Retirement, pension eligibility and home production*

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October 7, 2013

Abstract

I study the change in home production at retirement. Descriptive evidence from the 2007 Italian Survey on Income and Living Conditions shows that retirees spend much more time than workers on household chores, shopping and caring, even when the comparison is made for individuals of a given age. To account for the endogeneity of retirement, I exploit the discontinuity in pension eligibility generated by the Italian Social Security system. Estimates show that women increase time spent on household production at retirement by more than 400 minutes per week. No evidence of an equally large change is found for men.
JEL: J22, J26, D1
Keywords: retirement, house work, regression discontinuity

1 Introduction

The evidence of a drop in consumption at retirement spurred a large stream of research which tried to reconcile it with the permanent income hypothesis. Actually, Hurst (2008) argued that empirical studies found mostly a reduction in food expenditure,

*I wish to thank Marco Francesconi, Elena Stancanelli, Marcello Morciano, Carlo Mazzaferro, Patrick Nolen, Mark Bryan, Roberto Nistico, Jonathan James, Claudio Deiana, Ludovica Giua, Guglielmo Weber, Matthias Parey, and seminar participants at Essex for helpful suggestions. This paper was part of my PhD dissertation at the University of Essex, during which I received funding from the Economic and Social Research Council and from the Royal Economic Society Junior Fellowship, that are gratefully acknowledged. The views expressed in this paper are those of the author and do not necessarily reflect those of the Bank of Italy. CAPP stands for the Center for the Analysis of Public Policies (www.capp.unimore.it), to which the author is affiliated.

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which could be explained by an increase in home production. Following his argument, in this paper I provide additional evidence from Italy about the change in time spent on producing household goods at retirement. The main goal is to understand whether it can explain the drop in expenditure which was found, for the same country, by Battistin et al. (2009). This is of interest for policy makers, who may be willing to understand whether this decrease is rather due to inadequate retirement benefits.

The interest in estimates of the increase in home production at retirement is not limited to testing the permanent income hypothesis. Rogerson and Wallenius (2012) argued that this change in time use can be utilised to recover the relative value of two elasticities that are crucial in modeling labour supply: the intertemporal elasticity of substitution of leisure, and the elasticity of substitution between time and market good inputs in home production. However, differences between employed and retired individuals at any given age can differ from the quantity of interest, because retirees may have different preferences for leisure and house work (Rogerson and Wallenius, 2012, p. 14). To manage this problem, I identify and estimate the change in time use exploiting the discontinuities in pension eligibility induced by the Italian Social Security system. 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 the subject of several reforms in the last two decades, so that different rules apply to individuals who have retired in different years.

The Italian case is also 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, p. 16-19). I present separate estimates for men and women, in order to understand whether retirement has an equalizing effect.

One of the first pieces of empirical research on home production and retirement is Aguiar and Hurst (2005), where the authors found that households with a retired head spend 18 minutes more per day on shopping for food and preparing it. Other papers that have provided supporting evidence of this increase are Szinovacz and Harpster (1994), Szinovacz (2000) and Hurd and Rohwedder (2005, 2006) for the U.S.; Schwerdt (2005) and Luhrmann (2007) for Germany; Stancanelli and van Soest (2012) for France; and Luengo-Prado and Sevilla (2012) for Spain.

Studies attempting to estimate the drop in consumption after retirement have recently began to use fuzzy Regression Discontinuity Designs (RDD). Among others, Battistin et al. (2009) analysed data from the Italian Survey on Household Income and Wealth, exploiting the discontinuity around the time of eligibility for a pension.

They found a drop in nondurable expenditure to be about 9.8%, due to the retirement of a male head of household. This could be explained by a reduction in work-related expenses, in particular food, and a decrease in household size. However, they could not estimate the change in time spent on home production, because their dataset did not include such information. Stancanelli and van Soest (2012) used a similar framework and data from the French 1998-99 Time Use Survey to estimate the causal effect of either partners' retirement on house work in couples.

This paper employs a fuzzy Regression Discontinuity Design (RDD) to estimate the causal effect of retirement on time spent on home production, with data from the 2007 cross-section of the Italian Survey on Income and Living Conditions (SILC).[1] To the best of my knowledge, only Stancanelli and van Soest (2012) used (fuzzy) RDD to address this question. While they exploited the discontinuity in retirement at age 60 induced by the French system, here the running variable is time to/from eligibility (as in Battistin et al., 2009), which depends on both age and social contributions. The nature of the home production information in SILC is also different from Stancanelli and van Soest (2012). While their data are collected from a single day diary, in SILC respondents are asked about time spent in house work during an average week.

Section 2 presents the identification strategy, while section 3 introduces the dataset. In section 4 I provide descriptive evidence about time use. The main results are reported in section 5. In section 6 I discuss them and I conclude.

2 Identification strategy

I follow the identification strategy outlined by Battistin et al. (2009), which exploits the discontinuity in retirement behaviour with respect to time to/from eligibility for a pension.[2] As they noticed, if I define an individual as retired only when s/he does not work and s/he is recipient of a retirement pension, I should not observe anybody in this state before meeting the requirements. Restricting the sample to individuals who are currently employed or retired from work, I observe a sensible increase in the proportion of retired individuals between one year before eligibility and one year after. This motivates a RDD.

Define S_i as time to/from eligibility to retirement, $D_i \equiv \mathbf{1}\left[S_i \geq 0\right]$ as the dummy for being eligible, R_i as a dummy for being retired from work, and Y_i as the (observed)

[1] At the moment of writing, I am aware of only one economic related study using SILC data on home production. Addabbo et al. (2011) studied time allocation within double and single-earner couples, but they did not analyze retirement.

[2] Differently from them, I also discuss the consequences of using a discrete running variable, given that distance to eligibility is rounded to years.

time spent on house work. Individuals are indexed by $i = 1, ..., N$. I can further define Y_{1i} as the time spent on home production if i was retired, while Y_{0i} if s/he was still employed. For each single individual, I actually observe only one or the other, so that (Hahn et al., 2001)

$$Y_i = \delta_i R_i + \epsilon_i, \tag{1}$$

$$\epsilon_i \equiv Y_{0i},\ \delta_i \equiv Y_{1i} - Y_{0i}. \tag{2}$$

One way to identify the causal effect is to exploit the RDD. First of all, I need a discontinuity in retirement:

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

Given that the majority of retirement benefits in Italy come from state-managed funds, the eligibility rules are expected to have a strong effect on retirement behaviour.[3] This prior is corroborated by previous results from Battistin et al. (2009), who found a 43.5 percentage points increase in the proportion of retired household heads at $s = 0$.

In order to exploit the discontinuity in retirement, the potential time spent on house work without retirement must not change discontinuously at $s = 0$:

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

However, there might be age-specific effects that force individuals to exit the labour market and spend more time on home production. For instance, their partners' health may deteriorate, demanding a considerable amount of caregiving. The probability of such an event is quite likely to be a function of age and seniority, but there is no particular reason to believe it to be discontinuous at the specific and rather arbitrary point of eligibility, given also that requirements have changed every year since 1992.

I start the discussion with a constant effect framework, where $\delta_i = \delta\ \forall i = 1, ..., N$. If I define $\eta_i \equiv \epsilon_i - h_Y(S_i)$, a full specification would be

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

$$R_i = \gamma_D D_i + h_R(S_i) + \xi_i. \tag{4}$$

R_i is allowed to be endogenous if ξ_i and η_i are correlated. Given the definition of η_i,

[3] In 2007 SILC dataset, 50.03% of individuals aged 50 or more received a pension managed by these funds, while only 0.40% were recipients of a private pension.

we have that $E[\eta_i|S_i] = 0$ by assumption (A2). Define the reduced form for Y_i as

$$Y_i = \beta_D D_i + g_Y(S_i) + \theta_i, \tag{5}$$

$$\text{with } g_Y(S_i = s) \text{ continuous at } s = 0; E[\theta_i|S_i] \equiv 0. \tag{6}$$

Under assumptions (A1) and (A2), δ is equal to β_D/γ_D. 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 three additional problems.

First of all, in SILC I can measure time/to from eligibility only in years. As argued by Lee and Card (2008), this discreteness forces us to choose a parametric specification.[4] The most popular alternative is to specify it as a polynomial of order J.[5] For ease of exposition, define the vector

$$P_i \equiv \left[D_i S_i, ..., D_i S_i^J, (1 - D_i) S_i, ..., (1 - D_i) S_i^J\right]. \tag{7}$$

which does not include the eligibility dummy D_i itself. The model can be rewritten as

$$Y_i = \alpha_0 + \delta R_i + P_i\alpha + \eta_i^* + \eta_i, \tag{8}$$

$$R_i = \gamma_0 + \gamma_D D_i + P_i\gamma + \xi_i^* + \xi_i, \tag{9}$$

where $\eta_i^* \equiv h_Y(S_i) - P_i\alpha$ and $\xi_i^* \equiv h_R(S_i) - P_i\gamma$ can be interpreted as specification errors. As noticed by Lee and Card (2008), I do not need the polynomial to be the correct model. It is sufficient that the induced approximation errors do not bias the OLS estimator for the true discontinuities. The reduced form for Y becomes

$$Y_i = (\alpha_0 + \delta\gamma_0) + \delta\gamma_D D_i + P_i(\alpha + \delta\gamma) + \delta(\xi_i^* + \xi_i) + \eta_i^* + \eta_i. \tag{10}$$

If I interpret the polynomial specification as the Best Linear Projection ($BLP[\cdot]$) of the true functions h_R and h_Y on the vector P_i, it follows that by construction $BLP[\xi_i^*|P_i] \equiv 0$, and $BLP[\eta_i^*|P_i] \equiv 0$. However, suppose that I want to estimate eq.

[4] Although I follow the discussion on Fuzzy RDD in the appendix of Lee and Card (2008), I expand it at some points. While the authors proposed a mean independence property for the specification errors (Lee and Card, 2008, pg. 673), I prefer to focus on their linear projection. I also discuss the implications of choosing the order for the polynomial either on the basis of the outcome or of the treatment reduced forms.

