,1\}$ and a simple Wald estimator, the estimate is 431 minutes/week, very similar to the main results, although clearly much less precise (s.e. 358).

Finally, an alternative would be to focus only on the discontinuities with respect to age, neglecting social contributions. This can reduce the problems related to measurement error. Retirement is smooth with respect to age for men, and therefore a RDD cannot be implemented. Women's behaviour is instead more affected by the NRA at 60. Results for housework are broadly in line with those presented here, but they tend to be more influenced by the introduction of covariates. See Appendix B for a discussion.

7 Possible explanations

In the next subsections I discuss different mechanisms and explanations for the main results: (i) the presence of substantial heterogeneity across groups characterized by different opportunity costs; (ii) the relevance of other changes in health or household composition that may reduce the demand for household goods; (iii) the importance of caregiving; (iv) the role played by the absence of "semi-leisure" chores in explaining the different results for men; (v) the distribution of housework between partners.

7.1 Heterogeneity

Theoretically, it is interesting to understand whether those groups that had a larger opportunity cost of housework while in working age, due to a higher market wage, are also more likely to increase the time spent on it after retiring, given that this opportunity cost becomes irrelevant. Table 9 shows the 2SLS estimates splitting by education, area and category. Because of higher average wages, the opportunity cost is likely to be larger for more educated individuals, living in the North and in densely populated areas (because of the urban wage premium, see Addario and Patacchini, 2007). With respect to work category, public employees are expected to have smaller opportunity costs, because their contracts offer more possibilities to take paid and unpaid days off if they have family needs, such as an elderly parent with impaired health. Given the small sample size, I keep the simplest specification, with only S as additional covariates. Nevertheless, results with both $(1-D)\times S$ and $D\times S$ are quite similar (see Appendix A), although they are less precise.

For men, the estimated effect is indeed economically significant for college graduates (176 minutes/week) and in the North (148 minutes/week), although not statistically significant. For private and public employees, the estimated effect of retirement is larger than for the self-employed (113 and 105 minutes/week as opposed to -21), but not far from that which is estimated for the whole sample. The only estimate which is statistically significant, even if only at the 10% level, is the one for men living in densely populated areas (34% of the sample), which is approximately 225 minutes/week, similar to the OLS results. This is again in line with the opportunity cost explanation.

The educational heterogeneity for women is less in line with the interpretation related to wage differentials. While women with a high school degree exhibit estimates for δ larger than those with a lower degree, the change in time spent on home production is negative and large for college graduates (Table 9). However, it is probably driven by the weakness of the instrument and by the small sample size. In contrast, the geographical pattern is more in line with the differences in the housework opportunity cost. The effect is stronger in the North, where wages are higher and women do more paid work (Table 3). It is also stronger in densely populated areas and in intermediate ones (more than 600 minutes/week), while it is negative, but not statistically significant in thinly populated areas. With respect to job type, the increase for public sector employees (325 minutes/week) is smaller than for other categories, in line with the fact their contracts already allow them to take days off for family needs.

An alternative explanation of the heterogeneity is that the increase in housework with retirement should be larger for those individuals with higher substitutability between home production and consumer expenditure (see Rogerson and Wallenius, 2012). First of all, in densely populated areas there is a more developed market for goods and services that could replace, at least partially, housework (e.g. prepared food or helpers/housekeepers).

This could explain the larger estimates found in those areas. Secondly, looking at both genders it seems that the increase in housework at retirement is larger in richer households, usually residents in the North and in densely populated areas. More wealthy households may have larger substitutability between home production and consumer expenditure, so that they engage less in housework when active in the labour market. For instance, they may be more able to acquire services or goods on the market which reduce their need for housework. I further split the sample, still separately by gender, between households with equivalent income above and below the median.[12] For men there are small differences, while they are quite large for women, with those in higher income households showing a much larger increase. However, this group starts from a lower level of housework, as the counterfactual prediction for a female worker at $S = 0$ (derived from the same estimates) is 1472 minutes/week, against 1711 minutes/week for the below median income sub-sample. These results are in line with the explanation in terms of different substitutability, because part of the larger increase found for the first group can be interpreted as a catching-up with the other. Still, sizeable effects are only found in the female sample.[13]

For men, the higher coefficients for those who are more educated or living in the North may also be related to differences in social norms and bargaining power within couples, which are likely to be less gender-biased among these groups. Indeed, from descriptive regressions (Table 3) we know that men in the North are actually more likely to do housework given other characteristics, although there are not large differences by education.

7.2 Other changes

Other changes caused by retirement may have an off-setting effect on home production (see Appendix C for full results). Coe and Zamarro (2011), using data from the Survey of Health, Ageing and Retirement in Europe (SHARE), found evidence of a health-preserving effect of retirement for men, with a decrease of 35 percentage points in the probability of reporting fair, poor or very poor health. Using the same variable in my dataset, I also find a negative effect for men (using 2SLS and S as covariate), but much smaller (6 percentage points) and not statistically significant at conventional levels. For women the estimate is also negative, but still smaller (4 percentage points), and not statistically significant.

Household size may change if co-resident adult children leave when one of the parents retires. This may result in a lower demand for home produced goods, such as food. Battistin et al. (2009) found a reduction in household size by 0.3 with the retirement of male household heads, explained by adult children leaving the parental home. In SILC, the estimated

[12] Household income refers to the year previous to the interview. I used the simplest OECD equivalence scale, i.e. the square root of household size.

[13] It would be interesting to see whether households reduce their use of helpers/housekeepers. However, although the information was collected in SILC 2007, it has not been made available in the microdata.

effect of retirement is actually positive for men and negative for women, but very small.[14] The estimates of Battistin et al. (2009) refer to years 1993-2004. On the one hand, between 1993 and 2007 there was a rise in retirement age. This may imply that, at the time that parents retired, children were older in 2007 than in 1993, so that they were more likely to leave the household. On the other hand, the deterioration of expectations about economic growth after 2007 could have reduced the incentives for adult children to form independent households. The latter trend may have offset the former.

7.3 Caregiving

It would be important to understand whether the changes in housework estimated for women is likely to be due only to caregiving. On the one hand, this could provide useful direct evidence for the rising stream of literature which suggests that the presence of retired grandparents (in particular grandmothers) has a strong effect on younger mothers labour supply (Arpino et al., 2010, 2012; Battistin et al., 2014). On the other hand, caregiving also includes elements of leisure, in particular for grandchildren care, and therefore it may be interesting to analyse it separately.

As already discussed in Section 4.2, it is not clear whether the SILC question picks up caregiving for non-coresident members. Assuming it does, it is still not possible to directly breakdown housework into its different components, nor we have information on the extended family network, such as the presence of grandchildren. However, some tentative indications can be obtained from the TUS sample, by comparing retirees and workers around retirement age.[15]

For women between 55 and 65, around 11 percent of the higher average daily housework among retirees is related to taking more care of children (mainly outside the household), while 5 percent to more adult care. Caregiving is, therefore, relevant in explaining the difference between retirees and workers, but it does not seem to be the main component of it.[16] If these fractions are applied to the 430 minutes/week increase estimated in SILC, nearly 70 minutes would go to caregiving, of which 47 for children. Obviously, these calculations are only indicative, as nothing ensures that the proportions recovered in the simple comparison made in the TUS sample can be directly applied to the RDD estimates from SILC.

[14] There is also no evidence of a change in the probability of being married. Stancanelli (2014) found, for France, that the likelihood of divorce increases at retirement.

[15] In TUS data it is still not possible to identify only care for grandchildren, as the help for children/adults is only differentiated by whether the recipient is a member of the household or not.

[16] It must be mentioned that information collected by single daily diaries (TUS), or from questions about "average" use of time, tend to underestimate caregiving. The reason is that this activity often takes place only in specific days or periods (e.g. summer for grandparenting).

7.4 "Semi-leisure" chores

From a theoretical point of view, it is strange that men do not significantly increase their home production after retirement, given the strong increase of their available time. One possible reason, supported by the comparison with TUS data, is that at retirement men put most of their effort into "semi-leisure" chores, such as gardening or house-repair, which are not in the SILC definition.

The best that can be done on the SILC data is to focus on some specific groups where there is not much scope for these chores, so that retired men cannot spend most of their increase in available time on them. For instance, men in densely populated areas may have less opportunities for gardening or similar activities. This can explain why I find a positive increase in their housework, although still much smaller than the related estimate for women. To further inspect this, I also split the sample for men between those living in households with a private garden and those who are not. The estimate for the subgroup without a garden is indeed larger (142.8 vs -18.7), although both are not statistically significant at conventional levels (with large standard errors, 114.1 and 134.2 respectively) and still small if compared to the other gender.

These results suggest that part of the difference with respect to women may be explained by the absence of these activities in SILC definition. Indeed, Stancanelli and van Soest (2012) found that, in France, men's increase in time spent on home production was mostly in this category.

7.5 Gender, marital status and couples

Another explanation for the different results for men is that, within couples, the unequal division of household chores between partners is not levelled-off at retirement. As discussed in Section 2, comparative analysis with other countries suggests the presence of a stronger gender-gap in Italy. This may persist when the men leaves work, although this may depend also on the opportunity cost of the wife's housework.

To provide some evidence, I first split the sample between those who live with a partner and those who do not. Among the former, I also distinguished between those who are married and the few cases in which they only cohabit. The change is very close to zero for married men, while it is large for those who are not living with a partner (413 minutes/week), although statistically significant only at the 10% level (p-value 0.069).[17] Those who are not married but cohabit show quite a large increase. One may speculate that less traditional families have a different distribution of household chores, but the number of observations is far too limited to draw any conclusion. Differently from men, married women living with their

[17] Among married men living with a partner, there are 14 who actually report to be *de facto* separated from their spouse, so that I can infer that they are cohabiting with a different person. Removing them has a very small effect on the estimates. This is similar for women, though there are only 3 cases.

partner show an increase (around 380 minutes/week), though this is smaller than for those not living with a partner (approximately 750 minutes/week, p-value 0.060).[18] Estimates for women living with their partner, but not married, are quite imprecise due to a very low first stage. Clearly, 2SLS estimates on these smaller samples are quite imprecise. Nevertheless, a statistical test for the equality of the effect for men who are married (and living with a partner) and for those not living with a partner rejects the null, although still only at the 10% level (p-value 0.094). In contrast, for women the test fails to reject the null, with a high p-value (0.443).

For married couples living together, Table 10 shows the effect of individual retirement not only on his/her own housework, but also on their partner's and on the gap between them. The first three columns (the "husbands' sample") select, from the overall sample, only married men living with their partners, while the last three columns do the analogous selection for wives. Both subsamples are further restricted to those couples where both partners have non-missing housework, and they also exclude cases where one of them is interviewed via proxy (but results are qualitatively similar if I keep cases where the partner has a proxy interview).

Results for the "husbands' sample" do not show an increase in their housework associated with own retirement (column 1), with no differences by their wives' working status.

If the wife does not work (middle panel), it seems that she decreases her engagement in housework (column 2) when the husband retires, resulting in a decrease in the housework gap (column 3). This could be explained by wives assigning a higher value to joint leisure with their partner, so that their opportunity cost of housework increases when he has more time available. It is more difficult to interpret this as a change in their bargaining power within the couple, because in this case we would expect a rebalancing of housework between the two partners, with the husband increasing his engagement. Furthermore, although the gender gap in housework narrows down, the drop in home production for the (non-employed) wife is smaller than the drop in market work for the husband (2449 minutes/week), so that there is an increase in their leisure gap. Results are very similar if we focus only on housewives or only on retired wives, so that it does not seem to depend on their past engagement in formal work.

