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ABSTRACT

We investigate the impact of more generous terms for delaying state pensions on claiming and labor supply in the United Kingdom using a 2005 policy change. First, we find that the more generous delay terms reduced the fraction of males receiving pensions at the earliest eligibility age and shortly after. While there are also post-policy changes in women's claiming behavior, further investigation reveals that these changes do not coincide with the start of the policy and are therefore less likely to be causal effects. Second, we find post-policy increases in labor supply around the earliest pension eligibility age, followed by post-policy decreases in labor supply at older ages. While these labor supply changes cannot easily be separated from longer-term trends, they are consistent with some individuals choosing to work longer to finance pension delay, followed by some individuals retiring earlier due to the income effect from more generous pension benefits. Finally, we find that among individuals who delayed pensions for up to 5 years, about 3 percent of individuals took their gains from delay as lump sums, an option made available under the policy changes.

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I. Introduction

Throughout developed countries, aging populations have been putting pressure on public pensions and other social programs. In response, policy makers have taken steps to raise pension eligibility ages and boost tax revenue by encouraging longer working lives. In the United Kingdom, policy makers have begun gradually increasing state pension eligibility ages for both men and women.[1] In addition, starting in April 2005, the terms for delaying pension claiming became more generous. These more generous terms included providing a larger increase in benefits for each month of delay, allowing individuals to delay for an unlimited number of months, and allowing individuals to receive the gains from delay as a lump sum. In this paper, we empirically examine the impact of these more generous deferral options on pension claiming and labor supply choices among older workers.

To our knowledge, we are the first to address this question empirically for the United Kingdom. Numerous papers have examined the impact of other changes to the U.K. pension system, including increases in state pension ages (Cribb, Emmerson, and Tetlow 2016; Blundell and Emmerson 2007) and the repeal of the earnings rule that reduced pension benefits for individuals who continued to work (Disney and Smith 2002), as well as general pension system incentives (Blundell, Meghir, and Smith 2004). Some researchers have quantified the financial incentives for delay or explored optimal claiming strategies under the claiming incentives that were introduced in 2005 (Kanabar and Simmons 2016; Dagpunar 2015; Farrar, Moizer, and Hyde

[1] The pension age for women, which was 60 for women born before April 5, 1950, has slowly risen to 65, to match that of men. In addition, the state pension age for both men and women began rising this year and will reach 67 for individuals born after March 6, 1961.

2012).[2] They all conclude that deferral is an optimal strategy for some of the population, but heterogeneity exists depending upon the predicted survival and tax rates that people face.

Similar questions have been explored empirically in other countries. Amilon and Nielsen (2010) examine the impact of pension deferral on labor supply in Denmark; they find a small increase in labor supply at age 65. In the United States, Social Security retirement benefits can be claimed at any age between ages 62 and 70, with higher monthly benefits payable for delayed claims. While numerous studies have established that the terms of delay are actuarially generous (see, e.g., Meyer and Reichenstein 2010, 2012; Sass, Sun, and Webb 2013; Shoven and Slavov 2014a, 2014b), the vast majority of individuals claim benefits well before age 70. Maurer, Mitchell, Rogalla, and Schimetschek (2016; 2017) examine the impact of paying the gains from delay as a lump sum on claiming behavior. Their survey evidence (Maurer, Mitchell, Rogalla, and Schimetschek 2016) suggests that paying the gains from delay as a lump sum makes delaying both claiming and retirement more attractive.

The policy experiment we focus on began in 2005. Between 1975 and 2005, individuals could delay their state pension for up to five years, receiving an additional 1/7 of 1 percent of the benefit payable at pension age for every week of delay. This amounted to an approximately 7.5 percent increase in benefits per year of delay. The Pensions Act of 1995 increased the generosity of deferral to 10.4 percent per year of deferral and removed the five-year limit, with changes to go into effect in 2010. However, in 2002, the Labour Government called for moving the implementation forward to 2006. Subsequently, the Pensions Act of 2004 brought the

[2] Moizer, Farrar, and Hyde (2018) examine the impact of deferral rules on the fiscal sustainability of the pension system.

implementation forward to April 6, 2005 and added a lump sum option (Thurley 2017). The lump sum is calculated by adding interest – at a rate that is at least 2 percent above the Bank of England base rate – to any forgone benefits during the delay period (Pension Service 2008). In addition, the lump sum is taxed at the tax bracket determined by other income, excluding the lump sum. Thus, delaying benefits to a period where individuals have lower earnings could results in significant tax savings.

Two key concerns with this policy experiment need to be addressed. First, state pension eligibility ages have also been rising. We separate the impact of the more generous adjustment for delay from changes in state pension eligibility ages by focusing on men born before 1953 and women born before 1950, whose state pension eligibility ages remained unchanged at 65 and 60 respectively. Second, subsequent legislation passed in 2014 made the terms of delay less generous and removed the lump sum option for individuals reaching state pension age on or after April 6, 2016. However, we limit our time period to 2015 and earlier. Thus, our sample restrictions allow us to focus on individuals who were affected only by the changes made in 2005.