[5] The same argument follows if I allow the order of the polynomial to differ to the right and left of eligibility.

(10) by OLS. In order for the estimated discontinuity in Y_i to equal $\delta\gamma_D$, we need:

$$BLP\left[\delta\xi_i^* + \eta_i^* | D_i, P_i\right] = BLP\left[\delta\xi_i^* + \eta_i^* | P_i\right] \equiv 0. \tag{11}$$

As long as this condition holds, we can recover δ by 2SLS, using D_i as an instrument for R_i. To be precise, the crucial assumption for 2SLS to be consistent is

$$\text{(A3).}\ BLP\left[\eta_i^* | D_i, P_i\right] = BLP\left[\eta_i^* | P_i\right] \equiv 0$$

which implies that the approximation does not introduce any discontinuity in the main equation of interest (8), so that D_i can be excluded from it. If $BLP\left[\xi_i^* | D_i, P_i\right] = BLP\left[\xi_i^* | P_i\right] \equiv 0$, then the discontinuity in the BLP of R_i on (D_i, P_i), call it γ_D^*, is also equal to the true jump in retirement (γ_D). However, the equation for retirement is only a first stage, and therefore we only need it to be a best linear approximation.[6]

One might prefer to look at the two reduced forms $E\left[Y_i | S_i\right]$ and $E\left[R_i | S_i\right]$ separately and then estimate δ as the ratio of the two discontinuities. In this parametric setting, however, I chose to employ 2SLS. This strategy has the advantage of being clearer, given that it is equivalent to an instrumental variable approach. The question then is how to choose the correct order for the polynomial. Given that the retirement equation is only a first stage, 2SLS are consistent as far as the chosen polynomial satisfies assumption (A3). Suppose, instead, that one decides to use the polynomial order which provides the best fit for R_i. This might not be the correct one for the time use equation (8), so that we would introduce a spurious discontinuity. As a result, eq. (8) would not satisfy the exclusion restriction on D_i, and the 2SLS estimator would not be consistent for δ. With this caveat in mind, in the following I will estimate δ by 2SLS on the best polynomial specification for Y_i.

The second problem of identification is caused by the fact that S is discrete because it is rounded in years. In theory, one should calculate the distance to eligibility in weeks or days, but my dataset does not allow me to do so. Dong (2012) 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. In particular, one must assume that the moments of the rounding error are independent from S and that the true

[6] The reason is that, under assumption (A3), the discontinuity in the BLP of Y_i on (D_i, P_i) would be equal to $\delta\gamma_D^*$, so that 2SLS would be still 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.

functional form for $h_R(S_i)$ and $g_Y(S_i)$ are polynomials of possibly unknown order.[7] Unfortunately, at the moment I do not have access to any additional archive that I can use to observe S in smaller intervals of time. Nevertheless, I calculated the bias-corrected estimates assuming an uniform distribution inside each year interval. This seems to be at least a good approximation, given that eligibility depends on a mixture of years of contribution and age, so that it is equivalent to assume that individuals started to work and were born more or less uniformly during the year.[8]

The last problem, discussed in Battistin et al. (2009), is that the process of recovering S from other survey information 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 the retirement status before being eligible, that is when $S < 0$. The reason is that I defined individuals as retired only if they received a pension. Following the discussion in Battistin et al. (2009), suppose that the observed distance to/from eligibility S is equal to the true value S^* when the (unobserved) indicator Z is equal to 1, while it is measured with error and equal to $\tilde{S}$ when $Z = 0$. Consistency of the 2SLS estimator requires an additional assumption. The measurement error process, defined by Z and the error-ridden measure S, must be statistically independent from (Y, R), conditional on the true value S^*. This assumption is non-refutable, because Z is not observed. One concern is that S is necessarily calculated differently for workers and retired. 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, I also show results based on age as running variable.

Although assumptions (A1)-(A3) are parametric, I based the whole discussion on the desire to approximate the underlying continuous functions. If there are heterogeneous effects, so that $\delta_i \neq \delta \ \forall i = 1, ..., N$, 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

[7] See Dong (2012) for the other assumptions. It must be added that the current literature does not discuss the potential problems arising from the presence of both rounding and misspecification. Note, however, that in the main results we always fail to reject the null of correct specification for the reduced form of Y.

[8] I estimated the distribution of date of birth within a year using data from the Italian administrative records (http://demo.istat.it/altridati/IscrittiNascita/, last access: 06/03/13). Unfortunately, they are available only for recent years, between 2001 and 2011. The first four empirical moments (0.507, 0.339, 0.255, 0.203) are similar to the theoretical ones from a uniform distribution (0.500, 0.333, 0.250, 0.200; see Dong, 2012, for a similar comment on the US). I also used data on the month of hire for employees, years 2009-2011 (*Comunicazioni Obbligatorie*, available only for some regions at http://www.venetolavoro.it/servlet/dispatcherServlet?load=/osservatorio/seco/SeCo_04_12.xls, last access: 09/03/13). Although there are downs in December and August, followed by picks in September and January, the first four empirical moments (0.492, 0.338, 0.259, 0.211) are not too far from the theoretical ones with equiprobability of being hired in each month (0.500, 0.348, 0.273, 0.228).

I also need that, *near* $s = 0$, R_i as a deterministic function of S_i is monotonic, while δ_i and $R_i(S_i)$ must be jointly independent of S_i (see Hahn et al., 2001). The monotonicity can be defended using the same institutional argument supporting $\gamma_R \neq 0$. The independence assumption allows the group of compliers to have different δ_i in a neighborhood of s = 0. However, it requires that neither the distribution of groups, nor the distribution of δ_i within each of them change discontinuously with S_i at $s = 0$. In the cross-sectional case, this would not hold if workers near retirement could manipulate S adding some years to their seniority (or age). It does not seem to be a problem for the present study, given that the National Social Security Institutions keep track of a worker's contribution history. Despite its local properties, the LATE at eligibility is of interest for a policy maker who is planning to strengthen the seniority and age requirements.

3 Data

The Italian component of the European Union Survey of Income and Living Conditions is a stratified sample of the households' population.[9] Conducted every year since 2004, its main target is to collect quantitative and qualitative information to provide measures of deprivation and inequality.

Here I discuss only the main steps I followed in generating the estimation sample, while details are provided in the online Appendix A (Ciani, 2013). 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". Conversely, I defined workers as individuals with "employed" as self reported employment status, excluding those who have not worked in the week prior to the interview because of being temporarily unemployed or under a temporary layoff public scheme called *cassa integrazione*.

Distance to/from eligibility S is calculated as age at interview minus age at eligibility. I recovered current age as year of interview minus year of birth, to which I subtracted one if the quarter in which the interview took place was before the quarter of birth. Age at retirement is instead calculated as age at first regular job, plus years spent in paid job, plus one. The final correction is taken to account for rounding.[10] The age at eligibility is then recovered simulating the rules that applied in the year

[9] SILC 2007 microdata must be requested directly from ISTAT (www.istat.it), which provides them free of charge to researchers working in public and/or not-for profit institutes.

[10] Without this correction, the estimated jump in retirement is smaller for men, while it is larger and clearer for women. The estimates of the effect on Y are in line with those discussed here, although $\widehat{\delta}$ for women is slightly smaller. See the online Appendix A (Ciani, 2013) for the full results.

in which the individual went into retirement, calculated as year of birth plus age at retirement, or plus the current one for workers.[11] Although the way in which I built S is inspired by Battistin et al. (2009), some steps are different, mainly because their dataset includes direct information about the year of retirement.

In 2007, people could retire following two alternative paths. The first was based on social contributions. The minimum requirement was 35 years, combined with 57 years of age for employees, or 58 for self-employed. This path could be taken without age limits if the worker had at least 39 years of contribution (40 for self-employed). For individuals who started working before 1995, the related retirement benefit was called *pensione di anzianità*. The second path was a National Retirement Age (NRA), fixed to 65 for men and 60 for women, plus a minimum requirement of social contributions (between 15 and 20 years). Individuals who followed this route could access the *pensione di vecchiaia*. We know from administrative data that there are strong gender differences in the proportion of individuals who retired following one or the other route. Among new pension recipients in 2007, 58% of men retired with a *pensione di anzianità*, exploiting their social contributions, while almost 80% of women with a *pensione di vecchiaia*.[12] Eligibility requirements were more generous in the past and they have been changed almost every year since 1992 (see Table 2 in Battistin et al., 2009, for details).

Table 1 reports sample selection by gender. I kept only workers or retirees, for two main reasons. First, I am not interested in comparing them to housewives or other inactive individuals. Second, we cannot define S for those who have never worked in a paid job. I also excluded all proxy interviews, which is the case when another household member provides the information on an individual who is not available at the time of interview.[13] The reason is that they are likely to increase measurement error and not to be particularly reliable for Y. There are few missing values for house work.[14]

As in Battistin et al. (2009), I kept only the window $S \in [-10, 10]$, in order to limit the influence of observations far away from the eligibility threshold, and I excluded

[11] In the process of recovering the rules, apart from institutional websites, I consulted Brugiavini and Peracchi (2004), Morciano (2007) and Intorcia (2011). I do not report a full table with requirements here, but the do-file generating S from raw data is available and documents each step with reference to the specific law.

[12] Source: National Institute for Social Security (INPS) - www.inps.it (last access: 28/11/2012).

[13] Including proxy interviews the graphical evidence is less clear for women, but all main estimates are quantitatively similar. Full results are available on request.

[14] The other variables employed here do not contain any missing for the sole reason of having been imputed by the ISTAT using multivariate methods. While for income data an imputation factor is available, no such information is reported for qualitative variables. Although this standard practice is debatable, ISTAT does not release the original raw data and therefore I cannot provide details.

observations with $S_i = 0$. The fact that contributions, age at first job and time spent in 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 (2012, pg. 25), is to discard observations with $S_i = 0$.[15]

I did not use sample weights, because they were designed for the original sample and it is not clear whether they would be appropriate in the selected one. Furthermore, the weights are highly dispersed: the 95 percentile is approximately 13 times larger than the 5 percentile, implying that some individual observations would be over-weighted in the final regression. Nevertheless, in section 5.4 I discuss what happens either including stratification variables in the regression or employing sample weights.

