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*Abstract*: Exploiting the 2011 abolishment of the mandatory-retirement provision that had allowed employers to force retirement of their 65-or-older employees in the UK, this paper estimates the effect of outlawing mandatory retirement on public-pension benefit claim behavior of the elderly through changing their employment and retirement rates. Abolishing mandatory retirement raises public-pension claim rate of non-household-head benefit-eligible individuals by making them retire more, while it lowers public-pension claim rate of household-head benefit-eligible individuals by lowering their retirement rate. On the other hand, the abolishment raises full-time employment rate of the elderly, regardless of household-head status or gender.

*Keywords:* mandatory retirement, public-pension claim, household head
*JEL:* H55, J26, J71, J21, K31

## I. Introduction

Due to rapidly aging population, an increasing number of countries are outlawing mandatory-retirement provision that allows employers to force their old employees to retire without a performance-based reason.[1] Abolishing mandatory retirement may help prevent population aging from shrinking the share of working population by leading increasing elderly individuals to work, as the abolishment eliminates the legally-allowed discrimination against their labor supply. Furthermore, policymakers expect that outlawing mandatory-retirement provision would reduce

*Correspondence: Peking University HSBC Business School, University Town, Nanshan District, Shenzhen, China, 518055. Tel: +86 755 2603 2293 E-mail: islee@phbs.pku.edu.cn

[1] Supported by stricter and more consistent applications of the EU equal treatment rules (Directive 2000/78/EC) and domestic laws against age discrimination, many European countries are outlawing their mandatory-retirement provisions in some or all sectors. For example, in 2015, Danish Parliament decided to abolish its 70-year compulsory retirement rule in the name of the Danish Anti-Discrimination Act. In Spain, labor reforms of February 2012 outlawed compulsory retirement. There is a similar legal change in the Czech Republic. In addition, the Court of Justice of the European Union (CJEU) applied EU equal treatment rules (Directive 2000/78/EC) strictly on increasing litigations against mandatory-retirement provisions such as Lufthansa pilot retirement rules, and Hungarian forced early retirement of judges, all of which were ruled by the CJEU to be violations of the EU law. This is contrasting to the previous lenient stance that the CJEU took by stating in 2007 that the mandatory retirement of workers at age 65 could be justified, although discrimination based on age was illegal. Furthermore, China and Korea also conducted similar legislative changes. Minister of Human Resources and Social Security of China plans to increase the mandatory retirement age gradually. In 2016, Korea implemented the law for protecting old workers from age-based terminations of employment until age 60.

public-pension benefit claims of the elderly by lowering their retirement rate. Labor supply decision of the elderly is clearly different from their pension benefit claim decision, although the two decisions are closely related to each other. Understanding public-pension benefit claim behavior is important not only for public-pension sustainability but also for finances of the elderly. Nevertheless, public-pension benefit claim behavior is not extensively studied and there is no rigorous study conducted regarding the effect of a labor-policy change on public-pension benefit claim behavior. This paper investigates how outlawing mandatory retirement affects public-pension benefit claim behavior of elderly individuals.

Because abolishing mandatory retirement is a legislative change to restrict *employers* from terminating employment of their old employees or from refusing to offer elderly individuals a job with age-based reason, it is an exogenous positive shock on demand for elderly individuals' labor at the extensive margin. Through elderly individuals' labor-supply response to the positive demand shock, abolishing mandatory retirement can entail a change in their public-pension benefit take-up behavior. Although numerous empirical studies (e.g., Coile and Gruber, 2007; Behaghel and Blau, 2012) found that retirement rate hikes at public-pension entitlement ages, they showed that retirement rate at the entitlement ages is far lower than 100%. As retirement decision is clearly separate from public-pension claim decision, once reaching public-pension entitlement ages, an individual can claim and receive public-pension benefits even when he (she) is not retired but working. Therefore, understanding the effect of abolishing mandatory retirement on employment or retirement of the elderly is not understanding the effect of the abolishment on their public-pension benefit claim. In fact, various empirical studies have shown *contradictory* findings on whether outlawing mandatory retirement actually makes elderly individuals be retired less or employed more. With pre-1967 data of white men, Adams (2004)

found positive effect on employment rate of abolishing age-discriminatory retirement mandates, while he failed to find consistent evidence regarding its effect on retirement rate.[2] Based on an imputation analysis of male workers who aged 62 – 64 in 1973, Burkhauser and Quinn (1983) reported positive effect of abolishing age-discriminatory retirement mandates on the labor force participation prospect. In contrast, with data of white men around 1978, Wachter (2002) found that outlawing age-discriminatory retirement mandates does not affect employment rate while it reduces retirement rate. Shannon and Grierson (2004) showed that outlawing mandatory-retirement provision in the early 1980s does not make a consistent difference in the employment rate of the elderly between age 65 and 69. Confined to one specific occupation (university professor), Ashenfelter and Card (2002) and Warman and Worswick (2010) found negative effect of outlawing mandatory retirement on the retirement rate of elderly professors, while Warman and Worswick (2010) also found that this negative effect disappears for those who have obtained their highest degree later than age 33.

Moreover, notice that the existing empirical studies are based on the data of elderly men only between the 1960s and the early 1980s. Because the elderly labor force has been fundamentally changed since then, we need empirical investigations with post-2000 data of *both* men and women. Firstly, as the elderly have become much healthier and better educated than before the 1980s, their labor productivity has increased substantially (OECD, 2006). Secondly, in the OECD countries, including the United Kingdom, the labor force participations of the elderly had declined from the 1970s and reversed to increase since the early 1990s particularly for the elderly women (Blöndal and Scarpetta, 1999). [3]

[2] According to Table 6 and 7 of Adams (2004), some of the estimates for the effect on retirement rate of pre-1967 legal protections for the elderly against age discrimination are not different from zero while others are negative.

[3] In the OECD countries, the average labor force participation rate of 65-or-older men increased from 15.1% in 1990 to 19.9% in 2016, while that of 65-or-older women rose from 6.9% in 1990 to 10.2% in 2016. Moreover, the increasing trend of elderly

So far, little research is conducted on how a labor policy change affects public-pension benefit claim behavior of elderly individuals, in contrast to a great number of studies done regarding how a public-pension policy affects their labor supply (e.g., Coile and Gruber, 2007; Behaghel and Blau, 2012; Atalay and Barrett, 2015). However, because individuals smooth their consumption over their lifetime, public-pension benefit claim and retirement are related to each other. Hence, a change in labor policy for elderly individuals can change public-pension benefit claim behavior of elderly individuals. With data of men of 1931 – 1943 birth cohorts, Neumark and Song (2013) investigated whether the effect of Social Security reform in 1983 was strengthened by enforcement aspects of existing anti-age-discrimination laws (firm size that the anti-age-discrimination laws can be applied to, compensatory and/or punitive damages on the firms caught, and length of the filing period). While Neumark and Song (2013) did not examine the effect of outlawing mandatory retirement, they found that stronger enforcements of anti-age-discrimination laws make statistically insignificant and inconsistent differences in the effect of the 1983 Social Security reform on public-pension benefit claim behavior of the elderly men.