Theoretically, making the terms for delay more generous should reduce the fraction of individuals claiming at the earliest eligibility age (65 for men and 60 for women). For later ages, the increased generosity of the actuarial adjustment should discourage claiming as well. On the other hand, the ability to receive the accumulated gains from delay as a lump sum may encourage claiming at older ages, particularly given that individuals seem to value lump sums more than actuarially equivalent annuities. Kanabar and Simmons (2016) suggest that, in the absence of borrowing constraints, those who defer would also choose less work. More generous deferral incentives increase lifetime income, which should increase leisure and decrease labor supply. The ability to finance consumption with borrowing makes it feasible to retire earlier while also delaying pension receipt to take advantage of the actuarial adjustment. However, the impact on labor supply is less clear if people face borrowing constraints. Delaying pension receipt may induce liquidity constrained individuals to stay in the labor force longer as a means to finance consumption during the delay period. The desire to work longer may induce labor supply changes at earlier ages as well. For example, younger individuals may increase their labor supply if continuous work experience improves their opportunities for remaining employed later in life.

Using a regression discontinuity approach and data from the British Labour Force Survey, we do not find evidence of an immediate, statistically significant impact on claiming or working at the earliest eligibility ages (60 for women and 65 for men). The regression discontinuity approach compares claiming and labor force status just before and after adoption of the policy, controlling for time trends in the before and after periods. However, an event study approach – comparing pension receipt and labor force status at each age during the full before and after periods – suggests that the policy may have had a longer-term effect, at least on male claiming behavior. In our preferred specification, which controls for a quadratic time trend, the post-policy probability of pension receipt is about 6.4 percentage points lower at age 65 for men. A year-by-year examination of the male claiming age distribution suggests that this shift is consistent with a causal effect of the policy, rather than longer-term trends in claiming ages. While there are also post-policy shifts in the probability of female pension receipt, these shifts do not coincide with the start of the policy and are less likely to be causal effects. We also observe post-policy increases in male and female labor supply around the earliest eligibility ages, followed by post-policy decreases in labor supply at older ages. However, the year-by-year disaggregation suggests that these changes in labor supply cannot easily be separated from longer-term trends that may not be adequately captured by the quadratic time trend.

Using the English Longitudinal Study of Ageing, we also provide evidence on the fraction of individuals who opt to take their gains from delay as a lump sum. Some of the previous literature suggests that lump sums may be particularly attractive (Maurer, Mitchell, Rogalla, and Schimetschek 2016; 2017). We find that among individuals who delayed pensions for up to 5 years about 3 percent of individuals took their gains from delay as lump sums. This proportion of individuals utilizing the lump sum option increases with the length of delay.

This paper is organized as follows. Section II describes our data and methodology. Section III presents our results. Section IV provides a discussion and conclusions.

II. Data and Methodology

a. Data

We use data from the UK Labour Force Survey (LFS) from 1993 to 2015. The LFS is the largest household survey in the UK, managed by the Office for National Statistics (ONS) in Great Britain and by the Northern Ireland Statistics and Research Agency (NISRA) in Northern Ireland. The survey provides detailed information on the UK labor market and is used as the source of official employment statistics for the UK. The LFS is conducted quarterly and the sample size in each quarter is about 90,000 individuals.

We restrict the sample to women born before 1950 and men born before 1954 so that the sample of women all face a pension age of 60 and men all face a pension age of 65.[3] Later cohorts faced increases in pension eligibility ages, which would confound the effect of the more generous actuarial adjustment. In addition, we exclude data from the first quarter of 2003 because of a data anomaly. The average share of individuals receiving a pension in our data is about 88 percent, whereas the share in the first quarter of 2003 is close to 62 percent, which is significantly different from the trends both before and after that quarter. We do not have an explanation for this anomaly.

The main outcome variables in this study are employment and pension claiming status. The LFS asks all individuals aged 16 and older about their economic activity. We define an individual as employed if he or she reports being an employee, self-employed, or a government worker. Individuals aged 69 and under are asked whether they are receiving any state benefits; follow-up questions request more specific information about what types of state benefits they are receiving. We define a pension claiming indicator based on whether the individual is receiving state pension benefits. Age is defined as the individual's age at the end of the reference week. Because the terms for delaying pensions became more generous in April 2005, we define the pre-intervention period from the first quarter of 1993 to the first quarter of 2005, and the post-intervention period from the second quarter of 2005 to the last quarter of 2015.

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[3] We do not know the exact birth year, so we estimate this as the survey year minus age. This will lead some individuals who are born in the cutoff years but interviewed before their birthday to be excluded. In addition, men born in December of 1953 faced a pension age that was up to 3 months later than age 65. Many of these men will be excluded since they are likely interviewed before their birthday, but there may be a few in the sample.

For our analysis of pension claiming, we restrict our sample to men aged 65 to 69, and women aged 60 to 69. Ages 60 and 65 are the earliest eligibility ages at which women and men, respectively, can claim pensions, and, as discussed above, the LFS questions about pension receipt are only asked of respondents age 69 and under. We further restrict the sample period to 1998 to 2015 since pension data are not consistently available prior to 1998. For our analysis of labor supply, we broaden our sample to men aged 60 to 80, and women aged 55 to 80. We perform the labor supply analysis on an expanded range of ages as the terms for pension deferral may affect labor supply both prior to eligibility and after claiming. As most people in the sample have retired by age 80, we drop individuals over the age of 80.

During our sample period, the LFS asks respondents about pension receipt, but not pension income. Thus, we cannot use it to determine how many individuals took the lump sum option in the post-intervention period. To understand lump sum choices, we use the English Longitudinal Study of Ageing (ELSA), a panel survey of individuals aged 50 and older and their spouses. The survey has been conducted every other year since 2002. We use data from 2002 through 2012. The ELSA does not directly ask about lump sum receipt; however, data are available on state pension income. We assume an individual took a lump sum if the pension amount in the first year of receipt is 150 percent or more than the average pension amount across all years for the same individual. This definition results in plausible values for the fraction taking a lump sum: almost nobody takes a lump sum prior to 2005. After 2005, significant shares of individuals above the earliest eligibility age take lump sums. We compute the fraction of first-time pension recipients at each age who take a lump sum.

