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The Displacement Effect of Public Pensions on the Accumulation of Financial Assets

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Abstract

The generosity of public pensions may depress private savings and provide incentives to retire early. While there is plenty of evidence supporting the latter effect, there remains considerable controversy whether public pensions crowd out private savings. This paper uses international micro-datasets collected over recent years to investigate whether public pensions displace private savings. The identification strategy relies not only on cross-country differences in generosity but also on differences in the progressivity or non-linearity of pension formulas across countries. We estimate that an extra dollar of pension wealth depresses accumulated financial assets around the time of retirement by 22 cents. An extra ten thousand dollars in public pension wealth reduces the average retirement age by roughly one month which implies an elasticity of retirement years with respect to pension wealth of -0.15 .

Keywords

saving; social security; displacement; international comparisons

1. Introduction

In his seminal article, Feldstein (1974) argues that the generosity of public pensions may affect the amount of private savings held at the time of retirement through two effects: a displacement effect due to the fact that a marginal increase in public benefits will lead to higher consumption over the entire life-cycle, hence leading to lower private savings; or an early retirement effect due to the increase in lifetime income which may imply larger savings at the time of retirement to finance consumption through retirement. Hence, the effect is ambiguous. Furthermore, although the standard model predicts perfect displacement (one additional dollar of pension assets reduces financial assets by one dollar), conditional on a given retirement age, this will not hold in more general models with borrowing constraints and uncertainty. Knowing the magnitude of these effects is important as it may directly affect consumption possibilities in retirement and helps predict the reaction of private savings and retirement age to public pension generosity changes.

Feldstein (1974) uses aggregate time series data from the U.S. and reports displacement effects ranging from –30 cents to –50 cents per dollar of pension benefits depending on the specification. Since then, a wide range of estimates has been reported using different identification strategies and measures of non-pension and pension assets. Estimates of the effect of pension wealth on non-pension wealth range from close to zero (Kotlikoff, 1979) to close to negative one (Gale, 1998). The bulk of those studies have been done on U.S. data but there is also international evidence exploiting pension reforms in several European countries. Attanasio and Brugiavini (2003) find evidence that public pension benefits affect negatively saving rates of Italian households following the 1992 pension reform. Similarly, Attanasio and Rohwedder (2003) analyze pension reforms in the U.K. and also find evidence of a displacement effect. Kapteyn, Alessie and Lusardi (2005) find that cohort differences in accumulated wealth can in part be explained by differences in social security wealth. Some studies have also exploited cross-country differences in assets. Feldstein (1979) uses aggregate data from 12 countries and finds a displacement effect of 37 cents per dollar of pension benefits. Kapteyn and Panis (2005) consider a case study using micro data from three countries (the U.S., the Netherlands and Italy) to show that in countries where public pensions are generous, individuals in general hold less savings. Disney (2006) uses a panel of OECD countries over time to estimate a displacement effect. He shows that the closer a system is to actuarial fairness, the more potential there is for an offset as pensions are then a clear substitute for private savings. The degree of actuarial fairness may vary across groups differentiated by lifetime income because of the progressivity of public pensions.

In this paper, we present new evidence on the displacement effect of public pensions by using cross-country variation in the generosity and progressivity of pensions. We use micro-datasets from many more countries than those considered in Kapteyn and Panis (2005) to construct income replacement rate and private saving measures by education level and marital status, which are good proxies for lifetime earnings. We estimate reduced-form equations derived from a standard life-cycle model in order to quantify the displacement effect and use the progressivity of the pension formula, as well as cross-country variation in average generosity, to estimate the effects.

We find evidence of a modest displacement effect of public pensions on accumulated financial assets of 22 cents for every additional dollar of public pension wealth. We also report a substantial income effect of pension wealth on the retirement age with an estimated elasticity of years of retirement with respect to pension wealth of 0.15. We find that these results are robust when accounting for differential mortality, bequest motives and including a measure of occupational/private pension benefits in the replacement rate. These results are important as they show that there is imperfect substitution between private savings and public pensions and that in general workers retire earlier when facing more generous pensions.

The paper is organized as follows. We first present within a simple theoretical model the factors that affect the displacement and early retirement effect and derive estimable equations that can be used to assess the size of such effects on private savings. We then discuss in Section 3 our empirical strategy. Section 4 reports the results while we conclude in Section 5.

2. Theoretical Model

We setup a very simple life-cycle model of consumption without uncertainty and derive optimality conditions guiding the choice of consumption and retirement which depend on

the generosity of public pensions. These conditions guide the empirical strategy we use. Our setup is closest to Gale (1998) and Laitner and Silverman (2007).

Assume an individual who is making decisions from time $t = 0$ to T , regarding consumption c_t and work. We assume that individuals work from period $t = 0$ to some time $t = R$ at which point they retire. For simplicity we assume individuals get constant income $y_t = y$ while working.

The public pension received once retired is given by $b_t(R) = \alpha(R)\varphi(y)$. The term $\alpha(R)$ is an actuarial factor, which adjusts the pension to the retirement (claiming) age and $\varphi(y)$ is the benefit that would be received at normal retirement age. For example, in the United States it is the primary insurance amount (PIA). In most countries, the earnings-related part of the pension formula is progressive such that $\varphi'(y) < 1$, and $\alpha'(R) > 0$ such that later retirement is rewarded by higher benefits. Assuming an interest rate, r , the present value of pension benefits upon retirement is given by

$$B(R) = \alpha(R)\varphi(y) \int_0^{T-R} e^{-rt} dt = \alpha(R)\varphi(y)\xi(R).$$

The present value of benefits at retirement is actuarially fair (does not depend on R) if $\alpha(R)\xi(R) = k$ where k is a scaling parameter. In that case, pension wealth at the time of retirement is $B(R) = k\varphi(y) = \varphi_k(y)$.