Exploiting the abolishment of mandatory retirement from age 65 (Default Retirement Age) in the United Kingdom, which took effect in 2011, this paper estimates the effects of outlawing mandatory retirement on public-pension benefit claim of elderly men and women. This paper finds that abolishing the mandatory retirement lowers public-pension benefit claim rate of household-head individuals of ages 65 – 70 who are benefit-eligible by reducing their retirement rate. In contrast, this paper also finds that the mandatory-retirement abolishment *raises* public-pension benefit claim rate of benefit-eligible elderly individuals who are not head of a household by increasing their retirement rate. Because most of the benefit-eligible elderly men are

women's labor force participation is likely to continue because the average labor force participation rate of women between age 15 and 64 increased from 58.1% in 1990 to 63.6% in 2016, while that of counterpart men dwindled from 82.1% in 1990 to 80.0% in 2016.

household head, as opposed to the female counterpart, the effects of abolishing the mandatory retirement on public-pension benefit claim and retirement are found to be negative for the elderly men and positive for the elderly women. On the other hand, outlawing the mandatory retirement raises full-time employment rate of the benefit-eligible elderly individuals. This effect on full-time employment rate remains positive for both household-head and non-household-head individuals as well as for both elderly men and women.

This paper is organized as follows. Section II provides information on the mandatory-retirement abolishment and public-pension policies in the United Kingdom. Section III elaborates on the strategy for identifying the effect of outlawing the mandatory retirement on public-pension benefit claim, retirement and employment of the elderly with the dataset introduced in Section IV. Section V presents the estimation results. Section VI concludes the paper.

## II. Institutional Background

In the United Kingdom, it was legally justifiable that employers force their employees to retire or refuse to offer employment to old job applicants based on age if age of employees or job applicants is equal to or higher than 65, Default Retirement Age (DRA). Facing criticisms against such age-discriminatory application of the anti-discrimination laws as well as rapidly aging population, the mandatory retirement was abolished uniformly for all sectors in 2011. Abolishing the mandatory retirement started with some exceptions from April 6th 2011, and it was fully implemented with no exception only after October 1st 2011 when the transitional provisions attached to the abolishment were expired. In detail, during the transitional period between April 6th 2011 and September 30th 2011, age-based termination of employment from DRA was still legally justifiable *if* an employer had reached his (her) own public-pension

entitlement age by September 30th 2011 or *if* an employer had given retirement notice to his (her) employees at DRA or above according to qualified notification procedures on or before April 5th 2011. Because the dataset used for this paper does not have the information on the birth date and gender of an individual's employer or on the date and detailed procedure of retirement notice, if any, that an individual received, the observations of the dataset collected between April 6th 2011 and September 30th 2011 are discarded because they are not proper for clearly identifying the effect of abolishing the mandatory retirement. Above all, the group to whom abolishing the mandatory retirement introduced legal protection from age discrimination is composed of elderly men and women whose age was 65 (DRA) or above when the abolishment was fully implemented. However, because the mandatory retirement had never been applicable to the self-employed before and after the abolishment, observations on them, which are negligibly small, are not used.

Because the mandatory-retirement provision addresses individuals who are of age 65 (DRA) or above, job-seeking and continuing paid-work of these 65-or-older individuals are affected by outlawing the mandatory-retirement provision of the UK in 2011. Although the mandatory-retirement provision of the UK is not requiring all employers to terminate all of their 65-or-older employees or to not hire them, there are various evidences showing that the number of elderly individuals who are affected by the abolishment of the mandatory retirement would not be small. According to Disney et al. (1998), 48.9% of workers report that their company imposes upper age limit of working to enforce retirement of old employees. Moreover, with the data of EU survey of 2011, Barret and Sargeant (2015) found that significantly higher portion of old workers of the UK want to keep working after 65 (DRA) than the average of 27 EU countries. According

to Barret and Sargeant (2015), 64% of non-manager white-collar workers of the UK want to keep working after DRA, while only 29% of non-manager white-collar workers of 27 EU countries do.

65-or-older individuals, who are affected by the mandatory-retirement abolishment, are also eligible to claim public-pension benefit. Like public pension program of United States, the United Kingdom also provides nationwide public pension (State Pension) to the UK citizens. From the public-pension entitlement age, individuals become eligible to claim and receive public-pension benefit, regardless of whether they actually stop working or not. The public-pension entitlement age stayed 65 for men before and after mandatory-retirement abolishment, while the public-pension entitlement age for women gradually increased from 60 to 65 between 2010 and 2018 according to Pensions Act 2007. This concurrent change in the public-pension entitlement age may raise a bias concern in estimating the effect of the mandatory-retirement abolishment. To address this bias concern, the time interval after abolishing the mandatory retirement is chosen short enough to let none of 65-or-older individuals in the estimation be affected by the increases in the public-pension entitlement age. On the other hand, unlike the US Social Security, the UK public pension program has not had *early* public-pension entitlement age from which discounted periodic public-pension benefits are claimable before reaching his (her) public-pension entitlement age. Most of all, even if an individual is benefit-eligible after reaching his (her) public-pension entitlement age, public-pension benefit is not automatically paid out to him (her) but requires him (her) to formally claim public-pension benefit. The amount of public-pension benefit per year is calculated according to a pre-set formula and is increasing with the total amount of contributions made before. If benefit-eligible individuals delay claiming and receiving their public-pension benefit after the entitlement age, their public-pension benefit per year increases more than actuarially fair rate. When benefit-eligible individuals work after

the public-pension entitlement age, they no longer need to pay public-pension contributions. As such, the UK public pension program gives incentives for benefit-eligible individuals to delay claiming public-pension benefit and to work longer. The main basic unit of the UK public pension program is individual, not household. Except for the public-pension entitlement ages, the public-pension benefit formula and other regulations of the UK public pension program are not gender-discriminatory but treat both genders equally.

As the basic unit that both the mandatory-retirement provision and the UK public pension program are applied is individual, the mandatory-retirement abolishment is expected to reduce public-pension benefit claim of all 65-or-older individual men and women by making them work more and retire less. When the abolishment leads 65-or-older individuals to supply more labor and earn more labor incomes, 65-or-older individuals are more able to delay claiming and receiving public-pension benefits, which will reward them with greater amount of periodic public-pension benefits. This is what policymakers expect, which assumes that all of 65-or-older individuals take into account of themselves only for making decisions of their labor supply and public-pension benefit claim for responding to the mandatory-retirement abolishment, which is also the standard approach of theoretical economic analyses. On the other hand, if 65-or-older individuals take into account of other household members as well, then how their public-pension benefit claim and labor supply respond to the abolishment of the mandatory retirement is not a priori clear.