Table 1 shows the summary statistics of the main variables for our sample. The number of observations for the entire sample is 2,297,250, among which about 28 percent have pension information. About 20 percent of the respondents are employed and 91 percent of pension aged respondents receive pensions. The average age is 67 years old.

The summary statistics are further divided by gender and pre-/post-policy in columns 2 to 5. We also provide employment and pension for those aged 60-64 and 65-69 which are the first 5 years after pension eligibility for women and men, respectively. For both men and women, the share of the employed increases for those aged 60-64 and 65-69 post-policy. The share receiving a pension declines at ages 65-69 for men, whereas it increases at ages 60-64 and 65-69 for women. In further analysis, we investigate whether the policy change is driving these differences.

b. Methodology

To test whether the policy impacts claiming or labor force participation at the earliest pension eligibility age, we begin with a regression discontinuity approach and estimate the following specification separately for men aged 65 and women aged 60:

$$y_i = \alpha + \beta_1 post_i + \beta_1 (yearq_i - 2005{:}Q2) + \beta_2 (yearq_i - 2005{:}Q2) * post_i + \beta_3 (yearq_i - 2005{:}Q2)^2 + \beta_4 (yearq_i - 2005{:}Q2)^2 * post_i + \epsilon_i \qquad (1)$$

In this equation, $post_i$ is equal to 1 if the observation is in the second quarter of 2005 or beyond and 0 otherwise, and $yearq_i$ is a time trend representing the year and quarter of the observation.

The dependent variable, y_i, indicates either pension receipt or work status and takes on a value of 1 if the individual is, respectively, receiving a pension or working. The coefficient of interest is β_1 which measures any discontinuous change in the dependent variable that occurred in the second quarter of 2005. That is, it measures the immediate post-policy change in the fraction of individuals at the earliest eligibility age who have claimed benefits or who are still working. We expect β_1 to be negative for claiming. It could be either positive or negative for working depending on whether the relative importance of the income effect versus the liquidity constraint effect. The equation above includes a quadratic trend, but we also estimate the equation with a linear trend.

The regression discontinuity approach measures the immediate impact of the policy on individuals who are newly eligible for pensions. However, because retirement is planned in advance, it is possible that individuals change their behavior gradually. It is also possible for the policy to affect the entire distribution of claiming and retirement ages. Effects at ages greater than 65 (for men) or 60 (for women) would most likely occur in the years after the policy, as most individuals reaching those ages at the time the policy was enacted had already claimed.[4] To examine these issues for claiming, we estimate the following specification:

$$y_i = \alpha + \sum_{a=a_0}^{a_1} \beta_a \, \mathbf{I}[age_i = a] * post_i + \sum_{a=a_0}^{a_1} \gamma_a \, \mathbf{I}[age_i = a] + year_i + year_i^2 + \varepsilon_i \quad (2)$$

[4] Individuals who already claimed their pensions could choose to stop receiving them (Pension Service 2008). However, previously made retirement or financial plans may be costly to change.

In this equation, y_i is our indicator for either pension receipt or working, and $post_i$ is equal to 1 if the observation falls in the second quarter of 2005 or beyond and 0 otherwise. The β_a's represent the impact of the more generous deferral rules on claiming a pension or working at different ages, a. In the equations describing claiming behavior, age (a) runs from 60 to 69 for women and 65 to 69 for men. In the equations describing work behavior, age runs from 55 to 80 for women and 60 to 80 for men. Finally, we include a quadratic time trend. Because the time trend is constrained to be the same both before and after the policy change, these coefficients pick up not only the immediate impact of the policy but also any post-policy deviation from the time trend. To check robustness, we estimate equation (2) with year dummies instead of the time trend, and with controls for quarterly GDP per capita and unemployment rate.

Equation (1) plausibly provides an immediate effect of the deferral policy under the assumption that no other relevant variables changed at the same time. However, it shows only the immediate impact. Equation (2) can provide some evidence of a longer-term average impact of the policy. However, it does not allow us to separate this impact from general time trends in the distribution of retirement and claiming ages. (The quadratic trends only capture trends in the fraction claiming across all ages.) To try to address this concern, we also estimate the following equation:

$$y_i = \alpha + \sum_{y=1999}^{2015} \sum_{a=a_0}^{a_1} \beta_{a,y}\, \mathbf{I}[age_i = a] * \mathbf{I}[year_i = y] + \sum_{a=a_0}^{a_1} \gamma_a\, \mathbf{I}[age_i = a] + \varepsilon_i \quad (3)$$

Variable definitions are the same as in equation (2), but this time we interact the age dummies with year dummies. The coefficients of these interaction terms, the $\beta_{a,y}$'s, show us how the probability of claiming or working at each age changes from year to year. These interactions allow us to pinpoint more precisely when the distributions changed. A post-policy shift in these coefficients would be consistent with the policy having a causal effect on claiming and working.

III. Results

Results from estimating equation (1), the regression discontinuity specification, are shown in Figures 1 (males) and 2 (females). In each figure, the top panel shows the results for claiming and the bottom panel shows the results for working. The figures show that, for both men and women, the immediate effect of the policy on pension claiming goes in the predicted direction but is statistically insignificant. The impact on employment is positive for men and negative for women; but again, both are statistically insignificant. Overall, therefore, we find no evidence of an immediate statistically significant impact of the policy change on claiming or working at the earliest eligibility ages. The results shown include quadratic trends; results with linear trends or local-polynomial estimators are similar and available upon request.

