THE GENDER GAP IN PRIVATE PENSIONS

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ABSTRACT

Older women in Britain receive considerably less private pension income than older men, on average. We analyse this differential by examining differences between the sexes both in private pension coverage and in pension income conditional on receipt. Using regression-based decompositions, we show that both gaps are associated mainly with gender differences in returns to personal characteristics rather than with gender differences in personal characteristics per se. In particular, although there are marked differences between elderly men and elderly women in their lifetime employment histories, these differences account for only a small fraction of the overall private pension income gap between the sexes.

NON-TECHNICAL SUMMARY

Differences in income from occupational and personal pensions and annuities ('private pension income'; PPI for short) are one of the most importance sources of *income inequality* among British pensioners. There is also a substantial difference between the *average income* of older men and of older women, much of which has been attributed to differences between the sexes in private pensions. Only one third of female pensioners over 65 received any PPI in 1993–94, compared with two-thirds of male pensioners. Moreover, according to the former Department of Social Security, '[t]he main difference [in the net income of men and women] occurs in occupational pension income. In 1997/8, single men received £44 a week on average from this source, compared with £29 for single women. Average incomes from the other sources were very similar for single men and women.

In this paper we investigate the sources of the differences in PPI between British men and women aged 66 and older, addressing several related questions. To what extent is the gender gap due to women having lower coverage rates than men or due to having lower pension incomes conditional on coverage? To what extent does the gender gap arise because women's characteristics are less well rewarded in pension terms than men's? And to what extent does the gap arise because women have less advantageous personal characteristics than men? In particular, what is the role played by women's lower lifetime labour market participation rates and their interrupted working lives?

Answers to these questions are relevant to the formulation of pension policy, including whether measures should be directed at the labour market or at pension schemes per se. For example, even if personal pension coverage rates among women rise, their subsequent PPI receipts will remain low if also earnings remain relatively low or working lives are relatively short or interrupted. In the past, this was not necessarily seen as a problem, as it was assumed that women would benefit from the pension entitlements of their husbands. This supposition is increasingly less appropriate, given rising divorce rates and growing expectations of financial independence for women regardless of marital status.

We examine the gender gap in private pensions using a regression-based decomposition framework, based on a joint model of the probability of PPI receipt and of PPI amounts conditional on receipt. Using data from the British Household Panel Survey, we fit the model separately for older men and women and use the estimates to decompose the differences between the sexes on both outcomes, combining applications of methods due to Gomulka-Stern and to Blinder-Oaxaca.

We show that there is a substantial gender gap in private pensions in Britain, with two distinct components. First, there is the shortfall for women in the likelihood of receiving any PPI at all (only one third of elderly women receive any PPI, compared with about three-quarters of elderly men) and, second, there is the lower PPI received by female recipients compared to male recipients (about £40 per week on average compared with £80 per week).

Both components of the overall gender gap arise mainly because women's characteristics are less well rewarded than men's, rather than because women have less advantageous personal characteristics than men. In particular, differences in returns account for at least four-fifths of the PPI gap among recipients. Put another way, the role of differences in characteristics is greatest in accounting for the gap in PPI receipt probabilities but, even in this case,

differences in characteristics are responsible for at most one half of this gap. The results mean that women's distinctly lower lifetime labour market participation rates and lower educational qualifications may be less responsible for the gender gap in private pensions that might have been imagined. Their effects on the gap, such as they are, primarily operate through the probability of receipt rather than PPI levels.

Our results about differences in returns to characteristics being particularly responsible for the gender gap in private pensions in Britain may be driven by several related factors. The returns may reflect the nature of jobs that these cohorts of women have taken (often part-time and without opportunities to join occupational pension schemes), or the lower lifetime earnings of women relative to men (generating smaller entitlements when a scheme was available). It may also reflect the preferences of women of this generation, choosing not to contribute to private pensions, in the expectation (rightly or wrongly) that they could rely on their husbands' entitlements. Identifying the roles of each influence is difficult. One way to control for differences in preferences would be to restrict analysis to a sub-sample of 'career' men and women (single never-married without children), but cell sizes were too small for us to follow-up this idea. Our estimates do suggest, however, that there is a negligible private pension gap between never-married career men and women in professional occupations. This underlines how the gender gap would be reduced were similar men and women to have similar jobs and similar earnings. Reducing these gaps is a key issue, but difficult to resolve as long as combining paid work and parenthood (or other caring responsibilities) affects women more than men.

I. Introduction

Differences in income from occupational and personal pensions and annuities ('private pension income'; PPI for short) are one of the most importance sources of income inequality among British pensioners (Titmuss 1955, Atkinson 1973, Johnson and Stears 1995, Department of Social Security 2000, Curry and O'Connell 2003).[1] There is also a substantial difference between the average income of older men and of older women, much of which has been attributed to differences between the sexes in private pensions (Ginn and Arber, 1996). Only one third of female pensioners over 65 received any PPI in 1993–94, compared with two-thirds of male pensioners (Ginn and Arber, 1999). Moreover, according to the former Department of Social Security, '[t]he main difference [in the net income of men and women] occurs in occupational pension income. In 1997/8, single men received £44 a week on average from this source, compared with £29 for single women. Average incomes from the other sources were very similar for single men and women.' (DSS, 2000, p. 30.)

In this paper we investigate the sources of the differences in PPI between British men and women aged 66 and older, addressing several related questions. To what extent is the gender gap due to women having lower coverage rates than men or due to having lower pension incomes conditional on coverage? To what extent does the gender gap arise because women's characteristics are less well rewarded in pension terms than men's? And to what extent does the gap arise because women have less advantageous personal characteristics than men? In particular, what is the role played by women's lower lifetime labour market participation rates and their interrupted working lives?

