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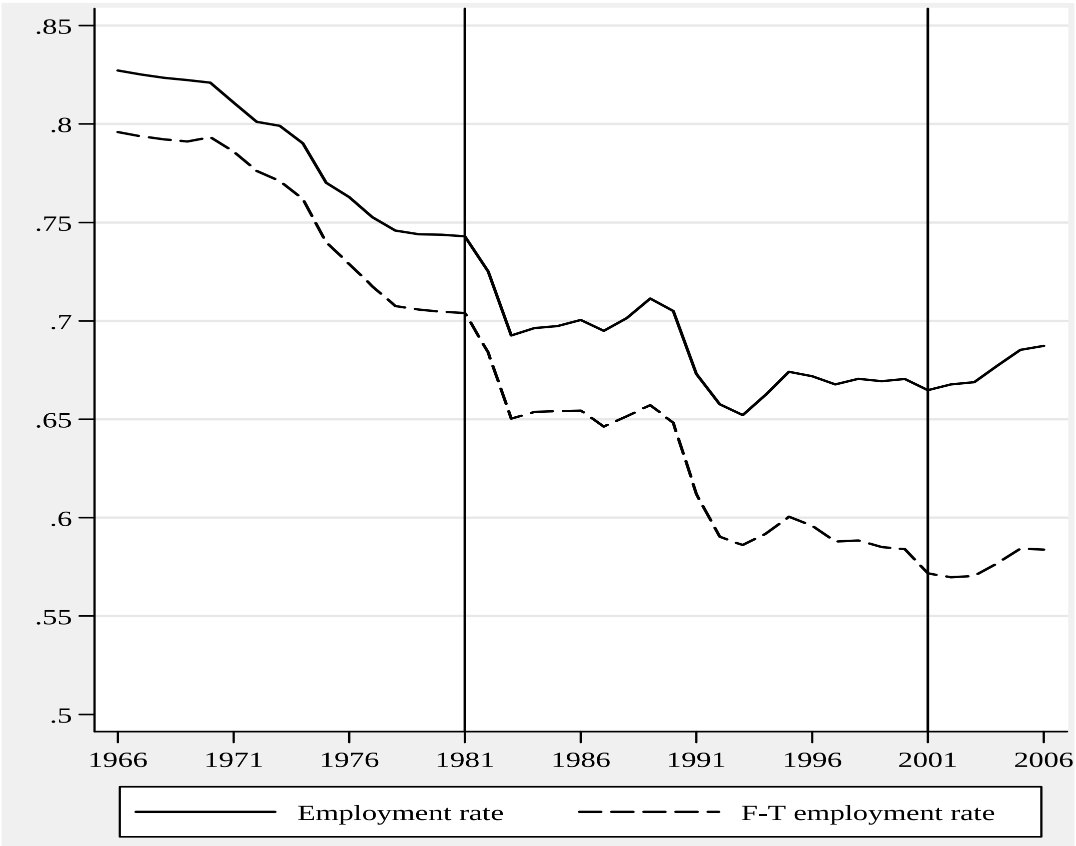
Abstract

Using Australian data spanning the period 1981 to 2001, we apply a propensity score re-weighting decomposition approach to investigate the extent to which the large decline in the male employment-population rate over this period can be attributed to changes in socio-demographic characteristics. We find that changes in observed characteristics account for little of the aggregate decline. However, changes in characteristics are found to be important for population sub-groups. In particular, changes in partner status and partner employment status have acted to decrease employment rates of younger males, but increase employment rates of older males. A further finding is that, holding observed characteristics constant, there has been a very large decline in the employment rate of 55-64 year olds with bachelor degree qualifications. In the course of applying the decomposition method, we illustrate that validity of inferences depends on 'appropriate' specification of the re-weighting function.

1. Introduction

Male employment – especially full-time employment – has declined substantially in recent decades in Australia. In 1966, 79% of males aged over 15 years were employed full-time. By 2006, this had fallen to 58% (Figure 1). The decline in full-time employment is only partially accounted for by growth in part-time employment, since the aggregate employment-population rate of males over 15 years of age fell from 83% to 68% over the same period. This is a pattern to some extent experienced by all OECD countries in the post-1970 period, although there is considerable variation in the extent and precise timing of the decline. For example, in the UK, the 15-64 year old male employment-population rate fell from 93% in 1973 to 72% in 1993, whereas in the US the decline was concentrated in the pre-1983 period, since when the male employment rate has been reasonably stable (OECD 1996, 2006).

Figure 1: Male employment rates 1966 to 2006



Source: ABS Labour Force Survey.

Given available data, it is extremely difficult to disentangle the effects of the many supply, demand and institutional factors that have potentially driven the decline in male employment. In this paper, we focus on investigating the nature of the decline in male employment by considering the changes in the observed characteristics of males between 1981 and 2001. Specifically, we decompose changes into those potentially attributable to changes in male personal and family characteristics and those due to changes in employment rates of males of given characteristics. We furthermore examine patterns of ‘unexplained’ changes by age, educational attainment and partner status to provide insights into the factors behind the changes.

Examination of the 1981-2001 period is dictated by availability of suitable data. While much of the decline in the employment rate pre-dates 1981, it is nonetheless the case that there was a significant decline between 1981 and 2001. There have also been substantial changes in the characteristics of males over this period, suggesting significant potential for much of the decline to be attributable to these characteristics changes. This includes population ageing, increased educational attainment, decreased incidence of partnering and dependent children, and increased educational attainment and employment of partners for those males who are partnered (see Appendix Table A2). In addition to examining aggregate employment rates, we also investigate the role of characteristics changes in producing changes in employment rates for male population subgroups defined by the key socio-demographic characteristics of age, education and partner status.

Existing research internationally has given little attention to the role of socio-demographic changes in declining male employment. However, a number of studies have considered the roles of other factors – notably incentive effects of welfare policies and retirement income policies (e.g., Gruber and Wise 1998, Parsons 1980, Ruhm 1996 and Autor and Duggan 2001 for the US, Blundell and Johnson 1998 and Campbell 1999 for the UK, Kapteyn and de Vos 1998 for the Netherlands, and Borsch-Supan and Schnabel 1998 for Germany), and of structural changes in labour demand away from low-skill labour produced by changes in technology and international trade patterns (e.g., Juhn 1992, Welch 1997). The emphasis on these potential explanations reflects the tendency for the employment rate declines to have been greatest for older and low-skilled males.

Australian research has likewise primarily looked to welfare policies, retirement income policies and decreased demand for low-skill labour to explain the decline in male employment since the early 1970s (Stricker and Sheehan 1981, Merrilees 1982, 1983, Moir 1982, Miller 1983, Hughes 1984, Kenyon and Wooden 1996, Connolly and Kirk 1996, Borland 1997, Kennedy and Hedley 2003). Borland (1995) and ABS (2003) use micro data to investigate changes in male employment rates in Australia, and come closest to the current study by considering the potential effects of changes in the age and education composition of the male working-age population. However, neither study considers the effects of other characteristics, nor allows for interaction effects between these characteristics. Particularly notable for its absence is consideration of changes in intra-household labour supply decision-making associated with the growth in female labour supply in recent decades. In this paper, we estimate the male employment rates that might have prevailed after 1981 if the distributions of both personal and family characteristics of Australian males had remained unchanged from 1981, taking a semi-parametric approach that allows for non-linearities and interaction effects between characteristics.

The decomposition technique we adopt is the propensity score re-weighting approach attributable to DiNardo, Fortin and Lemieux (1996) (DFL). This has become a popular decomposition technique in recent years, particularly for decomposing changes over time in earnings or income (e.g., Butcher and DiNardo 2002, Biewen 2001, Daly and Valletta 2006, Hyslop and Mare 2005, Lehmann and Wadsworth 2001). The approach is a generalisation of the Oaxaca (1973) and Blinder (1973) decomposition method (e.g., see DiNardo (2002) for the relationship between Oaxaca/Blinder decompositions and propensity score re-weighting) that is very flexible and has several advantages over alternative decomposition methods, including the ability to decompose measures (such as the Gini coefficient) that are not themselves decomposable.