Now assume the individual derives utility from consumption prior to retirement given by

$$u(c_t) = \frac{c_t^{1-\gamma}}{1-\gamma}, \gamma > 0$$

and after retirement $v(c_t) = u(c_t) + \Gamma$. The parameter γ captures the curvature of the utility function. The parameter Γ represents the utility of leisure in retirement. Our formulation of the utility function implies that consumption and leisure are strongly separable. This simplifies the exposition without altering the main mechanisms. Finally, assume the individual has a discount rate ρ .

Assuming the individual starts with zero assets, $a_0 = 0$ and has to die with non-negative assets, $a_T \geq 0$, he solves the following optimization problem

$$\begin{aligned} \max_{\{c_t\}_{t=0,\dots,R}} & \int_0^R e^{-\rho t} u(c_t) dt + \psi(a_R + B(R), R) \\ \text{s.t.} & \\ & \dot{a}_t = ra_t + y_t - c_t, a_0 = 0 \end{aligned} \quad (1)$$

The post-retirement indirect utility ψ is given by solving the maximization problem

$$\begin{aligned} \psi(a_R + B(R), R) = \max_{\{c_t\}_{t=R, \dots, T}} \int_R^T e^{-\rho t} v(c_t) dt \\ \text{s.t.} \\ \dot{a}_t = r a_t - c_t \\ a_R = A + B(R), a_T \geq 0 \end{aligned} \quad (2)$$

The solution for consumption, given the retirement age, can be directly derived from the first-order conditions to (1) and (2). The solution is given by

$$\frac{\dot{c}_t}{c_t} = \frac{r - \rho}{\gamma} = \eta, \quad c_0 = \delta [Y(R) + e^{-rR} B(R)] \quad (3)$$

where $\delta = \lambda / (e^{\lambda T} - 1)$, $\lambda = \eta - r$, and life-time discounted earnings is defined as $Y(R) = y \int_0^R e^{-rs} ds$. The term λ measures the difference between the rate of growth in consumption (η) and the real interest rate. In what follows, we focus on pension-induced saving and thus assume $r = \rho$. In this case, we see that consumption is constant over the life-cycle: in each period the individual consumes a constant fraction of his life-time income. A consequence of this assumption is also that the fraction of life-time income devoted to consumption does not depend on preferences. But life-time income may depend on preferences as it is affected by the choice of the retirement age.

The first-order condition for the timing of retirement takes the form

$$c_R^{-\gamma} [y + B'(R)] = \Gamma \quad (4)$$

where B' denotes the derivative with respect to R . The left-hand side represents the marginal benefit from delayed retirement (additional income converted into utility using the marginal utility of consumption) and the right-hand side represents the marginal cost of leisure foregone by retiring later.

Substituting optimal consumption and taking logs we get

$$\log[y + B'(R)] - \gamma \log[Y(R) + e^{-rR} B(R)] = \log \Gamma - \log[\delta^*] \quad (5)$$

where $\delta^* = \delta^{-\gamma}$. This equation implicitly defines the retirement age. The first term on the left-hand side captures the “accrual” effect of postponing retirement: more income directly from earnings, and the accrual in the present value of pension benefits. Gruber and Wise (1999) document this term and find that in most countries there is a penalty to continuing work (i.e. $B'(R) < 0$) past the earliest age when individuals are eligible for public pension benefits. The second term on the left-hand side captures the wealth effect. It increases with the concavity of the utility function.

A closed-form solution to R will generally not be possible. However, we can compute how the optimal retirement age reacts to the generosity of pensions. For simplicity, assume $\varphi_k(y) = k\varphi y$ where φ is the replacement rate. Differentiation of (5) with respect to φ , which amounts to an increase in the generosity of the pension system, shows that increasing generosity will reduce the retirement age provided $B'(R)$ is either small or negative as is the case in many countries.

The effect of changing the replacement rate on the consumption path can be found analogously. Then, differentiating the solution for initial consumption found in equation (3) yields,

$$\frac{dc_0}{d\varphi} = \delta e^{-rR} [ky + (y - r\varphi_k(y)) \frac{dR^*}{d\varphi}] \quad (6)$$

The first term within the brackets represents the displacement effect. The individual can consume the marginal increase in his retirement benefit (discounted) and spread it over the entire life-cycle. Hence savings decrease. However, there is a second effect due to induced earlier retirement. If the income effect is large, the individual will retire earlier which leads to a drop in lifetime income. This is financed by an increase in savings during the working life. The net effect of increasing the replacement rate on savings is therefore unclear and depends on preferences as well as the extent to which pensions replace earnings.

