Financial support for long-term elderly care and household saving behaviour

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Abstract

In 2002, Scotland introduced a set of reforms which increased the financial support for long-term elderly care. We study how these reforms affected households' propensity to save. Using a difference-in-differences estimator, we find that the policies reduced the household saving rate by 1.9 percentage points. This amounts to an annual reduction in savings of $\pounds 503$. The estimated effect is heterogeneous. The effect is particularly strong among potential care givers (head of household in his/her 40s) and potential care recipients less likely to receive informal care (singles older than 65 living alone).

JEL classifications: C21, D14, I18, J14

1 Introduction

How to financially support the elderly and their families during the period of their long-term elderly care needs is a policy question often debated in many developed countries. When designing a policy aimed at financially supporting the elderly with their care costs, it is imperative that we consider the potential behavioural changes among the elderly and their families. One concern related to the introduction of a more generous system of formal elderly care is that households may reduce the amount of accumulated assets over their life-cycle, since they anticipate that they will rely more heavily on public funds.

We exploit 2002 Scottish reforms, which increased the amount of in-kind financial support for long-term elderly care, to study their impact on the saving rate. Before the introduction of these reforms, Scotland and the rest of the UK shared the same public system for long-term elderly care. After 2002, Scotland experienced significantly higher level of financial support compared to the rest of UK. Therefore, UK households outside of Scotland can be used as a control group to disentangle the impact of these policies on saving behaviour from any other changes in assets induced by time effects common to all UK regions.

The existing literature investigated the impact of health insurance coverage on individual or household savings in the US (Gruber and Yelowitz, 1999; Maynard and Qiu, 2009; Lee, 2016; Gallagher et al., 2017), the UK (Guariglia and Rossi, 2004), and Taiwan (Chou et al., 2003, 2004). These papers typically used policy reforms that changed the eligibility conditions or the generosity of insurance coverage as natural experiments. Most of them found that generous health insurance coverage reduced the amount of savings, except for Guariglia and Rossi (2004) and Gallagher et al. (2017), who found no effects. Gruber and Yelowitz (1999) and Maynard and Qiu (2009) estimated that \$1000 Medicaid-eligible dollars reduced wealth by 2.5% and 5.5%, respectively. Chou et al. (2003) and Chou et al. (2004) found that the introduction of Taiwan's comprehensive health insurance coverage reduced wealth by between 8.6% and 19.3%. Another group of papers from the US studied the impact of medical expenditure uncertainties on individual/household savings among those aged 65 and above by estimating structural life-cycle models (Palumbo, 1999; De Nardi et al., 2010). They concluded that such uncertainties were an important factor for precautionary savings.[1]

[1] See Van Ooijen et al. (2015) for a survey of the theoretical literature on life-cycle models with uncertain health expenses and savings.

None of the above studies explicitly focused on medical expenditures related to long-term elderly care, since their measures also included all the other health-related costs. However, as many governments are attempting to find ways to support the elderly population, it is important that we understand the implications that long-term support for elderly care has for saving. One paper that addresses this issue and is closely related to our paper is Costa-Font and Vilaplana-Prieto (2017). They studied the effect of a 2007 Spanish policy, called *Sistema para la Autonomia y Atención a la Dependencia* (SAAD). SAAD is a non-means-tested governmental support for long-term elderly care that offers either cash or in-kind support. By focusing on individuals aged 55 and above, they found that the overall policy reduced savings by 13–39% of the average individual subsidy. The strongest effects were found for those aged between 55 and 75 and those receiving cash benefits, whereas most of the estimates suggest insignificant effect of in-kind support.

Our contribution to the existing literature is three-fold. First, we enrich the scarce empirical evidence on the impact of financial support for long-term elderly care on household saving rates by exploiting a natural experiment. Understanding an unintended impact of these policies on saving contributes to developing a more complete picture of the impact of the reforms and will allow policymakers to better anticipate the overall effects when planning future policies. Second, unlike Costa-Font and Vilaplana-Prieto (2017) or much of the existing literature focusing on those who are either younger than 65 or older than 65, we estimate the effect across age. Changes in the prospect for future care costs might have impacted not only those who immediately benefited from such policies but also younger generations as they planned their future consumption and savings. Finally, only a handful of studies have investigated the issue outside of the US. Our paper contributes to this limited literature by presenting the UK evidence.

We find that the Scottish policy reforms reduced the average household saving rate by 1.9 percentage points. If we take a Scottish household with the average gross income across the period under analysis as the reference, this effect amounts to an annual reduction in the saving flow of $£503$. In addition, the estimated effect is heterogeneous: it is more negative if the head of the household (HOH) is aged between 30 and 50, single, and childless.

This article proceeds as follows. Section 2 provides institutional information on the 2002 Scottish reforms. Section 3 discusses the theoretical predictions. Section 4 presents the data and the econometric model. Section 5 reports and comments on the estimation results. Section 6 concludes. The Online Appendix, available at the journal website, contains further details on the reforms and the full set of estimation results from the baseline model.

2 Background information

Prior to 2002, formal personal care[2] costs in the UK were paid almost entirely by individuals.[3] Such costs exposed individuals in need of long-term care to a significant financial burden. In 2000, an average individual in England required 7.6 hours of personal care per week, and the average hourly cost of personal care was approximately £12 (National Statistics, 2002). Approximately 39% of households benefiting from home care received 6 or more visits and more than 5 hours of care per week. Half of these households required intensive care, defined as more than 10 contact hours and 6 or more visits during the week. In addition to charges for formal personal care, local authorities often charged for meals delivered to the home or participating in day care sessions. The financial burden faced by the elderly, therefore, raised concerns among UK policymakers (Netten et al., 2003).

Amid growing concerns regarding the financing of elderly care, in December 1997, the Labour government established the Royal Commission on Long Term Care for the Elderly under the chairmanship of Sir Stewart Sutherland. The Commission reported back to the UK Parliament in March 1999 (Sutherland report), recommending that for those aged 65 and above, formal personal care should be provided free of charge after rigorous need-based assessment conducted by local authorities.

At the same time as the publication of the Sutherland report, the UK political system went through significant changes. Powers were transferred in 1999 from Westminster to devolved governments in Scotland, Wales, and Northern Ireland, while England remained under the direct control of Westminster. The establishment of devolved governments implied that each government acquired some scope to form its own health care policies, although the differential degree of devolution meant that some had more autonomy from Westminster compared to the others. In response to the Sutherland report, Scotland welcomed the idea of state-funded personal care. The Scottish Executive established the Care

[2]Examples of personal care are bathing, toileting, assistance with preparation of and eating food, and dressing. Family members may informally provide care to the elderly. Paid personal care is also available from social workers administered by local authorities or privately hired caretakers. Paid personal care is referred to as ‘formal’ care.

[3]Stringent means-tested subsidies were offered to the elderly once their wealth fell below £18,500 (2001 rate).