## III. Identification Strategy

Because 65-or-younger men are not legally allowed to claim their public-pension benefits, those who are younger than DRA (65) cannot be suitable control group for identifying the effect of outlawing the mandatory retirement on public-pension benefit claim. This legal restriction on

public-pension benefit claim below DRA leaves *only* the observations of 65-or-older individuals suitable for identifying the effect. Thus, we cannot use difference-in-difference regressions which are adopted by the related studies (e.g., Ashenfelter and Card, 2002; Wachter, 2002; Adams, 2004; Shannon and Grierson, 2004; Warman and Worswick, 2010). Only with the observations of 65-or-older individuals, we adopt matching estimator for identifying the effects of abolishing the mandatory retirement.

To elaborate on the identification strategy, let $Y_i(1)$ denote outcome (public-pension benefit claim, retirement, or employment) of an individual *i* if he (she) receives the treatment of abolishing the mandatory retirement, and $Y_i(0)$ denote his (her) outcome if he (she) does not. For any given individual *i*, while both of the two outcomes, $Y_i(1)$ and $Y_i(0)$, have the potential to be realized, only one of the two is actually realized and observed with the other unrealized one remaining as potential outcome. Under this framework, the effect of abolishing the mandatory retirement is stated as

$$\alpha = E[Y_i(1) - Y_i(0) \mid W = 1], \tag{1}$$

where $W$ takes the value of one (zero) if an individual is treated (not treated) by the abolishment. Because we cannot observe $Y_i(0)$ for treated 65-or-older individuals ($W = 1$), we need to estimate $E[Y_i(0) \mid W = 1]$ with the observations of not-treated 65-or-older individuals ($W = 0$). For any given treated 65-or-older individual *i*, one or multiple not-treated 65-or-older individuals are matched to him (her), based on the relevant characteristics, $\mathbf{X}$, for estimating the counterfactual outcome of what the treated 65-or-older individual *i* would do if he (she) were not treated. Specifically, propensity-score matching (Rosenbaum and Rubin, 1983) is utilized. The identification conditions for unbiased propensity-score matching estimation are

(2) $$E[Y_i(0)\,|\,p(\mathbf{X}), W=1] = E[Y_i(0)\,|\,p(\mathbf{X}), W=0] = E[Y_i(0)\,|\,p(\mathbf{X})]$$

(3) $$0 < \Pr(W=1\,|\,\mathbf{X}) < 1$$

where $p(\mathbf{X}) = \Pr(W=1\,|\,\mathbf{X})$ is the propensity score.

Firstly, the conditional mean-independence condition of (2) can be violated when individuals can choose whether to be treated. As an individual cannot choose whether his (her) age is equal to or higher than 65 (DRA) when the mandatory retirement is abolished, the treatment assignment is exogenously given. Secondly, the overlap condition of (3) is to ensure that, for each treated 65-or-older individual, there exists at least one not-treated 65-or-older individual who is identical to him (her) in terms of the covariates $\mathbf{X}$. The condition of (3) gives rise to the need for a sufficiently large number of observations.

Because the mandatory retirement was abolished simultaneously across all sectors and all regions, it is inevitable that observations of not-treated 65-or-older individuals for matching are obtained from the same population before abolishing the mandatory retirement. The time window for the estimation is chosen to maintain both treated individuals and not-treated individuals similar except for the treatment. Nonetheless, still one may as well concern about the potential existence of time trend that might generate a difference between not-treated individuals and treated individuals.[4] If the time-trend effect exists, it can violate the condition of (2). To address this concern, we estimate the time-trend effect, notated as $\delta$, and purge it out from the difference between outcome of treated individuals and their counterfactual outcome estimated from matched not-treated individuals that are observed before outlawing the mandatory retirement. By allowing the time-trend effect $\delta$ to exist, the estimation of $E[Y_i(0)\,|\,W=1] =$

[4] Because time indicator variable overlaps with indicator of treatment of abolishing the mandatory retirement, it is inappropriate to include time indicator variable in the set of the covariates **X** which are used for matching treated and not-treated individuals. However, there are no considerable macroeconomic changes or shifts that may affect retirements.

$E_{\mathbf{X}|W=1}\{E[Y_i(0)|W=0]\}$ is modified to

(4) $$E[Y_i(0)|W=1] = E_{\mathbf{X}|W=1}\{E[Y_i(0)+\delta|W=0]\} = E_{\mathbf{X}|W=1}\{E[Y_i(0)|W=0]\}+\delta .$$

As a consequence, the estimation of the key parameter (1) is re-stated as

(5) $$\alpha = E_{\mathbf{X}|W=1}\{E[Y_i(1)|W=1]-E[Y_i(0)|W=0]\}-\delta .$$

We estimate the time-trend effect $\delta$ by difference between observed outcome of not-treated 65-or-older individuals (observed before the mandatory-retirement abolishment) and counterfactual outcome of what matched treated 65-or-older individuals (observed after the abolishment) would do *without* the abolishment treatment.

(6) $$\delta = E_{\mathbf{X}|W=0}\{E[Y_i(0)|W=1]-E[Y_i(0)|W=0]\} .$$

Notice that $E_{\mathbf{X}|W=1}\{E[Y_i(0)|W=1]\}$ is not necessarily equal to $E_{\mathbf{X}|W=0}\{E[Y_i(0)|W=1]\}$. While we allow the possibility that $\delta$ can take a non-zero value, the condition of (2) restricts $\delta$ to be zero. For unbiased identification of $\alpha$, after purging out the time-trend effect $\delta$ the condition of (2) is replaced by

(7) $$\mathbf{X} \perp W \,|\, p(\mathbf{X}) .$$

The balancing condition of (7) generates the identification environment of random assignment of the treatment, conditioned on the propensity score (Rosenbaum and Rubin, 1983). In fact, the balancing condition of (7) is stronger than the conditional mean independence condition of (2).

For unbiasedly estimating the effect of outlawing the mandatory retirement, it is useful to test whether the two key identification conditions of (3) and (7) are met or not with the covariates $\mathbf{X}$ chosen for the matching estimation. First, we can directly check the overlap condition of (3) by drawing histograms of the propensity score $p(\mathbf{X})$ for treated and not-treated individuals. Second, utilizing the formula of Rosenbaum and Rubin (1985) that measures potential bias from failing to

meet the balancing condition of (7), we select the set of the covariates **X** only if the entailed standardized potential bias is less than or equal to 5% after matching.[5]

The estimation procedure is summarized as follows. First, based on Probit model, we obtain the propensity score $p(\mathbf{X})$ for treated and not-treated 65-or-older individuals. Second, using the obtained propensity score, to each treated 65-or-older individual, the most similar not-treated 65-or-older individuals are matched. Third, over the population of treated 65-or-older individuals we average difference between observed outcomes (public-pension benefit claim, retirement, or employment) of treated 65-or-older individuals and their counterfactual outcomes (estimated from matched not-treated 65-or-older individuals). Fourth, for estimating the time-trend effect $\delta$, over the population of not-treated 65-or-older individuals, we average difference between observed outcomes of not-treated 65-or-older individuals and counterfactual outcomes of what matched treated 65-or-older individuals would do with no treatment. Lastly, for an estimate of $\alpha$, the difference obtained from the fourth step is subtracted from the difference obtained from the third step. For robust inference, we utilize heteroscedasticity-robust standard errors, according to Abadie et al. (2004).