4 Descriptive evidence on time spent on home production

4.1 Does SILC differ from the Time Use Survey?

In SILC 2007, Italian respondents aged 15 or more were asked *"On average, how much time per week do you spend on domestic and family-related work (household chores, shopping, caregiving), in hours and minutes?"*.[16] Hereafter, Y_i is equal to the individual answer to this question, measured in minutes per week.

To better understand the content of this information, I compare it with the Italian Time Use Survey 2008-2009 (TUS), where "family related" work consists of cooking, cleaning dishes, cleaning the house, doing the laundry, sewing, knitting, shopping, and general administrative work. It also includes gardening, taking care of pets, maintenance of the house and vehicles. Lastly, it accounts for time spent on caring for children or adults. Unfortunately, the TUS does not collect information on years of contribution, so that it is not possible to replicate the RDD.

In Table 2, I compare data from the overall SILC sample to the Time Use Survey 2008-2009.[17] At the aggregate level, house work time is lower in SILC with respect to

[15] Results including the zeros are reported in the online Appendix A (Ciani, 2013). The estimates for the jump in retirement are generally smaller than those presented in section 5, but the main results on the effect on home production lead to the similar conclusions. For men I still do not find evidence of an increase in house work at retirement. For women the estimated effect is larger, but comparable with the main estimates.

[16] The question was also asked in the following year. However, the 2008 cross-section contains a large number of missing values (18.05%) which casts doubts on its validity.

[17] All estimates and figures are obtained using Stata$^{\text{TM}}$12. OLS regressions are estimated using the `regress` command with `robust` standard errors. 2SLS is estimated using the `ivreg2` command

"family related" work from column TUS (A). The difference is proportionally larger for men. After age 65, both samples display a drop in participation and average minutes per day for women. However, the decrease is larger in SILC. For men I observe an increase in average minutes using both datasets, but SILC shows a drop in participation rate against an increase in TUS (A). Comparing retired and employed individuals, in both samples retirees spend more time on house work, but the difference is larger in the Time Use survey. Moreover, participation slightly drops for women in SILC while it increases in the other dataset.

One might conclude that there is a substantial under-reporting in SILC. However, the difference with TUS data, which is stronger among the elderly, is more likely to be related to a different definition. The general question posed in SILC might exclude some activities. While caring and shopping are explicitly mentioned, "household chores" is 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". However, it might exclude "semi-leisure" chores, such as gardening. To provide indirect evidence in favour of this hypothesis, in columns labeled TUS (B) I redefined the variable in the Time Use Dataset, keeping only shopping, cooking, caring and "core" household work. As expected, the averages for men are generally closer to SILC, in particular for those aged 65 or over and for retirees.

4.2 Age profile of house work

Figure 1 shows how time use changes at different ages.[18] In order to provide an overall picture, I focus here on the original sample, without selecting only workers or retirees. For both men and women time spent on paid work is quite stable until the late forties. The average for women is lower, because of their lower employment rate and higher amount of part-time work. For both genders, after age 50 time spent in paid work decreases at a steep rate, down to almost zero at age 65. Home production exhibits two peaks for men, around age 40 and after 65. The first is related to the presence of young children in the household. I then observe a decrease while they grow up,

with `robust` standard errors (Baum et al., 2007). Lee and Card (2008) standard errors, the G test, the RESET and Dong (2012) corrected estimates are produced using commands that I wrote, available with the do-files. Estimates are reported with the help of `esttab` command (Jann, 2007). Do-files that produce final results starting from raw data are available upon request.

[18] Using age in quarters the figure is noisier, because the sample size is not large enough in each cell. The line plots a 95% confidence interval for a 6th order polynomial in age, estimated using OLS with heteroskedasticity robust standard errors. Higher orders tend to produce numerical problems due to near collinearity.

followed by a steep increase after age 55, which mimics the opposite pattern of paid work. For women the first peak is slightly earlier, at the end of their thirties, and it is not followed by a decrease. The second peak is between age 60 and age 65, when almost all of them go into retirement. Then I observe a decline, which could be due to health deterioration and the shrink in household size due to the partner's death.

Figure 2 draws attention to the window 50-70, during which the majority of individuals retire. I kept only workers and retirees, as defined in the previous section. For workers of both genders I notice a decline in the average Y with age. This could be due to demographic factors, or to a selection effect, where only workers with less productivity in household work are left in employment. The most important observation is that, at any age, the average Y is larger for retirees than for workers. It is also interesting to notice that not only men spend much less time on house work than women, but for females the difference between retirees and workers is almost double that for males.

5 The change in house work at retirement

5.1 Graphical analysis and discontinuities

I now focus on the selected sample. Figure 3 plots the mean value of R and Y for each S. All fitted polynomials are interacted with the eligibility dummy. For both genders I observe a small proportion of individuals who retired before meeting the eligibility criteria, although for women I notice some increase at $S = -1$. 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 $S = 10$. The jump at eligibility is confirmed by linear and quadratic polynomials.

Time spent on house work is slightly increasing before eligibility. Women spend almost four times the minutes per week dedicated by men to home production, despite the fact that the proportion of retirees at $S < 0$ is almost the same. After eligibility is met, I notice a progressive increase in the average Y for men, but there is no clear evidence of a discontinuity. I observe an increase at $S = 0$ around 50 minutes/week, but it is followed by alternate falls and rises. For women, time spent in home production is quite constant before eligibility. I then observe a jump at $S = 0$ by nearly 160 minutes/week, followed by an increase up to $S = 4$. 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. 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.

I know from McCrary (2008) that a discontinuity in the density function at eligibility might be a sign of individuals sorting around the threshold, even if a continuous density function is neither a sufficient nor a necessary condition for identification. Density plots are reported in the online Appendix B (Ciani, 2013). I observe no change in the density at $S = 0$ for men. For women I observe a drop of around 1%, 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 compulsory at $S = 0$ according to 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. Hence I do not take the observed drop as evidence of sorting.

5.2 The jump in retirement at eligibility

Table 3 shows the results of regressions of R_i on the eligibility dummy D_i, a polynomial in S_i and their interactions.[19] I focus on regressions up to the 3rd order because graphical evidence, available on request, shows that 4th order polynomials tend to overfit at $S_i = 0$. For model selection, I focus on minimizing the Akaike (AIC) and Bayesian (BIC) information criteria. The first is suggested by Lee and Lemieux (2010), while the second is useful in this context as it puts more weight on the number of parameters to be estimated. I also discuss Ramsey's RESET test of correct specification, obtained testing the significance of the square and cube of fitted values as additional covariates.[20] 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 that includes a dummy for each value of S.

For men (columns (1)-(3) in Table 3), both a cubic and quadratic polynomial estimate a jump in retirement (γ_D) around 30% at eligibility, this being statistically significant at the 1% level.[21] The Akaike criterion favours the highest order, though all 3rd order terms are not statistically significant at conventional levels and the Bayesian criterion is minimized with the quadratic. Ramsey's RESET test does not reject the

[19] Results with no interactions, available on request, are stronger for the retirement discontinuity and more precise for the effect on Y.

[20] This is equivalent to testing the significance of additional higher order terms in the polynomial.

[21] I can also compare $\widehat{\gamma_D}$ with results from Battistin et al. (2009), who estimated an increase in the proportion of retired male heads at $s = 0$ by 0.435 (s.e. 0.038), using a quadratic polynomial with no interactions. If I run the same regression on SILC, I obtain $\widehat{\gamma_D} = 0.398$ (s.e. 0.027). A t-test for equality fails to reject the null with p-value 0.468. If instead I use a quadratic polynomial with interactions on their dataset, $\widehat{\gamma_D}$ is 0.252 (s.e. 0.069), closer to the equivalent result in SILC (0.313, s.e. 0.048). I used Battistin et al. (2009) files available on the American Economic Review website.

null of correct specification at the 5% level. Differently, the G test strongly rejects the constraints imposed by the polynomials. Lee and Card (2008) argued that this is not a problem, as far as the best linear projections of the specification error does not bias the estimator of the discontinuity. In this case, they proposed to correct the standard errors by clustering on S. The p-values for the test of $\gamma_D = 0$ is still less than 1%.[22] Lastly, Dong's (2012) corrected estimates are smaller for the 3rd order polynomial, with a p-value 0.055, but they do not differ much in the 2nd order.

For women (columns (4)-(6) in Table 3), 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% and statistically significant. The statistical tests do not give a clear indication. The G test is passed at the 5% level with the cubic and not with the quadratic, but the RESET test gives the opposite result. The Akaike information criteria leads us to choose the cubic regression, but the Bayesian is minimized for the second order, and it should be noted that the R^2 does not change up until 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 (0.182), but still statistically significant at conventional levels.

5.3 The effect of retirement on home production

OLS regressions of Y on R and S show a large increase in house work at retirement. After leaving employment, men (Table 4, column (2)) spend 287 more minutes per week on this activity, which is a large change given that the average of Y among workers is 371 minutes/week. For women (Table 4, column (5)) the increase associated with retirement is larger in absolute value (610 minutes/week), but smaller in proportion to the average worker (1533 minutes/week).