Answers to these questions are relevant to the formulation of pension policy, including whether measures should be directed at the labour market or at pension schemes per se. (Cf. stakeholder pensions, intended to increase access to private pension schemes for those with moderate earnings.[2]) For example, even if personal pension coverage rates among women rise, their subsequent PPI receipts will remain low if also earnings remain relatively low or working lives are relatively short or interrupted. In the past, this was not necessarily seen as a problem, as it was assumed that women would benefit from the pension entitlements of their

[1] Occupational pensions are by far the most important component of PPI (it was only in 1988 that personal pension schemes were allowed as an alternative to the State Earnings Related Pension Scheme). For descriptions of the current British pension system and future prospects, see for example Emmerson (2002), Curry and O'Connell (2003), and Pensions Commission (2004).

[2] Stakeholder pensions were introduced in 2001 with the aim of extending private pension membership among people on moderate earnings without access to occupational pension schemes. Because they allow workers to stop and restart contributions without penalties, they are intended to help many women to make their own private pension provision.

husbands. This supposition is increasingly less appropriate, given rising divorce rates and growing expectations of financial independence for women regardless of marital status (Ginn 2003). Our results are of relevance to countries outside Britain as well: a sharp divide between affluent pensioners who have PPI (disproportionately men) and poor pensioners without PPI (disproportionately women) arises in most countries in which private pensions play an important role in the pension system (Behrendt, 2000). In the US 1982 Newly Entitled Beneficiary Survey, for example, two-thirds of elderly men received private pensions but only just over one third of elderly women, and women's pensions and total incomes were much lower than men's (Even and Macpherson, 1994).

Our paper is the first we are aware of that examines the gender gap in private pensions using a regression-based decomposition framework based on a joint model of the probability of PPI receipt and of PPI amounts conditional on receipt. We fit the model separately for older men and women and use the estimates to decompose the differences between the sexes on both outcomes, combining applications of methods due to Gomulka and Stern (1990) and to Blinder (1973) and Oaxaca (1973). Ginn and Arber (1996) analysed the relationship between work history and older women's non-state pensions, but did not decompose differences between the sexes.[3] Potential reasons for the gender gap in PPI are usefully reviewed by Ginn (2003, chapter 2), but without a quantitative assessment of the roles played by different factors. Our research is closest to that of Even and Macpherson (1994) who examined private pensions in the USA. Compared to them, we consider Britain and employ more appropriate estimation and decomposition methods.[4]

There are many potential reasons why elderly women today are less likely than elderly men to receive a private pension. The norm of a traditional family, with a non-working wife caring for children and a working husband generating pension entitlements that would cover them both, was much more prevalent. Related to this, women of this generation spent a large proportion of their employment career working part-time, and part-time workers have only recently been given the same rights as full-time workers to join their employer's pension

[3] Stewart (2003) examined pensioners' total incomes (rather than non-state pensions) and related them to working-life earnings levels and to wealth levels. Bardasi and Jenkins (2002) related differences in low-income rates between men and women aged 60+ to differences in lifetime employment histories. Similar topics were also addressed by Rake et al (2000, chapter 6), using simulated life histories for women with different characteristics.

[4] Even and Macpherson modelled the probability of receipt using a probit model. However, rather than decomposing differences in fitted probabilities, they decomposed differences in the linear index function that is the argument of the probit function (1994, fn. 12). Moreover, when estimating their pension amount equation, they did not allow for potential selection biases.

scheme if there is one.[5] In any case, most part-time workers are found in low-skilled low-paid occupations without the opportunity to join a scheme. The low portability of occupational pensions between jobs has made it less worthwhile for women, with breaks in their employment careers, to join a scheme. But even if women join a scheme, their lower earnings than men on average – reflecting both lower hourly pay rates and lower work hours – translate into lower pension income as well.[6] Flatter age-earnings profiles also mean that women do worse from occupational schemes that calculate entitlements as an average of earnings in the last few work years.

Our data set, derived from the British Household Panel Survey (BHPS), contains measures of many of the factors mentioned in the preceding paragraph, and which we use as explanatory variables in our regressions. These include the proportion of the total work life spent working in specific occupations, out of work, or unemployed, and we also distinguish between spent in full-time work, part-time work, and self-employment. We also have data about marital and fertility histories. We do not have direct measures of lifetime pension contribution histories, however, nor of lifetime earnings (few surveys have these). The positive impact on PPI income of, say, working in a well-paid job with a good occupational scheme, has to be inferred from information about the time spent in a 'good' occupation. As will be seen below, our results are consistent with prior expectations, but the reduced form nature of the modelling implies that conclusions can be drawn only with some caution, and we also report the results of some sensitivity analysis.

The econometric model is set out in Section II, together with the decomposition framework. Section III describes the BHPS sample and the measures of PPI outcomes and explanatory variables. Model estimates are presented in Section IV. The roles of differences in characteristics and differences in returns to characteristics in accounting for the gender gap in PPI, in a disaggregated manner (Section V) and using summary Gomulka-Stern and Blinder-Oaxaca decompositions (Section VI). Section VII provides concluding comments.

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[5] The European Court of Justice ruled in September 1994 that the exclusion of part-time workers from occupational pension schemes may represent indirect discrimination against women (DSS, 2000).