In the second step of the estimation procedure, we allow the possibility that there exist *multiple* not-treated individuals who are the closest (and thus equally similar) to a given treated individual in terms of the propensity score. For averaging the outcomes of the matched not-treated individuals to get one counterfactual outcome of the given treated individual, we adopt local linear matching of Heckman, Ichimura, and Todd (1997) because Fan (1992 and 1993) showed that local linear matching is more robust to different data densities than other alternatives.

Although the mandatory retirement was abolished equally for men and women, the effects of

[5] The standardized potential bias will be 0% only in a rather unrealistic case that for every treated individual, there exists a perfectly identical replica of him (her) who is not treated.

abolishing the mandatory retirement on public-pension benefit claim, employment and retirement can be different for both genders. Firstly, public-pension entitlement ages have been different for men and women. Secondly, as various empirical studies (e.g., Azmat, Güell and Manning, 2006; Chevalier, 2007) have shown, there have long existed substantial gender gaps in wages and employment. These two kinds of gender gaps can beget a gender-difference in the effects of the mandatory-retirement abolishment. Furthermore, such gender-difference in the effects is unlikely to be controlled simply by including an indicator of gender in the covariates of $\mathbf{X}$ in the propensity-score matching estimation. As a result, putting the observations of men and women together in the same matching estimation can beget an uncontrollable bias. For preventing potential bias problems, we run the same matching estimation with the same covariates $\mathbf{X}$ separately for men and women.

## IV. Data Overview

For the estimation, we utilize the Family Resources Survey (FRS) that is a nationally representative micro-level cross-sectional data annually collected by the Office for National Statistics and National Centre for Social Research.[6] The FRS provides the information of individual respondents such as their labor status, public-pension benefit take-up, marital status, education, health condition, and the like. We use data on individuals of ages 65 – 70 from the three (2010, 2011 and 2012) waves of the FRS. In the FRS dataset, the age of an individual refers to his (her) age at the time of survey. As noted by the previous empirical studies on labor supply of the elderly (e.g., Ruhm, 1990; Wachter, 2002; Adams, 2004; Coile and Gruber, 2007; Atalay

[6] The FRS was downloaded from http://www.data-archive.ac.uk/ after obtaining the permission from the distributor of the UK Data Archive. As an alternative to the FRS that is a cross-sectional data, ELSA (English Longitudinal Study of Ageing) that is a biennial panel data of the elderly of the UK is also considered. However, the size of available observations of 65-or-older individuals from the ELSA is less than 50% of the observations available from the FRS, although the necessary key information such as labor status, public pension benefit claim, education, and the like are all available for both data sets. The small number of observations from ELSA causes failures in meeting the overlap condition of (3) and the balancing condition of (7).

and Barrett, 2015; Cribb, Emmerson and Tetlow, 2016), in terms of employability, those who are of ages 71 and above are fairly different from those who are of ages 65 – 70. Hence, for a clear identification, we do not use the observations of 71-or-older individuals in the estimation.

The length of the intervals before and after abolishing the mandatory retirement, respectively, is chosen as 1.25 years so that both not-treated and treated 65-or-older individuals from the same population are under very similar macroeconomic environment except for the abolishment. With this length of the interval, the growth rate of GDP and unemployment rate of the UK remain quite stable[7]. Moreover, we choose the length of the interval to make both not-treated and treated 65-or-older individuals alike unaffected by Pensions Act 2007. Expanding the interval will make 65-or-older individuals before and after abolishing the mandatory retirement be affected also by Pensions Act 2007 and face quite different macroeconomic environments. The pre-abolishment period is from January 1st 2010 to April 5th 2011 and the post-abolishment period is from October 1st 2011 to December 31th 2012. For the estimation, we use the data of individuals of ages 65 – 70 observed in the pre-abolishment period and the post-abolishment period, not in the transitional period. For identifying the effects of abolishing the mandatory retirement, the control group is composed of individuals who are of ages 65 – 70 and observed during the pre-abolishment period, while the treatment group is composed of individuals who are of ages 65 – 70 and observed during the post-abolishment period. Because the FRS data is a pooled cross-

[7] The GDP growth rate and aggregate unemployment rate remain very stable throughout the sample periods (between 2010 and 2012). According to World Bank database, the GDP growth rate of the UK was 1.9% for 2010; 1.5% for 2011; 1.3% for 2012. According to World Bank database, the unemployment rate of the UK was 7.8% for 2010; 8.0% for 2011; and 7.9% for 2012. According to the OECD database, in the UK, the employment rate of men of ages 15 - 64 was 75.2% for 2010; 75.4% for 2011; and 75.6% for 2012. And, the employment rate of women of ages 15 - 64 was 65.3% for 2010; 65.2% for 2011; and 65.4% for 2012. According to the OECD database, the labor force participation rate of men of ages 15 - 64 was 82.5% for 2010; 82.5% for 2011; and 82.7% for 2012 in the UK. And, the labor force participation rate of women of ages 15-64 was 70.2% for 2010; 70.4% for 2011; and 70.7% for 2012. Moreover, according to the UK government statistics, there is no discernible change in 'healthy life expectancy' of the elderly of the UK between 2010 and 2012. As we choose the time interval to be short enough to avoid substantial changes in economic conditions, the estimated effect of the mandatory-retirement abolishment is not long run effect. With adjustments of young workers' retirement plans and savings, its long-run effect can be different.

sectional data, instead of a panel data, a 65-or-older individual can belong to one of these two groups.

Above all, for estimating the effect of outlawing the mandatory retirement on public-pension benefit claim, outcome variable of $Y_i$ is defined as a binary indicator that takes the value of one if individual *i* takes up the UK public pension (State Pension) benefit[8] and the value of zero otherwise. To better understand the channel of how the mandatory-retirement abolishment affects public-pension benefit claim behavior, we also estimate the effects on labor supply of benefit-eligible individuals. First, for identifying the effect on their retirement of abolishing the mandatory retirement, we define outcome variable of $Y_i$ as a binary indicator that takes the value of one if individual *i* is retired and the value of zero otherwise. Second, for estimating the effect on of benefit-eligible individuals' employment of abolishing the mandatory retirement, outcome variable of $Y_i$ is defined as a binary indicator that takes the value of one if individual *i* is employed and the value of zero otherwise. Since elderly individuals often take part-time jobs as the interim between full-time employment and retirement (Ruhm, 1990), we additionally conduct the same estimation with binary indicators of being full-time employed and being part-time employed, respectively, as the outcome variable $Y_i$.

