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The Effect of Free Personal Care for the Elderly on Informal Caregiving

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Contents and

Abstract

Population forecasters have predicted that the proportion of people in the UK aged 65 and above will rise significantly in coming decades. This shift in demographics will put increasing pressure on the NHS and providers of social care. However, older people do not rely only on care provided by the state; informal care of the elderly is often supplied by family and friends. Therefore, the relationship between formal and informal care and the reaction of informal carers to institutional changes is an important policy issue. This study uses individual-level data from the British Household Panel Survey to estimate the effects of the introduction of free personal care for the elderly in Scotland on informal care behaviour. As the change in policy applied only to Scotland, a natural experiment is formed allowing a difference-in-differences approach to be used. This paper finds that the introduction of the policy increased the probability of supplying informal care by 3 to 5 percentage points. In addition, it reports evidence of a shift in the hours of care distribution towards the lower tail as a result of the change in policy.

Introduction

In 2010, 10 million people in the UK were 65 and over. By 2050, this number is predicted to double, constituting a quarter of the population. The ageing population is expensive—average NHS spending on retired households is nearly double that on non-retired households (Cracknell, 2010). However, formal care provided by the state is not the only resource on which the elderly rely. England alone has nearly six million informal carers, a large number of whom are over 65 themselves (Beesley, 2006).

Following the Scotland Act 1998, the Scottish Parliament was given responsibility for devolved matters, which include health, education, local government and social work. Foreign policy, national defence, fiscal and monetary policy, immigration and social security, among others issues, are known as “reserved matters” and remain the responsibility of Westminster.

In 2002, the Scottish government implemented a policy in which free personal care would be provided to everyone aged 65 and over who needed it. As a result, Scottish elderly residents became eligible for a payment to help with personal care either within their own homes or in a care institution.

This paper will assess the effects of this policy on the provision of informal care using data from the British Household Panel Survey (BHPS, 2009). As Scotland is the only part of the UK to implement the policy, its introduction provides a convenient natural experiment in which we may investigate the relationship between formal and informal care. We use a difference-in-differences approach to evaluate its effects, with the rest of the UK as the control group.

The observed relationship between formal and informal care is important in informing policy. If increased state provision of care lowered the incentive to supply informal care, the effect of the policy may have been crowded out and the net benefits low. However, if the policy encouraged more individuals to supplement the state-provided care with care of their own, there may have been a sizeable benefit to those receiving care.

The work most closely related to this paper is the study by Bell, Bowes and Heitmueller (2006) who, using the BHPS, found no statistically significant change in the probability of supplying informal care. The current paper expands on these findings in a number of ways. First, it uses data from a greater number of years to account for the potential for lagged effects of the policy. It also uses an arguably cleaner definition of care-supply which helps to prevent problems that may arise from the relatively small Scottish sample. Whereas Bell, Bowes and Heitmueller (2006) use individual fixed effects and exploit longitudinal variation at the individual level to assess the effects of the policy, this paper performs longitudinal analysis at the aggregate level by using the representative survey structure of the BHPS.

In addition, this study investigates care at the intensive margin and estimates the effect of the policy on the number of hours of care supplied by individuals. In order to avoid potential bias that could have occurred as a result of a violation of the common trends assumption, a difference-in-difference-in-differences (DDD) estimator is used for a robustness check. Finally, it simplifies the theoretical model of Bell, Bowes and Heitmueller (2006) so that it includes only one type of care and may be represented graphically.

The remainder of the paper is organised as follows. Section 2 explains the background and details of the policy. Section 3 discusses the existing literature on the relationship between formal and informal care, including studies on the policy introduced in Scotland. Section 4 outlines the motivating theory, and Section 5 describes the data used. Section 6 explains the difference-in-differences methodology and presents the results. In Section 7, a number of robustness checks are performed and Section 8 concludes.

Free Personal Care

Background

Shortly after the Labour Party was elected in 1997, the government created the Royal Commission on Long Term Care. The purpose of the Commission was to assess the current system of the provision of care for the elderly and to provide recommendations for establishing a sustainable system of funding long term care in the home and in other settings.

The conclusion of the Commission was that the current system was in clear need of reform. It was particularly concerned with the inequity in the way individuals broadly classed as requiring "social care" were treated. A new system was proposed in which personal care costs (the additional costs of care due to frailty or disability) were distinguished from living and housing costs and would be funded by general taxation (Sutherland, 1999).

Although these recommendations were for the whole of the UK, only Scotland acted on them. In 2001, a Care Development Group (CDG) was created in order to bring forward the proposal. The CDG also focused on the notion of equity:

> Free personal care is right in principle because it will remove the current discrimination against older people who have chronic or degenerative illnesses and need personal care. It will bring their care in line with medical and nursing care in the NHS where the principle of free care based on need is almost universally applied and accepted. (CDG, 2001, p. 10)

In July 2002, The Community Care and Health (Scotland) Act (Scottish Executive, 2002) was implemented and Scotland became the only region in the UK to provide care to people aged 65 and over without means testing, both in care homes and in a domiciliary setting. This was a "flagship" policy for the newly devolved government, attracting a great degree of political and media attention.

Details

The Regulation of Care (Scotland) Act 2001 defines personal care as:

> . . . care which relates to the day to day physical tasks and needs of the person cared for (as for example, but without prejudice to that generality, to eating and washing) and to mental processes related to those tasks and needs (as for

> example, but without prejudice to that generality, to remembering to eat and wash). (Scottish Executive, 2001, p. 6)

Under The Community Care and Health (Scotland) Act, people aged 65 and over are entitled to a flat rate payment of £145 per week and those who receive care in a nursing home receive an additional £65 per week. The payment is delivered to care homes and individuals via local authorities (Bell and Bowes, 2006).

Uptake of the payments has been high and costs have increased steadily since the policy was first introduced. In the policy's first full fiscal year (2003/04), the Scottish Parliament estimated the full cost of free personal care to be £144m (Bell and Bowes, 2006). The amount spent on care for older people in that year accounted for around 0.2 percent of Scottish GDP and the additional costs of providing free personal care for the elderly increased this amount by around 10 percent (Bell and Bowes, 2006). By 2011, the cost of providing personal care in the home had increased to £342m per year. The cost of providing the care in nursing homes rose by 25 percent from £86m per year to £108m per year over the same period. The large increases in spending are a result of increased demand. The number of elderly people receiving care in the home increased by 42 percent over these eight years. In total, around 77,000 people now receive care in either setting, compared with 64,000 in 2002 (BBC News, 2012).

Existing Literature

There is an extensive literature that investigates the relationship between formal and informal care in different settings. Often, researchers are concerned with whether the two types of care are complements or substitutes in the production of care for older people.

Theoretical literature

Economic models of the relationship between formal and informal care often use the neoclassical model of family decision-making with household production, first proposed by Becker (1981). In addition, many of the studies focus on the endogeneity of living arrangements and the key role that this plays in the optimisation problems of various family members.

For example, Hoerger, Picone and Sloan (1996) design a model in which the parent has three options: to live independently, live with his child or live in a nursing home. Each state is associated with a utility function and a corresponding budget constraint. It is assumed that the parent and child choose the living arrangement jointly to maximise family utility, subject to the budget constraint.

Family utility in each case is a function of formal care, informal care, the parent's consumption, the child's consumption, the level of services in the nursing home, the level of services in the family (or separate) home and the severity of the parent's disability. The budget constraints depend on the prices of informal and formal care, the "units of housing" in each state, and the incomes of both parent and child. The family budget is larger when parent and child choose to cohabit (as opposed to living separately, not in a nursing home) because overall housing costs are lower. It also is cheaper for the child to provide informal care because travel costs are reduced to zero. The income of the disabled parent is assumed to be a subsidy for care, received from the state.

One living arrangement will be chosen over another if the resulting family utility is higher from choosing the first. This set-up allows the authors to model the effect of institutional changes (in particular, changes in the price of formal care) on the type of living arrangement selected. A similar approach is taken in Pezzin, Kemper and Reschosky (1996) and Van Houtven and Norton (2004).

In the studies discussed above, care is assumed to be an undifferentiated service. In practice, however, formal and informal carers may have preferences over types of care or specialise in one over another. Bell, Bowes and Heitmueller (2006) build on existing models of caring behaviour to include different types of care.