Development Group in January 2001, with the aim of pursuing options on how to implement state-funded personal care and to evaluate the estimated cost of introducing such a policy. After several revisions, the Bill passed and received Royal Assent on 12 March 2002 to become the Community Care and Health (Scotland) Act 2002 (CCHA), which in turn was implemented on 1 July 2002. In contrast to Scotland, the rest of the UK did not follow the Commission's recommendation to offer formal personal care for free, and they continue to charge individuals for this type of care.

The CCHA introduced Free Personal Care (FPC) in Scotland, but it distinguished care offered at home from that received in residential care homes. If an individual received formal care at home, all personal care costs were covered as long as the local authority assessed the individual and approved the amount of care.[4] Cost coverage for formal care provided in residential care homes was instead fixed at a flat rate.[5]

The FPC policy pays the specified amount as either in-kind support or a cash allowance. However, the overwhelming majority of recipients chose to receive the support in-kind and requested their local authorities to arrange for formal personal care (Gillespie, 2017). Moreover, even among those who opted for the direct-payment option, there is no actual cash transfer into their bank accounts (Direct Payments Scotland, 2003). Once the preferred care provider is found, the local authority would pay that part of the care cost directly to the care provider.

The maximum weekly amount of FPC in 2002 was £145. This amount translates to an annual amount of £7,540. As the annual basic pension during this period was on average approximately £3,940 (DWP, Annual abstract of Statistics, 2013), the individuals would have been required to spend their entire basic pension to cover their care costs in the absence of FPC. This implies that the introduction of FPC is likely to have significantly reduced the financial burden faced by households.

At the same time as the introduction of FPC, two other changes related to long-term elderly care were implemented for those staying in residential homes. The first is related to Attendance Allowances. The Attendance Allowance (AA) is a non-means-tested weekly benefit for severely disabled people aged 65 or over who need help with personal care. It had been paid out to all UK individuals whom the local authority assesses as being in need. However, Scottish individuals in residential care homes were no longer entitled

[4]On average, individuals received £80 per week for formal personal care received at home (National Statistics, 2012).

[5]Individuals are still asked to pay other costs such as those of cleaning, day care, laundry or meals on wheels.

to AA after 2002.

Another change was related to nursing care (NC) cost coverage. NC is medical care offered by registered nurses. Individuals that receive care in their own homes have always received this care for free at the point of delivery. In contrast, those in residential care homes needed to cover the cost. To eliminate the differential treatments based on care location, NC allowances were introduced throughout the UK for those in residential care homes between October 2001 and October 2002. Further background details on these policies are included in Online Appendix A.

2.1 Summary of financial gains

Table 1 highlights individuals' financial gains due to all the reforms by care setting and region of residence. For each region, we calculate the maximum possible amount of weekly allowances given to individuals. When we focus on those in residential homes, the largest increase in the amount of allowances is observed for Scottish individuals, but the increase is also affected partially by the introduction of NC and the withdrawal of AA. In contrast, the increase in the allowances for Scottish individuals receiving care at home is exclusively due to the 2002 FPC reform.

Table 1 may give the impression that those interventions other than the FPC reform may have affected the relative attractiveness of going into a residential home. This is because those in care homes outside of Scotland receive approximately £100 more than those at home after 2002. However, it is important to remember that the NC cost is free if received at home. This implies that assessing the effective amount of allowances received at home requires adding approximately £100 to every 'care received at home' cell in Table 1. See Online Appendix A.4 for a more detailed discussion.

Our estimates reflect the joint effects of changes in care related financial support. However, the majority of individuals in the UK receive care at home: approximately 70% of care recipients in England received care at home in 2010–2011 (AgeUK, 2014). Therefore, the most likely source of financial support information that both recipients and their family members are exposed to is the one for the home care setting. For this reason, the 2002 FPC reform is likely to be the most relevant policy change for the majority of the population.

Table 1: Maximum weekly allowance calculations (£ per week)

Care received in care homes	Before the reforms £ per week	After the reforms £ per week
England	54.88 (AA)	56.25 (AA)+110 (NC)=166.25
Wales	54.88 (AA)	56.25 (AA)+107.63 (NC)=163.88
Northern Ireland	54.88 (AA)	56.25 (AA)+100.00 (NC)=156.25
Scotland	54.88 (AA)	145.00 (FPC)+65.00 (NC)=210.00
Care received at home	**Before the reforms £ per week**	**After the reforms £ per week**
England	54.88 (AA)	56.25 (AA)
Wales	54.88 (AA)	56.25 (AA)
Northern Ireland	54.88 (AA)	56.25 (AA)
Scotland	54.88 (AA)	56.25 (AA)+145.00 (FPC)=201.25

Notes: This table illustrates how the maximum amounts of weekly allowances changed before and after the reforms depending on which region the elderly reside in and where they receive care. The amounts of allowances are adjusted for inflation and converted to 2002 rates using the Office for National Statistics consumer price index. AA stands for Attendance Allowance; FPC means Formal Personal Care allowance; and NC is the Nursing Care allowance. Since there is no upper limit on the amount of FPC for Scottish individuals at home, we use the maximum amount provided to those in residential care homes, i.e. £145. The take-up rate of AA in the UK overall is 13.7% of the 65+ population (Department for Work and Pensions statistics, 2016). The take-up rate of FPC is approximately 9.5% of the Scottish 65+ population. Information on the take-up rate of NC is available only for Scotland, amounting to 3% of the Scottish 65+ population (National Statistics, 2007).
Sources: Bell et al. (2006), Institute for Fiscal Studies (2012), King (2018).

3 Predicted directions of the effect of the policies

Section 2 highlighted that all households in Scotland experienced an increase in financial support for elderly care with the largest increase received by those at home. According to life-cycle theory, agents plan their consumption and saving behaviour over their entire life-cycle on the basis of their lifetime wealth, i.e. the discounted sum of expected future lifetime income. However, when agents become aware of an unexpected shock to future income,[6] they update their consumption and saving to return to an optimal consumption-saving path. For example, when individuals predict an increase in future income, they reduce their saving at all ages to smooth their consumption.

The reforms could differentially affect households across age groups. On the one hand, Krueger and Perri (2011) show that the magnitude of the reduction in saving depends on the length of time since the shock: the farther away the income shock is, the smaller the reduction in saving. Therefore, we would expect the magnitude of the negative effect on household saving to be increasing with age. On the other hand, Gourinchas and Parker (2002) note that the effect of the precautionary saving motive is particularly prevalent during the early years: households younger than 40 years accumulate wealth

[6]The reforms cut the costs of personal care for the elderly. Therefore, it can be interpreted as an expected increase in the income available for consumption when aged 65 or above.

for the uncertainty of future income and, were it not for income uncertainty, they would instead borrow against future labour income.