An issue that we confront, which does not appear to have received any attention in the literature applying this decomposition technique, is the validity of the weights obtained from the re-weighting process. Clearly, the key requirement for the re-weighting function is that it achieve its goal of replicating the distribution of observed characteristics in the original (say, 1981) sample. Correspondingly, it would seem incumbent on researchers to test the functional form of the re-weighting function for decomposition analyses based on propensity score re-weighting, something which does not currently appear to be common practice.

We therefore undertake tests of the re-weighting function which compare the distribution of observed characteristics in the reweighted sample with the original sample. Based on these tests, we increase the flexibility of the functional form by incorporating additional interaction terms. We further show that it is indeed important to inferences to allow for these interaction effects. The intuition for this finding is that characteristics and their effects can ‘operate in different directions’ for different population sub-groups. For example, the incidence of partnering with an employed female may be decreasing over the period for low-educated males, but increasing for highly-educated males. Furthermore, effects associated with partnering with an employed female may differ for these two groups of males. Failing to allow for such interaction effects will, naturally enough, lead to failure to fully identify the effects of the socio-demographic changes.

The plan of the remainder of the paper is as follows. Section 2 describes the data and Section 3 describes the decomposition method and the issues that arise in implementing it in the context of male employment rates. Results are presented in Section 4 and Section 5 concludes.

2. The census data

The data we use comprise the publicly released unit record files for the one per cent samples of the Australian population censuses over the period 1981 to 2001. The samples contain individual-level information on a variety of personal and household characteristics, including age, educational attainment, partner status, the presence of dependent children, immigrant status, student status, English proficiency and partner characteristics (including educational attainment and employment status).

Several issues in relation to these datasets warrant specific mention. The first issue concerns the timing of the censuses. The censuses to 1986 were conducted on 30th June, while the last three censuses were conducted in early August (see Table A1). There may therefore be seasonal effects for the first two censuses compared with the last three. There may furthermore be 'day of the week effects' present, although all censuses were conducted on a weekday, and generally mid-week. Second, the labour force status information available from the census data does not precisely match the labour force survey. The census is a snapshot of Australia at the date of the census, with labour force status and hours worked defined with respect to the week preceding the census date. An implication is that some employed persons will have reduced or zero hours due to sickness or holidays. We do not attempt to adjust for this.[1]

Table 1 presents employment-population and full-time employment-population rates. Panel A compares the census data with estimates derived from the labour force surveys. The censuses give very similar estimates of the total employment rate, but give consistently lower estimates of the full-time employment rate, in the order of 4 to 5 percentage points lower. This is

[1] Between 3 and 5 per cent of employed persons in each census have hours recorded as 'not stated'. These workers are assumed to have the same distribution of hours worked as workers who reported working hours.

expected given the approach taken to defining the employment status and hours worked when using the census data. The relative consistency of the differential means that changes over the period are quite similar for the two data sources. Panel B presents estimates for the 15-64 years age-range on which we focus. Changes over the period are little-affected, suggesting population ageing explains little of the decline in the employment rate among all males (over 15).

Table 1: Males employment rates, 1981 to 2001

	1981	1986	1991	1996	2001	Change 1981 to 2001
A. Comparisons between labour force survey and census estimates						
Employment rates – Males aged 15+ years						
Labour force surveys	74.40	70.55	66.58	66.93	66.78	-7.62
Census	74.19	70.03	66.07	65.98	67.00	-7.19
Full-time employment rates – Males aged 15+ years						
Labour force surveys	70.34	65.83	60.41	59.01	56.96	-13.38
Census	65.74	61.09	55.90	54.07	53.15	-12.59
B. Effect of age restriction						
Census employment rate estimates – Males aged 15-64 years						
Employment rate	81.30	77.50	73.31	74.08	75.03	-6.27
Full-time emp. rate	72.46	67.93	62.30	61.15	60.02	-12.44

Notes: The estimates from the labour force surveys are not seasonally adjusted and are June estimates in 1981 and 1986 and August estimates in 1991, 1996 and 2001. The June labour force survey employment rate estimates were approximately 0.5 of a percentage point higher than the corresponding August estimates in 1981, and were approximately 1 percentage point higher in 1986.

3. Propensity score re-weighting method

3.1 The DFL decomposition method

The employment rate at a point in time, $E_t(e)$, is a mean of an individual employment rate variable and, following the decomposition approach of DFL, may be expressed as the integral of the mean conditional on a set of characteristics x and on a date t_x, $E(e \mid x, t_e)$, over the distribution of individual characteristics $F(x \mid t_x)$ at date t_e:

$$\begin{aligned} E_t(e) &= \int_{x \in \Omega_x} E(e \mid x, t_e = t)\, dF(x \mid t_x = t) \\ &\equiv E(e; t_e = t, t_x = t) \end{aligned} \qquad (1)$$

where Ω_x is the domain of definition of the individual characteristics.

The notation in the second line of Equation (1) allows us to express equations for counterfactual employment rates, with t_e denoting the date from which the set of employment rates for each 'characteristics bundle' is drawn, and t_x denoting the date from which the distribution of characteristics is drawn. For example, while $E(e;t_e = 2001, t_x = 2001)$ represents the actual employment rate in 2001, $E(e;t_e = 2001, t_x = 1981)$ represents the employment rate that would have resulted in 2001 had the distribution of individual characteristics remained as it was in 1981. This hypothetical employment rate is identified as follows:

$$\begin{aligned} E\left(e;t_e = 2001, t_x = 1981\right) &= \int E\left(e \mid x, t_e = 2001\right) dF\left(x \mid t_x = 1981\right) \\ &= \int E\left(e \mid x, t_e = 2001\right) \psi_x(x) dF\left(x \mid t_x = 2001\right) \end{aligned} \tag{2}$$

where $\psi_x(x)$ is a "re-weighting" function:

$$\psi_x(x) \equiv \frac{dF\left(x \mid t_x = 1981\right)}{dF\left(x \mid t_x = 2001\right)} \tag{3}$$

The equation for the counterfactual employment rate is identical to the equation for the 2001 employment rate except for the function $\psi_x(x)$, so that once an estimate of this function, $\hat{\psi}_x(x)$, is obtained, the counterfactual employment rate can be estimated as:

$$\hat{E}\left(e;t_e = 2001, t_x = 1981\right) = \sum_{i=1}^{n_{2001}} \frac{\hat{\psi}_x(x_i)e_i}{n_{2001}} \tag{4}$$

where n_{2001} is the 2001 sample size and the summation is over observations in the 2001 sample. Applying Bayes' rule to the ratio $\frac{dF\left(x|t_x=1981\right)}{dF\left(x|t_x=2001\right)}$ gives the following equation for the re-weighting function:

$$\psi_x(x) = \frac{\Pr(t_x = 2001)}{\Pr(t_x = 1981)} \cdot \frac{\Pr(t_x = 1981 \mid x)}{\Pr(t_x = 2001 \mid x)} \tag{5}$$

The probability of being in period t given characteristics x $\left(\Pr(t_x = t \mid x)\right)$ can be estimated non-parametrically, by identifying the proportion of individuals with each characteristic combination at each date, or by a discrete choice model like the logit, with the x's entered in a reasonably flexible way. For the case where *x* is a set of dummy variables, DFL is identical to Oaxaca/Blinder with the *x* variables used in an equation for the employment rate (DiNardo, 2002).

3.2 The re-weighting function

An issue in the implementation of the DFL method that has received little attention to date concerns the appropriateness of the specification (functional form) of the model used to re-weight observations. The potential exists for the re-weighting function to fail to produce a re-weighted sample with the same distribution of attributes as the original sample, in turn potentially leading to incorrect inferences. While it is understood that a 'reasonably flexible' specification of the model used to produce the re-weighting function is required (DiNardo *et al*, 1996), existing research appears not to actually evaluate the re-weighting procedure.

