Occupational Sex-Segregation, Specialised Human Capital and Wages: Evidence from Britain

Introduction

The gender wage gap is one of the most enduring features of the labour market (Olsen and Walby, 2004). However, there is still ample debate about the mechanisms which create and perpetuate it. In this article, panel data and panel data methods are used to assess the impact of the sex-composition of occupations on the wages of men and women and on the gender wage gap, building upon different theoretical perspectives.

The literature documents extensive and pervasive occupational segregation between men and women within modern labour markets (Charles and Grusky, 2004; Rubery and Fagan, 1995). There is also general agreement that working in an occupation in which a large proportion of workers are women incurs a wage penalty (Petersen and Morgan, 1995), as well as lower prestige (Magnusson, 2009), worse working conditions (Glass, 1990) and slower career progression (Petersen and Saporta, 2004). From a sociological standpoint, devaluation theory argues that work in female-dominated occupations is undervalued due to institutionalised bias against women, with wages being lower in such occupations as a result (England, 1992). In contrast, the specialised human capital approach, advocated chiefly by labour economists, suggests that the gender composition of an occupation has no direct effect on wages, but that the low wages observed in female-dominated occupations are caused by their lack of skill specialisation (Tam, 1997).

This article makes three contributions to the literature. First, it examines in detail the relationship between the proportion of workers in individuals' occupations (i.e. occupational feminisation) and wages in Britain using panel data and evaluates whether this is linear or non-linear. Second, following the work of Tam (1997) on the US, it examines the role of specialised human capital in explaining the association between occupational feminisation and wages in Britain, which has yet to be investigated. Third, it uses regression models which allow controlling for individual-specific unobserved heterogeneity (i.e. person-level unique wage differentials). Results indicate that workers in female-dominated occupations receive lower wages than those in other occupations after controlling for skill specialisation and allowing for individual-specific unobserved heterogeneity and that a sizeable portion of the gender gap in wages is due to occupational sex-segregation. These findings support sociological notions which attribute gender pay differentials to the undervaluation of 'women's work'.

Explaining the impact of occupational sex-segregation on wages: devaluation or skill specialisation?

A number of theories are relevant in explaining the effect of the sex-composition of occupations on wages and its contribution to the gender wage gap. The focus of this research is on two competing explanations which have received support in empirical research, the devaluation and human capital (HC) explanations.

There is ample sociological evidence that the distribution of power between sexes in society is not balanced and men have historically dominated almost all spheres of social life (Walby, 1986). Patriarchal systems generate sex-bias in the social construction of value, which also operates in relation to the work performed by men and women. Ideology and tradition play a role in defining which skills are valuable, desirable and profitable within a given society. As McGrath and DeFilippis (2009) put it *"some of the most critical work in society is often the least rewarded"* and *"this is ideologically justified in gendered [...] ways"* (p.68). A higher value is attributed to jobs or occupations typically carried out by men or associated with male-stereotyped skills and so discrimination does not take place against individuals but against the types of jobs that they do (Maume, 1999). Based on these arguments, devaluation theory suggests that male-dominated occupations are more highly-rewarded than female-dominated occupations because 'women's work' is devalued by social structures. As argued by comparable-worth theories, the rewards systems in male- and female-dominated occupations operate differently and occupations which require comparable – though different – skills are paid worse if they are female rather than male dominated (England, 1992; Grimshaw and Rubery, 2007). For example, market work which involves caring and nurturing skills or emotional labour is usually undervalued (Bolton and Muzio, 2008; England, 1992). It has been argued that, where some tasks are usually performed as unpaid labour in the domestic sphere, similar work undertaken for pay tends to be devalued (McGrath and DeFilippis, 2009). Findlay et al. (2009) provide an example for Britain using the occupation 'nursery nurses': while the work requires considerable effort, knowledge, initiative, communication skills and relationship management, it is poorly remunerated because traditionally-female skills are poorly weighted in job evaluation schemes. Even when women manage to permeate male-dominated occupations, they concentrate into a narrow range of 'female-friendly' occupational branches with comparable skill requirements but less favourable career ladders, working conditions and remuneration packages - see Bolton and Muzio (2008) for recent case studies for the law, teaching and management professions.

HC theories are the most common economic explanations for differences in pay received by workers. HC is defined as the stock of knowledge and skills accumulated by an individual and is acquired through education, training and experience. Becker's work effort/rational choice theory (1985) and Polachek's (1981) theory of occupational self-selection are relevant when discussing wage differences between men and women. Becker's well-known argumentation suggests that, because women are naturally prepared to bear children, men have higher incentives to specialise in market work while women are predisposed to specialise in non-market work. Thus, women allocate fewer resources in acquiring and updating their HC, tend to work in occupations with lower skill requirements and consequently earn less than men. Similarly, Polachek suggests that women are more likely than men to interrupt their work and careers due to household commitments, which results in less labour market experience, forgone training and skill depreciation. To maximise their lifetime earnings, women choose to work in occupations in which work arrangements are flexible, starting wages are highest, depreciation rates are lowest and wages are less dependent on experience, at the expense of receiving lower pay than men in the long run.[1] From a sociological standpoint, Hakim (2000) argues that, while women's circumstances are becoming progressively more heterogeneous, a fair share of women still display preferences for traditional work-life arrangements.