Table 2 shows results for estimating equation (2) with pension receipt as the dependent variable. Each cell reports the β_a for the age given in the row. These coefficients show that after the policy change, the fraction of men receiving pensions at age 65 fell by 6.4 percentage points. There was a smaller decline – 3.0 percentage points – in the fraction receiving pensions at age 66. There was an increase in the fraction receiving pensions at ages 68 and 69. For females, there was a 2.1 percentage point decrease in the fraction receiving pensions at the earliest eligibility age of 60. There was a statistically significant increase at ages 62-64, and statistically significant decreases at ages 65-67 and 69. The coefficients reported in Table 2 for ages 60 and 65 contrast with the regression discontinuity results because the regression discontinuity approach only captures the immediate impact of the policy on women and men who were 60 and 65, respectively, at the time of the policy change. Equation (2), on the other hand, compares the average difference in claiming rates among newly eligible individuals before and after the policy, controlling for a quadratic trend that is assumed not to change at the time the policy is implemented.

Table 3 shows results from estimating equation (2) with working as the dependent variable. It shows that for men, the fraction working increased between 0.7 and 6.5 percentage points at ages 60-69 following the deferral policy change. For women, the fraction working increased between 1.4 and 7.2 percentage points at ages 56-66 following the deferral policy change. For both men and women, there is evidence of a decrease in labor supply at older ages (69+ for women and 71+ for men). The results in Tables 2 and 3 include quadratic trends. Results with year fixed effects, and results with controls for GDP and unemployment, are similar and available upon request.

While this evidence is consistent with the policy having an impact on claiming and labor supply, we cannot rule out the possibility that the results in Tables 2 and 3 simply reflect time trends in the distribution of claiming and retirement ages that are not captured with the quadratic trend. To try to investigate this issue, Table 4 shows results from estimating equation (3) for men, with claiming as the dependent variable. Each cell in the table represents the $\beta_{a,y}$ with the year (y) specified in the rows and age (a) in the columns. The first column of the table indicates that there was a 2.4 percentage point decrease in claims at age 65 in 2005 compared to 2004. However, the other coefficients in the column suggest a falling probability of claiming at age 65 even before the policy. For example, in 1998, the probability of receiving a pension at 65 was 3.2 percentage points greater than it was in 2004. In 2000, the difference had dropped to 1.7 percentage points, and by 2001, it became a statistically insignificant 1.1 percentage points. The first panel of Figure 3 provides a visual depiction of these coefficients and illustrates the time trend. Thus, we cannot be sure that the drop in claiming at age 65 is not part of a general time trend that is not fully captured by the quadratic trend included in the regressions.

At age 66, we observe a significant drop in the probability of pension receipt starting in 2007, and the coefficients prior to 2004 do not indicate a clear time trend (illustrated visually in the second panel of Figure 3). This evidence is consistent with a causal effect for the policy. Most individuals aged 66 in 2005 would have already claimed; thus, the policy is likely to affect claiming at age 66 at least one year following its implementation. For ages 67-69 as well, we observe a decrease in the probability of pension receipt in the years following the policy. The decrease at ages 67 and 68 appear to start in 2010 or 2011, and the decrease at age 69 appears to start in 2013. This pattern is consistent with the policy impacting claims at these ages further down the road. As illustrated in the last three panels of Figure 3, there is no clear pre-policy trend at these ages. The results for ages 68 and 69 indicate that the positive post-policy coefficients shown in Table 2 do not appear to be causal effects of the policy. Instead, they reflect the sharply lower probability of pension receipt in 1998 and 1999 (for age 69) relative to 2004. We do not have an explanation for these negative coefficients but do see that disability claims are proportionally higher in these years for ages 68 and 69.

Table 5 shows the same results for women. Again, each cell shows the $\beta_{a,y}$ specified in the row (y) and column (a).[5] For women, the pattern of coefficients is less consistent with a causal effect of the policy. For example, claims at age 60, the earliest eligibility age, only appear to drop in 2007, two years after the policy. The increase in pension receipt at age 64, shown in Table 2, begins in 2005. It seems unlikely for the policy to immediately impact the behavior at age 64, but have a delayed impact for 60 year old.

We perform similar age by year analyses with employment as the dependent variable for men and women. In both cases, we see increases in labor supply after the policy change, but there is an increasing pattern in coefficients prior to the policy change as well. This could be consistent with either the policy having a causal effect or simply an increasing trend in labor supply over time.[6]

Table 6 shows the fraction of men and women taking their gains from delay as a lump sum at each age. For those who claim within 5 years of eligibility, 3.3 and 2.4 percent of men and women, respectively, take their gains as lump sum. These percentages increase to 9.2 and 16.7 percent for men and women, respectively, for those who delay claiming beyond 5 years.

IV. Discussion and Conclusions

Overall, the results suggest that more generous pension delay policies incentivize delayed claiming among men. The results for women's claiming are less consistent with a causal effect. Overall, after the policy change, women are less likely to claim their state pension at the earliest

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[5] Note that women aged 60 are not in the sample after 2010, and older ages drop out later. This reflects the fact that we exclude women born 1950 and later since they are subject to different retirement ages.