In principle, the problem can be avoided by adopting a completely non-parametric specification of the logit; but in practice the 'curse of dimensionality' means this is usually not viable. The question therefore essentially reduces to "what are the deviations from a non-parametric specification that don't cause any significant differences in distribution of attributes between the re-weighted sample and the original sample?" There is no guarantee that such a parametric specification exists, but of course this does not alter the fact that it is important that a specification achieve its intended function if inferences are to be valid. Inability to find a specification that can be estimated with the available data and correctly-re-weights observations simply means the data available do not support the making of valid inferences.

In the current context, failure to employ a sufficiently flexible specification may lead to understatement or overstatement of the roles played by changes in characteristics in the observed changes in male employment rates. At a general level, the intuition for this is fairly straightforward: if a sample for one year (x) has not been *correctly* re-weighted to have the same distribution of observed characteristics as the sample in the other (comparison) year (y), then we of course cannot say that we have fully identified the effects of changes in the observed characteristics composition of the population between years x and y when we compare the actual year-x sample with the re-weighted year-x sample.

The need for interactions could apply to any group of characteristics, but possibly most important is to allow for interactions between variables for partner's employment activity and other characteristics. It is likely that growth in female labour force participation has had implications for both male labour supply (particularly if labour supply decisions are made at the family level) and demand for male labour. We furthermore hypothesise that changes in family status (especially partner status and partner employment status) have not been uniform across age-education groups, and that the extent to which given family status changes can account for changes in employment also differ by age-by-education group. For example, family status may have different influences on the labour supply decision for old and young men. Interacting family status with age and education when obtaining the re-weighting function would allow for differences by age and education in both the implications of family status for male employment and the changes in family status.

To illustrate, consider the simple scenario where there are only three age groups (*y*oung, *p*rime, *o*ld), two family states (*m*arried, *s*ingle), and no other characteristics. The employment rate in 2001 can then be expressed as a weighted average of the employment rates of the six age-by-family-status groups:

$$E^{01} = p^{01}_{y,m}E^{01}_{y,m} + p^{01}_{y,s}E^{01}_{y,s} + p^{01}_{p,m}E^{01}_{p,m} + p^{01}_{p,s}E^{01}_{p,s} + p^{01}_{o,m}E^{01}_{o,m} + p^{01}_{o,s}E^{01}_{o,s}$$
$$\text{where} \quad p^{01}_{y,m} + p^{01}_{y,s} + p^{01}_{p,m} + p^{01}_{p,s} + p^{01}_{o,m} + p^{01}_{o,s} = 1 \tag{6}$$

If age and family status are not interacted in the logit model used to obtain the re-weighting function, then there is no requirement that each of the population shares (denoted by p's) in Equation (6) be equal to the 1981 values in the re-weighted distribution. Any combination that satisfies the following five conditions is possible:

$$\begin{aligned} p^{cf}_{y,m} + p^{cf}_{y,s} &= p^{81}_{y} \\ p^{cf}_{p,m} + p^{cf}_{p,s} &= p^{81}_{p} \\ p^{cf}_{o,m} + p^{cf}_{o,s} &= p^{81}_{o} \\ p^{cf}_{y,m} + p^{cf}_{p,m} + p^{cf}_{o,m} &= p^{81}_{m} \\ p^{cf}_{y,s} + p^{cf}_{p,s} + p^{cf}_{o,s} &= p^{81}_{s} \end{aligned} \tag{7}$$

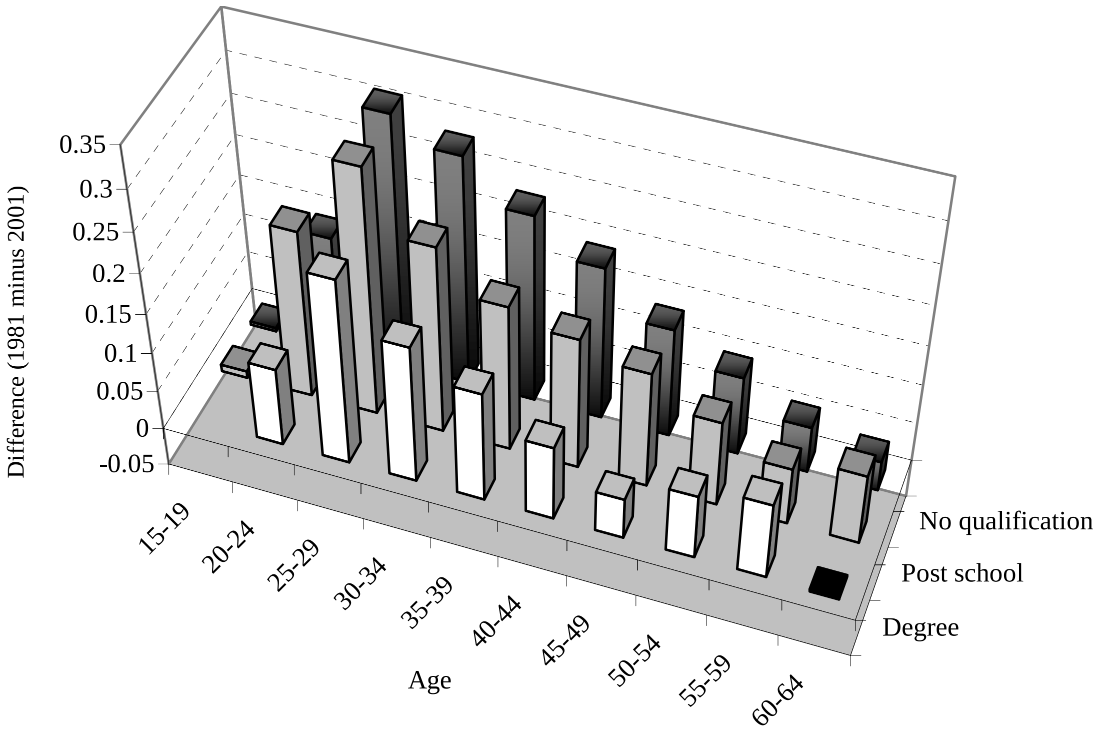
Since there are five equations and six unknowns, multiple solutions exist. We are therefore not guaranteed that the $p_{i,j}$'s $(i = y, p, o; j = m, s)$ obtained in the re-weighted 2001 sample will in fact be equal to the $p_{i,j}$'s found in the 1981 sample.[2] For example, if we have a decrease in $p_{y,m}$ and an increase in $p_{y,s}$ but no aggregate change in p_y, p_m or p_s, the "un-interacted" approach does not change the weighting given to $p_{y,m}$ and $p_{y,s}$. An interacted

[2] More generally, if there are *K* characteristics, with each characteristic *k* containing I_k values, there will be $\prod_{k=1}^{K} I_k$ unknowns and $\sum_{k=1}^{K} I_k$ equations if all characteristics are fully interacted.

specification, by contrast, will result in all six population shares being adjusted to 1981 levels, allowing for differential effects on employment rates of each of the six cells, differential changes in population shares of each of the six interacted cells, and differential changes in employment rates within each of the six cells.