Another reason for potentially finding a differential effect across age is that the 2002 policies may have induced informal carers to change labour supply.[7] Hollingsworth et al. (2017) demonstrated that, in response to the Scottish reforms, individuals aged between 25 and 54 did not change their labour supply, whereas those who are older than 55 increased it, at both the extensive and intensive margins. The increase in labour supply and, consequently, in income could impact the saving rate either positively or negatively. Since the saving rate is the ratio between saving and income, the increase in income could either increase or reduce the saving rate, depending on whether saving reacts with a larger or a smaller growing rate than the one of income.

There are other household characteristics across which we might expect a heterogeneous impact of the reforms on saving. For example, if precautionary motives are important in determining households' propensity to save, we might expect a stronger impact on those households less endowed with safety nets. Similarly, if households without potential informal carers face more precautionary motives, the introduction of the Scottish reforms would imply that the magnitude of the reduction in saving for these households would be larger.[8] Childless households and single households are more likely to be vulnerable to health shocks and are more likely to self-insure against the cost of future personal care. Low-income households are at a higher risk of poverty in the event of health shocks, and therefore, their dissaving reaction to the reforms is expected to be more pronounced. However, personal care was subsidised even before the reforms for households with low wealth (see footnote 3). Hence, we may observe a smaller estimated effect of the policies for this subgroup.[9]

[7] Employment-related policies such as those regarding the taxation and benefits or the retirement/pension systems are all administered by Westminster and not by the Scottish government (GOV.UK, 2013). Similarly, the large-scale workforce programmes, such as the New Deal, which targeted the young and the other age groups, were implemented throughout the UK.

[8] De Nardi et al. (2010) find that retired single individuals in the US keep a large amount of assets to respond to the risk of expensive medical care.

[9] Our dataset does not have information on wealth. We therefore cannot assess the heterogeneous effect of the policies across this dimension or use low-wealth households as a further control group.

4 Data and econometric specifications

4.1 Data, sample, and variable definition

This study employs the repeated cross-sectional dataset of the UK Expenditure and Food Survey (EFS). EFS has been collected by the Office for National Statistics (ONS) on a yearly basis since 2001. Prior to 2001, EFS was called the Family Expenditure Survey, and the same set of information has been collected by the ONS since 1961. Every year, approximately $10,000$ households are interviewed, and information is collected at the household and personal level. The EFS contains extensive information on expenditure and income at both the household and individual levels. Such expenditure information is reported during the two-week window in which the survey is conducted. The information on income and expenditures included in EFS makes this dataset ideal for the purpose of our study. We exploit this information to compute our outcome variable, the saving rate, i.e. a flow variable defined as the ratio between weekly saving and weekly gross income. Since it is easier to adjust the flow than the stock of saving, it is more likely for us to observe the effect of the policies on the former. Other widely used datasets, such as the Family Resources Survey or the British Household Panel Survey (BHPS), do not include all the relevant information. For example, the Family Resources Survey only reports the stock of savings. Similarly, the BHPS only contains information on the positive amount of savings per month and not the amount of borrowing; thus, the calculated saving rate would be truncated at zero.

The analysis is carried out at the household level. Not all the expenditures are collected at the individual level in the EFS. The expenditures that cannot be linked to a particular individual, such as those for a vehicle or a package holiday, come from questions at the household level. Hence, by aggregating all the variables at the household level, we do not need to identify the sharing rule governing the intrahousehold distribution of household-level expenditures.

The final sample ranges from 1998 until 2007. The choice of the starting year is made because variables used to control for regional time-varying heterogeneity are available beginning in 1998. Data were only included until 2007 to avoid the 2008 financial crisis, which may confound the effect of the 2002 policies. Northern Ireland is excluded from our sample due to its small sample size. This implies that, in our data, we have 11 official regions, 10 in England and Wales, which are the controls, plus Scotland as the treated region. We further exclude households with an HOH younger than 30. This is done not only to ensure that the sample is composed of households that are likely to look after elderly members but also to minimise the chance of including households whose main breadwinner is still in education. Finally, we eliminate observations reporting zero gross income (66 households) and cut the bottom and the top percentiles of the saving rate distribution ($1,204$ households). This is to prevent our findings from being driven by outliers. Our initial sample size was $76,218$ households. After applying these selection criteria, we are left with a sample of $55,831$ households.

We define the post-policy period to begin on 1 April 2002. This is because the FPC bill, which most significantly increased the formal care subsidy out of all policies studied in this paper, passed on 12 March 2002. The progression of the bill was closely followed by the UK media and received wide coverage. Therefore, it is likely that households in Scotland were aware of the policy even prior to its implementation. To test for potential anticipation effects, we conduct a sensitivity analysis by eliminating observations interviewed in 2001. This sensitivity analysis is reported in Section 5.2. Households interviewed in 2002 are assigned to the post- or pre-policy period according to the month of interview. In the EFS, interviews take place almost uniformly over all the months of the year.[10]

The dependent variable in our model is the saving rate, defined as the fraction of the weekly household gross income not spent on goods or services, income taxes, or employee national insurance contributions.[11]

In a sensitivity analysis, we also use the consumption rate as the dependent variable, i.e. the ratio between the weekly household expenditure and the weekly household gross income.[12] We show in Section 5.2 that our estimated effect is not sensitive to the choice between gross and net income.

Table 2 presents descriptive statistics of the household saving rate before and after the end of March 2002 in Scotland and in the rest of the UK. It also includes the raw

[10]Table B-2 in the Online Appendix reports the distribution of the month in which the households were interviewed.

[11]The expenditures are on both non-durable and some durable goods (e.g. vehicles and furniture). Although durable goods could significantly contribute to the expenditure value, they are unlikely to be included during the two-week period in which households complete the survey, since they are infrequently purchased.

[12]An alternative way to define the dependent variable is to use the net instead of the gross income. We opted for the gross income since households might take particular consumption choices during the fiscal year to manipulate the amount of taxes and, therefore, the amount of net income.

double difference, which shows how the difference in the amount of saving between the two regions changed over time. Three points emerging from Table 2 are worthy of mention. First, Scottish households on average have a higher propensity to save (0.93 pp) than those in England and Wales (0.68 pp). Second, the regional difference in saving becomes negligible after the reforms. Third, Scottish households' overall propensity to save declined over time compared to those in England and Wales. The unconditional difference-in-differences of the average saving rate is equal to -0.56 pp, although it is not significant.