Specialised human capital (SHC) theories evolve directly from the above. Although the concept of SHC has been discussed for a long time (see Parsons, 1972 or Jovanovic, 1979), Tam (1997) was the first to apply this within the literature linking occupational sex-segregation and wages. Unlike general HC, investments in SHC are occupation, industry, or firm-specific and have little or no economic value outside the setting in which they were acquired. The expectation of career breaks and the higher opportunity costs of training due to the unequal distribution of non-market work lead women to avoid jobs which require large amounts of SHC and invest instead in more 'portable' skills. At the macro-level, findings from Estevez-Abe (2005) support these arguments by showing that countries with 'national skill regimes' which lean towards specialised rather than general skills have highly sex-segregated labour markets. Highly specialised jobs are risky for both the employer (who bears additional training costs) and the employee (who forgoes the possibility to apply the relevant skills in other contexts). Thus, both workers and employers have incentives to maintain the employment relationship. To prevent highly specialised workers from leaving their jobs, firms may offer long-term contracts with upward sloping wage-tenure profiles (Polavieja, 2007). Therefore, at high levels of tenure individuals working in highly specialised jobs receive wages which are comparatively higher than those offered by other jobs. As a result, the SHC approach pay differences attributed to occupational sex-segregation may actually be due to differences in skill specialisation across occupations.

Empirically, the impact of occupational feminisation on wages is captured through the inclusion of an indicator of the proportion of workers in an occupation who are women in a wage equation. Commonly, results are reported as the predicted wage differences between individuals working in occupations in which all of workers are men and individuals working in occupations in which all of workers are women. Most existing research focuses on the US and uses cross-sectional regression data and methods. Studies can be divided into two groups: those which do not incorporate SHC and those which incorporate SHC. The common finding in studies which do not control for SHC is that working in feminised occupations reduces the wages of men and women. Men working in fully female-dominated occupations earn wages which are between 7 percent (England et al., 1988) and 26 percent (Cotter et al., 1997) lower than those of men working in fully male-dominated occupations, while the effects range between 4 percent (Gerhart and El Cheikh, 1991) and 42 percent (US Bureau of the Census, 1987) for women.

Recent studies have incorporated SHC to models of the impact of occupational feminisation on wages. These use data from the US (Tam, 1997; Tomaskovic-Devey and Skaggs, 2002), a group of European countries (Polavieja, 2007, 2009) and Spain (Polavieja, 2008) and different ways to operationalise SHC. Tam (1997) uses the specific vocational preparation required for each occupation imputed from the American Dictionary of Occupational Titles (DOT) while the other studies use self-

[1] It must be noted that HC theories have been heavily questioned – see for example England (1982, 1984), but it is beyond the scope of this article to present a comprehensive critique.

reported job-learning time. Results show that introducing measures of SHC into wage equations reduces the negative impact of occupational feminisation on wages, which in these studies is no longer statistically different from zero.

However, results from studies including measures of SHC are questionable. None of them allow for the potentially biasing effects of unobserved heterogeneity, which if correlated with the variables of interest can bias their associated coefficients. The indicator of SHC used by Tam (1997) comes from the DOT, which has been criticised for underestimating the skills required to do jobs mainly performed by women (Phillips and Taylor, 1980; Steinberg, 1990) and the effect of occupational feminisation on wages returns when adding a control for general educational development (England et al., 2000). Tomaskovic-Devey and Skaggs (2002) use the proportion of women in the job (rather than in the occupation) and a very small sample (700 individuals) which is not nationally-representative. The measures of occupational sex-composition in Polavieja (2008, 2009) are based on a very small number of workers, while Polavieja (2007) pools all available European countries and makes the strong assumption that the sex-composition of each occupation is equal across nations. Furthermore, none of these studies focuses on Britain. This is important, as recent research has demonstrated that the gender composition of occupations has substantial consequences for the welfare of British workers and that wage growth is slower in female-dominated occupations (Dex et al., 2008; Olsen and Walby, 2004).

All available studies constrain the relationship between occupational sex-segregation and wages to be linear. This is a strong assumption, which if violated may lead to misleading results. For example, it is well-known that public sector jobs and certain professions which offer relatively high wages are (or are becoming) female dominated, while many poorly-paid blue-collar jobs are held by men. Furthermore, theoretical and empirical research suggests that the impact of the sex-composition of *jobs* and *workplaces* on wages may be non-linear (Kanter, 1977; Pfeffer and Davis-Blake, 1987; Tomaskovic-Devey, 1993).

This research adds to the literature by focusing in Britain; evaluating whether the relationship between occupational feminisation and wages is linear or non-linear; and using panel data models which allow controlling for individual-specific unobserved heterogeneity.

Data and methods

Three different nationally-representative datasets are used: the British Household Panel Survey (BHPS), the UK Labour Force Survey (LFS) and the Skills Surveys (SS). Analyses presented here use the first seventeen waves of the BHPS covering the period 1991-2007 and are based on a sample of employees within statutory working age and not in full-time education. The resultant sample size is of 3,968 men (26,365 observations) and 4,359 women (29,446 observations). To maximise sample sizes, LFS quarters are pooled into annual files to construct occupation-level variables which are then matched to BHPS respondents by their standard occupational classification (SOC) code and year. Due to the change from SOC90 to SOC2000 in LFS data, the proportion of workers who are women in each respondent's occupation is calculated using occupational units from SOC90 (n=371) from 1991 to 2000 and occupational units from SOC2000 (n=353) from

2001 to 2007. The third data source is the 2001 and 2006 SS. These are cross-sectional surveys of up to 8,000 individuals which provide data on skill and job requirements in the British labour market. These are pooled to derive two occupation-level measures of SHC at the minor occupational group (n=81) using SOC2000.