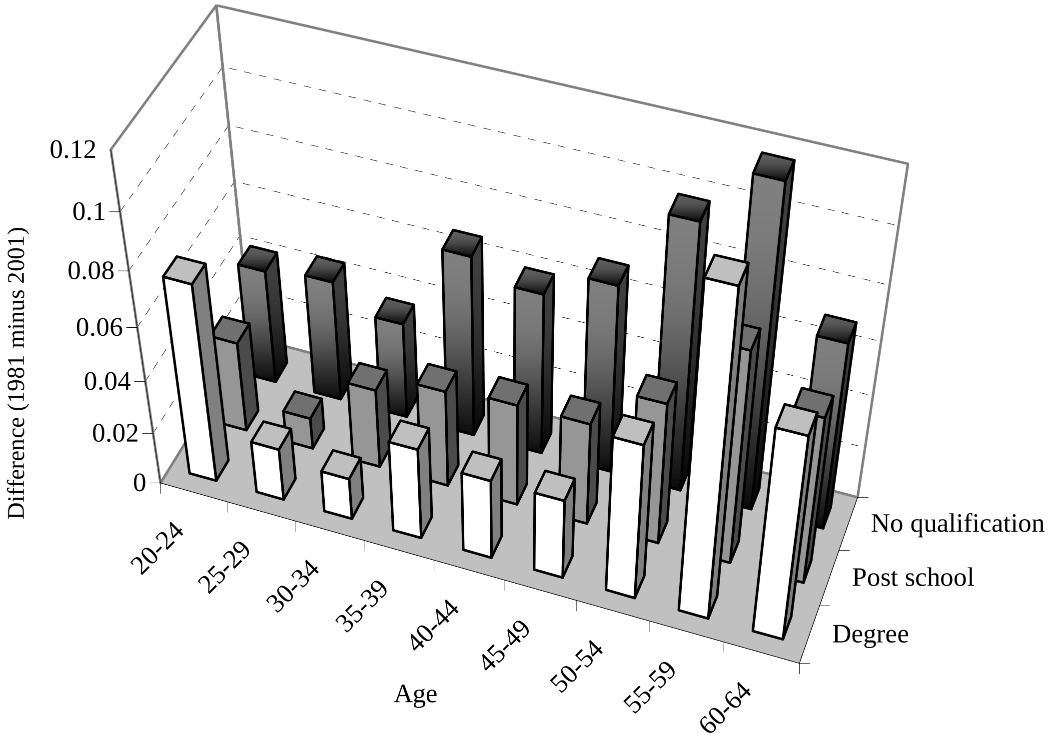
Figures 2 and 3 provide some support for the hypothesis proposed above with regard to family status changes. They show that changes in partner status have not been uniform across age-by-education groups, and that changes in employment rates among those partnered differ by age-by-education group.[3] They also illustrate the more general point of the potential importance of testing the re-weighting function and allowing for interactions. Figure 2 shows changes in the partner status composition of the male population are quite variable across education and age groups – the trend decline is not uniform across age-education groups, and indeed is non-monotonic in age and education. This does not of itself have implications for the re-weighting function, but it does in the context of Figure 3, which shows quite different changes in the employment rate of partnered males across age-education groups.

Figure 2: Proportion of individuals who are partnered: differences between 1981 and 2001



[3] Note that Figures 2 and 3 consider only one dimension of family status: whether the person is partnered or not. In particular, they do not consider partner employment status.

Figure 3: Proportion of partnered men who are employed: differences between 1981 and 2001



Note: The age group 15-19 years is excluded because of the very small number of observations.

In particular, Figures 2 and 3 show that older more-educated men experience relatively little change in partnering rates, but relatively large drops in employment rates of those who are partnered. Thus, in the absence of interactions between partner status and age-by-education group, one might attribute a decline in the employment rate of older educated males to a decline in partnering of males (as reflected by the average aggregate decline in the partnered rate) when in fact it is due to a 'real' decline in the employment rate. Conversely, an inferred 'real' decline in the employment rate of younger and less-educated males from an uninteracted specification might at least in part derive from a change in characteristics of these males in the form of a decline in partnering (in excess of the aggregate average partnering rate decline, the effect of which is captured by uninteracted partner status variables).

3.3 Identifying an appropriate re-weighting function

Identifying an appropriate specification of the re-weighting function is an iterative process that contains an element of arbitrariness, and the approach taken will depend on the specific circumstances. Nonetheless, in all cases, at a minimum we should test that the mean value of each element of x is the same in the two samples. t-tests of individual variable means, or

Hotelling tests of all characteristics simultaneously, can be undertaken. If there are variables that differ in mean values, the researcher then needs to experiment with modifications to functional form and/or interactions between variables. The process also potentially involves testing whether specific combinations of characteristics – the mean values of specific interactions between variables – are the same in the two samples, if the researcher believes these important to the outcome under consideration.[4] It is, in general, a labour-intensive process.

The particular process we employ for choosing the re-weighting function in the current context is as follows. We estimate logit models that treat 1981 as the base year, so that each sample is re-weighted to contain the same distribution of attributes as 1981. We first estimate the logit models to obtain the weights using a set of elementary demographic characteristics: age, educational attainment, immigrant status, student status, and the presence of dependent children. In addition, we include a set of interactions of partner status, partner's employment status, and the difference in education level between the person and their partner, which produces indicator variables such as 'has full-time employed partner who is more educated' and 'has non-employed partner who is less educated'.

We apply the weights and conduct t-tests for the equality of the mean values for each of these characteristics across the samples. Statistically significant inequalities – which indicate the reweighting function has not achieved its goal of reproducing the distribution of observed characteristics – form the basis for inclusion of additional interaction terms. Decisions on additional interactions to include are based on both the t-test results and (largely intuitive) reasoning on the potential drivers of the inequalities. For example, upon observing that education levels are not equal across the samples we can speculate that this derives from changing levels of educational attainment over time, and so a suitable remedy might be to interact education levels with age. The logit model is then re-estimated with the additional interaction terms to generate new weights. The t-tests are re-run and further interaction terms are added if significant inequalities in means are again observed. This iterative process is repeated until no significant inequalities are obtained.

Once replication of the distribution of uninteracted characteristics has been achieved (such that the average value of each characteristics variable in the overall sample is the same in the re-weighted year as in the base year), we then proceed to consider the distribution of

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[4] The researcher could, of course, include all such interactions in the original logit specification, but this is likely to lead to the same identification problems that preclude a completely non-parametric specification.

characteristics within groups defined by the key socio-demographic characteristics of age, educational attainment, and partner status. That is, we examine the mean values of characteristics variables within each age group, within each education group, and within each partner status group. When inequalities in the means of characteristics are encountered within these disaggregated samples, an often beneficial strategy is to interact the characteristics in question with each of the categories of the disaggregation. For example, inequalities in the immigrant status variables within each of the age categories may be resolved by interacting these immigrant status variables with each age category. As before, the logit model is then re-estimated (on the entire sample) with the additional interaction terms and the resulting re-weighting function tested. This process may involve numerous iterations, with each consisting of the inclusion extra interaction variables. Once all of these t-tests are satisfied, we produce sets of weights for each of the samples of data for 1986, 1991, 1996, and 2001 that enable us to re-weight these samples such that they exhibit the same distributions of characteristics as the 1981 sample.[5]