Table 2: Summary statistics of the saving rate before and after the reforms for the treatment and control groups

	Mean	Std. Dev. (Std. Err.)	Min.	Max.	Observations
Scotland					
Overall, 1998-2007	0.0093	0.4533	-3.0669	0.6905	5,107
Before, 1998-2001	-0.0042	0.4432	-3.0669	0.6905	2,196
After, 2002-2007	0.0195	0.4606	-2.9300	0.6862	2,911
Mean difference after − before	0.0237	(0.0127)*			5,107
England & Wales					
Overall, 1998-2007	0.0068	0.4683	-3.1746	0.6910	50,724
Before, 1998-2001	-0.0098	0.4628	-3.1746	0.6903	21,957
After, 2002-2007	0.0194	0.4722	-3.1743	0.6910	28,767
Mean difference after − before	0.0293	(0.0042)***			50,724
Difference-in-Differences	-0.0056	-0.0054			55,831

Notes: *** Significant at 1%; ** significant at 5%; * significant at 10%.
Source: Authors' calculations using the 1998–2007 EFS.

Table 3 reports descriptive statistics of the saving rate by selected household characteristics, across which we will study the heterogeneity of the effect. Overall, the household saving rate amounts to 0.7%, meaning that households were able to save 0.7% of their gross income on average during the observed time window. The saving rate is higher for older (6.2%) households and when the HOH is single (1.5%) or childless (3.0%). In Table B.1 of the Online Appendix, we report the descriptive statistics of the regressors used in the econometric analysis.

4.2 Difference-in-differences model

Identification of the effect of the policies on household saving behaviour is achieved by exploiting the fact that the Scottish reforms were introduced only for a specific group of individuals and that both the treated population (those in Scotland) and the untreated population (those in the rest of Britain) are observed before and after the reforms. Simply

Table 3: The saving rate by household characteristics

	Relative frequency	Total sample Mean	Total sample Std. Dev.	Scotland Mean	Scotland Std. Dev.	England & Wales Mean	England & Wales Std. Dev.
Saving rate	1.000	0.007	0.467	0.009	0.453	0.007	0.468
Saving rate by the age of the HOH							
Between 30 and 39 years old	0.222	-0.019	0.436	-0.032	0.459	-0.018	0.434
Between 40 and 49 years old	0.221	-0.013	0.453	-0.027	0.428	-0.012	0.456
Between 50 and 64 years old	0.280	-0.012	0.479	-0.018	0.455	-0.011	0.481
Older than 64 years	0.277	0.062	0.465	0.089	0.436	0.059	0.468
Saving rate by HOH's marital status							
Single	0.385	0.015	0.431	0.015	0.411	0.015	0.433
Living in a couple	0.615	-0.005	0.519	0.001	0.509	-0.006	0.520
Saving rate by the presence of children in the households							
No children	0.681	0.030	0.473	0.040	0.454	0.029	0.475
Children	0.319	-0.042	0.449	-0.060	0.445	-0.040	0.449
Saving rate by gross income							
Income $\geq$ 75th percentile	0.250	0.141	0.297	0.132	0.291	0.141	0.297
25th percentile $\leq$ Income $<$ 75th percentile	0.500	0.015	0.413	0.031	0.382	0.013	0.416
Income $<$ 25th percentile	0.250	-0.142	0.634	-0.123	0.615	-0.144	0.637
Observations		55,831		5,107		50,724	

Source: Authors' calculations using the 1998–2007 EFS.

comparing household saving behaviour in Scotland before and after 2002 is problematic since there may have been many economic influences other than the policies that affected the household saving rate over time. Similarly, a simple difference between the average household saving rate in Scotland and in the rest of Britain after 2002 also poses a problem because there might be fundamental differences in the household propensity to save between the two groups of regions. As a result, we employ a difference-in-differences (DD) estimator and estimate changes in the differences in the household saving rate between Scotland and the rest of Britain before and after the reforms. The identification of causal effects requires several assumptions. In what follows, we conduct statistical tests for each of these assumptions to check whether they are supported by the data.

Our empirical evaluation will be in a repeated cross-sections framework. We specify the following model for the saving rate y of household i living in region r in tax year t.

$$y_{irt} = \mathbf{x}'_{irt}\boldsymbol{\beta} + \boldsymbol{\gamma}_r + \boldsymbol{\phi}_t + \delta_{DD} I_{rt} + \varepsilon_{irt}, \tag{1}$$

- $\mathbf{x}_{irt}$ is the $K \times 1$ vector of regressors, and $\boldsymbol{\beta}$ is the conformable vector of coefficients. The regressors in $\mathbf{x}_{irt}$ include those at both the household and regional levels. The former are a cubic function of the age, gender, race, and marital status of the HOH, housing tenure, the education of the HOH and of the spouse (if present), and the number of children. They capture differences in labour market attachment, earnings, and relevance of the precautionary or bequest motives and are therefore expected to influence saving behaviour. The controls at the regional level are the unemployment rate (by the gender of the HOH), the per capita gross value added, the per capita gross disposable income, and the Halifax house price index. These controls are included to remove time-varying regional heterogeneity in saving induced by differential evolutions over time of the business cycle, of the state of the labour market, and of the housing market.

- γ_r is a set of regional fixed effects. There are 11 regions in our data. The treated region is Scotland. The control group comprises households living in the remaining 10 government office regions of England-Wales.[13]
- ϕ_t is a set of tax year fixed effects. The unit of time is the tax year, i.e. from 6 April until 5 April of the next year, as the post-introduction period is defined to start in April 2002, which corresponds to the beginning of the 2002 tax year.
- I_{rt} is the regressor of interest. It is an indicator variable equal to 1 if the household resides in Scotland after the reforms, i.e. after March 2002. The corresponding parameter δ_{DD} is the effect of the introduction of the 2002 reforms in Scotland on the saving rate.
- ε_{irt} is the error term at the household level.

The parameters of Equation (1) are estimated using ordinary least squares (OLS). Inference is problematic. In our DD application, the identification of the effect of the policies is based on variations across regions and years. The regressor of principal interest, i.e. the treatment dummy after 2002, is therefore correlated within clusters (i.e. regions), and inference should take this into account. The cluster-robust variance estimator (CRVE) is a simple way to deal with correlation within-groups (Liang and Zeger, 1986). However, this approach is unbiased only when the number of clusters is large enough and the asymptotic results can be safely invoked. In our application, the number of regions is only 11, and therefore the cluster-robust standard errors are likely to suffer from small sample bias, resulting in a type I error.[14] Cameron et al. (2008) proposed a

[13]North West and Merseyside, Yorkshire and the Humber, East Midlands, West Midlands, Eastern, London, South West, South East, South West, and Wales.

[14]See Cameron and Miller (2015) for an overview of the problems in performing inference when the

wild cluster bootstrap-t procedure to obtain critical values when the number of clusters is small. However, MacKinnon and Webb (2017) showed that with unbalanced clusters and a small number of treated clusters (only one in our analysis), the wild cluster bootstrap fails: the wild cluster bootstrap based on unrestricted residuals (WCBUR) tends to over-reject, also resulting in type I errors as in the CRVE t statistics; the wild cluster bootstrap based on restricted residuals (WCBRR) tends instead to under-reject just as severely, resulting in type II errors.[15] To the best of our knowledge, there is currently no method to safely obtain critical values in a DD model with a small number of untreated clusters and one treated cluster. Hence, we report p-values based on the CRVE t statistics[16] and the wild cluster bootstrap procedure of Cameron et al. (2008) with both unrestricted and restricted residuals.[17]

Following the discussion in Section 3, we are interested in estimating potentially heterogeneous effects of the policies across different dimensions of household characteristics. We do this by splitting the sample based on various household characteristics and by replicating the estimation of Equation (1) for each subsample.

