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BREAKS IN WOMEN'S CAREERS DUE TO FAMILY REASONS: A LONG-TERM PERSPECTIVE[*]

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Abstract

We analyse whether family-related quits present long-term effects upon women's careers, which are summarized in three measures of occupational prestige. There is an association between intermittent attachment to the labour market and being engaged in occupations with lower prestige levels. In causal terms, we find that women choose jobs with lower occupational prestige anticipating future family-related quits. The database consists of the retrospective information of the British Household Panel Survey.

Keywords: women's labour career, family-related quits, occupational prestige
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1. Introduction

The objective of this article is to analyze the long-term relationship between family-related quits and women's labour careers. To measure the impact of this type of career breaks we do not use wage changes, but occupational prestige score changes[1]. As Sicherman and Galor (1990) have previously remarked, using wage changes to measure (up)downward career mobility is troublesome. An increase in wages related to occupational mobility might reflect a transition towards a job with negative characteristics compensated (partially or totally) by a higher wage; i.e., a transition towards a worse job. Therefore, we need a measure which unambiguously increases (decreases) with higher (lower) job quality. Here we follow one of the proposals of these authors: the use of occupational prestige scores. Specifically, a negative relationship between family-related breaks from work and the average occupational prestige of women's labour career is expected. Our database is the British Household Panel Survey. The second and third waves include retrospective information on the whole range of employment statuses —including unemployment and inactivity periods— from the first job held to the year 1993. This allows us to analyse women's employment histories during the twentieth century in Great Britain (the North of Scotland is excluded from the survey). Thus, our data are particularly suitable for studying the association between family-related breaks and women's labour careers.

The historical increase in women's participation in the labour market has been widely documented (Mincer, 1962). In spite of this, women not only spend less time overall in the labour market than men, but they are also less likely to work continuously (Mincer and Polachek, 1974 and 1978; Corcoran and Duncan, 1979; Goldin, 1989; Hill and O'Neill, 1992). Therefore, it is not only important to consider total work experience during their life-cycle, but also their intermittent attachment to paid employment. For instance, for a 45 year-old woman, enjoying a continuous 15-year career from the age of 30 may be rather different from a broken career as the following: working for five years from the age of 16, then stopping work from age 21 to 35 and, finally, going on to work for an additional 10-year period. The former case corresponds to a much delayed entry into the labour market but with continuous attachment, while the second one seems to be a typical family-related break due to marriage or child care. The impact of these situations may potentially be rather different.

[1] Previous work about the effect of women's mobility on wages is, for instance, Keith and McWilliams (1995) or Jacobsen and Levin (1995).

A great challenge of this literature is to disentangle the effects of family related quits on career –here, the occupational prestige– from the effects of choosing an occupation by anticipating future family related quits. This 'chicken-or-egg' problem has been studied by, among others, Gronau (1988). We will deal with this version of the endogeneity bias assuming that individuals have rational expectations about their future careers, and, consequently, we will use the observed 'future' family related quits as a proxy of expectations when women chose their 'current' occupations.

Our results show that women who have breaks due to family reasons experience a long-term negative impact in terms of lower average occupational prestige, and this association varies according to the timing of the quits. Nevertheless, we find evidence of an endogeneity bias, confirming that the election of jobs by women is affected by expected family related quits.

The remainder of the article is as follows. In the next section, we present a review of the literature on women's mobility due to family reasons. In the third section, we describe the main characteristics of the data base. The fourth section presents the econometric estimations. The final section summarises the main conclusions of the article.

2. Women's Mobility Due to Family Reasons: a brief review

One of the most important historical changes in Western labour markets throughout the twentieth century has been the increase in labour market participation by women, especially married ones (Goldin, 1989). However, as many authors have stressed (Smith and Ward, 1984; O'Neill, 1985; Moulton, 1986; Goldin, 1989), women's average years of work experience have increased very little. The key to such a weird combination of facts lies on the analysis of work experience throughout the life cycle. According to Goldin (1989), the greater the tendency of women to remain in the workforce over the life cycle, the more their increase in labour-force participation will reduce employed women's accumulated work experience. The reason is that the more heterogeneous women are with regard to labour supply, the more increases in participation will bring less experienced women into the labour force. Therefore, career interruptions become potentially a key issue in understanding women's labour history from a long-term perspective.

There is an extensive amount of literature stressing the importance of childbearing decisions, family formation and family care in order to understand the labour supply behaviour of women (see, for example, Killingsworth and Heckman, 1986, for an overview). One of the most important effects of family care on women's labour opportunities is their intermittent attachment to the labour market (Mincer and Polachek, 1974; Gronau, 1973; Corcoran and Duncan, 1979; Even, 1987). Relevant works exist on the effect of intermittency on wages (Stewart and Greenhalgh, 1984; Mincer and Offek, 1982; Stratton, 1995; Jacobsen and Levin, 1995, Keith and McWilliams, 1995). A remarkable result gleaned from this literature is that women who interrupt their careers and leave the labour market due to family responsibilities often return to find that their wages lag behind those of women at comparable stages in their careers who did not leave the labour force. Many reasons account for this lag. First, women who leave the labour force do not build up seniority, which, by itself, leads to higher wages. Second, women who return to the labour force are less likely to receive on-the-job training to increase their productivity and thereby raise their pay. Third, when women are not in the work force, their job skills may depreciate. Finally, employers may view gaps in work history as a signal that women who leave may do so again, and, therefore, some employers would therefore hire them for less important, low-paid jobs to limit the impact of a future decision to leave. Nevertheless, there is an inconclusive discussion in this literature about whether there is a rebound effect or not.

The use of wage changes to study the effects of career interruptions on labour market outcomes has some disadvantages, some of which are discussed by Sicherman and Galor (1990): if positive characteristics of jobs are compensated by negative wage differentials, upward occupational mobility may not be detected by merely computing wage differentials (Sicherman and Galor 1990). Since some aspects of job quality are better captured by occupational structure, the use of occupational prestige scores might help overcome this problem[2]. These scores have a direct and unambiguous relationship with occupational mobility: upward (downward) mobility towards an occupation with better characteristics is always related to a higher (lower) score, because positive (negative) characteristics of the job are always related to higher (lower) occupational prestige[3]. Furthermore, there are two practical reasons to prefer occupational prestige scores to wages in this research. First, as we are using retrospective data on individuals'

[2] Sicherman and Galor's (1990) ranking of occupations (pp. 189), is very similar to measures of occupational status or prestige developed by sociologists. Indeed, their index is highly correlated with the Duncan socio-economic status index and the NORC occupational prestige index.

[3] In addition to Sicherman and Galor (1990), occupational prestige scores have been also used in Economics in order to analyse the risk of fatal injury (Marin and Psacharopoulos, 1982)

life course, wages are not available for every job, since the quality of the answers would be very low (due to recall error). Instead, the only information needed to include each job's occupational prestige is the type of occupation held in every past job, which is much easier information to remember than former wages for every job. Second, unless one is able to observe the complete wage profile following an interruption, looking only at immediate post interruption wages might give a misleading picture of the effect of the interruption on future earnings. Based on these premises, occupational information could serve as a substitute for a long-run wage profile analysis, allowing coverage of the complete life course. Nevertheless, using occupational prestige indicators is not the panacea, mainly because life-cycle models proposed by economists are based on the crucial relevance of lifetime earnings (but not lifetime prestige). On the other hand, it is unlikely that any difference in occupational prestige which is not captured by differences in long-term earnings will exclusively reflect compensating wage differentials (unless we define any such differences as 'compensating wage differentials'). However, as collecting information about wages for the whole life course in surveys or administrative databases is highly problematic, the use of occupational prestige might be considered as a reasonable and useful 'second-best' solution[4].

