Should I Care Or Should I Work? The Impact Of Working In Older Age On Caregiving

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ABSTRACT

While there is converging evidence that having informal carer responsibilities has an adverse effect of labour supply, far less is known about the effect of working longer on the informal supply of care to family and friends. In this paper, we explore whether prolonging working careers reduces the amount of help provided by women around the pensionable age (55-65 years old), which are a crucial source of informal caregiving in most advanced economies. Exploiting a major pension reform in the UK, we use eligibility to the State pension to overcome the endogeneity of labour supply decision. We find that working longer as result of not being eligible for State pension significantly reduces the time devoted to offer help outside the household. Our results suggest that policy makers should account for the unintended consequences on informal care supply when regulating labour markets and social and health services for older people.

1. INTRODUCTION

The demand of care by older Europeans is large and growing fast, due to an unprecedented demographic transition fuelled by declining mortality and fertility rates (OECD/EU, 2016). In OECD countries, family and friends are the most important source of care for people who require help with everyday tasks, living inside or outside of the caregiver household (OECD, 2017a). With tightening public expenditure budgets, increasing dependency ratios and socioeconomic changes in family contexts, the risk that the demand of care would fail to be met by effective, responsive and high-quality social protection is high and worrisome (AgeUK, 2017; European Commission, 2014). Another potential threat to care system may be that as an increasing proportion of older adults are in work, the supply of informal care declines sharply. Over the last decade, most OECD countries have introduced important work-incentive policies, with the aim of raising the employment rate of older workers, and particularly women, to reinforce the financial sustainability of pension systems (OECD, 2016). Little attention has been paid to whether these the incentives to work longer in older age may have adverse consequences on informal care provision: if labour market activities crowd out informal caregiving (e.g., increasing the opportunity costs of caring), new challenges for social protection would need to be addressed by policy makers. If, on the other hand, caregiving activities are inflexible and hard to combine with the extended time in paid work, time allocation for other activities could be reduced (e.g., leisure, home-tasks, exercise), while keeping the informal help provision constant (He and McHenry, 2016).

Evidence on the consequences of working longer on the supply of informal care is sparse. The majority of the literature that focus on informal care and work explores the causal relationship between informal care provision and employment status. There is converging evidence that having informal carer responsibilities makes people more likely to withdraw from the labour force, choose part-time employment and have lower wages (see, e.g., Schmitz and Westphal (2017), Van Houtven *et al.* (2013), Crespo and Mira (2014), Bolin *et al.* (2008b), Mentzakis *et al.* (2009), Michaud *et al.* (2010), and the review in Jacobs *et al.* (2014)).

However, far less is known about the potential effects of working longer on care provision. Some studies exploit longitudinal data to analyse how caregiving and employment interact. Using data from the British Household Panel Study (BHPS) Michaud *et al.* (2010) examine the pathways through which caregiving and employment interact. Their results suggest that moving to non-employment increases the probability of providing co-residential care by 0.9 p.p. in the following year and while this effect is 2.3 p.p. for extra-residential care. Using data from the same survey but for a longer time period, Carmichael *et al.* (2010) focus on people who, while not yet carers in a given year, will become so one year later and investigate the relationship between employment status and transitions into informal care. Their results suggest that people who are already in employment and particularly those with higher wages are less willing to undertake care. The main limitations of these two studies is that employment status is likely to be endogenous: unobservable individual and family characteristics may influence both labour market status and care provision. Failing to fully account for this may threaten the causal interpretation of the results. To account for the endogeneity of employment status, He and McHenry (2016) exploit state-level unemployment rate as an instrumental variable to identify the effect of working on the provision of informal care using SIPP data from the US. Their results indicate that working 10% more hours per week reduces the probability of providing any informal care by about 2 percentage points; moreover, the effect of formal employment on caregiving to household members is found to be stronger than for extra-household settings. Their results should be interpreted as the effect of working triggered by a change in unemployment rate for the working age population. However, no analysis has so far analysed the effect of working longer amongst older adults.

In this paper, we aim to fill this gap by exploiting a pension reform in the UK to estimate the causal effect of increasing labour supply on informal care provision, using an instrumental variable approach. We use eligibility to the State Pension Age (SPA) as an instrument for labour supply. The State Pension Age (SPA) in the UK is the minimum age at which people can claim the Basic State Pension, which provides an almost-flat minimum level of retirement income[1]. SPA for women was gradually increased from 60 to 65 between 2010 and 2018, with the exact SPA for women born after March 1950 depending (non-linearly) on their date of birth. The pension reform generates variation on eligibility for women of the same age. This allows us to flexibly control for age in our models which is key for identifying for credibly using pension eligibility as an instrument. Using data from Understanding Society collected between 2009 and 2016, we focus on women aged 55 to 65, who are likely to be affected by a change in SPA. This is a particularly relevant population for the informal care supply, since the proportion of caregivers is particularly high amongst women (Rodrigues, 2013), and especially among older working age adults (AgeUK, 2017; Centre for Policy on Ageing, 2014; Verbakel *et al.*, 2017). Thus, our analysis will be particularly relevant for policymakers involved in the regulation of labour market, health and social care for older workers. We differentiate our analysis by care setting, looking separately at help given in and outside the caregiver's household, as the two types of assistance (and the care recipients) are likely to differ (Carmichael *et al.*, 2010; Michaud *et al.*, 2010). We focus on both the incidence and intensity of care giving by using a two-part model, as is common in the informal care literature (Carrino *et al.*, 2018b; Mentzakis *et al.*, 2009). We use log of working hours as our main measure of labour supply, as being over SPA may not necessarily lead everybody to retire but could instead trigger a reduction in the number of hours. Because our approach relies on a specific source of variation in labour supply, i.e., a change in pension rules which increased incentives to stay at work longer, our analysis differentiates from other papers which exploit other sources of variation, e.g., unemployment rates, and consequently focus on different (usually, younger) population and compliers.

We find that working longer as result of not being eligible for State pension has no effect on in-household care giving but reduce informal care provided outside the household. More precisely, we find an effect on the intensive margins (the number of care hours) but not on the extensive margins (the probability to provide care). We find that a 1% increase in work-hours leads to a 0.3% reduction in care-hours: given that the average amount of weekly working hours if 15.3, and that non-residential caregivers provide an average of 9 weekly care-hours, an increase of 1 work-hour induced by being below the State Pension Age reduces caregiving by 0.177 hours (10 minutes). Using two occupational indices of physical and psycho-social burden built by Kroll and Lampert (2011) we show that the adverse effects of working are stronger for those in physically and mentally demanding jobs.

Our study makes several important contributions.
While literature has studied how care may affect employment, we contribute to the emerging literature looking into the effect of labour supply on informal care and providing reliable evidence that causality goes both ways, as working longer is found to have has a negative effect on the provision of informal care.
Second, we specifically look at the implication of working longer for the help provided by the female population aged 55-65, which constitutes a crucial backbone of the caregiving in the UK and Western world. Moreover, we are the first to provide evidence of the impact of policies extending working

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[1] The specific amount depends National Insurance contribution-years. The full Basic State Pension amounted to £5,077.8 for a single individual and £8,119.8 for a couple; corresponding figures in 2016 were £6,029.4 and £9,643.

lives on situations of work-care conflicts (Burr and Colley, 2017), by exploiting a major pension reform in the UK. More generally, our results suggest that policies that aim to increase labour supply of older adults may have unintended negative consequences on the supply of informal care, especially for people engaged in demanding occupations. A reduction of informal care provision may be compensated by an increase in the use of formal care, which may involve significant public and private costs. However, existing evidence suggest that formal and informal care are not perfect substitute[2]. Therefore, these policies may have detrimental effects on the wellbeing not only of the carers but also on the unmet needs (thus, on the welfare) of the recipients (AgeUK, 2017). Further research should focus on understanding the full welfare effects of pension reforms.
Finally, we are able to study the heterogeneous impact of working longer by type of care, differentiating between residential and extra-residential settings.