**Table 1** summarizes the public-pension benefit claim, retirement and employment rates of treated and not-treated benefit-eligible individuals of each gender before matching. Comparing treated and not-treated individuals of ages 65 – 70 shows that public-pension benefit claim rate of benefit-eligible elderly men decreases after abolishing the mandatory retirement, while public-pension benefit claim rate of benefit-eligible elderly women increases. In line with this, after the

[8] Notice that State Pension benefit is for those who are entitled by their own past contributions to National Insurance of the UK. Thus, this does not include benefit from transfer programs such as Guarantee Credit and Savings Credit.

abolishment, retirement rate of benefit-eligible elderly men decreases, while retirement rate of benefit-eligible elderly women increases. On the other hand, full-time employment and overall employment rates of both benefit-eligible elderly men and women increase together after outlawing the mandatory retirement.

**Table 1**] **Summary Statistics of Public-Pension Claim, Retirement, and Employment by the Abolishment of the Mandatory Retirement**

| | Elderly Men | | | Elderly Women | | |
|---|---|---|---|---|---|---|
| | Treated by Abolishment | Not-treated by Abolishment | Δ | Treated by Abolishment | Not-treated by Abolishment | Δ |
| | Mean (SD) | Mean (SD) | | Mean (SD) | Mean (SD) | |
| Public-Pension | 0.963 | 0.968 | −0.005 | 0.978 | 0.976 | 0.002 |
| Benefit Claim | (0.189) | (0.175) | | (0.146) | (0.153) | |
| Retirement | 0.801 | 0.812 | −0.011 | 0.821 | 0.810 | 0.010 |
| | (0.399) | (0.391) | | (0.384) | (0.392) | |
| Part-time | 0.061 | 0.062 | −0.001 | 0.073 | 0.076 | −0.003 |
| Employment | (0.240) | (0.242) | | (0.261) | (0.264) | |
| Full-time | 0.067 | 0.056 | 0.011 | 0.025 | 0.017 | 0.008 |
| Employment | (0.251) | (0.230) | | (0.156) | (0.130) | |
| Overall | 0.129 | 0.119 | 0.010 | 0.098 | 0.093 | 0.005 |
| Employment | (0.335) | (0.323) | | (0.298) | (0.290) | |
| No. of Obs. | 2193 | 1713 | | 2518 | 1973 | |

**Table 2** displays the summary of the covariates $\mathbf{X}$ of treated and not-treated individuals, prior to matching. As elaborated in Section III, the covariate variables $\mathbf{X}$ are chosen to meet the identification conditions of (3) and (7) and used for obtaining the propensity score in the first step of the estimation procedure. The binary indicator variable of 'Head of a Household' takes the value of one if an individual is a household head so that he (she) is responsible for the accommodation of his (her) household by paying the rent or owning the house and the value of zero otherwise. While a household head takes financial responsibility for the accommodation of his (her) own household members, household-head status is defined more by family relation[9]

[9] The variable of 'Head of a Household' is from the FRS data of whether an individual respondent is a householder or not. Originally, a householder in British culture is responsible for disciplines within the household and represents other household

than by income earnings. As household head is social status, a head of the household is not necessarily the highest income earner in the household.[10] Therefore, this variable is not necessarily indicating employment or retirement. The variable of 'No. of Children Taken Care of' is the number of children to whom an individual provides financial supports; so, it is smaller than total number of his (her) children. The binary indicator of 'Having a Health Problem' takes the value of one if an individual suffers an illness that limits daily activities and the value of zero otherwise. As 'Non Real-Estate Financial Wealth' refers to the value of liquid assets such as savings account, stock, bonds, and the like, it does not include the value of real estate asset.

The data of 'Per Capita GDP of the Region of Residence' and 'Unemployment Rate of the Region of Residence' are from the OECD dataset. In particular, the UK government office regions are North East, North West, Merseyside, Yorks and Humberside, East Midlands, West Midlands, Eastern, London, South East, South West, Wales, Scotland, and Northern Ireland. All the monetary values of different years are converted into year-2013 British pound based on the CPI (Consumer Price Index).

As the FRS dataset is not a panel data, it does not trace the same individual before and after abolishing the mandatory retirement. Thus, individuals observed before and after the abolishment are different persons from different households. Even though the FRS is carefully designed to be nationally representative with finding the survey respondents from random sampling based on the zip codes, it does not necessarily ensure that the group of 65-or-older respondents stays similar in every aspect over time before and after outlawing the mandatory retirement. Nonetheless, as shown in **Table 2**, most of the relevant characteristics of 65-or-older

members by paying rents and bills. Before the recent advancements of equal human rights that gave women more economic opportunities, a householder exercised a patriarchal authority over other household members.

[10] In particular, the variable of 'Head of a Household' is based on the raw data of 'HHOLDER' of the FRS. In the dataset, there is another variable of 'HRP' (household reference person) that is identified by whether a respondent individual earns the highest income in his or her own household, which should not be confused with the variable of 'HHOLDER'. An indicator of highest income earner in the household, by definition, is inherently affected by employment and retirement.

individuals are quite similar before and after the abolishment, except for skin color and women's non real-estate financial wealth. In the matching estimation, these differences are controlled to make treated and matched non-treated individuals as similar as possible for identifying the effect of the mandatory-retirement abolishment.

**Table 2] Descriptive Statistics for Characteristics of the Elderly Men and Women by the Abolishment of the Mandatory Retirement**

| | Elderly Men | | Elderly Women | |
|---|---|---|---|---|
| | Treated by Abolishment | Not-treated by Abolishment | Treated by Abolishment | Not-treated by Abolishment |
| | Mean (SD) | Mean (SD) | Mean (SD) | Mean (SD) |
| Age | 67.315 | 67.389 | 67.308 | 67.418 |
| | (1.724) | (1.693) | (1.689) | (1.701) |
| White (=1) | 0.778 | 0.890 | 0.773 | 0.894 |
| | (0.415) | (0.313) | (0.419) | (0.308) |
| Married (=1) | 0.754 | 0.744 | 0.651 | 0.656 |
| | (0.431) | (0.437) | (0.477) | (0.475) |
| Head of a Household (=1) | 0.958 | 0.945 | 0.079 | 0.085 |
| | (0.201) | (0.229) | (0.270) | (0.278) |
| No. of Children Taken Care of | 0.000 | 0.000 | 0.0004 | 0.001 |
| | (0.000) | (0.000) | (0.020) | (0.032) |
| Having a Health Problem (=1) | 0.340 | 0.347 | 0.356 | 0.340 |
| | (0.474) | (0.476) | (0.479) | (0.474) |
| Education (years) | 11.357 | 11.295 | 11.241 | 11.093 |
| | (2.464) | (2.446) | (2.131) | (2.013) |
| Owning a House (=1) | 0.654 | 0.621 | 0.678 | 0.644 |
| | (0.476) | (0.485) | (0.467) | (0.479) |
| Non Real-Estate Financial | 1.515 | 1.473 | 1.618 | 1.290 |
| Wealth (1000 £) | (5.479) | (2.905) | (5.962) | (1.606) |
| Per Capita GDP of the Region | 25.128 | 25.175 | 25.373 | 25.441 |
| of Residence (1000 £) | (5.782) | (5.357) | (6.104) | (5.719) |
| Unemployment Rate of the | 7.816 | 7.742 | 7.838 | 7.769 |
| Region of Residence (%) | (1.207) | (1.137) | (1.204) | (1.145) |
| No. of Observations | 2193 | 1713 | 2518 | 1973 |