In order to obtain robust results, we will use three occupational prestige scores: the Camsis score, the Hope-Goldthorpe score, and the Cambridge score. Out of these, the most widely known is the Hope-Goldthorpe one. We include the other two because they consider differences by gender (Camsis) or life-styles (Cambridge), which may potentially be important for our analysis. It is important to remark that the three scores where obtained using information originally collected for the United Kingdom. The details of the three scores are described in Appendix C. All occupational prestige indexes exhibit strong correlation indexes (correlation coefficients of 0.8 and 0.9 were found by Wegener, 1992). Moreover, they have great stability over time: since 1925, the structure of occupational prestige has remained almost constant in Western countries (see Hauser and Featherman, 1977). Thus, the use of these occupational prestige

[4] The earnings information in the BHPS is only collected in the panel questionnaire but not in any of the three retrospective life-course questionnaires. As the retrospective information matches with the first years of the British Household Panel Survey, it is only possible to use the earnings for the last observed employment spells (when they end and/or begin between 1990 and 1993). This type of earnings information is totally unsuitable for our research.

indicators is especially appropriate for detecting long-term effects with retrospective data covering the most part of women's careers in the twentieth century[5].

3. Database

Our data come from the first three waves of the British Household Panel Survey (BHPS) and three special retrospective questionnaires passed along the second and third waves. The first wave was designed as a nationally representative sample of the population of Great Britain living in private households in the autumn of 1991 (the north of Scotland is not included). Approximately, 5,500 British households (containing about 10,000 persons) were interviewed. See Taylor (1997) for the technical details of the BHPS.

Information is recorded on labour market status at the time of each interview, and for the period between 1st September a year before and the interview date. Thus, for respondents present at waves 1 to 3, we have a complete and detailed record of their labour market status from 1st September 1990 (or before: the start date of a job held at that date is known) to at least 1st September 1993. In addition, for our analysis, it is also necessary to have information on each woman's entire career. In order to fill the gap between leaving full-time education and the beginning of the panel-derived labour market history, retrospective data were also collected in waves 2 and 3. In wave 2 a complete employment status history was collected, recording non-employment states in detail, as well as histories for child bearing and union formation for all respondents in the panel. In wave 3 a complete job history was collected with detailed information on every job held. These retrospective data are matched to the within panel data to construct detailed marriage, fertility and work histories for every woman in the survey from her first job up to 1993. This enables us to provide estimates for several cohorts of the UK population, and also avoids the problem of left hand censoring, which often arises when using the panel component only. A comprehensive description of the retrospective modules in the BHPS can be found in Halpin (1997).

Our analysis uses a sub-sample consisting of all women aged at least 34 years-old at 1st December 1993, so as to avoid very short life histories. Given that most

[5] The sociological literature about occupational prestige scores is very wide. In addition, to the Hope-Goldthorpe score, there are other scores very popular as the Duncan index. As we have explained, usually all scores present very high correlations and great long-term stability, but we have preferred scores based on information originally collected in the United Kingdom (as our data base), and, among them, a score (the Camsis scale) which explicitly includes the differences in prestige when the same occupation is held by men or women. See Appendix C and Malo and Muñoz-Bullón (2007) for additional details.

women's family-related breaks from work occur at the beginning of their labour careers and that our interest lies in whether or not they have any long-term impact on their occupational prestige, we will compare the group of women who have labour force breaks during their first ten years of labour experience with the group of women who do not. This way, enough time is allowed for women to have at least one work interruption. Finally, in order to be sure of comparing two groups of women who are actually different, we erase from the sample those women without family-related breaks during their first 10 years of labour market experience who have ever left their job from the tenth year onwards (they are only 90 individuals). Thus, the group of women with family-related quits must have at least one break from work due to family reasons between their first job held and their tenth year of labour market experience.

As cohabitation is very important in the UK (either as a precursor to legal marriage or as a substitute), we include marriage and cohabitation in only one variable (addressed to as 'unions' in tables). We have the individual's marriage history from the age of 16 up to the data of interview in wave 2. The month and the year of cohabitations leading to marriages are provided, as well as dates for which marriages ended as separations. Similar information is provided about cohabitations that are never made into legal marriages.

As regards birth events, the retrospective history collected in 1992 records the dates of birth of all the respondent's children to that date. These data are recoded into a monthly panel of data covering births or adoptions in each individual's life up to the time of their interview in wave two. These data are then merged with the within panel data to create one event history file, where we have explicitly taken into account when (and whether) children (either natural or adopted) leave home.

The sub-sample used in the empirical analysis consists of 1,833 women. We have considered five birth cohorts as follows: the first cohort includes women who were born between 1906 and 1919; the second cohort refers to women who were born between 1920 and 1929; the third one collects those born between 1930 and 1939; the fourth one, the ones who were born between 1940 and 1949; and the last one, women who were born between 1950 and 1959.

Table 1 presents some cohort characteristics. Most women in the first two cohorts —and partially those in the third one— are above the mandatory retirement age (60 years for women in the UK). Thus, we are able to observe the complete life-cycle evolution of their employment status dynamics. On the contrary, life cycles must be considered as 'right-censored' in the remainder cohorts. In principle, recall bias is a potential problem for any retrospective analysis. However, in practice, previous research attempting to assess the magnitude of recall effects in the BHPS has not found this kind of bias in particular (Elias, 1997). In addition, the BHPS has also attempted to minimize recall error by asking sample members to detail marital and fertility events (which tend to be well remembered) prior to their employment histories, thereby providing a chronological ordering of personal histories aiding the recall of employment events. This procedure has been shown to work well in other surveys. Hence we argue that the recall error in the BHPS labour histories is less of a problem than in most other retrospective data sets.

Table 1 shows descriptive statistics of the set of variables collecting quits due to family reasons (i.e., leaving to have a baby, and due to child/home care): two dummy variables indicating, respectively, whether the woman has ever left the job during her first ten years of labour market experience and whether she has ever left the job from the 10th year of work force experience onwards; the number of quits, and, finally, the ratio of the number of quits over the number of employment spells. As can be observed, around 70% of women on average leave the job due to family reasons during their first 10 years of labour market experience, while only around 10% of women do so from the 10th year onwards. Besides, women have on average one quit, although there are some of them with up to nine quits (Figure 1 shows the frequency of the number of family-related quits). The ratio between the number of quits and the number of employment spells shows how frequent family-related quits are throughout women's labour career. The mean shows that the proportion of employment spells ending in quits is decreasing as we advance from the first to the last cohort. This reduction is the joint result of a rather stable number of family-related quits and an increase in the number of employment spells. Thus, the pattern of quits has changed very little (from 1.07 to 0.99) in comparison to total women mobility (as regards the latter, the mean of employment spells has passed from around 3 employment spells to above 6). This implies that women in the youngest cohorts are less likely to interrupt their employment spells when they marry or have children than the ones in the eldest cohorts.

Two variables that are likely to be important in explaining the potential occupational prestige losses arising from family breaks are whether or not women have ever been married, and whether or not they have children. Comparing two similar women, one of whom has never been married, the has-married one will tend to have more family breaks throughout her career, even more if she has had children. This is confirmed in Table 2, which shows the means and distributions of some of the family-related quit variables collected in Table 1. As can be observed, only 16.15% of never-married women have suffered at least one family break from work during the first 10 years in the labour force, while this proportion rises to 72.44% among women who have been married. There is, therefore, a vast difference between married and non-married women in their rate of family-related quits. Moreover, the distribution of the number of family-related quits throughout the life-cycle is concentrated on very low values for the former group of women, while the opposite happens for women who have been married. The latter have on average five times as many family-related quits as never-married women (1.05 as opposed to 0.21). Finally, on average, the proportion of quits over the number of employment spells is substantially larger among the group of women who have at some point got married. As regards child care, the greater the number of children, the greater is the proportion of women who suffer family-related breaks from work, as well as the mean number of quits and the ratio of quits over the number of employment spells.