In their "care production function" they include types of care produced, informal care inputs and formal care inputs. Whether the care inputs are formal or informal is not assumed to affect the utility of the older person directly. Instead, the authors include both types of care in the utility function as well as a preference parameter over who provides the care.

The utility functions of carer and caree are used to derive an expression for the marginal rate of technical substitution between formal and informal care of each care type which varies across care types as formal and informal carers have different skill sets. The authors then derive an expression for the marginal product of informal care in the production of each care type which can be compared to the marginal product of formal care in the production of the same type. This set-up allows them to model changes in behaviour due to changes in the relative prices of types of care.

For example, suppose there are only two types of care, c_1 and c_2, and the price of c_1 decreases. In this model, the carer would be motivated to either substitute towards greater participation in the labour market or substitute towards increased provision of another type of care. If the latter effect offsets the former, no change in the total level of informal care provided will occur. Therefore, a fall in the price of one type of formal care will not necessarily lead to a fall in informal care, as family members may substitute towards providing more of another type.

Empirical literature

The relationship between formal care, informal care and family living arrangements also has been extensively studied empirically. Hoerger, Picone and Sloan (1996) use The National Long Term Care Survey to estimate the effect of various factors that affect the relative prices of formal and informal care, for example Medicaid subsidies for formal care. Using multinomial probit specifications, the authors found that increased Medicaid subsidies meant that disabled parents were more likely to live independently than with their children.

The relationship between formal and informal care also has been studied specifically in the context of the introduction of free personal care for the elderly in Scotland. For example, McNamee (2006) uses the Scottish Household Survey to examine the trends in the use of formal and informal care following the introduction of the policy. Using Mann-Whitney non-parametric tests, McNamee (2006) finds no significant change in either the level or intensity of informal care after 2002. However, he finds evidence that the probability of informal care being provided after the policy change was less dependent on income than it had been: before the introduction of free personal care, a higher income was associated with a higher likelihood of receiving informal care, but afterwards there was no significant association between the two variables.

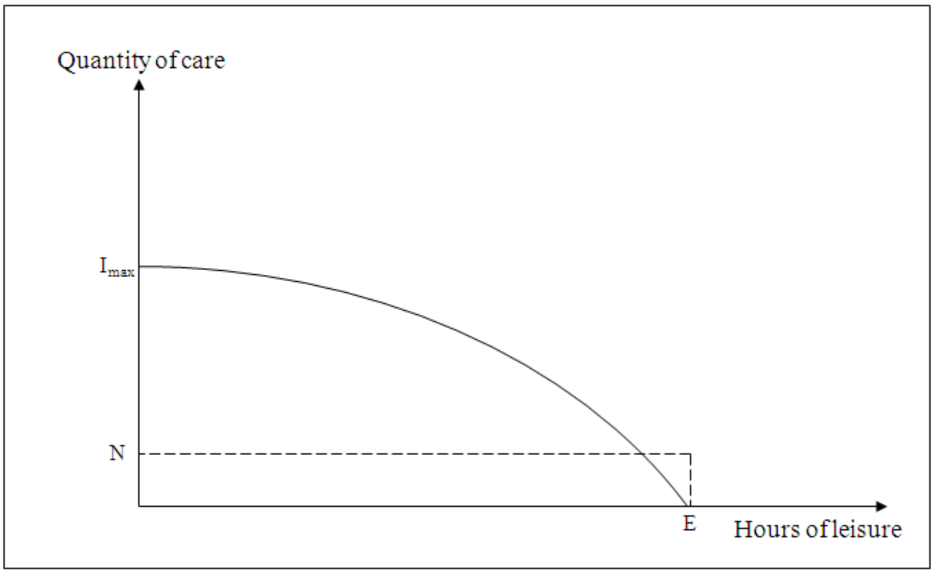
Finally, use the BHPS to calculate a difference-in-differences estimate of the effect of free personal care in Scotland on the level of informal care. They use a linear probability model with fixed effects and examine the effect of the policy on co-residential and extra-residential care separately. Their data are from 1999 to 2003—two years before and after the policy change. The authors found no statistically significant change in the probability of individuals providing co-residential care, suggesting that the offsetting effect implied by their model may exist. It is suggested that the carer may wish to observe the health of the dependent or the quality of care, or desire to see herself as remaining a carer to her parent. As explained in their model, a fall in the price of one type of formal care will not necessarily lead to a fall in informal care, as family members may substitute towards providing more of another type. There was a decrease, however, in the probability of providing extra-residential care, although this result was of a small magnitude and significant only at the 10 per cent level.

The authors also look for changes in the distribution of hours of care. This is done by simply comparing the proportions of people providing certain numbers of hours before and after the policy change without using regression analysis. They find a reduction in the intensity of care provided.

Modelling Framework

The purpose of this paper is to examine the consequences of a simple natural experiment: the introduction of free personal care for the elderly, implemented in Scotland in 2002, but not elsewhere in the UK. The model used in this paper is a simplified version of the models of care giving described in the third section of the paper and focuses specifically on the trade off between labour supplied for informal care versus other uses of the carer's time. It is a straightforward, two-good diagrammatical model, where living arrangements are implicit. Unlike the model in Bell, Bowes and Heitmueller (2006), all types of care are combined into one composite care variable. Such a simplification allows straightforward predictions and interpretations of care behaviour. The single indifference curve in the diagram may be interpreted as belonging to the carer or the shared household (carer and caree).

Figure 1. The production possibility frontier



We assume that the carer chooses an optimal bundle of two "goods", total quantity of care and leisure hours, to maximise her utility. An individual's time is split between care hours and "non-care hours" so if an individual gives up one hour of leisure, this translates into supplying one more hour of care.

Each hour of care supplied is an input into the carer's "care production function" that transforms hours of care into quantity of care delivered to the parent. This is an example of a household production function. We assume diminishing returns to scale, so that each hour of care is less productive than the last. This gives the production possibility frontier (PPF) a concave shape, as shown in Figure 1.

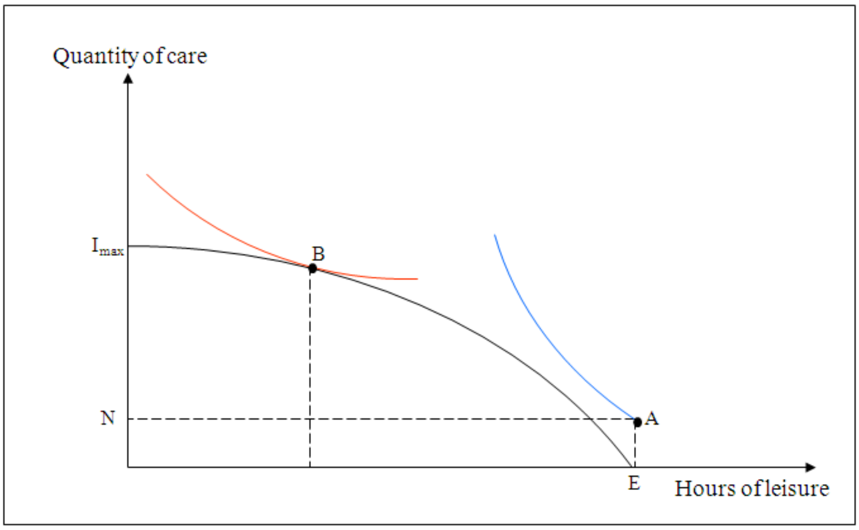
The concave shape also is motivated by the idea that the value, to either parent or child, of an additional hour of care input by the child is diminishing. For example, the first visit in a week by the child to make a meal for the parent is valued more highly than the second, etc. The slope of the indifference curves of the carer represents her preferences over the two goods or the carer-caree joint preferences.

Although this model requires only one type of care, in reality the types of care provided formally and informally may differ. State provision of formal care could allow for specialisation in different types of care that could lead to efficiency gains. Formal carers could specialise in bathing or personal hygiene maintenance, for example, whereas informal carers could specialise in more appealing types of care such as meal preparation.

For simplicity, other uses of the carer's time such as labour supply are ignored. This is a realistic assumption as more than half of the sample used in the empirical analysis in this paper does not work. We also ignore money and thus the ability to buy additional care in the home. In addition, we assume zero care in the home before the policy change, ignoring the existing social services network.