(1) The variables with (=1) are binary indicators that take the value of one if the statement of the variable name is true and the value of zero otherwise.
(2) All the monetary values are converted in year-2013 British pounds based on the CPI.

## V. Estimation Results

To begin with, **Table 3** reports the results of the first step of the estimation (i.e., obtaining the

propensity score). Because the treatment assignment (whether an individual is affected by the abolishment of the mandatory retirement) is exogenously given, in the Probit regression estimation of the propensity score, the coefficient of the right-hand-side variables of **X** does not need to have economic interpretations for how they affect the treatment assignment. As there is no variation in the variable of 'No. of Children Looked After' for treated and not-treated elderly men (as **Table 2** shows), its coefficient is not able to be estimated for the elderly men. In addition, we test the validity of the Probit regression equation for obtaining the propensity score. The bottom part of **Table 3** displays the results of Pregibon specification test, indicating that the null hypothesis of misspecification is rejected at the 1% statistical significance level. Above all, the obtained propensity score is used for matching treated and not-treated individuals.

**Table 3] Propensity Score Estimation**

| Dependent Binary Variable : | Elderly Men | | Elderly Women | |
|---|---|---|---|---|
| Being Treated by the Abolishment of the Mandatory Retirement | Coefficient Estimate | (SE) | Coefficient Estimate | (SE) |
| Age | −1.622 | (1.112) | 0.225 | (1.047) |
| $Age^2$ | 0.012 | (0.008) | −0.002 | (0.008) |
| White | −0.539*** | (0.057) | −0.584*** | (0.054) |
| Married | 0.027 | (0.048) | 0.009 | (0.042) |
| Head of a Household | 0.169 | (0.096) | −0.063 | (0.071) |
| No. of Children Taken Care of | NA | | −0.366 | (0.754) |
| Having a Health Problem | −0.020 | (0.044) | 0.073 | (0.042) |
| Education | 0.002 | (0.009) | 0.019** | (0.010) |
| Non Real-Estate Financial Wealth | 0.054*** | (0.007) | 0.113*** | (0.010) |
| Owning a House | 0.099** | (0.045) | 0.110** | (0.044) |
| Per Capita GDP of the Region | 0.003 | (0.004) | 0.002 | (0.003) |
| Unemployment Rate of the Region | 0.045*** | (0.017) | 0.035** | (0.016) |
| P-value of Pregibon Test Against Misspecification | 0.000 | | 0.000 | |
| No. of Observations | 3906 | | 4491 | |

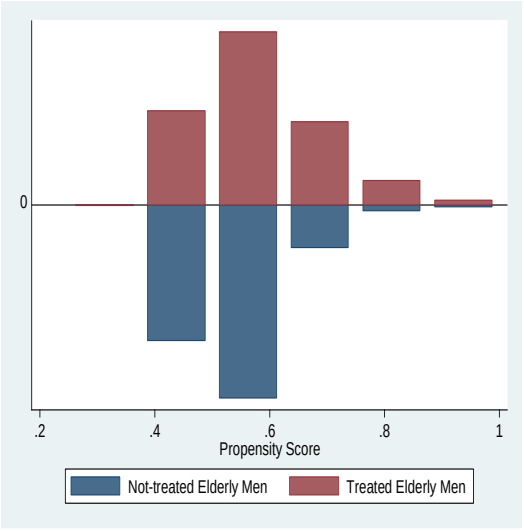
Note: ** and *** refer to being statistically significant at the 5% and 1% levels, respectively.

Before presenting the estimated effects of abolishing the mandatory retirement on public-pension benefit claim, retirement and employment, we report the results from testing the two key

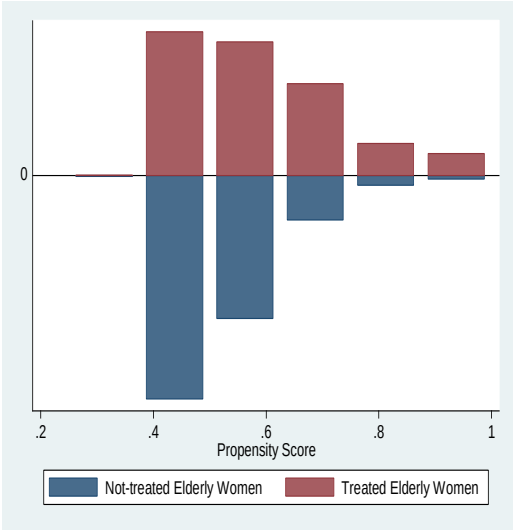
identification conditions of (3) and (7). Firstly, to test the overlap condition of (3), we examine the values of the estimated propensity scores for treated and not-treated individuals. As illustrated in the two panels of **Figure 1**, for both treated and not-treated men and women, the propensity score $p(\mathbf{X})$ is strictly greater than zero and lower than one, which shows that the identification condition of (3) is met.

**Figure 1] Distribution of Propensity Scores of Treated and Not-treated Individuals**

[a] Distribution of Propensity Scores of Treated and Not-treated Elderly Men




[b] Distribution of Propensity Scores of Treated and Not-treated Elderly Women




Secondly, we test whether the covariates of $\mathbf{X}$ used for the propensity score matching satisfy the balancing condition of (7). **Table 4** reports the balance test results obtained after matching. For each of the covariates $\mathbf{X}$, the difference between treated and *matched* not-treated individuals is not different from zero at any sensible statistical significance levels, which shows that each of the covariates $\mathbf{X}$ is balanced between treated and not-treated individuals after matching based on the propensity score $p(\mathbf{X})$. Notably, calculated according to Rosenbaum and Rubin (1985), the standardized potential bias of the set of the covariates $\mathbf{X}$ after matching is 4.8% for men of ages 65 – 70 and 5% for the female counterpart. Thus, **Table 4** suggests that the covariates of $\mathbf{X}$

satisfy the balancing condition of (7) for random assignment of the treatment of the mandatory-retirement abolishment. Overall, **Figure1** and **Table 4** show that the crucial conditions of (3) and (7) for identifying the effects of abolishing the mandatory retirement (i.e., estimating the parameter $\alpha$) are met.