4. Empirical results

4.1. The determinants of average occupational prestige scores

In this section we assess the role played by taking a break from work due to family reasons in the first 10 years of the labour career on the measures of women's occupational prestige described above. Since our focus is on the women's entire career, our occupational prestige variable has been obtained by constructing the weighted average of each prestige scale in the different occupations held during their lives. These weights are the proportions of time that sample members spend in their respective occupations[6]. Specifically, the dependent variable for each woman in the sample is the logarithm of the following weighted average:

$$AvgP = \left(\frac{\sum_{i=1}^{N} \left(\text{Prestige of Occupation}_i * \text{Time in Occupation}_i \right)}{\text{Total time occupied}} \right) \quad (1)$$

[6] Also arithmetic averages of the prestige scales in the different occupations held have been calculated. Results obtained with the arithmetic averages are similar to the ones presented in the paper, though the fitness of the different specifications of the empirical model is substantially lower.

This average becomes meaningful if the occupational prestige differs for each of the groups of women under consideration. Figure 2 shows the evolution of this average measure for the Camsis scale score across the different employment spells, distinguishing between women who exhibit family-related breaks in the first 10 years of labour experience and those who have not. Women who have not left any job due to family reasons in general enjoy a larger average prestige measure. In addition, this gap between both groups is larger during the first employment spells, i.e., at the beginning of the career. Finally, the larger the number of employment spells, the lower the average occupational prestige is. Therefore, women who experience more employment spells seem to attain jobs associated, on average, with lower prestige levels.

As the distribution of family-related quits at different moments of the career seems to be important, we have analysed whether there is a family-related quit in the first ten years of the career. While some women have already accumulated ten years of experience at the end of their second employment spell, others do not do so until their third employment spell, or even later. Figure 2 also plots the evolution of the average measure of the Camsis occupational prestige for women who accumulate ten years of labour experience at the beginning of their second and third employment spells, respectively. As can be observed, before accumulating this 10-year experience, women who eventually abandon the labour force enjoy a similar or even greater prestige than the other subgroup of women. However, this trend changes from that moment onwards: the average occupational prestige of those who have left the labour force due to family reasons is usually below the prestige curve of the other subgroup. As we can conclude from those figures, it is interesting to distinguish between the first 10-year period of labour market experience, and the one ranging from the tenth year of labour market experience until the end of the observation period[7].

The empirical model, in addition to the aforementioned variables collecting family-related quits, takes into consideration the following explanatory variables (those variables are described in Appendix A):

- Personal characteristics: dummies for ethnic origin, sex, birth cohort, educational level, the number of unions experienced (either marriages or cohabitations), the number of children (either natural or adopted), a dummy denoting whether or not

[7] In the empirical analysis, we must confront with a potential bias arising from the fact that in our sample there may be some women who do not have any employment spell along the observed period. However, this is the case for only 98 women in the original database. Therefore, this small sample size does not allow us to correct an eventual selectivity bias.

women have currently reached the mandatory retirement age (i.e., 60 at the date of interview) and women's age at their first spell.

- Labour market experience characteristics: continuous variables such as the proportion of time that women have spent in a situation of unemployment or inactivity.

The final specification of the model can then be written as:

$$\mathrm{Ln}(AVgP_i) = \beta_0 + \beta_1 FQ_i + \beta_2 PS_i + \beta_3 LM_i + \varepsilon_i \qquad (i=1,2,\ldots N) \qquad (2)$$

where the subscript *i* refers to each woman, *AvgP* represents the average occupational prestige measure (either the HGS, the Cambridge or the Camsis one), *FQ* collects a family-related quit variable —either the dummy indicating whether the woman has ever left the job due to family reasons, the number of quits, or the ratio of the number of quits over the number of employment spells—, *PS* collects personal characteristics, *LM* collects labour market experience characteristics and ε_i is the error term with $E[\varepsilon_i]=0$. The parameter of primary interest is β_1, the effect of family quits on the outcome variable.

The OLS parameter estimates are presented in Tables 3, 4 and 5 for each of the three measures of prestige and three specifications of the model. These different specifications correspond to the different variables collecting family-related quits described above.

For any of the definitions, the family-related quit variables are, in general, statistically significant and with the expected negative sign. If we keep the remainder variables constant, those women who have quitted from their jobs due to family reasons during their first ten years of labour market experience suffer a reduction in their estimated prestige levels of around 4 percent during their life-course[8].

In a similar way, significant negative impacts are also associated both with the ratio of quits over the number of employment spells and with the number of quits. As observed in Table 3, for instance, a unit-increase in the number of family-related quits

[8] Predictions of the dependent variable for reference women offer a Camsis scale score equal to 29.17 for women who have suffered no quits due to family reasons, and 28.01 for those have ever suffered at least one family-related quit. Looking up for the occupations leading to this predicted impact, according to the Standard Occupational Classification, the change from the occupation named as "All other occupations in farming & related" (with a Camsis Scale score of 31.49 in group 902) to that named as "Packers, bottlers, canners, fillers" (with a Camsis Scale score of 28.35 in group 862) is the one which better approximates the 4-percent reduction in the average occupational prestige. In addition, a histogram of the average camsis scale score by family-quit status shows that the distribution is more concentrated around lower values for the group of women who have ever quitted from work due to family reasons during their first ten years in the work force (not shown, but available from the authors upon request).

presents a negative impact on the Camsis scale score of 5.1 percent[9]. This result, therefore, implies that the effects of family-related quits depend on the existence of additional quits following an initial workforce gap. Finally, results are very similar for the other two prestige scales.

As a robustness check we have estimated the effects of previous family-related quits on the current employment spell controlling for unobserved heterogeneity. For each job held, we gather its duration, the individual's age at the beginning of that job, the duration of the intermediate non-employment spell existing between the previous employment and the current job, and whether or not the woman has quit the previous job due to family reasons. Our approach is to use a fixed-effects estimator to control for unobserved characteristics that may be correlated with displacement probabilities. For instance, if less able or less labour-market-motivated women are more susceptible to quitting due to family reasons, estimates of displacement effects that fail to control for individual-specific heterogeneity will be biased toward finding larger prestige losses[10].

More specifically, the effects of family-related quits observed for woman *i* at employment spell *t-1* on prestige levels associated with the current occupation at employment spell *t* can be modelled in the following way:

$$\mathrm{Ln}(P_{it})=X_{it}\,\beta + Z_{it-1}\,\alpha + \lambda_{it} + \varepsilon_{it} \qquad (i=1,2,\ldots N)\ (t=1,2,\ldots,T) \qquad (3)$$

where P_{it} is the individual i's prestige score associated with the current job; X_{it} and Z_{it-1} are two vectors of observable variables associated with, respectively, the current and the previous job, which potentially influence a woman's prestige at the present occupation; λ_{it} is a time invariant individual specific error that captures the effects of unobservable characteristics; and ε_{it} is assumed to have a constant variance and to be uncorrelated across individuals and jobs. The parameters of interest (α, β, λ) are estimated using the within-group technique, which is equivalent to a simple least squares estimation of the model in which the variables are defined as deviation from their means (it consists of a generalisation of the "differences-in-differences" technique). In estimating the model,

[9] For the reference women, the predicted Camsis prestige score equals 33.02 when no family-related quits are experienced, and 31.33 when one quit is suffered. The nearest associated occupations according to the Standard Occupational Classification correspond to those named as "Clothing cutters, milliners, furriers" (with a Camsis Scale score of 32.61 in group 557) and "All other occupations in farming & related" (with a Camsis Scale score of 31.49 in group 902).

[10] In fact, without including fixed effects, the predicted negative impact of the dummy which collects family-related quits is even larger (results of the pooled regressions are available from the authors upon request). However, this pooled-OLS regression does not take into account the unobserved heterogeneity present in the data.

some of the terms in X_{it} and Z_{it-1} such as education or ethnic origin have been eliminated from the equation since they do not vary with time[11].