[6] Tables can be provided upon request.

eligible age. However, a year-by-year disaggregation suggests that the shift does not coincide with the start of the policy. Both men and women experience a post-policy increase in employment around their earliest eligibility ages, and a post-policy decrease in employment at older ages. This is consistent with increased incentives to work to finance pension delay in the presence of borrowing constraints, followed by an income effect from more generous pension benefits. While the pattern of results is suggestive of a causal impact of more generous policy upon labor market behavior, a year-by-year disaggregation suggests that we cannot rule out that some of the results are simply capturing long-run time trends.

References

Amilon, Anna and Torben Heien Nielsen. 2010. "How Does the Option to Defer Pension Payments Affect the Labour Supply of Older Workers in Denmark?" in Working and Aging: Emerging Theories and Empirical Perspectives. Luxembourg: Publication Office of the European Union. Retrieved from: https://www.cedefop.europa.eu/files/3053_en.pdf

Blundell, Richard and Carl Emmerson. 2007. "Fiscal Effects of Reforming the UK State Pension System." in *Social Security Programs and Retirement around the World: Fiscal Implications of Reform* (eds. Jonathan Gruber and David A. Wise). Chicago: University of Chicago Press.

Blundell, Richard, Costas Meghir, and Sarah Smith. 2004. "Pension Incentives and the Pattern of Retirement in the United Kingdom." in *Social Security Programs and Retirement around the World: Micro-Estimation* (eds. Jonathan Gruber and David A. Wise). Chicago: University of Chicago Press.

Cribb, Johnathan, Carl Emmerson, and Gemma Tetlow. 2016. "Signals Matter? Large Retirement Responses to Limited Financial Incentives." *Labour Economics* 42: 203-212.

Disney, Richard and Sarah Smith. 2002. "The Labour Supply Effect of the Abolition of the Earnings Rule for Older Workers in the United Kingdom." *Economic Journal* 112: 136-150.

Dagpunar, John. 2015. "Deferring a State Pension - Is It Worthwhile?" *Significance* 12 (2): 30-35.

Farrar, Sue, Jonathan Moizer, and Mark Hyde. 2012. “The Value of Incentives to Defer the UK State Pension.” *Pensions* 17 (1): 46-62.

Kanabar, Ricky and Peter Simmons. 2016. “To Defer or Not Defer? UK State Pension and Work Decisions in a Lifecycle Model.” *Applied Economics* 48 (58): 5699-5716.

Maurer, Raimond, Olivia S. Mitchell, Ralph Rogalla, and Tatjana Schimetschek. 2018. "Will They Take the Money and Work? People's Willingness to Delay Claiming Social Security Benefits for a Lump Sum." *Journal of Risk and Insurance* 85 (4): 877-909.

Meyer, William, and William Reichenstein. 2010. “Social Security: When to Start Benefits and How to Minimize Longevity Risk.” *Journal of Financial Planning* 23 (3): 49–59.

Meyer, William, and William Reichenstein. 2012. “Social Security Claiming Strategies for Singles.” *Retirement Management Journal* 2 (3): 61–66.

Moizer, Jonathan, Sue Farrar, and Mark Hyde. 2018. "UK State Pension Deferral Incentives and Sustainability." *Applied Economics* 50 (21): 2356-2368.

The Pension Service. 2008. *State Pension Deferral – Your Guide*. Department for Work and Pensions.

Sass, Steven A., Wei Sun, and Anthony Webb. 2013. “Social Security Claiming Decision of Married Men and Widow Poverty.” *Economics Letters* 119 (1): 20–23.

Shoven, John, and Sita Slavov. 2014a. “Does It Pay to Delay Social Security?” *Journal of Pension Economics and Finance* 13 (2): 121–144.

Shoven, John, and Sita Slavov. 2014b. “Recent Changes in the Gains from Delaying Social Security.” *Journal of Financial Planning* 27 (3): 32–41.

Thurley, Djuna. 2017. “State Pension Deferral.” Briefing Paper No. CBP-02868. House of Commons Briefing Library.

Table 1: Descriptive Statistics

	Overall	Male		Female	
		Pre	Post	Pre	Post
Employment	0.20	0.20	0.26	0.20	0.16
Employment at Age 60-64	0.40	0.47	0.55	0.25	0.31
Employment at Age 65-69	0.14	0.14	0.22	0.08	0.14
Pension	0.91	0.95	0.92	0.87	0.91
Pension at Age 60-64	0.84	-	-	0.81	0.86
Pension at Age 65-69	0.94	0.95	0.92	0.94	0.95
Age	67.69	68.70	68.32	66.39	68.27
Year	2003	1999	2010	1999	2009
Observations	2,297,250	559,354	438,826	843,845	455,225

Note: The sample includes women born before 1950 and men born before 1953. The sample period is 1993 Q1-2015 Q4. The overall sample includes women with ages 55-80 and men with ages 60-80. For the pension variable, the period begins in 1998 Q2 and ages cover women 60-69 and men with 65-69.