The original endowment is 24 hours of leisure and 0 hours of care, labelled as point E in Figure 1. We assume an exogenous level of care that the client would receive if he were to enter a nursing home (N), expressed in the same care units as those provided by the child. We also assume that if the client is in a nursing home, the carer does not provide any informal care. The maximum level of care that can be received by the parent before the policy change is I_{max}. If we denote the care production function as I(.), then for any level of informal care supplied by the child, E_0, the parent receives quantity of care $I(24-E_0)$.

Figure 2. Preferences in the supply of care



As shown in Figure 2, whether a child supplies informal care depends on the shape and location of her indifference curves. If, like Person A, the individual has steep indifference curves due to a high preference for leisure (or a low PPF due to low care productivity), it is likely that her highest indifference curve will not touch the PPF and she will supply zero hours of care. This means that the parent will be in a nursing home (a corner solution). If the indifference curves are flatter (Person B), it is likely that she will reach her highest indifference curve by supplying some informal care.

Figure 3. The shift in the PPF

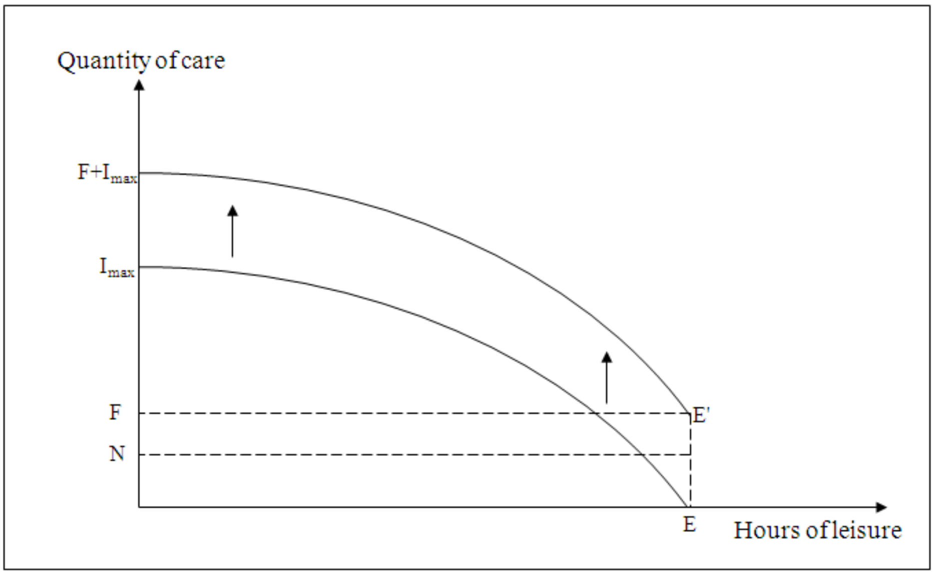
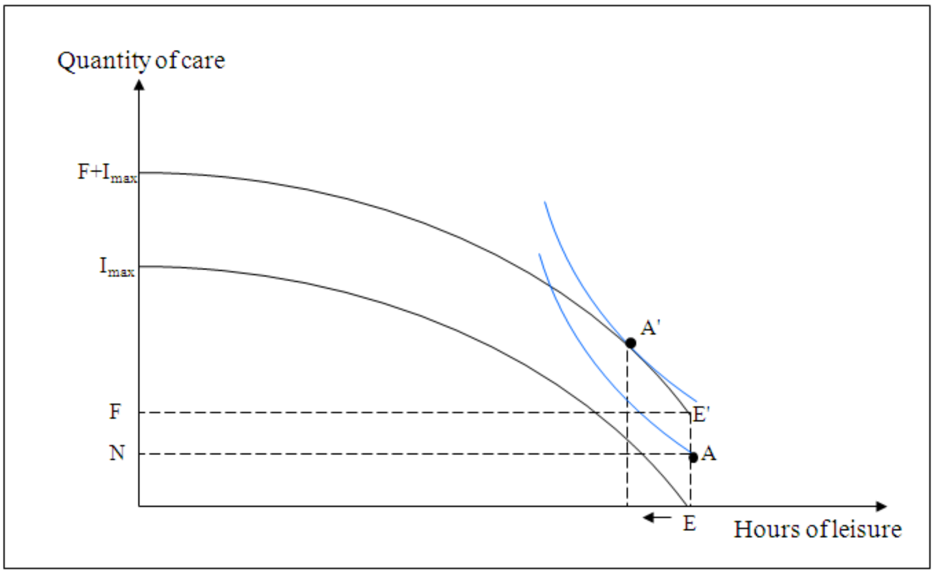


Figure 3 shows the upward shift in the PPF as free personal care is introduced. We assume that the level of care provided in the home by the state (F) is greater than that available in a nursing home (N). This reflects the high value that both carer and client put on the client remaining in the home. The new endowment point is E'. Now, instead of receiving $I(24-E_0)$ for each level of informal care input, the parent receives $F + I(24-E_0)$.

The outcome of the re-optimisation of the carer in response to the policy change depends on the shape of her indifference curves. The response of Person A is shown in Figure 4. Provision of free domiciliary formal care now allows the client to stay in his own home. The carer may then supplement the formal care by providing some hours of informal care herself. The new equilibrium is therefore point A'. Thus Person A participates in care provision as a result of the policy change. Later on in this paper, the participation decision is explored empirically.

Figure 4. The participation decision

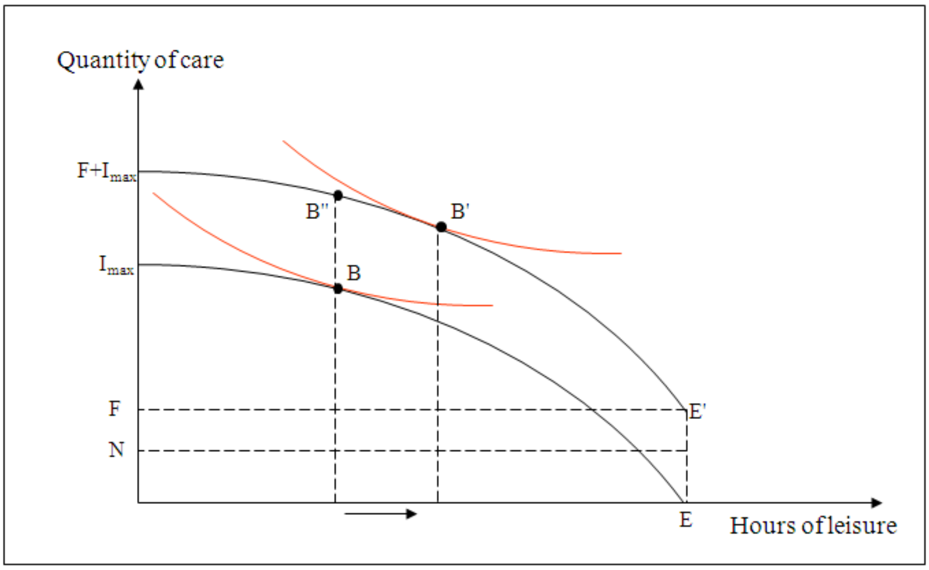


The reaction of Person B to the policy change is modelled in Figure 5. The decision of Person B depends on the interaction between the income and substitution effects and their relative strengths. The upward shift of the PPF is like an increase in total wealth, deriving from the state provision of care in the home, F. If leisure is a normal good, this will cause the carer to consume more of it. This substitution away from care and towards leisure is known as the income effect. The highest indifference curve of the carer may now lie tangent to the production PPF at B', for example, where she supplies less care, but the parent receives more in total.

The concave shape of the PPF means that its slope increases as we move to the right. For this reason, if there is a positive income effect, there will also be a substitution effect. This occurs because each hour of care supplied by the child produces a higher quantity of care for the parent. This gives the parent the incentive to increase her care supply. Therefore, the income and substitution effects work in opposite directions. If there is a positive income effect it will be attenuated by the substitution effect.

Depending on preferences, the new consumption bundle might lie anywhere on the new PPF between B and E'. This includes B'' where the income effect is equal to zero and the child supplies the same amount of care as she did previously. However, if the parent moves to B'' where the income effect is zero, the substitution will also be zero because the slope of the new PPF is the same as the slope of the original PPF at that point. In the empirical analysis, the intensity of hours supplied is analysed using the BHPS for individuals who begin a different points on the diagram.