Table 6 provides the estimation results of the prestige equation (3) for the three different prestige scales used. As before, we find significant negative impacts associated with the variables collecting quits. In particular, women who have left their previous jobs due to family reasons present a significant reduction in the prestige level associated with the current job. This reduction is approximately 3 percent when the Hope-Goldthorpe scale is used, and nearly 2 percent in case that the Camsis scale is taken as the dependent variable. In addition, as the number of accumulated quits increases, the reduction in the prestige levels from the previous to the current job is greater.

4.2. Endogeneity analysis

In this section, we analyze whether the negative impact of family-related quits is only a mere association or a causal relationship. As we do not have enough information to estimate a structural model inspired in Gronau (1988), we propose a different strategy.

The cornerstone of our problem is that when women choose a job (with a certain occupational prestige) they may consider the different costs of future family-related quits, which are potentially associated with different occupations. As higher occupational prestige is associated with jobs related to long-term attachments and/or higher qualifications (from education or training), women who anticipate that they will experiment future family-related quits will choose occupations with a relatively lower prestige. Here, our problem is to find a proxy of this anticipation of future family-related quits. Assuming rational expectations about future prospects of labour career, we will use the observed family-related quits as a proxy of the expectations regarding the future when women choose an occupation before taking any quit[12].

We have run three OLS regressions on the following three measures of occupational prestige: first, the prestige associated with the first job in their career; second, the average prestige associated with occupations held before the first observed

[11] Given that the variable collecting marital status (whether or not women have ever been married) would also be eliminated from the equation, we estimated separate equations for each group of women: those women who have never been married, and those women who have been married. However, the former subgroup of women does not contain enough observations so as to offer confidence in the estimation results.

[12] In addition, we tried a most conventional strategy applying a Hausman test (Hausman, 1978) to check whether family-related quits are exogenous or not. These results rejected (with only one exception) the endogeneity of family-related quits variables. However, we are not confident of these results because of the problems in finding valid instruments in our database. These estimations are available upon request.

family-related quit; and, third, the prestige associated with the occupation immediately previous to the first observed quit. In these three cases, family-related quits variables are always proxies of future events, because the corresponding occupational prestige was measured before any family-related quit. Table B.1 shows estimation results for only the coefficients of future quits. As can be observed, future family-related quits have only a non-significant impact on the occupational prestige associated with the first job. It is likely that the lack of significant influence on the first occupation is related to the tentative character of most of first jobs: for many workers, the first job is not a very 'significant' job, in the sense that it is rather different from the occupations they will hold during the greatest part of their lives. However, a negative impact arises in the other two cases. Although the size of this effect differs depending on the prestige score, the lowest figures indicate a decrease of around 20 percent, which is a much larger figure than the size of the coefficient for quits in Table 6 (which were around 4 percent). When quits are defined as continuous variables, we have introduced a quadratic term whose estimated coefficient is positive: this negative impact is not linear, but decreasing when family-related quits rise. Thus, there exists evidence showing that future family-related quits exert some influence on the election of jobs by women: if women anticipate a higher number of family-related quits, this fact is associated with their choosing jobs with lower occupational prestige levels (which presumably have lower costs related to interruptions). Finally, we want to remark that the accuracy of these results rests on our assumption on rational expectations regarding future career prospects and the suitability of our proxy for these expectations.

4.3. Other variables

In Tables 3 to 5, unions show a non-significant impact on occupational prestige levels, even though the estimated parameters for the dummies collecting these events present, in general, a negative coefficient for the first union (either cohabitation or marriage)[13]. However, as we would expect, the larger the number of children, the larger the negative impact on prestige levels is. For example, having only one child reduces prestige levels by nearly 10-percent. And having more children lowers the rate of occupational prestige even more (to the extent that the third child represents a 40-percent reduction).

[13] Estimation results have been implemented using the number of marriages and the number of cohabitations separately. In any case, results are robust: non-significance remains. A similar result is obtained by Gronau (1988, pp. 282).

Therefore, having children is associated with a lower average occupational prestige for women in two ways: first, by means of family-related breaks and, second, by children themselves.

Women without studies have lower average prestige levels, as well as a greater proportion of time spent unemployed or inactive. In fact, the highest educational levels —especially university education and higher— are associated with greater prestige levels. In addition, the average prestige scale score is reduced when belonging to the birth cohorts 1906-19 and 1920-29.

Controlling for unobserved heterogeneity (Table 6), two or more unions have a significant positive effect. This result suggests that women with more unions are probably more engaged with their working career. The number of children and the variation in prestige levels are negatively associated, particularly when having two children and for three or more. The negative effect from having two, or three or more children is rather similar, suggesting that the negative impact on women's labour careers is mainly associated with having at least two children (while it is not clear for having one child). Therefore, in general, the results of family variables are consistent with those obtained in previous estimations (Tables 3 to 5), where unobserved heterogeneity was not properly controlled for.

As regards the remainding variables, a positive relationship is found between tenure in the previous position and current prestige gains. Moreover, the longer the permanence in non-employment, the greater the relative prestige loss is. However, the longer the time spent with the current employer, the larger the prestige gain is. Considering the size of these effects, although the impact of past non-employment duration implies the existence of prestige losses, this non-employment incidence is found to have a temporary penalty effect, since it tends to disappear after women re-enter into employment.

Finally, compared to the youngest women (up to 35 years-old), those over 35 are able to enjoy occupations associated with significantly higher prestige levels, and especially those over 45 years-old. This improvement ranges from 3 to 7 percent for those aged from 35 to 45 years-old, while it reaches a nearly 8 percent increase for the oldest women.

5. Conclusions

In this article we have used work-history data from the British Household Panel Survey in order to empirically analyse the effects arising from interruptions in women's labour careers due to family reasons. Our analysis casts new light on the long-term effects of family-related quits and complements in a fruitful way the negative impacts of family-related quits on women's wages found in previous literature. As a novelty, several occupational prestige scales have been applied —in particular, the Camsis Scale, the Hope-Goldthorpe Scale and the Cambridge Scale— as measures of the different positions held by women throughout their life-cycle.

We have estimated the determinants of the average occupational prestige during the woman's entire career. In addition, we have presented a fixed-effects model in order to control for the existence of unobserved heterogeneity. Results show a robust negative association between family-related quits and the average career occupational prestige. This result remains when controlling for unobserved heterogeneity.

We have checked whether these results hide an endogeneity bias or not, since accepting a job might be influenced by the expectations of experiencing future family-related quits, and the eventual higher (lower) costs of these quits for higher (lower) occupational prestige jobs. We have used observed family-related quits as proxies of expectations of future career interruptions when estimating the determinants of the occupational prestige in the first job, the average prestige of all jobs before the first family-related quit, and the prestige of the job held immediately before the first family-related quit. With the exception of the estimation of the occupational prestige of the first job, results confirm that the anticipation of future interruptions has a negative impact on 'current' occupational prestige. Therefore, there is a causal influence of family-related quits on the election of occupations: women who anticipate more interruptions choose jobs with lower occupational prestige. Note that these results do not eliminate the possibility that discriminatory occupational segregation exists. In such a case, there may exist a long-term prestige penalty following any family-related quit. Nevertheless, this research is useful to amplify not only the existing knowledge on how women's careers are affected by their central role in families by providing care (and, in fact, the most part of home production), but also how interruptions due to family reasons may exert long-term consequences on their careers.