For those cases in which the wife is in employment (lower panel), she does not seem to change her housework with the husband's retirement. Similar regressions do not show significant changes in their market work either. Given that in this subsample we are conditioning on the wives being employed (both for employed and retired men), this is not

[18] One explanation for this difference could be that, for married women who can rely on their spouses' income, there is a stronger selectivity out of the labour market with respect to their relative productivity in housework. This would imply that those who are more likely to engage in it are not included in the current analysis, as they never entered the labour market or they left it very early. Differently, this selection may be weaker across single women, so that their increase in housework is larger when their time constraint is relaxed at retirement. This mechanism can also partially explain the differences between women in richer household found in Section 7.1.

necessarily inconsistent with the previous findings that suggest an increase in the value of leisure when their partners leave work. The reason is that there may be joint retirement. A 2SLS regression, run on the overall husbands' sample (as in the upper panel) but with a "retired partner" dummy as dependent variable, shows that the retirement of the husbands leads to a 10.7 percentage points increase in wives likelihood to be retired.

The change in employed wives' housework associated with their husband's retirement may also depend on the opportunity cost in terms of her market wage. I further split the sample between wives with annual earnings (referred to 2006) below and above the median (for the selected sample of wives). Results (available on request) show that in the case of wives with higher earnings the retiring husband seems to increase his housework, although by much less than the average result for women. The (employed) wife decreases it, so that the gender gap narrows down, in line with an opportunity cost argument. Differently, when the wife has earnings below the median the retiring husband reduces his housework, while the wife increases it. This is more difficult to reconcile with the theory, as we would expect at most no change in their behaviour. Anyway, these further sample splits lead to very imprecise estimates, and therefore it is difficult to draw clear conclusions.

For women in married couples (the "wives' sample") there is an increase in their housework associated with own retirement (column 4), as discussed before, although the reduction in sample size makes it less significant. The husband seems to reduce his effort on home production (column 5), so that the gap in housework between partners increases by around 500 minutes/week (column 6). The differences between women with or without an employed partner are not sizeable, once we account for the lower precision associated with smaller sample sizes.[19] Similar regressions actually show that the employed husbands increase their market work by around 650 minutes/week with their wives' retirement (results available on request). The estimates from a regression with a "retired partner" dummy as dependent variable suggest no effect of women's retirement on the likelihood that their husband is retired.[20] The increase in available time for the retiring wife seems, therefore, to be partly used to decrease the husband's housework, possibly in favour of market work.

To summarize, results are in line with Italian husbands having enough bargaining power to avoid a significant engagement in housework. Indeed, singles tend to increase their housework when they retire, while married men do not, at least on average. When their husband retires, wives tend to reduce the engagement in housework if they are not employed (either housewives or in other condition), or to retire if they are employed. This seems to be explained by an increased marginal value of (joint) leisure, rather than by a shift in bargaining

[19]In this case, it is not possible to separately analyse employed male partners' with earnings above and below the median, because the sample size becomes even smaller, leading to low first-stage F statistics and to a weak instrument problem.

[20]This is consistent with recent findings from Hospido and Zamarro (2014), who found a positive joint retirement effect on women's retirement but no effect on men's, using a sample of individuals from different European countries (SHARE).

power. On the opposite, previously working women increase their housework when they retire, while their husbands tend to reduce their own engagement in home production and to increase, if employed, their market work.

8 Discussion and conclusions

I used an RDD that exploits the discontinuity in retirement behaviour induced by the Italian Social Security System. Although the proportion of men leaving employment at eligibility is quite large, the strong discontinuity in retirement is not associated with a jump in time spent on home production. Conversely, for women I observe an increase in both retirement and housework at eligibility. The resulting estimate for the causal effect of retirement on housework is between 430 and 490 minutes per week (nearly an hour per day), depending on the introduction of covariates and on whether or not we interact S with D.

The strong gender difference found for Italy seems to have no parallel in the US, France, Germany or Spain. Hurd and Rohwedder (2005, 2006), using data from the Health and Retirement Study, showed that women who retired between 2001 and 2003 increased by 309 minutes per week the time spent in activities with close market substitutes. However, they found a sensible increase for men as well, of around 361 minutes/week. The gerontology literature provides similar evidence (Szinovacz and Harpster, 1994). Szinovacz (2000), using US panel data, found that husbands' increase their relative contribution not only in "male tasks (outdoor tasks, repairs, paying bills)", but also in "female tasks (preparing meals, doing the dishes, cleaning house, laundry)" (Szinovacz, 2000, p. 82). For France, Stancanelli and van Soest (2012) estimated that at retirement wives spend 2 hours 40 minutes per weekday more on housework, but they found that husbands also increased housework by around 3 hours per weekday. Furthermore, there is evidence for Germany (Schwerdt, 2005; Luhrmann, 2010; Bonsang and van Soest, 2015) of an increase in housework at the retirement of the household heads, who are mostly men. Bonsang and van Soest (2015) study retirement in couples using panel data from German SOEP. They find that both partners increase home production upon own retirement, but decrease it when the other retires as well. Lastly, Luengo-Prado and Sevilla (2013) provided evidence that in Spain the retirement of the household head causes a reallocation of household duties, with men increasing their involvement in shopping and cooking. They also suggested that this equalizing effect is the result of a move towards more egalitarian social norms.

One explanation for the different result in Italy is that, after retirement, men mostly focus on "semi-leisure activities", such as gardening, which are not included in the SILC definition of home production. This argument is consistent with the descriptive comparison with TUS and with results from Stancanelli and van Soest (2012), who showed that (in France) the increase for men is concentrated in these activities. Furthermore, it must be noted that some weak evidence of an increase is found for men residing in densely populated areas, who are probably less likely to specialise in these “semi-leisure activities”.

Another explanation is that married men have a strong bargaining power and leave most of the housework to their wives, even after retirement. Indeed, when I focus on this group the estimate is very small, while it is around 400 minutes/week for those living without a partner, even if statistically significant only at the 10% level. Differently, for women the estimate is positive and statistically significant both for singles and for those who are married, in line with the drop in the opportunity cost.

Overall, the results suggest that retirement does not lead to a more egalitarian distribution of housework between genders, at least if we focus on “core” household chores. Social norms and/or differences in bargaining power that may explain the strong differences found in Italy seem, therefore, quite persistent to the individual transitions out of paid employment.

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Tables and Figures

Table 1: Pension requirements

	Men				Women			
	Seniority pension: age (with 35 years of contributions)[a] or contributions alone			Old-age: age plus minimum contributions	Seniority pension: age (with 35 years of contributions)[a] or contributions alone			Old-age: age plus minimum contributions
	Private	Public	Self-employed	[b]	Private	Public	Self-employed	[b]
1997	52 or 36	52 or 36	56 or 40	63	52 or 36	52 or 36	56 or 40	58
1998	54 or 36	53 or 36	57 or 40	63	54 or 36	53 or 36	57 or 40	58
1999	55 or 37	53 or 37	57 or 40	64	55 or 37	53 or 37	57 or 40	59
2000	55 or 37	54 or 37	57 or 40	65	55 or 37	54 or 37	57 or 40	60
2001	56 or 37	55 or 37	58 or 40	65	56 or 37	55 or 37	58 or 40	60
2002	57 or 37	55 or 37	58 or 40	65	57 or 37	55 or 37	58 or 40	60
2003	57 or 37	56 or 37	58 or 40	65	57 or 37	56 or 37	58 or 40	60
2004	57 or 38	57 or 38	58 or 40	65	57 or 38	57 or 38	58 or 40	60
2005	57 or 38	57 or 38	58 or 40	65	57 or 38	57 or 38	58 or 40	60
2006	57 or 39	57 or 39	58 or 40	65	57 or 39	57 or 39	58 or 40	60
2007	57 or 39	57 or 39	58 or 40	65	57 or 39	57 or 39	58 or 40	60

Note: The table combines data from Brugiavini and Peracchi (2004); Morciano (2007); Battistin et al. (2009); Intorcia (2011). Requirements for the years before 1997 are not reported in the Table, but are available on request. [a]There were exceptions for blue-collars and for those who started their career (and paid some contributions) before 1992 (these have been taken into account in calculating S). [b]For the self-employed the minimum age for the old-age path was always 65 for men and 60 for women. For the public sector it depended on the specific categories, but the general rule for the workers of the central administration was 65 for men and 60 for women (I apply this rule in defining S, as I cannot identify each single category). The old-age path required between 15 and 20 years of contributions for those who started working before 1996 (depending on year of retirement), but only 5 years for those who started later.

Table 2: Average minutes per day spent in housework, by gender and self-defined employment status, SILC 2007 and TUS 2008-2009, individuals aged [50,70].

		SILC	TUS (A)	TUS (B)	TUS (C)
Men	Employed	54	79	58	54
	Retired	89	202	132	117
Women	Employed	213	242	231	223
	Retired	282	378	350	327

Note: estimated on original microdata using sample weights. In SILC I excluded missing values in housework. TUS (A) refers to total “family related” work, while column (B) contains only shopping, cooking, caring and “core” household work. Column (C) further excludes caregiving to non-coresident individuals. To calculate average minutes per day in SILC, I divided Y by 7. TUS data refer to an average day calculated from averaging diaries collected in different days of the week (one diary per respondent). All estimates and figures are obtained using Stata™12, plus programs `ivreg2` (Baum et al., 2007) and `esttab` (Jann, 2007).

Table 3: The association of housework and market work with employment status and other characteristics. Predicted values from OLS regressions, minutes per week, SILC 2007 (full sample)

	House work		Market work		Total work	
	Men	Women	Men	Women	Men	Women
			Employment status			
Worker	319	1155	2515	2098	2834	3254
Unemployed	421	1688	467	442	888	2130
Housewife		2136		24		2160
Student	330	1016	99	3	429	1019
Retired	566	1944	25[(a)]	13[(a)]	592	1957
Other	444	1687	213	70	658	1757
			Area			
North	426	1585	1466	757	1892	2342
Centre	368	1569	1405	738	1773	2307
South	377	1686	1363	697	1739	2382
			Population density			
Densely populated area	403	1491	1400	721	1803	2212
Intermediate area	397	1637	1433	743	1830	2379
Thinly populated area	389	1754	1424	736	1813	2490
			Education			
College	386	1382	1370	669	1757	2051
High school	409	1655	1429	751	1837	2316
Middle school	402	1680	1429	740	1831	2420
Primary school	375	1685	1413	729	1788	2414
			Marital status			
Live with partner, married	388	1873	1453	721	1842	2593
Live with partner, not married	483	1684	1496	756	1979	2441
Not living with a partner	404	1288	1360	748	1764	2036

Note: SILC 2007, original sample (excluding missing values in time spent in domestic work). The predictions are based on OLS regressions including a full set of dummies for each category, plus controls for age, age^2, household size, dummy for poor health, dummy for missing health information. Regressions have been run separately by gender. Each cell shows the predicted value obtained by fixing the selected variable at the indicated level and averaging over the distribution of the other covariates. The definitions of retirees and workers are described in Section 4.1. [(a)]Although retirees do not do any market work by construction, the linear model does not precisely fit this constraint and therefore the average value from other covariates still predicts a small positive time spent on housework.