[6] On gender wage differentials in Britain, see for example, Harkness (1996); Makepeace et al. (1999), and Dolton et al. (2002).

II. Model and methods

In order to *receive* any PPI in old age, one must have made contributions during one's working life. This contribution history is not observed directly in our data set, but we can model it as a latent propensity. For each individual $i = 1, \ldots, N$, we suppose that

$$C_i^* = Z_i\gamma + \varepsilon_i,\ \varepsilon_i \sim N(0, 1) \tag{1}$$

where contribution propensity, C_i^*, is a function of observed characteristics such as age, education, labour market attachment, occupation, and household characteristics (vector Z_i), plus normally-distributed unobserved differences (ε_i). We observe PPI receipt if contribution propensities are sufficiently high, i.e. $C_i = 1$ if $C_i^* > 0$, and $C_i = 0$ otherwise.

The potential *amount* of PPI received in old age, Y_i^*, is also a function of observed and normally-distributed unobserved characteristics (X_i, u_i, respectively) which determine lifetime earnings and hence contributions paid:

$$Y_i^* = X_i\beta + u_i,\ u_i \sim N(0, \sigma^2). \tag{2}$$

We assume that the unobserved components of (1) and (2) follow a bivariate normal distribution with covariance ρ. A positive covariance arises if individuals who are more likely to contribute and receive a pension are also more likely to receive a higher amount than the average PPI recipient, other things being equal. The amount of PPI generated by the contribution history is observed only if contribution propensities were sufficiently high ($C_i = 1$). Thus the observation rule is:

$$\begin{gathered} Y_i = \log(PPI_i) = Y_i^* \text{ and } C_i = 1 \text{ if } C_i^* > 0 \\ Y_i \text{ unobserved and } C_i = 1 \text{ if } C_i^* \leq 0 \end{gathered} \tag{3}$$

We model the logarithm of pension income because the distribution of PPI is highly skewed with a long right tail and taking logarithms makes the data approximately normally distributed and the Blinder-Oaxaca decomposition methods used below are based on regressions that require that the dependent variable be expressed in logarithms.

How can the model be used to account for differences between the sexes in PPI receipt probabilities and levels? The predicted probability of PPI receipt is given by $\Phi(Z_i\gamma)$, where $\Phi(.)$ is the normal distribution function, and the expected value of log(PPI) for individuals with positive PPI is $X_i\beta$.[7] The expressions show that differences between the sexes arise because of differences in observed characteristics (differences in Z_i and X_i), and because of

[7] One could also calculate expected PPI without conditioning on receipt, in which case $E(Y \mid X_i, Z_i) = [X_i\beta + \rho\sigma\phi(Z_i\gamma)/\Phi(Z_i\gamma)]\Phi(Z_i\gamma)$. We focus on the conditional calculations in this paper, following the precedent of the

differences in the impact that these characteristics have on PPI receipt probabilities (differences in γ) and on PPI levels (differences in β). We provide information about each of these elements in Section V.

The relative importance of differences in characteristics and differences in 'returns' to characteristics is assessed in Section VI using Gomulka-Stern and Blinder-Oaxaca decompositions. The former decomposes the difference in average probabilities of PPI receipt into terms reflecting differences in average characteristics holding coefficients ('returns') constant and differences in coefficients holding characteristics constant:

$$
\begin{aligned}
\hat{\Pr}(C_i = 1 \mid j=1) - \hat{\Pr}(C_i = 1 \mid j=0) &= \sum_{i\in 1} \frac{\Phi(\hat{\gamma}^{1\prime} Z_i)}{N_1} - \sum_{i\in 0} \frac{\Phi(\hat{\gamma}^{0\prime} Z_i)}{N_0} \\
&= \left(\sum_{i\in 1} \frac{\Phi(\hat{\gamma}^{0\prime} Z_i)}{N_1} - \sum_{i\in 0} \frac{\Phi(\hat{\gamma}^{0\prime} Z_i)}{N_0} \right) \\
&\quad + \left(\sum_{i\in 1} \frac{\Phi(\hat{\gamma}^{1\prime} Z_i)}{N_1} - \sum_{i\in 1} \frac{\Phi(\hat{\gamma}^{0\prime} Z_i)}{N_1} \right)
\end{aligned}
\tag{4}
$$

where $\hat{\Pr}(C_i = 1 \mid j)$ indicates the average predicted probability of receiving a private pension for group j = 0 (women) and 1 (men), and N_j is the number of individuals in group j. The Blinder-Oaxaca decomposition of the gender gap in PPI is given by:

$$
\overline{\log P^1} - \overline{\log P^0} = \sum_i \hat{\beta}^0 \left(\overline{X}^0 - \overline{X}^1 \right) + \left[\left(\hat{\alpha}^0 - \hat{\alpha}^1 \right) + \sum_i \overline{X}^1 \left(\hat{\beta}^0 - \hat{\beta}^1 \right) \right] . \tag{5}
$$

In (4) and (5), differences in characteristics are weighted by women's coefficients and differences in coefficients are weighted by men's characteristics. Equally, one could weight by men's coefficients and women's characteristics (the choice is essentially arbitrary). Calculations using both sets of assumptions yield bounds on the decompositions.

III. Data

We use data from the British Household Panel Survey (BHPS), combining panel data from waves 1–10 (survey years 1991–2000) and the retrospective lifetime employment, marital and fertility histories.[8] We analyse men and women aged 66+ at the date of interview. The age-based sample selection rule was used to avoid endogeneity problems arising because

gender wage gap literature. Also, the principal policy relevance of our results concerns the correlates of the PPI receipt probabilities, and PPI levels among those who have pensions.

