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Occupational Feminization, Specialized Human Capital and Wages: Evidence from the British Labour Market

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Non-technical summary

The gender pay-gap is one of the most enduring features of the labour market, although the debate about what mechanisms create and perpetuate it is still open. Previous literature has shown that the separation of men and women into different occupations is one of the drivers behind gender pay-differences. In particular, similar employees receive lower wages in occupations in which a higher proportion of workers are women. However, the reasons for such an effect are unclear. Cultural theories maintain that lower wages in female-dominated occupations are the product of societal bias against the work typically carried out by women and that the sex-composition of occupations affects wages directly. In contrast, recent human capital theories maintain that the wage-penalties associated with working in female-dominated occupations result from different requirements in specialized training and that the effect is indirect.

In this article, we explore how wages are affected by the proportion of employees in an occupation who are women and then test whether this association is mediated by skill specialization, job amenities, managerial responsibilities, socialization, domestic work or unobserved factors. We also estimate how much of the gender gap in wages occurs because of the separation of men and women into different occupations. Our empirical analyses use longitudinal data on individuals from the British Household Panel Survey from 1991 to 2007 and occupational-level information derived from the Labour Force Survey.

Results show that wages are lower in occupations dominated by women, and that these wage differences cannot be explained by human capital factors or any of the observable or unobservable characteristics considered. An individual working in an occupation in which all workers are men has wages which are around 10% higher than those of an otherwise identical individual who works in an occupation in which all workers are women. The segregation of men and women intro different lines of work explains around 15% of the gender pay-gap.

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Abstract

Research has consistently demonstrated a negative and significant relationship between occupational feminization and wages. This has traditionally been attributed to societal mechanisms undervaluing the work mainly performed by women. More recently, empirical evidence from the US and Europe has supported theories based on the concept of specialized human capital. We examine whether lower wages in female-dominated occupations in Britain are explained by differences in specialized human capital, allowing for other potentially mediating factors. We also explore the functional form of the relationship between occupational feminization and wages and estimate the contribution of occupational sex-segregation to the gender pay-gap.

Keywords: sex-segregation, human capital, skills, specialization, devaluation

JEL Codes: C23, J16, J24, J31, Z13

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1. Introduction

The gender pay-gap is one of the most enduring features of the labour market. Although there has been extensive research on gender inequality in the work context since the early 1970s, scholars are far from achieving consensus about what mechanisms create and perpetuate it. In this article, we use panel data and panel data methods to examine the impact of the sex-composition of occupations on the wages of men and women and on the gender pay-gap and test the importance of human capital and other theories in explaining this.[1]

The literature documents extensive and pervasive distributional inequality in respect to occupation within modern labour markets and also about the impact such segregation has on wages. There is general agreement in sociology and economics that working in an occupation in which a large proportion of workers are women incurs a wage penalty and is associated with other non-pecuniary costs, such as a loss of prestige or slower career progression. Wage effects are typically investigated by including the proportion of workers in an occupation who are women in a wage-equation. The sign and magnitude of the coefficient are interpreted as the impact of the gender-composition of an occupation on wages. Studies have also quantified the role of occupational feminization in explaining the gender pay-gap through different decomposition techniques.

We make three contributions to this literature. First, we provide a detailed examination of the relationship between occupational sex-segregation and wages in Britain using panel data and investigate whether the relationship is non-linear. Second, we examine the role of specialized human capital (SHC) in explaining the association between occupational feminization and wages, which has yet to be explored using British data. The SHC approach (Tam, 1997; Polavieja, 2007, 2008a, 2009) proposes that occupational sex-segregation does not directly affect wages, but that this association is caused by the lack of specialization of work in female-dominated occupations. This argument contrasts with the well-established devaluation hypothesis, which maintains that wage inequality is socially constructed and work in female-dominated occupations is undervalued as a result of institutionalized bias against women (Treiman and Hartman, 1981; Kilbourne et al, 1994; England et al, 2007). Thirdly, we use models which allow for individual-specific unobserved heterogeneity, which potentially biases previous estimates.

[1] We use the terms 'occupational sex-segregation', 'occupational sex-composition' and 'occupational feminization' interchangeably to refer to the proportion of employees in the individual's occupation who are female.

Determining the causes of any wage penalties to occupational sex-segregation has important policy implications for gender equality in the labour market and for the narrowing of the gender pay-gap. If these penalties are found to be caused by lower SHC in female-dominated occupations, policies should aim at enhancing women's opportunities to develop and/or be able to exploit their skills. For instance, individual-oriented policy responses such as improved maternity leave and government funded childcare would be suitable approaches. However, if the devaluation theory holds and work of equivalent value is remunerated differently in male- and female-dominated occupations, comparable worth policies should be favoured.

2. Literature Review

In this section we discuss (a) empirical evidence on the impact of occupational feminization on wages; (b) analyses on the contribution of occupational sex-segregation to the gender pay-gap and; (c) theories which explain these processes.

The impact of occupational feminization on wages

Many studies examine the extent to which occupational feminization affects wages. Estimates from different studies are, however, only partially comparable as they vary across datasets, years, units of analysis (i.e. individuals or occupations), model specifications, estimation methods and measures of occupational feminization.[2]

The empirical literature typically illustrates the impact of occupational feminization on wages by reporting the fall in wages due to a hypothetical move from a fully male-dominated to a fully female-dominated occupation. Following this tradition, Table 1 presents a review of results from previous studies. Overall, we can see that studies are spread evenly across time and focus primarily on the US. Most analyses use cross-sectional regression methods and census data, and the unit of analysis is more often individuals rather than occupations.

The common finding from US studies is that occupational feminization reduces the wages of men and women. Men working in fully female-dominated occupations earn wages

[2] Due to collinearity, the impact of occupational feminization on wages is confounded when models also include measures of industrial segregation, establishment segregation, occupation-establishment segregation, workplace segregation and/or job-level segregation. In these cases, the coefficient is often positive and significant or

between 7% (England et al, 1988) and 26% (Cotter et al, 1997) lower relative to working in fully male-dominated occupations, while the effects range between 4% (Gerhart and El Cheikh, 1991) and 42% (US Bureau of the Census, 1987) for women. The classic exception is Filer (1989) who finds positive returns to occupational feminization in his fully specified model. However, his study has been heavily criticised for the use of a large number of explanatory variables (over 220), as collinearity may obscure the real effect of occupational feminization on wages.

Results from studies for a variety of other countries are presented at the bottom of Table 1 and show that the effect of occupational feminization on wages varies across countries. Estimates from OLS analyses indicate that working in a fully female-dominated occupation is associated with wage penalties of between 2% (Walby and Olsen, 2003; UK) and 20% (Magnusson, 2009; Sweden) relative to a fully male-dominated occupation, while multilevel models report penalties of up to 54% of average wages (Haberfeld et al, 1998; Israel).

All 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. We know, for example, that women tend to be employed in public sector jobs which offer relatively high wages. They are also progressively entering well-remunerated professional occupations. Similarly, most blue-collar jobs are held by men, and they are relatively poorly paid. Furthermore, previous theoretical and empirical evidence suggest that the impact of the gender-composition of jobs and workplaces on wages may be non-linear (see Kanter, 1977; Pfeffer and Davis-Blake, 1987 or Reskin et al, 1999), but this branch of the literature has so far neglected occupational sex-segregation. One of our contributions is to examine whether or not the relationship between occupational feminization and wages is linear.

negative but very small in magnitude. Therefore, we do not discuss studies including multiple segregation measures in the review.

Table 1. Results from studies of the impact of occupational sex-segregation on wages.

Author/s	Year	Unit	Data	Method	Effect of a 'full switch' in occupational feminization on wages
United States					
Ferber & Lowry	1976	Occ.	1970 Census	OLS	Annual wages > Women: -1438$ / Men: -5008$
Snyder & Hudis	1976	Occ.	1970 Census	OLS	Annual wages > Women: -2070$ / Men: -3900$
Treiman & Hartmann	1981	Occ.	1970 Census	OLS	Annual wages > Women: -1630$ / Men: -2960$
England et al	1982	Occ.	1970 Census	OLS	Annual wages > Women: -1682$ / Men: -3005$
O'Neill	1983	Occ.	1980 CPS	OLS	Hourly wages > Women: -15.8% / Men: -14.8%
Johnson & Solon	1984	Ind.	1978 CPS	OLS	Hourly wages > Women: -9% / Men: -16.8%
England	1984	Ind.	1974 PSID	OLS	Hourly wages > Women: -18% to -19% / Men: No significant effect
US Bureau of the Census	1987	Ind.	1984 SIPP	OLS	Hourly wages > Women: -21.1% to -41.7% / Men: -18.9% to -24.1%
England et al	1988	Ind.	1968-1980 NLS	FE	Hourly wages > Women: -6% to -13% / Men: -7% to -25%
Sorensen	1989	Ind.	1984 PSID	OLS	Hourly wages > Women: -23% / Men: -23.9%
Parcel	1989	Occ.	1980 Census	OLS	Annual wages > Women: -762$ / M > -3527$
Filer	1989	Occ.	1980 Census	OLS	Hourly wages > Women: +30 cents / Men: +31 cents
Groshen	1991	Ind.	1974-1978 IWS	OLS	Hourly wages > Women & Men: -24.2% to -85.2% in different industries
Gerhart & El Cheikh	1991	Ind.	1983 & 1986 NLS	OLS+FE	Hourly wages > Women: -3.6% to -9.9% / Men: -19.2% to -13.2%
England	1992	Occ.	1980 US Census	OLS	Hourly wages > Women & Men: -40 to -60 cents
England et al	1994	Occ.	1980 US Census	OLS	Hourly wages > Women: -58 to -67 cents / Men: -1.28$ to -1.88$
Kilbourne et al	1994	Ind.	1968-1981 NLS	FE	Hourly wages > Women & Men: -4% to -10%
Macpherson & Hirsh	1995	Ind.	1983-1993 CPS	OLS	Hourly wages > Women: -10.64% / Men: -11.70%
Tomaskovic-Devey	1995	Ind.	1989 NCEHS	OLS	Monthly wages > Women & Men: -315$ to -889$
Elliott & Parcel	1996	Ind.	1979-1987 NLSY	OLS	Hourly wages > Women & Men: -22%
Cotter et al	1997	Ind.	1990 Census	OLS	Hourly wages > Women: -24% to -29.1% / Men: -16% to -26.1%
Other countries					
Semyonov & Lewin-Epstein (Israel)	1989	Occ.	1972 & 1983 Census	OLS	Hourly wages > Women & Men: -11% (lagged 10 years)
Heberfeld et al (Israel)	1998	Ind.+Occ.	1983 Census	OLS+MLM	Monthly wages > Women & Men: -19.1% (OLS) & -54.4% (MLM)
deRuijter et al (Netherlands)	2003	Ind.+Occ.	1997 LSO	MLM	Hourly wages > Women & Men: -7% (discrete female-dominated occ.)
deRuijter & Huffman (Netherlands)	2003	Ind.+Occ.	1997 LSO	MLM	Hourly wages > Women & Men: -5% (discrete female-dominated occ.)
Walby & Olsen (UK)	2003	Ind.	1991-1999 BHPS	OLS	Hourly wages > Women & Men: -2%
Olsen & Walby (UK)	2004	Ind.	1991-2002 BHPS	OLS	Hourly wages > Women & Men: -13%
De la Rica & Amuedo-Dorantes (Spain)	2005	Ind.	1995 & 2002 ESS	OLS	Hourly wages > Women & Men: -10% to -15.9%
Magnusson (Sweden)	2009	Ind.	2000 LNU	OLS	Hourly wages > Women & Men: -10% to -20%

Notes: A 'full switch' in occupational feminization is defined as a hypothetical move from an all male to an all female occupation. Occ.=Occupations; Ind.=Individuals. OLS=Ordinary least squares models; FE=Fixed effects models; MLM=Multilevel models. Data abbreviations: CPS= Current Population Survey; PSID= Panel Study of Income Dynamics; SIPP=Survey of Income and Program Participation; IWS=Bureau of Labor Statistics Industry Occupational Wage Surveys; NLS= National Longitudinal Survey; NCEHS=North Carolina Employment and Health Service Survey; IWS= PATC=National Survey of Professional, Administrative, Technical and Clerical employees; BHPS= British Household Panel Survey; LNU= Level-of-Living Survey; ESS=Spanish Wage Structure Survey; LSO=Netherlands Structure of Earnings Survey; NLSY= National Longitudinal Survey of Youth.