Table 2: Impact of Deferral Policy on Claiming

VARIABLES	Males	Females
Age 60 x post		-0.021***
		(0.003)
Age 61 x post		-0.003
		(0.003)
Age 62 x post		0.006*
		(0.003)
Age 63 x post		0.006*
		(0.003)
Age 64 x post		0.013***
		(0.003)
Age 65 x post	-0.064***	-0.009***
	(0.003)	(0.003)
Age 66 x post	-0.030***	-0.011***
	(0.003)	(0.003)
Age 67 x post	-0.012***	-0.013***
	(0.003)	(0.003)
Age 68 x post	0.007**	-0.005
	(0.003)	(0.003)
Age 69 x post	0.033***	-0.010***
	(0.003)	(0.003)
Observations	209,560	424,838
R-squared	0.040	0.055

Notes: Robust standard errors in parentheses. Regressions also include age dummies and a quadratic year trend.
*** p<0.01, ** p<0.05, * p<0.1

Table 3: Impact of Deferral Policy on Working

VARIABLES	Males	Females
Age 56 x post		0.045***
		(0.007)
Age 57 x post		0.071***
		(0.005)
Age 58 x post		0.068***
		(0.004)
Age 59 x post		0.072***
		(0.004)
Age 60 x post	0.034***	0.054***
	(0.003)	(0.003)
Age 61 x post	0.036***	0.032***
	(0.003)	(0.003)
Age 62 x post	0.047***	0.028***
	(0.003)	(0.003)
Age 63 x post	0.048***	0.030***
	(0.003)	(0.003)
Age 64 x post	0.046***	0.024***
	(0.003)	(0.003)
Age 65 x post	0.065***	0.034***
	(0.003)	(0.003)
Age 66 x post	0.049***	0.014***
	(0.003)	(0.003)
Age 67 x post	0.029***	-0.001
	(0.003)	(0.003)
Age 68 x post	0.025***	0.002
	(0.003)	(0.003)
Age 69 x post	0.007**	-0.011***
	(0.004)	(0.003)
Age 70 x post	0.004	-0.017***
	(0.004)	(0.003)
Age 71 x post	-0.008**	-0.020***
	(0.004)	(0.003)
Age 72 x post	-0.016***	-0.030***
	(0.004)	(0.003)
Age 73 x post	-0.020***	-0.035***
	(0.004)	(0.003)
Age 74 x post	-0.018***	-0.036***
	(0.004)	(0.003)
Age 75 x post	-0.027***	-0.039***
	(0.004)	(0.003)
Observations	998,180	1,299,070
R-squared	0.226	0.248

Notes: Robust standard errors in parentheses.
*** p<0.01, ** p<0.05, * p<0.1

Table 4: Impact of Deferral Policy on Male Claiming Ages

Year/Age	65	66	67	68	69
1998	0.032***	-0.002	0.002	-0.085***	-0.112***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
1999	0.022***	0.000	0.012*	-0.001	-0.089***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
2000	0.017**	0.005	0.012*	0.010	-0.005
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
2001	0.011	0.011	0.020***	0.008	-0.001
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
2002	0.006	-0.005	0.009	0.004	0.006
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
2003	0.002	-0.011	-0.008	-0.002	-0.000
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
2004			Base Year		
2005	-0.024***	-0.005	0.001	0.008	0.002
	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
2006	-0.013*	-0.010	0.003	0.001	0.004
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
2007	-0.034***	-0.029***	0.007	0.004	0.010
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
2008	-0.067***	-0.038***	-0.005	-0.009	-0.004
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
2009	-0.051***	-0.054***	-0.010	-0.013*	-0.004
	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
2010	-0.058***	-0.044***	-0.028***	-0.009	-0.011
	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
2011	-0.052***	-0.037***	-0.023***	-0.033***	-0.010
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
2012	-0.078***	-0.048***	-0.012*	-0.006	-0.012
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
2013	-0.092***	-0.054***	-0.035***	-0.015**	-0.022***
	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
2014	-0.097***	-0.072***	-0.035***	-0.038***	-0.029***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
2015	-0.134***	-0.072***	-0.045***	-0.031***	-0.027***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)

Notes: Robust standard errors in parentheses. Regression based on 209,560 observations and also includes age dummies.
*** p<0.01, ** p<0.05, * p<0.1

Table 5: Impact of Deferral Policy on Female Claiming Ages

Year/Age	60	61	62	63	64	65	66	67	68	69
1998	-0.052***	-0.072***	-0.052***	-0.079***	-0.078***	-0.035***	-0.046***	-0.044***	-0.060***	-0.046***
	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
1999	-0.066***	-0.059***	-0.040***	-0.033***	-0.040***	-0.026***	-0.043***	-0.028***	-0.033***	-0.020**
	(0.007)	(0.007)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
2000	-0.017**	-0.037***	-0.039***	-0.020**	-0.018**	-0.004	-0.025***	-0.031***	-0.011	-0.018**
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)
2001	0.007	-0.022***	-0.026***	-0.028***	-0.002	0.006	-0.021**	-0.017**	-0.021**	-0.017**
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)
2002	-0.017**	-0.017**	-0.018**	-0.010	-0.007	0.002	-0.020**	-0.012	-0.016*	-0.017**
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
2003	-0.018**	0.004	-0.006	-0.020**	-0.014	-0.004	-0.003	-0.017*	-0.003	0.001
	(0.008)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.010)
2004						Base Year				
2005	-0.011	0.015**	0.003	0.018**	0.018**	0.002	-0.006	-0.002	0.001	0.003
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)
2006	0.009	0.024***	0.034***	0.014*	0.025***	0.018**	-0.004	-0.009	0.010	-0.001
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)
2007	-0.029***	0.001	0.024***	0.030***	0.019**	0.014*	-0.003	-0.007	0.007	0.010
	(0.007)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)	(0.009)
2008	-0.012	-0.012*	-0.011	0.022***	0.028***	0.011	-0.005	0.001	0.005	0.007
	(0.007)	(0.007)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.008)	(0.009)
2009	-0.023***	0.007	0.007	0.011	0.030***	0.022***	-0.003	-0.018**	-0.008	-0.006
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
2010		-0.025***	0.014*	-0.002	0.014*	0.014*	0.002	-0.009	-0.001	0.002
		(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
2011			0.011	0.004	0.022***	0.017**	-0.008	0.003	-0.004	0.003
			(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
2012				0.009	0.026***	0.012	-0.006	-0.003	0.011	0.005
				(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
2013					0.022***	-0.002	-0.011	-0.005	0.004	0.004
					(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
2014						-0.002	-0.009	-0.008	-0.001	-0.004
						(0.008)	(0.008)	(0.008)	(0.009)	(0.009)
2015							-0.026***	-0.018**	-0.010	-0.012
							(0.008)	(0.008)	(0.008)	(0.009)