Figure 5. The hours of care decision



Data

This paper uses the British Household Panel Survey from the years 1996 to 2008. This includes six years before and after the policy change. The BHPS is a well-known longitudinal study with a common structure for all regions of the UK and a boosted sample for Scotland (BHPS, 2009).

The selection of households for the survey was made using an approximately equal probability of selection method (EPSEM) design. This is achieved by implicitly stratifying the population of addresses into an ordered listing by region and a number of socio-demographic variables, derived from the 1981 census (Taylor et al, 2010).

Despite this, the survey includes design weights to adjust for unequal selection probabilities of addresses. As this paper performs longitudinal analysis at the aggregate level for Scotland versus the rest of the UK, cross-sectional probability weights are applied to the data. These weights are equal to the inverse of the probability of an observation being selected into the sample and are used to make the statistics computed from the data representative of the population with respect to the sampling strategy. This is important because the Scottish population is severely underrepresented in the years 1996 to 1998; the BHPS began the boosted sample for Scotland in 1999 (Taylor et al, 2010).

Table 1 shows the sample size by year and region, before and after the probability weights have been applied. Of these individuals, 38 percent provide either type of care at some point in the period.

The survey questionnaire requests information on whether or not one individual provides care to another. The question of supplying informal care is split into two categories: co-residential care, where the caree lives in the same household as the carer, and extra-residential care, where the caree lives independently. Where the type of care is co-residential, there are data on the age of the caree, but this information is not available for those who are cared for extra-residentially. Because the Scottish sample size is fairly small, those providing co-residential and extra-residential care have been aggregated into one “care” variable in the following analysis.

Table 1. Weighted and unweighted sample sizes

	Scotland		Rest of UK	
Year	**Weighted**	**Unweighted**	**Weighted**	**Unweighted**
1996	312	295	3947	3492
1997	423	401	4740	4283
1998	389	383	4618	4195
1999	584	1396	6702	5342
2000	572	1435	6650	5244
2001	708	1439	8163	6620
2002	607	1331	6695	5811
2003	582	1289	6724	5751
2004	542	1257	6037	5669
2005	578	1228	6490	5633
2006	608	1222	6433	5607
2007	596	1193	6272	5528
2008	661	1154	7123	5451

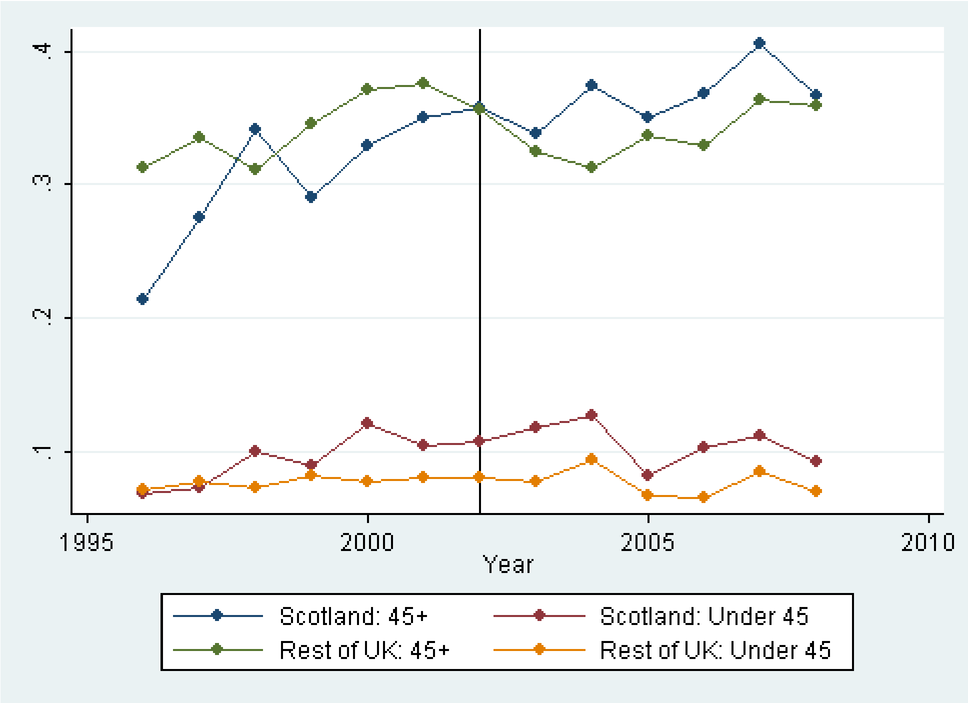
Source: Author’s calculations based on BHPS (2009)

The BHPS does not contain exact measures of the amount of time spent caring but contains information on the interval in which the number of hours supplied lies: 1-4, 5-9, 10-19, 20-34, 34-49, 50-99 or 100+.

As it is clear from the literature that the most common carer-caree relationship is between middle-aged children and their parents, the analysis is performed on a sample of individuals aged 45 or over, with no children in the household. This is different from the approach used in Bell, Bowes and Heitmueller (2006) and has the advantage of allowing the simple aggregation of co-residential and extra-residential carers, which is helpful given the small sample sizes.

Figure 6 shows the proportion of the sample that provides care in Scotland and the rest of the UK by age (using weighted data), where the black line indicates the start of the policy. For the difference-in-differences analysis, the blue line shows the proportion of carers in the treatment group and the green line shows the proportion of carers in the control group. The key identifying assumption for difference-in-differences analysis is that the two groups followed common trends before the introduction of the policy. Although this assumption appears to be satisfied from 1999 onwards, the proportion of carers is increasing at a faster rate in Scotland than in the rest of the UK before then. This issue is addressed in the final section of this paper

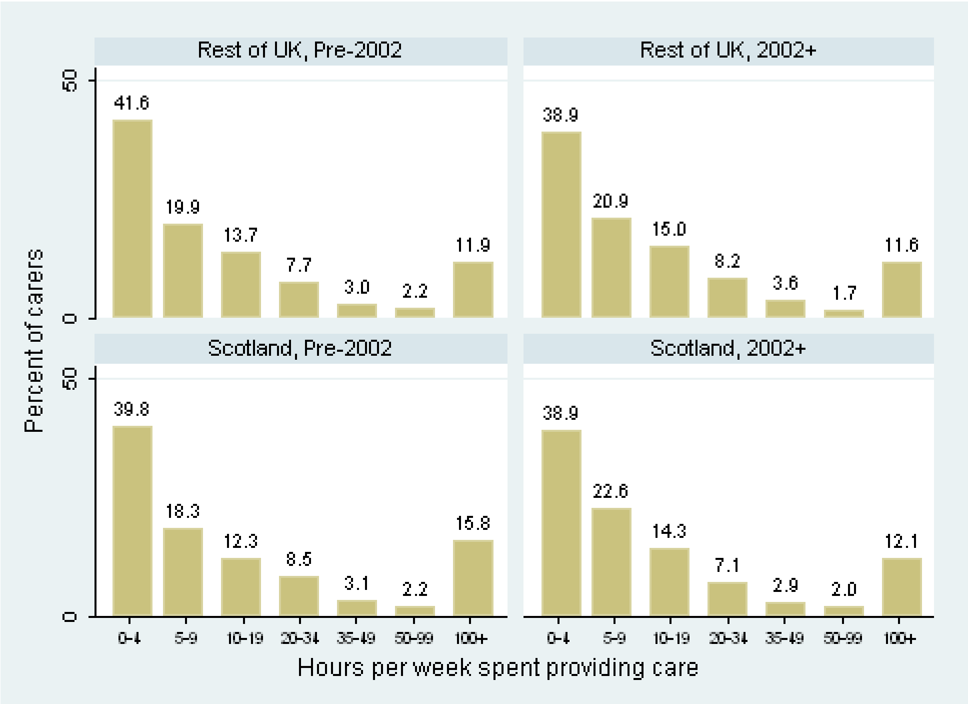
Figure 6. Proportion of carers in the BHPS



Source: Author's calculations based on BHPS (2009)

Figure 7 depicts the distribution of hours of care before and after the policy change in Scotland and the rest of the UK for those aged 45 and over. The distributions are similar in both areas, but there is a shift towards the lower tails in Scotland that is not apparent in the rest of the UK. From simply looking at the data, this appears to have come partly from a shift away from the upper tail, where the proportion of carers supplying 100 or more hours fell from 16 percent to 12 percent. The distributional changes are explored econometrically in the discussion of the intensity of care in this paper.