**Table 4] Balance of Treated and Matched Not-treated Individuals**

| | Elderly Men | | Elderly Women | |
|---|---|---|---|---|
| Covariate | Mean Difference | (SE) | Mean Difference | (SE) |
| Age | 0.006 | (2.347) | −0.045 | (2.263) |
| $Age^2$ | 0.713 | (316.815) | −6.123 | (305.069) |
| White | 0.017 | (0.329) | 0.012 | (0.349) |
| Married | 0.002 | (0.603) | 0.011 | (0.683) |
| Head of a Household | 0.026 | (0.277) | −0.032 | (0.403) |
| No. of Children Taken Care of | 0.000 | (0.000) | 0.000 | (0.000) |
| Education | −0.364 | (3.680) | −0.314 | (2.842) |
| Having a Health Problem | −0.0005 | (0.645) | −0.031 | (0.673) |
| Owning a House | 0.028 | (0.597) | 0.020 | (0.596) |
| Non Real Estate Financial Wealth | 0.140 | (3.871) | 0.273 | (3.976) |
| Per Capita GDP of the Region | −0.559 | (8.600) | −0.296 | (9.240) |
| Unemployment Rate of the Region | 0.001 | (1.428) | −0.010 | (1.590) |
| Standardized Potential Bias (%) | 4.8 | | 5.0 | |

Note: The standardized potential bias is measured according to Rosenbaum and Rubin (1985).

Now, **Table 5** presents the results of estimating the effects of abolishing the mandatory retirement on public-pension benefit claim, retirement and employment. Notably, outlawing the mandatory retirement does not lower public-pension benefit claim rate of all the benefit-eligible individuals. In particular, abolishing the mandatory retirement *increases* public-pension benefit claim rate of the benefit-eligible elderly women by 0.61% points, while the abolishment lowers public-pension benefit claim rate of the benefit-eligible elderly men by 0.15% points. To understand this effect of outlawing the mandatory retirement on benefit-eligible individuals' public-pension benefit take-up behavior, we also estimate the effects on their retirement and employment, because public-pension benefit claim decisions of benefit-eligible individuals are

closely related to their labor supply decisions. First, **Table 5** shows that abolishing the mandatory retirement increases retirement rate of the benefit-eligible elderly women by 0.69% points, while the abolishment lowers retirement rate of the benefit-eligible elderly men by 3.19% points. This is consistent with the changes in public-pension benefit claim rate brought by the mandatory-retirement abolishment. Second, in contrast to the effects of the abolishment on public-pension benefit claim, retirement and part-time employment, which turn out to take opposite signs differing by gender, its effects on full-time and overall employments are positive for both genders alike. According to **Table 5**, the mandatory-retirement abolishment raises full-time employment rate by 1.64% points for the benefit-eligible elderly men and by 1.81% points for the benefit-eligible elderly women.

**Table 5] Effects of Outlawing the Mandatory Retirement**

| | Elderly Men | | Elderly Women | |
|---|---|---|---|---|
| Outcome | Estimated Effect | (SE) | Estimated Effect | (SE) |
| Public-Pension Benefit Claim | $-0.0015^{***}$ | (0.0003) | $0.0061^{***}$ | (0.0002) |
| Retirement | $-0.0319^{***}$ | (0.0007) | $0.0069^{***}$ | (0.0006) |
| Part-time Employment | $0.0013^{***}$ | (0.0004) | $-0.0063^{***}$ | (0.0004) |
| Full-time Employment | $0.0164^{***}$ | (0.0004) | $0.0181^{***}$ | (0.0002) |
| Overall Employment | $0.0177^{***}$ | (0.0006) | $0.0118^{***}$ | (0.0004) |
| No. of the Treated | 2193 | | 2518 | |
| No. of the Not-treated | 1713 | | 1973 | |

(1) ** and *** refer to being statistically significant at the 5% and 1% levels, respectively.
(2) In the parentheses are heteroscedasticity-robust standard errors, based on Abadie et al. (2004).

Notably, **Table 5** shows that responding to the mandatory-retirement abolishment, public-pension benefit claim rate and retirement rate move in the same direction, while public-pension benefit claim rate and employment rate do not always do so. This suggests that public-pension benefit claim is more closely related to retirement than to employment, which resonates with

various studies regarding the effect of public-pension policy on labor supply of the elderly (e.g., Coile and Gruber, 2007; Atalay and Barrett, 2015). Under the rationale of life-time consumption smoothing, public-pension benefit claim is more synchronized with retirement that entails no labor incomes afterwards than with being unemployed that can entail positive labor incomes afterwards. In addition, from **Table 5**, notice that an increase in retirement rate does not necessarily entail a decrease in employment rate. The relationship between these two rates is not one-to-one. When more individuals stop job searches for good or stop working for their own job (if they have one) for good, retirement rate increases. An individual does not necessarily have a job when he or she retires to exit the labor force. On the other hand, employment rate increases only when more individuals land a new job or retain their own job. Therefore, it is fully possible that both retirement rate and employment rate increase together.

Moreover, **Table 5** shows that abolishing the mandatory retirement increases public-pension benefit claim rate and retirement rate of women of ages 65 – 70, while it reduces these two rates of the male counterpart. To explain this contrast by gender, one may well consider *de facto* gender-discriminatory application of the mandatory-retirement abolishment. However, this explanation is not convincing because **Table 5** also shows that the abolishment increases full-time employment of the elderly women and men *alike*.

As another attempt to understand the findings of **Table 5**, we decompose our data by household-head status and then re-conduct the same matching estimation with the same covariates of $\mathbf{X}$. **Table 6** reports the results of this additional estimation. As demonstrated in the panel [a] of **Table 6**, abolishing the mandatory retirement reduces public-pension benefit claim rates of both benefit-eligible elderly men and women who are household head by increasing their employment rate and lowering their retirement rate. For both men and women alike, the effect of

abolishing the mandatory retirement on the elderly household-head individuals' labor supply is consistently positive, entailing decreases in their public-pension benefit claim rates. In particular, outlawing the mandatory retirement lowers public-pension benefit claim rate of the benefit-eligible household-head women by 12.85% points and that of the benefit-eligible household-head men by 0.57% points by making both household-head men and women work more.