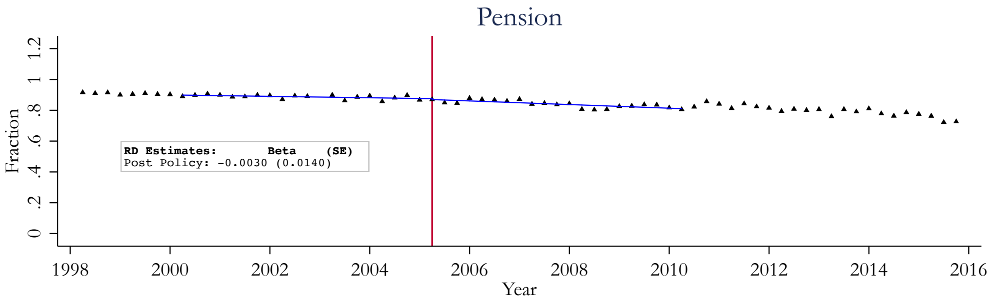
Notes: Robust standard errors in parentheses. Regression based on 424,838 observations and also includes age dummies. Women aged 60-65 not present in some years due to sample restrictions on birth years.

*** p<0.01, ** p<0.05, * p<0.1

Table 6. Share of First Time Pensioners Receiving Lumpsum After April 2005 by Gender and Age

	Male			Female		
Claiming Age	Lump Sum	No Lump Sum	Lump Sum Share	Lump Sum	No Lump Sum	Lump Sum Share
60				4	243	1.62
61				5	386	1.28
62				5	124	3.88
63				4	54	6.90
64				3	44	6.38
65	13	417	3.02	6	27	18.18
66	15	432	3.36	0	34	0.00
67	3	59	4.84	2	18	10.00
68	0	24	0.00	3	17	15.00
69	1	11	8.33	1	21	4.55
70	0	13	0.00	4	18	18.18
71	0	22	0.00	3	14	17.65
72	0	15	0.00	1	12	7.69
73	2	16	11.11	1	11	8.33
74	2	12	14.29	3	9	25.00
75	1	10	9.09	3	14	17.65
76	0	11	0.00	2	12	14.29
77	0	9	0.00	1	6	14.29
78	1	11	8.33	3	7	30.00
79	0	8	0.00	0	4	0.00
80	1	6	14.29	1	2	33.33
81	1	7	12.50	1	5	16.67
82	0	4	0.00	0	0	-
83	2	5	28.57	1	5	16.67
84	1	2	33.33	6	5	54.55
85	3	4	42.86	3	5	37.50
86	0	3	0.00	0	6	0.00
87	0	3	0.00	1	1	50.00
88	1	1	50.00	2	0	100.00
89	1	2	33.33	0	0	-
90	1	3	25.00	4	7	36.36