[8] For detailed information about the BHPS, see http://www.iser.essex.ac.uk/bhps/index.php, and about the retrospective employment and job history data, see Halpin (1997). We combined these histories, and the retrospective partnership and fertility histories, with the corresponding histories built up within the panel itself.

retirement decisions are likely to be based upon the expectation of the amount of pension, and the vast majority of individuals have retired by age 65 (the state retirement pension age for men). There were 1119 men and 1499 women aged 66+, contributing 6109 and 8709 person-wave observations. Analysis was based on the sub-sample for whom complete retrospective history data were available, i.e. 891 men and 1229 women.[9]

PPI is the sum of any occupational or personal pension or annuity reported by the individual as being received at the time of the interview, converted to a weekly amount pro rata and expressed in January 2001 prices.[10] Only pensions received by the individual in his or her own right were included (survivor pensions were excluded). For the regression analysis, the dependent variables were based on PPI averaged over all panel years that a positive amount was observed. Longitudinal averaging was used to reduce potential measurement error. (Non-averaged estimates provided similar results.)

Average PPI among all men in the sample was £59 per week, almost six times larger than the average among all women, £10 per week. Among those receiving some PPI, the differential was still large: average PPI for men was £82 per week, which is double the average for women, £41 per week. The differences between the conditional and unconditional averages draw attention to the large differences between the sexes in the probability of PPI receipt. Although 77 per cent of men received some PPI, only 34 per cent of women did.

The explanatory variables in the log(PPI) equation (X_i) included birth cohort and educational qualifications and variables specifically intended to measure differences in contribution records (work history variables described shortly). The explanatory variables in the contribution propensity equation (Z_i) included all the regressors in the log(PPI) equation, plus variables thought to affect contribution propensities but not PPI amounts (discussed in Section IV).

Differences in lifetime earnings and contribution records were summarised using a set of variables describing labour market attachment and occupation over the working life (more specifically, the fraction of time spent in each category between the ages of 20 and 60). The variables recorded the fraction of the working life that had been spent working in full-time work, part-time work, self-employment, unemployment or inactivity. Nine occupational

[9] Selecting only those with complete histories may arguably introduce a selection bias. To check this, we compared the age and educational qualifications of the selected sample with those of all men and women aged 66+. The two samples had very similar characteristics, whether all individuals were considered or only those with PPI. There was no difference in the distribution across birth cohorts. However, the selected sample (men and women) had slightly lower educational qualifications on average.

groups could be distinguished without cell sizes becoming too small: managerial, professional, associate professional and technical; clerical and secretarial, craft and related, personal and protective services, sales; plant and machine operatives, unknown, and 'other' occupations (mostly unskilled).

We would expect that the longer the time spent in the labour market, and the longer the time spent in higher-skilled occupations, the greater the lifetime earnings, and hence a greater likelihood of contributing to a private pension contribution, and higher PPI in old age conditional on having contributed. The variables summarising the fraction of the work life spent in different occupations do not allow us to distinguish between the separate effects on PPI entitlements of belonging to a relatively generous pension scheme, spending a longer time contributing to any particular scheme, or having higher earnings. To help distinguish between the latter two effects, we constructed an additional variable to summarise occupational wage differentials for sample members during their 40s. We pooled five cross-sections of the Family and Expenditure Survey (1968 to 1972, the period that maximizes the percentage of our sample in their 40s) and computed the average wage, by occupation and sex, for men and women aged 40–49, separately for full-time, part-time and self-employed workers. Each BHPS sample member was then assigned the occupational wage corresponding to his or her characteristics when the individual was 45 (or the closest age to when he or she was last in work). When included as an explanatory variable in our regressions, this variable did not have a statistically significant association with the probability of PPI receipt or PPI levels. This could mean that the accumulation of pension contributions over time is the main source of the private pension differential (and that the distinction of the time spent in different statuses – full-time, part-time and self-employment – is able to fully capture the differential pace at which contributions are accumulated), but of course it could also be that the imputed occupational wage variable was too crude.[11] The occupational wage variable was not included in the specifications of the regression models that are reported here.

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[10] Arguably income from the State Earnings Related Pension scheme should also be included in a measure of PPI as SERPS is a potential substitute for a private pension. In the BHPS, however, SERPS income cannot be distinguished separately from basic state pension income.

[11] Stewart (2003) generated lifetime earnings variables for his BHPS sample using the large samples available from the New Earnings Survey. This option was not available to us (Stewart used a special proprietary version of the NES).

IV. Model estimates

The model described in Section III is a Tobit II model, otherwise known as a sample selection model (Heckman, 1979), which we estimated by maximum likelihood separately for men and women.

To identify the model parameters using exclusion restrictions, we require instruments for the PPI contribution propensity equation that do not appear in the log(PPI) equation. For men, the instrument was a binary variable indicating whether there had been a switch between employment and self-employment (or vice versa) at any time during the working life (ages 20–60). The argument is that a change in the type of contract can entail loss of occupational pension rights, because of the change in 'regime' (self-employed people do not have occupational pensions, and are not obliged to contribute to a personal pension), while at the same time not affect the amount of pension somebody is receiving if the individual is able and willing to continue to contribute to a pension scheme. This variable was always significant in the contribution propensity equation and never significant in the pension equation, whether the two equations were estimated separately or jointly.