The contribution of occupational feminization to the gender pay-gap

A number of studies estimate the contribution of occupational feminization to the gender gap in wages. Again most of them focus on the US. Table 2 provides a summary.

US estimates from OLS regressions vary widely, even within studies, and suggest that occupational sex-segregation explains between 6% (England, 1992) and 69% (Gerhart and El Cheikh, 1991) of the gender-wage gap. Cotter et al (1997) use individuals and metropolitan areas as hierarchical units of analysis in a multilevel framework, and claim that occupational sex-segregation is responsible for almost all the gender pay-gap. Few studies use panel data. The only analysis using US panel data (Kilbourne et al, 1994) suggests that the gender composition of occupations explains between 8% and 17% of the gender pay-gap.

There are only a handful of studies that focus explicitly on Britain. Walby and Olsen (2003) and Olsen and Walby (2004) use data from the British Household Panel Survey (BHPS) and find that occupational feminization explains 8% and 10% of the gender wage-gap respectively. However, these figures are calculated using aggregated two-digit occupational categories, which obscure the degree of sex-segregation observed in the labour market and consequently may underestimate the real effect. More recently, Mumford and Smith (2007) use the 1998 British Workplace Employee Relations Survey (WERS) and find that occupational sex-segregation explains around 5% of the gender pay-gap. Again, this is likely to be downward-biased, because their models include both workplace and occupational feminization and these are highly correlated. No UK studies use panel data models to estimate the contribution of occupational sex-composition to the gender gap in wages.

In summary, between 5% and 40% of the gender-pay gap is attributed to occupational sex-segregation, with 20% to 25% regarded as a reasonable estimate (Hakim, 1992).

Table 2. Studies of the contribution of occupational sex-segregation to the gender pay-gap.

Author/s	Year	Units	Country	Data	Method	Contribution
Treiman & Hartmann	1981	Occ.	US	1970 US Census	OLS	19% to 41%
England et al	1982	Occ.	US	1970 US Census	OLS	21% to 38%
Johnson & Solon	1984	Ind.	US	1978 CPS	OLS	11% to 21%
US Bureau of the Census	1987	Ind.	US	1984 SIPP	OLS	17% to 43%
Sorensen	1989	Ind.	US	1984 PSID	OLS	24%
Goldin	1990	Occ.	US	1980 US Census	Other	19%
Groshen	1991	Ind.	US	1974-1978 IWS	OLS	11% to 26%
Gerhart & El Cheikh	1991	Ind.	US	1983 & 1986 NLS	OLS	10% to 69% (45%)
England	1992	Occ.	US	1980 US Census	OLS	6% to 11%
England et al	1994	Occ.	US	1980 US Census	OLS	7% to 16%
Kilbourne et al	1994	Ind.	US	1968-1981 NLS	FE	8% to 17%
Cotter et al	1995	Occ.	US	1980 & 1990 US Census	OLS	14% to 15%
Macpherson & Hirsh	1995	Ind.	US	1983-1993 CPS	OLS	12% to 19%
Tomaskovic-Devey	1995	Ind.	US	1989 NCEHS	OLS	23% to 43%
Petersen & Morgan	1995	Ind.	US	1974-1983 IWS & PATC	OLS	64%
Cotter et al	1997	Ind./Ind.+ MAs	US	1980 & 1990 US Census	OLS / MLM	25% to 38% / *"Almost the totality"*
Walby & Olsen	2003	Ind.	UK	1991-1999 BHPS	OLS	8%
Olsen & Walby	2004	Ind.	UK	1991-2002 BHPS	OLS	10%
Mumford & Smith	2007	Ind.	UK	1998 WERS	OLS	2% to 12%

Notes: Occ.=Occupations; Ind.=Individuals. OLS=Ordinary least squares models; FE=Fixed effects models; MLM=Multilevel models. MA=Metropolitan areas. Data abbreviations: CPS=Current Population Survey; PSID= Panel Study of Income Dynamics; SIPP=Survey of Income and Program Participation; IWS=Bureau of Labor Statistics Industry Occupational Wage Surveys; NLS= National Longitudinal Survey; NCEHS=North Carolina Employment and Health Service Survey; PATC=National Survey of Professional, Administrative, Technical and Clerical employees; BHPS= British Household Panel Survey; WERS= British Workplace Employee Relations Survey.

Theories of occupational sex-segregation and the gender pay-gap

There are a number of theories to explain occupational sex-segregation, its effect on wages and its contribution to the gender pay-gap. We discuss (and later test) the devaluation, human capital, socialization, gender gap in authority and compensating differentials hypotheses. Although demand-side factors like discrimination, labour market segmentation and social closure processes are also likely to be important, tests of these require information which is rarely available in survey datasets (e.g. on the characteristics, attitudes and hiring and promotion practices of both employees and employers). Most of these theories of occupational sex-segregation are not mutually exclusive. Many authors acknowledge that sex-segregation in the labour market and the gender pay-gap are multi-dimensional phenomena and warn against using simple single-theory explanations (England, 1984, 2005, Grimshaw and Rubery, 2007). Consequently, wide-ranging analyses and the simultaneous testing of alternative theories are important for a deeper understanding of work valuation processes.

Cultural explanations

Cultural explanations emphasise the role of ideology and tradition in defining which skills are valuable, desirable and profitable. Sociological theory has established that the distribution of power between sexes in society is not balanced and men dominate in almost all spheres of social life. The institution of patriarchy reflects this (Walby, 1986). Sex-bias in the social construction of value also operates in relation to the work performed by men and women. In line with the male-centred order of society, a higher value is attributed to jobs or occupations mainly carried out by men or associated with male-stereotyped skills.

The devaluation hypothesis offers a simple explanation for why male-dominated occupations receive higher wages than female-dominated occupations. Women's work is devalued by social structures and discrimination does not take place against individuals but against the types of jobs that they perform (Maume, 1999). A whole literature is devoted to the comparison of the rewards systems operating behind male- and female-dominated jobs or occupations (e.g. England, 1992; Kilbourne et al, 1994; Magnusson, 2009). These studies show that, although different in nature, the skills required to fulfil lower paid female-dominated jobs are comparable to those in better paid male-dominated jobs. Much attention in the comparable worth literature has been attributed to the devaluation of caring and nurturing skills traditionally associated with women (see England, 1992, 2005; England et al, 1994). The undervaluation of job tasks when these are performed by women has also been demonstrated in experimental research (Bose and Rossi, 1983; Major et al, 1984).

The devaluation hypothesis maintains, therefore, that the work commonly performed by women will be undervalued in respect to the work usually performed by men due to deep-rooted traditionally-established beliefs privileging men.

Human capital theories

Human capital (HC) theories are the most common and established economic explanations for differences in pay between men and women. Human capital is the stock of knowledge and skills accumulated by an individual and is acquired through education, training and experience. According to HC theories, gender differences in participation and wages are the result of gender-specific preferences regarding labour market investments and in the allocation of resources between the household and the workplace.

Becker's work effort/rational choice theory (1957, 1981, 1985, 1991) applies a utility-maximising standpoint to specialization in the household suggesting that if men expect to receive higher pay in the market than women, men will decide to work and women to stay at home. If women expect to spend less time in the labour market, they allocate fewer resources to their education, job-related training and duties at work than men and instead invest their efforts in family and household-related activities. Mincer and Polachek (1974) and Polachek, (1976, 1979, 1981) suggest that women are more likely than men to interrupt their work and careers due to family responsibilities. Intermittent employment leads to less labour market experience, forgone training and skill atrophy or depreciation. Therefore, to maximise their lifetime earnings, women may choose to work in positions and sectors of the economy in which work arrangements are more flexible, starting wages are highest, depreciation rates are lowest and wages are less dependent on experience, but which offer comparatively lower wages in the long run. Therefore, women become concentrated in a narrow range of occupations which offer such benefits and this explains the observed patterns of occupational sex-segregation and wage differences.

Although the concept of specialized human capital (SHC) has been present in the economics literature for a long time (see e.g. Parsons, 1972; Jovanovic, 1979; Corcoran and Duncan, 1979), Tam (1997) was the first to introduce it in the occupational sex-segregation field. Unlike general HC, investments in SHC are occupation, industry or firm-specific. Highly specialized jobs are risky for both the employer (who bears additional training costs) and the employee (who forgoes the possibility to apply the obtained skills in other job contexts). To prevent highly specialized workers from leaving their jobs, firms may offer long-term contracts with upward sloping wage-tenure profiles (Polavieja, 2007). So at high levels of tenure employees in highly specialized jobs receive wages which are comparatively higher than those offered by other jobs. Therefore, both workers and employers have incentives to maintain the employment relationship. 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 larger amounts of SHC. As a result, pay differences attributed to occupational sex-segregation may actually be due to differences in levels of specialization between occupations.

Tam (1997) uses 1988 US data to show that introducing SHC into wage equations reduces the negative effects of occupational feminization. SHC is measured using the years of specific vocational preparation required for an occupation, imputed from the American Dictionary of Occupational Titles (DOT). In his base models, moving from a completely male-dominated to a completely female-dominated occupation is associated with a wage penalty of 16% for men and 22% for women. These penalties disappear when including the SHC control. However, these results have been questioned. For example, England et al (2000) use the same dataset and show that the penalty associated with occupational feminization returns when adding a control for general educational development from the DOT (see Tam (2000) for a reply). Furthermore, the data and estimation methods do not allow for the potentially biasing effects of unobserved heterogeneity. A further criticism of Tam (1997) is that it uses an externally defined measure of SHC. These, and the DOT in particular, have been criticised for carrying an implicit sex-bias against the work performed mainly by women (Phillips and Taylor, 1980; Steinberg, 1990; Rees, 1992).

Tomaskovic-Devey and Skaggs (2002) use the 1989 North Carolina Employment Survey to define SHC in terms of job-learning time (as reported by the respondent) and use the proportion of females in the job (rather than in the occupation). Results are consistent with Tam (1997). Once job-learning time is added to the model, a switch from a fully male-dominated to a fully female-dominated job results in a 6% wage penalty which is not statistically significant. However, the sample in this study is very small (700 individuals) and not nationally representative and therefore the results are not generalisable.

Interrelated articles by Polavieja (2007, 2009) shift attention from the US to a group of European countries (including the UK). These use the 2004 European Social Survey (ESS) and a measure of job-learning time to operationalize SHC. Results, again consistent with those of Tam (1997), indicate that a switch from a fully male-dominated to a fully female-dominated occupation has no statistically significant impact on wages once SHC controls are included. However, these results are based on a measure of occupational sex-composition that is derived directly from the ESS and is therefore based on a very small sample. In addition, the 2007 article assumes a common occupational distribution across the countries under study when deriving occupational feminization levels. The later article corrects for this and uses country-specific measures, but does so at the expense of further limiting the sample size. A third paper, Polavieja (2008a), uses the same dataset but restricts the focus to Spain. This shares similar weaknesses and has a smaller sample size (n=1,100), although results are comparable. Moving from a completely male-dominated to a completely female-dominated occupation results in a 6% wage penalty, which is not statistically different from zero.