Figure 1: Male Regression Discontinuity Estimates



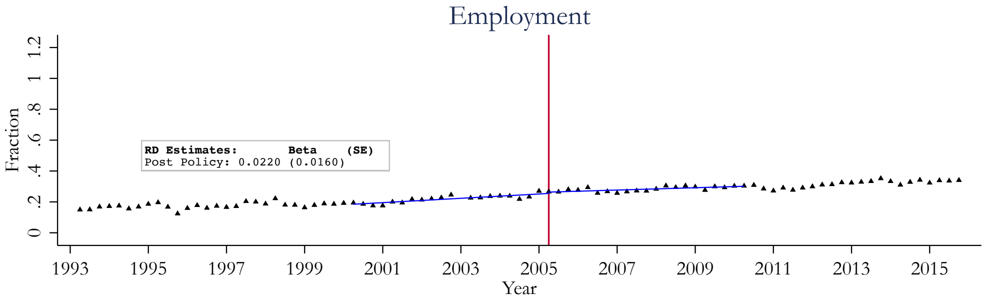


Figure 2: Female Regression Discontinuity Estimates

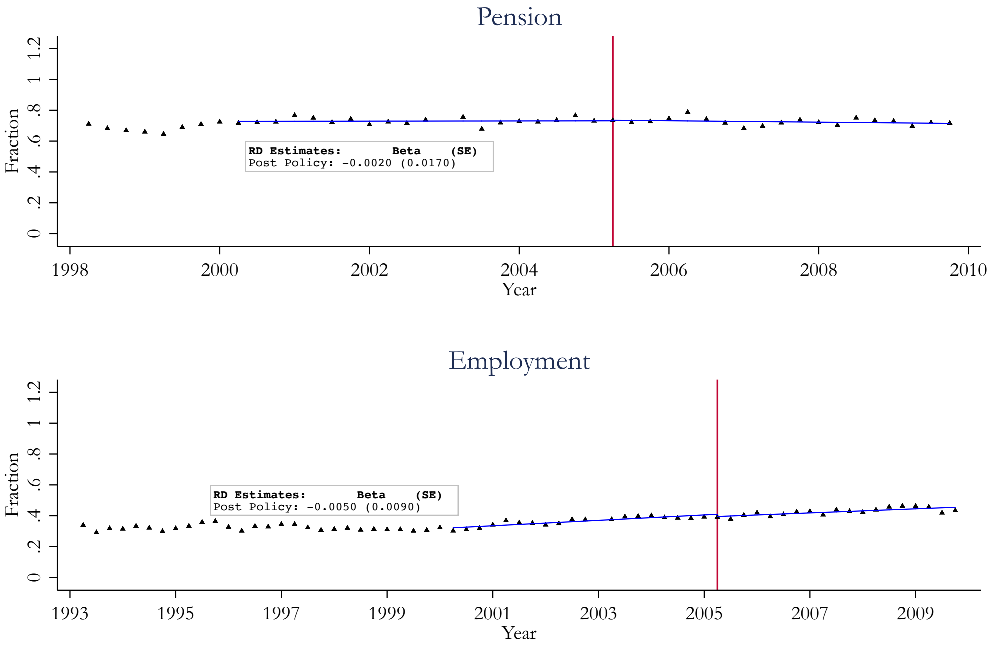
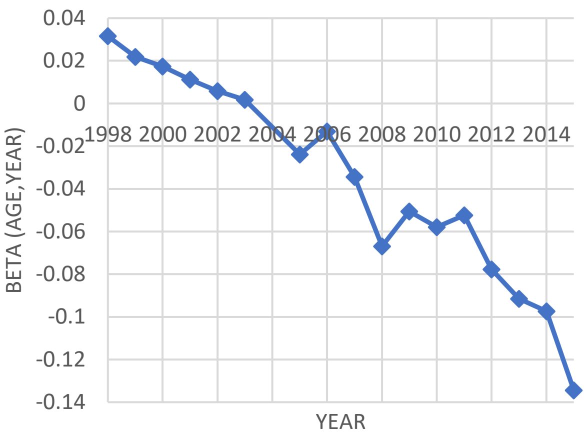
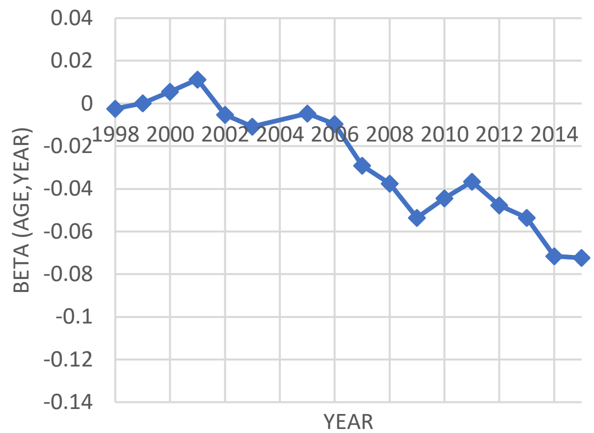


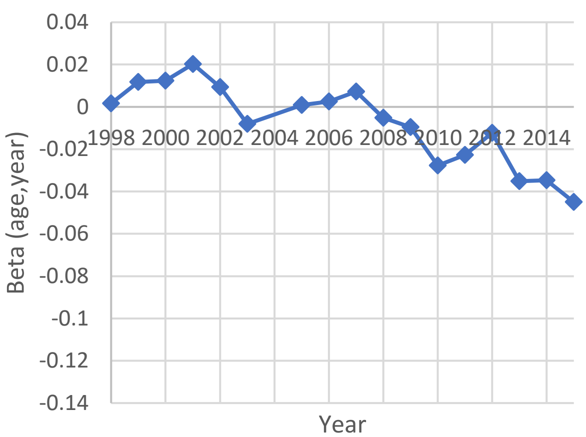
Figure 3. Change in Male Pension Claiming Relative to 2004



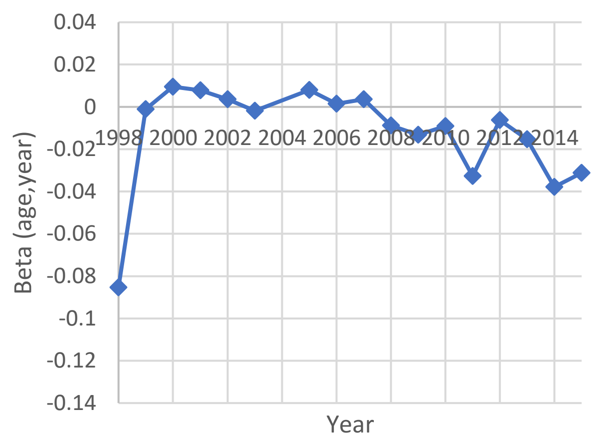
(A) Age 65



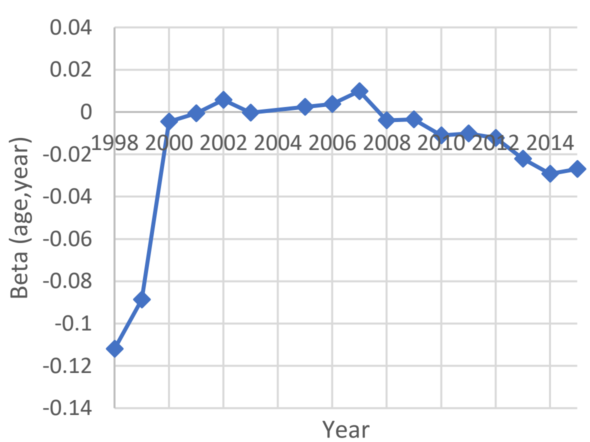
(B) Age 66



(C) Age 67



(D) Age 68



(E) Age 69