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Table 1. Birth cohort characteristics

	Cohort 1 (1)	Cohort 2 (2)	Cohort 3 (3)	Cohort 4 (4)	Cohort 5 (5)
Age at 3rd wave	74-92	64-73	54-64	44-53	34-43
Starting average year of 1st spell	1920	1930	1940	1949	1957
Avg. age at starting year of 1st spell	15	15	16	17	17
QUIT VARIABLES (std. dev. between brackets)					
Have left job due to family reasons:					
During 1st 10 years in work force	.69	.72	.70	.71	.68
	(.46)	(.45)	(.46)	(.45)	(.47)
From year 10th in work force onwards	.13	.13	.09	.07	.07
	(.34)	(.33)	(.29)	(.26)	(.26)
Avg. number of quits due to family reasons	1.07	1.03	0.99	1.02	.99
	(1.04)	(0.98)	(.84)	(.89)	(.91)
Avg. ratio of quits/employment spells	.32	.24	.19	.18	.16
	(.27)	(.22)	(.17)	(.17)	(.16)
Avg. number of employment spells	3.66	4.53	5.86	6.38	6.47
	(1.49)	(2.08)	(2.84)	(2.77)	(2.60)
Number of observations	205	324	366	527	411

Notes: "Avg." means Average; (1) 1906-19; (2) 1920-29; (3) 1930-39; (4) 1940-49; (5) 1950-59. **Source:** British Household Panel Survey.

Table 2. Family-related quit variables by marital status and number of children

		MARITAL STATUS		NUMBER OF CHILDREN			
		Never-married women	Have Been-married women	0	1	2	>=3
Have left job due to family reasons during first 10 years in work force (%)		16.15	72.44	16.46	67.10	74.66	80.25
Distribution of family-related quits							
	0	83.85	27.55	83.54	32.90	25.34	19.75
	1	12.25	47.91	12.29	49.57	50.73	49.06
	2	2.68	18.63	4.17	15.12	19.27	21.36
	3	1.22	4.55	-	2.41	3.56	7.40
	4	-	0.85	-	-	0.81	1.40
	5	-	0.32	-	-	0.29	0.56
	6	-	0.11	-	-	-	0.30
	7	-	-	-	-	-	0.18
	8	-	-	-	-	-	-
	9	-	0.07	-	-	-	-
Avg. number of quits due to family reasons *		0.21	1.05	0.21	0.87	1.04	1.26
		(0.55)	(0.92)	(0.49)	(0.75)	(0.84)	(1.02)
Avg. ratio of quits/employment spells *		0.04	0.21	0.04	0.20	0.21	0.24
		(0.11)	(0.19)	(0.11)	(0.21)	(0.19)	(0.19)
Number of observations		78	1755	169	292	720	652

Notes: *(std. dev. between brackets). **Source:** British Household Panel Survey.

Table 3. Prestige variable: Log(Camsis Scale)

	Coef,	t	Coef,	t	Coef,	t
Have left job due to family reasons (1=Yes)	-0.041	-1.670	-	-	-	-
Number of Quits	-	-	-0.051	-2.440	-	-
(Number of Quits)^2	-	-	-0.001	-0.290	-	-
Number of Employment Spells	-	-	0.141	10.880	-	-
(Number of Employment Spells)^2	-	-	-0.006	-7.760	-	-
Ratio Quits/Empl. Spells	-	-	-	-	-0.091	-0.720
(Ratio Quits/Empl. Spells)^2	-	-	-	-	-0.566	-3.250
Age at first spell	0.101	1.970	0.083	1.710	0.111	2.220
(Age at first spell)^2	-0.002	-1.420	-0.001	-1.040	-0.002	-1.630
White (1=Yes)	0.166	2.100	0.101	1.340	0.150	1.930
Birth Cohort 1906-1919	-0.402	-8.070	-0.281	-5.830	-0.331	-6.690
Birth Cohort 1920-1929	-0.130	-2.890	-0.069	-1.600	-0.104	-2.370
Birth Cohort 1940-1949	0.006	0.180	-0.012	-0.350	0.009	0.270
Birth Cohort 1950-1959	-0.023	-0.610	-0.051	-1.440	-0.021	-0.590
Higher Education	0.466	8.020	0.391	7.030	0.440	7.720
Teaching. nursing and other univ. ed.	0.319	9.120	0.273	8.150	0.307	8.950
GCE A level Education	0.171	3.110	0.155	2.950	0.157	2.910
GCE O level or equivalent	0.262	8.380	0.236	7.910	0.248	8.070
Vocational Training education	0.265	7.020	0.221	6.140	0.236	6.380
Currently above mandatory retirement age (1=Yes)	-0.048	-1.010	-0.021	-0.470	-0.034	-0.730
Proportion of time unemployed	-0.580	-2.410	-0.626	-2.730	-0.610	-2.590
Proportion of time inactive	-0.663	-1.070	-0.787	-1.330	-0.808	-1.330
One child	-0.119	-2.370	-0.079	-1.670	-0.096	-1.980
Two children	-0.221	-4.740	-0.173	-3.920	-0.194	-4.290
Three or more children	-0.404	-8.500	-0.351	-7.780	-0.367	-7.940
One union	-0.047	-0.670	-0.077	-1.150	-0.042	-0.610
Two or more unions	0.013	0.190	-0.027	-0.420	0.027	0.400
Constant	2.274	4.830	1.926	4.290	2.173	4.710
R^2	0.356		0.417		0.382	

Reference individual: Non-white; birth cohort 1930-39; no studies; below the mandatory retirement age (65 for men and 60 for women), no children, no union. **Note:** “union” refers to either a marriage or cohabitation. **Source:** British Household Panel Survey. **Number of observations:** 1,833

Table 4. Prestige variable: Log(Hope-Goldthorpe Scale)

	Coef,	t	Coef,	t	Coef,	t
Have left job due to family reasons (1=Yes)	-0.036	-1.410	-	-	-	-
Number of Quits	-	-	-0.049	-2.230	-	-
(Number of Quits)^2	-	-	0.000	0.000	-	-
Number of Employment Spells	-	-	0.145	10.540	-	-
(Number of Employment Spells)^2	-	-	-0.007	-7.980	-	-
Ratio Quits/Empl. Spells	-	-	-	-	-0.085	-0.640
(Ratio Quits/Empl. Spells)^2	-	-	-	-	-0.493	-2.680
Age at first spell	0.107	2.000	0.091	1.760	0.116	2.200
(Age at first spell)^2	-0.002	-1.630	-0.002	-1.290	-0.003	-1.800
White (1=Yes)	0.143	1.720	0.083	1.040	0.128	1.570
Birth Cohort 1906-1919	-0.450	-8.620	-0.335	-6.580	-0.387	-7.430
Birth Cohort 1920-1929	-0.176	-3.750	-0.120	-2.640	-0.154	-3.320
Birth Cohort 1940-1949	0.003	0.090	-0.014	-0.400	0.006	0.170
Birth Cohort 1950-1959	-0.047	-1.210	-0.075	-2.010	-0.046	-1.200
Higher Education	0.483	7.930	0.412	7.000	0.460	7.640
Teaching. nursing and other univ. ed.	0.331	9.040	0.288	8.140	0.321	8.870
GCE A level Education	0.116	2.000	0.102	1.830	0.103	1.810
GCE O level or equivalent	0.173	5.270	0.148	4.700	0.160	4.950
Vocational Training education	0.130	3.300	0.090	2.360	0.106	2.710
Currently above mandatory retirement age (1=Yes)	-0.018	-0.370	0.006	0.130	-0.006	-0.130
Proportion of time unemployed	-0.635	-2.520	-0.684	-2.820	-0.662	-2.660
Proportion of time inactive	-0.406	-0.620	-0.527	-0.840	-0.534	-0.830
One child	-0.154	-2.940	-0.116	-2.310	-0.134	-2.620
Two children	-0.252	-5.160	-0.207	-4.440	-0.228	-4.790
Three or more children	-0.454	-9.110	-0.404	-8.470	-0.420	-8.630
One union	-0.006	-0.080	-0.034	-0.480	-0.001	-0.020
Two or more unions	0.061	0.840	0.023	0.330	0.073	1.030
Constant	2.131	4.320	1.773	3.730	2.043	4.200
R^2	0.320		0.374		0.340	