Figure 7. Hours of care distribution in the BHPS



Source: Author's calculations based on BHPS (2009)

Carers in the sample are typically low-educated and aged between 45 and 65, as shown in Table 2. They are most commonly retired, though a large proportion is in paid employment; 8.6 percent state their employment status to be full-time family care (see Figure 8).

Table 2. Characteristics of carers

Highest academic qualification	Percent	Age	Percent
Higher degree	1.85	45-54	29.51
First degree	5.57	55-64	34.44
HND, HNC, teaching qual	6.25	65-74	22.95
A level	9.57	75+	13.07
O level	17.46		
CSE	1.10		
None of these	58.19		

Source: Author's calculations based on BHPS (2009)

Methodology and Results

Following Angrist and Pischke (2009), we begin by examining the conditional expectation of the amount of care provided by individuals. The conditional expectation of the number of hours of care ($\mathrm{H_{it}}$) can be expressed as a combination of a participation decision (whether to supply hours) and an intensity decision (how many hours to supply):

$$E[H_{it}|D_{it}] = E[H_{it}|H_{it} > 0, D_{it}\,]P[H_{it} > 0, D_{it}] \qquad [1]$$

where D_{it} = 1 if the observation is from the treatment group (Scotland after 2002) and D_{it} = 0 if it is from the control group.

The difference in the expected number of care hours between the treatment and control groups is thus:

$$E[H_{it}|D_{it} = 1] - E[H_{it}|D_{it} = 0] \qquad [2]$$

$$= E[H_{it}|H_{it} > 0, D_{it} = 1]P[H_{it} > 0, D_{it} = 1] - E[H_{it}|H_{it} > 0, D_{it} = 0]P[H_{it} > 0, D_{it} = 0]$$

$$= \underline{\{P[H_{it} > 0, D_{it} = 1] - P[H_{it} > 0, D_{it} = 0]\,\}} \quad E[H_{it}|H_{it} > 0, D_{it} = 1]$$

$$+ \underline{\{E[H_{it}|H_{it} > 0, D_{it} = 1] - E[H_{it}|H_{it} > 0, D_{it} = 0]\}} \quad P[H_{it} > 0, D_{it} = 0]$$

Therefore, the overall difference in average hours may be broken down into two separate effects. The first is the participation effect, which is the difference in the probability that the hours are positive (dotted underline). The second is the "conditional on positive effect" (COP) effect, which is the difference in means conditional on participation (wavy underline) (Angrist and Pischke, 2009).

According to the model outlined in the previous chapter, we would expect the participation effect to be positive (Person A) and the COP effect to be ambiguous (Person B).

The participation decision

The following equation is used to estimate the effect of free personal care for the elderly on the probability of supplying informal care:

$$p_{it}^{*} = \delta t + \varphi S + \gamma (t * S) + \varepsilon_{it} \qquad [3]$$

where p_{it}^{*} is the latent probability that individual *i* provides care to someone aged 65 or over in period *t*, *S* is a dummy variable which indicates whether the observation is from Scotland, *t* is a dummy variable which indicates whether the observation is after 2002 and ε is an idiosyncratic error term. All regressions are clustered at the individual level.

This is the classic difference-in-differences formulation. The coefficient of interest is γ: a positive coefficient would suggest that the policy caused the probability of providing care to increase and a negative coefficient would suggest the opposite.

We begin with a probit specification, where ε_{it} is distributed $\mathcal{N}(0, \sigma_{\varepsilon}^{2})$ and the latent index model is $p_{it} = 1[p_{it}^{*} > 0]$. This type of non-linear estimation is appropriate as the probability of providing care is a limited dependent variable—it may only take non-negative values and may not be greater than one.

We then estimate the equation using a linear probability model. Although the LPM is criticised for producing biased marginal effects if the conditional expectation function (CEF) is non-linear, its approximations are often very similar to those obtained using maximum likelihood methods (Angrist and Pischke, 2009). The results therefore serve as a useful check of robustness.

Both the probit and the LPM are first estimated without controls for personal characteristics. The results are report in columns (1) and (3) of Table 3, below. The equation is then estimated using a vector of controls for personal characteristics including marital status, employment status, education level, household size, sex, age and income. The coefficients on these controls are specified as common to all individuals in the sample. These results are reported in columns (2) and (4) of Table 3.

As the policy may have affected certain groups of individuals with particular characteristics differently, the estimations are repeated with first-order interactions between covariates and the various dummy indicators. This structure allows for nonlinearity in the treatment effect due to differences in the covariates.

Using fully interacted controls is important if the Scottish sample differs compositionally from the rest of the UK. For example, suppose the policy affects the care-supply decision of educated and non-educated people differently. If there is a change in the share of educated people in either Scotland or the rest of the UK after the policy, the average treatment effect will be biased if the controls are not interacted.

The regressions now take on the following form:

$$p_{it}^{*} = \delta t + \varphi S + \gamma(t * S) + \alpha(X_{it} * t) + \tau(X_{it} * S) + \mu(X_{it} * S * t) + \varepsilon_{it} \quad [4]$$

where X is the vector of personal characteristics.

The treatment effect for person *i* is estimated by

$$\hat{\beta}(X_{it}) = \hat{\gamma} + \hat{\mu}(X_{it}|S_{it} = 1) \quad [5]$$

and the average treatment effect of the policy may then be estimated by the sample average of $\hat{\beta}(X_{it})$:

$$ATE = E\left[\hat{\beta}(X_{it})\right] = \hat{\gamma} + \hat{\mu} * E(X_{it}|S_{it} = 1) \quad [6]$$

We then may do a Wald test of Equation 6 where the means of X for the trended population are replaced with the sample means. The results are reported in column (5) of Table 3. As the treatment effects from the LPMs using the two different types of controls (columns 4 and 5) are very similar, no probit model with interacted controls is estimated.

All specifications produce very similar results: an average treatment effect of around 5 percent, significant at the 1 percent level. This suggests that the introduction of free personal care for the elderly in Scotland resulted in an increase of 5 percentage points in the probability of over-45s providing informal care.

These findings may be interpreted in the context of the model outlined previously in the section on the modelling framework. It appears that some of the individuals in the sample exhibited behaviour similar to Person A, whose steep indifference curves located far to the right meant that her utility was maximised by not providing care before the policy change. After 2002, carers provided by the state could supply some care, while informal carers supplemented this with care of their own, allowing more elderly people to stay in their own homes.

Table 3. Participation results

	(1) Probit	(2) Probit	(3) LPM	(4) LPM	(5) LPM
Treat	**0.0514***** **(3.29)**	**0.0480***** **(3.22)**	**0.0522***** **(3.40)**	**0.0487***** **(3.34)**	**0.0451***** **(2.91)**
Scotland	-0.0111 (-0.78)	-0.0122 (-0.87)	-0.0111 (-0.78)	-0.0123 (-0.90)	0.1977 (0.96)
After	-0.0155** (-2.35)	-0.0135** (-1.98)	-0.0154** (-2.35)	-0.0126* (-1.92)	0.0874 (0.53)
Observations	60122	55155	60122	57424	57472
R^2 /Pseudo R^2	0.000	0.017	0.000	0.119	0.124
Controls	No	Common	No	Common	Interacted

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Notes: Marginal effects; *t* statistics in parentheses; observations are weighted

The positive treatment effects found in this work contradict the results of who found no statistically significant impact on participation. Possible explanations for this include the greater number of years used. We use data from the years 1996-2008, whereas Bell, Bowes and Heitmueller (2006) use only 1999-2003. Using this extended sample may have allowed us to pick up any lagged effects of the policy change. As the proposed motivating theory includes changes to living arrangements, such a delayed effect might be expected.

In addition, there are differences in the definitions of the supply of care and the samples used. Bell, Bowes and Heitmueller (2006) analyse the effect on the supply of co-residential and extra-residential care separately, whereas this paper aggregates the two types of care into one variable. This is justified in this study by the small Scottish sample and the small number of people supplying care within it. Treating the two care settings separately, as is done by Bell, Bowes and Heitmueller (2006), may have led to larger standard errors and thus a more imprecise estimate.