The impact of occupational feminisation on individuals' wages is first modelled using standard linear regression (ordinary least squares, OLS). The dependent variable is the natural log of respondent's gross hourly wages, deflated to 2007 prices using Consumer Price Indices. Additionally, the top and bottom 1 percent of the wage distribution have been dropped to exclude outliers. The key independent variable is the proportion of employees who are women in each respondent's occupation. Models control for a set of observable time-varying individual-, job-, establishment- and occupation-level variables known to affect wages similar to those used in Tam (1997), including year, region, age, marital status, education, establishment size, contract type, hours of work, job tenure, sector and industry.

However, survey data rarely contain all desirable information on individuals' characteristics necessary to establish causal relationships. Thus, there may be unobserved individual-specific characteristics which correlate both with wages and occupational sex-segregation that are missing from the data. For example, individuals may possess different unmeasured productivity-related factors such as unobserved skill which affect their wages (England et al., 1988) as well as tastes, preferences and motivations which affect their choice of occupation (Hakim, 2000). Such individual-specific factors are important because, if not suitably allowed for in statistical analyses, they can bias model estimates. Fortunately, panel data allow controlling for these unobserved time-invariant individual-specific effects. In particular, within-group fixed effects (FE) regression models are well-suited to understand the relationships between gender segregation and wages. Instead of relying on comparison *between* individuals, FE models base their predictions on comparisons of the same individuals over time (Allison, 2009) – that is, *within* individuals. Consequently, their predictions *"are not contaminated with spurious effects of any stable, unmeasured individual characteristics"* such as *"cohort, socioeconomic background [...] unchanging aspects of intelligence, preferences resulting from early socialisation, life cycle plans, and unmeasured human capital"* (England et al., 1988, p.548). Accounting for unobserved heterogeneity is particularly important for the aims of this article, since both observable and unobservable characteristics should be controlled for to be able to make claims of devaluation.

FE models are therefore estimated as complements to OLS models to evaluate the extent to which the impact of occupational feminisation on wages is explained by time-invariant individual-specific unobserved traits. To allow the wage effects of all covariates to differ by sex gender-specific models are estimated.[2]

[2] The OLS model can be represented as:

$$\log(W)_{it} = F_{it}\,\beta_1 + X_{it}\,\beta_2 + Z_i\beta_3 + v_{it}$$

where the i and t subscripts designate individual and time respectively, log(W) represents logged hourly wages, F is an indicator of occupational feminisation, X is a vector of observable time-

The impact of occupational feminisation on wages in Britain

Figure 1 provides preliminary descriptive evidence of the relationship between occupational sex-segregation and wages by plotting average wages of individuals by the proportion of workers in their occupation who are women. The graph shows a similar concave relationship for men and women. Wages are highest in more gender-integrated occupations than in sex-segregated occupations and male-dominated occupations pay more than female-dominated occupations. These results suggest that there may be a negative non-linear relationship between occupational sex-segregation and wages.

[FIGURE 1 HERE]

Table 1 presents estimates of the impact of occupational feminisation on wages for men and women from multivariate regression models which include the control variables. The estimated coefficients on occupational feminisation from OLS models are –0.171 and –0.321 for men and women respectively. These are statistically different from zero and indicate that working in a completely female-dominated occupation is associated with wages 17 percent and 32 percent lower relative to working in a completely male-dominated occupation. Since average wages are £12.33 for men and £9.53 for women, this is equivalent to £2.11 per hour for men and £3.06 per hour for women. Thus, occupational feminisation is not only strongly and negatively associated with wages but its effects are also larger for women. Furthermore, women earn 13.4 percent lower wages than otherwise similar men (results not shown), which equates to £1.42 per hour, everything else – including occupational feminisation – being equal. When unobserved heterogeneity is allowed for in FE models, the wage penalty associated with occupational feminisation is substantially lower. Working in a fully female-dominated occupation is associated with wages 12.8 percent (£1.58) and 16.7 percent (£1.60) lower relative to working in a fully male-dominated occupation for men and women respectively. This suggests that unobserved characteristics of individuals (e.g. ability, motivation and preferences) play an important role in allocating workers within the occupational feminisation distribution and that individuals with unmeasured characteristics positively associated with wages (i.e.

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varying variables and Z is a vector of observable time-invariant characteristics The error term v_{it} can be decomposed into

$$v_{it} = \upsilon_i + \varepsilon_{it}$$

where υ represents individual-specific time-constant unobservable characteristics and ε is a random error. The FE model is estimated by taking deviations from individual-specific means over time in dependent and explanatory variables:

$$\log(W)_{it} - \overline{\log(W)}_i = (F_{it} - \overline{F}_i)\beta_1 + (X_{it} - \overline{X}_i)\beta_2 + (\varepsilon_{it} - \overline{\varepsilon}_i).$$

This removes unobserved time-invariant characteristics (υ_i) and allows for arbitrary correlation between observable and unobservable factors.

workers who are more 'able' for reasons unobserved in the data) tend to work in less feminised occupations.