Table 4: First stage OLS regressions for retirement status, SILC 2007 (selected sample)

Dependent variable R	(1) Men	(2) Men	(3) Men	(4) Women	(5) Women	(6) Women
D	0.270***	0.313***	0.398***	0.110	0.236***	0.470***
	(0.082)	(0.048)	(0.027)	(0.097)	(0.059)	(0.034)
$(1-D)\times S$	-0.004	0.021***	0.011***	0.081***	0.034***	0.008***
	(0.024)	(0.007)	(0.001)	(0.028)	(0.009)	(0.002)
$(1-D)\times S^2$	-0.004	0.001		0.012**	0.002***	
	(0.005)	(0.001)		(0.005)	(0.001)	
$(1-D)\times S^3$	-0.000			0.001**		
	(0.000)			(0.000)		
$D\times S$	0.138***	0.075***	0.043***	0.201***	0.137***	0.048***
	(0.052)	(0.017)	(0.004)	(0.059)	(0.019)	(0.004)
$D\times S^2$	-0.016	-0.003**		-0.022*	-0.008***	
	(0.010)	(0.001)		(0.011)	(0.002)	
$D\times S^3$	0.001			0.001		
	(0.001)			(0.001)		
Constant	0.089**	0.117***	0.097***	0.188***	0.134***	0.078***
	(0.036)	(0.021)	(0.012)	(0.046)	(0.028)	(0.014)
Observations	3970	3970	3970	2701	2701	2701
R^2	0.570	0.569	0.568	0.704	0.703	0.696
$H_0:\gamma_D=0$ (p-value)	0.001	0.000	0.000	0.259	0.000	0.000
- (p-val cluster)	0.004	0.001	0.000	0.112	0.002	0.000
Dong's $\widehat{\gamma_D}$	0.197	0.285	0.382	0.044	0.183	0.450
Dong's $\widehat{\gamma_D}$ (p-value)	0.055	0.000	0.000	0.712	0.005	0.000
G (p-value)	0.005	0.003	0.001	0.186	0.078	0.000
AIC	1949.554	1949.817	1954.425	378.492	380.463	439.276
BIC	1999.846	1987.536	1979.571	425.703	415.872	462.881

Note: * p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. R is a dummy for retirees, D is a dummy for eligible, S is distance to/from eligibility, γ_D is the coefficient for the discontinuity at eligibility. G (p-value) is Lee and Card (2008) statistic. Dong's refer to Dong (2014) correction, while the p-value cluster is calculated by clustering on S. AIC is the Akaike criterion; BIC is the Bayesian criterion.

Table 5: Regressions for time spent on house work (in minutes per week), SILC 2007 (selected sample)

Dependent variable Y	(1) Men	(2) Men	(3) Women	(4) Women
		Reduced form OLS regressions		
D	21.3	30.9	205.2**	218.6**
	(38.2)	(37.1)	(89.3)	(87.0)
$(1-D)\times S$	3.9		8.7	
	(3.4)		(8.3)	
$D\times S$	17.1***		22.3*	
	(5.5)		(11.5)	
S		9.4***		14.2**
		(3.0)		(6.8)
Constant	417.3***	449.3***	1621.6***	1654.2***
	(22.6)	(20.5)	(54.5)	(46.6)
R^2	0.017	0.016	0.035	0.035
$H_0:\beta_D=0$ (p-value)	0.578	0.404	0.022	0.012
- (p-value cluster)	0.547	0.463	0.016	0.017
Dong's $\widehat{\beta_D}$	14.7		198.4	
Dong's $\widehat{\beta_D}$ (p-value)	0.708		0.030	
G (p-value)	0.200	0.101	0.796	0.789
AIC	61613.328	61615.615	45063.107	45062.065
BIC	61638.474	61634.475	45086.712	45079.769
		2SLS regressions		
R	53.4	73.2	436.7**	429.1***
	(95.1)	(86.9)	(184.4)	(166.1)
$(1-D)\times S$	3.3		5.2	
	(4.0)		(9.1)	
$D\times S$	14.7*		1.3	
	(8.8)		(18.4)	
S		7.6		3.8
		(4.9)		(10.4)
Constant	412.1***	436.4***	1587.5***	1579.4***
	(28.9)	(34.6)	(63.7)	(72.8)
$H_0:\delta=0$ (p-value)	0.574	0.399	0.018	0.010
- (p-value cluster)	0.505	0.384	0.001	0.000
First Stage F	216.080	265.562	196.336	253.943
OLS est. for δ	285.7***	287.3***	655.6***	637.0***
	(28.6)	(27.5)	(72.0)	(68.6)
Observations	3970	3970	2701	2701

Note: * p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. Y is time spent on housework. β_D is the coefficient for the discontinuity at eligibility in the upper panel, while δ is the coefficient on R in the lower panel. In 2SLS, R is instrumented by D. See Table 4 for other definitions.

Table 6: 2SLS regressions for time spent on house work (in minutes per week), including covariates, SILC 2007 (selected sample)

Dependent variable Y	(1) Men	(2) Men	(3) Women	(4) Women
R	79.9	100.4	493.1***	466.0***
	(93.2)	(85.5)	(176.9)	(160.1)
$(1-D)\times S$	2.2		-0.3	
	(4.0)		(8.8)	
$D\times S$	14.2*		-14.4	
	(8.5)		(17.5)	
S		6.7		-5.3
		(4.8)		(9.8)
Constant	427.0***	451.5***	1549.8***	1519.2***
	(41.5)	(47.4)	(85.5)	(93.1)
Observations	3970	3970	2701	2701
$H_0 : \delta = 0$ (p-value)	0.391	0.240	0.005	0.004
- (p-val cluster)	0.286	0.232	0.000	0.000
First Stage F	229.085	277.760	206.277	264.778
OLS est. for δ	274.4***	277.8***	611.8***	588.4***
	(29.1)	(28.1)	(70.8)	(67.8)

Note: * p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. R is instrumented by D. δ is the coefficient on R. See Tables 4 and 5 for other definitions. The p-val cluster is calculated by clustering on S. Covariates include dummies for geographic area, population density, education and employment category. Coefficients are available on request.

Table 7: Regressions for different socio-economic variables, SILC 2007 (selected sample)

Dependent variable	Men			Women		
	$\widehat{\gamma}_D$	p-value	G (p-value)	$\widehat{\gamma}_D$	p-value	G (p-value)
Geographical dummies						
North	-0.046	0.196	0.249	-0.039	0.374	0.962
Centre	0.059*	0.052	0.508	0.016	0.686	0.060
South	-0.014	0.656	0.042	0.023	0.537	0.347
Densely populated area	-0.030	0.368	0.680	0.083*	0.053	0.340
Intermediate area	0.026	0.458	0.048	-0.065	0.135	0.805
Thinly populated area	0.004	0.881	0.229	-0.018	0.619	0.297
Test for joint significance		0.368			0.263	
Educational dummies						
College	-0.015	0.522	0.627	0.013	0.668	0.895
High school	0.084**	0.013	0.051	0.004	0.925	0.360
Middle school	-0.065**	0.041	0.048	0.007	0.848	0.961
Primary school	-0.004	0.889	0.460	-0.024	0.527	0.366
Test for joint significance		0.064			0.922	
Additional educational variables						
Years of schooling	0.282	0.318	0.769	0.479	0.196	0.801
Age highest degree - 6	-1.180	0.102	0.318	0.722	0.376	0.007
Test for joint significance		0.046			0.411	
Employment categories						
Private employee	-0.013	0.704	0.020	-0.069	0.104	0.543
Public employee	-0.055**	0.044	0.469	0.019	0.656	0.189
Self-employed	0.068**	0.031	0.005	0.050	0.195	0.180
Test for joint significance		0.035			0.216	

* p<.10 ** p<.05 *** p<.01. The regressions include $(1-D)\times S$, $D\times S$ and a constant. γ_D is the coefficient for the discontinuity at eligibility. See Tables 4 and 5 for other definitions. The null hypothesis for the test for joint significance is that there is no discontinuity in all variables of each group, and it is run by using Stata command `suest` with robust standard errors. In the case of mutually exclusive dummies (for instance North-Centre-South), one constraint is removed, but the result does not depend on which one is chosen.

Table 8: Sensitivity analysis

Dependent variable Y	(1) S calculated without adding one to age at retirement	(2) Regressions including $S=0$	(3) Definition of S based on *first* age of eligibility	(4) Removing individuals with non-consistent info on pension income	(5) Using self-defined definition of retirement
			Men		
R	50.6	74.1	22.6	124.1	105.7
	(93.9)	(108.4)	(151.8)	(77.5)	(88.5)
Obs	3864	4139	2988	3614	4105
$H_0:\delta=0$ (p-value)	0.590	0.495	0.882	0.109	0.232
			Women		
R	404.5**	457.6**	636.3**	365.4**	371.0**
	(168.3)	(196.7)	(267.3)	(150.0)	(180.2)
Obs	2638	2795	2115	2488	2779
$H_0:\delta=0$ (p-value)	0.016	0.020	0.017	0.015	0.039

Note: * p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. R is instrumented by D. The regression includes S and a constant as additional covariates. See Tables 4 and 5 for other definitions.

Table 9: 2SLS estimates splitting the sample by education, area, employment category and marital status, SILC 2007 (selected sample)

		Men				Women			
		$\widehat{\delta}$	p-value	First stage F	obs	$\widehat{\delta}$	p-value	First stage F	obs
By education:	Middle school or less	72.2	0.579	138	2020	485.7	0.041	141	1354
	High school	28.1	0.830	101	1492	551.1	0.024	115	1006
	College	176.2	0.448	30	458	-385.0	0.526	15	341
By area:	North	148.4	0.176	157	1953	551.4	0.009	146	1369
	Centre	86.2	0.619	61	966	357.5	0.286	61	718
	South	-105.1	0.635	50	1051	204.0	0.598	51	614
By degree of urbanization:	Densely populated	225.2	0.067	108	1360	634.4	0.003	96	985
	Intermediate area	-28.4	0.855	97	1689	677.5	0.015	107	1110
	Thinly populated	-24.0	0.896	63	921	-197.7	0.608	53	606
By category:	Private employee	112.8	0.222	268	2080	482.4	0.028	207	1073
	Public employee	104.9	0.690	34	761	325.4	0.269	75	981
	Self-employed	-21.3	0.934	25	1129	569.4	0.182	33	647
By household equivalent income:	Above median	100.8	0.406	120	1985	605.7	0.006	130	1350
	Below median	77.9	0.511	177	1985	228.8	0.362	137	1351
By marital status:	Living with partner, married	0.9	0.992	211	3132	384.8	0.032	237	1828
	Living with partner, not married	679.2	0.077	11	105	-346.5	0.840	2	55
	Not living with partner	413.4	0.069	46	733	750.8	0.060	37	818

Note: all estimates include only a constant and S, while R is instrumented by D. δ is the coefficient on R (i.e. the effect of retirement on housework). The p-value and the first stage F are calculated using robust standard errors.