This paper is organised as follows: Section 2 describes the data, the empirical model and the instrumental variable approach; while section 3 discusses the results and provides sensitivity analyses. Section 4 concludes with a discussion of the implications of our findings for both research and policy.

2. METHODS

2.1 DATA, SAMPLE SELECTION AND DESCRIPTIVE EVIDENCE

This analysis exploits data from the *Understanding Society* survey, collected annually between 2009 and 2016 (seven waves) on a sample of household members aged 16+ in the UK (see Lynn (2009) and Knies (2016) for technical details). Our main selection includes women aged between 55 and 65, i.e. close to the statutory pension age which ranges between 60 and 63 years old in the selected time-frame. Although an all-female sample is not uncommon in this literature - women being the majority of the caring population (e.g., He and McHenry (2016); Michaud *et al.* (2010)) – our sample choice is driven by empirical reasons, namely, the availability of a credible instrumental variable for the employment status, which is not available for men (see Section 2.2). Similarly, we exclude respondents who never worked, as they would not be directly affected by our instrument, namely, an exogenous variation in pension eligibility. After dropping entries with only proxy interviews or missing information on any variable of interest, our final sample consists of an unbalanced panel of 24,297 observations (6,904 women).
Labour market status is defined as either “in paid work”, “unemployed”, “retired”, “looking after family or home” or “long-term sick/disabled”. We define as working those whose have “paid work” as their main activity (Bonsang *et al.*, 2012; Hessel, 2016; Mazzonna and Peracchi, 2017). The survey also includes information on working hours usually worked per week, for main and second occupation. As the overall working hours distribution is strongly skewed on the right, and sometimes exceed plausible values, we trim this variable at the 99th percentile. We additionally build an “intensive” definition of working, with a dichotomous variable identifying those working 20 or more hours a week (Michaud *et al.*, 2010).

[2] Numerous studies have investigated the relationship between informal-care and formal-care use, with most evidence showing a complementary relationship for high-skills care (Balia and Brau, 2013; Bonsang, 2009; Carrino et al., 2018b), and a substitution relationship for low-skill care (Bolin et al., 2008a; Bonsang, 2009; Stabile et al., 2006; Van Houtven and Norton, 2004).

We identify as informal caregivers those reporting to "look-after or give special help" to someone who is "sick, disabled or elderly" (Carmichael and Ercolani, 2016). We can further disaggregate between in-household (residential) and extra-household (non-residential) care settings, depending on whether the caregiver co-resides with the recipient. Respondents are asked to report the overall weekly number of care-hours, in brackets, and regardless of care-type (Figure 1). Due to the questionnaire design, it is not possible to disentangle care-hours by care-type, for women who provide both in- and extra-household care. However, given that this only affects 1% of the sample, and that care intensity is significantly higher for in-household caregiving, we classify such respondents as in-household caregivers (see, e.g, Michaud *et al.* (2010)). Sensitivity tests confirm this assumption does not affect our results.[3] We additionally build dichotomous variables for providing, respectively, 5 or more, or 20 or more care-hours per week.

Socioeconomic status, derived from the hierarchical National Statistics Socio-economic Classification (NS-SEC), distinguishes between managerial/administrative/professional, intermediate (including small employers and own-account workers), and routine/manual-workers respondents.[4]

The survey includes several additional respondents' characteristics such as living arrangements, the year at which respondents left their last job, number of children, and educational attainment (categorised as A-level or higher, GCSE level, no education).

Descriptive evidence

Table 1, column (a), provides descriptive statistics on the main variables included in our analysis. The average age is 60, the majority of the population is working, living with a partner, not eligible for the State Pension, and owning their own house. Nearly half of the population has one child, and an A-level education certificate or higher. Thirty-eight percent of the sample is classified among the lowest Socio-Economic Status (routine). With respect to the non-working population, those in paid-work are slightly younger, less likely to belong to the routine SES, more likely to be childless and with higher education (columns b and c).

Table 1, Descriptive statistics for the whole sample and by employment status

	(a)	*(b)*	*(c)*
	Whole population	*Non-working pop.*	*Working pop.*

[3] Two categories, namely, "varies under 20 hours" and "varies 20 hours or more", which are selected by 6% of caregivers, are purposely vague, which makes it harder to compare them with the remaining brackets. When dropping respondents who selected these categories, our results are confirmed.

[4] The NS-SEC coding is based on a cross-reference between individuals' current or last occupational category (based on the Standard Occupational Classification, SOC2000), firm size, and employment status (employer, self-employed or employee).

	Mean	SD	Mean	SD	Mean	SD
Age	59.84	3.166	61.15	2.997	58.75	2.875
working	0.544	0.498	0	0	1	0
retired	0.326	0.469	0.715	0.451	0	0
Below State Pension Age	0.592	0.492	0.398	0.489	0.754	0.431
Living with a partner	0.703	0.456	0.686	0.463	0.716	0.450
Widowed	0.242	0.428	0.255	0.436	0.231	0.421
Single	0.055	0.228	0.058	0.234	0.052	0.223
No children	0.251	0.433	0.202	0.402	0.291	0.454
One child	0.478	0.500	0.502	0.500	0.458	0.498
Two children or more	0.271	0.445	0.295	0.456	0.251	0.434
No education	0.176	0.381	0.258	0.438	0.108	0.310
Education at GCSE level	0.338	0.473	0.342	0.474	0.335	0.472
Education at A-level or higher	0.485	0.500	0.399	0.490	0.557	0.497
SES routine	0.382	0.486	0.430	0.495	0.342	0.475
SES intermediate	0.270	0.444	0.254	0.435	0.284	0.451
SES managerial	0.347	0.476	0.316	0.465	0.373	0.484
house owned	0.807	0.394	0.761	0.427	0.846	0.361
Provides informal care	0.288	0.453	0.288	0.453	0.288	0.453
- inside household	0.081	0.273	0.104	0.305	0.062	0.242
- outside household	0.206	0.405	0.185	0.388	0.225	0.417
Provides 5+ hours informal care	0.185	0.388	0.200	0.400	0.172	0.377
- in household	0.068	0.253	0.090	0.287	0.050	0.219
- outside household	0.116	0.321	0.110	0.313	0.122	0.327
N	24,257		11,251		13,206	

Note: the sample includes women aged 55-65 between 2009 and 2016, having been engaged in paid work in their life, excluding proxy respondents. The SES classification follows the National Statistics SEC-3 taxonomy.

Around 28% of the sample provides any type of informal care, less than one-third of which is performed in-household (8%). Care-intensity appears to be higher for residential settings: 6.8% of the sample provide in-household care for more than 5 hours a week (84% of in-household caregivers), while 11.6% provide non-residential care for more than 5 hours (56% of non-residential caregivers). Details on care-hours disaggregation can be further appreciated from Figure 1: extra-household care is less intensive and highly skewed, with around 42% providing less than 5 hours per week. Furthermore, among in-household caregivers, 50% provide between 1 and 34 weekly hours, almost uniformly distributed.