Socialization and domestic labour supply theories

Socialization theories use concepts such as sex-role socialization, stereotypes, and the gender-typing of jobs to explain sex-segregation at work. The patriarchal order of society produces differences in the socialization of boys and girls (Clausen, 1968), as they are raised to conform to a gendered identity determined by cultural standards. Consequently men and women are unequally distributed within the labour market because they display preferences for jobs and occupations which match their traits and skills and because crossing gender boundaries is socially penalized (Jacobs, 1989). The degree of internalization of traditional gender roles may be thought of as a process which sorts men and women into: (a) market and non-market work and, (b) particular jobs within the occupational structure. Furthermore, non-market work has traditionally been reserved for women in virtually every society at any point in time reducing women's commitment to the paid labour market.[3]

The time invested in non-market work and the choice of occupation may be mediated by the social context, and are expected to be highly correlated with a person's internalized perceptions about the role of women in relation to work and the family (Marini and Brinton, 1984). Women holding conservative attitudes may display real preferences for a certain kind of work ('women's work') and interrupt their career more often for household-related reasons. Hakim (2000) distinguishes between three different types of women according to their preferences towards labour market and homemaking activities: 'home centred', 'adaptive' and 'work-centred'. To the extent that the dual role of both domestic and market production affects women's capabilities at and preferences towards work, increased domestic labour supply and traditional attitudes may have a negative impact on wages for women.

Therefore, if more traditional women and those bearing a higher household burden concentrate in female-dominated lines of work, it is possible that these aspects explain the relationship between occupational feminization and wages.

<u>The gender gap in authority</u>

Since jobs and occupations involving authority over others have historically offered a wage premium, the gender gap in workplace authority could also explain the impact of occupational feminization on wages (England et al, 1994). Men tend to occupy managerial and supervisory positions more often than women and the integration of women into managerial occupations (especially higher managerial occupations) has been slow (ONS, 1961-1991). If women occupy such positions less often than men, and this work offers monetary benefits, the degree of authority exerted by the individual may explain the relationship between occupational sex-composition and wages.

<u>The compensating differentials hypothesis</u>

The compensating differentials hypothesis suggests that male-dominated occupations may enjoy a wage premium because they involve more hazardous working conditions (Schaffner and Kluve, 2006; Kilbourne et al, 1994). Unpleasant working conditions can take several forms, such as exposure to hazards, extremes of hot and cold, outdoor work, significant risk of injury or death and confrontation with high stress situations (Jacobs and Steinberg, 1990). Other unfavourable working conditions include travel time to one's workplace, number of unpaid overtime hours or times of the day in which the work takes place (Filer, 1989; Sloane et al, 2005). Thus, if (a) employees in more male-dominated occupations experience these conditions to a higher extent or more often than those in more female-dominated occupations,

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[3] However, time-use data suggests that trends may be shifting. Some examples are Gershuny and Fisher (2000)

and (b) these job characteristics are indeed associated with wage bonuses, compensating differentials may mediate the relationship between occupational feminization and wages.

In the remainder of the article we identify the impact of occupational feminization on wages in Britain and explore the extent to which this is accounted for by these various theories.

3. Data and methods

Datasets and sample

We use three different nationally representative datasets: the British Household Panel Survey (BHPS), the Labour Force Survey (LFS) and the 2006 Skills Survey.

The BHPS is a panel survey in which the same respondents have been interviewed on an annual basis every autumn since 1991. The first wave of the panel consisted of around 10,000 respondents from nationally representative randomly selected households across Britain. Continued representativeness of the British population is ensured by following panel members wherever they move within Britain and by a comprehensive weighting system which accounts for non-random panel attrition.[4] New members are added to the panel when original survey members form new households. While most available British datasets lack important dimensions for the study of occupational sex-segregation, are not representative of women in the UK or are extremely out of date, the BHPS provides an excellent source of data (Walby and Olsen, 2003).

Our analyses are based on a sample of British resident employees of working age (men aged 18 to 64 and women aged 18 to 59).[5] Those in full-time education are excluded because their choice of occupation is likely to be tied to their studies and unlikely to be final. Seventeen waves of the BHPS covering the period 1991-2007 are used. The resultant sample size is 8,326 individuals (3,968 men and 4,358 women) and 55,805 person-year observations (26,362 for men and 29,443 for women).

for the UK and Sayer (2005) for the US.

[4] Particular efforts have been made to avoid attrition in the BHPS and when this occurs is dealt with by the yearly updating of respondents weights. As an illustration of the degree of attrition in the BHPS, ten years after the beginning of the survey over 70% of eligible households still remained in the panel, with around 60% of them giving full interviews in each of the years (Lynn et al, 2006).

[5] As a sensitivity test, the analyses in this article were replicated excluding part-time workers. Results did not differ to the ones presented.

The LFS is a quarterly survey of individuals living at a random sample of about 60,000 private addresses in Britain. Its main purpose is to supply detailed information on individuals' labour market activity and employment and help to develop, manage, evaluate and report on labour market policies. Although the LFS can be used as a longitudinal resource, its rotational nature (with a maximum of 5 quarters per individual) is restrictive.

We use the LFS to construct occupation-level variables which are then matched to individuals in the BHPS by their SOC90/SOC2000 code and year.[6] In particular, the proportion of females in each three-/four-digit occupation is calculated for each year using LFS data.[7] The main advantage of the LFS is its large sample size, which allows for more accurate measurement of occupational feminization than possible using the BHPS. Due to data availability the proportion of females in each occupation is calculated using 371 three-digit SOC90 codes from 1991 to 2000, and 353 four-digit SOC2000 codes from years 2001 to 2007.[8] The quarterly LFS data are pooled into annual files to ensure sufficient sample sizes.[9] Occupational feminization is computed after excluding employees outside standard working age and the self-employed. Other occupation-level variables are also derived in the LFS and matched to individuals in the BHPS in a similar way.

The third dataset used is the Skills Survey (2006). This is a cross-sectional study of 8000 individuals which provides data on skill and job requirements in the British labour market. We use this to derive a measure of SHC at the three-digit occupational level using SOC2000.[10]

[6] There is some methodological debate about whether occupation-level variables can safely be appended to individual-level data (see Appendix 4 for a discussion).

[7] Typically in the literature the three-digit (or equivalent) level of occupation is used as it provides sufficiently detailed decomposition of occupational groups while maintaining cell sizes.

[8] There is no way of mapping SOC90 and SOC2000 codes onto each other, forcing this artificial break in the data series. The number of occupations is equivalent across both classifications and no major bias occurs from this change. Values for 1991 are replaced by their 1992 counterparts, since the 1991 LFS had insufficient respondents to calculate the proportion of females in the smallest occupations accurately.

[9] LFS figures for smaller SOC and SOC2000 occupations may still suffer from measurement error. Results from sensitivity tests using occupational sex-segregation derived from the 1991 and 2001 censuses did not differ substantially. Other studies also found the LFS to be a robust and reliable alternative to the use of the Census for measuring occupational sex-segregation (see e.g. Hakim, 1992, 1994 and Blackwell and Guinea-Martin, 2005).

[10] Since this indicator is derived from a cross-sectional survey and BHPS data containing SOC2000 codes stretches for a period of six years it is assumed that degrees of specialization across detailed occupational groups during this period of time are constant.

Model specification

The dependent variable in our analyses is the natural log of gross hourly wages, derived as follows:

$$\frac{\text{Usual gross monthly earnings from main job}}{(52/12) * \text{Weekly working hours in main job}}$$

Wages have been deflated to 2007 prices using Consumer Price Indices reported by the ONS and the top and bottom 1% of the distribution have been dropped to exclude outliers. The resultant average wages are £12.33 for men, £9.53 for women and £10.85 for the whole sample. The key explanatory variable is the proportion of females in an individual's occupation. In addition, all models control for year, region of residence, age, marital status, highest educational qualification, establishment size, permanent job, private sector job, part-time work, hours of work, job tenure and industry. We limit the number of control variables to avoid issues of collinearity (see Tam, 1997, 2000 and England et al, 2000 for a discussion) and to make results comparable with Tam (1997).[11]

Model estimation

Our models include a wide range of individual and job-related characteristics to help to identify accurately the impact of occupational feminization on wages. Despite this there may also be unobserved (or unobservable) individual-specific characteristics which influence wages. Such individual unobserved heterogeneity, if not suitably allowed for, can bias the coefficients of interest. This is important, since individuals may possess different unmeasured productivity-related factors which affect their wages, and personal tastes, preferences or psychological traits may shape individuals' decisions when choosing an occupation, industry or firm to work in. Also firms may base their hiring decisions on factors such as perceived ability or commitment, which are difficult to capture in survey data. Panel data allow us to control for unobserved time-invariant individual-specific effects.

Within our context the model to be estimated can be written as:

$$\log(WAGE_{it}) = FEM_{it}\beta + X_{it}'\delta + Z_i'\theta + v_{it} \quad \textbf{(1)}$$

[11] Only a measure of total labour market experience for the whole sample is missing. Sensitivity checks using information from the work-life history files in wave 3 of the BHPS only available for a subset of respondents show that the absence of this covariate does not qualitatively change the results.

where i and t subscripts designate individual and time respectively; log(WAGE) represents logged hourly wages; FEM is an indicator of the proportion of females in each respondent's occupation; X is a vector of observable time-varying individual-, job-, establishment- and occupation-level variables; Z is a vector of observable time-invariant characteristics; and β, δ, and θ are coefficients of interest. The error term v_{it} can be decomposed in the following way:

$$v_{it} = \upsilon_i + \varepsilon_{it} \qquad \textbf{(2)}$$

where υ_i represents individual-specific time-constant unobservable effects; and ε_{it} is a stochastic error term.

Estimating (1) using OLS ignores any time-invariant individual-specific characteristics (υ) which, if correlated with the observables (FEM, Z & X), will produce biased results. Within-group fixed effects (FE) models are estimated by taking deviations from individual-specific means over time in both dependent and explanatory variables. This removes the effect of unobserved time-invariant characteristics (υ) and allows for arbitrary correlation between observables and unobservables. However, the observed time-invariant covariates (Z) are also removed and their effects on the dependent variable cannot be directly estimated. The model to be estimated becomes:

$$\log(\text{WAGE})_{it} - \overline{\log(\text{WAGE})}_i = (\text{FEM}_{it} - \overline{\text{FEM}}_i)\beta + (X_{it} - \overline{X}_i)\,\delta + (\varepsilon_{it} - \bar{\varepsilon}_i) \qquad \textbf{(3)}$$

We employ FE regression as a complement to OLS regression to evaluate the extent to which the relationship between occupational feminization and wages is robust to controlling for unobserved individual heterogeneity.