[TABLE 1 HERE]

The above models assume that the relationship between occupational feminisation and wages is linear, but descriptives in Figure 1 suggest that this is not the case. Therefore, an additional specification which allows for non-linearities by adding a quadratic and a cubic term of occupational feminisation is estimated. Results are presented in the form of a graph in which predicted wages for representative male and female employees are plotted as a function of occupational feminisation (Figure 2). For men, the OLS model reveals a concave relationship between occupational feminisation and wages. Segregated occupations pay less than the linear prediction while integrated occupations pay more, with predicted wages being highest in occupations in which 28 percent of employees are women. The FE model again suggests a non-linear relationship, but this is less pronounced than that observed in the OLS model. For women, the OLS model also suggests a non-linear relationship: wages increase until 25 percent of workers in the occupation are women and then fall. However, non-linearities are less pronounced than for men. As for men, estimates from FE models for women show less evidence of non-linearities than those from OLS models.

[FIGURE 2 HERE]

Altogether, results suggest that the relationship between occupational feminisation and wages is virtually linear once time-invariant individual-specific unobserved heterogeneity is allowed for.

Accounting for specialised human capital

This section examines the extent to which SHC mediates the impact of occupational feminisation on wages in Britain. This is done using models inspired by those found in Tam (1997) which add controls for skill specialisation. However, rather than relying on a single indicator of SHC, five alternative approaches are used. The first approach uses variables indicating whether or not the individual received any on-the-job training in the year before the interview, the proportion of employees in the individual's occupation who participated in education or training courses connected with their present or future job in the last four weeks and the modal time spent in such training in the individual's occupation. The last two variables are derived from pooled LFS data for years 1992-1994 and 1997-2007.

The second approach uses a condensed version of the Eriksson, Goldthorpe and Portocarero (EGP) class schema derived from BHPS data which captures two dimensions of task specificity: firm-asset specificity and monitorability (Polavieja, 2005). Classes I (higher managerial and professional employees) and II (lower managerial and professional employees) have *high* SHC; classes IIIa (routine clerical employees) and V (manual supervisors) have *medium* levels of SHC; and classes IIIb (employees in routine service and sale jobs), VI (skilled manual job), VIIa (semi- and unskilled manual jobs) and VIIb (agricultural jobs) have *low* SHC.

The third approach aggregates occupations into four major skill groups following Elias and McKnight (2001). The least skilled employees in level 1 display *'competence associated with a good general education, usually acquired by a time a person completes his/her compulsory education'* and may also get involved in *'short periods of work-related training'* (p.511-512). Occupations in skill level 2 require *'the knowledge provided via a good general education'* but *'typically have a longer period of work-related training or work experience'* (p.512). Occupations in the third level of skill *'require a body of knowledge associated with a period of post-compulsory education but not to degree level'* as well as *'a significant period of work experience'* (p.512). Finally, the highest level of skill (4) includes occupations for which *'a degree or equivalent period of relevant work experience'* is needed (p.512).

Two more indicators of SHC are constructed using pooled data from the 2001 and 2006 SS. The fourth measure uses the mean response by minor occupational group to a question in which workers are asked to quantify the importance of 'specialist knowledge or understanding' for their jobs on a scale from zero (not important at all) to four (essential). The fifth measure uses median responses to a question on job-learning time by minor occupational group.[3] These variables are derived using SOC2000. Since no direct conversion is possible between SOC90 and SOC2000, this information can only be matched to waves 11 to 17 of the BHPS and sample sizes are therefore smaller when using these variables in estimation.

Table 2 presents pairwise correlations between occupational feminisation, wages and the measures of SHC. In principle, a given factor will mediate the impact of occupational feminisation on wages if it is negatively associated with occupational feminisation (i.e. it is scarce in more feminised occupations) and positively associated with wages (i.e. it increases pay), or *vice versa*. As predicted by SHC theory, most of the variables capturing skill specialisation are negatively associated with occupational feminisation and positively associated with wages, which suggests that they may mediate the impact of occupational feminisation on wages. However, training-related variables are positively correlated with occupational feminisation: employees in more feminised occupations are more likely to receive training (r=0.04) and to work in occupations in which a higher proportion of employees receive training (r=0.27). This is inconsistent with SHC theory and suggests that these variables do not mediate the effect of occupational feminisation on wages. Interestingly, longer periods of training are associated with receiving lower wages (r=–0.20).

[TABLE 2 HERE]

[3] Possible responses to the job-learning time question are: less than one week; less than one month; one to three months; three to six months; six months to one year; one to two years; and more than two years.

These relationships are now explored within a multivariate framework by introducing SHC variables to the base regression models. If SHC mediates the effect of occupational feminisation on wages, then adding the SHC variables will move the estimated coefficient on occupational feminisation towards zero. Table 3 summarises results from OLS models. As a benchmark, the estimated coefficients on occupational feminisation in the base models were −0.171 for men and −0.323 for women. Model 2a adds to the base specification the first set of SHC controls: whether or not the respondent received any on-the-job training in the past 12 months, the proportion of employees in the minor occupational group that received on-the-job training and the modal length of such training. The receipt of training has a positive and statistically significant impact, increasing wages by 1.3 percent for men and by 2.9 percent for women. The proportion of trainees in an occupation has a large positive effect on wages for both sexes: a ten percentage-point increase in the proportion trained is associated with 17 percent higher wages for men and 20 percent higher wages for women. However, including these variables increases rather than decreases the wage penalty associated with working in female-dominated occupations – the estimated coefficients become −0.232 for men and −0.366 for women.