For women, we used instruments summarising the fraction of time between ages 0–60 spent married, single, divorced, separated or widowed, and whether children aged less than five were present in each case. The time spent being mothers of pre-school children was expected to have a negative impact on their ability and opportunity to contribute toward an occupational or personal pension. In practice, we found no statistically significant association and, at the same time, the estimate of ρ did not differ significantly from zero. We experimented with a number of alternative instruments including the total number of children, the age of the mother at first birth, her age at first marriage, and a binary variable indicating whether the woman had two children or more, but with the same results. We therefore rely on non-linearities in functional form to identify the model for women.[12]

Estimates of the model parameters are shown in Table 1. As expected, work history variables had strong statistical associations with the probability of PPI receipt for both men and women. For both sexes, the probability was substantially larger, the more of the work life that was spent in professional, technical, clerical, managerial, and personal and protective service occupations. Spending more time spent in the remaining occupations also had a positive, although smaller, effect on the probability of PPI receipt. Put another way, the longer the time spent in economic inactivity, the lower the probability of receiving a private pension. For women, the longer the time spent in part-time work, the lower the probability of PPI receipt and, for both sexes, the more time spent in self-employment, the lower the probability. There were no statistical significant associations between women's PPI receipt probabilities and the life time marital status variables. It appears that the effects of marriage and children worked entirely through their impact on work histories.

< Table 1 near here >

The estimates of the log(PPI) equation are also shown in Table 1. Compared to the contribution propensity equations, fewer variables had statistically significant associations with PPI levels, for both men and women. For both sexes, the more time that was spent in higher skilled occupations (professional, technical, clerical and personal and protective service occupations) rather than out of the labour force, the higher was PPI. The coefficients for other occupations were never statistically significant. For women, the more of the working life spent working part-time, the lower was PPI. The penalty is substantial, about 2 per cent for each extra year spent working part-time.

Having a degree is strongly associated with larger PPI levels (so too is having A-levels, but for women only). Average PPI for university graduates of either sex is about twice as high as the average for individuals with no formal qualifications. The results are consistent with the idea that graduates receive higher lifetime earnings which translates into higher PPI entitlements. (If earnings differences were captured by educational differences, this would also explain why our imputed occupational wage differential variable was not statistically significant.)

For women, the longer the time that was spent in marriage, separated and widowhood before age 60, the lower the PPI. This result careful interpretation because it refers to the women' personal income, and may have been offset by transfers of income from their (current or former) husbands. (The spouses may have decided that the husband would specialise in paid work, and wife would specialise in domestic production.) We have little information about these transfers. Regardless of this caveat, the apparent 'partnership penalty' to PPI is relevant to the financial independence for women, an issue of policy interest.

For both sexes, there is little evidence of a correlation between unobserved characteristics that increase both the probability of receiving a private pension and the pension

[12] We also estimated the model for women imposing the restriction $\rho = 0$. The results were very similar to those reported.

level (the estimate of ρ does not differ statistically from zero). In addition, birth cohort has little significant association with PPI outcomes.

V. The impact of differences in characteristics versus differences in coefficients

In order to provide the background to the decomposition summaries, we first look in detail at the impacts on outcomes of differences in characteristics and differences in coefficients, in turn.

The impact of differences in characteristics can be assessed by an examination of the average values of the explanatory variables used in the regressions for men and women. The averages are shown in Table 2, for all men and women, and also separately for the subsamples who received PPI. Several patterns emerge. First, PPI recipients have better educational qualifications, spent more years in full-time employment and in highly skilled occupations, and were out of the labour market for a shorter period of time than non-recipients. Second, the average female PPI recipient is very different from the average woman. For example, women with PPI spent 13 years on average out of the labour market (as compared with almost 18 for the average woman), they were in full-time employment for 20 years (rather than 14 years). Also, women with PPI spent more than 12 years working in the four highest-status occupational categories (managers, professionals, associate professional and technical, clerical and secretarial), about twice as long as for the average woman. These contrasts between recipients and others are less pronounced for men.

< Table 2 near here >

Third, although differences in characteristics between the subsamples of men and women with PPI are substantially smaller than those for all men and women, they still exist and are large in some cases. Men with PPI were out of the labour market for fewer than 3 years out of 40, as compared with 13 years for female recipients. Also, men spent more than 34 years working full-time, compared to 20 years on average for female recipients.

Finally, although the marital and fertility characteristics of men with PPI differ little from those for all men, women with PPI differ somewhat from all women. For example, women with PPI spent a larger proportion of their life never married and a smaller proportion married, with and without pre-school children. In part this is because women who receive private pensions are younger (and therefore less likely to have married young, as was more common for older cohorts), but it is also because family commitments reduce female labour market participation, and this reduces women's likelihood of contribution to a private pension.

We explore the implications of differences in coefficients on differential receipt and differential amount received by looking at predicted PPI receipt probabilities and log(PPI) for a set of (hypothetical) individuals with specified characteristics. The predictions are shown in Table 3. The sets of characteristics were chosen to reflect a range from combinations more typical of men (rows nearer the top of Table 3) to those more typical of women (rows nearer the bottom).

< Table 3 near here >

Consider predicted PPI receipt probabilities first (columns 1 and 2). The reference person has A-levels and worked 37 years between ages 20–60. Even in higher-status occupations (rows 1–3), the predicted receipt probability is at least 20 percentage points lower for women than for men. In other unskilled occupations the gap is even larger: some 40 percentage points (row 4). Of course, 37 years in work is unusual for women. If one compares an individual working in a clerical occupation for 37 years with one working in the same occupation for 20 years (rows 3 and 8) the PPI receipt probability falls for both men and women (especially the latter), and the gender gap increases from 20 to 35 percentage points. The gap increases further if ten of the work years were spent in part-time rather than full-time work (row 9).