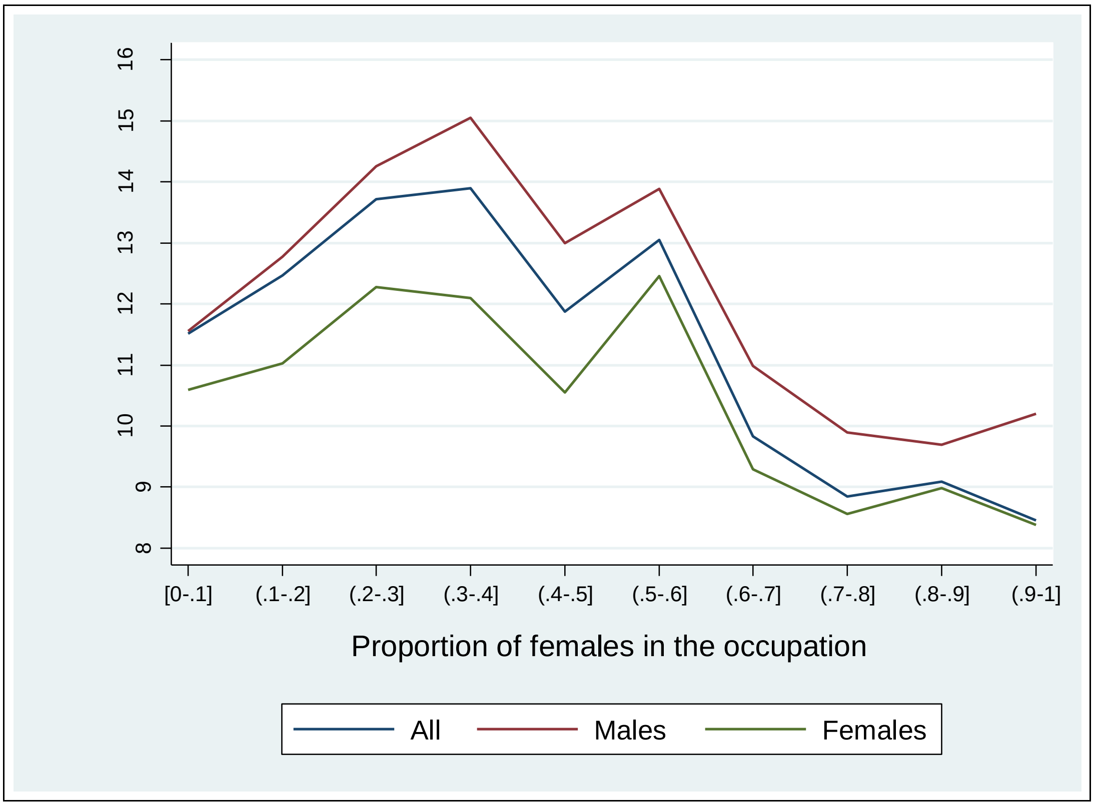
We estimate models pooling men and women and also gender-specific models. The former provide a general overview of the relationship between occupational feminization and wages and allow the impact of gender on wages to be estimated, while the latter are more flexible and allow the effect of all covariates to differ by sex. The standard errors in all models are adjusted to control for the clustering of observations within individuals. We first examine the relationship between occupational feminization and wages in detail before investigating the extent to which this can be explained by SHC, socialization, authority and compensating differentials theories.

4. The impact of occupational feminization on wages

Descriptive analysis

Graph 1 plots average wages of men and women by the proportion of females in their occupation. This indicates a similar relationship for men and women, with average wages highest in more integrated occupations than in sex-segregated occupations. There is also evidence that male-dominated occupations pay more than female-dominated occupations. The remainder of the article examines this relationship in more detail within a multivariate framework, and then explores potential explanations for why it emerges.

Graph 1. Wages by occupational feminization.



Base multivariate models

Tables 3 and 4 present estimates of the impact of occupational feminization on wages for men and women from models which include the control variables.[12] We initially discuss results from OLS models.

In model 1 the proportion of employees within the respondent's occupation that are women enters as a linear term. The R^2 in this model is .434 for men and .436 for women, indicating that observables explain around 44% of the total variance of wages. The estimated coefficients on occupational feminization 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% and 32% 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 feminization is not only negatively associated with wages but its effects are also more harmful to women. The coefficient on the female dummy in the pooled model indicates that women earn 13.4% lower wages than otherwise similar men (Appendix 2). This suggests that on average women earn about £1.42 per hour less than men, everything else (including occupational feminization) being equal.

Estimates from base FE models are consistent with those obtained from OLS, while the within R^2 falls to .299 and .218 for men and women respectively. Results indicate that once unobserved heterogeneity is allowed for, the wage penalties associated with female-dominated occupations are considerably lower. Moving from a fully male-dominated occupation to a fully female-dominated occupation is associated with a wage penalty of 12.8% (£1.58) for men and 16.7% (£1.60) for women relative to working in a completely male-dominated occupation. This suggests that unobserved characteristics of individuals (e.g. ability, motivation or taste) play an important role in allocating workers within the occupational feminization distribution, as those with unmeasured characteristics positively associated with wages tend to work in more male-dominated occupations.

However, these specifications assume that the relationship between occupational feminization and wages is linear. Descriptives in Graph 1 suggest that this is not the case. We now introduce models which allow for non-linearities in a number of ways.

[12] Table A1 in Appendix 1 presents a full set of estimates. The coefficients on control variables are in line with other studies and are not discussed here.

Table 3. The impact of occupational feminization on wages: Men.

	OLS					FE				
	1	**1b**	**1c**	**1d**	**1e**	**1**	**1b**	**1c**	**1d**	**1e**
Occupational feminization	-0.171***	0.550***	0.688***			-0.128***	0.087**	0.043		
Occupational feminization^2		-0.920***	-1.370***				-0.261***	-0.119		
Occupational feminization^3			0.361$^+$					-0.113		
Absolute segregation				-0.294***					-0.039*	
Female side				-0.175***					-0.070***	
Integrated occupations					*Ref. cat.*					*Ref. cat.*
Male-dominated occupations					-0.008					0.021***
Female-dominated occupations					-0.196***					-0.068***
N (observations)	26362	26362	26362	26362	26362	26362	26362	26362	26362	26362
N (individuals)	3968	3968	3968	3968	3968	3968	3968	3968	3968	3968
R^2	0.434	0.447	0.447	0.446	0.444	0.299	0.301	0.301	0.299	0.3

Notes: Dependent variable = Natural log of hourly wages. Significance levels: *** 0.01, ** 0.05, * 0.1, $^+$ 0.2. Models include a full set of controls –see Appendix 1 for details.

Table 4. The impact of occupational feminization on wages: Women.

	OLS					FE				
	1	**1b**	**1c**	**1d**	**1e**	**1**	**1b**	**1c**	**1d**	**1e**
Occupational feminization	-0.321***	-0.047	0.852***			-0.167***	0.034	0.525***		
Occupational feminization^2		-0.234***	-2.065***				-0.178***	-1.199***		
Occupational feminization^3			1.069***					0.608***		
Absolute segregation				-0.308***					-0.178***	
Female side				-0.122***					-0.058***	
Integrated occupations					*Ref. cat.*					*Ref. cat.*
Male-dominated occupations					0.047***					0.014$^+$
Female-dominated occupations					-0.139***					-0.072***
N (observations)	29443	29443	29443	29443	29443	29443	29443	29443	29443	29443
N (individuals)	4358	4358	4358	4358	4358	4358	4358	4358	4358	4358
R^2	0.436	0.436	0.438	0.434	0.435	0.218	0.219	0.22	0.218	0.218

Notes: As for table 3.

Exploring non-linearities

We estimate four additional specifications to examine the relationship between occupational feminization and wages in more detail. The first and second extensions include quadratic and cubic terms of the proportion of employees in an individual's occupation who are women (models 1b and 1c). The third identifies the pure effect of segregation on wages by using the absolute level of occupational sex-segregation. This is defined as the absolute difference between 0.5 (equality) and the proportion of females in an individual's occupation, and is introduced together with a variable which indicates the dominant sex in the occupation. The fourth approach (model 1e) follows previous literature and divides occupations into occupational sex-types (male-dominated, integrated and female-dominated occupations) (see e.g. de Ruijter and Huffman, 2003).[13] This is less restrictive than imposing the linearity assumption on the effect of occupational sex-composition on wages, but the effects within occupational sex-types are assumed to be equal.

<u>Men</u>

Table 3 summarises the results for men. We focus on OLS models first. The second and third columns (models 1b and 1c) present the results of including squared and cubic terms. These are consistent with Graph 1 and indicate a non-linear relationship. Wages initially increase with the proportion of females in the occupation but do so at a declining rate. The shape of the relationship is better illustrated in Graph 2 in which the average wages of a representative man are plotted as a function of occupational feminization.[14] The plots for the squared and cubic specifications are virtually identical and suggest that men receive highest wages in occupations in which about one third of employees are women. Results from model 1d show a significant wage penalty associated with working in sex-segregated occupations (an estimated coefficient of –0.294) and an additional penalty (of –0.175) if such occupations are segregated towards the female side. Results from model 1e indicate that working in

[13] Male-dominated occupations are characterized as those in which less than 35% of employees are women; integrated occupations as those in which between 35.01% and 65% of employees are women; and female-dominated occupations as those in which over 65.01% of the employees are women.

[14] The representative man is not involved in care duties, has completed a-levels, is married, has a permanent full-time job in the private sector, works in a firm which has between 100 and 1000 employees in the manufacturing industry and lives in London in 1999. The representative woman differs from the representative man in the industry in which she works, which is set as public administration. The values for continuous variables (age, job tenure and job hours) are the gender-specific sample means.

female-dominated occupations is associated with a wage penalty of 19.6% relative to working in integrated occupations.

The right side of Table 3 shows the equivalent models estimated using FE. Results from entering squared and cubic terms again suggest a non-linear relationship between occupational feminization and wages. However, plots of these estimates shown in Graph 2 indicate that the non-linearities are reduced once unobserved individual-specific fixed effects are allowed for. Segregated occupations still pay less than integrated occupations, but the differences are smaller than in OLS. The estimates from model 1d are consistent with those from OLS, but the wage penalties to absolute segregation (of 3.9%) and to segregation towards the female side of the distribution (of 6.8%) are much lower. The results from model 1e show that there is a small wage premium (of 2.1%) for working in male-dominated rather than integrated occupations when individual unobserved heterogeneity is controlled for. The wage penalty for working in a female-dominated rather than integrated occupation persists, but is considerably smaller than in OLS models (6.8% compared with 19.6%). The smaller effects of occupational feminization on wages in fixed effect models indicates that men working in more female-dominated occupations have unobserved characteristics that are associated with lower wages.