Model 2b includes the condensed version of the EGP class schema. For both men and women the degree of specialisation is statistically significant and positively associated with wages. Being in the highest specialisation group is associated with a wage premium of 43.9 percent for women and 36.5 percent for men relative to being in the lowest specialisation group. Consistent with SHC theory, including this specialisation measure increases the wage penalty associated with occupational feminisation for men – the estimated coefficient becomes −0.243. However, such penalty becomes less negative for women (−0.139), which is inconsistent with SHC theory.

Model 2c introduces the classification denoting skill requirements of occupations from Elias and McKnight (2001). Respondents working in occupations requiring higher skill levels earn considerably more than respondents in occupations requiring lower skill levels. Women gain relatively more than men from working in skilled occupations at any level of skill and this difference increases with skill levels. More importantly, the introduction of these SHC indicators reduces the negative effect of occupational feminisation on wages more than any other set of controls. The coefficient is now −0.094 for men. For women, it falls to −0.044 and is only significant at the 10 percent level.

Model 2d uses the first measure of specialisation derived from the SS. Wages increase with average self-reported levels of specialisation: in a five-point scale, a one-unit increase in specialisation is associated with a wage premium of 28.9 percent for men and 33.6 percent for women at sample means. Including this variable also mediates the negative impact of occupational feminisation on wages, although the coefficient is still negative and statistically significant (−0.1 for men and −0.163 for women).

Model 2e uses the second skill specialisation indicator from the SS. Wages increase with job-learning time: a one-unit increase in the median job-learning time in respondent’s occupation is associated with wages 8.6 percent and 12.9 percent higher for men and women respectively. Including this variable explains a large share of the negative effect of occupational feminisation on men’s wages, although it has little effect on women's. The coefficient on occupational feminisation for men becomes −0.044 and is only statistically significant at the 90 percent level.

An interesting discussion point is why the results differ when using different measures of SHC. Introducing training-based measures does not reduce the wage penalty for working in female-dominated occupations, possibly because women in Britain now undertake training almost as often as men – although of a shorter duration. This is in line with the results from the pairwise correlations in Table 2 and with findings from recent literature (Felstead et al., 2007; Jones et al., 2008). One of the measures of SHC with the largest effect on the impact of occupational sex-composition on wages is the skill-based subdivision of SOC from Elias and McKnight (2001). It is widely-accepted that a sex-bias in skill conceptualisation and evaluation affects occupational classifications (Grimshaw and Rubery, 2007; Phillips and Taylor, 1980; Steinberg, 1990). Therefore, to the extent that such biases are embedded into this measure, it is possible that the effect of occupational feminisation on wages is downward-biased in specifications which use it. Finally, although they undertake training as often as men, individuals working in female-dominated occupations report lower levels of specialist knowledge required for the job and job-learning time. This suggests that the training received in female-dominated occupations concentrates on more transferable skills or that there are sex-differences perceiving or reporting skill specialisation (Correll, 2001; Horrell et al., 1994). Also, it must be kept in mind that these indicators of SHC are imperfect and suffer from measurement error. They are based on occupational-means rather than on individuals' reports of their skills and therefore cannot capture within-occupation differences in SHC.

In a final model, a more stringent set of covariates than usually found in the relevant literature is used. These capture other factors which may mediate the relationship between occupational sex-segregation and wages, such as domestic labour supply, socialisation, authority at work and compensating differentials. Domestic labour supply is measured using information on weekly housework hours and caring for children under 11 and for sick or elderly relatives. Socialisation is measured using an index of gender-role attitudes constructed from a battery of questions on individuals' perceptions of the roles of men and women in society and the labour market (Swaffield, 2000). Workplace authority is measured by combining information on whether individuals' report having managerial or supervisory duties at work. Compensating differentials are measured using occupational injury rates calculated using LFS data, unpaid overtime, travel to work time and working shifts or unsociable hours. The indicator of SHC which had the largest impact on the effect of occupational feminisation on wages without reducing the estimating sample size is also included. This is the occupational skill level variable from model 2c. The estimated coefficients on the additional variables are in line with expectations and due to space constraints will not be discussed further here. The coefficient on the SHC variable remains virtually unchanged, but the parameters on occupational feminisation become substantially less negative relative to those in base models. Coefficients move from −0.171 to −0.126 for men and from −0.323 to −0.084 for women, but remain negative and statistically significant. Therefore, the inclusion of controls for domestic labour supply, socialisation, authority at work and compensating differentials does not qualitatively change previous results.

[TABLE 3 HERE]

So far, the wage penalties associated with working in a completely female-dominated occupation relative to a completely male-dominated occupation remain even when controlling for theoretically-relevant controls. However, the OLS estimates could be biased if there are unobserved factors correlated both with worker's wages and their sorting into occupations. This is addressed by estimating wage equations using FE. Results from FE estimation of models including SHC controls are presented in Table 4. For men, the estimated coefficient on occupational feminisation in the base model is –0.127. The addition of SHC indicators in models 2c, 2d and 2e reduces this to around –0.09. Hence, SHC explains some, but not all, of the impact of occupational feminisation on wages. The estimated coefficients on occupational feminisation for women are less negative in models 2c, 2d and 2e (–0.07 to –0.10) than in the base model (–0.17). As for men, SHC explains some of the impact of occupational feminisation on wages. For both men and women, the wage returns to SHC among men diminish and even disappear, suggesting that unobservable characteristics are important in allocating workers to different training schemes and influence their SHC accumulation. More motivated and able employees have both higher wages and higher levels of SHC.

