European Sociological Review VOLUME 24 | NUMBER 2 | 2008 199–213

DOI:10.1093/esr/jcm043, available online at www.esr.oxfordjournals.org
Online publication 19 November 2007

The Effect of Occupational Sex-Composition on Earnings: Job-Specialization, Sex-Role Attitudes and the Division of Domestic Labour in Spain

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Important theoretical controversies remain unresolved in the literature on occupational sex-segregation and the gender wage gap. These controversies can be summarized as a debate between cultural-socialization arguments and economic or rational-action theories of specialization. The article discusses these theories in detail and carries out a preliminary test of the relative explanatory performance of some of their most consequential predictions. This is done by drawing on the Spanish sample of the second round of the European Social Survey (ESS). Empirical results suggest that the effect of occupational sex-segregation on wages could be explicable by workers' sex-role attitudes, their relative input in domestic production and the job-specific human-capital requirements of their jobs. Of these three factors, job-specialization seems clearly the most important one.

Introduction

In all OECD countries median wages for full-time employed men are higher than those for women. The average difference is around 15 per cent and exceeds 20 per cent in several countries (OECD, 2006). Various empirical analyses suggest that occupational sex-segregation could actually explain more than 90 per cent of this difference (Groshen, 1991; Tomaskovic-Devey, 1993; Petersen and Morgan, 1995; Petersen *et al*., 1997; Meyersson-Milgrom *et al*., 2001). In other words, it seems well-established that women earn less than men mainly because they are more likely to occupy jobs that entail lower rewards—whichever the sex of their incumbents. Two questions immediately follow from these stylized facts: (i) why are women sorted into particular jobs? and (ii) why do jobs mostly occupied by women tend to offer lower rewards than those mostly occupied by men? The first question refers to what Petersen and Morgan (1995) called *allocative* processes, whereas the second refers to what they called *valuative* processes.

Research on sex segregation and the gender wage gap has been very prolific both within sociology and economics. Yet and despite the large body of empirical evidence produced in the last three decades, crucial theoretical controversies remain unresolved to date (see e.g. Reskin, 1993, 257–260; Tam, 1997, 2000; England *et al*., 2000). One useful way of summarizing these theoretical controversies is seeing them as a conflict between two main competing views on the factors or mechanisms that govern *allocative* and *valuative* processes. The first view argues that both these processes can be explained mainly by socio-cultural factors. Men and women are socialized in different gender roles and this in turn shapes their occupational choices and career aspirations. Moreover, society at large, values differently the work performed by each sex so that the jobs mostly occupied by women are rewarded less, regardless of their intrinsic economic value. The second view, however, sees both *allocative* and *valuative* processes as mainly driven by purely economic determinants. This latter perspective stresses the importance of cost-benefit calculations, economic specialization, and, crucially, the consequences that human capital investments and contractual hazard might have for the allocation and valuation choices of both individual employers and employees. Hence, whilst the former view focusses on the existence of 'gendered' rationalities acquired through socialization processes, the latter seeks to explain both segregation and the gender wage gap by drawing on rational action principles of individual behaviour that are taken to be shared by all actors, regardless of their sex.[1]

It is apparent that not all theoretical disputes can be settled empirically. What empirical analysis can do is evaluate the relative explanatory performance of those indicators, cultural and economic, that *are* testable. The main contribution of this article lies in testing the empirical consequences of three factors that have always been central to the existing theoretical debates on the gender wage gap but have rarely been measured *directly* and *simultaneously*, namely, the specific human-capital requirements of jobs, individuals' sex-role attitudes, and their supply of domestic work. The empirical focus of this article is on the observed association between occupational sex-composition and individual earnings. The goal is to *explain away* this association by introducing the aforementioned indicators in the wage equations. Hence, although allocation and valuation processes are closely connected from a theoretical point of view, the empirical part of this study focusses *mainly* on the latter, and this is where the interpretation of the findings should *mainly* circumscribe to.

The empirical part of this article exploits the analytical possibilities of the second round of the European Social Survey (ESS) carried out in 2004 (Jowell and CCT, 2005). In order to eliminate institutional variation, this article focusses on one single country, namely Spain ($N = 1{,}663$). This allows us to concentrate on the micro-level mechanisms linking occupational sex-composition to earnings, whilst holding constant all the institutional macro-level effects that could be mediating *allocative* and *valuative* processes. Yet for this very reason, the empirical findings of this article should not be automatically extrapolated to other societies.[2]

The second round of the ESS seems to offer an extraordinary chance to fill the existing gaps in empirical research on the gender wage gap because it brings together a wealth of relevant indicators pertaining to the attitudinal, domestic, and productive spheres. Yet despite these very promising features it must be noted that ESS country samples are rather small for the purposes of earnings research. Small sample size is bound to yield large standard errors and hence wide confidence intervals around the parameter estimates.[3]

The article is organized as follows. In the first part, the main theoretical arguments of socio-cultural and economic perspectives are discussed. Hakim's preference theory is treated as a special case that stands somewhere in between these two main approaches. The second part of the article describes the methodology used in the empirical analysis, which focusses on explaining the observed association between occupational sex-composition and earnings. This section includes a discussion of the variables used, as well as of the imputation techniques and the selection-bias correction procedures applied. Empirical findings are then presented. These findings are discussed in a concluding section.