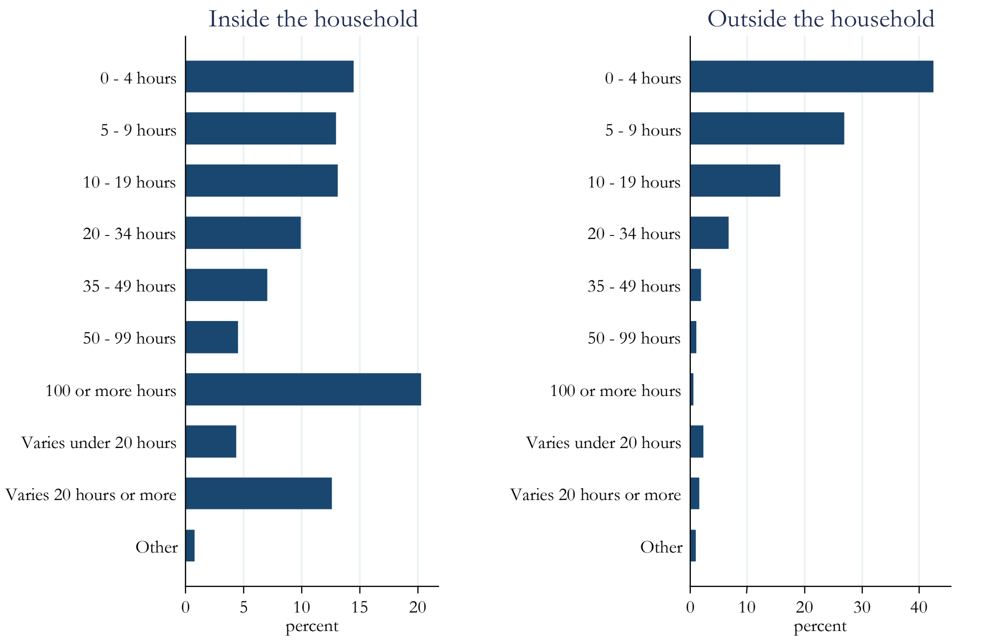
Although the overall prevalence of caregiving is almost identical across employment status, Table 1 suggests that women in paid work provide slightly less in-household care than non-working respondents, yet more extra-household care. Care-intensity does not seem to vary much by employment-status: e.g., out of 10.4% non-working respondents who provide residential care, 9% do so for 5+ hours (86%); similarly, out of 6.2% working women involved in residential caregiving, 5% provide 5+ hours care (80%).

In our sample 4,934 respondents provide care outside of the household, whilst 1,947 provide in-household care. The care recipients tend to be different for in- and out-household care. Over two third (68.8%) of respondents who provide care out of the household look after their parents, 17.3% provide care for a friend or neighbour and 16.5% for another relative. In contrast, nearly two-third (65.4%) of those who provide in-household care look after their partner, 17.1% after their parents and 13.5% care for their child.

We now turn to our econometric model to identify the causal impact of working on caregiving provision.

Figure 1 here

Figure 1 Hours spent providing care, inside and outside the household



2.2 ECONOMETRIC SPECIFICATION

We investigate the causal effect of being in paid work for longer spells on the supply of informal home-care. Similarly to previous studies on informal-care provision (Carrino *et al.*, 2018b; Mentzakis *et al.*, 2009), we adopt a two-part model (2PM) and model the informal-care provision separately at the extensive and the intensive margin: first, the individual decides whether to provide care or not; then, conditional upon caregiving, the amount is determined. This model is largely adopted in the health economics literature to estimate actual outcomes, i.e. fully-observed variables. In our context, "zero" caregiving hours indicate that no care was provided, i.e., a corner solution, as care-hours cannot be negative (see, e.g., Duan *et al.* (1983), Van Houtven and Norton (2004) and Bonsang (2009)).

In the first part of the 2PM we estimate a probit regression for the probability of being a caregiver for individual *i* at time *t*:

$$(1) \qquad Pr(ICH_{it} > 0) = \gamma_0 + \gamma_1 WH_{it} + \sum\nolimits_a \delta_a (A_i = a) + \sum\nolimits_y \vartheta_y (Y_t = y) + X_{it}\beta + \varepsilon_{it}$$

where *ICH* is the number of weekly care-hours provided. The type of care provided is likely to be different depending on living arrangements (i.e., extra-residential vs co-residential care), and this can in turn affect the interaction between care and employment. Recent literature has highlighted that caregiving has a stronger negative effect on employment when the care provider is co-residing, perhaps due to a stronger pressure on co-residing caregivers to care for in-household members rather than for extra-households members. Indeed, care-settings (in- or extra-household) highly characterise both care-intensity (higher in co-residence) and care-recipients (mostly spouses in co-residing care, mostly parents in extra-household care (Carmichael *et al.*, 2010; Heitmueller, 2007; Mentzakis *et al.*, 2009; Michaud *et al.*, 2010). Thus, we examine heterogeneous effects of employment on caregiving by characteristics of the care recipient, i.e., estimating separate regressions for in- and extra-household caregiving.

The main regressor *WH* is a continuous variable for working hours, transformed using a hyperbolic sine transformation which provides a coefficient interpretation equivalent to the logarithmic transformation, while being able to deal with zero hours corresponding to not working (Burbidge *et al.*, 1988). In an alternative specification we employ a dichotomous variable for being in paid work.. We flexibly control for age and time effects by adding fixed effects for age in years (A) and for interview year (Y).[5] The matrix X includes additional individual characteristics as fixed effects for living arrangements and marital status (single/never married, living in couple, widowed/divorced/separated, married), country, number of children (zero, one/two, three or more children). As the opportunity cost of caregiving may differ by socioeconomic status, we control for highest educational attainment, NS-SEC socioeconomic classification (routine, intermediate, managerial), and for a binary home-ownership information. Finally, albeit it is tempting to account for respondent's health status as explanatory variables, there is large and growing evidence emphasising its endogenous relationship with caregiving. First, unobservable (thus, omitted) variables like personality may affect both respondents' health and her care decisions (Bom *et al.*, 2018); second, caregiving may impact the provider's physical and mental health, both in the short and in the long-run (see Bom *et al.* (2018), Zwart *et al.* (2017) and Carrino *et al.* (2018a) for a review). Therefore, we exclude health from the main set of regressors, although we discuss its potential mediator in section 3.1.3.

The second part of the model is estimated on the conditional sample of caregivers where the dependent variable is interval log-hours of informal-care, with same regressors as in (1).[6] Because in Understanding Society the weekly care hours is interval coded (see Figure 1), we estimate an interval regression model (Bettin and Lucchetti, 2012; Stewart, 1983):

$$\log(ICH_{it}^{*}) = \gamma_0 + \gamma_1 WH_{it} + \gamma_2 AQ_i + \gamma_3 AQ_i^2 + \sum\nolimits_{yq} \delta_y (Y_t = y) + X_{it}\beta + v_{it} \qquad (2)$$

with $ICH_{it} = j$ if $\xi_{j-1} < ICH_{it}^{*} \leq \xi_j$, *j=1,…,m* where we account for *m* cut-off points ξ.

2.2.1 Endogeneity issues and the instrumental variable approach

The causal effect of working on caregiving can be consistently identified at both margins only if the error terms ε_{it} and v_{it} are uncorrelated with the employment status. This assumption is unlikely to hold for several reasons (Carmichael *et al.*, 2010; Crespo and Mira, 2014; He and McHenry, 2016). First, as individuals make caregiving and employment decisions jointly, a two-way causality may occur, as caring choices are likely to be *affected by*, while at the same time being *impacting on* employment decisions. Second, the causal effect may be confounded by observable and unobservable individual and family characteristics influencing both decisions, e.g., the degree of dependency of the care recipient, the caregiver's preferences between caregiving and leisure. Thus, in order to identify the causal effect, we use eligibility to the State Pension to instrument employment status.