The identification of the effects of the policies through a DD approach is based on some underlying assumptions.

Assumption 1 (Parallel trends): Conditional on observables, households residing in Scotland would experience similar trends in the saving rate to those in the rest of Britain if the 2002 reforms had not been implemented.

We test the validity of Assumption 1 by comparing the trends in the household saving rates of England-Wales and Scotland. Figure 1 reports the least squares estimates of the coefficients of the tax year dummies for Scotland and of the tax year dummies for England-Wales. We obtain these estimates by regressing the saving rate on a set of time dummies, the coefficients of which are allowed to differ for Scotland and England/Wales, and on all the other regressors in the baseline equation. In other words, we estimate the

number of clusters is small.

[15]In MacKinnon and Webb (2017), the WCBRR is the procedure in which the model is re-estimated under the null hypothesis of no treatment effect in the bootstrap algorithm. When the procedure is based on the unrestricted residuals, the null hypothesis is instead not imposed.

[16]Given R, the number of regions, we will compute $\sqrt{R/(R-1)}$-clustered robust standard errors and t_{R-1} critical values as suggested in Brewer et al. (2013).

[17]We bootstrapped the residuals 2,500 times using the Webb six-point distribution as weights (Webb, 2014).

following equation

$$y_{irt} = \mathbf{z}'_{irt}\boldsymbol{\omega} + \boldsymbol{\gamma}_r + \boldsymbol{\phi}_t^{EW} + \boldsymbol{\phi}_t^{Sc} + u_{irt}, \quad (2)$$

where $\boldsymbol{\phi}_t^{EW}$ are tax year dummies if individual i lives in England-Wales, and $\boldsymbol{\phi}_t^{Sc}$ are tax year dummies if individual i lives in Scotland; $\mathbf{z}_{irt}$ includes all the other regressors except for the constant. The estimated coefficients of these indicators are plotted in Figure 1. If the saving behaviour in Scotland followed the same trend as that in England and Wales, the two lines depicted in Figure 1 would be parallel before 2002. We formally test whether these trends are parallel by jointly testing for $\forall t = 1998, \ldots, 2001$, $\boldsymbol{\phi}_t^{Sc} - \boldsymbol{\phi}_t^{EW} = k$, where $k \in \Re$ is some constant. Our results indicate that the null hypothesis cannot be rejected, and thus, the parallel trends assumption seems to be fulfilled.[18] [19]

Assumption 2 (Exogeneity of the intervention): Conditional on observables, the Scottish reforms are exogenous with respect to and not motivated by demand for formal personal care in Scotland. Rather, it is politically determined.

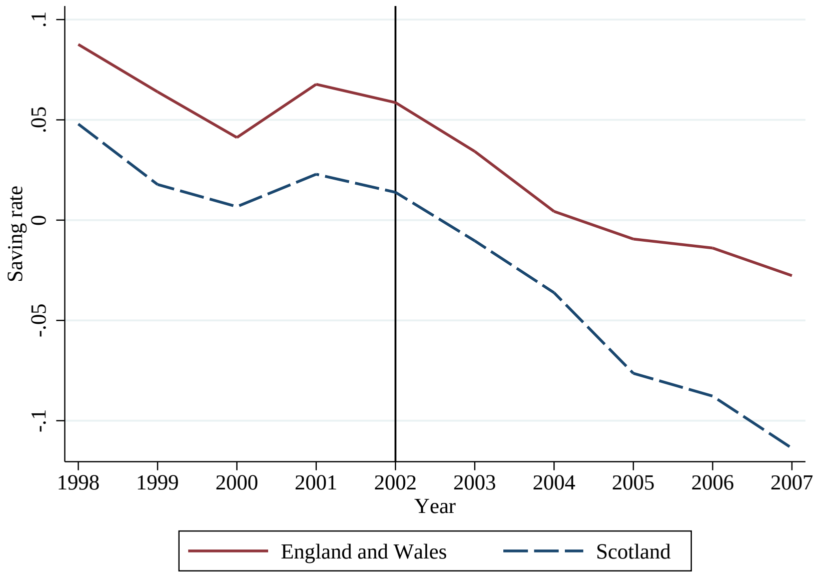
If Scotland had implemented the 2002 reforms in response to increasingly stronger demand for formal care, we would have an endogeneity problem, as the policy variable after 2002 would capture the effect of both the reforms and of the differential trends in the demand for personal care. This would then translate into potentially diverging trends in saving rates between the two groups of regions not because of the policy introduction but rather due to differential underlying demand for formal elderly care. Figure 2 presents the trends of the use of formal personal care in England and Scotland.[20] The left and right panels of Figure 2 illustrate the trends of 1–5 hours and 6 or more hours of personal care usage per week, respectively. Although the aggregated data used to plot these graphs do not allow us to formally test whether these lines are parallel to each other, the plotted

[18]The p-values were as follows: 0.614 from the CRVE t statistic, 0.742 from the WCBUR, and 0.833 from the WCBRR.

[19]We also conducted two alternative tests for Assumption 1. First, we estimated a placebo test by including lags of order one, two, and three of the policy indicator I_{rt} and tested the significance of the associated coefficients. We rejected the null hypothesis of joint significance of these lagged policy indicators. Panel a) of Table D.1 in the Online Appendix reports these findings. Second, we ran a set of placebo tests by pretending that the policy reforms took place in other regions of the UK. Panel b) of Table D.1 in the Online Appendix shows that the coefficients of these placebo policy dummies are not significantly different from 0 at the usual 5% level.

[20]Data for England are taken from the 1999–2007 Community Care Statistics and the 2000-2007 Home Care Services. Similarly, the 2000-2007 Home Care Services, Scotland and 2014 Social Care Statistics were used for the Scottish data. Unfortunately, data for Wales are not available in the same format and thus are not included in the calculation of this figure.