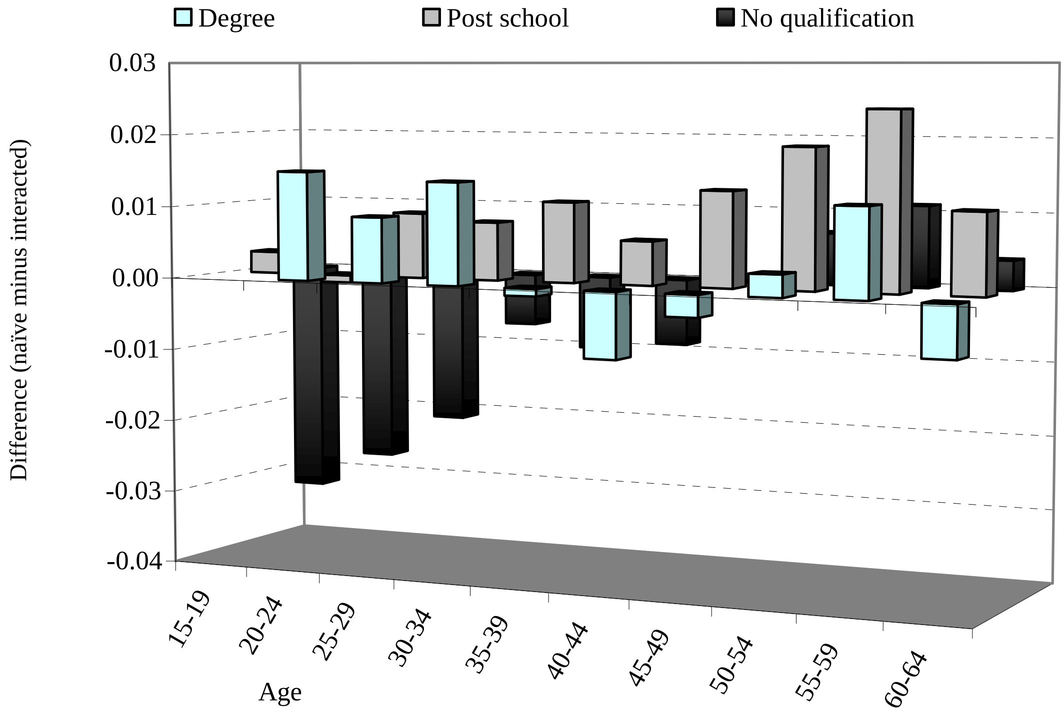
Figure 4 illustrates the point that inclusion of interaction terms in the re-weighting function does indeed matter to inferences on employment rate changes – in this case, for individual age-by-education cells. It presents, for each age-by-education cell, the difference between the 2001 employment rate holding characteristics constant at 1981 levels using the naïve (uninteracted) re-weighting function and the 2001 employment rate change holding characteristics constant at 1981 levels using the re-weighting function containing interaction terms. That is, for each age-by-education cell, the bar represents:

$$\hat{E}(e;t_e = 2001, t_x^n = 1981) - \hat{E}(e;t_e = 2001, t_x^i = 1981) \qquad (8)$$

where t_x^n is based on the naïve re-weighting scheme and t_x^i is based on the interacted re-weighting scheme. A positive value indicates that the naïve weighting scheme positively over-states the change. Figure 4 demonstrates that, firstly, including interaction terms does generally matter, since non-zero differences are evident. Furthermore, the manner in which it matters differs substantially across age-by-education groups.

[5] We do not report the interactions resulting from this process because they are too numerous (the total number of variables in the logit models ranges from 143 to 181), and they differ across years. Details are, however, available from the authors on request. Appendix Table A3 presents the means of the uninteracted variables for the re-weighted samples in 1991 and 2001.

Figure 4: Employment rate changes 1981-2001 by age-education cell, holding characteristics constant at 1981 distribution – Differences between interacted and naïve specification of the re-weighting function



4. Decomposition results

Table 2 presents counterfactual employment rate changes when the 1981 distribution of characteristics prevails. Changes in observed characteristics account for none of the aggregate decline in male employment. This is somewhat surprising in the context of the large changes in characteristics over this period. Indeed, it would seem to be a significant finding, since – aside from the increase in educational attainment – most changes in characteristics would perhaps have been expected to decrease employment.

Table 2 also presents counterfactual changes when the re-weighting function is obtained from a 'naïve' logit specification that contains all the characteristics variables but no interaction terms. Although re-weighting does not achieve the 1981 distribution of characteristics, inferences are in this case little-affected. Nonetheless, it is notable that in moving from the naïve re-weighting scheme to the interacted re-weighting scheme, the (employment-increasing) effects attributable to changes in characteristics become slightly smaller.

Table 2: Effects of characteristics changes on aggregate employment rate changes, 1981-2001

	1981	1986	1991	1996	2001
Employment rate					
Raw	0.810	0.768	0.729	0.734	0.746
Change from 1981	-	-0.042	-0.081	-0.076	-0.064
Change from 1981, keeping observed characteristics at 1981 level	-	-0.038	-0.081	-0.086	-0.067
Change from 1981, keeping observed characteristics at 1981 level (naïve)	-	-0.041	-0.086	-0.091	-0.073
Full-time employment rate					
Raw	0.702	0.657	0.590	0.593	0.579
Changes from 1981	-	-0.045	-0.112	-0.109	-0.123
Changes from 1981, keeping observed characteristics at 1981 level	-	-0.040	-0.105	-0.116	-0.125
Changes from 1981, keeping observed characteristics at 1981 level (naïve)	-	-0.043	-0.110	-0.121	-0.132

Note: 'Naïve' means re-weighting the sample using inverse predicted probabilities obtained from the logit model without interactions. This sets of weights does not pass 'balancing' tests.

In Table 3, we consider 'within-cell' changes in employment rates for cells defined by key demographic characteristics. Specifically, for the 1981-1991 and 1981-2001 periods, we present changes in employment rates for individual age groups, education groups and partner status groups, and identify the changes attributable to (explained by) characteristics changes within these groups. Interpretation of the table is aided by noting that the change in the employment rate between year 1981 and year y (where y is 1991 or 2001) can be expressed, using our earlier notation, as:

$$\underbrace{\hat{E}(e;t_e = y, t_x = y) - \hat{E}(e;t_e = 1981, t_x = 1981)}_{\text{Total change}}$$
$$= \underbrace{\hat{E}(e;t_e = y, t_x = y) - \hat{E}(e;t_e = y, t_x = 1981)}_{\text{Explained change}} + \underbrace{\hat{E}(e;t_e = y, t_x = 1981) - \hat{E}(e;t_e = 1981, t_x = 1981)}_{\text{Unexplained change}} \qquad (9)$$

The 'explained' changes are those attributable to changes in characteristics and are given by the difference between the actual employment rate in year y and the counterfactual employment rate for the year-y sample. A negative value implies that, keeping characteristics at the 1981 distribution, the employment rate would have been higher than that actually realised – that is, changes in characteristics have acted to lower the employment rate.

Table 3: Male employment rates by characteristics – Total change and change due to changes in observed characteristics

	Level in 1981	Change 1981 to 1991 (1991 minus 1981)		Change 1981 to 2001 (2001 minus 1981)	
		Explained	Total	Explained	Total
Employment rate					
By Age					
15-19	0.521	-0.066	-0.139	-0.070	-0.108
20-24	0.821	-0.026	-0.120	-0.042	-0.082
25-29	0.885	-0.016	-0.084	-0.023	-0.062
30-34	0.919	-0.003	-0.071	-0.012	-0.056
35-39	0.925	0.006	-0.075	-0.004	-0.078
40-44	0.914	0.009	-0.060	0.004	-0.058
45-49	0.906	0.013	-0.055	0.015	-0.069
50-54	0.872	0.017	-0.076	0.039	-0.070
55-59	0.773	0.025	-0.099	0.056	-0.098
60-64	0.525	0.024	-0.062	0.056	-0.055
By education					
Degree+	0.935	0.001	-0.041	-0.012	-0.044
Post school qualification	0.910	0.006	-0.065	-0.012	-0.073
No PS qualification	0.755	-0.017	-0.105	-0.029	-0.102
By partner status					
Not partnered	0.693	0.001	-0.075	0.018	-0.033
Partnered	0.890	0.018	-0.066	0.020	-0.057
Full-time employment rate					
By Age					
15-19	0.411	-0.067	-0.196	-0.069	-0.251
20-24	0.704	-0.040	-0.164	-0.067	-0.191
25-29	0.771	-0.022	-0.105	-0.035	-0.107
30-34	0.817	-0.008	-0.098	-0.012	-0.109
35-39	0.819	-0.002	-0.096	-0.009	-0.109
40-44	0.811	0.003	-0.084	0.008	-0.099
45-49	0.811	0.005	-0.098	0.016	-0.120
50-54	0.754	0.008	-0.096	0.040	-0.102
55-59	0.666	0.014	-0.120	0.048	-0.141
60-64	0.420	0.011	-0.085	0.037	-0.092
By education					
Degree+	0.808	-0.003	-0.048	-0.014	-0.061
Post school qualification	0.806	-0.001	-0.093	-0.017	-0.114
No PS qualification	0.647	-0.024	-0.140	-0.035	-0.180
By partner status					
Not partnered	0.574	0.000	-0.106	0.028	-0.107
Partnered	0.789	0.009	-0.096	0.020	-0.098

It turns out that including interaction terms most matters within population sub-groups, rather than for aggregate changes. Sizeable negative effects of characteristics changes are evident for males under 35 years of age and males without post-school qualifications, while sizeable positive effects of characteristics changes are evident for those over 45 years of age and within both partner-status cells. How do we reconcile the findings in Table 3 with those in Table 2? For the analysis by age group, it appears that negative effects for young males are offset by positive effects for older males, to produce a near-zero net aggregate effect. For the analyses by educational attainment and by partner status, however, it must be that compositional changes in the proportion in each education and partner status group have acted to offset the within-group effects. For example, the education composition has shifted towards higher-employment-rate groups (i.e., higher educational attainment groups). This has acted to increase the aggregate employment rate, thereby offsetting the negative within-education-group changes in characteristics. (By its nature, the analysis by education group does not account for effects of changes in education composition, since it examines each education group separately.)