OLS results may be biased by the endogeneity of retirement status, given that at any age workers leaving employment may have different preferences or constraints with regard to house work. Therefore we focus on the change in Y at $s = 0$, because eligibility is exogenously defined by institutional rules. In the regressions of Y on D and a polynomial in S, I discuss only the linear and quadratic cases, because information criteria invariably lead us to prefer the simplest specification and graphical analysis

[22] I report these additional p-values in all tables, but in the rest of the paper I do not comment them as they do not lead to different conclusions. Lee and Card (2008) also proposed to correct the s.e. by estimating the variance of the specification error. In all tables, using this correction, the p-values are similar to the clustered ones. Results are available on request.

did not show large differences. I also present and compare a simple linear regression, with no interaction between D and S.[23]

Despite the strong evidence of a jump in retirement at eligibility for men, none of the estimated models show a parallel discontinuity in the average time spent on home production (Table 5, columns (1)-(3)). Regression analysis is therefore in line with the intuitions resulting from graphical inspection. To recover the causal effect δ of retirement on house work, I use 2SLS, instrumenting R with D. The highest estimate (Table 6, column (3)) is 73 minutes/week, obtained including only S. It is around 25% of the relative OLS estimate and it is not statistically significant at the 5% level.[24]

To understand whether results differ sensibly across different groups, Table 8 shows 2SLS estimates splitting by education, area and category. I do it separately, because of sample size.[25] The estimated effect is economically significant for college graduate (176 minutes/week) and in the North (148 minutes/week), but not statistically significant. For private and public employees it is larger than for self-employed (113 and 105 minutes/week against -21), but not far from the one estimated for the whole sample. The only estimate which is statistically significant, though 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.

From a theoretical point of view, it is strange that the effect for men is, at least overall, quite small and not sensibly different from zero. Given the strong increase in available time associated with retirement, I should expect at least a partial increase in time spent on home production.[26] There are two possible reasons. The first is that men, at retirement, usually put the most of their effort on "semi-leisure" chores, such as gardening or house-repair. Indeed, Stancanelli and van Soest (2012) found that men's increase in time spent on home production was mostly in this category. Furthermore, there seems to be some effect for men in densely populated areas, where probably there is less scope for these activities. Another explanation is that, within couples, the unequal division of household chores by gender is not levelled-off at retirement.

[23] In this case, Dong's correction is zero. The reason is that the bias is due to the presence of a kink at eligibility, but using only S there is no change in the slope at $S = 0$.

[24] From graphical inspection it seems that there is a kink in house work at eligibility. I tried to exploit it instrumenting retirement with both D and $D \times S$, where the latter captures the kink (Dong, 2011; Card et al., 2009, see). The only exogenous regressor included in the equation of interest is S. Although $\widehat{\delta}$ becomes 135.5, with p-value 0.063, it is not stable to the inclusion of covariates, where it drops down to 90.7 (p-value 0.329).

[25] Results are obtained with no other covariates from X. However, adding the covariates not used in each split sample lead to similar conclusions (see the online Appendix A in Ciani, 2013).

[26] To get a magnitude of the increase in available time, I can use 2SLS with time spent on a paid job as dependent variable. Including $(1 - D) \times S$, $D \times S$, the estimated drop in working time is 2489 minutes/week (s.e. 77), almost equal to the average time spent by workers (2527 minutes/week). Results with a quadratic polynomial are quantitatively similar.

To provide some evidence, I also split the sample by marital status. The interesting result is that the change is almost zero for married men, while it is large for those who are not married (382 minutes/week, p-value 0.085).[27]

Results in columns (4)-(6) of Table 5 provide evidence on the presence of a discontinuity in Y at $s = 0$ for women, around 222 minutes/week using a linear polynomial without interactions. Although a second order polynomial shows no discontinuity, the information criteria indicate a preference for the simpler polynomials, for which both the G and the RESET tests fail to reject the null of correct specification.[28] Dong's correction does not lead to different conclusions.

Using the linear polynomial with interactions, and instrumenting R with D, the 2SLS estimate for δ (Table 7, column (2)) is 444 minutes/week, statistically significant at the 5% level, while coefficients associated with the polynomial terms are not statistically significant.[29] Including only S as additional regressor, $\widehat{\delta}$ is 435 minutes/week (Table 7, column (3)). Compared to the equivalent OLS regression, it is 32% smaller. It is interesting to note that a similar increase of 382 minutes/week was found for non married men.

While women with a high school or lower degree exhibit estimates for δ similar to that obtained in the main 2SLS regressions, the change in time spent on home production is negative for college graduates (Table 8). The magnitude is very large (289 minutes/week), but it is probably driven by the weakness of the instrument and by the small sample size. The effect is stronger in the North (569 minutes/week), though still relevant in the Center and the South (347 and 204 minutes/week respectively). 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 category, the increase for public sector employees (349 minutes/week) is smaller than for other categories, probably because their contracts already allow them to take paid and unpaid days off if they have family needs, such as an elderly parent with impaired health. Differently from men, married women still show an increase, around 329 minutes/week, though this is smaller than not married (936 minutes/week).

Results might be driven by the choice of the window size. I checked how they change when it is decreased, using 2SLS regressions including $(1 - D) \times S$, $D \times S$ as covariates,

[27] It includes never married, widowed and separated/divorced.

[28] A very similar estimate (211 minutes/week, p-value 0.018) is obtained by a regression of Y on D, S and S^2, with no interactions as in Battistin et al. (2009).

[29] It might be that using a linear polynomial we are confounding a kink with a jump. An alternative would be to use a quadratic specification (including S and S^2), and instrument R with D and $D \times S$, where the latter picks up the kink. The estimated δ would be 470.2 (s.e. 178.8), and it is robust to the inclusion of covariates.

and using D as an instrument for R. The online Appendix A (Ciani, 2013) includes a graph that depicts the different results. The estimates for men oscillate around zero and they are never statistically significant at the 10% level. When $|S| \leq 2$, the first stage F statistic becomes too small to reject the null of weak instruments.

For women, going down to $|S| \leq 5$ I find that $\widehat{\delta}$ is quite stable. However, the 95% confidence interval becomes larger and includes zero. 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 4 points are probably not enough to obtain precise estimates of the linear fit with interactions. Another is that, given that measurement error smoothes down the discontinuity in retirement, I need other points away from $S = 0$ to partially correct for it. Nevertheless, I tried to exploit only the data point close to eligibility, limiting the sample to $S \in \{-1, 1\}$ and using a simple Wald estimator.[30] In this case $\widehat{\delta}$ is equal to 423 minutes/week, very similar to the main results, but much less precise (s.e. 366).

5.4 Discontinuities in other covariates

One way to check the plausibility of the continuity assumption (A2) is to inspect whether some baseline characteristics exhibits discontinuities at eligibility. I focus on three sets of variables:

1. Geographical dummies for area of residence and population density (which are used for sample stratification);
2. Dummies for highest educational achievement;
3. Variables used to build S.

Geographical dummies are relatively smooth (Table 9, for graphs see Appendix B in Ciani, 2013). We observe an increase at eligibility in the proportion of men residing in the Centre (p-value 0.052) and in the proportion of women in densely populated area (p-value 0.042). However, a test for joint significance of all the discontinuities in geographical dummies fails to reject the null at 5% level for both genders.[31]

Educational dummies are fairly smooth 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

[30] In other words, I used 2SLS instrumenting R with D without adding any other covariate.

[31] The same applies if we separately test the joint significance of area dummies and of population density dummies

Italy university graduates are allowed to pay-back social contributions to cover the years of higher education and become eligible 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, that 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. Firstly, by construction $S = 0$ does not include a single cohort: the proportion of cohort 1949 at $S = -1$ is 0.725, quite close to the proportion at $S = 1$ (0.621).[32] Secondly, 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.

It must be added that, although both discontinuities found for men are statistically significant at the 5% level, the joint test for all educational dummies has a p-value 0.064. To further inspect the change in overall educational level, I calculated the total years of schooling by attributing the official lenght to each degree. This allows to account for some shorter vocational training degrees that are included in "high school" dummy.[33] As shown in Table 9, there is no evidence of a discontinuity for both gender. I also calculated the difference between age 6 and the age at which the individual has taken 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 exibith 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.

Among variables used to build S, age, years of contribution and age at retirement are fairly smooth. Differently, for men we observe a decrease at eligibility in the proportion of public employees, compensated by an increase in self-employed. This is related to an increase in years spent in paid job, which makes sense given that some self-employed individuals may have not contributed for some years to the retirement scheme, because they were included in the national insurance only at the end of the fifties.[34] For women we observe a decrease in private sector employees at eligibility,

[32] If I run a regression of $\mathbf{1}[cohort \geq 1949]$ on D, $(1 - D) \times S$ and $D \times S$ I estimate a drop in the proportion by 0.2979 (s.e. 0.0293). However, in this case using a linear polynomial is clearly not the best choice, because the indicator $\mathbf{1}[cohort \geq 1949]$ rapidly goes to zero after eligibility. With a quadratic the discontinuity is 0.0017 (s.e. 0.0522).

[33] Although they were usually 3 years long, some lasted only 2 years. I cannot distinguish them, but choosing either one or the other lenght does not affect the conclusions.

[34] Farmers from 1957, craftmans and small enterpreneurs from 1959, self-employed working in trade and retail from 1966 (Intorcia, 2011, p. 13). In the sample, 44.3% of the self-employed started working before 1966, and 9.4% before 1957.

though statistically significant only at the 10%, mostly compensated by an increase in self-employed (not statistically significant). We also observe an increase in years spent in a paid job, though largely imprecise (p-value 0.091), and an increase in the age at first job of around one year, statistically significant at the 5%.

To summarize, there seem to be discontinuities mainly related to some of the variables that enter in the definition of eligibility. Having excluded the possibility of sorting related to the educational qualification, 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.

One possible solution is to state all assumptions (A1)-(A3) conditional on the different covariates and employ the RDD on cells defined by employment category. Although this is not feasible given the sample size, I have already discussed how estimates different when the sample is split according to different characteristics (taking one variable at a time).

Another solution would be to adopt a parametric framework, where the counterfactual $\epsilon_i \equiv Y_{0i}$ depends linearly on these additional variables X, which therefore enter all regressions as covariates.[35] To understand how the introduction of X affects the estimates I obtained the fitted values for R and Y from a regression on dummies for education, geographical area and employment category, plus age, years of social contribution, years spent in paid work and age at which the respondent began the first regular job.[36] For men, fitted values for retirement exibith a drop at eligibility, though not significant at the 5% level, while fitted hours of domestic work show a small drop, estimated around around 18 minutes/week per men (p-value 0.001) using a linear polynomial (see the online Appendix B in Ciani, 2013, for the plot of fitted values on S). Also for women there is a drop in fitted retirement probability, not significant at the 5% level, and a drop of around 54 minutes/week in Y (p-value 0.033). Therefore I expect 2SLS regressions excluding these variables to underestimate the true effect.