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[5] We test alternative specifications in the robustness section.

[6] We transform the dependent variables using the inverse hyperbolic sine function

The State Pension Age (SPA) in the UK is the minimum age at which people can claim the Basic State Pension, which provides an almost-flat minimum level of retirement income (the specific amount depends National Insurance contribution-years). The full Basic State Pension amounted to £5,077.8 for a single individual and £8,119.8 for a couple; corresponding figures in 2016 were £6,029.4 and £9,643.4 (see OECD (2013), PPI (2015) and Lain (2016) for further details). The female SPA, historically set at 60 years old, has been gradually increasing since 2010, depending on year-and-month of birth, for women born after March 1950, and reached 65 in 2018.[7] The rate of increase is nonlinear in birth-date: for women born between 6th April 1950 and 6th April 1953, or between December 1953 and October 1954, SPA increases by one month for those born after the sixth day of each month. For women born between 6th April and 5th December 1953, SPA rises by three months for those born after the sixth day of each month. The impact of the reform is substantial: e.g., being born one year after March 1950 implies a one-year delay in SPA, while the SPA postponement exceeds 36 months for cohorts born after March 1953 (Figure 3 in the Appendix provides a graphical representation). To determine whether an individual lies above or below her SPA at time of interview, we combine available information on year-and-month of birth, and date-of-interview.[8]

To be valid, the instrumental variable should satisfy four conditions. First, it has to be strongly correlated with the endogenous variable, in our case the employment status. Recent estimates show that the recent change in pension rules boosted female employment by 6 to 10% (Cribb and Emmerson, 2017; Cribb *et al.*, 2016). More generally, eligibility to the State pension has been found to have a strong effect on retirement decision, and hence on the probability to be in work, in the UK, as well as in the US and in Europe (Behncke, 2012; Belloni *et al.*, 2016; Bonsang *et al.*, 2012; Eibich, 2015; Mazzonna and Peracchi, 2017; Rohwedder and Willis, 2010).

Second, the effect of the instrument on employment should be monotonic. This means that whilst the instrument may have no effect on some individuals, it must affect everybody who are affected in the same way. The eligibility to the State Pension may not affect the employment decision of some people (who may keep working despite having reached the SPA), but it is very unlikely that people would decide to start working as a result to being eligible to the State Pension.

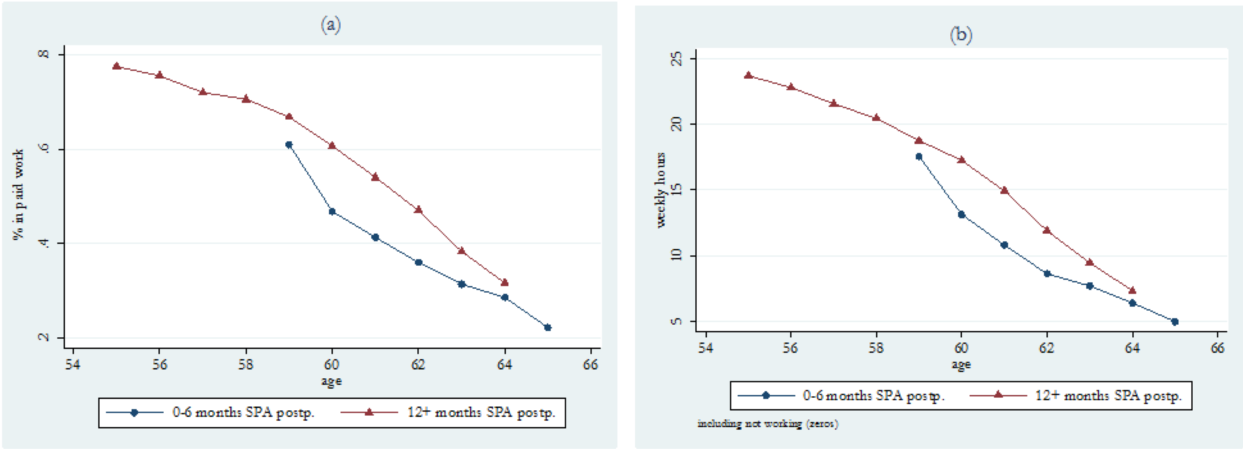
Figure 2 plots the employment rates and the working hours by age group, distinguishing between women (nearly) unaffected by the reform (having zero or up-to 6 months postponement in their SPA), and cohorts whose SPA was delayed by 12 or more months. Both graphs show, unsurprisingly, that both labour attachment and intensity rapidly decline between the age of 55 and 65. Most importantly for us, employment rates are shown to be much higher for the ‘affected’ group among people aged 60 to 62 (a gradient of 11 to 15 percentage points), whilst such difference is strongly reduced at ages 63 and 64 (gradient of 6 to 3 points), when all women (in this sample) are above SPA. Similarly, working hours are higher (by up to 4.3 hours per week) for affected women compared to their same age reform-unaffected control groups.

Figure 2 here

[7] Since 2019, both female and male SPA will rise from 65 to reach 67 in 2028 (see Thurley and Keen (2017)).

[8] SPA eligibility is based on whether individuals are born before the sixth day of each month. In our data, we have no information on day of birth. Therefore, our definition of SPA eligibility assumes that respondents are subject to the pension rules as if they were born from the 6th day of the month or later (as in, e.g., Cribb *et al.* (2016)). Although this leads to some misclassification of SPA eligibility, misclassification is relatively small: we estimate that if dates of birth are uniformly distributed within each month, we would misclassify about 1.2% of women above SPA as being below SPA at time of interview.

Figure 2, employment rates (a) and weekly working hours (b) by age, for women with 12+ months vs zero/six months postponement in SPA



Notes: panel (a) shows the percentage of women reporting to be working for pay; panel (b) shows the average weekly hours worked (non-working women are assigned zero hours).

Finally, the instrument must be uncorrelated with the error terms ε_{it} and υ_{it}. As pointed out by, e.g., Angrist and Pischke (2009), this assumption has two parts. First it implies that the instrument must be as good as randomly assigned. The State Pension Age is defined by the UK government and depends on age and date of birth. individuals have no say in whether they are eligible or not, as it is highly unlikely that they would be able to manipulate their date of birth.
The instrument must further satisfy the exclusion restriction, which, in our context, implies that being below SPA affects care giving only thought the effect on employment. Individuals who are below the SPA are younger than those who are not, and this could influence propensity to provide care directly. To prevent this, we extensively account for age and year effects by adding fixed effects for both age and year. This is only possible because, due to the reform, pension eligibility does not solely depend on age (in a non-linear way) but also on birthdate. Conditioning on age, being below the SPA is unlikely to have affect care giving through other channels than employment and retirement.
Using the pension eligibility status as an instrument implies that our estimates should be interpreted as a Local Average Treatment Effect (LATE), that is, the effect on caregiving of being in paid employment induced by being below the SPA. The impact on caregiving of becoming unemployed or leaving work for any other reasons and regardless of the pension eligibility status may be different to that of retiring. Therefore, even if there was no endogeneity, we would expect the uninstrumented model to yield different results than IV model.

We estimate model (1), essentially an IV-probit, through a 2 stage residual inclusion (2SRI), where the instrumented variable is the continuous amount of log working hours, as suggested in the recent literature (Deb *et al.*, 2017; Terza *et al.*, 2008). As per the instrumented interval regression model (2), we employ the eintreg command in Stata15. In all analyses, standard errors are clustered at the month-of-birth level (224 clusters) to account for the fact that treatment assignment varies by month of birth. However, findings are robust to standard errors being clustered at the individual level (Section 3.2).