3. Empirical Strategy

To derive testable implications it is first useful to rewrite the solution in terms of assets. Assets around the time of retirement, say $t = R$, is:

$$a_R = y\pi(r, R) - c_0 e^{Rr} \pi(\lambda, R) \quad (7)$$

where $\pi(s, t) = \frac{e^{st} - 1}{s}$. Equation 7 represents accumulated assets at the time of retirement as the difference between life-time income and life-time consumption up to that point. The adjustment factors account for interest accumulated and growth in consumption. Substituting the solution for initial consumption we obtain

$$a_R = (1 - q(\lambda, R)) y\pi(r, R) - q(\lambda, R) \varphi_k(y) \quad (8)$$

where $q(\lambda, R) = \frac{e^{\lambda R} - 1}{e^{\lambda T} - 1}$ is an adjustment factor. The term $q(\lambda, R)$ measures the offset in non-pension wealth at age R from augmenting pension wealth by one dollar at age 0. A change of one dollar in pension wealth does not translate in a change of one dollar in wealth at the time of retirement. Because consumption is increased at each age, the cumulative effect on non-pension wealth at an age less than T is less than unity. With a discount factor of 3%, life expectancy of 80, start age of 20, retirement age of 60, this factor q is 0.93. Obviously it is lower for earlier retirement ages t . Hence, by adjusting pension wealth by that factor, one gets the prediction of a perfect displacement effect at the time of retirement. A one dollar increase in adjusted pension wealth leads to a one dollar decrease in assets at the time of retirement (Gale, 1998).

An alternative specification involves dividing both sides of (8) by earnings, yielding a simple bivariate relationship

$$\frac{a_R}{y\pi(r, R)} = \alpha(\lambda, R) - q(\lambda, R) \frac{k\varphi(y)}{y\pi(r, R)} \quad (9)$$

where $\alpha(\lambda, R) = 1 - q(\lambda, R)$ is an intercept. The left hand-side is the ratio of assets to lifetime earnings. The right hand side is the adjusted ratio of pension wealth to lifetime earnings. We define the left-hand side of equation 9 as $A = a_R / y\pi(r, R)$. It is a measure of wealth at the

time of retirement normalized by income during the working life. We also define $REP = q(\lambda, R) \frac{k\varphi(y)}{y\pi(r, R)}$ as the adjusted replacement rate an individual gets in retirement.

We will exploit two sources of identification to estimate equation (9). First, we will exploit cross-country differences in the generosity of public pensions. This cross-country strategy is closest to the strategy used by Feldstein (1979). The second strategy we use will exploit within-country differences in the generosity of public pensions across socio-demographic groups. These differences are induced by the progressivity of the formula. We discuss the merit of both these identification strategies by first presenting the cross-sectional identification strategy used in the literature.

First assume we have access to cross-sectional data at the time of retirement where we observe $z_i = (y_i, a_i, R_i)$, $i = 1, \dots, N$ and can compute the quantities A , REP as defined below equation 9, and assuming $r = \rho$ so that $\lambda = -r$. Then the following regression

$$A_i = \alpha_0 + \alpha_1 REP_i + \varepsilon_i, \quad i = 1, \dots, N \quad (10)$$

should yield a coefficient estimate of $\alpha_1 = -1$ if the model is correctly specified and $\lambda = -r$. In practice, studies focusing on cross-sectional variation within a country identify the effect from covariation in z_i (the individual data) but not in $\varphi()$. The covariation in z_i is problematic if there is, for example, unobserved heterogeneity in tastes for leisure and consumption. An alternative source of identification is available if data from multiple countries are available. In that case variation across countries in REP could help identify α_1 . Following Feldstein (1979), we will estimate a cross-country regression of the form

$$A_c = \alpha_0 + \alpha_1 REP_c + \varepsilon_c, \quad c = 1, \dots, C \quad (11)$$

This strategy deals with unobserved heterogeneity across individuals. However, other country differences might explain the cross-country differences in the replacement rate and the asset-to-income ratio, resulting in bias due to omitted cross-country heterogeneity.

To address this possibility we will pool cross-sectional observations across countries. This strategy exploits similarities across individuals in different countries in terms of tastes, leisure, and economic opportunities. We group individuals by education level and marital status. This allows us to restrict the source of variation used to identify α_1 to within-country and group differences in A and REP . We estimate the following specification

$$A_{i,c} = \alpha_{0,c} + \alpha_{0,i} + \alpha_1 REP_{i,c} + \varepsilon_{i,c}, \quad i = 1, \dots, N, c = 1, \dots, C \quad (12)$$

where both group and country fixed effects are included. The key identifying assumption is that differences in unobserved tastes across groups are the same across countries. Assuming that k and R are the same across countries, the key source of variation in replacement rates is given by deviation in $\varphi(y)/(y)$ from their country mean and group means. If replacement rates do not depend on earnings, there will be no variation in replacement rates in response to deviation from their country or group means. Hence, progressivity is crucial.

Estimation of the asset equation above effectively controls for the retirement age since we substituted the solution for optimal consumption conditional on the retirement age. In that sense it allows us to estimate the partial displacement effect. However, the net effect depends on the magnitude of the wealth effect of pensions on retirement age. We also

estimate the wealth effect by following a similar strategy with respect to retirement age. In particular, we use the following specification:

$$RET_{i,c} = \beta_{0,i} + \beta_{0,c} + \beta_1 SSW_{i,c} + \beta_2 y + v_{i,c}, i=1, \dots, N, c=1, \dots, C \quad (13)$$

where $RET_{i,c}$ is the average retirement age for cell (i,c) and $SSW_{i,c} = k_c \phi_c(y_{i,c})$.

Because both retirement age and asset equations form a system of equation, we could in principle use features of the Social Security systems as exclusion restrictions to estimate parameters. However, the inclusion of country fixed effects makes it virtually impossible to find exclusion restrictions that vary across individuals and countries. We also abstract from potential efficiency gains from estimating both equations jointly. As it turns out, our estimates are relatively precise given the small sample used in estimation.