[TABLE 4 HERE]

Overall, estimates from FE models are consistent with those from OLS specifications. The inclusion of SHC variables reduces the impact of occupational feminisation on wages relative to the base model, but the coefficients on occupational feminisation remain negative, relatively large and statistically significant. Hence, results indicate that the gender-composition of occupations has an impact on wages that is independent of SHC. Predicted wage penalties associated with working in a completely female-dominated relative to a completely male-dominated occupation range from 8 percent 13 percent for men and from 7 percent to 15 percent for women in the preferred specifications. Thus, they are larger than those found in previous studies incorporating SHC: less than 1 percent in Tam (1997), around 3 percent in Polavieja (2007, 2009) and 6 percent in Tomaskovic-Devey and Skaggs (2002) and Polavieja (2008), none of which was statistically significant. There are several potential explanations for these disparities. There may be differences between Britain and other countries in the mechanisms or institutions which allocate workers across occupations or wages across occupations and workers. They may also be caused by the use of less detailed measures of occupational feminisation in previous research, by how SHC has been operationalised in these studies, or by the inclusion of individual-specific time-invariant unobserved-heterogeneity. More homogeneous international comparative studies would shed more light on this.

<u>The contribution of occupational sex-segregation to the gender wage gap</u>

The contribution of occupational sex-segregation to the gender wage gap can be calculated following a simple decomposition technique (Tomaskovic-Devey, 1995).[4] Intuitively, this gives the percentage of the gender wage gap which is explained by occupational sex-segregation, controlling for variables in the model. The estimates from conducting this analysis are presented in Table 5.

In OLS models which control for SHC, occupational feminisation explains between 16 percent (model 2c) and 52 percent (model 2a) of the gender wage gap. In the preferred FE specifications, estimates are lower and range from 14 percent (model 2c) to 25 percent (model 2a). Therefore, results suggest that occupational sex-composition accounts for between 14 percent and 25 percent of the gender wage gap in Britain. These estimates are similar to those from the US literature and higher than those from previous studies using British data.

[TABLE 5 HERE]

Discussion and conclusion

The aim of this article was to examine the impact of the sex-composition of occupations on the wages of workers in Britain and, for the first time in this literature, evaluate the role of SHC in explaining this relationship. The analyses first replicated those of Tam (1997) on the US, and then offered additional insights by exploring whether the relationship between occupational feminization and wages was linear or non-linear and whether it was robust to accounting for unobserved characteristics of workers.

Results indicate a strong negative relationship between occupational feminisation and wages. This remains in models that introduce controls for SHC, domestic labour supply, socialisation, workplace authority and compensating differentials. In the preferred FE specifications which allow for time-invariant individual-specific unobserved heterogeneity, such penalty is of 7 percent to 13 percent for men and of 9 percent to 15 percent for women. Thus, it can be concluded there is a negative impact of occupational feminisation on wages which is independent of all other factors. This provides support for the devaluation theory, as the hypothesis that deep-rooted societal mechanisms contribute to the undervaluation of 'women's work' cannot be rejected. However, there is also support for the SHC thesis: skill specialisation increases wages net of education, age, job tenure and other important drivers of pay and also reduces the effect of occupational feminisation on wages. Other findings indicate that women receive wages which

[4] Mathematically, such contribution can be expressed as:

$$\mathrm{abs}\left(\frac{\left[\left(\bar{X}_f - \bar{X}_m\right)\beta_{all} * (AW_{all})\right]}{AW_{wom} - AW_{men}}\right) * 100$$

where $\bar{X}_f$ and $\bar{X}_m$ are sample averages of occupational feminisation for men and women respectively; β_{all} is the coefficient on occupational feminisation in a regression model pooling men and women; and AW_{all}, AW_{men} and AW_{wom} are sample average wages for all respondents, men and women respectively.

are around 15 percent lower than those of otherwise similar men, that the sex-segregation of occupations accounts for at least 14 percent of the gender wage gap, that unobserved individual traits play an important part in allocating workers to occupations with different gender-compositions and that the relationship between occupational feminisation and wages is linear.

Taken together, these findings that the sex-composition of occupations affects wages and that SHC partially explains this imply that a multifaceted policy approach should be applied to reduce gender inequality in the labour market and narrow the gender wage gap (Rubery and Fagan, 1995). On the one hand, future policy should target women's work-life balance and support their accumulation and preservation of specific skills, for instance through statutory childcare, maternity leave, creation of flexi-time positions in male-dominated occupations and enhanced awareness of non-traditional career paths. On the other hand, comparable-worth strategies which ensure that work of equivalent value is equally remunerated regardless of the sex-composition of the workforce should also enter policy debates (Findlay et al., 2009).