3. RESULTS

Table 2 reports the estimation outcomes for the two-part model at both the extensive (columns 1-2) and intensive margin for weekly care hours in logs (columns 3-4) in our main specification, where the independent variable of interest is the amount of log weekly work-hours. At the extensive margin the coefficient of interest should be interpreted as the impact on the caregiving prevalence of a 1%

increase in working hours (extensive magin); at the intensive margin, the coefficient is already an elasticity (% change in caregiving hours for a 1% change in working hours), among caregivers. We report the effect on care provision outside the household in Panel A and on care provision inside the household and in Panel B.

At the extensive margin, in the uninstrumented model (column 1), higher working rates are associated with higher probability of providing residential care, yet no effect is found for non-residential care. When endogeneity is accounted for, no causal effect of working longer on caregiving prevalence is found (column 2). The instrument is informative and valid, as the F-test of excluded instrument (being below SPA) is highly significant. Full results for the first-stage equations are available from the authors. At the intensive margin, a negative and relatively small statistical association is confirmed between work intensity and caregiving hours in the uninstrumented model, in both care settings. Conversely, the causal estimates suggest a larger impact of working longer on non-residential help: specifically, a 1% increase in work-hours would lead to a 0.3% reduction in care-hours: given that the average amount of weekly working hours if 15.3, and that non-residential caregivers provide an average of 9 weekly care-hours, an increase of 1 work-hour induced by being below the State Pension Age reduces caregiving by 0.177 hours (10 minutes). As highlighted in the previous section, the IV estimates should be interpreted as a LATE, in our case as the effect on caregiving of working longer because of the pension eligibility status. Those who keep working because they are below the SPA (the compliers) are likely to have less time to devote to providing care. This may be different from the effect of being in or out of a job for any other reason. For instance, those who are out of paid work because they are in poor health may also be unable to provide care because of their health, which would attenuate the relationship between working and caregiving. Similarly, those who become unemployed may devote most of their time to look for another job, or may see their health deteriorate as a result of a long spell of unemployment, preventing them to increase the time they spend providing care.

No statistically significant results are found for in-household help. However, because of the small number of observations, our instrument becomes weak and we may lack power to detect the causal effect of working longer on care intensity.

Table 2 Effect of working hours on informal care provision

	Any amount of care		log weekly hours among carers	
	(1) probit	(2) IV-probit	(3) Interval reg	(4) Endog Int reg
A – extra-household caregiving				
Log working-hours	0.003	0.264	-0.054***	-0.316**
	(0.002)	(0.212)	(0.011)	(0.127)
N	24,055	24,055	4,887	4,887
F-test excl. instr.		42.31		24.19
Mean outcome	0.206	0.206	9.046	9.046
B – in-household caregiving				
Log working-hours	-0.010***	-0.191	-0.118***	-0.430
	(0.002)	(0.143)	(0.020)	(0.441)
N	24,137	24,137	1,938	1,938
F-test excl. instr.		42.30		3
Mean outcome	0.081	0.081	31.376	31.376

Note: In columns 1 and 2 the dependent variable is a binary variable indicating if the respondent provides any care outside the household (Panel A) or inside the household (Panel B). In columns 3 and 4 the dependent variable is the log of weekly hours spent providing care coded as intervals. In columns 2 and 4, working is instrumented by the status of being above/below State Pension Age (SPA). All models control for fixed effects for age (in years), interview year, Socio-economic Status (NS-SEC), home-ownership, marital status (dummies for being single or widowed) and number of children (dummies for one child, two or more). We report marginal coefficients computed at means for a probit (column 1), iv-probit (column 2), interval regression (3), and instrumented interval regression (4) models. Standard errors are clustered at the month-year of birth level.

We then test whether the above findings are confirmed when the independent variable is defined as a binary indicator for being in paid work (Table 3). Results from the instrumented-probit model[9] in column 2 highlight no causal relationship between being in paid-work and providing any amount of long-term care among our sample, regardless of the care setting (extra / in-household). As before, the instrument is informative and valid.

At the intensive margin, results from the interval regression suggest that, without correcting for endogeneity (column 3), being in work is statistically associated with a 20% lower supply of extra-household care-hours;[10] the effect is larger among residential caregivers, with a reduction of 36%. The magnitude of causal effects (instrumented model, column 4) are larger, in absolute terms, than the statistical associations: working leads to a substantial 72% reduction of weekly hours for non-residential caregivers. Given an average amount of 9 care-hours per week, the reduction due to the extended working career would amount to 6.5 hours when evaluated at averages.

As before, the result for in-household care is non-significant.

[9] We estimate the model through the biprobit command in STATA, due to the presence of a dichotomous dependent and independent (endogenous) variables.

[10] To get the percentage change, we compute: $(e^{-0.221} - 1) \cdot 100$

Table 3, Effect of working longer on care provision, extensive and intensive margin

	Any amount of care		log weekly hours among carers	
	(1)	(2)	(3)	(4)
	probit	biprobit	Interval reg	Endog Int reg
A – extra-household caregiving				
working	0.013	-0.078	-0.221***	-1.308***
	(0.009)	(0.132)	(0.041)	(0.501)
N	24,297	24,297	4,948	4,948
F-test excl. instr.		35.9		24.66
Mean outcome	0.206	0.206	9.046	9.046
B – in-household caregiving				
working	-0.038***	0.000	-0.450***	-1.555
	(0.006)	(0.019)	(0.077)	(1.628)
N	24,379	24,379	1,960	1,960
F-test excl. instr.		35.9		3.3
Mean outcome	0.081	0.081	31.376	31.376

Note: In columns 1 and 2 the dependent variable is a binary variable indicating if the respondent provides any care outside the household (Panel A) or inside the household (Panel B). In columns 3 and 4 the dependent variable is the log of weekly hours spent providing care coded as intervals. In columns 2 and 4, working is instrumented by the status of being above/below State Pension Age (SPA). All models control for fixed effects for age (in years), interview year, Socio-economic Status (NS-SEC), home-ownership, marital status (dummies for being single or widowed) and number of children (dummies for one child, two or more). We report marginal coefficients computed at means for a probit (column 1), iv-probit (column 2), interval regression (3), and instrumented interval regression (4) models. Standard errors are clustered at the month-year of birth level.

3.1.1 Heterogeneity by caregiving intensity

The results above are based on the entire distribution of caregiving hours, which are self-reported in brackets by respondents. We now investigate whether our findings are heterogeneous by intensity of care. There are at least two reasons why being in paid work may have different effects depending on the level of care giving intensity. First, higher intensity levels may reflect in higher commitment and responsibility attached to the caring activity, which could imply lower substitutability between care and work for intensive caregivers. Those providing a large number of care hours may thus not adjust their care supply to an increased working horizon. On the other hand, higher time-consuming caregiving may be more difficult to sustain while remaining engaged longer in the labour market (He and McHenry, 2016). To detect higher intensity caregiving we adopt a threshold of 20+ hours a week, as suggested by Carmichael *et al.* (2010) (in our dataset, 50% of in-household caregivers provide less than 20 care-hours a week). Second, the distribution of care-hours is highly skewed for nonresidential caregiving, as almost 50% of caregivers provide less than 5 hours a week (see Figure 1). Thus, we set a threshold of 5+ weekly care-hours to identify a minimum meaningful caregiving activity in non-residential settings (Michaud *et al.*, 2010). We thus run a series of iv-probit models as in (1), but with the dependent variable being a binary indicator for providing 20+hours of care (in- or extra-household), and for providing 5+ hours of care extra-household (Table 4).