It is important to consider whether any unobserved differences across countries and groups might invalidate the identification strategy we use. First, for this to be the case these differences would have to affect differentially each group across countries and be correlated with the replacement rate from public pensions. Hence, pure differences across countries, such as differences in the discount rate, rate of return, should not affect our estimates. Nevertheless, we will allow in the empirical analysis mortality rates to be different across countries. Second, the unobserved differences would have to be unobserved differences in wealth normalized by earnings rather than in levels of wealth. The theoretical model tells us that normalized wealth levels should be independent of lifetime earnings once we condition on the replacement rate measure. It is much more difficult to think of plausible unobserved differences in normalized wealth that correlate with the replacement rate than it is to think of unobserved differences in the levels of wealth which would correlate with the replacement rate. We investigate empirically other potential challenges to the identification strategy, such as cross-country variation in differential mortality and bequest motives, in section 4.3.

3.1. Data Sources

We use data from a combination of micro-data sources to implement our empirical strategy. To merge these data sources together, we aggregated data by cells defined by education level and marital status (defining observations i). We used three education categories as defined by the ISCED classification (primary, secondary, and tertiary) and two marital status categories (married/partner, single). The rationale for using those characteristics is that education level and marital status are powerful measures of lifetime earnings. Variation in lifetime earnings ultimately allow us to exploit cross-country differences in the progressivity of pension formulas. Furthermore, we use net rather than gross measures because additional variation is created through differences in the progressivity of the income tax function across countries.

We aggregate data on assets and retirement ages from aging surveys in 12 countries covering the age 50+ population: the Health and Retirement Study (HRS), the English Longitudinal Study of Ageing (ELSA) and the Survey of Health, Ageing and Retirement in Europe (SHARE) which cover 10 continental European countries (Austria, Belgium, Denmark, France, Greece, Germany, Italy, Netherlands, Spain, and Sweden). We use data collected in the 2004 wave of each survey (wave 7 of HRS, wave 2 of ELSA, and wave 1 of SHARE). The sample includes retired males aged 65 to 75. By that age, most males have retired and thus provide an adequate sample to measure assets and retirement age (a_R , R).

We computed a measure of median financial assets, a_R , comparable across countries. The data appendix gives details on the construction of the variable. Financial assets such as savings, stocks, and retirement accounts are relatively liquid and one would expect the

displacement effect to operate for those in particular. Other assets, most notably housing, are much less liquid, and participation in the housing market likely differs across countries and groups for reasons other than the generosity of pensions. We constructed the retirement age, R , using self-reports from HRS, ELSA, and SHARE. For SHARE, no direct question asked about the retirement age. We therefore used respondent reports of the date when the last job ended to construct the retirement age.

The aging surveys have limited measures of lifetime earnings. Hence, we constructed pre-retirement net earnings from longitudinal surveys in the same set of countries. For the U.S., we used the 1994 to 1997 Panel Study of Income Dynamics (PSID), while for the European countries, we used the 1994 to 1997 waves of the European Community Household Panel (ECHP). We derived net individual and household earnings and top-coded the 99th percentile of each cell. The ECHP asks directly for net household and individual annual earnings. The PSID asks only about gross household and individual annual earnings. We used the OECD tax calculator to compute net earnings. We defined our measure of permanent earnings, y , as average earnings from age 25 to 55. We constructed the lifetime measure of earnings using an interest rate of 0.03, a start age of 20, and the average retirement age within each cell.

Aging surveys ask retired individuals about their current gross pension benefits. Getting a measure of net pension benefits is not straightforward as benefits are taxed at the household level in some countries. Instead, we used data from the OECD on net replacement rates as a function of relative net average lifetime earnings. The OECD calculates these net replacement rates, defined as the net benefit divided by net average lifetime earnings. We obtained a complete mapping for the entire range of lifetime earnings for each country. These replacement rates are calculated on an individual basis and do not include spouse benefits. We used net earnings computed from the ECHP and PSID to map a replacement rate to each spouse and compute household net pension benefits, $\phi(y)$.

For couples, we calculate what each spouse is entitled to separately. The OECD does not publish replacement rates at the household level. Hence, this implies that we do not take account of the possibility that some (few) particular programs might provide spouse benefits to partners with a limited earnings history. Our analysis thus ignores program entitlements to couples, which may affect retirement behavior (e.g. Blau, 1997). In the U.S., Blau (1997) finds weak effects of the spouse benefit provision on labour supply, a provision that pays a benefit based on the maximum of the individual's benefit and half the spouse's benefit. According to the U.S. Social Security Fast Facts of 2004, fewer than 7% of new beneficiaries were receiving spouse benefits in 2004. Other measures commonly found include differential basic and supplemental pensions for singles and couples (e.g. Denmark).

We computed pension wealth assuming a discount factor of 3% and using country and gender specific survival probabilities. We used mortality rates from the Human Mortality Database for the year 2004 (www.mortality.org). This formed our estimate of k . Pension wealth was then defined as $SSW = k\phi(y)$. The factor $q(-r, R)$ was estimated using the average retirement age in each cell, a discount rate of 3% and T equal to life expectancy as of age 20 (roughly 56–60 years). The estimated factor ranges from 0.91 to 0.96. The remaining variables A and REP were constructed by dividing each quantity by lifetime earnings. Table 1 gives descriptive statistics by country on the main variables used to derive A and REP .

In our analysis, we do not include measures of private pension replacement rates. These are notoriously difficult to measure, particularly because of the heterogeneity found in plans even within a country. Another problem is that in some countries, individuals can cash out

the value of their pension upon retirement, which implies that the pension will already show up in wealth. It is generally difficult to get information on such cash-outs. Another problem is that tax treatment of pension income varies across countries and type of pensions such that net measures are difficult to compute. In our robustness analysis, we assess the sensitivity of the results by including an imperfect measure of private pension wealth in the replacement rate calculation.