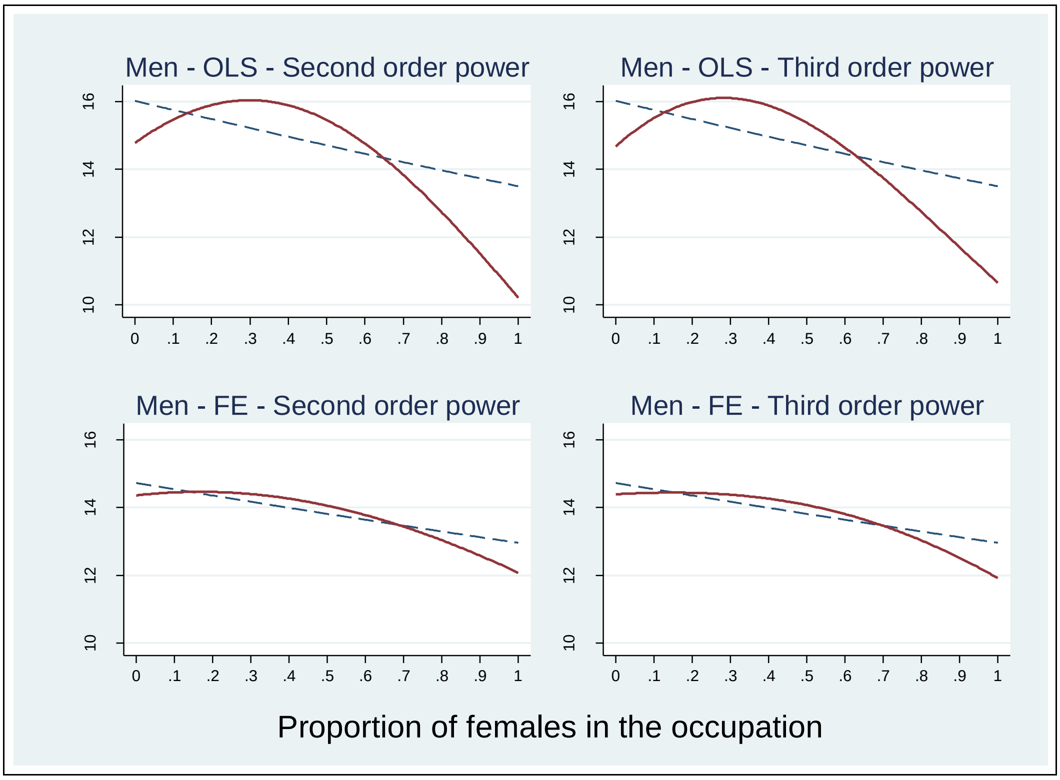
Women

We now consider the results for women shown in Table 4, and initially focus on OLS specifications. Estimates from introducing squared and cubic values of occupational feminization (models 1b and 1c) indicate a non-linear relationship, with wages initially increasing with the proportion of females in the occupation and then falling. However the plots in Graph 3 suggest that these are not as pronounced as for men. Estimates from model 1d show a wage penalty for women associated with working in a segregated occupation (of 30.8%) which is further increased (by 12.2%) if female segregated. Results from model 1e indicate that women working in male-dominated occupations have wages which are 4.7% higher than those of women in integrated occupations while, there is a wage penalty of 13.9% associated with working in female-dominated occupations.

The results from FE models for women are shown in the right side of Table 4 and in Graph 3. Estimates from the specifications with squared and cubic terms again indicate a non-linear relationship. As for men, Graph 3 shows that this is less pronounced in FE than in

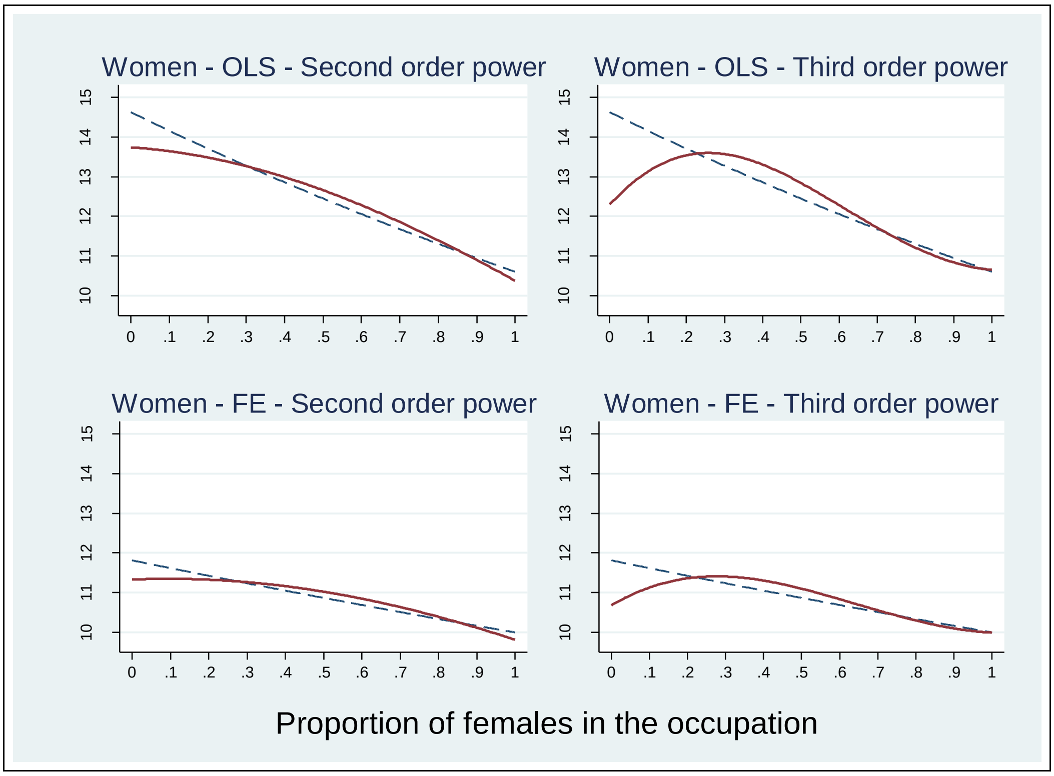
OLS specifications. The results from model 1d are consistent with those from OLS, although once more the sizes of the effects are smaller. Absolute segregation is associated with a wage penalty of 17.8% while working in the female side of the distribution reduces wages by an additional 5.8%. Unlike for OLS, in FE model 1e working in a male-dominated occupation is not significantly associated with wage gains relative to working in an integrated occupation. However, women in female-dominated occupations still have wages which are 7.2% lower than those of similar women employed in integrated occupations. Therefore, as for men, these results indicate that women working in female-dominated occupations have unobserved characteristics that are associated with lower wages. Even allowing for this, a wage penalty emerges for working in female-dominated occupations, with highest wages predicted for those in occupations where about 30% of employees are women.

Graph 2. Wages by occupational feminization: Men.



Notes: Results for including two and three power terms of occupational feminization. The dashed line represents the linear prediction.

Graph 3. Wages by occupational feminization: Women.



Notes: Results for including two and three power terms of occupational feminization. The dashed line represents the linear prediction.

5. Introducing potential mediating factors

In the following sections we examine the extent to which variables measuring SHC, socialization and domestic labour supply, authority at the workplace and compensating differentials mediate the impact of occupational feminization on wages. We do this using nested models in a model-building framework as in Tam (1997), Tomaskovic-Devey and Skaggs (2002) and Polavieja (2007, 2008a, 2009). Subsets of variables are added to the base model in a stepwise approach.[15] As several of the key variables of interest are time-invariant, and the main focus of this article is on SHC, we estimate fixed effect models for specifications that include SHC variables only. We first describe the variables relevant to each of the theories.

SHC variables

There are a number of ways to capture SHC using these data, and we use five different approaches. The first is to include a variable indicating whether or not the individual received any on-the-job training in the year before the interview, which is collected annually in the BHPS (see also Polavieja, 2008b). However, this is a relatively weak measure of SHC, given that it is individual-specific and may depend on career position.

The second approach uses the proportion of employees in each occupation that undertook education or training connected with their present or future job in the last four weeks. This information is available in the LFS for 1992-1994 and for 1997-2007.[16] The modal time spent in such training is also included.

The third approach follows Polavieja (2005) and uses a condensed version of the Eriksson, Goldthorpe and Portocarero class schema (EGP), which is derived from BHPS data. Classes I (higher managerial and professional employees) and II (lower managerial and professional employees) of this scheme include highly specialized employees, while classes IIIa (routine clerical employees) and V (manual supervisors) employ employees with medium levels of SHC. Employees in routine service and sale jobs (class IIIb), skilled manual jobs (class VI), semi and unskilled manual jobs (class VIIa) and agriculture (class VIIb) have low SHC.

The fourth approach aggregates the SOC90 and SOC2000 classifications into major skill groups, following Elias and McKnight (2001). Although, this reflects skill levels rather than specialization, distinguishing between the two is difficult. Four major skill levels can be identified. The least skilled employees in level 1 must 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"* (ibid). Finally, the highest level of skill (4) includes occupations for which *"a degree or equivalent period of relevant work experience"* is needed (ibid). Appendix 3 shows how occupational groups are allocated across the four skill levels.

[15] In these analyses we revert to the linear specification for our results to be comparable to those of previous research.

We construct the final measure of SHC from the 2006 Skills Survey, in which employees where asked to quantify the importance of particular tasks and skills for their jobs on a scale from zero (not important at all) to four (essential). One question relates to their 'specialist knowledge or understanding'. The mean response to this question by occupation is matched to respondents in the BHPS using three-digit SOC2000 occupational classification (this consists of 81 occupational groups, which is all that is available within the 2006 Skills Survey). 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 including this variable in our models (n=21,475).

Domestic labour supply and socialization variables

A second theory that may explain the relationship between occupational sex-segregation and wages relates to domestic labour supply and socialization. We capture domestic labour supply using the number of self-reported hours each respondent dedicates to housework per week, a variable identifying whether or not the respondent is responsible for caring for a sick or elderly person and a variable which identifies parents of children aged less than 11 who report doing most of the childcare. To measure socialization we use a variable which captures attitudes towards gender, work and the family and assesses individuals' perceptions of the roles of men and women in society and the labour market (see also Swaffield, 2000). This is calculated from responses to nine questions from the BHPS in which individuals are asked, on a scale from 0 to 4, the extent to which they believe that: (a) a pre-school child suffers if the mother works; (b) the family suffers if the mother works full-time; (c) a woman and a family are happier if she works; (d) husband and wife should both contribute to the household income; (e) a full-time job makes a woman independent; (f) husbands should earn and wives stay at home; (g) children need their father as much as their mother; (h) employers should help with childcare; and (i) a single parent can bring up children as well as a couple. The resulting index ranges from 0 to 36 where higher values indicate more traditional views. These variables are only collected biannually in the BHPS. Therefore, we calculate the average score for each individual using the information for all the waves in which the

[16] This indicator is derived from data pooled across years, and assumes that the proportion of people undergoing training in each occupation is time-invariant.

response is observed, and treat individual beliefs towards work and the family as time-invariant.[17]

Authority variables

The gender gap in workplace authority could also explain why people working in more feminised occupations suffer a wage penalty. To examine this we construct from the BHPS variables to measure an individual's managerial duties and working in a managerial occupation (SOC90 codes 100-199). The different categories identify people who have managerial duties in a managerial occupation, managerial duties in a non-managerial occupation (e.g. a head nurse of a particular hospital section or a school headmaster), non-managerial duties in a managerial occupation (e.g. a senior officer without direct influence over others) and no managerial duties in a non-managerial occupation.

Compensating differentials variables

Finally, we construct a range of variables to test whether the relationship between occupational sex-segregation and wages is explained by compensating differentials. The first is the average injury rate per 10,000 employees in the three-digit occupation, obtained from the LFS and matched to the BHPS using occupation codes. This variable is available in winter quarters of the LFS between 1993 and 2006 and asks respondents whether they had a work-related accident in the past 12 months. For sample size reasons we pool all winter quarters and construct a time-invariant indicator. Other measures of compensating differentials are obtained directly from the BHPS and include the number of unpaid overtime hours per week, the number of minutes spent travelling to work (one-way journey) and whether or not the respondent works shifts or unsociable hours. Although the latter are rather crude controls, they may be sufficient to identify any mediating effects of compensating differentials on the influence of occupational feminization on wages.

[17] These attitudes are, in fact, highly persistent. The mean difference in the values of the index for every observation from the time-invariant average was 1.21. An alternative construction of this scale using the first value for each respondent was tested and results did not change substantially.

6. Results

We present results from models which add variables to the base specifications to identify the extent to which SHC, domestic labour supply, socialization, authority and compensating differentials explain or mediate the relationship between occupational feminization and wages. If the negative effect of occupational feminization on wages is explained or mediated by these various theories, then adding the appropriate variables to the specifications will move the estimated coefficient on the occupational feminization indicator towards zero. Models 2a to 2e include measures of SHC: Model 3 adds a range of socialization and domestic labour supply variables; Model 4 introduces authority-related variables; and Model 5 incorporates compensating differentials. A final model includes all variables simultaneously. Models 2a to 2e, in which the additional variables are time-varying, are then re-estimated using FE.