It is only if the hypothetical individual has a university degree and worked in a top occupation that the gender gap in the probability of PPI receipt all but disappears, the probability of receipt being about 100% for both men and women (rows 5 and 6). The prevalence of such career women is low, however. For the types of women more commonly found, the predicted PPI receipt probability rarely exceeds 40 per cent, i.e. substantially lower than for men.

The impact of differences in the coefficients on the expected log(PPI) is summarised in columns (3) and (4) of Table 3. Expected log(PPI) is always smaller for women that for men with otherwise-equal characteristics, with the exception of the professional career woman in rows 2 and 6. As the number of years in work is reduced and includes a greater fraction of part-time work, the gender gap actually decreases slightly, from 0.62 (37 years full-time work in a clerical occupation; row 3) to 0.49 (20 years full-time work; row 8) to 0.42 (20 years work, half of which was part-time; row 9). This pattern contrasts with that for predicted PPI receipt probabilities, and suggests that differences in returns to characteristics play a smaller role in accounting for the gender gap in log(PPI) than they do in accounting for the gender gap in PPI receipt probabilities. We can check this more systematically using the decompositions, to which we now turn.

VI. Accounting for the gender gap: decomposition estimates

We now use the Gomulka-Stern and Blinder-Oaxaca decomposition methods to assess the relative contributions of differences in characteristics and differences in returns to the gender gaps in predicted probabilities of PPI receipt (Table 4, top panel) and in average PPI levels (Table 4, bottom panel). The estimates are based on the formulae given in (4) and (5), with standard errors computed using the delta method.[13] In each case, we used the alternative weighting assumptions described earlier.

The gender gap in predicted probabilities of PPI receipt is 43 percentage points (0.77–0.34, i.e. equal to the sample proportions). Regardless of the weighting scheme used, it is differences in coefficients (returns to characteristics) rather differences in characteristics that account for the majority of the gender gap in the probability of receiving PPI: 53% when women are assigned men's coefficients (decomposition 2), and 77% when women's coefficients are assigned to men (decomposition 1), and precisely estimated in both cases. Certainly, women have fewer educational qualifications and less lifetime labour market attachment than men (Table 2) but, even adjusting for differences in characteristics such as these, the differences between the sexes in the likelihood of PPI receipt is substantial. Even if men and women had had exactly the same characteristics, the gap in probabilities would have been between 23 and 29 percentage points.

< Table 4 near here >

Consider now the decompositions of the gender gap in log(PPI) shown in the bottom panel of Table 4. The observed gap is 0.776 log points (PPI for the average man is about twice that of the average woman). The striking result is that almost all this gap is explained by differences in coefficients: according to decomposition (1) the fraction is 95%; according to decomposition (2) the fraction is 83%. Thus, even if women had had the same characteristics as men, the gender gap in PPI would decline by very little.

Even though the fraction of the gender gap in PPI accounted for by observable characteristics is small, there is interest nonetheless in the role played by differences in lifetime work histories. After all, these are the differences between men and women that are

[13] Cf. Gomulka and Stern (1990, Appendix 2) and Oaxaca and Ramson (1998).

perhaps the most commonly remarked upon in the pensions context. We therefore decomposed the gap associated with differences in characteristics further, into four components corresponding to differences in birth cohort and education, marital and fertility, work history variables, and other variables (the last of which is not shown in Table 4).[14] The point estimates from both decompositions (1) and (2) suggest that differences in work histories account for virtually all of the gender gap attributable to differences in characteristics but, at the same time, the estimates are also imprecisely estimated and do not differ statistically from zero. It remains the case, however, that even if differences in characteristics essentially represent differences in work histories, their contribution to the gender gap in PPI is modest in comparison with the contribution of differences in returns to characteristics.

VII. Summary and conclusions

There is a substantial gender gap in private pensions in Britain that has two distinct components. First, there is the shortfall for women in the likelihood of receiving any PPI at all (only one third of elderly women receive any PPI, compared with about three-quarters of elderly men) and, second, there is the lower PPI received by female recipients compared to male recipients (about £40 per week on average compared with £80 per week).

Both components of the overall gender gap arise mainly because women's characteristics are less well rewarded than men's, rather than because women have less advantageous personal characteristics than men. In particular, differences in returns account for at least four-fifths of the PPI gap among recipients. Put another way, the role of differences in characteristics is greatest in accounting for the gap in PPI receipt probabilities but, even in this case, differences in characteristics are responsible for at most one half of this gap. The results mean that women's distinctly lower lifetime labour market participation rates and lower educational qualifications may be less responsible for the gender gap in private pensions that might have been imagined. Their effects on the gap, such as they are, primarily operate through the probability of receipt rather than PPI levels.

These findings differ from those for the US by Even and Macpherson (1994). They estimated that most of the gender gap in private pension receipt probabilities (between 69% and 81%) was contributed by differences in characteristics, rather than differences in returns

[14] In principle we could have also further decomposed the fraction of the gap in probabilities associated with differences in characteristics using a method proposed by Gomulka and Stern (1990). However, the method isinfeasible when there are many continuous explanatory variables as there is in our case.

as we found. The contribution of differences in characteristics to the gap in PPI levels was also large (70%) if a final-earnings variable was included in the relevant regressions. There are several potential reasons why these results differ from ours, including different sample selection criteria, different regression specifications and decomposition methods, different time periods (their survey referred to 1982) and of course the different labour markets and pension systems in the USA compared to Britain. A proper cross-national analysis would require data that were more comparable.