Indeed, when using covariates, estimates for the discontinuity in retirement are basically unchanged for men, while there are slightly smaller, but still statistically significant for women. For both men and women the estimated discontinuity in Y is

[35]Formally, assumption (A2) becomes $E[\epsilon_i|X_i, S_i] = X_i\alpha + h_Y(S_i)$, with h_Y continuous at $s = 0$. Frölich (2007) proposed a strategy that accounts for covariates non-parametrically.

[36]Age at retirement is not included because it is a non linear function of the other variables (see the online Appendix A in Ciani, 2013). 2SLS results are basically unchanged when adding it.

larger. For men, the highest estimate for δ is 89 minutes/week (Table 6, column (4)), but not statistically significant at the 5% level and still far from the OLS results. For women, the estimates are larger than the results without X: using a linear polynomial with interactions the results is 528 minutes/week, while it is 483 minutes/week including only S.

Finally, the discontinuities in covariates can be due to decision not to use sample weights. For men, using them I still find the discontinuities in geographical area and educational dummies, while those for employment category have similar size but are not statistically significant (full results available on request). The estimated jump in the proportion of retirees at eligibility is smaller, while the increase in domestic work is larger. 2SLS estimates for δ are therefore larger, with a maximum of 139 minutes/week using covariates, but never statistically significant. For women the discontinuities in baseline covariates become all statistically not significant when using weights. The estimated discontinuity in retirement is smaller, and statistically significant only at the 5% level using a quadratic polynomial, while the estimated increase in Y is almost unchanged without covariates. Estimates for δ are quite in line with those presented in the main text and more similar to the results adding covariates. In a nutshell, overall conclusions are confirmed using sample weights.

5.5 Using age as a running variable

In 2007, 80% of women who went into retirement exploited the rules for the NRA path (*pensione di vecchiaia*). In most of the cases, they required at least 60 years of age and 15 of social contributions, although for younger workers the latter requirement was increased up to 20 years.[37] One possibility is to ignore all contributory requirements and focus only on the discontinuity at age 60, although this threshold does not correspond to a precise eligibility condition. In this way, I can avoid measurement error in S and I also reduce the influence of rounding, because age is available in quarters.[38]

Results and figures are reported in the online Appendix C (Ciani, 2013). For retirement, graphical analysis (figure 15 in the online Appendix C) clearly indicates a jump at age 60. The Akaike and Bayesian criteria indicate a preference for the quadratic regression, which passes the RESET and the G test (Table 31 in the online Appendix C Ciani, 2013). The estimated discontinuity is 0.170 (s.e. 0.046), not far from $\widehat{\gamma_D}$ obtained using S as running variable.

[37] To be precise, the requirement was increased for younger workers who started paying social contributions before 1995, while it was decreased to 5 years to those who started later.

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

The evidence for house work is less clear (figure 15 in the online Appendix C Ciani, 2013). The figure with age in quarters has a large dispersion, while if I aggregate age at intervals of one year I observe a jump at age 60 by around 200 minutes/week. In both cases, fitted linear polynomials predict a similar discontinuity (Table 32 in the online Appendix C Ciani, 2013). Using quarters, the point estimate is 179.5 (p-value 0.018). The resulting 2SLS estimate for δ is 503.8 minutes/week (p-value 0.016, see Table 33).

Differently, 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 shows already a large increase from 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 (see Card et al., 2009; Dong, 2011). In the online Appendix C (Table 36, Ciani, 2013) I also show regressions using $(age - 50)$ as running variable, and using the kinks at 57 and 60 together with the jump at 60 as instruments for R. The point estimate is 660.0 (s.e. 179.0), quite large and more similar to OLS results.

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 eligible. Differently, 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 some discontinuities in baseline covariates at age 60, which we did not find at the time of eligibility (see Tables 34 and 35 in the online Appendix C Ciani, 2013). In particular, a linear polynomial predicts an increase in the proportion of private employees and a decrease in public employees. Furthermore there is a drop in the education level and an increase in the proportion living in the North, with a reduction in the South. 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 418.6 minutes/week (s.e. 215.6), very similar to my main result using S, though significant only at the 10% level.

5.6 Changes in health and household composition

Other changes caused by retirement may have an off-setting effect on the increase in home production. Coe and Zamarro (2011), using data from the Survey of Health, Ageing and Retirement in Europe (SHARE), found evidence of an health preserving effect of retirement for men, with a 35% decrease in the probability of reporting fair, poor or very poor health. In my dataset, I used 2SLS with, as a dependent variable, a dummy for these answers to the general health question. In the linear specification including $(1-D) \times S$ and $D \times S$, the estimated effect is indeed negative, but much smaller (8.4%) and not statistically significant at conventional significance levels.[39] For women the estimate is also negative, but still smaller (6.7%), and not statistically significant. The reason for such a difference is not clear, and further research might try to extend the analysis using other waves.

Battistin et al. (2009) found a reduction in household size by 0.3, explained by adult children leaving the parental home. The estimated effect of retirement in SILC is actually positive for men, but very small (0.0600, s.e. 0.2046).[40] For women it is negative, with magnitude 0.0659 (s.e. 0.1922), still far from Battistin et al. (2009). There is also no evidence of a change in the probability of being married. Battistin et al. (2009) estimates 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.

6 Discussion

I used a RDD that exploits the discontinuity in retirement behaviour induced by the Italian Social Security System. The proportion of men leaving employment at eligibility is quite large, and most of them follow the path determined by a mixture of social contribution and age requirements. Nevertheless, 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 house work at eligibility. The resulting estimate for the causal effect of retirement on house work is between 430

[39] Results are reported in the online Appendix D (Ciani, 2013). Estimates with polynomials up to the 3rd order are never statistically significant.

[40] I refer here to 2SLS regressions including $(1-D) \times S$ and $D \times S$, reported in the online Appendix D (Ciani, 2013).

and 530 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 and 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.[41] 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, washing dishes, cleaning house, laundry)" (Szinovacz, 2000, p. 82). For France, Stancanelli and van Soest (2012) estimated that wives at retirement spend 2 hours 40 minutes per weekday more on house work, but they found that husbands also increased house work by around 3 hours per weekday. Furthermore, there is evidence for Germany (Schwerdt, 2005; Luhrmann, 2007) of an increase in housework at the retirement of the household heads, who are mostly men.[42] Lastly, Luengo-Prado and Sevilla (2012) 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 supported by the descriptive comparison with Time Use Data and by results from Stancanelli and van Soest (2012), who showed that the increase for men is concentrated in these activities. Furthermore, it must be noted that an increase around 225 minutes per/week is found for men residing in densely populated areas (34% of the sample), who are probably less likely to specialize in these "semi-leisure activities". Another explanation, confirmed by separate regressions, is that married men do not increase their participation in household chores or caregiving, leaving them to their wives.

[41] Activities included house cleaning, yard work/gardening, food preparation, home improvements, washing/ironing, shopping and finances related.

[42] Schwerdt (2005), analyzing data from the German SOEP (1994-2000), studied home production (errand, housework and yardwork) around a window of 2 years before and after retirement. Distinguishing low and high income replacement household heads, he estimated that the former spent around 714 minutes more per week on house work after retiring, while the latter 504 minutes. Similarly, Luhrmann (2007) found an increase of home production (cooking, preparing meals, paperwork and gardening) in households with a retired head by about 574 minutes per week, using German Times Use surveys 1991/92 and 2001/02.

Is the increase in house work sufficient to explain the change in consumption at retirement? Battistin et al. (2009) found that retirement of male household heads was associated with a drop in expenditure on non-durable goods by 9.8% and on food (including meals out) by 14.1%. Part of this change was explained by adult children leaving the household, and the effect of retirement on equivalized expenditure was 4.1% for non-durable goods and 8.4% on food, only the latter statistically significant at the 10% level. Miniaci et al. (2010) provided additional results, using a cohort analysis on data from the 1985-1996 Italian Survey of Family Budgets. They found that total consumption expenditure drops by 5.4%.[43] The fall could be explained by increased home production of meals and decreased purchase of work-related goods.

In order to understand the fraction of the drop that could be explained by increased house work, I need to recover the monetary value of home production. With respect to the expenditure on food, some calculations can be performed using the results by Aguiar and Hurst (2007). They estimated that in the U.S. a good's price decreases by 7 to 10 percent if one doubles shopping frequency. Although I do not have equivalent evidence from Italy, from the Time Use Survey I estimated that the time spent on shopping is 189 minutes/week for employed women aged 55-59, while 119 minutes/week for men with the same characteristics. Doubling the time input for women is compatible with the increase found in the RDD analysis. Therefore, if doubling the shopping frequency requires an equivalent increase in time, this could explain a drop in expenditure on consumption up to 10%. It must be added that Miniaci et al. (2010) actually found an increase in the food at home share on total expenditure, and a decrease in meals out. This would still be compatible with individuals spending less per unit purchase of food and keeping stable their actual calories intake. Furthermore, part of the women's increase in house work is likely to be devoted to other activities, in particular cooking, that could further explain the drop in expenditure.

However, the result from Battistin et al. (2009) refers to the retirement of male household heads. But for married men the point estimate is close to zero. On the one hand, the increase for them can be positive, but small in absolute value and therefore difficult to detect. Indeed, the estimate is quite imprecise, so that we cannot statistically reject the null hypothesis that it is compatible with doubling the shopping time. On the other hand, the finding of an economically significant increase for non married men suggests that retirement may induce only a small change for those who can rely on their partners' house work. In this case, my findings could explain the drop in consumption only for dual earners couples and as long as partners' match

[43] This could be a lower bound if individuals retired earlier mainly because of work-related costs (Miniaci et al., 2010, p. 265).

each other's retirement. Differently, Luengo-Prado and Sevilla (2012) found that, in Spain, the drop in food expenditure at the retirement of household heads could be explained by higher involvement of men in cooking and shopping. My results do not seem consistent with this explanation, but further research focusing on couples is warranted.