OLS

Tables 5 and 6 summarise the results of OLS models for men and women (results for selected coefficients in pooled specifications are presented in Appendix 2). As a benchmark, the estimated coefficients on the occupational feminization term in the base models were –0.171 for men, –0.321 for women and –0.253 for the whole sample. The estimated R^2 in the new models ranges between .43 and .57.

Specialized human capital

Model 2a adds the first SHC control, whether or not the respondent received any on-the-job training in the past 12 months. Results show that, as predicted by SHC theories, training has a positive and significant effect, increasing wages by 4.3% for men and by 6.4% for women. However, including this variable does not reduce the negative and statistically significant effect of feminization on wages, the coefficient on which remains almost unchanged from the base models.

Model 2b includes the proportion of employees in the three-digit occupation that received on-the-job training in the past 12 months and the modal length of such training.[18] The

[18] As a robustness check, these were also entered separately. Results show that those in occupations with a modal training between one week and one month had a more severe wage penalty relative to those in occupations where the modal training is less than a week.

proportion of trainees in an occupation has a large positive effect on wages for both sexes. A ten percentage-point increase in the proportion of employees is associated with 17% to 20% higher wages, the effect being larger for women (Table 6) than for men (Table 5). Individuals in occupations in which the modal duration of training is less than a week have significantly higher wages than those in occupations with longer modal training periods.[19] Surprisingly, including these variables increases the penalty associated with working in female-dominated occupations. The estimated coefficients are now –0.233 and –0.366 for men and women respectively.

Model 2c includes the condensed version of the EGP classification.[20] For both sexes the degree of specialization is statistically significant and positively associated with wages. Being in the highest specialization group is associated with a wage premium of 44% for women and 36.4% for men relative to being in the lowest specialization group. For men, including this specialization measure actually increases the wage penalty associated with occupational feminization – the estimated coefficient becomes –0.245. For women, it becomes less negative (–0.137). Therefore, it seems that women are indeed sorted into different occupations in light of their SHC levels and that this explains part of the wage penalty associated with working in a female-dominated occupation. But why does the impact differ by sex? Two factors are important here: (a) how men and women are distributed across specialization groups and, (b) where specialization groups fall within the occupational distribution. The explanation may be that more specialized occupations are on average more integrated than other occupations for either sex. This can be seen as a mean reversal effect: for men, working in a highly-specialized occupation means moving to more feminized occupations than the average while the opposite holds true for women. If specialization is always correlated with higher wages, this may explain the diverging effects observed in these models.

Model 2d introduces variables denoting skill requirements of occupations. The results highlight the pecuniary advantages of working in more skilled occupations. Respondents

[19] One explanation compatible with the SHC thesis is based on a ‘career point’ effect. If a career is defined as a hierarchical succession of occupations, it is possible that most SHC (in terms of training time) is obtained at lower career levels. This idea introduces further theoretical and methodological challenges for the analysis of sex-segregation and SHC, since the benefits of on-the-job training may not only apply to the current occupation but to future positions on an occupational ladder. Also, not all occupations offer clearly defined career ladders. It is possible that ‘nested’ and ‘independent’ occupations may coexist within the occupational structure. Further work should pursue the question of whether this analytic strategy is compatible with a career-oriented view of occupations and acknowledge the different nature of occupations in relation to the availability of defined career ladders.

working in occupations requiring higher levels of skill earn considerably more than respondents in those requiring the lowest 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 feminization on wages more than any other set of controls. The coefficient is now –0.094 for men. For women, the coefficient falls to –0.044 and is only significant at the 10% level. Thus, a full switch from 0 to 1 in occupational feminization incurs a wage penalty of 9.4% for men and 4.4% for women.

Model 2e uses the measure of specialization derived from the 2006 Skills Survey. As with other SHC controls, wages increase with average self-reported levels of specialization. In a 5-point scale, an increase in one unit of specialization is associated with a wage premium of 29.8% for men and 34.5% for women at sample means. Including this variable also mediates the negative effect of feminization, although its impact is still negative and highly significant: the estimated coefficients are –0.102 for men and –0.204 for women.

Overall, these results suggest that some of the impact of occupational feminization on wages can be explained by SHC, and that the extent of this mediating effect varies with the measure of SHC used.

Socialization and domestic labour supply

Model 3 presents results from including socialization-related and domestic labour supply variables. Consistent with socialization theories, more traditional men earn higher wages (an estimated coefficient of 0.003) while more traditional women earn lower wages (–0.005). Hours spent on housework per week are, as expected, significantly and negatively associated with wages (each additional hour reduces wages by 0.7% for men and women). Likewise, caring for an infirm or elderly person reduces wages for both sexes by between 2% and 3%. Being responsible for most of the childcare of a child aged 0 to 11 has no significant impact on wages *ceteris paribus*. Although introducing these measures has little effect on the estimated coefficients on the occupational feminization variable, the coefficient on the female dummy in the pooled model becomes considerably less negative (–0.094 compared to –0.134 in the base model, see Appendix 2). Therefore, socialization and domestic labour supply

[20] Results using a more detailed 8-group Goldthorpe scale were similar.

explain little of the effect of occupational feminization on wages, but do contribute to the gender pay-gap.[21]

Gender gap in authority

Model 4 adds authority variables to the specification. The results show that, as expected, all managerial occupations and responsibilities are associated with higher wages holding other characteristics constant. This is true for both men and women and the effects are statistically different from zero. Having managerial duties in a managerial or non-managerial occupation has the largest effect, yielding a wage premium of 25% to 34%. For women, including these measures reduces the negative effect of occupational feminization on wages from –0.321 in the base model to –0.209. In contrast, the coefficients for men become more negative (from –0.171 to –0.208). These gender-differences result from a positive correlation between the proportion of females in the occupation and managerial responsibilities for men and a negative correlation for women. Therefore, for women male-dominated occupations are more often positions which require the exercise of authority, which partly explains the lower wages in female-dominated occupations.

Compensating differentials

Model 5 introduces variables capturing compensating differentials. The results suggest a wage-premium associated with working shifts or unsociable hours of 6.6% for men and 4.1% for women. Wages also increase with the number of weekly unpaid overtime hours for both men and women (around 0.2% per additional hour) and with the time spent travelling to work each way (from 12% to 18% per hour), but fall with the proportion of employees in an occupation that report work-related accidents (the estimated coefficients are –0.018 and –0.023 for men and women respectively). When introducing these compensating differentials, the occupational feminization coefficients become more negative for all specifications,

[21] These estimates could suffer from sample selection bias if people observed in employment are a non-random selection from the working age population (e.g. more traditional women will more often be found out of employment). Applying a Heckman selection model provides some evidence of sample selection for women but not for men. However, the coefficients of interest do not change.

especially for men (–0.305). Thus, compensating differentials do not explain the negative effect of occupational sex-composition on wages and in fact they exacerbate it.[22]

Fully specified model

In model 6 all covariates are introduced, although only one SHC measure is included because of collinearity. We use the occupational skill level variable from model 2d as it had the largest impact on the effect of occupational feminization on wages. The estimated coefficients on the SHC, socialization and domestic labour supply variables remain virtually unchanged in this specification. The key result is that, when including all controls, there is still a wage penalty associated with occupational feminization. The estimated coefficients become considerably less negative in all specifications from –0.171 to –0.126 for men and from –0.321 to –0.084 for women.

The main finding from OLS models is that the negative impact of occupational feminization on wages cannot be completely explained by other factors. The wage penalty associated with moving from a completely male-dominated to a completely female-dominated occupation ranges from 9.4% to 24.5% for men and from 4.4% to 36.6% for women in models which control for SHC. The inclusion of controls related to SHC, authority, socialization and domestic labour supply tends to reduce this penalty for women while having mixed effects for men. SHC is found to have a positive and statistically significant impact on wages for both men and women.

Differences with the results from previous studies are apparent even when using the most conservative estimates. Once SHC is accounted for, Polavieja (2007, 2008a, 2009) finds evidence that moving from a fully male to a fully female occupation reduces wages by between 1% to 3.1% in Europe and by 6% in Spain. Tomaskovic-Devey and Skaggs (2002) report a comparable figure of 6% in their US study. Tam (1997), using US data finds that moving from a completely male-dominated to a completely female-dominated occupation is associated with a wage penalty of 0.2% and 0.7% for men and women respectively when including SHC controls. None of the effects in these studies were statistically different from zero. These disparities in results may be caused by structural differences in the mechanisms

[22] As a sensitivity test we use ratios of workers reporting a *major* accident and death rates at the two-digit occupational level reported in Grazier and Sloane (2008). The first variable behaves in a similar way to the proportion of work-related accidents, while the death rates show no significant effect on wages. Neither affects the coefficient on occupational feminization.

which link workers to wages in Britain and other countries; by a downward bias in the estimates from previous studies due to the use of less elaborated measures of occupational feminization and/or by differences in the operationalization of SHC.

Table 5. OLS Estimates: Men.

	Model 2a	**Model 2b**	**Model 2c**	**Model 2d**	**Model 2e**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Occupational feminization	-0.172***	-0.233***	-0.245***	-0.094***	-0.102***	-0.167***	-0.208***	-0.305***	-0.126***
Received on-the-job training in the past 12 months	0.043***								
Proportion of trainees in the occupation		1.705***							
Modal training “less than a week”		*Ref. cat.*							
Modal training “more than a week but less than a year”		-0.042**							
Modal training “more than a year but less than three years”		-0.153***							
Modal training “more than three years”		-0.095***							
Modal training “indefinite, continuously”		-0.080***							
Low occupational specialization			*Ref. cat.*						
Medium occupational specialization			0.124***						
High occupational specialization			0.364***						
Level 1 of skill (lowest)				*Ref. cat.*					*Ref. cat.*
Level 2 of skill				0.130***					0.125***
Level 3 of skill				0.261***					0.221***
Level 4 of skill (highest)				0.466***					0.332***
Average self-reported “specialist knowledge”					0.298***				
Attitudes towards women, family and employment						0.003*			0.003**
Hours dedicated to housework per week						-0.007***			-0.005***
Caring for an ill or elderly person						-0.021*			-0.020*
Doing most of the childcare of an infant (aged 0 to 11)						-0.015			-0.025
No managerial duties							*Ref. cat.*		*Ref. cat.*
Managerial duties in managerial occupation							0.336***		0.183***
Managerial duties in a non-managerial occ.							0.287***		0.198***
No managerial duties in a managerial occupation							0.147***		0.033*
Ratio of accidents per 10000 employees in the occupation								-0.018***	0.003^{+}
Works shifts or unsociable hours								0.066***	0.082***
Number of unpaid overtime hours								0.015***	0.008***
Minutes spent traveling to work								0.002***	0.001***
N (observations)	26362	26362	26362	26362	9949	26362	26362	26362	26362
N (individuals)	3983	3983	3983	3983	2506	3983	3983	3983	3983
R^2	0.435	0.477	0.512	0.502	0.468	0.439	0.505	0.482	0.542

Notes: Dependent variable = Natural log of hourly wages; other non-reported controls: year, region, education, establishment size, job tenure and its square, permanent contract, age and its square, marital status, part time work, job hours and its square, private sector and industry. Significance levels: *** 0.01, ** 0.05, * 0.1, $^{+}$ 0.2.

Table 6. OLS estimates: Women.