Table 10: The effect of retirement on the individual and on his/her partner's housework

	(1)	(2)	(3)	(4)	(5)	(6)
		Husbands' sample			Wives' sample	
		Dependent variable:			Dependent variable:	
	Y	Partner's Y	Y-Partner's Y	Y	Partner's Y	Y-Partner's Y
				All couples		
R	-0.3	-216.8	216.5	354.1*	-193.3	547.4**
	(97.7)	(233.6)	(243.0)	(201.5)	(143.6)	(233.6)
obs	2747	2747	2747	1489	1489	1489
				Non-employed partner		
R	-4.5	-443.6	439.2	250.0	-225.8	475.9*
	(129.4)	(327.0)	(333.3)	(238.8)	(189.8)	(285.8)
obs	1641	1641	1641	793	793	793
				Employed partner		
R	16.4	-7.6	23.9	364.9	-160.1	524.9
	(148.3)	(257.0)	(273.7)	(538.0)	(229.6)	(577.8)
obs	1106	1106	1106	696	696	696

Note: * p<.10 ** p<.05 *** p<.01. The individual under analysis (husband in columns (1)-(3) and wife in (4)-(6)) is either worker or retiree with $S \in [-10, 10]$. The main explanatory variable (R) refers to his/her retirement. The sample includes only individuals living together with their partner, married, where partners have no proxy interviews or missing values of Y. The results show estimates from 2SLS where R is instrumented by D and S plus a constant are included as covariates. Robust standard errors in brackets. The first-stage F-statistic, not shown in the table, is always above 15.

Figures

Figure 1: Average minutes/week of housework by employment status (circles for workers, triangles for retirees) and age (in years), SILC 2007, only age-employment cells with at least 20 obs. Lines are fit from a 2nd order polynomial (with 95% confidence interval)

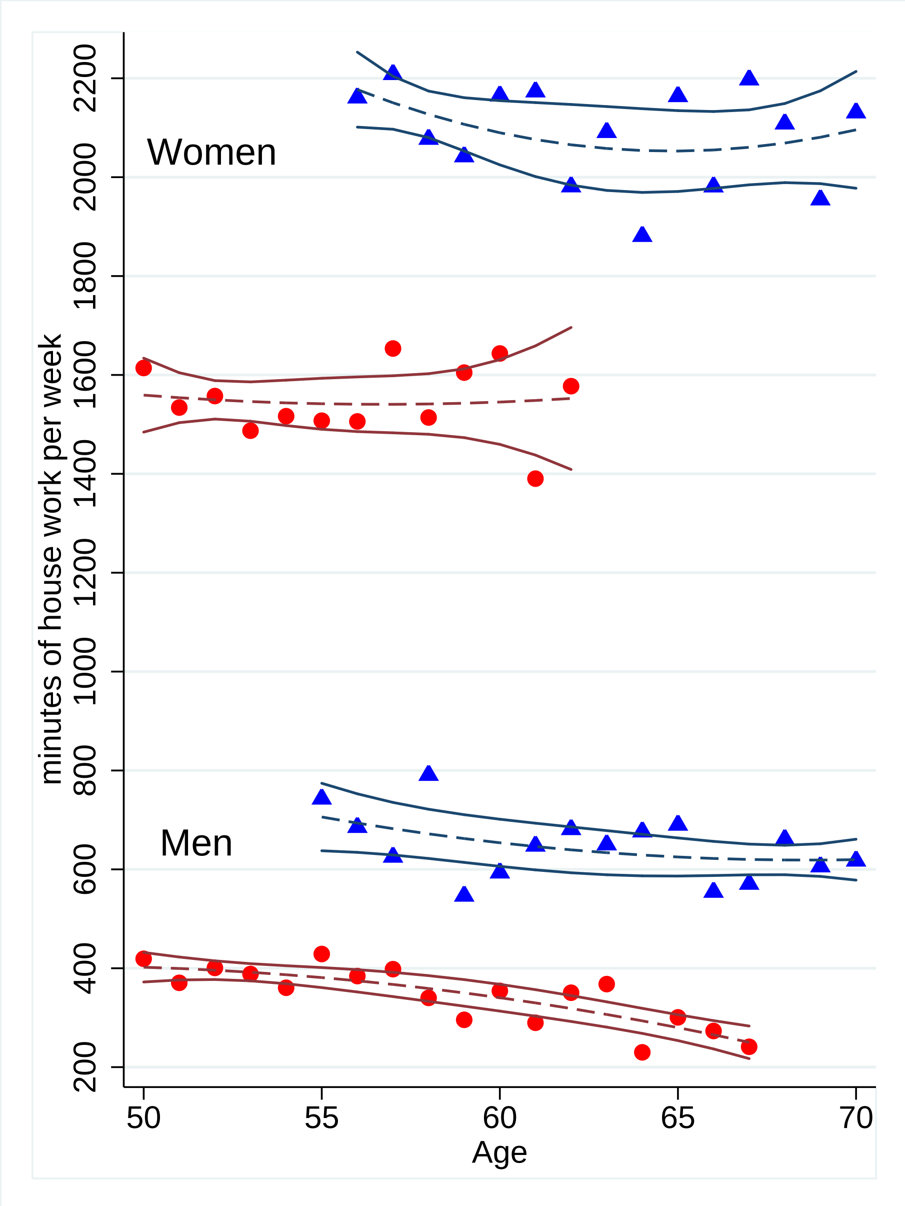
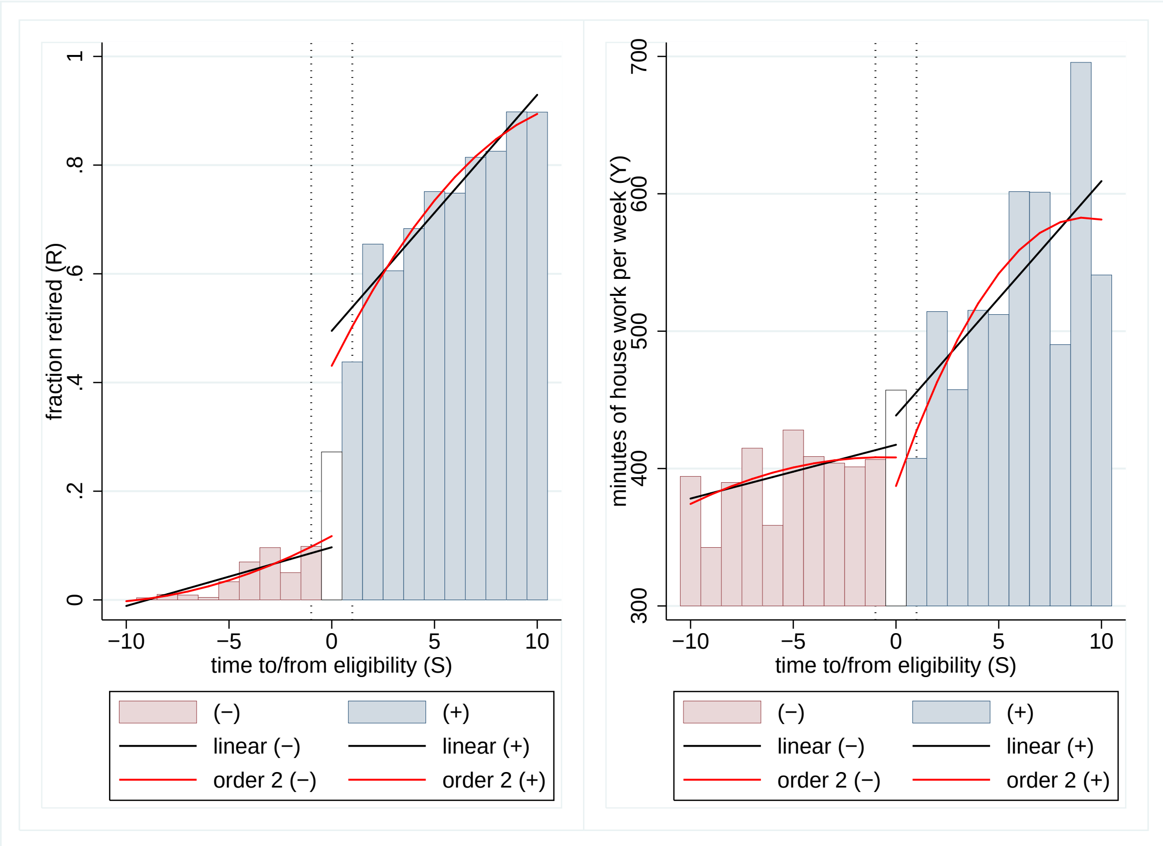
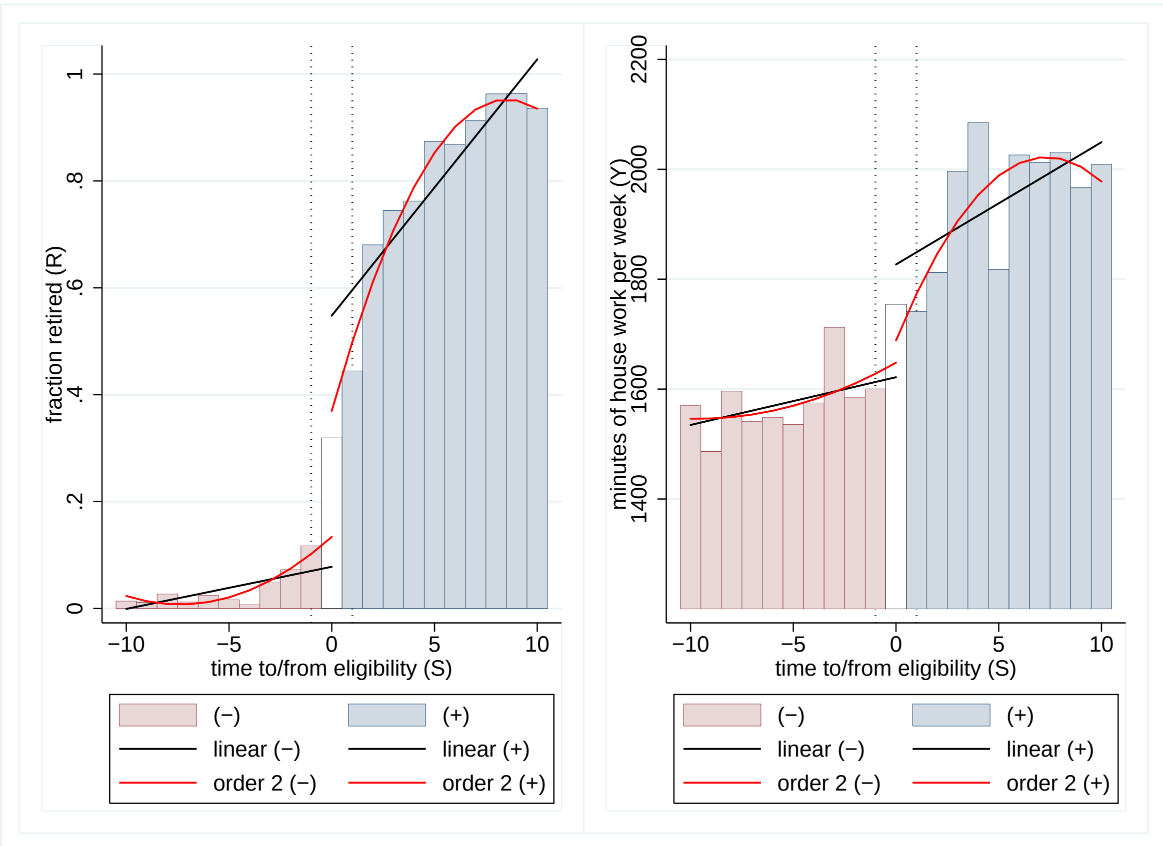


Figure 2: Retirement and housework with respect to S, SILC 2007 (selected sample). Lines are fit from polynomials allowing for different slopes at the right and left of eligibility, estimated excluding $S = 0$.