Tables and Figures

Table 1: Sample selection

	Male		Female	
	obs	% change	obs	% change
Raw data	21,522		23,611	
Worker or Retired	16,958	-21%	12,162	-48%
Non Proxy	13,979	-18%	10,856	-11%
Missing house work	13,437	-4%	10,546	-3%
$S \in [-10, 10]$	4,139	-69%	2,795	-73%
$S \neq 0$	3,970	-4%	2,700	-3%

Table 2: Participation and average minutes per day spent in house work, by gender and characteristics, SILC 2007 and TUS 2008-2009, all individuals aged 15+.

	Women						Men					
	Avg minutes/day			% participants			Avg minutes/day			% participants		
	SILC	TUS (A)	TUS (B)	SILC	TUS (A)	TUS (B)	SILC	TUS (A)	TUS (B)	SILC	TUS (A)	TUS (B)
TOTAL	238	285	271	0.93	0.93	0.92	64	96	70	0.63	0.68	0.64
AGE												
15-24	92	94	91	0.75	0.72	0.71	23	27	23	0.36	0.37	0.35
25-44	245	292	285	0.96	0.94	0.94	60	75	65	0.66	0.66	0.64
45-64	284	332	315	0.97	0.98	0.98	66	112	79	0.67	0.73	0.68
65+	219	306	278	0.89	0.93	0.92	80	150	97	0.63	0.84	0.76
SELF DEFINED EMPLOYMENT STATUS												
Employed	194	219	212	0.96	0.94	0.93	56	73	60	0.65	0.64	0.61
Unemployed	252	296	285	0.94	0.92	0.92	59	94	76	0.51	0.63	0.62
Housewife	338	389	370	0.97	0.97	0.97						
Student	62	70	68	0.70	0.68	0.67	25	23	20	0.40	0.39	0.37
Retired	236	325	296	0.92	0.96	0.95	85	168	108	0.67	0.87	0.78
Other	214	271	253	0.84	0.87	0.87	77	107	81	0.57	0.64	0.61

Note: estimated on original microdata using sample weights. In SILC I excluded missing values in house work. TUS (A) refers to total "family related" work, while column (B) contains only shopping, cooking, caring and "core" household work. 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.

Table 3: First stage OLS regressions for retirement status, SILC 2007 (selected sample)

Dep var R	(1) Men	(2) Men	(3) Men	(4) Women	(5) Women	(6) Women
D	0.270***	0.313***	0.398***	0.104	0.236***	0.471***
	(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***	0.206***	0.137***	0.048***
	(0.052)	(0.017)	(0.004)	(0.059)	(0.019)	(0.004)
$D\times S^2$	-0.016	-0.003**		-0.023**	-0.008***	
	(0.010)	(0.001)		(0.011)	(0.002)	
$D\times S^3$	0.001			0.001		
	(0.001)			(0.001)		
Observations	3970	3970	3970	2700	2700	2700
R^2	0.570	0.569	0.568	0.703	0.703	0.696
$H_0:\gamma_D=0$ (p-val)	0.001	0.000	0.000	0.286	0.000	0.000
- (p-val cluster)	0.004	0.001	0.000	0.168	0.003	0.000
Dong's $\widehat{\gamma_D}$	0.197	0.285	0.382	0.036	0.182	0.451
Dong's $\widehat{\gamma_D}$ (p-val)	0.055	0.000	0.000	0.767	0.005	0.000
RESET2	0.064	0.129	0.048	0.021	0.109	0.000
RESET23	0.138	0.055	0.139	0.070	0.244	0.000
G (p-value)	0.005	0.003	0.001	0.093	0.032	0.000
AIC	1949.554	1949.817	1954.425	380.124	382.628	441.741
BIC	1999.846	1987.536	1979.571	427.332	418.034	465.345

* p<.10 ** p<.05 *** p<.01. Robust standard error in brackets. γ_D is the coefficient for the discontinuity at eligibility. G (p-value) is Lee and Card (2008) statistic. Dong's refer to Dong (2012) correction. AIC is the Akaike criterion; BIC is the Bayesian criterion; RESET2 is the p-value from the RESET test adding square of fitted values, while RESET23 adding squares and cubes.

Table 4: OLS regressions for time spent on house work (in minutes per week), SILC 2007 (selected sample)

Dep var Y	(1) Men	(2) Men	(3) Men	(4) Women	(5) Women	(6) Women
R	286.7***	287.3***	249.4***	657.9***	638.2***	547.3***
	(28.6)	(27.5)	(21.7)	(71.5)	(68.6)	(43.4)
S	-3.9*	-4.0**		-10.2*	-8.6*	
	(2.1)	(1.9)		(5.3)	(5.0)	
S^2	0.0			-0.7		
	(0.3)			(0.6)		
Observations	3970	3970	3970	2700	2700	2700
R^2	0.043	0.043	0.042	0.065	0.065	0.064
RESET2	0.405	0.489		0.347	0.569	
RESET23	0.223	0.205		0.340	0.171	

* p<.10 ** p<.05 *** p<.01. Robust standard error in brackets. RESET2 is the p-value from the RESET test adding square of fitted values, while RESET23 adding squares and cubes.

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

Dep var Y	(1) Men	(2) Men	(3) Men	(4) Women	(5) Women	(6) Women
D	-20.802	21.280	30.915	16.112	209.189**	221.914**
	(64.606)	(38.201)	(37.069)	(149.527)	(89.015)	(86.830)
S			9.391***			13.881**
			(3.029)			(6.783)
$(1-D)\times S$	-0.456	3.916		21.163	8.689	
	(15.672)	(3.411)		(36.313)	(8.334)	
$(1-D)\times S^2$	-0.384			1.097		
	(1.341)			(3.139)		
$D\times S$	42.465*	17.065***		104.541**	21.533*	
	(23.501)	(5.474)		(50.732)	(11.424)	
$D\times S^2$	-2.307			-7.502*		
	(2.153)			(4.547)		
Observations	3970	3970	3970	2700	2700	2700
R^2	0.018	0.017	0.016	0.036	0.035	0.034
$H_0:\beta_D=0$ (p-val)	0.747	0.578	0.404	0.914	0.019	0.011
- (p-val cluster)	0.661	0.547	0.463	0.799	0.018	0.017
Dong's $\widehat{\beta_D}$	-42.583	14.705		-27.010	202.767	
Dong's $\widehat{\beta_D}$ (p-val)	0.546	0.708		0.867	0.026	
RESET2	0.799	0.278	0.069	0.539	0.117	0.414
RESET23	0.756	0.554	0.130	0.723	0.259	0.259
G (p-value)	0.166	0.200	0.101	0.760	0.649	0.654
AIC	61615.827	61613.328	61615.615	45048.293	45047.607	45046.465
BIC	61653.546	61638.474	61634.475	45083.699	45071.211	45064.168

* p<.10 ** p<.05 *** p<.01. Robust standard error in brackets. β_D is the coefficient for the discontinuity at eligibility. G (p-value) is Lee and Card (2008) statistic. Dong's refer to Dong (2012) correction. AIC is the Akaike criterion; BIC is the Bayesian criterion; RESET2 is the p-value from the RESET test adding square of fitted values, while RESET23 adding squares and cubes.

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

Dep var Y	(1) No X	(2) No X	(3) No X	(4) With X	(5) With X	(6) With X
R	-66.374	53.429	73.225	-96.849	81.312	89.194
	(208.601)	(95.124)	(86.878)	(220.787)	(95.816)	(93.195)
$(1-D)\times S$	0.909	3.339		22.190	11.839	
	(18.423)	(4.037)		(28.001)	(10.382)	
$(1-D)\times S^2$	-0.327			-0.034		
	(1.441)			(1.440)		
$D\times S$	47.462	14.745*		62.865	17.065	
	(36.742)	(8.809)		(42.160)	(10.595)	
$D\times S^2$	-2.499			-3.180		
	(2.603)			(2.722)		
S			7.606			14.668
			(4.946)			(9.743)
Observations	3970	3970	3970	3970	3970	3970
$H_0:\delta=0$ (p-val)	0.750	0.574	0.399	0.661	0.396	0.339
First Stage F	42.196	216.080	265.562	43.794	247.004	269.478

* p<.10 ** p<.05 *** p<.01. Robust standard error in brackets. R is instrumented by D. δ is the coefficient on R. Covariates X include a constant, age, 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.

Table 7: 2SLS regressions for time spent on house work (in minutes per week), women, SILC 2007 (selected sample)

Dep var Y	(1) No X	(2) No X	(3) No X	(4) With X	(5) With X	(6) With X
R	68.362	444.351**	434.811***	389.471	527.824***	483.067***
	(628.995)	(183.315)	(165.434)	(653.412)	(188.988)	(179.980)
$(1-D)\times S$	18.825	5.186		19.743	20.639	
	(51.406)	(9.113)		(74.186)	(21.139)	
$(1-D)\times S^2$	0.939			-0.709		
	(4.054)			(3.756)		
$D\times S$	95.166	0.271		43.092	-5.571	
	(126.991)	(18.234)		(143.451)	(24.246)	
$D\times S^2$	-6.950			-3.543		
	(8.670)			(8.735)		
S			3.428			8.237
			(10.308)			(21.288)
Observations	2700	2700	2700	2700	2700	2700
$H_0:\delta=0$ (p-val)	0.913	0.015	0.009	0.551	0.005	0.007
First Stage F	15.930	197.380	255.250	15.431	192.845	222.697

* p<.10 ** p<.05 *** p<.01. Robust standard error in brackets. R is instrumented by D. δ is the coefficient on R. Covariates X include a constant, age, 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.