Only individuals aged 65 or over are entitled to free personal care in Scotland. As the BHPS does not include information on the age of individuals who are cared for extra-residentially, attempt to infer the age of the caree using the age of the carer and the relationship between caree and carer. Instead, as the most common carer-caree relationship is between middle-aged children and their parents, this paper uses only individuals aged 45 and over with no children in the household.

Finally, we use individual fixed effects and look for variations in care behaviour at the individual level whereas this paper uses aggregate longitudinal variation. The latter type of analysis may produce more precise estimates as fixed-effects estimates can suffer from large measurement error problems in attempting to detect changes at the individual level (Angrist and Pischke, 2009). Measurement error in the explanatory variable leads to attenuation bias, meaning that the coefficients may be biased towards zero. This could explain the insignificant findings in the study by Bell, Bowes and Heitmueller (2006).

The intensity of care decision

To analyse the decision of how many hours of care to provide and the response to the change in policy, it is necessary to decompose the COP effect derived at the beginning of Section 6 (Angrist and Pischke, 2009):

$$\{E[H_{it}|H_{it} > 0, D_{it} = 1] - E[H_{it}|H_{it} > 0, D_{it} = 0]\} \qquad [7]$$

$$= E[H_{1it}|H_{1it} > 0] - E[H_{0it}|H_{0it} > 0]$$

$$= \quad E[H_{1it} - H_{0it}|H_{1it} > 0] \quad + \quad \{E[H_{0it}|H_{1it} > 0] - E[H_{0it}|H_{0it} > 0]\}$$

It contains two separate terms. The first is the causal effect on hours for the subpopulation of the treatment group that supplies care (dotted underline). The second term is the difference between the expected hours of those in the control group given they would supply positive hours in the treatment group (an unobservable counterfactual) and given they would supply positive hours in the control group. This is the selection effect (wavy underline).

The presence of the selection effect means that the COP effect does not necessarily have a causal interpretation (Angrist and Pischke, 2009). As we cannot observe the care hours of those who choose not to participate, regressions using a sample limited to individuals who supply positive hours are likely to produce biased results. This selection bias is likely to occur because the policy changes the composition of the group who supply positive care hours. As those who begin to participate as a result of the policy are likely to supply fewer hours (they have a lower “taste for care”), the selection bias term is expected to be negative.

The censored regression model

The traditional method by which this problem is often addressed is to use a censored regression model such as Tobit. As the data do not contain information on the exact number of care hours, it is not possible to use a straightforward Tobit model with a continuous latent dependent variable. Instead, a generalisation of the Tobit model–interval regression–is used here.

In the first of these censored specifications, it is assumed that the continuous latent dependent variable (number of hours of care) may take values below zero, but that we only observe hours of care which are zero or greater, that is, the dependent variable is left-censored. This may be likened to the traditional labour supply case in which some individuals have preferences such that they would prefer to supply negative hours of labour, but we only observe them supplying zero (being unemployed).

Because the result on latent hours may be difficult to interpret as the concept of negative latent hours of care is quite obscure, a second specification is estimated where only real hours are used, that is, care hours are assumed to be non-negative.

A more detailed exposition and discussion of the interval regression model and its results are presented in the Appendices. To summarise, the results of the left-censored model suggest that the introduction of free personal care for the elderly had a positive effect on the average number of hours of care supplied by informal carers of around four additional hours. However, the left-uncensored model yielded results insignificantly different from zero.

The censored regression models discussed above produce estimates of the average effect of the policy on the number of hours of care supplied. However, it is likely that the increased provision of formal care will have affected individuals differently depending on where in the hours of care distribution they lay before the policy change. For example, those at the high end of the distribution may have chosen to decrease their weekly hours as they were able to enjoy more leisure time without compromising the quantity of care their dependent receives. On the other hand, those who did not supply care before the policy change are likely to have supplemented the formal care with a relatively low number of hours of care. These effects are masked in the censored regression models.

The participation distribution model

An alternative approach to analysing the hours of care decision is to estimate distributional effects directly using a procedure suggested by Angrist and Pischke (2009). This is done by defining a series of binary indicators for providing hours of care per week greater than or equal to 5, 10, 20, 35, 50 and 100. These become the dependent variables in a series of regressions of the same form as those outlined in the participation section (Equation 3) (Angrist and Pischke, 2009). Both probit and LPM models are estimated, with and without covariates. Only the simple (common) controls are used as the interactions made very little difference to the participation results, as discussed elsewhere in this paper.

Another advantage of this approach, pointed out by Angrist and Pischke (2009), is that there is no selection effect as all individuals are included in the analysis. There is also no longer ambiguity over the way left-censored observations should be treated. The probit results with controls are reported in Table 4. The LPM results and the probit results without controls are not reported as they are very similar.

These results suggest that there was an increase in the probability of supplying more than 5 hours of informal care per week of about 3 percentage points, significant at the 5 percent level (column (2)). Note that this is smaller than the 5 percentage point increase in the probability of supplying any positive care hours (column (1)). There does not appear to be a statistically significant change to the higher parts of the distribution (columns (3)-(7)).

Table 4. Hours of care distribution

	(1) Partici- pation	(2) 5+	(3) 10+	(4) 20+	(5) 35+	(6) 50+	(7) 100+
Treat	**0.0480***** **(3.22)**	**0.0273**** **(2.45)**	**0.0086** **(0.94)**	**0.0014** **(0.07)**	**-0.0001** **(-0.01)**	**-0.0007** **(-0.13)**	**-0.0011** **(-0.23)**
Scotland	-0.0122 (-0.87)	0.0086 (0.80)	0.0117 (1.36)	0.0146** (2.26)	0.0127** (2.51)	0.0108** (2.31)	0.0094** (2.15)
After	-0.0135** (-1.98)	0.0023 (0.44)	0.0045 (1.04)	0.0026 (0.73)	0.0024 (0.81)	0.0017 (0.64)	0.0018 (0.73)
Obs	55155	55729	55729	55715	55701	55701	55668
Pseudo R^2	0.017	0.061	0.070	0.089	0.110	0.114	0.117
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes

* p<0.10, ** p<0.05, *** p<0.01
Notes: Marginal effects; *t* statistics in parentheses; observations are weighted

These results may be explained as the combination of a number of different effects. First, those who chose to participate in informal care as a result of the policy change enter the hours of care distribution at the low end. In the context of the model in the previous section, this decision corresponds to the shift of Person A from point A to A', where A' is a small number of hours (in either the 0-4 or 5-9 hours range). Such a decision is likely given Person A's low "taste for care" or low care productivity.

In addition, some A-type individuals may begin to supply care at a higher number of hours, but this effect may be masked on average by other individuals who reduce their hours of care due to the income effect (Person B moving from point B to point B', for example). This would lead to an average treatment effect insignificantly different from zero at higher points in the distribution.

The Common Trends Assumption: Robustness Checks

For difference-in-differences estimation to produce unbiased results, the common trends assumption must hold. In this case, the key identifying assumption is that trends in care-giving would be the same in both Scotland and the rest of the UK if it was not for the change in policy. From Figure 2, there is reason to believe that care-giving trends may have been different in the years leading up to the introduction of free personal care for the elderly. The proportion of carers in Scotland appears to have been increasing at a faster rate than in the rest of the UK between 1996 and 1998. This observation, however, may be the product of noisy data resulting from of the small Scottish sample size in these years. From 1999 to 2002, the trends in the two parts of the UK are very similar.

Omitting 1996-1998

The first method by which the common trends problem may be resolved is simply to omit the years where the trends appear to be different from the sample. Therefore, the analysis from the Methodology and Results section is repeated without the years 1996, 1997 and 1998 (excluding the censored regressions and the specifications with interacted controls).

The results of the participation regressions are reported in Table 5. The positive and significant treatment effect is robust to omitting the years 1996 to 1998, suggesting that the introduction of free personal care did indeed increase the probability of individuals supplying informal care. However, the magnitude of the coefficient falls from around 5 percent to between 3 and 4 percent.