Reference individual: Non-white; birth cohort 1930-39; no studies; below the mandatory retirement age (65 for men and 60 for women), no children, no union. **Note:** "union" refers to either a marriage or cohabitation. **Source:** British Household Panel Survey. **Number of observations:** 1,833

Table 5. Prestige variable: Log(Cambridge Scale)

	Coef,	t	Coef,	t	Coef,	T
Have left job due to family reasons (1=Yes)	-0.030	-1.020	-	-	-	-
Number of Quits	-	-	-0.046	-1.820	-	-
(Number of Quits)^2	-	-	-0.003	-0.480	-	-
Number of Employment Spells	-	-	0.145	9.220	-	-
(Number of Employment Spells)^2	-	-	-0.007	-6.600		-
Ratio Quits/Empl. Spells	-	-	-	-	-0.023	-0.150
(Ratio Quits/Empl. Spells)^2	-	-	-	-	-0.664	-3.170
Age at first spell	0.165	2.700	0.147	2.500	0.177	2.930
(Age at first spell)^2	-0.003	-1.950	-0.003	-1.640	-0.004	-2.140
White (1=Yes)	0.320	3.380	0.254	2.770	0.303	3.250
Birth Cohort 1906-1919	-0.433	-7.270	-0.310	-5.300	-0.361	-6.070
Birth Cohort 1920-1929	-0.137	-2.560	-0.075	-1.430	-0.111	-2.100
Birth Cohort 1940-1949	-0.003	-0.060	-0.021	-0.500	0.001	0.020
Birth Cohort 1950-1959	-0.025	-0.560	-0.053	-1.240	-0.023	-0.530
Higher Education	0.680	9.790	0.603	8.940	0.654	9.540
Teaching. nursing and other univ. ed.	0.432	10.330	0.385	9.480	0.420	10.190
GCE A level Education	0.261	3.960	0.244	3.830	0.247	3.800
GCE O level or equivalent	0.352	9.410	0.325	8.990	0.337	9.140
Vocational Training education	0.375	8.320	0.331	7.590	0.347	7.780
Currently above mandatory retirement age (1=Yes)	-0.030	-0.530	-0.002	-0.040	-0.016	-0.280
Proportion of time unemployed	-0.925	-3.210	-0.977	-3.520	-0.955	-3.370
Proportion of time inactive	-0.886	-1.190	-1.022	-1.420	-1.035	-1.410
One child	-0.085	-1.420	-0.041	-0.720	-0.061	-1.040
Two children	-0.229	-4.110	-0.176	-3.300	-0.201	-3.690
Three or more children	-0.402	-7.080	-0.343	-6.280	-0.363	-6.540
One union	-0.023	-0.270	-0.053	-0.650	-0.018	-0.210
Two or more unions	0.044	0.540	0.004	0.050	0.059	0.730
Constant	0.890	1.580	0.530	0.970	0.782	1.410
R^2	0.374		0.418		0.392	

Reference individual: Non- white; birth cohort 1930-39; no studies; below the mandatory retirement age (65 for men and 60 for women), no children, no union. **Note:** “union” refers to either a marriage or cohabitation. **Source:** British Household Panel Survey. **Number of observations:** 1,833

Table 6. Log prestige equations (within-group technique)

	CAMSIS				HGS				CAMBRIDGE			
	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio
Family-related quit in previous job	-0.016	-2.190	-	-	-0.028	-2.930	-	-	-0.017	-1.230	-	-
Accumulated number of quits	-	-	-0.005	-0.610	-	-	-0.024	-1.930	-	-	-0.037	-2.100
(Accumulated number of quits)^2	-	-	0.005	2.110	-	-	0.003	0.910	-	-	0.016	3.640
Tenure previous job												
<=2 years	-	-	-	-	-	-	-	-	-	-	-	-
>2 & <=4 years	0.013	2.340	0.013	2.300	0.007	0.850	0.006	0.780	0.014	1.210	0.013	1.150
>4 & <=6 years	0.030	4.350	0.030	4.330	0.010	1.010	0.008	0.870	0.045	3.310	0.045	3.250
>6 years	0.011	1.970	0.012	2.080	0.026	3.350	0.026	3.340	0.015	1.320	0.015	1.330
Tenure current job												
<=2 years	-	-	-	-	-	-	-	-	-	-	-	-
>2 & <=4 years	0.017	3.050	0.017	2.950	0.020	2.610	0.020	2.600	0.030	2.750	0.029	2.620
>4 & <=6 years	0.023	3.080	0.022	2.980	0.011	1.080	0.011	1.130	0.033	2.300	0.031	2.170
>6 years	0.031	5.150	0.029	4.810	0.037	4.620	0.040	4.840	0.051	4.390	0.049	4.130
Non-employment duration												
<=1 month	-	-	-	-	-	-	-	-	-	-	-	-
>1 & <=6 months	-0.022	-2.230	-0.025	-2.550	-0.044	-3.320	-0.048	-3.710	-0.029	-1.530	-0.031	-1.650
>6 & <= 18 months	-0.026	-3.130	-0.031	-3.870	-0.040	-3.550	-0.048	-4.470	-0.035	-2.120	-0.036	-2.330
>18 months	-0.013	-1.900	-0.022	-3.910	-0.034	-3.610	-0.048	-6.300	-0.034	-2.530	-0.042	-3.810
Age current job												
<=35 years-old	-	-	-	-	-	-	-	-	-	-	-	-
>35 & <= 45 years-old	0.029	4.440	0.029	4.420	0.048	5.470	0.054	6.110	0.046	3.630	0.047	3.670
> 45 years-old	0.034	3.520	0.032	3.340	0.059	4.590	0.069	5.240	0.071	3.830	0.070	3.690
Number of unions												
No union	-	-	-	-	-	-	-	-	-	-	-	-
One union	0.008	1.070	0.007	0.950	0.001	0.150	0.001	0.080	0.016	1.130	0.020	1.420
Two or more unions	0.033	2.380	0.031	2.230	0.028	1.490	0.030	1.590	0.049	1.810	0.050	1.850
Number of children												
No children	-	-	-	-	-	-	-	-	-	-	-	-
One child	-0.009	-1.120	-0.011	-1.230	-0.025	-2.200	-0.022	-1.860	-0.008	-0.520	-0.004	-0.210
Two children	-0.032	-3.720	-0.034	-3.760	-0.047	-4.120	-0.039	-3.210	-0.053	-3.160	-0.049	-2.770
Three or more children	-0.032	-2.850	-0.034	-2.950	-0.049	-3.270	-0.039	-2.500	-0.039	-1.820	-0.036	-1.600
Constant	3.804	348.460	3.807	347.420	3.654	247.500	3.653	246.350	3.307	154.760	3.312	154.500

Notes: regressions control for individual fixed effects, as well as for three different temporary periods (up to the year 1950. from 1950 to 1975, beyond 1975). **Source:** British Household Panel Survey. **Number of observations:** 9870.