Focusing on non-residential assistance, results in column 6 indicates that a significant share of caregivers reduced their caring contribution below a threshold of 5 hours a week (column 6), among those induced to work longer by being below the SPA. [11] The extent of the reduction for the compliers is substantial: for an induced change in working hours equal to the population average (15 hours a week), the prevalence of 5+ hours caregiving drops by almost 6 percentage points, compared to a population average of 11.6%. This is not surprising when read in conjunction with the previously described large negative effect of working longer on caregiving intensity. When focusing on higher-intensity care (20+ weekly hours), no statistically significant effect of working on care is found for high-intensity caregivers, when controlling for endogeneity (column 4): the coefficient suggests a reduction by 1.78 percentage points, yet it is only significant at 14%.
For residential settings, after controlling for endogeneity, working does not seem to affect the prevalence of intensive caregiving (column 2). Results are confirmed when using the working dummy in place of working hours.

Table 4, heterogeneous effects of working hours, for different levels of caregiving intensity

	In household intensive care (20+ hrs week)		Extra-household intensive care (20+ hrs week)		Extra-household care 5+ hrs week	
	(1) probit	(2) biprobit	(3) probit	(4) biprobit	(5) probit	(6) biprobit
Log working hrs	-0.098***	-0.004	-0.002***	-0.013	-0.002	-0.058***
	(0.013)	(0.014)	(0.001)	(0.009)	(0.001)	(0.022)
N	24,015	24,015	24,015	24,015	23,635	23,635
F-test excl. instr.	45.7		45.7		45.6	
Mean outcome	0.0440		0.0240		0.116	

Note: In columns 1 to 4 the dependent variable is a binary variable indicating if the respondent provides 20+ weekly hours of care inside or outside the household. In columns 5 and 6 the dependent variable is a binary for 5+ hours care provision outside the household. In columns 2, 4 and 6, log-working hours are instrumented by the status of being above/below State Pension Age (SPA). All models control for fixed effects for age (in years), interview year, Socio-economic Status (NS-SEC), home-ownership, marital status (dummies for being single or widowed) and number of children (dummies for one child, two or more). We report marginal coefficients computed at means. Standard errors are clustered at the month-year of birth level.

Overall, the results presented in Table 2 to Table 4 highlight a consistent causal impact of being in employment on extra-household caregiving amongst those below the retirement age. This effect consists in a sensible reduction in care hours (below 5 hours per week), yet not in a complete withdrawal from care-provision, and it is mostly concentrated among medium-low intensity caregivers (the vast majority of the carer population). Conversely, caregivers providing higher intensity assistance do not reduce it under the threshold of 20 hours per week. Moreover, our results

[11] We exclude from the analysis on the 5+ hours cutoff the respondents who reported a generical amount of “20 hours or less”.

show that residential caregiving duties, which usually brings more responsibility and care-attachment than non-residential caregiving, are largely unaltered by being in work.

3.1.2 Heterogeneity by job-exposure

We now investigate whether the previous findings are heterogeneous across job types, as women in more physically or psychologically demanding occupations may find it harder to balance care duties while working in older age.

Understanding Society does not include information on job exposure, yet it includes the International Classification of Occupation (ISCO-88) code of each respondent's current or last occupation. This allows us to exploit two occupational indices of physical and psycho-social burden externally built by Kroll and Lampert (2011), which are based on data from a large scale representative survey on working conditions for 20.000 employees in Germany, and matched to the ISCO-88 codes. Namely, the PJI (physical job exposure index) is based on a job's ergonomic stress and environmental pollution while the PSI (psycho-social exposure index) accounts for mental/social stress and temporal loads. Both indices range from 1 (low exposure) to 10 (high exposure).[12] Both indices have been externally validated (see Santi *et al.* (2013) for further details on the adopted methodology) and recently adopted in an influential study on work/retirement transition (Mazzonna and Peracchi, 2017). This classification is better equipped to capture differences in job occupations than the widely used distinctions based on the assumed skill level from the first digit of the ISCO-code (Mazzonna and Peracchi, 2017). Indeed, the PJI and PSI are linked to 100% of all ISCO-88 2-digit codes, 94.8% of the 3-digit codes and 78.5% of the 4-digit codes, and can be built for 99% of our sample.

Following Santi *et al.* (2013) and Mazzonna and Peracchi (2017), we identify as physically (psycho-socially) burdensome those occupations with a PJI (PSI) score of 6 or higher. We split our sample between those who are above or below each cut-off, and re-estimate our models for each sub-sample.[13]

Table 5 summarises our findings for extra-household care. Results for the extensive margin of the two-part model find no evidence of an increase or decrease in caregiving activity (defined as providing any amount of care) as a result of being in paid work due to pension eligibility status, as in our main specification. We thus exclude these results from the table.

Columns 1 to 4 report findings from the instrumented interval regression at the intensive margin (among caregivers). There is strong evidence of heterogeneous impact of working longer, as caregiving intensity is predicted to decrease for women in exposed occupations (both for physical and psycho-social exposure, columns 2, 4), with an elasticity of 0.7 computed at means (15 job-hours and 9 caregiving hours per week). This suggests that an increase of 1 working hour among caregivers leads to a reduction of 0.41 hours, 25 minutes, in caregiving, which is more than twice the effect observed in the analysis for the whole sample. Conversely, the elasticity of working and caregiving for non-exposed jobs is almost zero in magnitude, and non-significant (1, 3). Columns 5-12 complements these findings by looking at the alternative definitions of caregiving. No reduction in

[12] Among the higher levels of PJI we find, e.g., craft, agricultural/fishery, transport and sales occupations, plant/machine operators, travel attendants and restaurant services workers; among the less demanding we find accounting clerks, librarians, business, legal, social science and engineering professionals, government officials and corporate managers. Higher PSI levels comprise, e.g., plant/machine operators, protective service and social workers, travel attendants, general managers, nursing and health professionals, client information clerks; among the lower levels we find administrative, government, legal, social science and engineering professionals, secretaries and office clerks.

[13] Should people select into less straining jobs as a result of their prolonged working horizon, this would affect our mechanism analysis. Reassuringly, when running linear probability models for the probability of having a physically or a psychosocially straining job, we find that being below SPA has no effect (coefficient of -0.007 with s.e. of 0.016 for physical strain; 0.008 (s.e. 0.017) for psychosocial strain).

caregiving is found for women in non-exposed jobs. Conversely, the probability of providing 5+ care hours reduces significantly for those in exposed jobs, with a reduction of around 7-8 percentage points (compared to an average of 11%) for an increase of working hours by 15 hours per week (the population average). This effect, much higher than the one found in the overall sample, suggests that women in burdensome occupations are those driving the results. Interestingly, we also find a significant drop in the prevalence of intensive extra-household caregiving by around 2.5 percentage points for women in exposed jobs only (columns 8-12).

As for the whole sample, no significant effect is found for in-household care, for both exposed and non-exposed jobs. We thus do not present a table for these results, which are available upon request.