We use the number of observations in each cell to weight observations in the regressions. This was done to reflect the uncertainty in the cell estimates for those with few observations. We discard cells with fewer than 5 observations. The point estimates were generally robust to the cut-off used but slightly more dependent on whether we weighted or not. The HRS and ELSA have much larger samples than other surveys, which tended to increase the influence of these cells relative to smaller cells representing some European countries (particularly among the highly educated and the unmarried). The results were however robust to scaling U.S. and England counts back to the average cell size among other European countries.

4. Results

We first present evidence that OECD net replacement rates reveal substantial differences in progressivity. In Figure 1 we present net replacement rates, defined as net pension benefits divided by annual net earnings $\phi(y)/y$, as a function of relative annual earnings y for some of the OECD countries in our analysis (OECD Pension at a Glance, 2005). In general, replacement rates decline with earnings, reflecting progressivity but the degree of progressivity is also quite different across countries. This is crucial for the identification strategy in equation (9).

4.1. Assets

We estimate three different regressions of A on REP . First, we aggregate up at the country level and perform a country-level regression. This is quite similar to the cross-country regressions done by Feldstein (1977, 1978) although we use age-group-specific asset (those age 65–75) and earnings information (from age 25–55 in 1994). Figure 2 plots the estimates of A and REP for each country in the analysis along with a regression line. A negative relationship is clearly visible and the regression estimate reported in column 1 of Table 2 suggests an estimate of $\hat{\alpha}_1 = -0.13$ [t=2.79]. Using a similar strategy, Feldstein (1977) reports an estimate of -0.37 .

We then estimate a pooled regression where we include education and marital status fixed effects but no country fixed effects. This is more akin to empirical specifications using cross-sectional variation in earnings although it also uses variation in replacement rates across countries. Column 2 of Table 2 reports the results. The estimate is now $\hat{\alpha}_1 = -0.09$ [t=2.8]. Both these strategies assume the absence of a correlation between unobserved differences in normalized assets and replacement rate across countries and groups. The last column of Table 2 reports estimates of equation 9 with both country and group fixed effects. These estimates rely primarily on differences in the progressivity of the pension benefit formula across countries. The estimate of the displacement effect is $\hat{\alpha}_1 = -0.218$ [t=2.4]. For each dollar increase in pension wealth, this estimate suggests that financial wealth decreases by roughly 22 cents.

Under a simple life-cycle model α_1 should be unity. There are many possible reasons why our estimate is much below that value. For example, individuals may not be strictly life-cycle savers; they may operate under uncertainty; or they may be liquidity constrained. Accounting for these reasons is much beyond the scope of this paper. An additional reason,

which is likely important, is measurement error: our calculation of the actual Social Security wealth of each individual may have error; but in addition individuals may use different values of Social Security wealth from actual, true values when deciding on the amount to save. Measurement error will bias downward estimates of α_1 .

In Figure 3, we check whether the estimate effect is due to any particularly large value. To accomplish this we first regress the left-hand variable (asset to income ratio) and the right-hand variable (replacement rate) on country and cell-specific fixed effects. We then take the residuals from those regressions. Due to properties of least-squares, regressing the asset residual on the replacement rate residual yields the same estimate of α_1 . We plot those residuals in Figure 3 along with the regression line. We also include an indication of the size of each cell. The larger the “bubble” around a cell, the larger is the number of observations used in the cell. One can see that the results are not driven by any particular observations. Rather, there is a very clear negative relationship between these residuals. Overall, using different identification strategies, we estimate an effect ranging from -9 to -22 cents per additional dollar of pension wealth with the preferred estimate being -22 cents.

4.2. Retirement Age

In Table 3, we present regression results for the effect of pension wealth on the timing of retirement. The first column uses country-level observations as in Table 2. The point estimate is 0.003 but fairly imprecisely estimated ($t=1.53$). In the second column, we pool cells from all countries and obtain an estimate of -0.007 , now statistically significant at the 5% level ($t=2.05$). Finally, the last column of Table 3 reports estimates with country and group fixed effects. The estimate is somewhat larger (-0.011) and still precisely estimated ($t=-3.24$). Figure 4 shows that this is not due to any particular outlier. This suggests a substantial wealth effect. For a ten thousand dollar change in pension wealth, the retirement age decreases by roughly one month. The estimated elasticities of expected years of retirement with respect to pension wealth are reported below the point estimates in Table 3. For column 3, the elasticity is: -0.146 . This implies that a 10% change in pension wealth leads to 1.5 percent change in the retirement years. However given the relatively small differences in retirement age across countries (order of 5%) and large difference in pension wealth, even small elasticities have the potential of explaining the heterogeneity in retirement age. Take two countries with very different generosity of public pensions, the Netherlands and England. Our estimate of mean pension wealth in the Netherlands is \$317.6 thousand while it is \$222.2 thousand in England. The Dutch retire earlier than their British counterparts: the average retirement age is 59.1 in the Netherlands compared to 61.2 in England. The effect from the third specification in Table 3 yields a predicted difference of 1.06 years while the actual difference is 2.1 years. Hence, this would explain 50% of the gap in the retirement age between the two countries

4.3 Robustness

As we mentioned in section 3, our identification strategy might be invalid if there were omitted variables, which vary differently across groups and countries. Constant differences across countries are less problematic because of the inclusion of country fixed effects. Similarly, constant differences across groups are also controlled by group dummies. Below we discuss three possible omitted differences and assess the robustness of our results to those differences.