Future research could further the results in this article by developing similar analyses in an international comparative perspective or by identifying other factors that contribute to the relationship between occupational feminisation and wages. The operationalisation of SHC should be improved by investigating the extent to which on-the-job training, job specialisation and job-learning time capture the same underlying process; the relationships between skill specificity at the job, firm and occupation levels; and the degree to which SHC is transferable across occupations in a career ladder. In Britain, two datasets could help achieving this. First, the SS could be used to gain insights into the nature, content and duration of on-the-job training and the types and amounts of work-related skills in male- and female-dominated occupations using individual reports rather than aggregated data. Second, the British Workplace Employee Relations Survey could be used to explore the mediating effect of firm-level characteristics and the relative contributions of occupation, establishment and job sex-composition to the gender wage gap. However, these datasets are relatively small, do not follow individuals over time and lack some important information available in the BHPS and so their use implies several trade-offs.

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Tables and figures

Figure 1. Average wages by decile of occupational feminisation

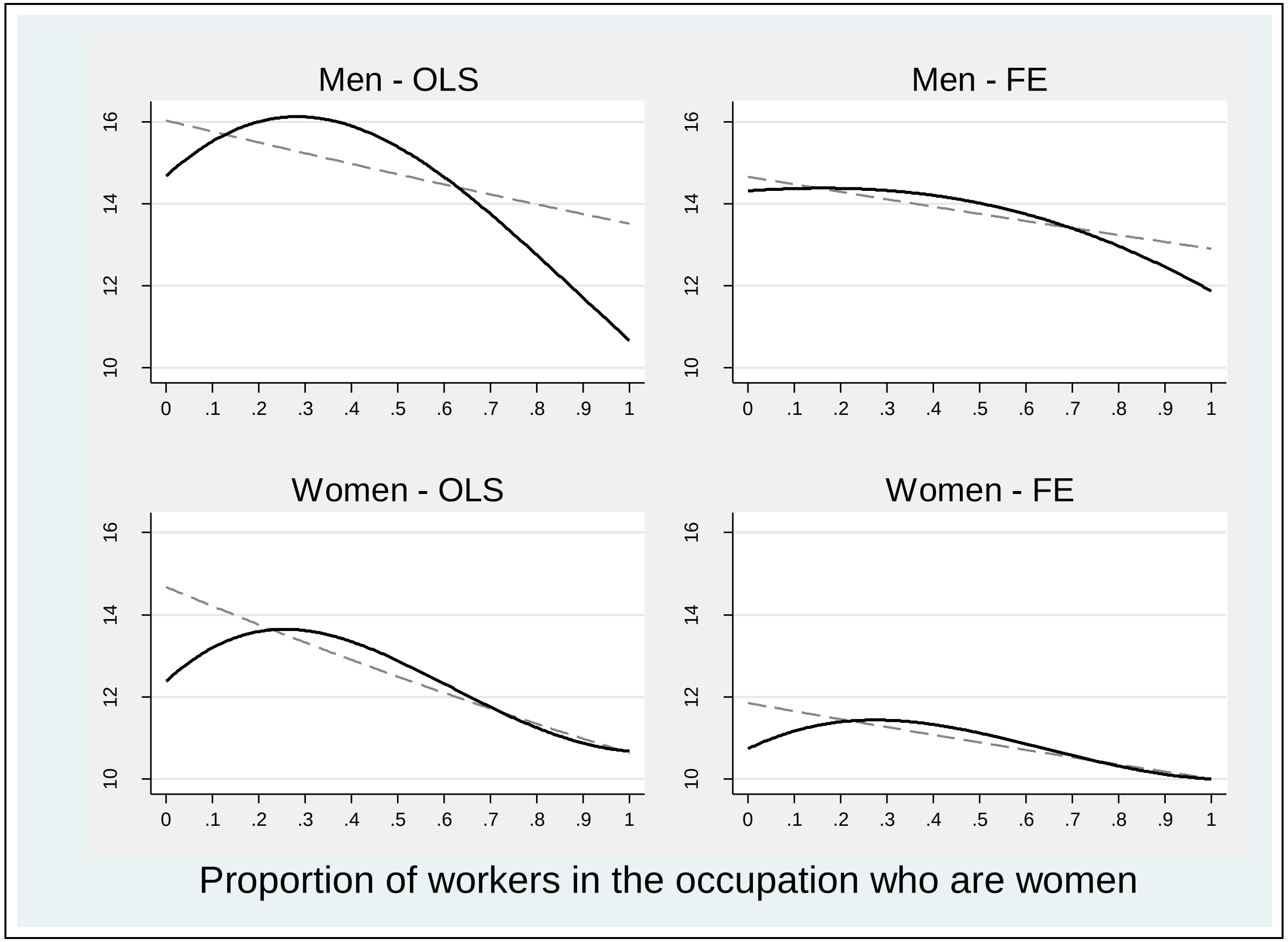
Notes: UK 1991-2007.

Table 1. The impact of occupational feminisation on wages: Base models

	Men		**Women**	
	OLS	**FE**	**OLS**	**FE**
Occupational feminisation	-0.171	-0.127	-0.323	-0.169
Base controls	Yes	Yes	Yes	Yes
Additional controls	No	No	No	No
N (observations)	*26,365*	*26,365*	*29,446*	*29,446*
N (individuals)	*3,968*	*3,968*	*4,359*	*4,359*
R^2	*0.434*	*0.299*	*0.436*	*0.220*

Notes: UK 1991-2007. Dependent variable = Logged hourly wages. All coefficients are statistically significant at the 1 percent level. OLS: Ordinary least squares. FE: Fixed effects.