**Table 6] Effects of Outlawing the Mandatory Retirement by Household Head Status**

| | Estimated Effect | (SE) | Estimated Effect | (SE) |
|---|---|---|---|---|
| [a] Head of Household | Elderly Men | | Elderly Women | |
| Public-Pension Benefit Claim | $-0.0057^{***}$ | (0.0004) | $-0.1285^{**}$ | (0.0034) |
| Retirement | $-0.0311^{***}$ | (0.0007) | $-0.1254^{***}$ | (0.0065) |
| Part-time Employment | 0.0004 | (0.0004) | 0.0219 | (0.0041) |
| Full-time Employment | $0.0123^{***}$ | (0.0004) | $0.0566^{***}$ | (0.0032) |
| Overall Employment | $0.0128^{***}$ | (0.0006) | $0.0785^{***}$ | (0.0051) |
| No. of the Treated | 2101 | | 199 | |
| No. of the Not-treated | 1618 | | 167 | |
| [b] Not a Head of Household | Elderly Men | | Elderly Women | |
| Public-Pension Benefit Claim | $0.3186^{***}$ | (0.0062) | $0.0188^{***}$ | (0.0002) |
| Retirement | $-0.0139$ | (0.0135) | $0.0132^{***}$ | (0.0006) |
| Part-time Employment | $-0.0239^{***}$ | (0.0087) | 0.0008 | (0.0004) |
| Full-time Employment | $0.0834^{***}$ | (0.0080) | $0.0135^{***}$ | (0.0002) |
| Overall Employment | $0.0594^{***}$ | (0.0113) | $0.0127^{***}$ | (0.0005) |
| No. of the Treated | 92 | | 2319 | |
| No. of the Not-treated | 95 | | 1806 | |

(1) ** and *** refer to being statistically significant at the 5% and 1% levels, respectively.
(2) In the parentheses are heteroscedasticity-robust standard errors, based on Abadie et al. (2004).

The panel [b] of **Table 6** shows that abolishing the mandatory retirement raises public-pension benefit claim rate of the benefit-eligible non-household-head men and women. In particular, the

abolishment raises public-pension benefit claim rate of the benefit-eligible non-household-head women by 1.88% points and that of the benefit-eligible non-household-head men by 31.86% points. Moreover, outlawing the mandatory retirement raises retirement rate of the benefit-eligible non-household-head women and reduces part-time employment rate of the benefit-eligible non-household-head men, while it increases full-time employment rates of both.

The findings from **Table 6** reveal that the effects on public-pension benefit claim of outlawing the mandatory retirement differ by household-head status, rather than by gender. Moreover, **Table 6** is not inconsistent with **Table 5** because most of the elderly men are household head, as opposed to the female counterpart. In a household, members of the household share consumption together and are usually led by a household head. Although a household-head individual is not always the only breadwinner of his (her) household, household-head individuals take the responsibility of financial supports for their own household members than non-household-head individuals. As the mandatory-retirement abolishment enables a head of a household to work more and to yield more labor income, other members of the household become able to consume *more* than before the abolishment. When the mandatory retirement is abolished, some of non-household-head individuals already have enough savings so that an increase in the resources by more work of their own household head enables them to maintain consumption smoothing if they retire now. The mandatory-retirement abolishment increases retirement rate of these non-household-head individuals to increase their public-pension benefit claim rate. At the same time, other non-household-head individuals who do not yet have enough savings still want to work more, even if their own household heads work more and earn more labor incomes. The mandatory-retirement abolishment also enables these non-household-head individuals to work more, because the abolishment is applied equally for both household-head individuals and non-

household-head individuals. As such, the mandatory-retirement abolishment increases retirement rate of non-household-head individuals increasing their public-pension benefit claim rate, while it increases full-time employment rate of both household-head individuals and non-household-head individuals. This logic for **Table 6** also can explain the seemingly-puzzling findings of **Table 5** because most of household-head individuals are male and most of non-household-head individuals are female.

Although both retirement decision and public-pension benefit claim decision fundamentally are individual decisions, the empirical findings of **Table 6** and **5** suggest that individual decision makers consider not only themselves but also other members of their own household with whom they share consumption. **Table 6** shows that how individual decision makers factor in other household members depends on household-head status. With other household members taken into account, 65-or-older individuals' responses to the mandatory-retirement abolishment are not purely individual, although the unit that the mandatory-retirement abolishment is applied is individual, not household. As a consequence, deviating from policymakers' expectation that the mandatory-retirement abolishment will reduce public-pension benefit claim rate of all 65-or-older individuals by lowering their retirement rate, the abolishment increases public-pension benefit claim rate and retirement rate of 65-or-older non-household-head individuals.

When we decompose the data by marital status (instead of household-head status)[11] and then re-conduct the same matching estimation with the same covariates, its results are not much different from those in **Table 5**. In fact, marital status is not an accurate indicator for household-

[11] On the other hand, decomposing the data by whether an individual is highest income earner within his or her own household is not suitable for identifying the effect of the mandatory-retirement abolishment on retirement and employment. As labor incomes are the major source of incomes, zero labor income due to either being unemployed or being retired will not make an individual be the highest income earner in his or her household. Thus, decomposing the data by whether an individual earns the highest income in his or her own household, instead of household-head status, generates a selection bias in estimating the effects of the mandatory-retirement abolishment on retirement and employment.

head status. That is, an elderly individual who is not married is not necessarily a head of household. In our data, 39.26% of the 65-or-older individuals who are not married are household head.

Most importantly, this paper newly discovers that the factor of household-head status needs to be taken into account for predicting the effect of abolishing the mandatory retirement on public-pension benefit claim. This paper finds that the mandatory-retirement abolishment reduces public-pension benefit claim of benefit-eligible household-head individuals, while the abolishment increases public-pension benefit claim of benefit-eligible non-household-head individuals. The policy implication of theses empirical findings is that outlawing the mandatory retirement does not necessarily achieve its goal of reducing public-pension benefit claim rate by reducing retirement rate. Thus, whether the impact of the mandatory-retirement abolishment on public-pension expenditure is negative or positive is not clear. Whether the mandatory-retirement abolishment reduces aggregate retirement rate of all benefit-eligible individuals or public-pension expenditure is not clearly determinable but needs to be calculated with taking the factor of their household-head status into account.

## VI. Concluding Remarks

Capitalizing upon the abolishment of the mandatory retirement from age 65 which took effect in 2011 in the United Kingdom, this paper estimates the impact of outlawing the mandatory retirement on public-pension benefit claim of benefit-eligible individuals of ages 65 – 70 through changes in their retirement and employment. Using the propensity-score matching estimator that purges out the time-trend effect, this paper finds that the mandatory-retirement abolishment raises public-pension benefit claim rate of benefit-eligible non-household-head individuals by raising their retirement rate, while the abolishment reduces public-pension benefit claim rate of

benefit-eligible household-head individuals by lowering their retirement rate. Because most of the elderly men are household head, as opposed to the female counterparts, the effects of abolishing the mandatory retirement on public-pension benefit claim and retirement turn out to be negative for benefit-eligible elderly men but positive for the female counterparts. Above all, this paper newly discovers that household-head status plays an important role in predicting whether the mandatory-retirement abolishment lowers public-pension benefit claim rate by reducing retirement rate of benefit-eligible elderly individuals.

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