Our results about differences in returns to characteristics being particularly responsible for the gender gap in private pensions in Britain may be driven by several related factors. The returns may reflect the nature of jobs that these cohorts of women have taken (often part-time and without opportunities to join occupational pension schemes), or the lower lifetime earnings of women relative to men (generating smaller entitlements when a scheme was available). It may also reflect the preferences of women of this generation, choosing not to contribute to private pensions, in the expectation (rightly or wrongly) that they could rely on their husbands' entitlements. Identifying the roles of each influence is difficult. One way to control for differences in preferences would be to restrict analysis to a sub-sample of 'career' men and women (single never-married without children), but cell sizes were too small for us to follow-up this idea. Our estimates do suggest, however, that there is a negligible private pension gap between never-married career men and women in professional occupations (Table 3). This underlines how the gender gap would be reduced were similar men and women to have similar jobs and similar earnings.[15] Reducing these gaps is a key issue, but difficult to resolve as long as combining paid work and parenthood (or other caring responsibilities) affects women more than men.

Given the key contribution of private pensions in creating a sharp divide between men's and women's individual incomes, the growing emphasis on the financial independence of women, and the difficulties of closing gender gaps in the labour market, there is an important role for the state to play in pension provision. Compared to the private pension system, the public system is better able to be redistributive towards women – to take account of the pension contributions foregone by their greater provision of unpaid family care.

[15] The Pensions Commission has noted that rates of membership of occupation pension schemes are now higher among women in full-time employment than among men, and membership among part-time female employees has increased substantially over the last fifteen years, stating: 'Taking full-time and part-time employees together, female employees are still less likely to be in a pension scheme, but the gap is closing' (2004, p. 276).

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TABLE 1
Regression estimates, by sex

	Men				Women			
	Pr(PPI>0)		Log(PPI)		Pr(PPI>0)		Log(PPI)	
Covariate	Coefficient	SE	Coefficient	SE	Coefficient	SE	Coefficient	SE
Born before 1919	–0.211	(0.189)	–0.344 **	(0.140)	–0.126	(0.136)	–0.023	(0.185)
Born 1919–1923	–0.110	(0.195)	–0.104	(0.131)	–0.060	(0.138)	–0.133	(0.206)
Born 1924–1928	–0.205	(0.182)	–0.021	(0.129)	0.016	(0.130)	0.094	(0.159)
Vocational education	0.123	(0.157)	–0.131	(0.145)	–0.036	(0.153)	0.155	(0.172)
O-level	0.551 ***	(0.195)	0.196	(0.150)	0.230 *	(0.136)	0.133	(0.147)
A-level	0.464 *	(0.278)	0.158	(0.206)	–0.034	(0.273)	0.650 **	(0.263)
Other higher education	0.384 **	(0.165)	0.732 ***	(0.110)	0.408 ***	(0.158)	0.333	(0.212)
Degree	6.766 ***	(0.303)	1.029 ***	(0.198)	1.633 ***	(0.586)	0.836 ***	(0.277)
Proportion of time between the ages 20–60:								
Self-employed	–0.938 ***	(0.269)	–0.215	(0.266)	–1.192 ***	(0.414)	–0.339	(0.548)
Part-time work	–0.527	(1.018)	–0.638	(0.480)	–1.052 ***	(0.210)	–0.769 ***	(0.246)
Unemployment	–2.357	(1.658)	–1.612	(1.810)	–1.264	(1.185)	–1.160	(1.926)
Unknown occupation	1.923 ***	(0.444)	0.455	(0.465)	1.414 ***	(0.209)	0.576 *	(0.311)
Managerial occupation	1.862 ***	(0.510)	0.787	(0.523)	2.021 ***	(0.500)	0.465	(0.599)
Professional occupation	4.745 ***	(1.400)	0.980 *	(0.518)	1.813 ***	(0.532)	1.835 ***	(0.434)
Technical occupation	3.142 ***	(0.837)	1.043 **	(0.494)	1.796 ***	(0.381)	0.976 ***	(0.347)
Clerical occupation	2.288 ***	(0.508)	1.220 **	(0.491)	1.959 ***	(0.223)	0.919 ***	(0.296)
Craft occupation	1.563 ***	(0.441)	0.178	(0.479)	0.321	(0.302)	0.597 *	(0.333)
Personal and protective services occupation	2.259 ***	(0.550)	1.122 **	(0.502)	2.052 ***	(0.326)	0.841 **	(0.340)
Sales occupation	1.842 ***	(0.576)	–0.264	(0.544)	0.823 ***	(0.310)	0.542	(0.448)
Plant/machine operator occupation	1.760 ***	(0.459)	0.110	(0.481)	0.889 ***	(0.287)	–0.250	(0.404)
Other unskilled occupation	1.506 ***	(0.466)	0.213	(0.477)	1.016 ***	(0.276)	–0.655	(0.681)
Both employee and self-employed	–0.444 ***	(0.152)						

Proportion of time between ages 0–60:								
Married with no child aged<5					–0.052	(0.270)	–1.112***	(0.295)
Married with child(ren) aged<5					0.487	(0.514)	–0.476	(0.676)
Separated with no child aged<5					–0.508	(1.310)	–3.996***	(1.532)
Divorced with no child aged<5					0.681	(0.759)	0.006	(0.455)
Widowed with no child aged<5					0.190	(0.538)	–1.222*	(0.703)
Constant	–0.360	(0.436)	3.357***	(0.466)	–1.226***	(0.216)	3.022***	(0.329)
ρ	0.027	(0.089)			0.098	(0.069)		
σ	1.012***	(0.037)			0.999***	(0.057)		
Log-likelihood		–1379.41				–1253.15		
Number of observations	891		686		1229		416	

Notes. ***: $p \leq 0.01$. **: $0.01 < p \leq 0.05$. *: $0.05 < p \leq 0.10$. Reference categories for categorical variables: born after 1928, no educational qualifications, proportion of time between ages 20–60 spent in full-time work, proportion of time between ages 20–60 out of the labour market, no switch between self-employment and employment, proportion of time between ages 0–60 spent never-married with no child aged < 5. The regressions also included wave dummies.