Figure 5 presents similar information to Table 3, but for each of 30 age-by-education cells.[6] It displays, for each cell *i*, the proportion of the total employment rate change between 1981 and 2001 'explained' by changes in characteristics within the cell – i.e., adapting our earlier notation,

$$p_i = \frac{\hat{E}(e_i;t_{i,e}=2001,t_{i,x}=2001)-\hat{E}(e_i;t_{i,e}=2001,t_{i,x}=1981)}{\hat{E}(e_i;t_{i,e}=2001,t_{i,x}=2001)-\hat{E}(e_i;t_{i,e}=1981,t_{i,x}=1981)} \qquad (10)$$

$p_i < 0$ implies effects of characteristics changes have operated in the opposite direction to the actual change, while $p_i > 1$ implies the change due to characteristics changes is in the same direction as the actual change, but is larger.

Consistent with the estimates presented in Table 3, the graphs present a striking picture of the proportion of the employment rate change explained by characteristics changes decreasing in age, particularly for the total employment-population rate. The proportion explained also tends to be highest for those holding bachelor's degrees, and lowest for those holding other post-school qualifications. We furthermore note that changes in characteristics generally account for a high proportion of the full-time employment rate change for prime-age (25-44 years) degree-holders.

As is indicated by Table A2, the characteristics changes driving these results largely comprise changes in partner status and partner's employment status and relative education level. Thus, for younger and more educated males, partner-related changes have been important sources of the declines in rates of employment. For males over 50 years of age, by contrast, partner-related changes have either been unimportant, or have in fact acted to *increase* employment rates (reflected by negative values of p_i in the context of employment rate declines).

[6] Information on the distribution of the number of observations in each cell is presented in Table A4.

Figure 5: Proportion of employment rate change within age-by-education cells explained by changes in within-cell characteristics, 1981 to 2001

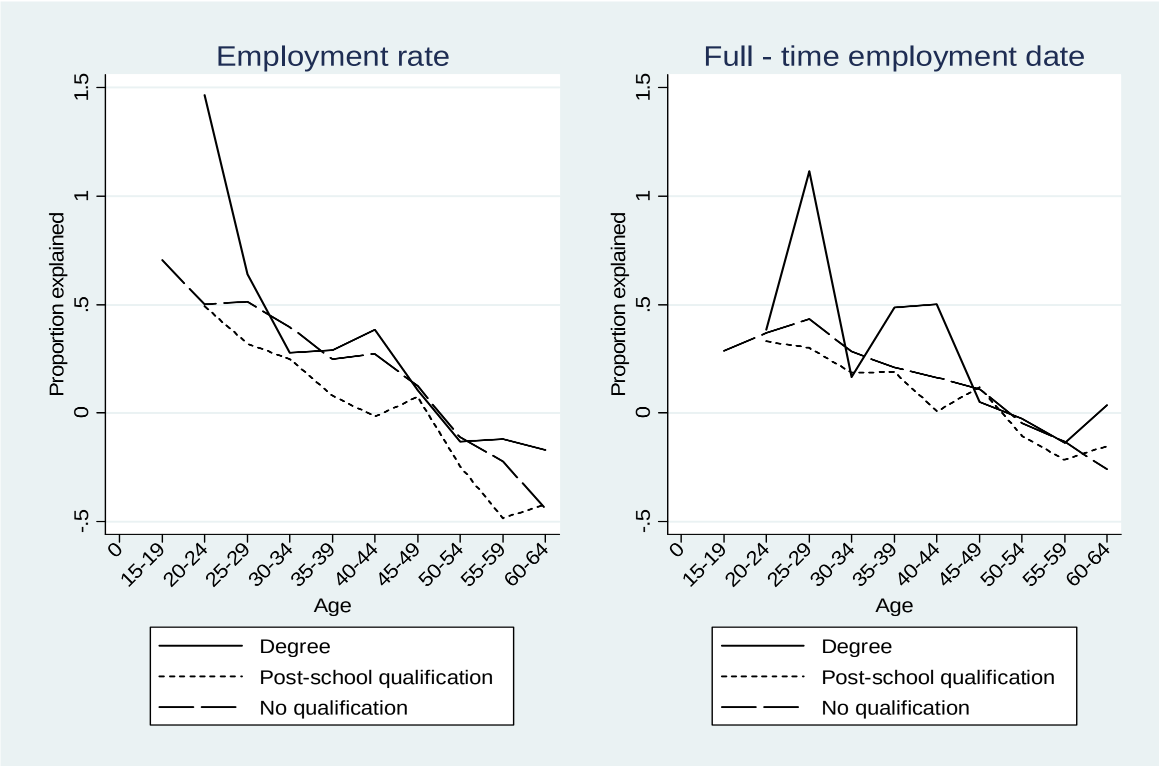
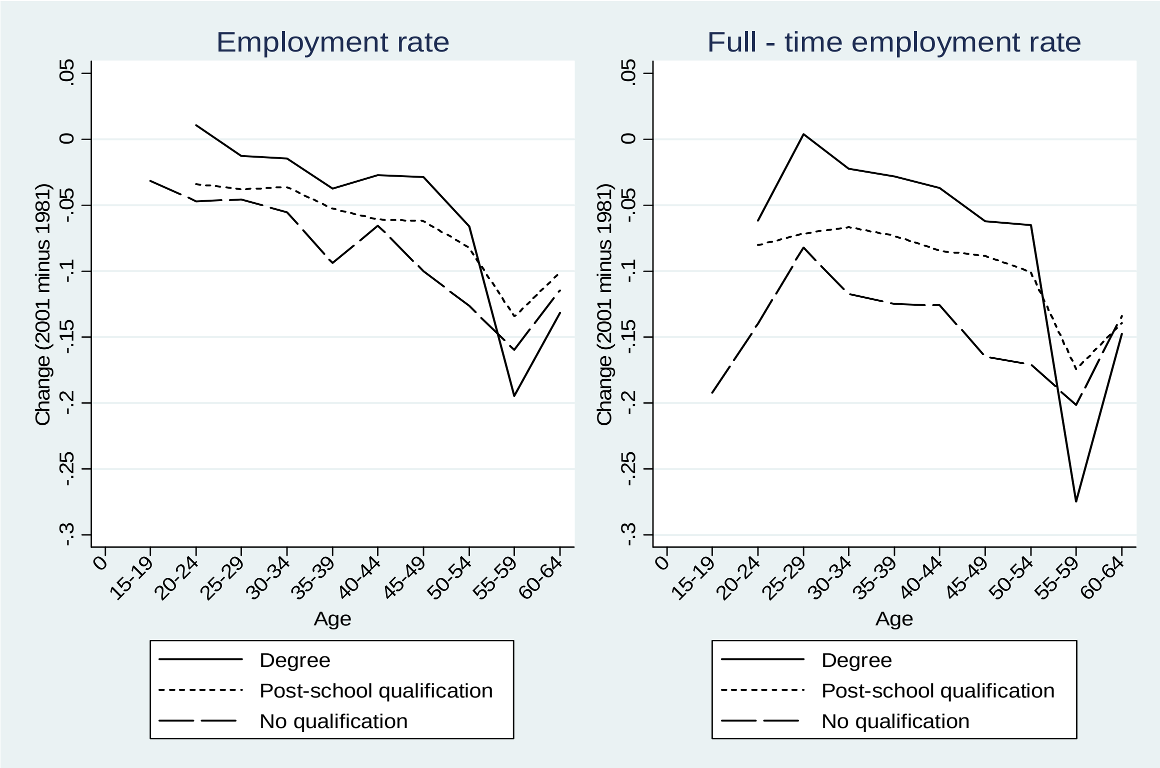