APPENDIX A. Descriptive statistics

Table A.1. Total sample, women who leave the workforce due to family reasons, and women who do not (OLS analysis)

	WHOLE SAMPLE		WOMEN WHO QUIT		WOMEN WHO DO NOT QUIT	
	Mean	***Std. Dev.***	***Mean***	***Std. Dev.***	***Mean***	***Std. Dev.***
Have left job due to family reasons	0.701	0.458	1.000	0.000	0.000	0.000
Number of Quits	1.016	0.920	1.450	0.760	0.000	0.000
Number of Quits (1st 10 years in labour force)	0.901	0.750	1.286	0.555	0.000	0.000
Number of Employment Spells	5.667	2.703	5.816	2.710	5.319	2.657
Ratio Quits/Empl. Spells	0.205	0.197	0.292	0.172	0.000	0.000
Ratio Quits/Empl. Spells (1st 10 years in labour force)	0.390	0.342	0.557	0.272	0.000	0.000
Age at first spell	16.269	2.140	16.141	1.955	16.571	2.498
White (1=Yes)	0.982	0.132	0.986	0.118	0.974	0.161
Birth Cohort 1906-1919	0.112	0.315	0.110	0.313	0.116	0.321
Birth Cohort 1920-1929	0.176	0.381	0.181	0.385	0.165	0.372
Birth Cohort 1940-1949	0.287	0.453	0.291	0.454	0.279	0.449
Birth Cohort 1950-1959	0.224	0.417	0.218	0.413	0.240	0.427
Higher and First Degree Education	0.051	0.221	0.040	0.196	0.079	0.269
Teaching. nursing and other univ. ed.	0.162	0.369	0.145	0.353	0.202	0.402
GCE A level Education	0.041	0.199	0.040	0.195	0.046	0.209
GCE O level or equivalent	0.187	0.390	0.196	0.397	0.167	0.373
Vocational Training education	0.092	0.289	0.100	0.300	0.074	0.262
Currently above mandatory retirement age (1=Yes)	0.366	0.482	0.376	0.485	0.342	0.475
Proportion of time unemployed	0.010	0.043	0.007	0.033	0.016	0.060
Proportion of time spent inactive	0.002	0.017	0.002	0.015	0.004	0.021
Average HGS occupational prestige	25.856	12.588	24.432	11.314	29.189	14.637
Average Cambridge occupational prestige	21.415	12.403	20.218	11.125	24.218	14.607
Average Camsis occupational prestige	31.214	14.442	29.627	13.128	34.931	16.557
No children	0.092	0.289	0.022	0.146	0.257	0.438
One child	0.160	0.366	0.153	0.360	0.176	0.381
Two children	0.393	0.489	0.419	0.494	0.333	0.472
Three or more children	0.355	0.479	0.407	0.491	0.234	0.424
No union	0.034	0.182	0.005	0.071	0.103	0.305
One union	0.223	0.417	0.226	0.418	0.217	0.413
Two or more unions	0.742	0.437	0.769	0.422	0.679	0.467
Number of observations	1833		1284		548	

Source: British Household Panel Survey. **Note:** “union” refers to either a marriage or cohabitation.

Table A.2. Total sample, women who leave the workforce due to family reasons, and women who do not (within-group analysis)

	WHOLE SAMPLE		WOMEN WHO QUIT		WOMEN WHO DO NOT QUIT	
	Mean	***Std. Dev.***	***Mean***	***Std. Dev.***	***Mean***	***Std. Dev.***
Family-related quit in previous job	0.224	0.417	1.000	0.000	0.000	0.000
Number of accumulated quits due to family reasons	0.714	0.835	1.379	0.721	0.522	0.765
Tenure previous job:						
<=2 years	0.436	0.496	0.325	0.469	0.468	0.499
>2 & <=4 years	0.199	0.399	0.239	0.426	0.188	0.391
>4 & <=6 years	0.119	0.324	0.160	0.366	0.108	0.310
>6 years	0.245	0.430	0.276	0.447	0.237	0.425
Non-employment duration						
<=1 month	0.616	0.486	0.084	0.277	0.770	0.421
>1 & <=6 months	0.055	0.227	0.062	0.240	0.053	0.224
>6 & <= 18 months	0.086	0.280	0.164	0.370	0.063	0.244
>18 months	0.243	0.429	0.691	0.462	0.114	0.318
Tenure current job						
<=2 years	0.447	0.497	0.368	0.482	0.469	0.499
>2 & <=4 years	0.202	0.401	0.193	0.394	0.204	0.403
>4 & <=6 years	0.106	0.308	0.108	0.310	0.106	0.308
>6 years	0.246	0.430	0.332	0.471	0.221	0.415
Age current job						
<=35 years-old	0.565	0.496	0.690	0.463	0.530	0.499
>35 & <= 45 years-old	0.272	0.445	0.235	0.424	0.283	0.450
> 45 years-old	0.163	0.369	0.076	0.264	0.188	0.391
Number of unions						
No unions	0.409	0.492	0.265	0.442	0.451	0.498
One union	0.539	0.498	0.693	0.461	0.495	0.500
Two or more unions	0.052	0.222	0.041	0.199	0.055	0.228
Number of children						
No children	0.509	0.500	0.301	0.459	0.569	0.495
One child	0.165	0.371	0.238	0.426	0.144	0.351
Two children	0.222	0.416	0.298	0.458	0.200	0.400
Three or more children	0.104	0.305	0.162	0.369	0.087	0.282
Number of observations	9870		2211		7659	

Source: British Household Panel Survey. **Note:** “union” refers to either a marriage or cohabitation.

APPENDIX B. Endogeneity analysis

Table B.1.: Estimated coefficients for family-quit variables. Log prestige OLS equations

	CAMSIS		HGS		CAMBRIDGE	
DEPENDENT VARIABLE :	Coef.	T-ratio	Coef.	T-ratio	Coef.	T-ratio
Log(prestige in first job held)						
Have left job due to family reasons (1=Yes)	-0.018	-1.170	-0.014	-1.010	0.011	0.410
Number of Quits	-0.013	-0.890	-0.013	-0.970	0.004	0.170
(Number of Quits)^2	-0.001	-0.310	0.000	0.090	-0.003	-0.450
Ratio Quits/Empl. Spells	-0.129	-1.550	-0.115	-1.510	-0.064	-0.450
(Ratio Quits/Empl. Spells)^2	0.047	0.400	0.095	0.890	-0.032	-0.160
Number of observations	1823		1815		1823	
Log(Average prestige in occupations previous to first quit)						
Have left job due to family reasons (1=Yes)	-0.304	-10.060	-0.315	-10.530	-0.304	-8.210
Number of Quits	-0.200	-7.110	-0.211	-7.570	-0.206	-6.020
(Number of Quits)^2	0.048	6.720	0.052	7.430	0.049	5.680
Ratio Quits/Empl. Spells	-1.424	-8.770	-1.490	-9.260	-1.496	-7.530
(Ratio Quits/Empl. Spells)^2	1.416	6.150	1.563	6.850	1.453	5.160
Number of observations	1823		1820		1823	
Log (Prestige for the occupation inmediately previous to first quit)						
Have left job due to family reasons (1=Yes)	-0.067	-4.120	-0.090	-4,690	-0,105	-3,690
Number of Quits	-0.062	-4.140	-0.085	-4,830	-0,111	-4,270
(Number of Quits)^2	0.008	2.120	0.011	2,410	0,017	2,530
Ratio Quits/Empl. Spells	-0.371	-4.270	-0.566	-5,540	-0,601	-3,950
(Ratio Quits/Empl. Spells)^2	0.262	2.120	0.540	3,730	0,438	2,030
Number of observations	1823		1822		1823	

Notes: regressions control for individual fixed effects, as well as for tenure in the previous job, non-employment duration, age at the current job, number of unions, number of children and three different temporary periods (up to the year 1950. from 1950 to 1975, beyond 1975). **Source:** British Household Panel Survey.

APPENDIX C. Definitions and characteristics of the occupational prestige scores used in the article

1. Hope-Goldthorpe scale score

The *Hope-Goldthorpe Scale* (HGS) score was derived from a survey on the social standing of occupations, whereby a ranking of occupations was made by a random sample of individuals interviewed throughout England and Wales in 1972. Although the HGS score is based on a survey launched in 1972 and our data cover the XX century, we want to remark that Hauser and Featherman (1977) have shown that there is great stability over time in occupational prestige: since the year 1925 the structure of occupational prestige has remained almost constant in Western countries, which is specially useful for our research (note that the oldest employment histories of the BHPS began around 1920). Furthermore, it turns out that in Britain the position of individuals in the occupational hierarchy is relatively stable over time. Therefore, the HGS score is also an adequate measure of people's permanent socio-economic status.