Figure 2. Predicted wages by occupational feminisation



<u>Notes</u>: UK 1991-2007. The dashed line represents linear predictions from base models. Estimates from fixed effects (FE) models (right-hand side) are preferred over estimates from ordinary least squares (OLS) models (left-hand side), as they account for individual specific unobserved factors.

Table 2. Pairwise correlations

Specialised human capital variables	Occupational feminisation	Hourly wages
Received on-the-job training	0.04	0.13
Proportion of trainees in the occupation	0.27	0.32
Modal training time	-0.15	-0.20
Specialisation level in EGP schema	0.01¥	0.49
Specialisation level in Elias and McKnight	-0.28	0.53
Average 'specialist knowledge'	-0.14	0.48
Median job-learning time	-0.23	0.46

Notes: UK 1991-2007. ¥ Not statistically significant at the 5 percent level.

Table 3. The impact of occupational feminisation on wages: Ordinary least squares models

	Men						**Women**					
	2a	**2b**	**2c**	**2d**	**2e**	**3**	**2a**	**2b**	**2c**	**2d**	**2e**	**3**
Occupational feminisation	-0.232	-0.243	-0.094	-0.100	-0.044*	-0.126	-0.366	-0.139	-0.044**	-0.163	-0.180	-0.084
Received on-the-job training	0.013*						0.029					
Proportion of trainees in occupation	1.689						1.947					
Modal training: <1 week (reference)												
Modal training: >1 week & <1 year	-0.042*						-0.112					
Modal training: 1-3 years	-0.153						-0.191					
Modal training: >3 years	-0.096						-0.254					
Modal training: indefinite or continuous	-0.078						-0.118					
Low occ. specialisation (reference)												
Medium occupational specialisation		0.122						0.161				
High occupational specialisation		0.365						0.439				
Level 1 of skill (reference)												
Level 2 of skill			0.130			0.124			0.153			0.137
Level 3 of skill			0.260			0.220			0.357			0.306
Level 4 of skill (highest)			0.466			0.331			0.603			0.486
Average 'specialist knowledge'				0.289						0.336		
Median job-learning time					0.086						0.129	
Base controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Additional controls	No	No	No	No	No	Yes	No	No	No	No	No	Yes
N (observations)	*26,365*	*26,365*	*26,365*	*9,953*	*9,953*	*26,365*	*29,446*	*29,446*	*29,446*	*11,528*	*11,528*	*29,446*
N (individuals)	*3,968*	*3,968*	*3,968*	*2,507*	*2,507*	*3,968*	*4,359*	*4,359*	*4,359*	*2,869*	*2,869*	*4,359*
R^2	*0.478*	*0.513*	*0.502*	*0.469*	*0.448*	*0.542*	*0.512*	*0.529*	*0.521*	*0.479*	*0.482*	*0.562*

Notes: UK 1991-2007. Dependent variable = Logged hourly wages. * = Statistically significant at the 5 percent level only. ** = Statistically significant at the 10 percent level only. All other coefficients are statistically significant at the 1 percent level.

Table 4. The impact of occupational feminisation on wages: Fixed effects models

	Men					Women				
	2a	**2b**	**2c**	**2d**	**2e**	**2a**	**2b**	**2c**	**2d**	**2e**
Occupational feminisation	-0.134	-0.123	-0.086	-0.087	-0.085	-0.154	-0.100	-0.068	-0.099	-0.094
Received on-the-job training	-0.004¥					0.012				
Proportion of trainees in occupation	0.400					0.447				
Modal training: <1 week (reference)										
Modal training: >1 week & <1 year	0.013¥					-0.008¥				
Modal training: 1-3 years	-0.047					-0.104				
Modal training: >3 years	-0.047					-0.146				
Modal training: indefinite or continuous	0.013¥					-0.012¥				
Low occupational specialisation (reference)										
Medium occupational specialisation		0.054					0.086			
High occupational specialisation		0.127					0.175			
Level 1 of skill (reference)										
Level 2 of skill			0.059					0.062		
Level 3 of skill			0.088					0.133		
Level 4 of skill (highest)			0.148					0.198		
Average 'specialist knowledge'				0.059					0.083	
Median job-learning time					0.016					0.033
Base controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Additional controls	No	No	No	No	No	No	No	No	No	No
N (observations)	*26,365*	*26,365*	*26,365*	*9,953*	*9,953*	*29,446*	*29,446*	*29,446*	*11,528*	*11,528*
N (individuals)	*3,968*	*3,968*	*3,968*	*2,507*	*2,507*	*4,359*	*4,359*	*4,359*	*2,869*	*2,869*
R^2 (within)	*0.303*	*0.311*	*0.308*	*0.164*	*0.162*	*0.230*	*0.240*	*0.233*	*0.134*	*0.134*

Notes: UK 1991-2007. Dependent variable = Logged hourly wages. ¥ Not statistically significant at the 10 percent level. All other coefficients are statistically significant at the 1 percent level.

Table 5. The contribution of occupational sex-segregation to the gender wage gap

Model	OLS	FE
1	42%	25%
2a	52%	25%
2b	38%	20%
2c	16%	14%
2d	22%	15%
2e	16%	14%
3	18%	

Notes: UK 1991-2007. OLS: Ordinary least squares. FE: Fixed effects.