Figure 6 shows changes in the employment rates between 1981 and 2001 for each age-by-education cell when characteristics are held constant at 1981 levels, i.e., $\hat{E}(e;t_e = 2001, t_x = 1981) - \hat{E}(e;t_e = 1981, t_x = 1981)$. We have labelled these as 'unexplained' changes, but they can equally be considered 'real' changes in employment rates, in the sense that they are changes in employment rates that are not simply artefacts of changes in the characteristics composition of the cell.

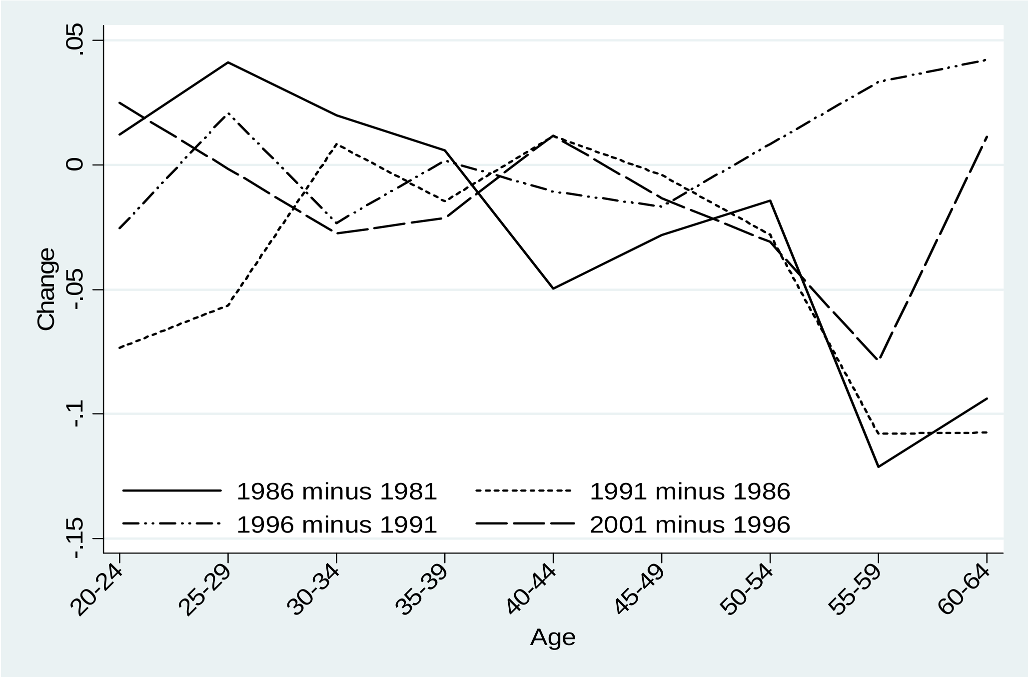
The extent of the decline in the employment rate (and the full-time employment rate) is broadly increasing in age and decreasing in educational attainment. The most notable feature of the graphs is the big drop for 55-59 year old degree-holders, which actually reverses the ordering of employment rate changes by educational attainment evident for younger age ranges. That is, once we control for changes in other characteristics of this age-education group, we find a large decline in the employment rate. This is highly suggestive of a labour-supply induced decline in employment among older males, deriving from an increase in early retirement.

Figure 6: Unexplained changes in employment rates



Based on the preceding results, in Figure 7 we break down the full-time employment rate changes for degree-holders by sub-period to pinpoint the timing of the decline for the 55-59 age group. The decline primarily occurred in the 1980s, although – following a small recovery between 1991 and 1996 – the employment rate again decreased between 1996 and 2001. Additionally evident is that a large decline occurred for 60-64 year old male degree-holders between 1981 and 1991. The decline for 55-64 year old males with bachelor's degrees may have been precipitated by the 1990 recession, but it is curious that the effect is so heavily concentrated on this narrow population group. Changes to retirement income policies may therefore also have played a role. It is also curious that, unlike 55-59 year olds, 60-64 year olds did not experience a further decline in the full-time employment rate between 1996 and 2001.

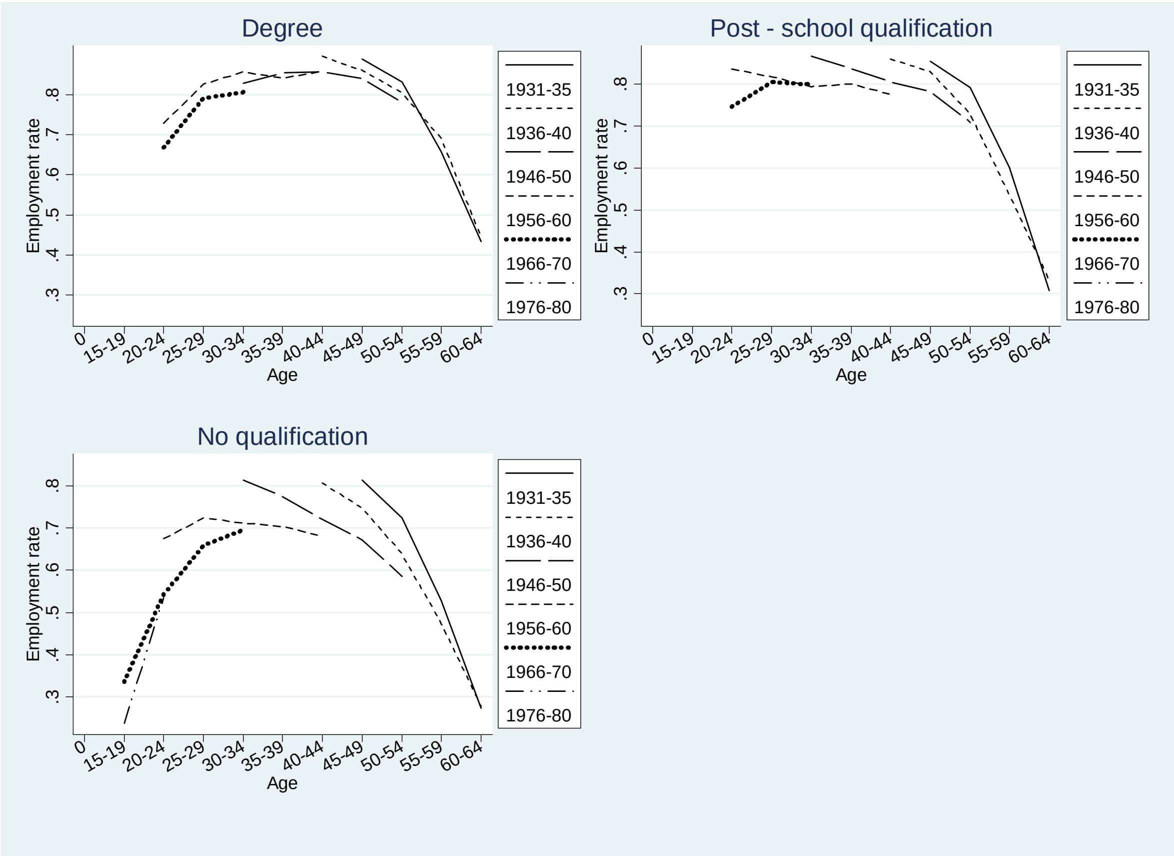
Figure 7: Unexplained changes in the full-time employment rate for degree-holders, by sub-period



The patterns evident in Figure 7 raise questions of the extent to which they reflect differences in employment rates across birth cohorts. Figure 8 takes a birth cohort perspective on the evolution of employment rates over time once effects of characteristics changes have been accounted for. For each of six cohorts comprising males born in a five-year window, the full-time employment rate is plotted against age, holding constant other characteristics at their 1981 distribution using the re-weighting function. The six cohorts are males born in 1931-35, 1936-40, 1946-50, 1956-60, 1966-70 and 1976-80. There are at most five data points for each cohort, corresponding to the five censuses conducted between 1981 and 2001.