Table 5: Probit participation results omitting 1996-1998

	(1) Probit	(2) Probit	(3) LPM	(4) LPM
Treat	**0.0392*** (2.81)**	**0.0327** (2.09)**	**0.0396*** (2.76)**	**0.0341** (2.52)**
Scotland	0.0014 (0.10)	0.0028 (0.21)	0.0015 (0.10)	0.0024 (0.18)
After	-0.0256*** (-3.83)	-0.0215** (-2.49)	-0.0256*** (-3.82)	-0.0208*** (-3.14)
Observations	50457	46734	50457	47840
R^2 /Pseudo R^2	0.000	0.000	0.053	0.122
Controls	Yes	Yes	Yes	Yes

* p<0.10, ** p<0.05, *** p<0.01'
Notes: Marginal effects; *t* statistics in parentheses; observations are weighted

The results of the regressions that investigate participation at different parts of the hours of care distribution are reported in Table 6. Only the probit regressions with common controls are reported as the results are very similar across the specifications. Again, the exclusion of the beginning years makes little difference to the results, suggesting that the violation of the common trends assumption may not have been a large problem. Like the participation results, the marginal effect on the probability of supplying more than 5 hours of care is reduced by around 1 percentage point.

Table 6: Hours of care distribution omitting 1996-1998

	(1) Participation	(2) 5+	(3) 10+	(4) 20+	(5) 35+	(6) 50+	(7) 100+
Treat	**0.0392*** (2.81)**	**0.0188* (1.80)**	**0.0079 (0.91)**	**0.0007 (0.10)**	**-0.0015 (-0.27)**	**-0.0027 (-0.54)**	**-0.0027 (-0.59)**
Scotland	0.0014 (0.10)	0.0170* (1.67)	0.0121 (1.45)	0.0147** (2.28)	0.0138*** (2.69)	0.0124*** (2.68)	0.0106*** (2.53)
After	-0.0256*** (-3.83)	0.0005 (0.09)	0.0017 (0.39)	0.0012 (0.33)	0.0025 (0.85)	0.0013 (0.47)	0.0014 (0.54)
Observations	50457	46407	46407	44602	45352	46380	46357
Pseudo R^2	0.000	0.063	0.072	0.097	0.118	0.118	0.119
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes

* p<0.10, ** p<0.05, *** p<0.01
Notes: Marginal effects; *t* statistics in parentheses; observations are weighted

Difference-in-difference-in-differences

Another way to address the potential bias resulting from the violation of the common trends assumption is to use a DDD estimator. In this specification, treatment and control groups are defined within Scotland.

As free personal care is available only to those aged 65 or over, we would expect only Scottish middle-aged "children" to be affected as their parents are eligible for the free care. This was the rationale behind keeping only those aged 45 and over in the sample in the previous estimations.

Here, individuals aged under 45 are added back into the sample. As the under-45s in the two regions capture the local trends in care giving, in effect this allows for different trends in the Scotland and the rest of the UK. To ensure that these individuals are caring for their parents and not disabled children, only those who report no children in the household are used in the analysis, as before.

The methodology used here is similar to that of Carpenter (2004) in which the effects of zero-tolerance alcohol laws for under-21-year-olds on various drinking behaviours are assessed (Carpenter, 2004). As the policy was implemented in some US states and not others, states that did not enact the laws served as the control group for the simple difference-in-differences estimation. However, as the states did not have common trends in the years leading up to the policy change, Carpenter (2004) defines those aged 21 to 24 as the treatment group within the treatment group. As drinking behaviours within states followed common trends, the coefficient on the interaction between the "zero-tolerance state" dummy, the "under-21" dummy and the "after policy implementation" dummy gives the DDD estimation of the effect of the policy.

Note, however, that Carpenter (2004) is not an ideal application of the DDD approach. Ideally, DDD contrasts groups that are assumed to have identical underlying trends. In Carpenter's study (Carpenter, 2004), for example, under-21s are assumed to follow the same trends in drinking behaviour as those aged 21 to 24. This is plausible given the similarities between the two groups. It is a stronger assumption in the application in this paper as the differences between under-45s and over-45s are substantial. For this reason, including controls is very important. In particular, it is necessary to allow for the covariates to affect over-45s and under-45s differently.

The estimating equation for the DDD participation regression is as follows:

$$p^*_{it} = \beta_1 t + \beta_2 S + \beta_3 G + \beta_4(t * S) + \beta_5(t * G) + \beta_6(S * G) + \beta_7(t * S * G) + \varepsilon_{it} \quad [8]$$

where G is a dummy variable that indicates whether the individual is aged 45 or older and the other variables are defined as before. The coefficient of interest is now β_7, as the treatment group consists of observations dated after 2002, in Scotland and aged 45 or older.

This equation is estimated using both a probit and linear model, again with and without covariates. As in the section in this paper on participation decisions, the control variables are interacted with the indicators for the different groups of individuals. Although this made little difference to the results in that section, it is important in the DDD specification as the covariates are likely to affect the behaviour of over-45s and under-45s differently. The results are reported in Table 7, below. Columns (2) and (4) give the results of the estimation where the covariates are interacted with only the over-45 indicator and columns (5) gives the results of the fully interacted model.

Table 7: DDD participation

	(1) Probit	(2) Probit	(3) LPM	(4) LPM	(5) LPM
Treat	**0.0231 (1.04)**	**0.0373* (1.65)**	**0.0381* (1.88)**	**0.0473** (2.46)**	**0.0800 (0.90)**
Scotland*After	0.0193 (1.06)	0.0014 (0.08)	0.0141 (1.08)	0.0014 (0.11)	0.3365 (1.12)
Scotland*Over 44	-0.0317* (-1.71)	-0.0425** (-2.22)	-0.0262 (-1.52)	-0.0321** (-1.97)	0.3741 (1.26)
After*Over 44	-0.0122 (-1.23)	0.0326** (2.18)	-0.0151* (-1.79)	0.0198** (2.12)	0.3287 (1.31)
Scotland	0.0225 (1.54)	0.0047 (0.59)	0.0151 (1.49)	0.0034 (0.69)	0.0726 (0.34)
After	-0.0006 (-0.07)	-1.4298*** (-5.79)	-0.0004 (-0.07)	-0.5164***	0.1467 (0.70)
Observations	103431	96587	103431	98856	98856
Controls	No	Under 45/over 45	No	Under 45/over 45	Fully interacted

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Notes: *t* statistics in parentheses; observations are weighted

The DDD results for participation are broadly similar to the difference-in-differences results. There is a positive treatment effect, indicating that the introduction of free personal care had a positive effect on informal caring in Scotland, net of the general Scottish trend.

However, there are some notable differences. The results in columns (1) to (4) of Table 7 are of a smaller magnitude and are less statistically significant. While the difference-in-differences probit marginal effects were around 5 percentage points and significant at the 1 percent level, the DDD probit estimates are around 3 percent and only significant at the 10 percent level when covariates are included. Similarly, the LPM DDD treatment effect is around 4 percent, compared to 5 percent from the straightforward difference-in-differences estimation. The LPM results are significant at the 10 percent level without controls and the 5 percent level with controls, as opposed to the previous 1 percent for both specifications.

The coefficient for the fully interacted model in column (5) of Table 7 is of a greater magnitude than the difference-in-differences results. However, it is not statistically significant. This is likely to be the product of the very high number of variables in the model (300) which has caused the standard errors to be large and the coefficient to be imprecisely determined[1].

Finally, the DDD estimator is used to examine the probability of supplying care at the different parts of the hours of care distribution. Because of the limitations of the fully interacted model described above, this specification is not estimated. As the probit and LPM results with and without covariates are very similar, only the probit results with covariates are reported in Table 8 below.

Table 8: DDD hours of care distribution

	Partici-pation	5+	10+	20+	35+	50+	100+
Treat	**0.0373* (1.65)**	**0.0324** (1.97)**	**0.0130 (0.91)**	**0.0068 (0.56)**	**0.0054 (0.48)**	**0.0028 (0.25)**	**-0.0003 (-0.03)**
Scotland *After	0.0014 (0.08)	-0.0117 (-0.84)	-0.0069 (0.91)	-0.0060 (-0.54)	-0.0052 (-0.51)	-0.0030 (-0.29)	-0.0004 (-0.04)

[1] The standard error of the estimated treatment effect is 0.089.