Table 8: 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	72.183	0.579	138	2020	468.949	0.049	141	1353
	High school	28.098	0.830	101	1492	573.677	0.018	116	1006
	College	176.195	0.448	30	458	-288.646	0.62	16	341
By area:	North	148.417	0.176	157	1953	568.688	0.007	148	1369
	Centre	86.169	0.619	61	966	346.976	0.302	61	717
	South	-105.052	0.635	50	1051	203.989	0.598	51	614
By degree of urbanization:	Densely populated	225.247	0.067	108	1360	641.716	0.002	98	985
	Intermediate area	-28.386	0.855	97	1689	690.372	0.013	107	1109
	Thinly populated	-24.008	0.896	63	921	-197.694	0.608	53	606
By category:	Private employee	112.762	0.222	268	2080	482.389	0.028	207	1073
	Public employee	104.9	0.690	34	761	349.389	0.226	77	980
	Self-employed	-21.271	0.934	25	1129	569.424	0.182	33	647
By marital status:	Married	32.469	0.724	220	3249	329.174	0.060	248	1914
	Not married	381.917	0.085	49	721	935.733	0.029	32	786

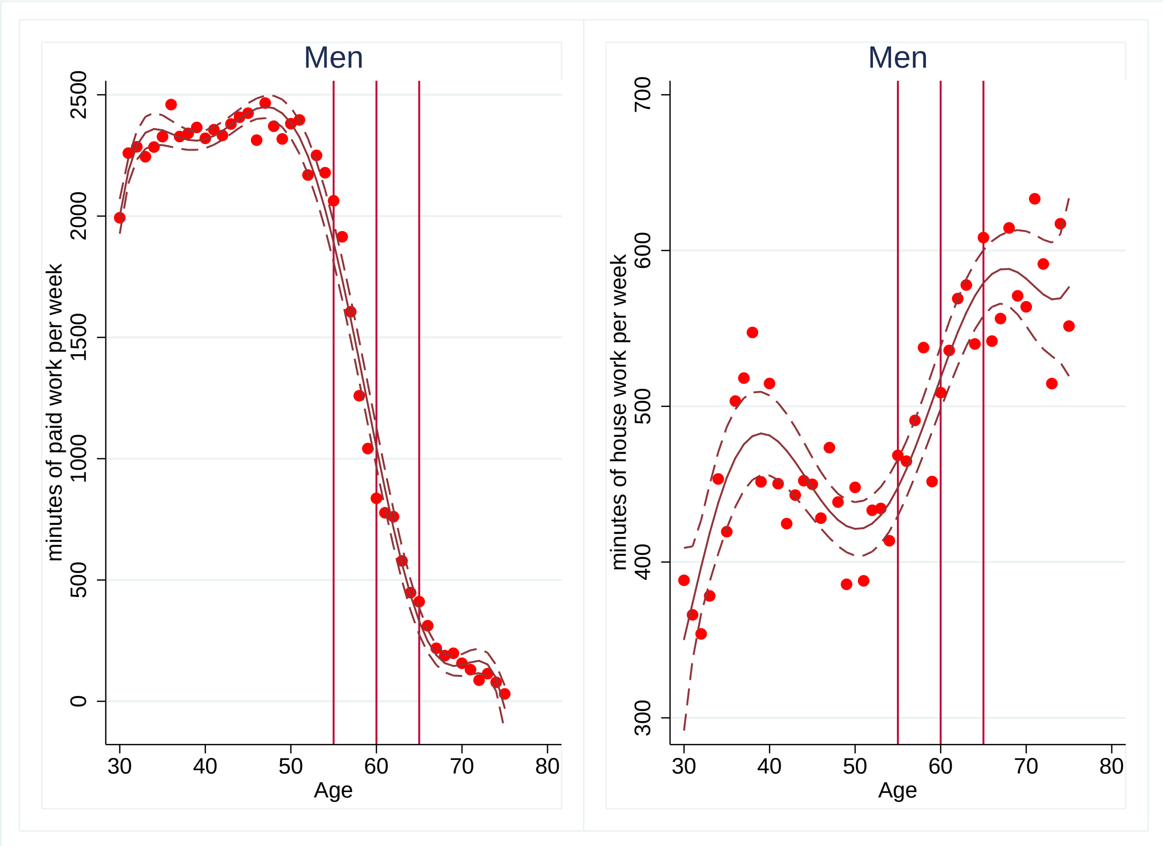
Note: all estimates include only S as additional covariate, while R is instrumented by D. The p-value and the first stage F are calculated using robust standard errors.

Table 9: Regressions for different socio-economic variables, SILC 2007 (selected sample)

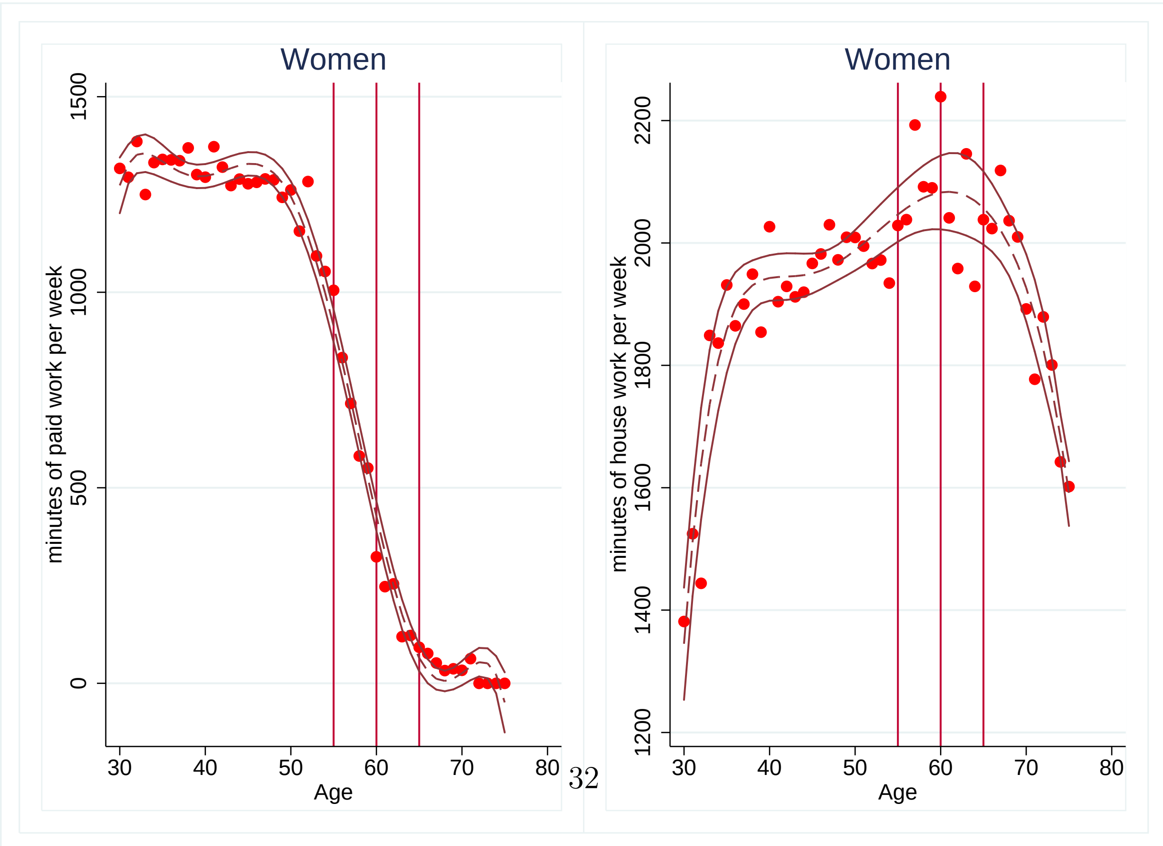
	Men			Women		
Dep. var.	$\widehat{\gamma}_D$	p-value	G (p-value)	$\widehat{\gamma}_D$	p-value	G (p-value)
Geographical dummies						
North	-0.046	0.196	0.249	-0.035	0.430	0.975
Centre	0.059*	0.052	0.508	0.013	0.743	0.060
South	-0.014	0.656	0.042	0.022	0.561	0.373
Densely pop area	-0.03	0.368	0.68	0.087**	0.042	0.284
Intermediate area	0.026	0.458	0.048	-0.068	0.120	0.790
Thinly pop area	0.004	0.881	0.229	-0.019	0.590	0.279
Test for joint significance		0.368			0.245	
Educational dummies						
College	-0.015	0.522	0.627	0.016	0.608	0.893
High school	0.084**	0.013	0.051	0.005	0.904	0.318
Middle school	-0.065**	0.041	0.048	0.004	0.921	0.970
Primary sch.	-0.004	0.889	0.460	-0.025	0.523	0.370
Test for joint significance		0.064			0.909	
Additional educational variables						
Years of schooling	0.282	0.318	0.769	0.508	0.170	0.828
Age highest edu - 6	-1.180	0.102	0.318	0.736	0.366	0.007
Test for joint significance		0.046			0.373	
Variables used in building S						
Age	-0.123	0.589	0.036	-0.131	0.516	0.670
Y. of contribution	0.175	0.592	0.284	-0.220	0.732	0.125
Age at retirement	0.064	0.828	0.064	-0.087	0.855	0.285
Age at first job	0.060	0.860	0.798	1.175**	0.032	0.725
Years spent in a paid job	0.976**	0.018	0.304	1.186*	0.091	0.325
Private employee	-0.013	0.704	0.020	-0.071*	0.091	0.485
Public employee	-0.055**	0.044	0.469	0.024	0.579	0.194
Self-employed	0.068**	0.031	0.005	0.048	0.212	0.210
Test for joint significance		0.000			0.000	

* p<.10 ** p<.05 *** p<.01. The regressions include $(1-D)\times S$ and $D\times S$. γ_D is the coefficient for the discontinuity at eligibility. 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.

Figure 1: Age profile of average minutes/week of paid work and house work, SILC 2007 (original sample, missing values in Y excluded). Age in years. Lines are fit from a 6th order polynomial in age (with a 95% c.i.)



(a) Men



(b) Women

Figure 2: Average minutes/week of house work by employment status (circles for workers, triangles for retirees) and age (in years), SILC 2007, only age-employment cells with at least 10 obs. Lines are fit from a 2nd order polynomial (with 95% c.i.).