Table 5, Effect of working hours on extra-household care: heterogeneous effects by job exposure (physical or psycho-social)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	log weekly hours among extra-hh carers				Extra-household care for 5+ hours a week				Caring for 20+ hours a week			
	Instrumented interval-regression				Instrumented probit				Instrumented probit			
Log working hours	-0.043	-0.736**	-0.150	-0.720**	-0.01	-0.084***	-0.035	-0.071***	0.001	-0.025**	-0.004	-0.022*
	(0.138)	(0.309)	(0.135)	(0.340)	(0.041)	(0.024)	(0.041)	(0.023)	(0.013)	(0.011)	(0.015)	(0.012)
N	2,193	2,630	2,708	2,115	9,430	13,940	11,889	11,481	9,516	14,053	11,998	11,571
F-test excl. instr.	15.8	8.1	17.3	6.5	15.44	25	15.3	29	16	25.5	14	29.1
Exposure	No psycho-social	Yes Psycho-social	No physical	Yes Physical	No psycho-social	Yes Psycho-social	No physical	Yes Physical	No psycho-social	Yes Psycho-social	No physical	Yes Physical

Note: In columns 1 to 4 the dependent variable is the log of weekly hours spent providing care outside the household among caregivers, coded as intervals. In columns 5-8 the dependent variable is a binary variable indicating if the respondent provides 5+ weekly hours of care outside the household. In columns 8-12 the dependent variable is a binary for 20+ hours care provision outside the household. Log-working hours are instrumented by the status of being above/below State Pension Age (SPA). All models control for fixed effects for age (in years), interview year, Socio-economic Status (NS-SEC), home-ownership, marital status (dummies for being single or widowed) and number of children (dummies for one child, two or more). Columns 1-4 report coefficients from an instrumented interval regression. Columns 5-12 report marginal probit coefficients computed at means. Standard errors are clustered at the month-year of birth level.

3.1.3 Testing for mediating health effects

As our estimation strategy instruments the probability of working (i.e., of retiring) with the pension eligibility status, should health be substantially worse for those working longer in older age with respect to those who retired, this could constitute an additional mechanism explaining our results: a worse health might impede working respondents from providing care. Several studies have documented that retirement may have a positive effect on mental (Belloni *et al.*, 2016; De Grip *et al.*, 2012; Eibich, 2015; Kolodziej and García-Gómez, 2017) and physical-health (Bertoni *et al.*, 2017; Bloemen *et al.*, 2017; Coe and Zamarro, 2011; Shai, 2018; Westerlund *et al.*, 2009). However, other studies have highlighted potential negative health (Behncke, 2012; Bonsang et al., 2012; Mazzonna and Peracchi, 2017), or no significant effects of retirement (Coe and Lindeboom, 2008; Coe and Zamarro, 2011; Hernaes *et al.*, 2013), with findings being sensitive to the choice of country, empirical strategy and health outcome (Avendano and Berkman, 2014; Mazzonna and Peracchi, 2017). In order to test whether caregivers' health is likely to drive our results, we run the two-part model as in (1) and (2) where we account for measures of physical and mental health and report the outcomes in Table 6, for the instrumented model. As before, we only show results for the intensive margin, as no effect is found at the extensive margin.

In columns 1 to 6 we employ measures of health based on the 12-Item Short Form Survey (SF-12) a generic health questionnaire based on the widely used Short-Form 36 Health Survey (SF-36). In column 1 and 2 we condition on the Mental Health Composite Score (MCS), while in column 3 and 4 we use the Physical Health Composite Score (PCS), both included as continuous variables. MCS and PCS are based on the same 12 questions, yet they are constructed using different weights, and range from 0 (poor health) to 100 (good health). We refer to Ware (2002) for extensive details on the SF-12. A MCS score of 45 or lower has been validated as a cutoff for clinical depressive disorders (Vilagut *et al.*, 2013), thus we include a dummy for being above or below such threshold in columns 5 and 6.

In columns 7 to 10 we include a different measure of mental health, the General Household Questionnaire (GHQ), a screening test for identifying psychiatric disorders in the general population, focusing on the inability to carry out normal functions and the appearance of distress (Goldberg *et al.*, 1997; Goldberg and Williams, 1988). The GHQ-12 consists of 12 items, each scored from zero to three using a Likert scale, or alternatively dichotomised between experiencing vs not-experiencing the symptom. The overall index is the sum of each item, ranging from 0 to 36 or from 0 to 12, with higher scores indicating worse health. We use the continuous 0-36 score in column 7 and 8. GHQ values of 3 or higher (in the 0-12 scale) have been shown to be a predictor of clinical depression (Makowska *et al.*, 2002), thus we include an indicator for being above/below this cut-off in columns 9 and 10. Finally, in columns 11 and 12 we jointly add the SF12 MCS and PCS and the GHQ scores in the model. Overall, our findings are not affected by the inclusion of caregivers' health status: the coefficients for being in paid work are very stable across all specifications, they are significant for non-residential care, and non-significant for residential care. We also re-estimated our models with the dichotomous dependent variable for intensive caregiving (20h/week) and minimum intensity extra-household caregiving (5h/week), obtaining very similar results (available upon request) as in Table 4.

Women in exposed jobs may be more protected by going into retirement (Coe *et al.*, 2012; Mazzonna and Peracchi, 2017), therefore we tested whether health could be a mechanism explaining the reduction in caregiving among this sub-population. When re-estimating the models separately by job-exposure index while adding variables for respondents' health, the results in Table 5 are entirely confirmed (available upon request), which suggests that health is unlikely to be the mechanism driving the effect of working on caregiving.

Table 6, Effect of working on care hours (among caregivers): controlling for health

	(1)	(2)	(3)	(4)	(5)	(6)
	A - Caring outside the household					
Log working hours	-0.334**	-0.341**	-0.323**	-0.331**	-0.316**	-0.318**
	(0.151)	(0.165)	(0.150)	(0.156)	(0.151)	(0.151)
N	4,601	4,601	4,873	4,588	4,873	4,873
F test excluded instrument	19.84	17.44	20.33	19.00	20.23	20.18
	B - Caring inside the household					
Log working hours	-0.606	-0.804	-0.489	-0.949	-0.487	-0.485
	(0.693)	(1.177)	(0.665)	(1.382)	(0.688)	(0.677)
N	1,763	1,763	1,925	1,747	1,925	1,925
F test excluded instrument	1.32	0.57	1.30	0.52	1.20	1.19
Health measures	MCS	PCS	MCS <= 45	GHQ	GHQ >= 3	MCS PCS, GHQ

Note: The dependent variable is the log of weekly hours spent providing care, among caregivers, coded as intervals. Log-working hours are instrumented by the status of being above/below State Pension Age (SPA). All models control for fixed effects for age (in years), interview year, Socio-economic Status (NS-SEC), home-ownership, marital status (dummies for being single or widowed) and number of children (dummies for one child, two or more). We report coefficients from an instrumented interval regression. Standard errors are clustered at the month-year of birth level.

3.2 ROBUSTNESS CHECKS

We run a number of alternative specifications to test the robustness of our findings, and summarise them in Table 7, for the instrumented models for care inside and outside the household. We only report the robustness tests for the second part of the two-part model, as the coefficients of interest at the extensive margin remain nonsignificant under all the alternative specifications. Thus, results in Table 7 can be compared to the estimates presented in column 3 and 4 of Table 2, which represents our preferred specification.

In the first column of Table 4 we show that results are robust to not including any covariates other than age and year of birth dummies (which are necessary for the exclusion restriction to hold). The estimate from the instrumented interval regression model is nearly identical to that of our preferred specification. In our main specification, we adopt a flexible functional form by adding dummies for each age (in years) and interview year. In columns 2 we show results from models that include fixed effects for age in quarter rather than in year. In column 3 we further include fixed effects for interview date in quarter instead of year. The coefficients are of the same magnitude as that of our preferred specification but are less precisely estimated. By using quarter dummies to control for age and interview year dummies we reduce the variation generated by the increase in SPA, hence lower F stat of the excluded instrument and the less precise estimates. We also tried less flexible specifications which included age as a second or third-order polynomial, and confirmed our main findings (available upon request).

In columns 4 we remove individuals who report providing a varying number of hours of care each week (20 hours or less; 20 hours or more), as the imprecise length of such categories could reduce the precision of our estimates. Results are entirely confirmed.