	Partici-pation	5+	10+	20+	35+	50+	100+
Scotland *Over 44	-0.0425** (-2.22)	-0.0206 (-1.58)	-0.0112 (-0.55)	-0.0106 (-1.18)	-0.0062 (-0.84)	-0.0042 (-0.61)	-0.0015 (-0.23)
After *Over 44	0.0326** (2.18)	0.0285** (2.77)	0.0224 (-1.01)	0.0226** (2.88)	0.0163** (2.62)	0.0005 (0.10)	-0.0017 (-0.39)
Scotland	0.0047 (0.59)	0.0083 (1.17)	0.0079 (-1.01)	0.0008 (0.16)	0.0040 (0.87)	0.0130** (2.16)	0.0093 (1.54)
After	-1.4298*** (-5.79)	0.0212 (0.23)	0.1581 ** (2.44)	0.0008*** (4.44)	0.3490 .	0.0006 (0.14)	0.0029 (0.75)
Obs	103431	96915	96851	96644	96212	93928	93828
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes

* p<0.10, ** p<0.05, *** p<0.01
Notes: Marginal effects; *t* statistics in parentheses; observations are weighted

Again, the results are very similar. Like the original difference-in-differences results and the results omitting the beginning years, the only part of the distribution at which there is a change in the probability of participation is providing 5 or more hours of care per week, aside from the probability of supplying positive hours. The DDD is almost identical to the difference-in-differences estimate--and is within the same 95 percent confidence interval, at around 3 percentage points. These results provide further evidence of a shift towards the lower tails of the hours of care distribution and are consistent with the theory of informal carers supplementing the care provided by the Scottish state.

Conclusions

When free personal care for the elderly in Scotland was introduced in 2002, policy critics were concerned that it would cause informal carers to reduce their supply of care resulting in a limited net benefit. This paper finds evidence of the opposite effect.

The difference-in-differences and DDD results suggest that the policy increased the probability of individuals supplying informal care by 3 to 5 percentage points.

Of the sample of Scottish individuals aged over 16 interviewed for the BHPS, around 45 percent were potentially "treatable" in 2002 as defined by this paper (aged 45 or over and with no children in the household)[2]. As the population of Scottish over-16s was 4.1 million in 2002 (Macniven, 2011), the potentially treatable population was around 1.8 million. Therefore, an increase in the probability of providing care of 5 percentage points amounts to 90,000 individuals opting into supplying care as a result of the policy. As the 95 percent confidence interval is from around 2 to 8 percentage points, this figure could be anywhere from 36,000 to 144,000 people.

The rise in informal care participation as a response to increased provision of formal care suggests that the two types of care act as complements in this setting. Although this contradicts much of the empirical literature discussed in above, it follows from the fact that the policy increased state provision of care in the home. As free personal care in the home allowed more elderly people to stay in their own homes, friends and family supplemented the formal care with care of their own.

Not only did informal care participation increase, but individuals who were supplying informal care before the policy change did not reduce their hours. The evidence suggests that those who were moved into participating in care supply as a result of the policy provided a small number of hours of care per week as the probability of supplying 5 or more hours of care increased by 1 to 3 percentage points.

If, as assumed in the discussion of the modelling framework, the gains to care are highest for the first hours of care supplied, such an increase in participation at the low end of the distribution may have resulted in substantial gross welfare gains. Whether

[2] This figure is derived using probability weights.

these gains outweigh the high costs described at the beginning of this paper is a topic for future work.

Appendices

A1. The left-censored model

In the standard Tobit model, it is assumed that the continuous latent dependent variable, H^*, may take values below zero but that we only observe hours of care that are zero or greater.

In the interval regression model, for each observation the dependent variable may be point data, right-censored or both left- and right-censored. This is a generalisation of the standard Tobit specification where the dependent variable may only be point data or left-censored.

In the application here, the dependent variable is right-censored if the number of hours of care is reported as 100+, it is both left- and right-censored if it is reported as within one of the intermediate intervals used for the survey responses and it is left-censored if it is reported as zero.

The observed interval, $\overline{H}_{it}$, is generated by

$$\overline{H}_{it} = 1[H_{it}^* > 0]H_{it}^* \qquad [A1]$$

where H_{it}^* is a normally distributed and continuous latent variable which may be negative. As H_{it}^* is not a limited dependent variable, it may be estimated linearly using the following specification:

$$H_{it}^* = \delta t + \varphi S + \gamma(t * S) + \beta X_{it} + \varepsilon_{it} \qquad [A2]$$

The above equation is defined for all values of H_{it}^*, including those that are negative and within the other various intervals. It is estimated using the Stata "intreg" command (StataCorp, 2007).

The coefficient is estimated by maximising the following log likelihood function:

$$\ln\mathcal{L} = -\frac{1}{2}\sum_{i\epsilon C} w_i\left\{\left(\frac{H_i - x\beta}{\sigma}\right)^2 + \log 2\pi\sigma^2\right\} + \sum_{i\epsilon L} w_i \log\Phi\left(\frac{H_{Li} - x\beta}{\sigma}\right)$$
$$+ \sum_{i\epsilon R} w_i \log\left\{1 - \left(\frac{H_{Ri} - x\beta}{\sigma}\right)\right\} + \sum_{i\epsilon X} w_i \log\left\{\Phi\left(\frac{H_{2i} - x\beta}{\sigma}\right) - \Phi\left(\frac{H_{1i} - x\beta}{\sigma}\right)\right\} \quad [A3]$$

where for observations $i\epsilon C$, we observe H_j (point data), for observations $i\epsilon L$, we know only that $H_i{}^*$ is less than or equal to H_{Li} (left-censored), for observations $i\epsilon R$, we know only that $H_i{}^*$ is greater than or equal to H_{Ri} (right-censored). $\Phi()$ is the standard cumulative normal and w_i is the weight for the ith observation (StataCorp, 2007).

A2. The left-uncensored model

However, as suggested by Angrist and Pischke (2009), the results of the effect on latent hours may be difficult to interpret as the concept of negative latent hours of care is quite obscure. The number of hours of care supplied by an individual really is zero in many cases, so a censored regression model may not be necessary. In addition, the Tobit formulation imposes strict distributional assumptions on H_{it}^*. Therefore, an alternative that is readily interpreted is to estimate the effect on real hours. Angrist and Pischke (2009) suggest treating observations of zero hours as zeros rather than as censored observations. The estimating equation takes on the same form as Equation A2, except the latent dependent variable may not be negative. The results of this specification are reported in columns (3) (no controls) and (4) (controls) of Table A.

A3. Interval regression results

Table A. Interval regression results

	(1) Left-censored	(2) Left-censored	(3) Left-uncensored	(4) Left-uncensored
Treat	**4.335*** **(1.95)**	**4.362**** **(1.97)**	**0.0423** **(0.06)**	**0.228** **(0.35)**
Scotland	1.839 (0.84)	1.779 (0.81)	1.429** (2.19)	1.398** (2.17)
After	-1.565 (-1.64)	-0.624 (-0.63)	-0.272 (-1.01)	0.139 (0.49)
Observations	58430	55838	58430	55838
Controls	No	Yes	No	Yes

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$
Notes: *t* statistics in parentheses; observations are weighted

The results of the left-censored model suggest that the introduction of free personal care for the elderly had a positive effect on the average number of hours of care supplied by informal carers. The interval regression results may be interpreted in a similar manner to OLS (StataCorp, 2009): those in the treatment group supplied an additional 4.3 hours of care on average than those in the control group, all else equal. This result is significant at the 10 percent level without controls and at the 5 percent level when covariates are included. The positive effect of the policy is a combination of increased participation in care and shifts in hours of those who were already providing care before the policy change.

However, there appears to be no statistically significant change in real hours of care following the introduction of free personal care in the home. A possible explanation for this is that when the model is limited to non-negative hours, a shift from supplying zero hours to supplying 1–4 hours, for example, is a small change in absolute terms, which may be masked by shifts in the higher part of the distribution. If latent hours instead are allowed to be negative, a shift from supplying a large "negative number of hours" to supplying 1–4 hours is a large change in absolute terms. Using the left-censored models allows us to incorporate into the estimated coefficient the effects of the policy on individuals with strong preferences against providing care.

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