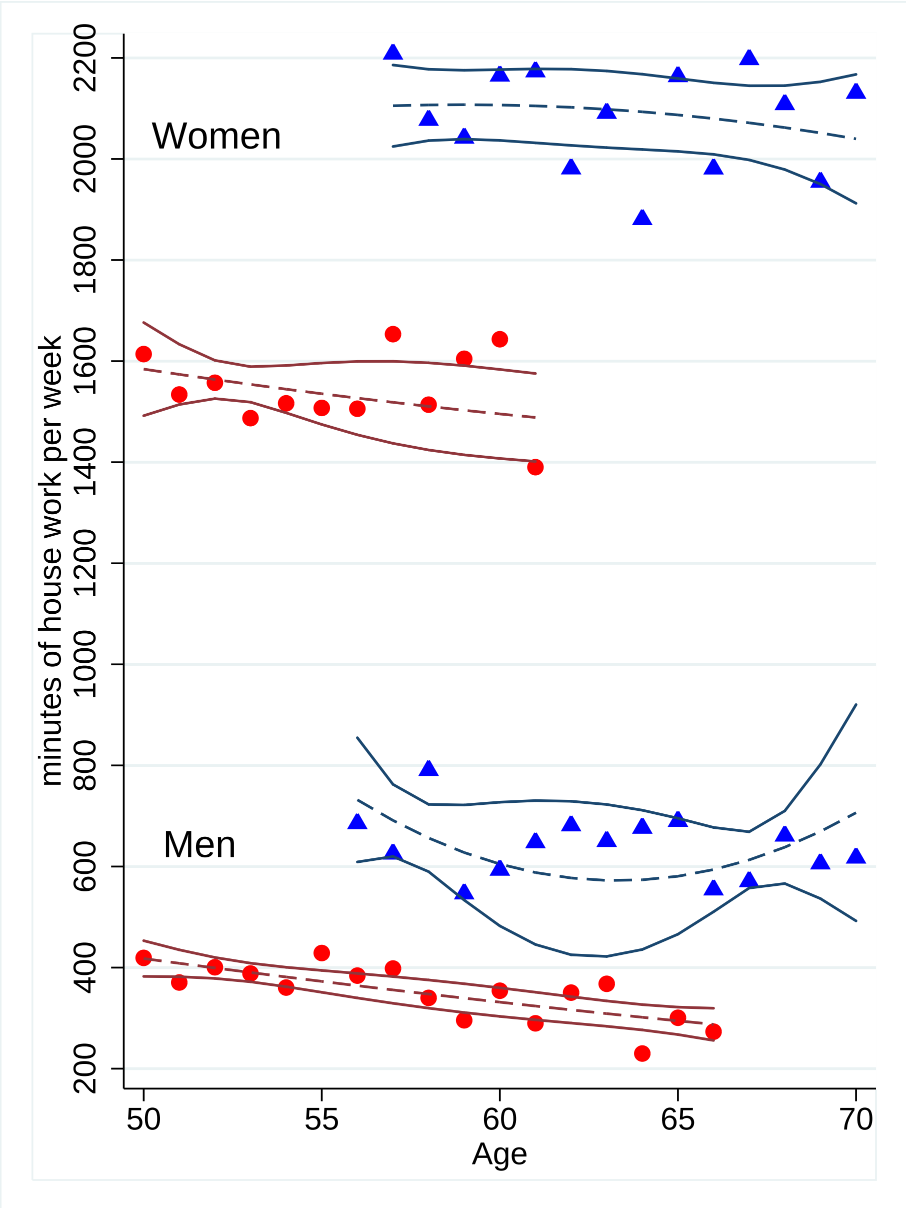
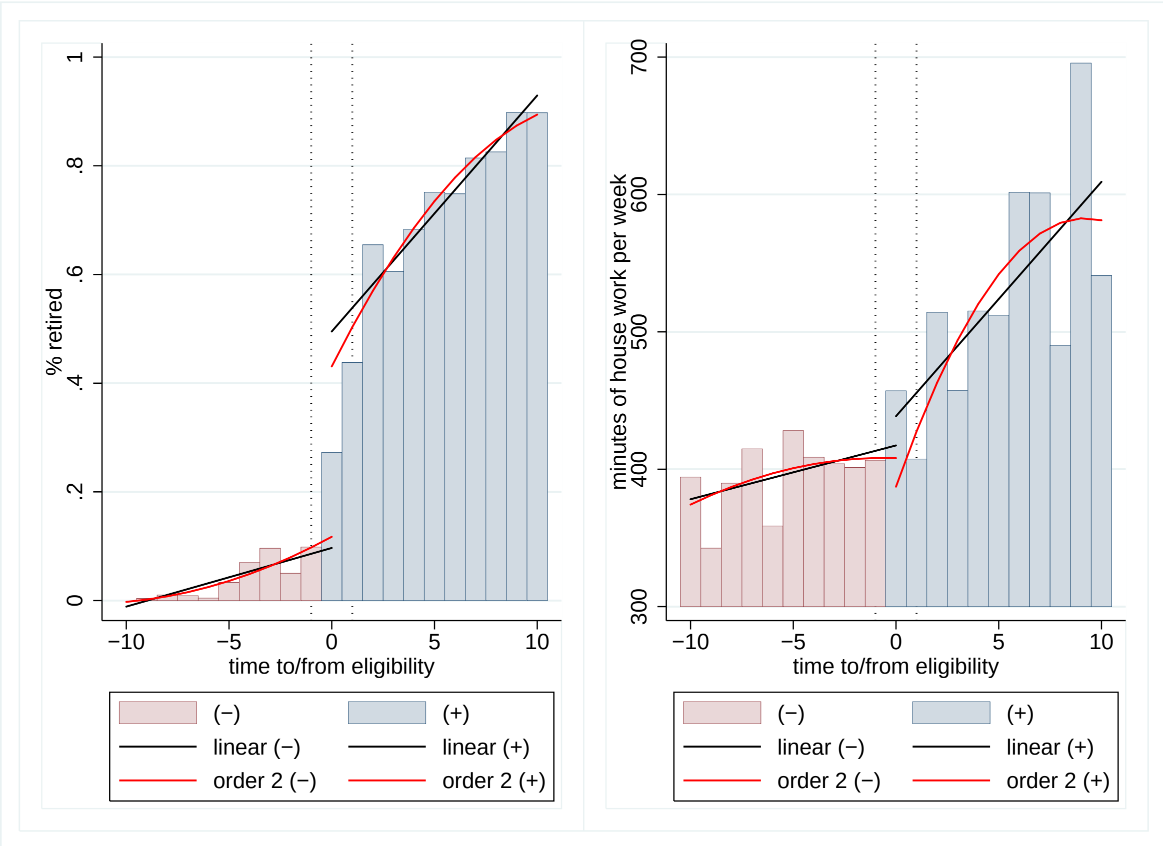
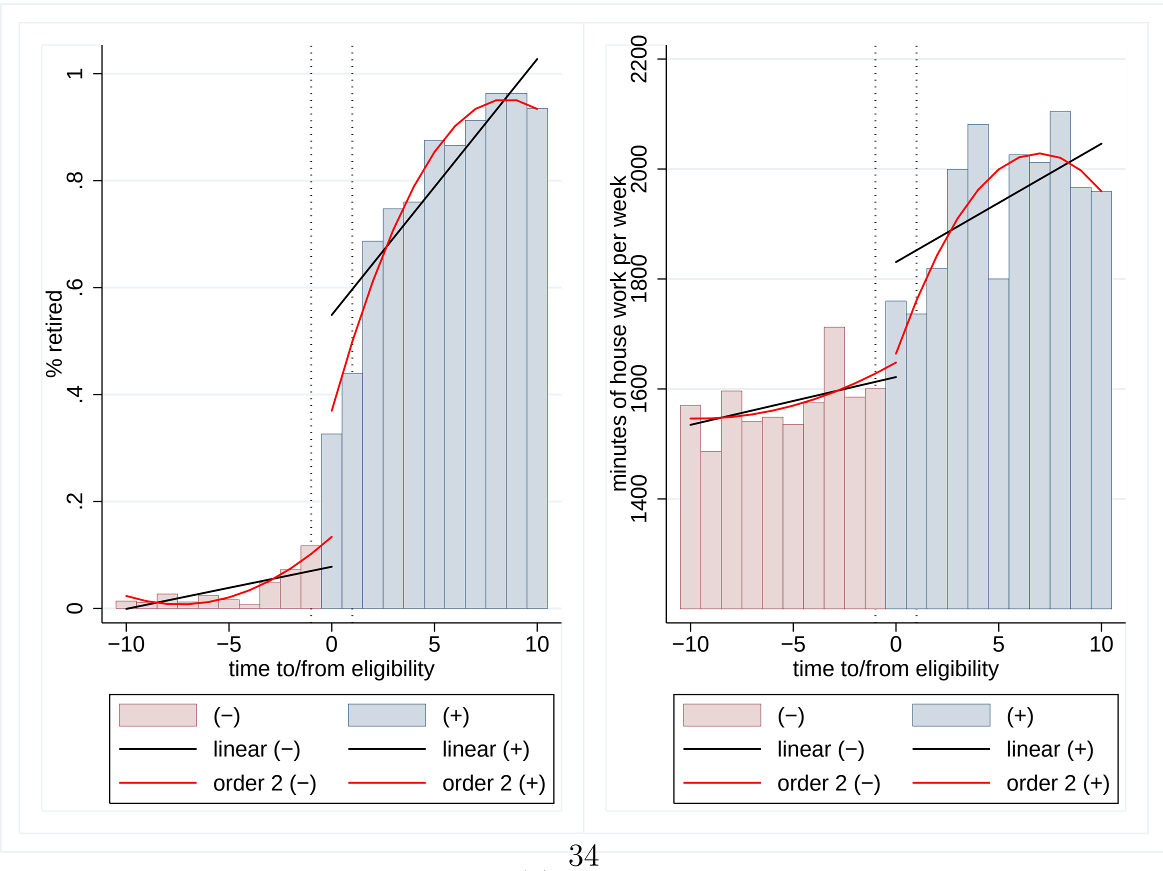


Figure 3: Retirement and house work with respect to S, SILC 2007 (selected sample).



(a) Men



(b) Women

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