Like virtually all other stratification measures, this score uses occupational groups as its basic units. The most important underlying assumption of the HGS score is that the social prestige of an occupation is based on various dimensions such as the living conditions it provides, the necessary knowledge it requires, the income earned in each occupation, and its social usefulness (see Stewart et al., 1980: 21-27, for the details about the construction of this score). Goldthorpe and Hope (1974) suggest that the scale which results from their occupational prestige grading exercise should be viewed as "a judgement which is indicative of what might be called the 'general goodness' or … the 'general desirability' of occupations" (p. 11-12).

This scale is included in the original BHPS data base in each wave and in all employment spells of the individuals' employment histories. The minimum (value 0) was set up for domestic housekeepers and related occupations[14]. Individuals were asked to assign numerical values to the remainder of occupations. The maximum corresponds to medical practitioners. It is widely documented that this score is highly correlated with earnings. Using data from the British New Earnings Survey, Phelps Brown (1977) reports a strong relationship between median gross weekly earnings by occupation and the HGS score, with a rise of 1 unit in the scale of occupational status being associated with an increase of 1.031 percent in earnings. Nickell (1982) also reports a correlation between the HGS score and the average hourly earnings by occupation of 0.85 using data from the National Training Survey. Thus, to the extent that labour income represents a

[14] The information about the Standard Occupational Classification in the BHPS has been obtained from Taylor et al. (2001). We use the 1990 version of the UK SOC, and not the SOC 2000. We use the coded information provided by the survey. Therefore, all occupations along the life course are coded using the SOC 1990. The use of an occupational classification closer to the time when information was collected minimises the problem of how consider new occupations. However, it does not consider that some occupations have dramatically changed along the XX century. This is a limitation inherent to any research using this data base.

substantial fraction of total income, the HGS score is likely to be a good measure of people's socio-economic position.

This prestige scale has been used before in topics closely related to Industrial Relations to measure the labour market success of individuals —Bond and Saunders (1999)— and to analyse the risk of fatal injury —Marin and Psacharopoulos (1982). The latter authors find that the risk of fatal injury presented a clear negative effect on the occupational prestige. In this sense, the HGS is related to the desirability of different occupations.

2. Cambridge scale score

The Cambridge scale score resulted from the work of the Cambridge stratification group (Prandy 1990; Stewart et al. 1973; Blackburn and Stewart, 1975), which used a variety of close social relationships to investigate social proximity and distance. While the HGS score asked individuals to evaluate the social desirability of occupations in general, the Cambridge scale score is based on 'the occupations of persons with whom their incumbents interact' (Stewart et al. 1980: 28). The current version of this score uses friendship and marriage patterns as the basis for evaluating the occupations. That is, people do not evaluate occupations in general, but in terms of occupations held by their friends and spouses. The score assumes that those with similar lifestyles and resources tend to interact more with one another in terms of both friendship choices and intermarriage. Therefore, the relative social distances between people in different occupations reflect dissimilarities in lifestyles and resources and hence social inequality (Prandy 1990: 635). The minimum in this scale score corresponds to "glass products and ceramics makers", while the maximum corresponds to "other social and behavioural scientists".

As a measure of stratification arrangements or "generalized advantage of lifestyle", the scale has been used to look at the impact of social distance on educational outcomes (Blackburn and Marsh 1991), ethnic inequality (Blackburn et al. 1997) and occupational segregation by gender (Blackburn et al. 1999).

3. Camsis scale score

The idea behind the Camsis scale score is that social interaction will occur more frequently between persons who are socially close to each other and will be rarer between those who are socially distant. Thus, acquaintances, friends and marriage partners will all tend to be chosen much more frequently from within the same group than from without. The Camsis Scale is part of a wider project about an internationally comparative assessment of the structures of social interaction and stratification across a number of countries. Detailed information on the CAMSIS (Cambridge Social Interaction and Stratification) project can be found in the following address: http://www.cf.ac.uk/socsi/CAMSIS)

Prandy and Lambert (2003) discuss the development of the Camsis score for the UK, showing that it is very closely comparable to the Cambridge score. One major difference is that the Camsis one has been constructed solely on the basis of marriage patterns. An advantage of using exclusively marriage data is that they can be derived from censuses or very large-scale official surveys. Another difference is that they are to a substantial degree directly comparable across countries. This combination of being nationally (and even time-period) specific yet directly comparable is a major advantage of Camsis scales (Prandy and Jones, 2001).

Since the Camsis score is derived within the context of gender groupings, different scores are obtained for men and women. Thus, for instance, there is no necessary relationship between the values of an occupation on its male and female scores (although they are likely to share similar relative locations). The minimum value in this scale is assigned to “glass and ceramics, furnace operatives”, while the maximum is achieved for “university and polytechnic teaching professionals”. To sum up, the Camsis score evaluates the social positions of occupations held by spouses by explicitly considering the gender of the person who held each occupation.

Figure 1. Frequencies of quits by type

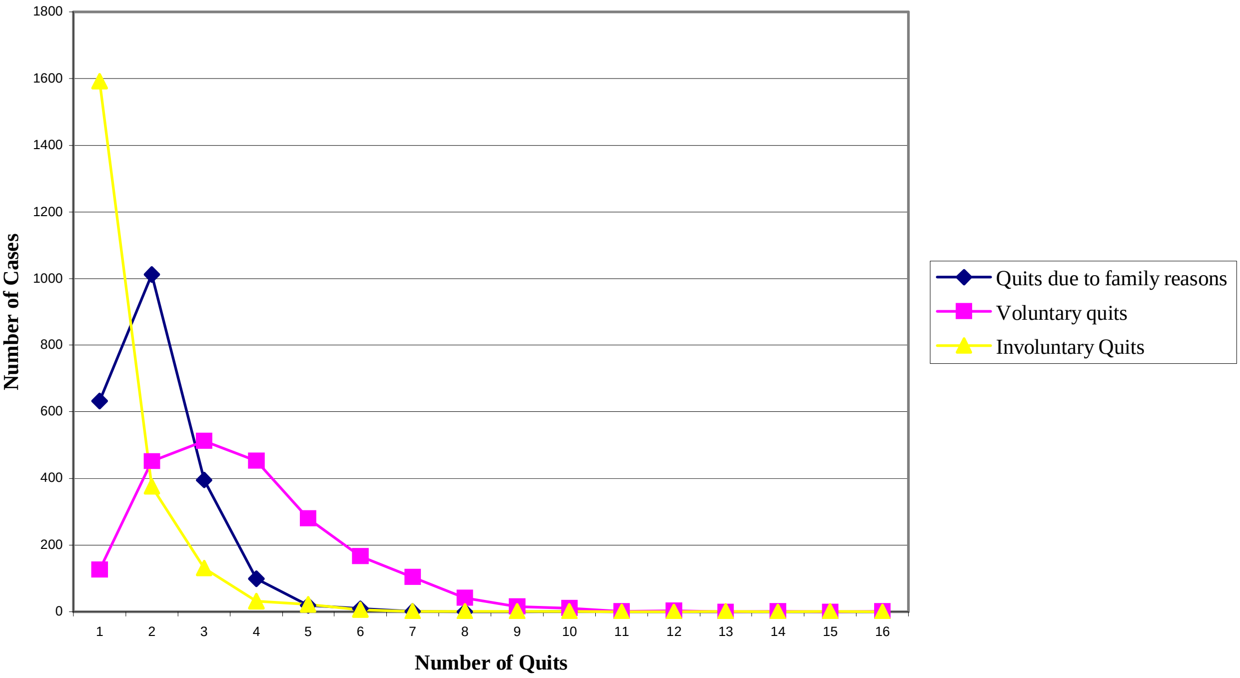

Figure 2. Average Camsis Prestige Score by Employment Spells
Note: 'Do not quit' and 'Quit' refer to quits during the first 10 years of labour experience