(a) Men



(b) Women

Appendices to “Retirement, pension eligibility and home production”

Appendix A: additional data description and robustness checks

Additional info on time to/from eligibility

Age at retirement is calculated as the age in which the respondent began the first regular job plus the number of years spent in paid job, plus one.[1] I added one year because it seems that respondents do not report the last year of work if it consisted of less than 12 months. To understand this step, define T as the difference between current age and age at retirement. Among retired, I expect to observe almost nobody with $T < 0$, then a positive jump in the frequency at $T = 0$, and a gradual decrease toward zero for larger T. However, if I do not add one year, there are very few retired with $T = 0$, and the discontinuity is at $T = 1$, implying that almost everybody retired at least one year before. Furthermore, if I build T for workers as well, I would expect the mode of the distribution to be $T = 0$. However, if I do not add one the mode of the distribution is at $T = 1$, with a frequency of only 1.21% at $T = 0$.

To distinguish self-employed and employees, I exploited information on the current job for workers and on the last job for retired. For some of the years before 2004, rules were somewhat more favourable to employees in the public sector. However, I have this information only for those currently working. Therefore I use the Statistical Classification of Economic Activities in the European Union (NACE code) for both workers and retired.[2] I define as employees in the public sector those working in “Public administration and defence, compulsory and social security”, “Education” or “Health and social work”. Among workers in 2007, only the 18.91% of those belonging to these three groups report to work for the private sector, and together these three categories accounted for the 84.2% of total public sector employees. One might argue that, given the availability of the public/private information for those currently working, I should use the NACE code only for retired individuals. Given that in 2007 rules for employees are independent from the sector of activity, it would make no difference.

[1] I also corrected the age at retirement to be equal to the current age for 0.30% (29 obs) of the retired for whom the first was larger than the second.

[2] See http://epp.eurostat.ec.europa.eu/ for details (last access: 12/07/2012).

Sample selection and descriptive statistics

Table A1: Sample selection

	Male		Female	
	obs	% change	obs	% change
Raw 2007 SILC data	21,522		23,611	
Worker or Retired	16,958	-21%	12,162	-48%
Non Proxy	13,979	-18%	10,856	-11%
Missing housework	13,437	-4%	10,546	-3%
$S \in [-10, 10]$	4,139	-69%	2,795	-73%
$S \neq 0$	3,970	-4%	2,701	-3%

Table A2: Descriptive statistics of covariates, Men, SILC 2007 (selected sample)

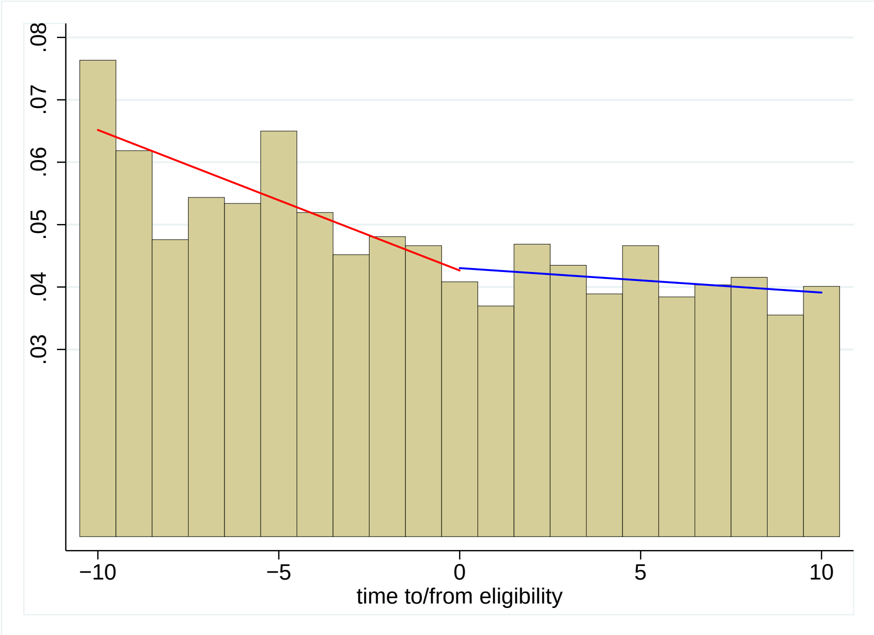
	mean	median	sd	min	max	count
Y	452.6914	300	571.778	0	5400	3970
Bad health	.0639798	0	.2447479	0	1	3970
Disabled	.0403023	0	.196692	0	1	3970
Missing health	.0244332	0	.1544094	0	1	3970
hsize	2.950126	3	1.191653	1	7	3970
S	-1.039295	-2	6.266241	-10	10	3970
Centre	.2433249	0	.4291437	0	1	3970
South	.2647355	0	.4412479	0	1	3970
intermediate area	.4254408	0	.494472	0	1	3970
thinly populated area	.2319899	0	.4221558	0	1	3970
Married	.8183879	1	.3855731	0	1	3970
Separated	.0277078	0	.164155	0	1	3970
Widowed	.0244332	0	.1544094	0	1	3970
Divorced	.0302267	0	.1712321	0	1	3970
College	.1153652	0	.3195024	0	1	3970
High school	.3758186	0	.4843946	0	1	3970
Middle school	.2979849	0	.4574304	0	1	3970
age	56.89421	57	6.128908	44	75	3970
ycontrib	32.5602	33	5.74599	5	50	3970
age first job	18.5204	17	4.959113	8	50	3970
years paid job	32.92393	33.5	7.500483	4	60	3970
employee public	.1916877	0	.3936782	0	1	3970
self-employed	.2843829	0	.4511768	0	1	3970
Observations	3970					

Table A3: Descriptive statistics of covariates, Women, SILC 2007 (selected sample)

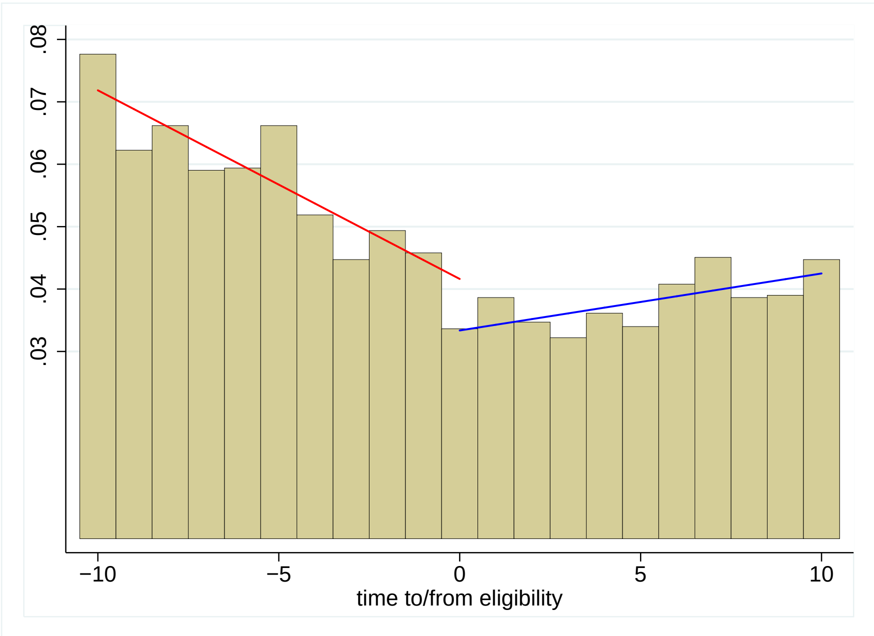
	mean	median	sd	min	max	count
Y	1722.445	1500	1032.236	0	6000	2701
Bad health	.0629397	0	.2428994	0	1	2701
Disabled	.0362829	0	.1870277	0	1	2701
Missing health	.028508	0	.1664497	0	1	2701
hsize	2.732692	3	1.135456	1	9	2701
S	-1.310255	-3	6.371929	-10	10	2701
Centre	.2658275	0	.4418546	0	1	2701
South	.2273232	0	.4191807	0	1	2701
intermediate area	.4109589	0	.4920989	0	1	2701
thinly populated area	.2243613	0	.4172383	0	1	2701
Married	.7089967	1	.4543092	0	1	2701
Separated	.0288782	0	.1674952	0	1	2701
Widowed	.1029248	0	.303917	0	1	2701
Divorced	.0529434	0	.2239619	0	1	2701
College	.1262495	0	.3321919	0	1	2701
High school	.3724546	0	.4835481	0	1	2701
Middle school	.2473158	0	.4315317	0	1	2701
age	56.49833	56	6.036593	43	79	2701
ycontrib	28.31988	29	7.377383	2	46	2701
age first job	20.04739	19	6.101728	8	50	2701
years paid job	28.82488	30	8.336003	2	60	2701
employee public	.3631988	0	.4810105	0	1	2701
self-employed	.2395409	0	.4268823	0	1	2701
Observations	2701					

Density plots

Figure A1: Density, SILC 2007 (selected sample)



(a) Men



(b) Women

Regression tables for Y including also a quadratic polynomial

Table A4: Reduced form OLS regressions for time spent on house work (in minutes per week), SILC 2007 (selected sample)

Dependent variable Y	(1) Men	(2) Men	(3) Men	(4) Women	(5) Women	(6) Women
D	-20.8	21.3	30.9	40.4	205.2**	218.6**
	(64.6)	(38.2)	(37.1)	(149.5)	(89.3)	(87.0)
$(1-D)\times S$	-0.5	3.9		21.2	8.7	
	(15.7)	(3.4)		(36.3)	(8.3)	
$(1-D)\times S^2$	-0.4			1.1		
	(1.3)			(3.1)		
$D\times S$	42.5*	17.1***		91.1*	22.3*	
	(23.5)	(5.5)		(50.8)	(11.5)	
$D\times S^2$	-2.3			-6.2		
	(2.2)			(4.6)		
S			9.4***			14.2**
			(3.0)			(6.8)
Constant	408.1***	417.3***	449.3***	1648.1***	1621.6***	1654.2***
	(39.0)	(22.6)	(20.5)	(92.0)	(54.5)	(46.6)
Observations	3970	3970	3970	2701	2701	2701
R^2	0.018	0.017	0.016	0.036	0.035	0.035
$H_0:\beta_D=0$ (p-value)	0.747	0.578	0.404	0.787	0.022	0.012
- (p-value cluster)	0.661	0.547	0.463	0.530	0.016	0.017
Dong's $\widehat{\beta_D}$	-42.6	14.7		4.2	198.4	
Dong's $\widehat{\beta_D}$ (p-value)	0.546	0.708		0.979	0.030	
G (p-value)	0.166	0.200	0.101	0.837	0.796	0.789
AIC	61615.827	61613.328	61615.615	45064.779	45063.107	45062.065
BIC	61653.546	61638.474	61634.475	45100.188	45086.712	45079.769

* p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. Y is time spent on housework. See Table **??** for other definitions. β_D is the coefficient for the discontinuity at eligibility. G (p-value) is Lee and Card (2008) statistic. Dong's refer to Dong (2014b) correction, while the p-value cluster is calculated by clustering on S. AIC is the Akaike criterion; BIC is the Bayesian criterion.