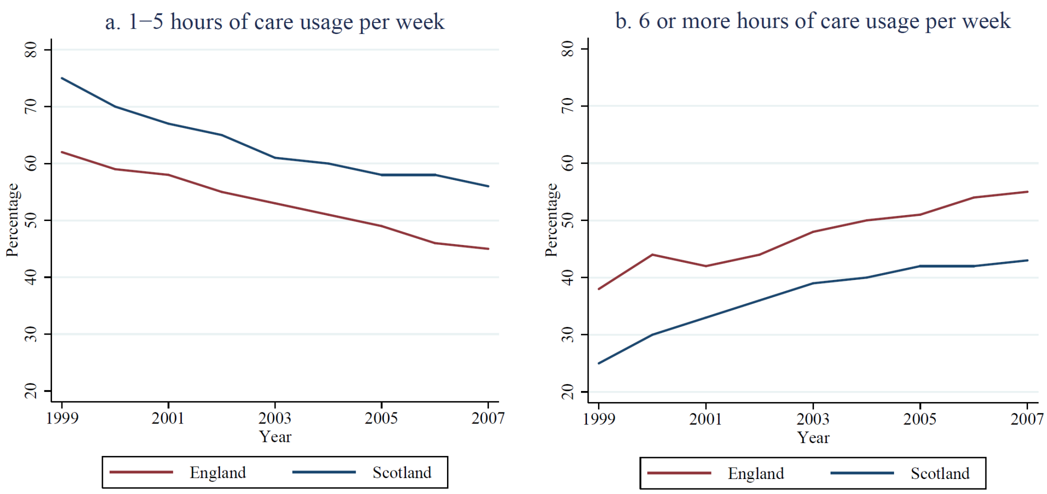
Figure 1: The parallel trends assumption



Notes: In this figure, we report the least squares estimates of the coefficients of the year dummies for Scotland and England-Wales. We obtained them by regressing the saving rate on a full set of time dummies for which the coefficients are different between Scotland and England-Wales and, as further control variables, all the other regressors reported in Table B-1 in the Online Appendix.
Source: Authors' calculations using the 1998–2007 EFS.

trends suggest that these two regions did not experience differential trends. A further piece of evidence to verify the validity of this assumption is presented by Hollingsworth et al. (2017). They analysed the impact of the Scottish reforms on the rate of informal personal care given to another adult. Although they did not consider trends in formal care, several studies show that informal care is a close substitute for formal care (Van Houtven and Norton, 2004; Bolin et al., 2008; Bonsang, 2009). If the trends of informal care usage before 2002 between the two groups of regions are parallel, this is another piece of suggestive evidence that the underlying reason for the introduction of the policies is something other than the demand for formal care. Using the 1999–2007 Family Resources Survey, they showed that the trends in the rate of informal personal care giving in Scotland and England-Wales were parallel before 2002.

Figure 2: The trends in the use of formal personal care in Scotland and England



Notes: We plot the trends in the demand for personal care in Scotland and England. The left-hand side figure shows the trends in the fraction of households using 1–5 hours of care at home per week. The right-hand side figure reports the trends for the fraction of households using 6 or more hours of care per week. Statistics for Wales are not available, as the data are not in the same format as those for England and Scotland. However, as the population of Wales represents less than 5% of the UK population, the exclusion of Wales should not substantially affect the calculated trends.
Sources: Community Care Statistics 1999, 2000, 2001, 2002, 2003, 2004, 2005, 2006, and 2007. Home Care Services, Scotland 2000, 2004, 2007, and Social Care Statistics 2014.

Assumption 3 (Stable sample composition): Conditional on observables, the composition of the treated and control groups is assumed to be stable before and after the policies.

Assumption 3 requires that the composition of the households residing in Scotland, England, and Wales is stable over the observation period, conditional on observed covariates.

Our findings would be biased if, for example, those who anticipate greater needs for formal care and those without much savings moved to Scotland from England or Wales due to the 2002 policies. Using the 1999–2007 BHPS, we analyse whether individuals' moving behaviour changed before and after 2002.[21] Table 4 presents estimates from a linear probability model, where the dependent variable equals 1 if individuals moved to Scotland from England or Wales. It is regressed on a dummy indicator of whether each individual moved to Scotland after 2002 and 0 otherwise. In all cases, we find that the reforms did not result in significant effect of individuals moving to Scotland.[22]

Table 4: Linear probability model regression estimates to test whether individuals moved to Scotland after the 2002 reforms

	Age 25-34	Age 35-44	Age 45-54	Age 55-64	Age 65-85
1 if observed after 2002	0.0003	0.0003	−0.0003	−0.0002	−0.0003
	(0.0004)	(0.0002)	(0.0002)	(0.0003)	(0.0002)
Constant	0.001***	0.0003**	0.0004**	0.0005**	0.0005**
	(0.0003)	(0.0002)	(0.0002)	(0.0003)	(0.0002)
Observations	29,064	31,852	26,971	22,769	28,197

Notes: We used the 1998–2007 BHPS and estimated separate linear probability regressions by the age of individuals. Standard errors robust to heteroskedasticity are reported in parentheses.
Source: Authors' calculations using the 1999–2007 BHPS.

Assumption 4 (No anticipation): Scottish households were not able to anticipate the introduction of the 2002 reforms.

The Scottish government's decision to take up the recommendation received widespread media coverage as early as January 2002. For example, the BBC announced that FPC for Scotland would be introduced in July of the same year on 15 January 2002. Similarly, the Guardian published an article after one of the bills passed in the Scottish Parliament (Inman, 2002). As a result of this widespread media coverage, Scottish households could have anticipated the introduction of the policy and might have faced incentives to alter

[21] BHPS is a UK longitudinal survey, which began in 1991. It collects approximately 5,500 households and 10,300 individuals drawn from 250 areas of the UK. It records detailed information on whether and when individuals moved to different parts of the UK.

[22] Since those living close to the border (i.e. Northern England) face lower costs of moving to Scotland, we may observe a significant effect of moving for those individuals. We conducted a separate analysis by only considering the movements from Northern England to Scotland. Nonetheless, we do not observe any significant change after 2002.

their consumption and saving decisions even before April 2002. If this were the case, the estimated effects would be biased towards zero. Therefore, we will present a robustness analysis in Section 5.2 by eliminating observations from tax year 2001. As we will see, removing 2001 from our sample does not alter our findings.

5 Estimation results

5.1 Baseline parameter estimates

The OLS baseline estimation results for the effect of the policies from Equation (1) are reported in panel a) of Table 5. The full set of estimation results are reported in Online Appendix C, Table C-1. The introduction of the 2002 policies meaningfully reduced Scottish households' propensity to save. The saving rate decreased by 1.9 pp over the observation period. As the average weekly gross income of Scottish households in our sample is $\pounds 506.09$, the estimated reduction in the amount of savings is approximately $\pounds 9.67$ per week or $\pounds 503$ per year. In terms of statistical significance, we obtain different results depending on the approach used to obtain the t statistics. The t statistics based on the CRVE and on the WCBUR indicate significance levels of 5% and 10%, respectively. With the WCBRR, by contrast, we find that the effect of the policies is not significantly different from zero.