	Model 2a	Model 2b	Model 2c	Model 2d	Model 2e	Model 3	Model 4	Model 5	Model 6
Occupational feminization	-0.322***	-0.366***	-0.137***	-0.044*	-0.204***	-0.314***	-0.209***	-0.353***	-0.084***
Received on-the-job training in the past 12 months	0.064***								
Proportion of trainees in the occupation		1.975***							
Modal training "less than a week"		*Ref. cat.*							
Modal training "more than a week but less than a year"		-0.112***							
Modal training "more than a year but less than three years"		-0.192***							
Modal training "more than three years"		-0.259***							
Modal training "indefinite, continuously"		-0.118***							
Low occupational specialization			*Ref. cat.*						
Medium occupational specialization			0.164***						
High occupational specialization			0.440***						
Level 1 of skill (lowest)				*Ref. cat.*					*Ref. cat.*
Level 2 of skill				0.154***					0.137***
Level 3 of skill				0.358***					0.306***
Level 4 of skill (highest)				0.604***					0.487***
Average self-reported "specialist knowledge"					0.345***				
Attitudes towards women, family and employment						-0.005***			-0.004***
Hours dedicated to housework per week						-0.006***			-0.004***
Caring for an ill or elderly person						-0.030***			-0.029***
Doing most of the childcare of an infant (aged 0 to 11)						0.004			-0.005
No managerial duties							*Ref. cat.*		*Ref. cat.*
Managerial duties in managerial occupation							0.255***		0.029+
Managerial duties in a non-managerial occ.							0.303***		0.170***
No managerial duties in a managerial occupation							0.135***		-0.042**
Ratio of accidents per 10000 employees in the occupation								-0.023***	-0.008***
Works shifts or unsociable hours								0.041***	0.033***
Number of unpaid overtime hours								0.021***	0.012***
Minutes spent traveling to work								0.003***	0.003***
N (observations)	29443	29443	29443	29443	11526	29443	29443	29443	29443
N (individuals)	4358	4358	4358	4358	2868	4358	4358	4358	4358
R^2	0.439	0.511	0.529	0.521	0.480	0.447	0.474	0.494	0.562

Notes: Dependent variable = Natural log of hourly wages; other non-reported controls: year, region, education, establishment size, job tenure and its square, permanent contract, age and its square, marital status, part time work, job hours and its square, private sector and industry. Significance levels: *** 0.01, ** 0.05, * 0.1, + 0.2.

Fixed effects models

Results from FE estimation of models including SHC are presented in Tables 7 and 8. For men, the estimated coefficient on occupational feminization in the base model was –0.128. The addition of the SHC measures in models 2d and 2e reduces this to –0.087 and –0.091 respectively. Compared to the OLS specifications, the wage returns to SHC diminish and even disappear, suggesting that unobservables are important in allocating employees to different training schemes and influencing their SHC accumulation. More motivated and able employees have both higher wages and higher levels of SHC.

As for men, SHC has a smaller effect on the wages of women in the FE specifications than in OLS. However, some differences by sex emerge. Receiving on-the-job training in the past 12 months attracts a wage premium of 1.4% for women while it has no effect for men. A larger impact of SHC on wages for women than men also emerges in all other models. The estimated coefficients on occupational feminization for women are considerably less negative in models 2c, 2d and 2e (–0.098, –0.066 and –0.102) than in the base model (–0.167). Therefore, SHC reduces the negative effect of occupational feminization on the wages of women in the FE models, although occupational feminization continues to have a statistically significant effect.

Overall, findings from FE models are consistent with those from OLS specifications. The inclusion of SHC variables reduces the impact of occupational feminization on wages relative to the base model, but the coefficients on occupational feminization remain negative, large and statistically significant. Controlling for unobserved heterogeneity also reduces the effects of SHC on wages. This suggests that more able or motivated employees tend to have higher levels of SHC and also tend to work in better paid male-dominated occupations.

Table 7. Fixed effects estimates: Men.

	Model 2a	**Model 2b**	**Model 2c**	**Model 2d**	**Model 2e**
Occupational feminization	-0.128***	-0.135***	-0.124***	-0.087***	-0.091***
Received on-the-job training in the past 12 months	-0.002				
Proportion of trainees in the occupation		0.396***			
Modal training "less than a week"		*Ref. cat.*			
Modal training "more than a week but less than a year"		0.012			
Modal training "more than a year but less than three years"		-0.045***			
Modal training "more than three years"		-0.046***			
Modal training "indefinite, continuously"		0.013			
Low occupational specialization			*Ref. cat.*		
Medium occupational specialization			0.054***		
High occupational specialization			0.125***		
Level 1 of skill (lowest)				*Ref. cat.*	
Level 2 of skill				0.060***	
Level 3 of skill				0.090***	
Level 4 of skill (highest)				0.150***	
Average self-reported "specialist knowledge"					0.061***
N (observations)	26362	26362	26362	26362	9949
N (individuals)	3968	3968	3968	3968	2506
R^2 Within	0.299	0.303	0.311	0.308	0.163

Notes: Dependent variable = Natural log of hourly wages; other non-reported controls: year, region, education, establishment size, job tenure and its square, permanent contract, age and its square, marital status, part time work, job hours and its square, private sector and industry. Significance levels: *** 0.01, ** 0.05, * 0.1, $^{+}$ 0.2.

Table 8. Fixed effects estimates: Women.

	Model 2a	**Model 2b**	**Model 2c**	**Model 2d**	**Model 2e**
Occupational feminization	-0.167***	-0.151***	-0.098***	-0.066***	-0.102***
Received on-the-job training in the past 12 months	0.014***				
Proportion of trainees in the occupation		0.450***			
Modal training "less than a week"		*Ref. cat.*			
Modal training "more than a week but less than a year"		-0.009			
Modal training "more than a year but less than three years"		-0.105***			
Modal training "more than three years"		-0.145***			
Modal training "indefinite, continuously"		-0.012			
Low occupational specialization			*Ref. cat.*		
Medium occupational specialization			0.087***		
High occupational specialization			0.174***		
Level 1 of skill (lowest)				*Ref. cat.*	
Level 2 of skill				0.064***	
Level 3 of skill				0.135***	
Level 4 of skill (highest)				0.200***	
Average self-reported "specialist knowledge"					0.092***
N (observations)	29443	29443	29443	29443	11526
N (individuals)	4358	4358	4358	4358	2868
R^2 Within	0.219	0.228	0.238	0.232	0.134

Notes: Dependent variable = Natural log of hourly wages; other non-reported controls: year, region, education, establishment size, job tenure and its square, permanent contract, age and its square, marital status, part time work, job hours and its square, private sector and industry. Significance levels: *** 0.01, ** 0.05, * 0.1, $^{+}$ 0.2.

7. The contribution of occupational feminization to the gender pay-gap

We now calculate the contribution of occupational feminization to the gender pay-gap following procedures proposed by Tomaskovic-Devey (1995). The decomposition can be expressed as:

$$ABS\left(\frac{[(\overline{X}_F - \overline{X}_M)\beta_{ALL}] * (AW_{ALL})}{AW_{WOM} - AW_{MEN}}\right) * 100$$

where $\overline{X}_M$ and $\overline{X}_F$ are sample averages in occupational feminization for men and women respectively; β_{ALL} is the coefficient on occupational feminization in the pooled model; AW_{ALL}, AW_{MEN} and AW_{WOM} are sample average wages for all respondents, men and women respectively.

By definition, this is calculated using information from models pooling men and women and results are presented in Table 9.[23] Focusing only on models which control for SHC, our measure of occupational feminization explains between 15.6% (model 2d) and 51.6% (model 2b) of the gender wage-gap in OLS specifications. In the preferred FE specifications, estimates are lower and range between 15.9% (model 2e) and 24.9% (models 2a & 2b). Therefore, we conclude that occupational sex-composition accounts for between 15% and 25% of the gender wage-gap in Britain.

Table 9. The contribution of occupational sex-segregation to the gender pay-gap.

Model	OLS	FE
1	42%	24.9%
2a	42%	24.9%
2b	51.6%	24.4%
2c	37.5%	19.8 %
2d	15.6%	13.4%
2e	25.2%	15.9%
3	40.2%	N/A
4	35.7%	N/A
5	57.4%	N/A
6	18.1%	N/A

Notes: Based on decomposition methods used by Tomaskovic-Devey (1995).

[23] Estimated coefficients on key variables from these models are presented in Appendix 2.

8. Discussion and conclusion

The aim of this paper was to examine the impact of occupational sex-segregation on the wages of men and women in Britain, and to evaluate the role of other factors (specialized human capital in particular) in explaining this. Previous work on the US, Spain and Europe has suggested that the long-established negative association between occupational feminization and wages can largely be explained by SHC (Tam, 1997; Tomaskovic-Devey and Skaggs, 2002; Polavieja, 2007, 2008a, 2009). This casts doubts on the devaluation thesis which interprets the negative association between feminization and wages in terms of societal undervaluation of work traditionally performed by women.

Our estimates indicate a strong negative relationship between occupational feminization and wages. Moving from a completely male-dominated to a completely female-dominated occupation is associated with wage penalties of 17% and a 32% for men and women respectively in base OLS models. The estimated penalties are lower in FE models (13% and 17% respectively), suggesting that unobservables play an important part in allocating employees to occupations in relation to their sex-composition. Further investigation suggests that this relationship is not linear, which is inconsistent with the devaluation hypothesis. However, these non-linear effects are much less pronounced when accounting for unobserved differences across individuals.

The wage penalty associated with working in more female-dominated occupations remains in models that introduce controls measuring SHC, domestic labour supply, socialization, authority at the workplace and compensating differentials. Moving from a completely male-dominated occupation to a completely female-dominated occupation is associated with a significant wage penalty of 7% to 9% for men and women in the preferred FE specifications. Additionally, women receive wages that are 10% to 15% lower than otherwise similar men and the sex-segregation of occupations is found to account for at least 16% of the gender pay-gap.

The impact of SHC varies across measures. For example, introducing training based measures of SHC did not reduce the wage penalty for working in female-dominated occupations. This may be because women in Britain now undertake training almost as often as men, although for a shorter duration (Greenhalgh and Mavrotas, 1994; Green and Zanchi, 1997; Jones et al, 2008). The measure of SHC which has the largest effect on the impact of occupational sex-composition on wages is a skill-based subdivision of the SOC. It is widelyaccepted that a sex-bias in skill conceptualization and evaluation affects occupational classifications (Grimshaw and Rubery, 2007; Steinberg, 1990; Phillips and Taylor, 1980). Therefore, to the extent that such biases are embedded into this measure, it is possible that the effect of occupational feminization on wages is downward-biased in this specification. Finally, although they undertake training as often as men, women report lower levels of specialist knowledge required for the job. This may suggest that the training received in female-dominated occupations concentrates on more transferable skills. There may also be sex-differences perceiving or reporting specialist knowledge (Horrell et al, 1994; Correll, 2001).

Overall, results in this article provide support for the devaluation theory. The negative and statistically significant relationship between occupational feminization and wages remains in the presence of theoretically relevant controls and unobserved individual-specific effects, which contrasts with findings for other countries. Thus, the hypothesis that deep-rooted societal mechanisms contribute to the devaluation of the work performed primarily by women cannot be rejected. However, there is also support for the SHC thesis. Measures of SHC increase wages net of education, age, job tenure and other important drivers of pay, and also reduce the effect of occupational feminization on wages.