TABLE 2

Mean values of characteristics

	Men		Women	
	All	With private pension	All	With private pension
Born before 1919	0.28	0.27	0.35	0.27
Born 1919–1923	0.24	0.25	0.23	0.21
Born 1924–1928	0.23	0.23	0.23	0.28
Born after 1928	0.24	0.26	0.20	0.24
No educational qualifications	0.54	0.49	0.68	0.55
Vocational education	0.13	0.12	0.08	0.10
O-level	0.11	0.12	0.10	0.14
A-level	0.04	0.05	0.02	0.02
Other higher education	0.16	0.19	0.10	0.17
Degree	0.02	0.03	0.01	0.03
Proportion of time between ages 20–60:				
Self-employed	0.09	0.06	0.03	0.03
Part-time worker	0.01	0.01	0.18	0.16
Full-time worker	0.82	0.86	0.35	0.48
Unemployed	0.01	0.01	0.01	0.00
Out of labour market	0.07	0.07	0.44	0.33
Unknown occupation	0.15	0.15	0.11	0.12
Managerial occupation	0.07	0.07	0.02	0.03
Professional occupation	0.03	0.05	0.01	0.03
Technical occupation	0.05	0.06	0.03	0.06
Clerical occupation	0.09	0.10	0.12	0.19
Craft occupation	0.24	0.22	0.05	0.03
Personal and protective services occupation	0.04	0.04	0.05	0.06
Sales occupation	0.03	0.03	0.05	0.04
Plant/machine operator occupation	0.14	0.14	0.05	0.04
Other unskilled occupation	0.08	0.07	0.07	0.06
Both employee and self-employed	0.23	0.18	0.10	0.09
Proportion of time between ages 0–60:				
Married with no child aged <5	0.37	0.39	0.38	0.36
Married with child(ren) aged < 5	0.12	0.12	0.12	0.11
Separated with no child aged < 5	0.00	0.00	0.00	0.00
Divorced with no child aged < 5	0.01	0.01	0.01	0.01
Widowed with no child aged < 5	0.01	0.01	0.03	0.03
Never married with no child aged < 5	0.49	0.48	0.45	0.48
Number of observations	891	686	1229	416

TABLE 3

Predicted probabilities of PPI receipt and expected log(PPI), by sex

	Type of individual	Pr(PPI>0)		Log(PPI)	
		Men	Women	Men	Women
		(1)	(2)	(1)	(2)
	Has A-level, worked 37 years and:				
1	Managerial occupation	0.92	0.77	4.12	3.47
2	Professional occupation	1.00	0.71	4.29	4.74
3	Clerical occupation	0.96	0.75	4.51	3.89
4	Unskilled other occupation	0.85	0.43	3.58	2.44
5	As 1, except with degree	1.00	0.99	4.98	3.66
6	As 2, except with degree and never married	1.00	0.99	5.16	5.52
7	As 3, except always worked part-time	0.90	0.39	3.92	3.18
8	As 3, except 20 years full-time work	0.79	0.44	3.99	3.50
9	As 3, except 10 years full-time and 10 years part-time work	0.75	0.34	3.84	3.31
10	A-level, never in work	0.36	0.13	3.39	3.04

Notes. Individuals are assumed to have been born in the period 1919–1923. The proportion of time between ages 20–60 spent unemployed, and the wave dummies are set at the mean values for men; so too are the marital/fertility variables, unless stated otherwise.

TABLE 4

Decomposition of the gender gap in private pensions

Difference	Gap	% of gap	Gap	% of gap
	(1)		(2)	
Pr(PPI>0)				
Total gap	0.43	100	0.43	100
	(0.02)		(0.02)	
Coefficients	0.23	53	0.29	77
	(0.08)		(0.03)	
Characteristics	0.20	47	0.14	33
	(0.08)		(0.03)	
Log(PPI)				
Total gap	0.776	100	0.776	100
	(0.097)		(0.097)	
Coefficients	0.740	95	0.643	83
	(0.138)		(0.138)	
Characteristics	0.036	5	0.133	17
	(0.171)		(0.108)	
Characteristics:*				
Birth cohort and education	0.014	39	0.017	13
	(0.008)		(0.013)	
Marital status and fertility	0.000	0	0.008	6
	(0.000)		(0.015)	
Work history	0.062	172	0.136	102
	(0.136)		(0.099)	

Notes. Decompositions based on equations (4) and (5). Decomposition (1): differences in characteristics weighted by men's coefficients; differences in coefficients weighted by women's characteristics. Decomposition (2): vice versa. Standard errors in parenthesis.
*: Decomposition of the differential associated with differences in characteristics. The sum of the contributions shown does not equal the total gap associated with differences in characteristics because the contribution of the wave dummy variables is not shown.