Table A5: 2SLS regressions for time spent on house work (in minutes per week), SILC 2007 (selected sample)

Dependent variable Y	(1) Men	(2) Men	(3) Men	(4) Women	(5) Women	(6) Women
R	-66.4	53.4	73.2	171.4	436.7**	429.1***
	(208.6)	(95.1)	(86.9)	(622.5)	(184.4)	(166.1)
$(1-D)\times S$	0.9	3.3		15.3	5.2	
	(18.4)	(4.0)		(51.1)	(9.1)	
$(1-D)\times S^2$	-0.3			0.7		
	(1.4)			(4.0)		
$D\times S$	47.5	14.7*		67.7	1.3	
	(36.7)	(8.8)		(125.4)	(18.4)	
$D\times S^2$	-2.5			-4.8		
	(2.6)			(8.6)		
S			7.6			3.8
			(4.9)			(10.4)
Constant	415.9***	412.1***	436.4***	1625.1***	1587.5***	1579.4***
	(57.3)	(28.9)	(34.6)	(157.0)	(63.7)	(72.8)
Observations	3970	3970	3970	2701	2701	2701
$H_0: \delta = 0$ (p-value)	0.750	0.574	0.399	0.783	0.018	0.010
- (p-value cluster)	0.672	0.505	0.384	0.506	0.001	0.000
First Stage F	42.196	216.080	265.562	15.950	196.336	253.943
OLS est. for δ	298.2***	285.7***	287.3***	680.9***	655.6***	637.0***
	(29.4)	(28.6)	(27.5)	(75.2)	(72.0)	(68.6)

* $p<.10$ ** $p<.05$ *** $p<.01$. Robust standard errors in brackets. R is instrumented by D. See Tables **??** and A4 for other definitions. δ is the coefficient on R. The p-value cluster is calculated by clustering on S.

Table A6: 2SLS regressions for time spent on house work (in minutes per week), including covariates, SILC 2007 (selected sample)

Dependent variable Y	(1) Men	(2) Men	(3) Men	(4) Women	(5) Women	(6) Women
R	-66.9	79.9	100.4	461.4	493.1***	466.0***
	(204.0)	(93.2)	(85.5)	(573.9)	(176.9)	(160.1)
$(1-D)\times S$	2.5	2.2		-6.9	-0.3	
	(18.1)	(4.0)		(47.7)	(8.8)	
$(1-D)\times S^2$	-0.1			-0.6		
	(1.4)			(3.8)		
$D\times S$	50.7	14.2*		1.8	-14.4	
	(35.9)	(8.5)		(115.4)	(17.5)	
$D\times S^2$	-2.8			-1.3		
	(2.6)			(8.0)		
S			6.7			-5.3
			(4.8)			(9.8)
Constant	451.0***	427.0***	451.5***	1539.6***	1549.8***	1519.2***
	(74.6)	(41.5)	(47.4)	(182.6)	(85.5)	(93.1)
Observations	3970	3970	3970	2701	2701	2701
$H_0: \delta = 0$ (p-value)	0.743	0.391	0.240	0.421	0.005	0.004
- (p-val cluster)	0.641	0.286	0.232	0.105	0.000	0.000
First Stage F	43.919	229.085	277.760	17.311	206.277	264.778
OLS est. for δ	284.5***	274.4***	277.8***	629.4***	611.8***	588.4***
	(30.0)	(29.1)	(28.1)	(75.0)	(70.8)	(67.8)

* $p<.10$ ** $p<.05$ *** $p<.01$. Robust standard errors in brackets. R is instrumented by D. δ is the coefficient on R. See Tables **??** and A4 for other definitions. The p-val cluster is calculated by clustering on S. Covariates include dummies for geographic area, population density, education and employment category. Coefficients are available on request.

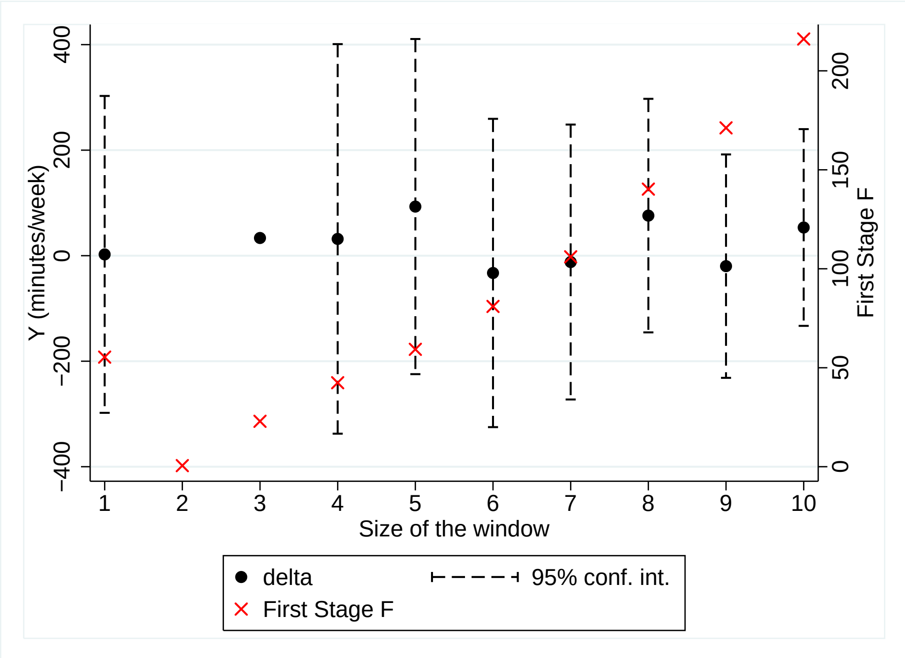
Table A7: 2SLS estimates splitting the sample by education, area, employment category and marital status, including both $(1-D)\times S$ and $D\times S$, SILC 2007 (selected sample)

		Men				Women			
		$\widehat{\delta}$	p-value	First stage F	obs	$\widehat{\delta}$	p-value	First stage F	obs
By education:	Middle school or less	70.0	0.602	132	2020	497.0	0.040	139	1354
	High school	-40.1	0.799	68	1492	599.1	0.053	62	1006
	College	130.9	0.596	20	458	-584.3	0.545	5	341
By area:	North	136.1	0.256	129	1953	587.5	0.013	112	1369
	Centre	64.4	0.728	54	966	339.5	0.386	45	718
	South	-139.4	0.570	38	1051	154.5	0.708	41	614
By degree of urbanization:	Densely populated	185.5	0.169	84	1360	654.1	0.004	86	985
	Intermediate area	-44.8	0.786	86	1689	755.5	0.016	78	1110
	Thinly populated	-17.0	0.934	46	921	-351.6	0.443	38	606
By category:	Private employee	107.2	0.267	249	2080	508.3	0.029	177	1073
	Public employee	96.3	0.780	17	761	327.4	0.394	40	981
	Self-employed	-84.9	0.770	19	1129	597.9	0.189	28	647
By household equivalent income:	Above median	85.8	0.520	93	1985	611.2	0.015	86	1350
	Below median	65.8	0.605	152	1985	258.1	0.330	120	1351
By marital status:	Married, living with partner	-21.0	0.837	177	3132	342.4	0.084	184	1828
	Not married, living with partner	1134.4	0.054	4	105	-2655.0	0.715	0.3	55
	Not living with partner	403.7	0.098	39	733	881.3	0.048	29	818

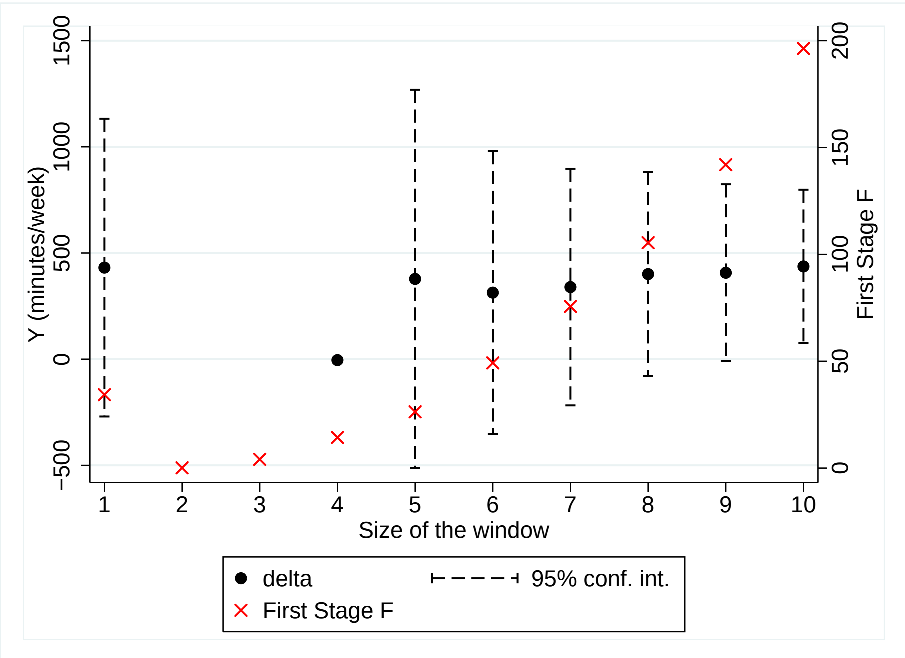
Note: all estimates include S plus other covariates, while R is instrumented by D. The p-value and the first stage F are calculated using robust standard errors.

Choice of window size

Figure A2: 2SLS estimates for different windows (for windows $|S| \in [2,10]$, regressions include $D \times S$ and $(1-D) \times S$; for $|S| = 1$ they are a Wald estimator with no covariates; when confidence interval or estimates are not shown they are larger than the graph interval).



(a) Men



(b) Women

Appendix B: regressions using age as running variable

In 2007, men mostly went into retirement exploiting the age-seniority combined path. Given that their NRA was higher (generally 65) most of them became eligible earlier than that. Indeed, in Figure B1 there is no evidence of a discontinuity in retirement behaviour at different age cut-offs (including 57 and 65).

In contrast, a large fraction of women (80% in 2007) who went into retirement exploited the rules for the NRA path (old age pension). In Figure B2 I ignore all contributory requirements and focus only on the discontinuity at age 60. In this way, I can avoid measurement error in S and also reduce the influence of rounding, because age is available in quarters.[3] The graphs clearly indicate a jump at age 60. The favourite specification according to tests and information criteria (a quadratic) shows a jump in retirement by 0.170 (s.e. 0.046) at age 60, not far from $\widehat{\gamma_D}$ obtained using S as running variable.

The evidence for house work is less clear. The figure with age in quarters has a large dispersion, while if I aggregate age at intervals of one year I observe a jump of around 200 minutes/week at age 60 (Table B1). In both cases, fitted linear polynomials predict a similar discontinuity. Using quarters, the point estimate is 180 (p-value 0.018, Table B2). The resulting 2SLS estimate for δ is 504 minutes/week (Table B3). On the other hand, a quadratic polynomial suggests no jump, and it is preferable according to the Akaike criterion (though not according to the Bayesian). Nevertheless, there is evidence of a kink at eligibility. Indeed, one possible reason for the different result is that the proportion of retired women already shows a large increase starting at age 57, because they can start going into retirement following the seniority path. This change is associated with a steeper slope in the average Y in the interval $[57,60]$, while the curve becomes flatter after age 60. One alternative would be to exploit this kink as an instrument, assuming that without retirement the average house work would have had a continuous slope at eligibility. In the presence of a jump, exploiting the kink can improve efficiency, although it may also induce a bias if the treatment effect varies linearly with S (see Dong, 2014a). Using $(age-60)$ and $(age-60)^2$ as covariates, and exploiting the kink and the jump at 60 together as instruments for R, we obtain a point estimate of 584 (s.e. 199), quite large and more similar to OLS results (Table B4). Similar results are obtained by using two dummies for $\mathbf{1}[age \geq 57]$ and $\mathbf{1}[age \geq 60]$, with or without interactions with $(age-60)$.