Our sample excludes individuals who never worked in their life. However, a potential concern is that women who were at some point involved in paid work, but stopped working several years ago, might be less affected by our instrument (pension eligibility). For example, women who have been inactive for a long time might have built smaller pensions over the lifecycle, and might therefore be more likely to be affected by an extension of the SPA. By contrast, those who left the labour marker earlier might comprise women from a higher socioeconomic status, which may be less affected by the pension reform. In columns 5 we remove individuals who left the labour force prior to 1999, and show that this does not alter our results.

Furthermore, as our sample includes repeated observations for some individuals, we run our models clustering standard errors at the individual level rather than at month-of-birth level. Coefficient in column 6 entirely confirm our main findings.

Overall, whatever the specification we use, the results remain the same: being in work because of being below the SPA reduces the number of hours of caregiving provided outside of the household but we find no statistically significant effects for in-household caregiving.

Results for models using a dummy for being at work are equally robust and available in the on-line appendix. Similarly, we ran the same robustness tests on the models having as dependent variables the binary indicators for intensive care (20+ hrs per week) and for 5+ hours extra-household care. Results are always confirmed (available upon request).

Table 7, Effect of working on care hours : different specifications

	(1) Interval reg	(2) Endog Int reg	(3) Interval reg	(4) Endog Int reg	(5) Interval reg	(6) Endog Int reg
	A - Caring outside the household					
Log working hours	-0.318**	-0.283*	-0.281*	-0.324**	-0.320**	-0.326**
	(0.151)	(0.166)	(0.170)	(0.156)	(0.155)	(0.152)
N	4,873	4,873	4,873	4,683	4,554	4,873
F test excluded instrument	20.18	14.81	14.35	18.80	19.25	14.70
	B - Caring inside the household					
Log working hours	-0.485	-0.347	-0.359	-0.528	-0.110	-0.395
	(0.677)	(0.500)	(0.481)	(0.605)	(0.466)	(0.496)
N	1,925	1,925	1,925	1,593	1,594	1,925
F test excluded instrument	1.19	2.32	2.60	1.96	1.86	2.07
Covariates	No	Yes	Yes	Yes	Yes	Yes
Age	Year f.e.	Quarter f.e.	Quarter f.e.	Year f.e.	Year f.e.	Year f.e.
Year	Year f.e.	Year f.e.	Quarter f.e.	Year f.e.	Year f.e.	Year f.e.
Sample	All	All	All	No left work<1999	No vague care-hours categories	All
Clustered SE	MoB	MoB	MoB	MoB	MoB	Individual

Note: "f.e."= fixed effects; "MoB" = month of birth. Model in column 1 only include age and year additional controls. Model in (2) controls for age-quarters fixed effects; model in (3) controls for age-quarters and year-quarters fixed effects;Model (4) excludes from the sample those who left their last job before 1999. Model (5) exclude respondents who reported vague care-hours categories (less than 20hrs, or more than 20 hrs). Model (6) clusters standard errors at the individual level.
The dependent variable is always the log of weekly hours spent providing care, among caregivers, coded as intervals. Log-working hours are instrumented by the status of being above/below State Pension Age (SPA). Models 2-6 additionally control for Socio-economic Status (NS-SEC), home-ownership, marital status (dummies for being single or widowed) and number of children (dummies for one child, two or more). We report coefficients from an instrumented interval regression. Standard errors are clustered at the month-year of birth level apart from column 6 (individual level).

4. Discussion and conclusion

In this paper, we exploit a major pension reform in the UK to estimate the causal effect of labour supply on informal care provision in older age, using an instrumental variable approach. Our analysis shows that working longer as result of not being eligible for State pension has no effect on in-household care giving but reduce informal care provided outside the household. We find no effect on the probability to provide care, but find that a 1% increase in work-hours leads to a 0.3% reduction in care-hours. Moreover,

working longer reduces by 6 percentage points the probability of providing 5 or more hours of extra-household care per week for a representative respondent.
Using two occupational indices of physical and psycho-social burden built by Kroll and Lampert (2011) we show that the adverse effects of working are stronger for those in physically and mentally demanding jobs. Our results should be interpreted as the effect on care giving of the extension in working-life because of ineligibility to the state pension. Since our approach relies on being eligible for State pension as an exogenous source of variation in labour supply, our results cannot readily be generalised to any source of variation in labour supply (e.g., the local unemployment rate as in He and McHenry (2016)), especially for younger groups, which are however not the focus of our research question.
Our results are important for several reasons. First, we establish that labour supply affect informal care provision for a group (women aged 55 to 65) that constitutes a crucial supply of care in the UK. Second, we provide the first evidence for policymakers on the negative side-consequences of a major pension reform in the UK. More generally, our results suggest that policy makers should account for the unintended consequences on informal care supply when regulating labor markets and social and health services for older people. Policies promoting flexible work schedules, a smooth transition to retirement, as well as other 'family-friendly' policies for workers in more demanding jobs may help to reach a more efficient work-family balance, with potential important consequences for individual wellbeing (OECD, 2017b). Indeed, Hassink and Van den Berg (2011) examined whether informal caregivers were able to reduce opportunity costs by joint production, combining informal care with other activities. Albeit informal care involves a substantial degree of simultaneity with household activities, it is very hard to combine it substantially with paid work.
Finally, the evidence that working longer affects the provision of informal help should stimulate further research exploring the consequences on the overall coverage (both in quantitative and qualitative terms) received by senior citizens in need of care, and the potential consequences on public budgets. Numerous studies have investigated the relationship between informal-care and formal-care use, with most evidence showing a complementary relationship for high-skills care (Balia and Brau, 2013; Bonsang, 2009; Carrino *et al.*, 2018b), and a substitution relationship for low-skill care (Bolin *et al.*, 2008a; Bonsang, 2009; Stabile *et al.*, 2006; Van Houtven and Norton, 2004). Public programmes of care in Europe are highly heterogeneous in their target population, and only a few countries account for the availability of informal care when determining eligibility for formal care (Brugiavini *et al.*, 2017; Gori and Fernandez, 2015). Moreover, programmes that directly support caregivers are often still limited to situations of high care-intensity and very low means, as in the UK where the Carer's Allowance targets caregivers earning less than £480 a month (net of taxes), and provide more than 35 hours of assistance per week (Brugiavini *et al.*, 2017).

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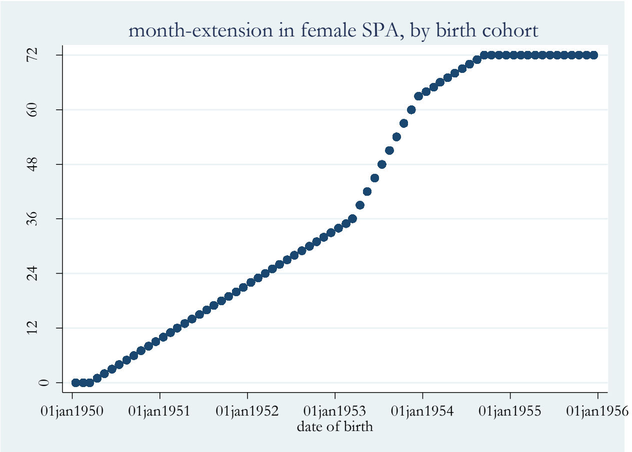
6. TABLES AND FIGURES

7. APPENDIX

7.1 ADDITIONAL FIGURES

Figure 3, postponement in the State Pension Age for women, as per Pension Acts 1995, 2011 and 2014



Note: authors' calculations based on Pension Acts 1995, 2011 and 2014

7.2 ADDITIONAL TABLES