To investigate whether the reforms imposed a temporary or a permanent shock to the saving rate, we modified Equation (1) to allow the effect of the policies to differ across years after the reforms. Panel b) of Table 5 shows that the reaction is immediate and increasing over time. The reforms generated a sudden decrease in the saving rate of almost 1 pp and then reached its maximum effect by the end of the period (-3.9 pp). One potential reason for the increasing effect is due to more people becoming aware of the policies over time. For example, back-of-envelope calculations using statistics from the Family Resources Survey and Gillespie (2006) reveal that the proportion of those receiving FPC increased from 27% to 41% of total home care recipients in Scotland between 2002 and 2005. This, in turn, may imply that an increased proportion of care recipients and caregivers became aware of the policies over the years.

As mentioned in Section 3, the effect of the policies may vary with the age of household members, although the predictions on the direction of the overall effect were mixed. Panel c) of Table 5 displays the effect of the reforms after splitting the sample according to

Table 5: The effect of the 2002 reforms on household saving rates

	Coeff.	p-values CRVE[(a)]	WCBUR[(b)]	WCBRR[(c)]	Observations
a) Homogeneous effect					
Scotland∗After (I_{rt})	−0.0191	0.039**	0.084*	0.368	55,831
b) Effect over time					
Scotland∗After∗$\mathbb{1}$(2002-2003)	−0.0094	0.048**	0.012**	0.414	
Scotland∗After∗$\mathbb{1}$(2004-2005)	−0.0115	0.392	0.429	0.532	
Scotland∗After∗$\mathbb{1}$(2006-2007)	−0.0385	0.014**	0.056*	0.316	55,831
c) Effect across age of the HOH					
Age $\in [30, 40)$	−0.0356	0.013**	0.017**	0.360	12,409
Age $\in [40, 50)$	−0.0478	0.000***	0.000***	0.488	12,340
Age $\in [50, 65)$	0.0163	0.195	0.225	0.614	15,629
Age ≥ 65	−0.0204	0.187	0.268	0.450	15,453
d) Effect by marital status					
Single	−0.0542	0.000***	0.000***	0.256	21,496
Couple	0.0044	0.623	0.674	0.727	34,435
e) Effect by presence of children in the household					
No children	−0.0246	0.024**	0.041**	0.357	38,035
With children	−0.0106	0.411	0.452	0.505	17,796
f) Effect for older couples and singles living without other household members					
Age of the HOH ≥ 65, single, no other household member	−0.0600	0.016**	0.060*	0.331	7,295
Age of the HOH ≥ 65, couple, no other household member	0.0077	0.695	0.795	0.758	6,835
g) Effect across spouse's education among younger households					
Age of both partners<65, spouse has low education	-0.0283	0.054*	0.040**	0.305	5,905
Age of both partners<65, spouse has high education	0.0101	0.363	0.455	0.587	20,697
h) Effect across gross income					
Income $\geq$ 75th percentile	0.0094	0.373	0.398	0.537	13,958
25th percentile $\leq$ Income $<$ 75th percentile	−0.0080	0.319	0.382	0.532	27,916
Income $<$ 25th percentile	−0.0263	0.077*	0.082*	0.311	13,957

Notes: *** Significant at 1%; ** significant at 5%; * significant at 10%. $\mathbb{1}(\cdot)$ denotes the indicator function, which is equal to 1 if the argument is true. 'After' is equal to 1 if the observation is collected after 2002 and 0 otherwise. 'Scotland' is equal to 1 if the household resides in Scotland and 0 otherwise.

[(a)] CRVE indicates that the p-values come from the CRVE t statistics.

[(b)] WCBUR indicates that the p-values come from the wild cluster bootstrap t statistics based on unrestricted residuals and 2,500 replications using the Webb six-point distribution as weights.

[(c)] WCBRR indicates that the p-values come from the wild cluster bootstrap t statistics based on restricted residuals and 2,500 replications using the Webb six-point distribution as weights.

[(d)] Spouse's low (high) education means that the spouse of the HOH left education when 15 or younger (16 or older).

Source: Authors' calculations using the 1998–2007 EFS.

the age of the HOH. When the HOH is between 30 and 49 years old, we find the strongest negative effect: -3.6 and -4.8 pp when the HOH is between 30 and 39 and between 40 and 49, respectively. Since the average weekly gross income of Scottish households is $\pounds 581.66$ when the HOH is between 30 and 39 years of age and $\pounds 662.70$ when the HOH is between 40 and 49, the magnitude of the estimated effects is non-negligible, amounting to an annual reduction in savings of approximately $\pounds 1,077$ and $\pounds 1,647$, respectively. Based on our estimates, the reduction in saving as a proportion of the maximum increase in allowances among those receiving care at home ranges between 8% and 22% for households with the head of household aged between 30 and 49. For households with an HOH between 50 and 64, the negative effect disappears.[23]

Panels d) to g) of Table 5 report the effect of the policies by subgroups defined according to different dimensions of households' characteristics. According to the discussion in Section 3, the households that are less likely to receive informal care are exposed to stronger incentives to dissave once the reforms are implemented. Consistently, we find that the dissaving is more important in single households (-5.4 pp) and those without children (-2.5 pp). Moreover, even though we did not find any effect among older households in panel (b) of Table 5, when we restrict the sample to single households with an HOH older than 64 without other household members, the reduction in the saving rate is 6 pp (panel (f)). This supports the hypothesis that the behavioural change is stronger for potential care recipients less likely to receive informal care. In contrast, the lack of any significant effect among other older households may point to the fact that those who had the option to rely on free informal care prior to the 2002 reforms may have simply switched to subsidised formal care as these two are found to be close substitutes (Van Houtven and Norton, 2004; Bolin et al., 2008; Bonsang, 2009).[24] It is also consistent with the results in De Nardi et al. (2010), who found that the US single elderly retain a large amount of assets to self-insure against the risk of incurring expensive medical

[23] Our findings seem to contrast with those of Costa-Font and Vilaplana-Prieto (2017), who studied the effect of the 2007 Spanish policy among those aged 55 and above, finding negative savings effects in the magnitude of 13% to 39% of the subsidy. However, these large and significant negative effects in their paper were consistently found only among those who received cash benefits. In contrast, in the majority of their estimates, they do not observe significant reductions in savings among those who received in-kind support. Since the large part of the increase in the Scottish financial support took the form of in-kind support (Gillespie, 2006), our findings need to be compared to the in-kind estimates reported in Costa-Font and Vilaplana-Prieto (2017).

[24] In fact, Hollingsworth et al. (2017) show that the 2002 reforms reduced informal caregiving to those receiving care at home. This piece of evidence also indicates that the older households with family support simply reduced reliance on family provided care after 2002.

expenditures.

The fact that the impact is driven by childless households is particularly reassuring for another reason. Scotland followed a separate path from the rest of Britain with respect to university tuition fees. In 1998, tuition fees were introduced across the UK (£1,000 per year). While England and Wales subsequently increased their university tuition fees to £3,000 in 2004 and £9,000 in 2009, Scotland abolished tuition fees in 2001 for Scottish individuals studying in Scotland. Instead of charging tuition fees, Scottish students were asked to repay £2,000 after they graduate and start earning at least £10,000 a year. The cheaper university tuition fees in Scotland compared to those in England-Wales may have further reduced the incentives to save for Scottish households with children, introducing a confounding effect in the interpretation of our findings. The results in panel e) of Table 5 suggest that this is not the case.