Theoretical Framework

Socio-Cultural Theories

Socio-cultural theories stress that values and stereotypes regarding women's and men's roles in society are carried over into the labour market (see e.g. England *et al.*, 1994; Crompton and Harris, 1997, 1998). As a result of these sex-role differences, which are acquired through socialization processes, women *on average* would be more likely to assign a greater value to home-caring than to pay-work and hence their job aspirations would be generally lower than men's. Women would consequently tend to self-select themselves into less demanding but also less rewarding jobs[4] (see e.g. Waite and Berryman, 1985; Vogler, 1994). Employers, on their part, would be prone to sex-typing and sex aversion/sex affinity in their contracting and promoting practices, hence incurring into more or less avert forms of gender discrimination. Employers' *gendered* tastes and the exclusionary practices that follow would also reflect their socialization in patriarchal values (see e.g. Bergmann, 1986; Reskin and Padavic, 1988; Goldin, 1990; Ridgeway, 1997). Both women's and employers' gender views would thus affect the allocation process.[5]

According to the so-called *cultural devaluation theory*, patriarchal attitudes and values can also affect the *valuative* process itself. Both experimental evidence as well as evidence from prestige surveys do indeed suggest that people tend to assign a lower value to work carried out mostly by women (Tam, 1997, 1655–1656). Sociologists endorsing the theory of cultural devaluation claim that such cultural bias also permeates the wage-setting process, making both employers and male-dominated trade unions prone to the pecuniary undervaluation of female-dominated jobs (see e.g. Reskin, 1988; England *et al.*, 1994; Kilbourne *et al.*, 1994). In support of such a claim, they draw on abundant survey evidence showing significant effects of the sex-composition of occupations on wages after controlling for a myriad of workers' characteristics that are intended to capture—and interpreted as capturing—workers' human capital (see e.g. Sorensen, 1990; England, 1992; Macpherson and Hirsch, 1995).

Empirical survey evidence in support of the cultural devaluation hypothesis follows a residual approach. That is, the existence of cultural devaluation is never tested directly, but only inferred indirectly from the size and statistical significance of the sex-composition coefficient in a multivariate context. The accuracy of the residual approach depends crucially on there being an adequate vector of explanatory variables in the underlying earnings function (Polavieja, 2005, 167). In other words, only if all the dimensions of individuals' human capital—as well as other economic factors influencing individual wages—are properly captured, one could interpret sex-composition effects as reflecting non-economic processes. Yet despite the crucial importance that controlling for economic factors has for these theories—let alone for economic theories themselves—human-capital variables are almost consistently miss-translated into operational indicators. One particular dimension of human capital that is typically disregarded in standard research is the specific human-capital requirements of jobs—i.e. of the amount of investments that employees have to make in order to learn to do their job well. To the extent that these investments are both linked to the degree of difficulty or effort required in the job and to the pecuniary compensation received, the specific human-capital requirements of jobs seem an obvious economic factor mediating the effect of occupational sex-composition on wages. It seems clear that without reliable indicators of this crucial dimension, the assumption that economic factors have been properly accounted for, which lies at the heart of the evidence in support of cultural devaluation hypothesis, seems unwarranted.

Economic Theories

The most influential economic theory of rewards is Beckers' human-capital theory. According to Becker (1993 [1964]), individuals are rewarded for the value of the additional productivity that investments in skills bring. Individuals acquire skills through schooling and on the job. Depending on their scope of applicability, skills are defined as general, industry-specific, occupational-specific, or firm-specific.

Rational employers have very few incentives to invest in skills that are not firm-specific, since nothing would prevent competing firms from poaching workers after a particular employer had incurred training costs. Employers will thus tend to shift the costs of training in general, industry-specific and occupation-specific skills to trainees themselves. This they will typically do by subtracting the costs of this type of training from employees' wages during the training period (Becker (1993 [1964], 34–35). Conversely, workers will be reluctant to bear with the costs of firm-specific specialization, since the skills that it provides have no value outside the firm. In order to ensure the adequate supply of firm-specific skills, rational employers will offer wage premiums that compensate for the additional costs of specialization. The wage premium for firm-specific training will reduce employees' quitting rates and consequently employers' risks of losing their investments. Hence firm-specific human-capital investments are expected to be positively correlated with earnings and inversely correlated with turn-over rates (Becker (1993 [1964], 44–49).