4.3.1 Private Pensions—The prevalence private pensions is quite heterogeneous across groups and countries. This in turn could imply that we are not adequately measuring financial resources and in particular replacement rates in retirement. For reasons we already discussed, we have not included private pension benefits in our measure of replacement

rates. Here, we include median self-reported private pension income of retirees which we add to the predicted net public pension benefits we computed earlier. This leads us to redefine *REP* as a replacement rate which includes both public and private pensions. As mentioned earlier, these private income receipts probably lump together cash-outs from defined contribution plans with annuity payments. Second, these are gross of taxes while public pension benefits measured are net of taxes. Nevertheless, it provides a useful check on the sensitivity of our results to private pension generosity and take-up.

4.3.2 Differential Mortality—Differential mortality across groups should not affect our estimates as long as those differences are similar across countries. There are few estimates of differential mortality by lifetime income, marital status or education *across* countries. Delavande and Rohwedder (2008) use subjective survival to derive estimates by *current* wealth, income and education. They find striking similarity in the direction of the gradient (richer more educated expect to live longer) although the magnitude of the differences differs across countries. Because the rich generally have longer life spans, this should undo some of the progressivity of public pensions, the key identifying variation we use. In countries where differential mortality is stronger, we could be over-estimating the degree of progressivity in public pensions and vice versa for countries where it is smaller. Following the analysis of Delavande and Rohwedder (2008), we add to the specification a variable measuring the average 10-year subjective survival probability within a cell.

4.3.3 Bequest Motives—Another challenge to our identification strategy might stem from the presence of other motives for savings. One such motive is saving to leave a bequest. Individuals with higher lifetime income are more likely to leave bequests and the propensity to leave a bequest may differ across countries because of estate tax laws or different social norms.

Each survey we use (HRS-SHARE and ELSA) contains questions on the subjective probability of leaving a bequest. One question asks about the probability of leaving a bequest above \$10,000 USD (HRS), 50,000 GBP (ELSA) and 50,000 Euros (SHARE). If the answer is zero to that first question then respondents are asked what is the probability they will leave any bequest. We construct a variable measuring the probability of any bequest which we define as the fraction of respondents within a cell who provide a positive probability to any of these two questions. The average probability is 80% and is larger for those with more education and for married individuals. We add this variable to the specification and expect a positive relationship with the wealth to assets ratio. It is less clear what effect a larger probability of a bequest has on the retirement age.

In Table 4, we report the regression estimates for both the asset and retirement age regression. In general, the parameter estimates reported in Table 2 and 3 are robust to the inclusion of subjective bequest and survival variables and the inclusion of private pension benefits in the definition of the replacement rate. In the wealth equation, the point estimate of the displacement effect goes up slightly (−0.24). Overall, these robustness checks suggest that the generosity of private pensions, mortality as well as differences in bequest motives are unlikely to bias our results.

Other motives for savings, in particular precautionary motives, are not as straightforward to measure. On the one hand, earnings risk is likely to be unimportant in explaining savings around the time of retirement. This precautionary motive is likely stronger at younger ages, when there is more uncertainty about wage growth, etc. Another potentially important risk which could vary across countries and groups is health spending risk. This could lead to a precautionary motive, which could differ across groups and countries. However, we were not able to construct a group specific measure, nor a proxy of out-of-pocket risk, which

would be comparable across countries. It is unclear how out-of-pocket risk correlates with lifetime earnings and replacement rates. Finally, we have not considered less liquid assets such as housing equity which is a form of saving for old age. The important task of introducing these risks along with bequest motives, differential mortality and housing wealth into our model of savings is left for future research.

5. Conclusion

This paper provides new evidence about the displacement and early retirement effects of public pensions using both cross-country and within-country variation in the generosity of public pensions. Earlier evidence either relied on within-country variation across individuals or over time, or on aggregate cross-country data. Using our preferred specification, our estimates suggest a displacement effect of roughly 22 cents of financial assets for every additional dollar of public pension wealth and an early retirement effect of approximately one month for every additional \$10,000 in public pension wealth, or an elasticity of expected retirement years with respect to pension wealth of -0.15 . These estimates provide new evidence which confirms that public pensions likely depress asset accumulation, although by less than what the standard life cycle theory predicts. There are many reasons why this offset is imperfect but little research has examined those in detail. One interesting avenue for further research is to exploit other institutional differences across countries, such as the degree of borrowing constraints and labor market regulations, to enrich our understanding of the reasons behind the imperfect displacement effect. Our modeling of retirement and saving decisions was simple for reasons of tractability and because of data limitations. But with additional waves of aging surveys, and retrospective life histories, one important avenue for further research is to enrich structural models of life cycle decision making with variation in institutions across countries.

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Data Appendix

Our analytic dataset is derived from a number of sources. Our household wealth measures are derived from “aging” surveys: Wave 1 (2004) of the Survey of Health, Ageing and Retirement (SHARE), Wave 2 (2004–2005) of the English Longitudinal Study of Ageing (ELSA), and Wave 7 (2004) of the Health and Retirement Survey (HRS). Our Household Pre-Retirement Net Earnings measure is derived from the European Community Household Panel (ECHP) (1994–1997) and the Panel Study of Income Dynamics (PSID) (1994–1997). Our pension information comes from replacement rate calculations as a function of net pre-retirement earnings made available to us by OECD. Finally, we use the 2004 period life tables for males from the Human Mortality Database (www.mortality.org).