The figure shows that earlier cohorts tend to have higher full-time employment rates at each age level, which could reflect unobserved cohort differences or 'time' effects. Differences across cohorts are for the most part greatest for those with no post-school qualifications and least for those holding bachelor's degrees. Also notable is that employment rates exhibit a tendency to decline latest (at the oldest ages) for degree-holders and earliest (at the youngest ages) for those with no post-school qualifications.. Indeed, sizeable declines in employment rates are evident for prime-age males with no post-school qualifications across all cohorts observed in this age range. Nonetheless, when the decline does commence for degree-holders – from around age 50-54 – it is dramatic. The decline in the full-time employment rate over the 50-54 to 60-64 years age range is very large for the 1936-40 birth cohort, and even larger for the 1931-35 birth cohort

Figure 8: Full-time employment rate age profiles, holding constant other characteristics – Selected birth cohorts



5. Concluding comments

The male employment-population rate, and more particularly the full-time employment-population rate, fell substantially over the 1981-2001 period examined in this study. We find that, at the aggregate level, this is not attributable to changes in the socio-demographic characteristics of males. The total decline is therefore ‘real’ in the sense that its magnitude would have been no different had observed characteristics remained unchanged. However, within cells defined by age group, educational attainment or partner status, we find that observed characteristics changes have potentially generated large employment rate changes.

Our ability to identify these effects to a large extent derives from our careful choice of the functional form of the logit equation used to re-weight the samples when applying the DFL decomposition method, which results in us accounting for a number of interaction effects. Indeed, a significant element of this source of employment rate change is the changes to partner status and partner educational attainment and employment status, which differ markedly in both their changes and their employment rate effects across males of different ages and levels of educational attainment. Specifically, partner-related changes have acted to decrease employment rates for younger males, more so for those with bachelor's degrees, but have acted to increase employment rates for males over 50 years of age, particularly among those who do not hold bachelor's degrees.

Perhaps the most striking finding of our analysis is the large decline in the employment rate of 55-64 year-old men with bachelor's degrees between 1981 and 1991. At in excess of 20 percentage points, this is an important development, particularly since there is no evidence of a (sustained) recovery in the employment rate of the 55-59 years age group after 1991. Thus, it appears that there has been a sustained increase in early retirement of educated males. Further investigation of this change, including the role played by government retirement income policies, would seem to be a valuable line of inquiry.

6. Appendix

Table A1: Census dates

Year	Date of census	Day of the week
1971	30 June	Wednesday
1976	30 June	Wednesday
1981	30 June	Tuesday
1986	30 June	Monday
1991	6 August	Tuesday
1996	6 August	Tuesday
2001	7 August	Tuesday

Table A2: Male Characteristics by Census Year (%)

	1981	1986	1991	1996	2001
Age group					
15-19	13.5	12.9	12.7	11.2	11.1
20-24	13.2	12.5	12.2	11.4	9.9
25-29	12.4	12.5	11.8	11.4	10.4
30-34	13.0	11.8	12.7	11.6	11.2
35-39	10.5	12.1	11.2	11.8	11.3
40-44	8.9	9.6	11.3	11.2	11.3
45-49	7.6	8.1	9.0	10.7	10.7
50-54	7.9	6.9	7.1	8.6	10.0
55-59	7.5	7.3	6.1	6.7	7.9
60-64	5.7	6.3	6.0	5.4	6.2
Place of birth					
Australian born	73.3	73.1	73.3	73.4	73.8
ESB	11.7	10.9	11.2	10.9	10.5
NESB	15.0	16.0	15.5	15.7	15.7
Educational attainment					
Bachelor degree or higher	6.2	7.3	9.8	13.0	14.7
Other post-school qualification	28.4	29.4	28.5	29.4	31.4
No post-school qualification	65.4	63.4	61.7	57.6	53.9
Family with dependent child/ren	49.0	47.8	47.0	44.3	43.1
Partner status					
Partnered	59.6	57.0	54.2	51.9	49.8
Single	40.4	43.0	45.8	48.1	50.2
Partner employment and education status					
Partnered, partner more educated and not employed	19.6	17.7	13.2	12.0	10.7
Partnered, partner more educated and full-time employed	10.8	10.2	9.7	10.0	10.1
Partnered, partner more educated and part-time employed	7.2	7.6	7.9	8.5	9.1
Partnered, partner less educated and not employed	8.5	7.2	5.9	5.7	5.1
Partnered, partner less educated and full-time employed	3.6	3.6	3.7	3.6	3.4
Partnered, partner less educated and part-time employed	3.2	3.6	4.0	4.4	4.4
Partnered, education or employment status 'missing'	6.7	7.0	9.8	7.7	7.1
Currently studying	13.3	13.7	17.2	16.5	16.5
Sample (N)	44,988	48,003	51,310	53,066	56,253

Table A3: Male characteristics re-weighted to 1981 distribution

	1981	1991		2001	
		Basic	Interacted	Basic	Interacted
Age group					
15-19	13.5	13.4	13.5	13.3	13.5
20-24	13.2	13.2	13.3	13.1	13.3
25-29	12.4	12.4	12.4	12.2	12.4
30-34	13.0	12.9	12.9	13.0	12.9
35-39	10.5	10.5	10.5	10.5	10.5
40-44	8.9	8.8	8.9	8.9	8.9
45-49	7.6	7.5	7.5	7.5	7.6
50-54	7.9	7.9	7.8	7.9	7.9
55-59	7.5	7.6	7.5	7.6	7.5
60-64	5.7	5.8	5.7	5.9	5.7
Place of birth					
Australian born	73.3	73.2	73.4	73.4	73.2
ESB	11.7	11.7	11.6	11.5	11.8
NESB	15.0	15.1	15.0	15.0	15.1
Educational attainment					
Bachelor degree or higher	6.2	6.2	6.2	6.2	6.2
Other post-school qualification	28.4	28.4	28.4	28.4	28.4
No post-school qualification	65.4	65.4	65.4	65.4	65.4
Family with dependent child/ren	49.0	48.5	48.8	48.3	48.9
Partner status					
Partnered	59.6	59.5	59.5	58.7	59.5
Single	40.4	40.5	40.5	41.3	40.5
Partner employment and education status					
Partnered, partner more educated and not employed	19.6	19.6	19.5	19.3	19.6
Partnered, partner more educated and full-time employed	10.8	10.7	10.8	10.4	10.8
Partnered, partner more educated and part-time employed	7.2	7.2	7.2	7.1	7.2
Partnered, partner less educated and not employed	8.5	8.5	8.5	8.4	8.5
Partnered, partner less educated and full-time employed	3.6	3.6	3.6	3.5	3.5
Partnered, partner less educated and part-time employed	3.2	3.3	3.2	3.3	3.2
Partnered, education or employment status 'missing'	6.7	6.7	6.7	6.7	6.7
Currently studying	13.3	13.5	13.4	13.6	13.3
Sample (N)	44,988	51,310		56,253	

Table A4: Distribution of number of observations in each age-by-education cell

	Median	Maximum	Minimum	5^{th} percentile
1981	1403	5863	79	143
1986	1657	5972	139	166
1991	1483	6340	176	184
1996	1794	5774	154	241
2001	1979	5966	280	380

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