Further research might focus on identifying any relevant mechanisms in the British labour market that differ from those in the US, Spain or Europe, or on whether there are other factors such as educational segregation (e.g. in degree disciplines or vocational training programmes) that could explain the persistent relationship between occupational feminization and wages. Better, more appropriate measures of SHC could also be defined. It is important to determine to what extent on-the-job training, job specialization and job-learning time capture the same underlying process. There also seems to be some confusion in the literature as to how to differentiate the specificity of skills at the firm, job and occupation (only Tomaskovic-Devey and Skaggs (2002) expand on this). Future research should attempt to disentangle these three types of skills, their contributions to wages and their interaction with the sex-composition of occupations and its effect on wages. Researchers may also wish to explore the possibility that ‘nested’ and ‘independent’ occupations may coexist within the occupational structure in relation to the availability of defined career ladders and its implications for SHC theory. If some SHC is transferable across occupations, the empirical formulation of SHC as it stands may be flawed. Further attention should also be paid to the factors motivating the functional form of the relationship between occupational sex-composition and wages. This may be important, for instance, for the identification of target-groups in the development of appropriate policy responses.

Any methodological advances require more accurate and appropriate data to measure SHC. The Skills Surveys may help to clarify factors which mediate the observed relationship between occupational sex-composition and wages, for example, by offering insights into the nature, content and duration of on-the-job training and by providing information on the types and amounts of work-related skills in male- and female-dominated occupations. Studies making use of matched employee-employer data which account for the effect of other firm-level characteristics and for the relative contributions of industry, establishment and job sex-composition to the pay-gap would also be welcome.

The fact that both the sex-composition of an occupation and SHC are shown to affect wages implies that a multifaceted policy approach should be applied to reduce gender inequality in the labour market and narrow the gender pay-gap. Future legislation should continue supporting and promoting women's investments in SHC (e.g. statutory childcare and maternity leave policies, creation of flexi-time positions in male-dominated occupations, enhancing awareness of non-traditional career paths for young women, etc) while also considering comparable-worth strategies which ensure that work of equivalent value is equally remunerated regardless of the sex-composition of the workforce.

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10. Appendices

Appendix 1 – Complete list of regression coefficients for base OLS and FE models

Model	OLS Men	OLS Women	FE Men	FE Women
Occupational feminization	-0.171***	-0.321***	-0.128***	-0.167***
Year = 1991	*Reference category*			
Year = 1992	0.009	-0.024**	0.021*	0.018*
Year = 1993	-0.004	0.030***	0.037**	0.049***
Year = 1994	0.007	0.038***	0.059**	0.063***
Year = 1995	0.004	0.042***	0.069**	0.083***
Year = 1996	0.013	0.033***	0.094**	0.090**
Year = 1997	-0.015+	0.035***	0.100**	0.103**
Year = 1998	-0.004	0.037***	0.128**	0.119**
Year = 1999	0.016+	0.069***	0.176***	0.169***
Year = 2000	0.016+	0.070***	0.200***	0.185***
Year = 2001	0.037***	0.089***	0.236***	0.222***
Year = 2002	0.052***	0.102***	0.269***	0.246***
Year = 2003	0.059***	0.118***	0.286***	0.273***
Year = 2004	0.045***	0.116***	0.308***	0.285***
Year = 2005	0.049***	0.121***	0.326***	0.304***
Year = 2006	0.037***	0.125***	0.331***	0.312***
Year = 2007	0.034**	0.090***	0.353***	0.304**
Region = London	*Reference category*			
Region = South East	-0.054**	-0.122***	-0.005	-0.089***
Region = South West	-0.162***	-0.230***	-0.092+	-0.216***
Region = East of England	-0.133***	-0.245***	-0.042	-0.240**
Region = East Midlands	-0.244***	-0.227***	-0.076+	-0.139***
Region = West Midlands	-0.211***	-0.230***	-0.131*	-0.232***
Region = Midwest	-0.149***	-0.203***	-0.156***	-0.095*
Region = Yorkshire	-0.219***	-0.233***	-0.100*	-0.145***
Region = North East	-0.194***	-0.225***	-0.280***	-0.166**
Region = Wales	-0.250***	-0.249***	-0.120+	-0.174***
Region = Scotland	-0.210***	-0.190***	-0.230***	-0.132*
Age	0.067***	0.048***	0.071***	0.037***
Age squared	-0.001***	-0.001***	-0.001***	-0.000***
Marital status = Single	*Reference category*			
Marital status = Married or cohabitating	0.144***	0.021+	0.039***	0.042***
Marital status = Divorced, separated or widowed	0.094***	-0.035+	0.019	0.037**
Education = None	*Reference category*			
Education = O-Levels	0.141***	0.143***	-0.008	-0.015
Education = A-Levels	0.202***	0.196***	-0.020+	-0.034**
Education = HND/HNC diploma	0.265***	0.304***	0.002	0.030**
Education = First degree	0.481***	0.557***	0.138***	0.103***
Education = High degree	0.604***	0.622***	0.209***	0.199***
Job tenure	0.001	0.003+	0.003***	0.001+
Job tenure squared	-0.000	-0.000	-0.000	0.000
Has a permanent job	0.184***	0.085***	0.103***	0.048***
Works for the private sector	-0.003	-0.126***	-0.012+	-0.032***
Works part-time (<30 hours per week)	-0.323***	-0.142***	-0.259***	-0.124***
Job hours	-0.023***	0.006***	-0.022***	-0.007***

Job hours squared	0.000***	-0.000***	0.000***	-0.000***
Establishment size = Less than 25	*Reference category*			
Establishment size = 25 to 99	0.083***	0.068***	0.044***	0.041***
Establishment size = 100 to 999	0.147***	0.120***	0.082***	0.079***
Establishment size = 1000 or more	0.199***	0.151***	0.102***	0.092***
Industry = Public administration	*Reference category*			
Industry = Agriculture, agriculture, hunting, forestry, fishing, mining and quarrying	-0.131***	-0.137**	0.032	-0.008
Industry = Manufacturing	-0.010	-0.031^{+}	0.028^{+}	0.048***
Industry = Electricity, gas and water supply	0.080**	0.177***	0.086**	0.123*
Industry = Construction	0.053**	0.097***	0.022	0.050^{+}
Industry = Wholesale and retail sale	-0.112***	-0.131***	-0.051**	-0.020^{+}
Industry = Transport, storage and communication	-0.031^{+}	0.037^{+}	0.037*	0.032
Industry = Financing, insurance, real estate and other business activities	0.103***	0.111***	0.034*	0.045***
Industry = Hotels, restaurants and other personal and community services	-0.153***	-0.170***	-0.060***	-0.050***
Constant	0.651***	0.508***	0.969***	0.988***
N (observations)	26362	29443	26362	29443
N (individuals)	3968	4358	3968	4358
R^2 / R^2 Within	0.434	0.436	0.299	0.218

Notes: Dependent variable = Natural log of hourly wages. Significance levels: *** 0.01, ** 0.05, * 0.1, $^{+}$ 0.2.

Appendix 2 – Selected coefficients from pooled OLS and FE models

	Coefficient	M1	M2a	M2b	M2c	M2d	M2e	M3	M4	M5	M6
OLS	Occupational feminization	-0.253	-0.253	-0.311	-0.226	-0.094	-0.152	-0.242	-0.215	-0.346	-0.109
	Female	-0.134	-0.147	-0.154	-0.154	-0.129	-0.094	-0.138	-0.126	-0.112	-0.134
FE	Occupational feminization	-0.150	-0.150	-0.147	-0.119	-0.081	-0.096	N/A	N/A	N/A	N/A

Notes: Dependent variable = Natural log of hourly wages; other non-reported controls as in previous tables. All coefficients are significant at the 1% level.

Appendix 3 – Occupations in each level of skill

Skill level	SOC90	SOC2000
4	1a – Corporate managers and administrators	11 – Corporate managers
	2a – Science and engineering professionals	21 – Science and technology professionals
	2b – Health professionals	22 – Health professionals
	2c – Teaching professionals	23 – Teaching and research professionals
	2d – Other professional occupations	24 – Business and public service professionals
3	1b – Managers/proprietors in agricultural services	12 – Managers and proprietors in agriculture and services
	3a – Science and engineering associate professionals	31 – Science and technology associate professionals
	3b – Health associate professionals	32 – Health and social welfare associate professionals
	6a – Protective service occupations	33 – Protective services occupations
	3c – Other associate professional occupations	34 – Culture, media and sports occupations
	7a – Buyers, brokers or sales representatives	35 – Business and public service associate professionals
	9a – Other occupations in agriculture, forestry and fishin	51 – Skilled agricultural trades

	5b – Skilled engineering trades	52 – Skilled metal and electrical trades
	5a – Skilled construction trades	53 – Skilled construction and building trades
	5c – Other skilled trades	54 – Textiles, printing and other skilled trades
2	4a – Clerical occupations	41 – Administrative occupations
	4b – Secretarial occupations	42 – Secretarial and related occupations
	6b – Personal service occupations	61 – Leisure and other personal service occupations
	9b – Other sales occupations	71 – Sales occupations
	8a – Industrial plant and machine operators, assemblers	72 – Customer service occupations
	8b – Drivers and mobile machine operators	81 – Process, plant and routine operatives
		82 - Transport and mobile machine drivers and operatives
1	9b – Other occupations	91 – Elementary trades, plant and storage related occ.
		92 – Elementary administrative and service occupations

Notes: Adapted from Elias and McKnight, 2001, p.513.

Appendix 4 - Issues arising when appending occupation-level variables to individual respondents in a regression framework

The literature on the impact of occupational feminization on wages covers three broad types of studies. The first treats occupations as units of analysis using occupational feminization as an explanatory variable and average wages in the occupation as the dependent variable. Second, the most popular and established approach is to model individual-level data and include occupational-level terms denoting occupational feminization as a regular covariate. Finally, multilevel (ML) models in which the individual is the unit of analysis and which use occupations as second-order groupings have been suggested.

The literature on ML models has highlighted potential problems with appending occupation-level variables to individual respondents. It has been argued that conventional models do not take into account the fact that individuals are nested within occupations and that these occupations are a sample of an underlying distribution of occupations (de Ruijter and Huffman, 2003). Ignoring the hierarchical structure of the data causes the disturbances for individuals in the same occupation to be correlated, resulting in a violation of OLS assumptions and in biased estimates of the standard errors of the parameters in the model (Bryk and Raudenbush, 1992; de Ruijter and Huffman, 2003; Luke, 2004).

Most studies of the impact of occupational feminization on wages append the sex-composition term to individual respondents (*inter alia* Johnson and Solon, 1984; England et al, 1988; Sorensen, 1989; Kilbourne et al, 1994; Petersen and Morgan, 1995; Tam, 1997) while occupation-level analysis is not rare (e.g. Treiman and Hartmann, 1981; Parcel, 1989; England et al, 1994). On the contrary, ML modelling techniques, although promising and rapidly evolving have not been used in this area of research. In practice, the only study which compares OLS and ML models is Haberfeld et al (1998). Their main finding is that using ML models the impact of occupational feminization on wages is consistent with that found in previous research using other modelling strategies, since more female-dominated occupations pay less than more male-dominated ones. Therefore, they conclude that the widely used estimates are unbiased. De Ruiter and Huffman (2003) and de Ruijter et al (2003) reach similar conclusions using occupational sex-types as their sex-composition variable.

As a robustness check, we have estimated the base model using two-level ML models for each of the waves of the BHPS where the two levels are individuals and occupations and found that the negative effects of occupational feminization on wages are comparable to those of OLS and FE models. In consequence, we discard the use of ML models for the present paper and agree with Goldstein when he asserts that *"[I]n some circumstances*, […] *they may be hardly necessary, and traditional single-level models may suffice"* (2003, p.12).