Sample

We define cells for each country by marital status (defined as married/partnered or single) and three education level (primary, secondary, tertiary) following the ISCED classification. The definition of these variables is comparable across the various datasets we used. We select men aged 65–75 and not working in the aging surveys and select men aged 25–55 in the ECHP/PSID surveys.

Financial Wealth

Our financial wealth variable includes bank account balances, the values of stocks and bonds, and money saved in individual retirement accounts. Values are expressed in 2004 dollars and are adjusted for purchasing power parity (PPP) using OECD comparative price levels. We compute the median wealth within each country-education-marital status cell. The original variables used are as follows:

	HRS	ELSA	SHARE
<i>Bank Accounts</i>	HQ344-HQ348	IASAVA	HBACCV
<i>Stocks</i>	JQ316-JQ320	IASS, IAUIT	HSTOCV, HMUTFV
<i>Bonds</i>	JQ330-JQ334, JQ356-JQ360	IANPB, IANS, IABG	HBONDV
<i>IRAs</i>	JQ162, JQ165_1 - JQ169_3	IASISA, IATI, IACISA, IAIP	HIRAV

Pre-Retirement Net Earnings

Our pre-retirement net earnings variable is derived from earnings data from four waves of ECHP and PSID, 1994 – 1997. Values are expressed in 2004 dollars, PPP-adjusted, using OECD wage growth data and PPP-adjusted exchange rates. The variable contains the mean value for net earnings within each country-education-marital status cell. The sample consists of households (single and married couples) headed by a male aged 25–55. Individuals with missing data on education, marital status, or work status were dropped. Earnings were top coded at the 99th percentile within each country. Gross earnings data for the United States from PSID was converted into net earnings using the OECD tax calculator.

Public Pension Net Benefits

Our public pension net earnings replacement rate variable is derived using replacement rates calculated by the OECD. The OECD analysis maps levels of pre-retirement earnings (expressed as a multiple of the earnings of a representative worker) to the replacement rate provided by each country's mandatory public pension system.¹

OECD has been publishing data using their current definition of representative worker earnings since 2000.² We converted cell level mean earnings from ECHP and PSID into 2000-level national currency using OECD wage growth data. This was then converted into a multiple of representative worker earnings using OECD earnings data.

The OECD mapping from earnings to replacement rate was then used to calculate the household replacement rate directly for unmarried men. For married men, additional steps were necessary, as the OECD mapping is given only with regards to individuals. We calculated individual replacement rates for the husband and wife separately, converted them into the implied individual benefits, added them together to get total household benefits, and divided the result by total household earnings to get a household earnings replacement rate. This does not include "spouse" benefits which are common in some countries.

Assumptions Regarding Mortality and Discounting

We compute the present value of pension benefits using the males' survival rates derived from the Human Mortality database for each country and assuming a discount rate of 3%. We discounted back to the average age at interview within each cell. Other present values assumed a discount factor of 3% as well.

¹This analysis is used in the OECD's "Pensions at a Glance" series; we received access to tables relating income to replacement rates with increments of 0.1*representative-worker-income.

²Reflecting the changing balance of economic sectors, the definition of representative worker now includes the full-time service workers in addition to the traditional full-time production workers.

Figure 1. Net Replacement Rates by Country

Notes: OECD net replacement rates as a function of multiple of average earnings as reported in the OECD publication “Pension at a Glance”. The net replacement rate is defined as the ratio of net pension benefits and net earnings.

Figure 2. Country-Level Relationship between Lifetime Replacement Rate and the Ratio of Financial Wealth to Lifetime Earnings

Notes: AU Austria, BE Belgium, DE Germany, DK Denmark, EN England, FR France, GR Greece, IT Italy, NL Netherlands, SP Spain, SE Sweden, US United States. The solid line represents a regression line from regressing the ratio of financial wealth to lifetime earnings (A) on the lifetime replacement rate (REP).

Figure 3. Relationship between Replacement Rate and the Ratio of Financial Wealth to Earnings: After Controlling for Group Differences

Notes: Each dot in the figure represents a cell's partial residuals for the ratio of financial wealth/life time earnings and the net replacement rate as calculated using the formula in the text. The residuals are computed by regressing each variable on a set of indicators for education, household type, and country. The size of the "bubble" around each dot is proportional to the size of the cell (in terms of observations). The solid line denotes the regression line from regressing the financial wealth ratio residual on that for the replacement rate. Its slope is equal to the point estimate reported in column 3 of Table 2 by construction.

Figure 4. Relationship between Pension Wealth and Average Retirement Age: After Controlling for Group Differences

Notes: Each dot in the figure represents a cell's partial residuals for average retirement age and pension wealth. The residuals are computed by regressing each variable on a set of indicators for education, household type, and country and household pre-retirement net earnings. The size of the "bubble" around each dot is proportional to the size of the cell (in terms of observations). The solid line denotes the regression line from regressing the average retirement age residual on that for pension wealth. Its slope is equal to the point estimate reported in column 3 of Table 3 by construction.