It is also possible that those who are more likely to provide informal care may respond strongly to the policies. Carmichael et al. (2010) present evidence that those who are less attached to the labour market are more likely to provide informal care. Therefore, we present evidence for younger households (i.e. those with both partners younger than 65) by the spouse's level of education. If spouses with lower education are less likely to be attached to the labour market and are more likely to provide informal care, we would observe a larger reduction in saving among those households. This is indeed what we find (-0.028 pp in panel (g)).

Finally, panel h) of Table 5 reports the effect by income quartiles. It shows that the effect is driven by low-income households. This is consistent with the hypothesis that their behavioural response is more pronounced because they face a higher risk of poverty in the event of negative health shocks and a need for personal care. However, since our dependent variable is defined using gross income, splitting our sample based on income will cause the truncation on the dependent variable problem (Heckman, 1979; Koenker and Hallock, 2001). Therefore, this heterogeneity analysis should be interpreted with caution.

5.2 Robustness checks

We conduct various sensitivity analyses to test the robustness of our baseline findings reported in Table 5.

In the first sensitivity analysis, we assess whether potential differences between the treatment and the control group drive our findings. If the effect is heterogeneous across observed characteristics, the estimated effect can be interpreted as an average effect across different impacts, but we must be certain that a suitable comparison group exists (Blundell et al., 2004). For this reason, we also estimate the effect using inverse probability weighting (IPW), which is similar to the method adopted by Albanese and Cockx (2018). The IPW DD estimation of the effect of the policies is reported in panel (a) of Table 6 and is in line with that from the benchmark model.

Table 6: Robustness checks

			p-values		
	Observations	Coeff.	CRVE[(a)]	WCBUR[(b)]	WCBRR[(c)]
a) Inverse probability weighting with trimming					
Scotland∗After	26,872	−0.0287	0.037**	0.061*	0.247
b) Excluding 2001					
Scotland∗After	49,854	−0.0197	0.033**	0.080*	0.388
c) Including interactions between control variables and 'After' dummy					
Scotland∗After	55,831	−0.0250	0.047**	0.166	0.416
d) Consumption rate as dependent variable					
Scotland∗After	55,831	0.0243	0.016**	0.049**	0.369
e) Saving rate as the ratio between saving and net income					
Scotland∗After	55,840	−0.0234	0.021**	0.064*	0.325
f) Excluding top and bottom 5% of the saving rate distribution					
Scotland∗After	51,273	−0.0129	0.017**	0.020**	0.468

Notes: *** Significant at 1%; ** significant at 5%; * significant at 10%. The estimated parameters of all the other regressors are not reported for the sake of brevity.
[(a),(b),(c)] See the corresponding footnotes of Table 5.
Source: Authors' calculations using the 1998–2007 EFS.

In the second sensitivity analysis, we exclude the year 2001 from our sample to test the existence of the policy anticipation effect as discussed in Section 4.2, Assumption 4. If households started changing their saving rate even in 2001 in response to the widespread media coverage, including the 2001 observations would positively bias our results.[25] However, panel b) of Table 6 indicates that excluding 2001 from our sample does not affect our conclusions.[26]

Third, we modified the baseline model by allowing the coefficients of $\mathbf{x}$ to vary over time. This ensures that the interacted covariates capture heterogeneity in the outcome variable dynamics over time. (Abadie, 2005). Operationally, we interact all the observables $\mathbf{x}$ with the indicator for the period after the policies and include this new set of

[25] In a further check, we changed the definition of the 'after period' from 1 April to 1 July. This change did not affect the estimated effect (-0.0187).

[26] After eliminating 2001 observations, we were left with 49,854 households, of which 4,583 are Scottish.

regressors in the baseline model specification. Panel c) of Table 6 reports the estimated effect, which is close to the baseline estimates.

Fourth, we used the consumption rate, defined as the ratio between the weekly household expenditure and the weekly household gross income, as an alternative dependent variable. As our results thus far indicate that households reduced the flow of saving, it would be interesting to determine whether this is also reflected by changes in the amount of consumption. Panel d) of Table 6 displays the estimation results of the impact of the 2002 reforms on the consumption rate. Scottish households' propensity to consume increased by 2.4 pp, an effect that is consistent with that from the baseline model.

Fifth, we defined the saving rate as the ratio between saving and net income, instead of gross income. The estimated effect reported in panel e) of Table 6 suggests that the main finding is robust to this alternative specification of the saving rate.

Finally, we trimmed the top and bottom 5% of the saving rate distribution (instead of the top and bottom percentiles) to assess the sensitivity of the benchmark estimate to a more conservative definition of outliers. The estimated parameter reported in panel f) of Table 6 is close to the baseline estimates.

6 Conclusions

We studied the impact of the 2002 reforms that offered more generous financial support for long-term elderly care in Scotland on British households' propensity to save. The Scottish reforms legislated that formal personal care be offered to the elderly free of charge. In contrast, the rest of UK continued to charge the elderly for these services. If households save to prepare for future elderly care expenditures, such a reduction in the care price may have led households to respond by reducing their propensity to save. This paper, therefore, studies an unintended consequence of the 2002 policies and evaluates whether and to what extent it crowded out private saving.

By using the households in England and Wales as the control group, we investigate how Scottish households' saving rate responded to the introduction of the policies for the elderly by using a difference-in-differences estimator.

We find that the Scottish policy reforms reduced the household saving rate by 1.9 pp on average, equalling approximately $\pounds 503$ per year. The effect is more negative when the HOH is aged between 30 and 50, single, and childless. A stronger precautionary saving motive for these types of households provides an explanation for the detected heterogeneity in the policy's impact. We also observe a large negative effect if the HOH is older than 64, single, and living alone. This supports the hypothesis that the saving behaviour of potential care recipients who are less likely to receive informal care is more sensitive to the change towards a more generous system.

Given the sizeable effect on saving, one may wonder whether households over-estimated the benefits introduced by the 2002 reforms, as noted by Bell et al. (2006). If so, the resulting reduction in precautionary saving might lead to a situation in which there is less than full insurance against long-term care for the elderly. In such a case, universal elderly care insurance schemes introduced in countries such as Japan or Germany may be more effective in addressing the large and volatile risks of long-term care for the elderly. These questions are left to be investigated in future studies.

Supplementary material

Supplementary material is available on the OUP website. The data used in this paper was made available to us by the UK Data Service. Due to the terms and conditions set out by the UK Data Service, the data cannot be shared on the OUP website. The supplementary material comprises the replication command files and the online appendix.

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