There are two main reasons to prefer the estimates using S as running variable. First of all, we can interpret them as the local average treatment effect for those individuals who go into retirement as soon as they are eligible. In contrast, the discontinuity at age 60 does not have such a clear interpretation, because a relevant group of women could go into retirement earlier than that. Secondly, there is evidence of discontinuities in baseline covariates at age

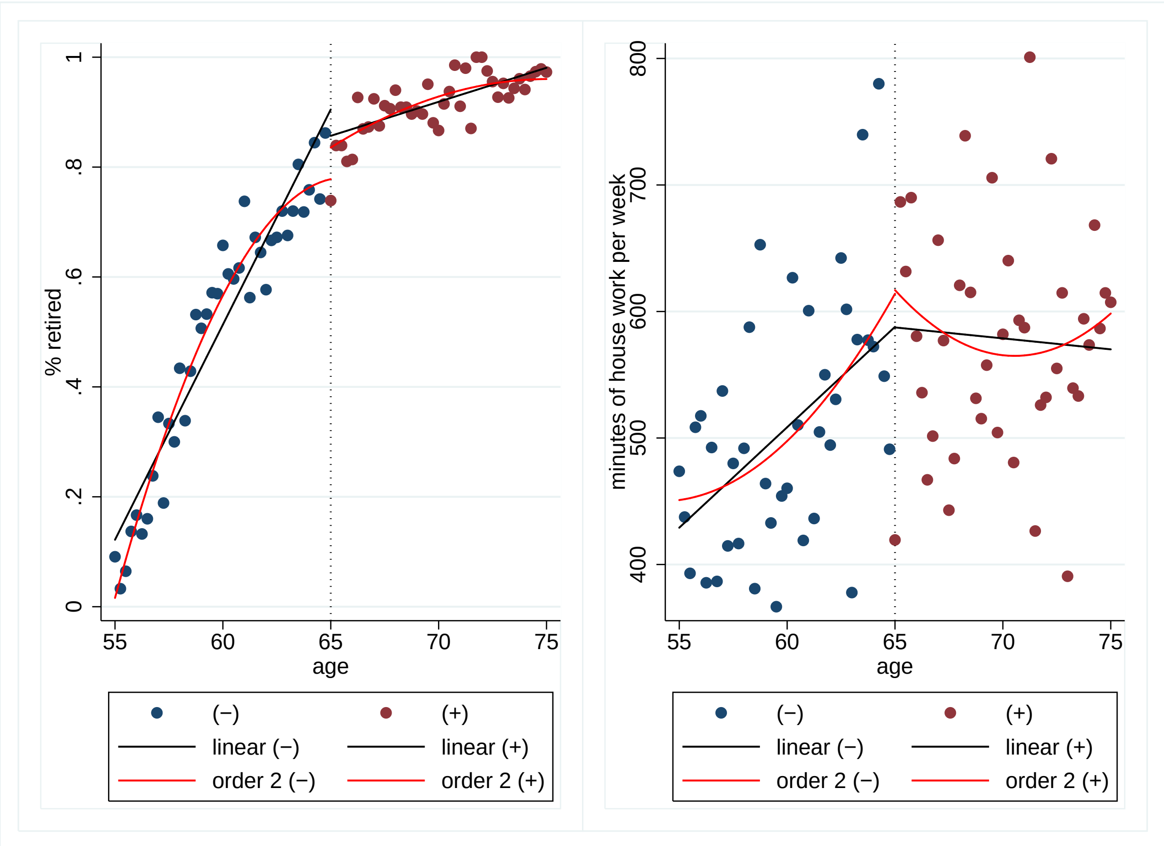
[3] I still do not consider observations at exactly age 60. I cannot exclude the presence of rounding at the quarter level. Moreover, the exact NRA in 2007 was actually 60 years and 2 months.

60, which are stronger than those found at the time of eligibility. If I introduce covariates in the 2SLS regression exploiting the jump at age 60 by mean of a linear polynomial, I obtain an estimate for δ of 430 minutes/week (s.e. 214), very similar to my main result using S.

The analysis with age as running variable could also be replicated also on the TUS sample. However, this would not be helpful in improving the gender comparison, because I would still not be able to obtain estimates for men for the reasons discussed above. Further refinements of the result for women only are of general interest, but I believe they exceed the purpose of the present work.

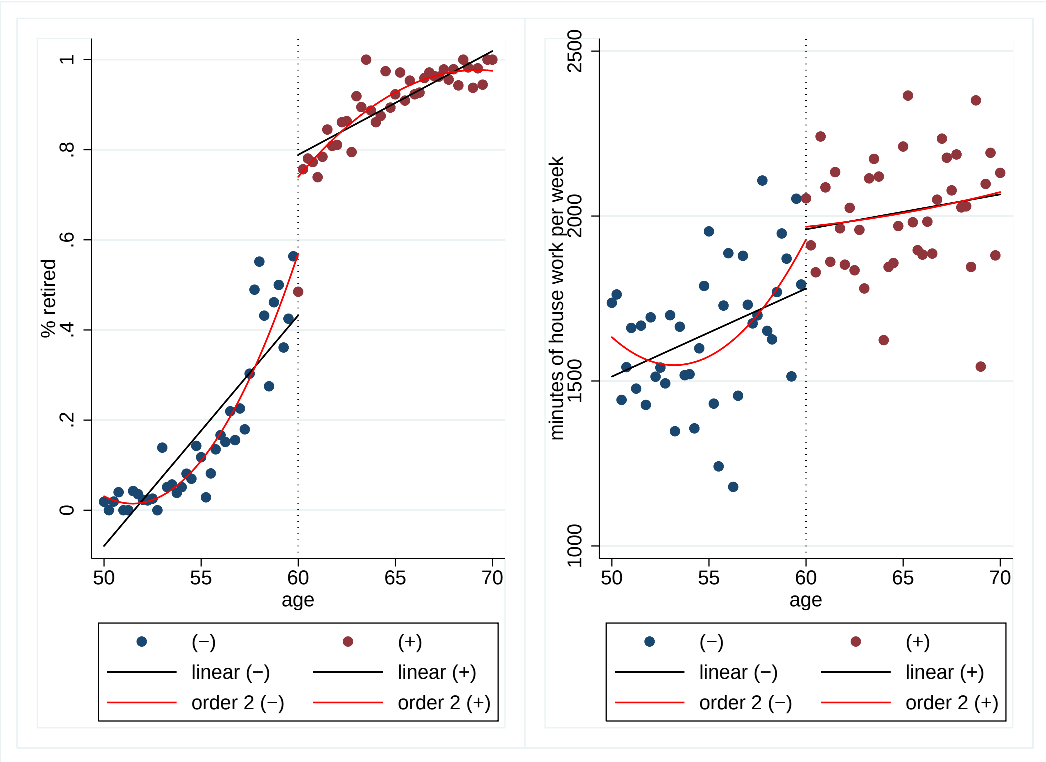
Graphs for men

Figure B1: Retirement and house work with respect to age, SILC 2007, men with age $\in [55, 75]$, either working or retired from work

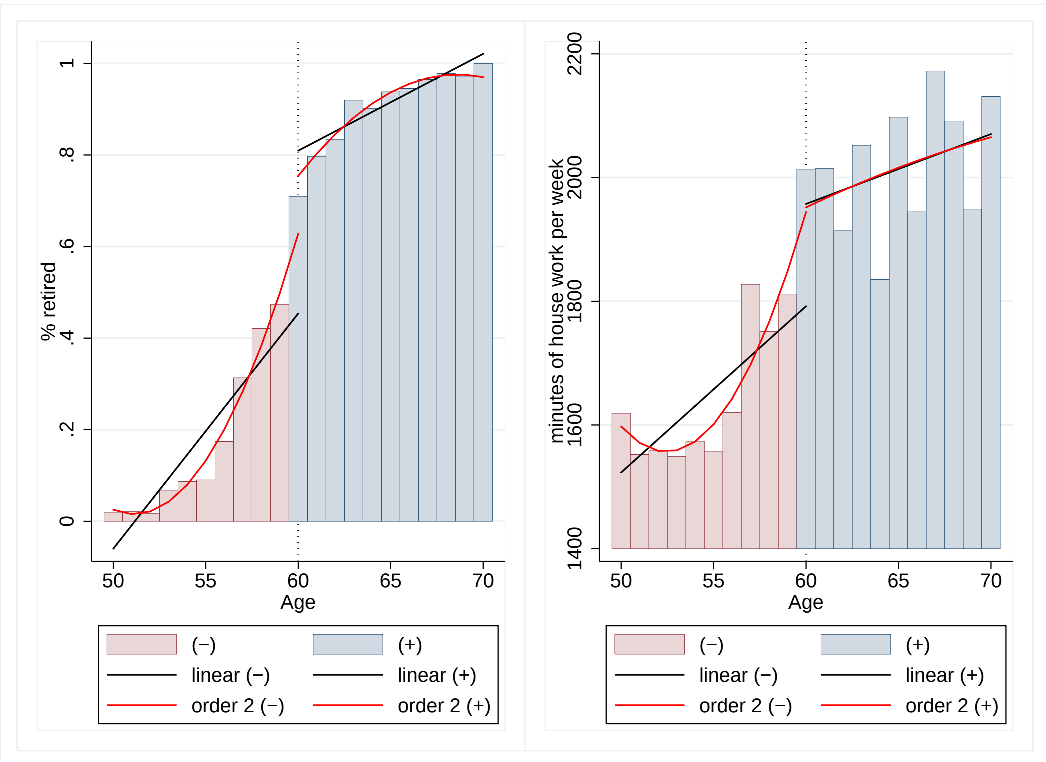


Graphs for women

Figure B2: Retirement and house work with respect to age, SILC 2007, women with age $\in [50, 70]$, either working or retired from work



(a) Age in years and quarters



(b) Age in years

Estimates for women

Table B1: First stage OLS for retirement status, SILC 2007, women with age between 50 and 70

Dep var R	(1)	(2)	(3)	(4)
$\mathbf{1}[age \geq 60]$	0.220**	0.155**	0.170***	0.356***
	(0.087)	(0.065)	(0.046)	(0.029)
$(1-\mathbf{1}[age \geq 60]) \times (age-60)$	0.007	0.033***	0.033***	0.013***
	(0.020)	(0.009)	(0.003)	(0.001)
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^2$	-0.002	0.001	0.000***	
	(0.002)	(0.000)	(0.000)	
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^3$	-0.000*	0.000		
	(0.000)	(0.000)		
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^4$	-0.000*			
	(0.000)			
$\mathbf{1}[age \geq 60] \times (age-60)$	0.013	0.017**	0.013***	0.006***
	(0.017)	(0.008)	(0.003)	(0.001)
$\mathbf{1}[age \geq 60] \times (age-60)^2$	0.000	-0.000	-0.000***	
	(0.002)	(0.000)	(0.000)	
$\mathbf{1}[age \geq 60] \times (age-60)^3$	-0.000	0.000		
	(0.000)	(0.000)		
$\mathbf{1}[age \geq 60] \times (age-60)^4$	0.000			
	(0.000)			
Constant	0.515***	0.571***	0.570***	0.432***
	(0.065)	(0.049)	(0.036)	(0.023)
Observations	3379	3379	3379	3379
R^2	0.624	0.624	0.623	0.616
$H_0 : \gamma_D = 0$ (p-value)	0.012	0.018	0.000	0.000
H_0 : ” (p-val clust)	0.001	0.001	0.000	0.000
Dong’s $\widehat{\gamma}_D$	0.218	0.163	0.180	0.360
Dong’s $\widehat{\gamma}_D$ (p-value)	0.013	0.012	0.000	0.000
G (p-value)	0.221	0.168	0.204	0.000
AIC	1599.369	1600.065	1596.416	1659.753
BIC	1660.623	1649.068	1633.168	1684.255

* p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. Age is measured in quarters. The selected sample includes only workers or retirees and excludes proxy interviews, missing house work and observations with age exactly equal to 60. γ_D is the coefficient for the discontinuity at eligibility. G (p-value) is Lee and Card (2008) statistic. Dong’s refer to Dong (2014b) correction. AIC is the Akaike criterion; BIC is the Bayesian criterion.