Table 1

Descriptive Statistics by Country

Country	Education Level (%)			Married (%)	Median Financial Wealth	Mean replacement rate $\phi(y)/y$	average net household earnings	Average Retirement Age
	Primary	Secondary	Tertiary					
Austria	23.5	53.9	22.6	80.7	10541	91	39889	59.13
Belgium	56.3	21.3	22.4	82.9	20241	66	32021	58.64
Denmark	18.9	52.7	28.4	72.3	43420	90	35727	61.39
England	56.5	32.4	11.1	77.7	21194	57	39478	61.22
France	55.7	27.1	17.2	83.3	24565	65	36571	58.37
Germany	6.6	65.2	28.3	86.9	12963	57	37208	60.46
Greece	72.5	15.5	12.0	85.8	4225	112	21889	60.87
Italy	86.7	10.5	2.7	92.9	8768	79	20365	58.77
Netherlands	56.0	28.2	15.8	90.1	15415	99	30420	59.17
Spain	90.9	4.3	4.7	86.6	5694	82	18609	61.64
Sweden	65.2	20.9	14.0	88.9	29809	71	29875	62.20
United States	27.9	36.1	36.0	81.8	85054	51	59913	61.74

Notes: The sample consists of retired men aged 65–75 in 2004, from Wave 1 of SHARE, Wave 2 of ELISA, and Wave 7 of HRS. Average net household earnings (pre-retirement) was calculated using a pooled sample of working men aged 50–55, using the 1994–1997 waves of the European Community Household Panel (ECHP) and Panel Survey of Income Dynamics (PSID). Sample weights used. Gross income measures in PSID were converted to net income using the OECD tax calculator. Education levels are based on the International Standard Classification of Education (ISCED 1997): Primary = ISCED 0,1; Secondary = ISCED 2,3; Tertiary = ISCED 4,5,6. The average retirement age does not include those not yet retired. Financial wealth includes bank account balances, the value of stocks and bonds, and money saved in individual retirement accounts. All monetary figures are expressed in 2004 dollars, PPP-adjusted.

Table 2**Regression Results – Wealth Ratio**

Model	(1)	(2)	(3)
Replacement Rate	−0.130 [2.79]	−0.099 [2.8]	−0.218 [2.4]
Fixed Effects			
Country	No	No	Yes
Education & Household Size	No	Yes	Yes
N	12	64	64
R-sq	0.181	0.652	0.791

Notes: Results are from regression of the ratio between median financial wealth and mean household earnings on the net replacement rate provided by mandatory public pensions. The first regression uses country-level aggregated measures. The second uses disaggregated data and fixed effects for each education*marital status combination. Education has three categories based on the International Standard Classification of Education (ISCED 1997): Primary (ISCED 0,1), Secondary (ISCED 2,3) and Tertiary (ISCED 4,5,6). Marital status is dichotomous for currently married or not currently married. The last specification includes both group and country fixed effects. Replacement rates are calculated using pension data from OECD and pre-retirement mean household earnings. Pre-retirement mean household earnings is derived from the European Community Household Panel (ECHP) and the Panel Survey of Income Dynamics (PSID). The sample consists of retired men aged 65–75 in 2004, from Wave 1 of SHARE, Wave 2 of ELSA, and Wave 7 of HRS. Absolute t statistics are below the parameter estimates in brackets. Regressions use cell counts (number of observations) as weights. Cells with fewer than 5 observations are discarded (8 cells).

Table 3**Regression Results – Retirement Age**

Model	(1)	(2)	(3)
Pension Wealth (in thousands)	0.003 [1.53]	−0.007 [2.05]	−0.011 [3.24]
Elasticity	0.034	−0.099	−0.146
Fixed effect controls			
Country	No	No	Yes
Education and household type	No	Yes	Yes
N	12	64	64
R-sq	0.067	0.299	0.915

Notes: Results are from regression of the average retirement age on the pension wealth provided by mandatory public pensions. The first regression uses country-level aggregated measures. The second uses disaggregated data and fixed effects for each education*marital status combination. Education has three categories based on the International Standard Classification of Education (ISCED 1997): Primary (ISCED 0,1), Secondary (ISCED 2,3) and Tertiary (ISCED 4,5,6). Marital status is dichotomous for currently married or not currently married. The last specification includes both group and country fixed effects. Pension wealth is calculated using pension data from OECD and pre-retirement mean household earnings. Pre-retirement mean household earnings is derived from the European Community Household Panel (ECHP) and the Panel Survey of Income Dynamics (PSID). The sample consists of retired men aged 65–75 in 2004, from Wave 1 of SHARE, Wave 2 of ELSA, and Wave 7 of HRS. Absolute t statistics are below the parameter estimates in brackets. The elasticities are of expected years of retirement with respect to pension wealth. Regressions use cell counts (number of observations) as weights. Cells with fewer than 5 observations are discarded (8 cells).

Table 4**Regression Results – Robustness Analysis**

Model	Financial Wealth	Avg. Retirement Age
Replacement Rate	−0.386 [1.76]	
Pension Wealth		−0.0124 [3.25]
P(survival 10 yrs)	1.00E-04 [0.73]	−0.015 [0.85]
P(any bequest)	0.015 [1.03]	1.368 [0.67]
Fixed effect controls		
Country	Yes	Yes
Education and household type	Yes	Yes
N	64	64
R-sq	0.733	0.875

Notes: Results are from regression of financial wealth and average retirement age with country level fixed effects (U.S. is the excluded category), and fixed effects for each education*marital status combination as in column 3 of Table 2 and 3. The replacement rate includes private and occupational pension benefits from HRS, ELSA and SHARE. Two variables are added: a) the cell-average subjective survival probability over the next 10 years, b) the cell-fraction of respondents who give a positive probability of giving any bequest. The sample consists of retired men aged 65–75 in 2004, from Wave 1 of SHARE, Wave 2 of ELSA and Wave 7 of HRS. t statistics below parameter estimates in brackets. Regressions use cell counts (# of observations) as weights. Cells with less than 5 observations are discarded (8 cells).