Table B2: Reduced form OLS for time spent in domestic work (minutes/week), SILC 2007, women with age between 50 and 70

Dep var Y	(1)	(2)	(3)	(4)
$\mathbf{1}[age \geq 60]$	146.362	188.426	38.621	179.504**
	(219.935)	(164.433)	(117.553)	(75.775)
$(1-\mathbf{1}[age \geq 60]) \times (age-60)$	1.636	7.856	28.030***	6.672***
	(46.496)	(22.107)	(8.516)	(2.083)
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^2$	-1.369	-0.700	0.516**	
	(4.567)	(1.267)	(0.200)	
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^3$	-0.045	-0.020		
	(0.166)	(0.020)		
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^4$	-0.000			
	(0.002)			
$\mathbf{1}[age \geq 60] \times (age-60)$	3.610	-20.034	1.534	2.648
	(55.214)	(26.333)	(10.163)	(2.365)
$\mathbf{1}[age \geq 60] \times (age-60)^2$	-1.182	1.318	0.027	
	(5.307)	(1.466)	(0.241)	
$\mathbf{1}[age \geq 60] \times (age-60)^3$	0.073	-0.021		
	(0.193)	(0.023)		
$\mathbf{1}[age \geq 60] \times (age-60)^4$	-0.001			
	(0.002)			
Constant	1844.494***	1857.992***	1928.870***	1780.308***
	(133.376)	(102.780)	(77.637)	(52.165)
Observations	3379	3379	3379	3379
R^2	0.035	0.035	0.035	0.033
$H_0 : \beta_D = 0$ (p-value)	0.506	0.252	0.743	0.018
H_0 : " (p-val clust)	0.443	0.247	0.737	0.033
Dong's $\widehat{\beta_D}$	145.406	202.707	51.787	181.516
Dong's $\widehat{\beta_D}$ (p-value)	0.530	0.230	0.663	0.017
G (p-value)	0.060	0.079	0.083	0.051
AIC	56705.229	56701.525	56699.321	56701.187
BIC	56766.482	56750.527	56736.073	56725.688

* p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. Age is measured in quarters. The selected sample includes only workers or retirees and excludes proxy interviews, missing house work and observations with age exactly equal to 60. β_D is the coefficient for the discontinuity at eligibility. G (p-value) is Lee and Card (2008) statistic. Dong's refer to Dong (2014b) correction. AIC is the Akaike criterion; BIC is the Bayesian criterion.

Table B3: 2SLS regressions for time spent on house work (in minutes per week), SILC 2007, women with age between 50 and 70

Dep var Y	(1) No X	(2) No X	(3) No X	(4) With X	(5) With X	(6) With X
R	1219.5	227.1	503.8**	1389.7	171.9	430.1**
	(1095.0)	(681.4)	(208.3)	(1134.1)	(691.2)	(214.2)
$(1-\mathbf{1}[age \geq 60]) \times (age-60)$	-32.4	20.6	0.2	-39.8	20.3	0.7
	(52.6)	(27.8)	(4.2)	(54.4)	(27.4)	(4.2)
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^2$	-1.3	0.4		-1.4	0.4	
	(1.6)	(0.5)		(1.7)	(0.4)	
$(1-\mathbf{1}[age \geq 60]) \times (age-60)^3$	-0.0			-0.0		
	(0.0)			(0.0)		
$\mathbf{1}[age \geq 60] \times (age-60)$	-40.4	-1.4	-0.3	-51.9	-4.3	-1.9
	(41.2)	(17.2)	(3.2)	(38.5)	(16.4)	(3.1)
$\mathbf{1}[age \geq 60] \times (age-60)^2$	1.8	0.1		2.3	0.1	
	(1.8)	(0.3)		(1.7)	(0.3)	
$\mathbf{1}[age \geq 60] \times (age-60)^3$	-0.0			-0.0		
	(0.0)			(0.0)		
Constant	1161.2*	1799.5***	1562.5***	907.1	1790.0***	1540.6***
	(701.6)	(442.4)	(130.4)	(836.8)	(512.8)	(161.1)
Observations	3379	3379	3379	3379	3379	3379
$H_0: \delta = 0$ (p-val)	0.265	0.739	0.016	0.220	0.804	0.045
First Stage F	5.646	13.485	151.490	5.534	12.852	138.021

* p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. R is instrumented by $\mathbf{1}[age \geq 60]$. δ is the coefficient on R. Age is measured in quarters. The selected sample includes only workers or retirees and excludes proxy interviews, missing house work and observations with age exactly equal to 60. Covariates X include a constant, plus dummies for geographic area, population density, education and employment category. Coefficients are available on request.

Table B4: 2SLS regressions using kinks, SILC 2007, women age between 50 and 70

	No X	With X	No X	With X	No X	With X
			FIRST STAGE: dependent variable R			
$\mathbf{1}[age \geq 57]$	0.203***	0.202***	0.384***	0.367***		
	(0.031)	(0.030)	(0.067)	(0.066)		
$\mathbf{1}[age \geq 60]$	0.341***	0.323***	0.224***	0.213***	0.355***	0.336***
	(0.030)	(0.029)	(0.053)	(0.052)	(0.029)	(0.029)
$(age-60)$	0.025***	0.024***	-0.001	0.003	0.078***	0.078***
	(0.002)	(0.002)	(0.013)	(0.013)	(0.009)	(0.009)
$(age-60)^2$	-0.000	-0.000	-0.002**	-0.002*	0.003***	0.003***
	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
$\mathbf{1}[age \geq 57]\,(age-60)$			0.080***	0.075***		
			(0.027)	(0.026)		
$\mathbf{1}[age \geq 60]\,(age-60)$			-0.037	-0.041	-0.081***	-0.084***
			(0.028)	(0.027)	(0.018)	(0.017)
Constant	0.241***	0.342***	0.149***	0.263***	0.479***	0.584***
	(0.024)	(0.029)	(0.050)	(0.053)	(0.030)	(0.032)
R^2	0.622	0.638	0.624	0.640	0.617	0.634
First Stage F	132.341	121.928	76.779	69.200	94.977	89.501
			SECOND STAGE: dependent variable Y			
R	618.97***	563.97***	598.10***	548.43***	584.36***	537.46***
	(182.48)	(186.63)	(178.11)	(182.34)	(199.47)	(203.73)
$(age-60)$	-7.27	-10.38	-5.96	-9.44	-5.09	-8.78
	(11.97)	(11.76)	(11.69)	(11.49)	(12.99)	(12.74)
$(age-60)^2$	0.23	-0.12	0.20	-0.14	0.19	-0.15
	(0.61)	(0.60)	(0.61)	(0.60)	(0.62)	(0.61)
Constant	1487.04***	1432.00***	1499.18***	1442.63***	1507.16***	1450.14***
	(107.22)	(135.05)	(104.96)	(132.45)	(117.38)	(146.51)
Observations	3379	3379	3379	3379	3379	3379
Hansen's test (p-val)	0.203	0.166	0.552	0.462	0.219	0.145

* p<.10 ** p<.05 *** p<.01. Robust standard error in brackets. Age is measured in quarters. The selected sample includes only workers or retirees and excludes proxy interviews, missing house work and observations with age exactly equal to 60. Covariates X include a constant, age at first job, years of contributions, years spent in a paid job, plus geographic area, population density, education and employment category dummies. Coefficients are available on request. Conclusions are similar if we drop $(age-60)^2$ from the regressions.

Table B5: Regressions for different socio-economic variables, SILC 2007, women with age between 50 and 70

	(1) North	(2) Centre	(3) South	(4) College	(5) High school
$\mathbf{1}[age \geq 60]$	0.074**	-0.018	-0.056**	-0.065***	-0.067**
$H_0:\gamma_D=0$ (p-val)	0.034	0.544	0.049	0.006	0.039
G (p-value)	0.199	0.087	0.443	0.254	0.307
	(6) Middle sch.	(7) Primary sch.	(8) Private	(9) Public	(10) Self-empl.
$\mathbf{1}[age \geq 60]$	0.059**	0.073**	0.119***	-0.098***	-0.021
	(0.029)	(0.032)	(0.034)	(0.034)	(0.029)
$H_0:\gamma_D=0$ (p-val)	0.044	0.024	0.001	0.004	0.470
G (p-value)	0.443	0.034	0.760	0.071	0.529

* p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. The regressions include $(1-D)\times(age-60)$, $D\times(age-60)$ and a constant. γ_D is the coefficient for the discontinuity at eligibility.

Appendix C: poor health, household size and marital status

Table C1: 2SLS regressions for health, household size and marital status, men, SILC 2007 (selected sample)

	(1)	(2)	(3)
	Health fair or poor	Hh size	Married
R	-0.0611	0.0162	0.0775
	(0.0799)	(0.1935)	(0.0637)
S	0.0180***	-0.0454***	-0.0009
	(0.0044)	(0.0109)	(0.0036)
Constant	0.3997***	2.8975***	0.7918***
	(0.0319)	(0.0771)	(0.0256)
Observations	3930	3970	3970
R^2	0.029	0.054	-0.006
Average dep var for $R=0$	0.2972	3.1288	0.8173
$H_0: \delta = 0$ (p-val)	0.4446	0.9332	0.2242
H_0 : " (p-val clust)	0.4046	0.9302	0.1339
First Stage F	258.8492	265.5622	265.5622

* p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. R is instrumented by D. δ is the coefficient on R. The selected sample for column (1) excludes missing values in general health. The dummy for health equals one for fair, poor or very poor health.

Table C2: 2SLS regressions for health, household size and marital status, women, SILC 2007 (selected sample)

	(1)	(2)	(3)
	Health fair or poor	Hh size	Married
R	-0.0433	-0.0003	-0.0383
	(0.0849)	(0.1778)	(0.0775)
S	0.0169***	-0.0475***	0.0013
	(0.0052)	(0.0113)	(0.0048)
Constant	0.4305***	2.6706***	0.7239***
	(0.0371)	(0.0790)	(0.0338)
Observations	2672	2701	2701
R^2	0.031	0.071	-0.002
Average dep var for $R=0$	0.3280	2.8833	0.7032
$H_0: \delta = 0$ (p-val)	0.6099	0.9988	0.6208
H_0 : " (p-val clust)	0.3849	0.9988	0.6048
First Stage F	243.9813	253.9427	253.9427

* p<.10 ** p<.05 *** p<.01. Robust standard errors in brackets. R is instrumented by D. δ is the coefficient on R. The selected sample for column (1) excludes missing values in general health. The dummy for health equals one for fair, poor or very poor health.

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