Most human capital acquired on the job actually consists of a mixture of general, industry, occupational and firm-specific components and hence its costs should be borne by both employers and employees, although in different degrees depending on which dimension predominates. In what follows I will refer to all types of human capital that is acquired on the job as job-specific human capital (JSK).

The crucial importance that JSK investments have in determining individual earnings has also been recognized by economic theories that depart from the human-capital approach. As such, both efficiency wage theories (Shapiro and Stigliz, 1984; Akerloff and Yellen, 1986), personnel economics (Milgrom and Roberts, 1992; Lazear, 1995) and transaction cost models of the employment contract (Williamson, 1985, 1996) stress the pecuniary effects of job specialization. The latter two approaches have, in turn, had a clear impact on rational-action theories of class and the employment contract (Sorensen, 1994; Breen, 1997; Goldthorpe, 2000, ch. X; Sorensen, 2000; Polavieja, 2003). All these theories see wages, not just as a mere reflection of individuals' human capital, but rather as an incentive device that is designed and implemented by employers to reduce contractual hazard in a context of asymmetric information.

A crucial implication of both human capital and contractual-hazard models is that the returns to job-specialization are linked to tenure so that employees must remain in their firms in order to recoup their job-specific investments.[6] Rational employees anticipating discontinuous careers will thereby be reluctant to incur such specialization investments. Similarly reluctant will employers be to invest in employees who they consider likely to quit their jobs. Given the existing distribution of childcare and domestic tasks, women will be *on average* more likely to experience job disruptions and hence more likely to choose, or be chosen for, jobs that have lower job-specialization requirements. There is both indirect and direct evidence showing important differences in job-specialization by sex.

Indirect evidence of the crucial importance of job-specialization in accounting for the gender wage gap was provided in Polavieja (2005). Drawing on the 1995 Spanish Survey on Wage Structure, it was found that gender wage gaps were significantly smaller in so-called service-class occupations, characterized by high levels of job-specialization and where the returns to tenure were shown to be higher.[7] The logical implication of these findings was that if women entered jobs requiring high levels of specialization in larger numbers, the wage gap in Spain would be reduced. It was also argued that the main factor increasing the costs of job-specialization for women was the existing distribution of domestic tasks (Polavieja, 2005, 171). Yet job specialization was not observed directly,[8] whilst the division of domestic labour was not observed at all. In both cases this was due to data limitations. Results were, therefore, considered as preliminary until direct indicators of these two crucial dimensions could become available (Polavieja, 2005, 176).

A more direct test of the crucial importance of job-specialization for the gender wage gap can be found in Tam (1997). Using cross-sectional data from the 1988 U.S. Current Population Survey, Tam showed that the observable impact of occupational sex composition on wages disappeared entirely once both information on the average length of specific training required in respondents' occupations and a set of industry dummies were introduced in the wage models. Tam interpreted his findings as evidence against the existence of a cultural devaluation of women's jobs and in support for the standard human capital model. Tam's contribution generated an important debate in the sociological field (see e.g. England *et al*., 2000; Tam, 2000; Tomaskovic-Devey and Skaggs, 2002). His approach illustrated the analytical pay-offs of getting closer to measuring the exact skill requirements of people's jobs.

Tam focussed only on valuation processes, but the human-capital approach that he endorsed also includes a theory of job allocation. The fullest development of the human-capital theory of allocation is contained in Becker's analysis of the family (Becker, 1981, 1985). The main thrust of Becker's model is considering that investments in household and labour market specialization produce increasing returns and thereby provide a strong incentive for a division of labour. Specialization would lead to men investing comparatively more in the labour market and women more in household production. According to Becker, this is because women have an intrinsic competitive advantage in the domestic sphere, which stems primarily, but not solely, from childbearing (Becker, 1981, 21–25, 1985, 41). As a result, working women would rationally economize on labour-market effort by seeking jobs that require less human capital investments (see also Polacheck, 1981).

Again, it is not necessary to endorse Becker's theory of the family in order to expect rational women choosing jobs that entail lower investments in JSK.[9] All that is required is that women anticipate discontinuous employment careers, regardless of whether higher risks of discontinuity for women are seen as the result of sphere-specialization or as a structural constraint (see below).

The Importance of Individual Work Orientations and Sex-Role Preferences: Hakim

Somewhere in between socio-cultural and economic perspectives, stands Hakim's view of women as 'self-determined' actors (Hakim, 1991, 1995, 1996a, b, 2000, 2003a). For Hakim, women's individual work orientations and preferences play a key role in determining their labour-market outcomes. These orientations are not only different *on average* to those of men, but also internally heterogeneous, reflecting a much higher degree of personal choice or 'agency' than standard cultural-socialization arguments typically concede.

Hakim differentiates between three qualitatively different types of women: '*work-centred*', '*home-centred*', and '*adaptative*' women and argues that these differences in core preferences are responsible for a large share of the observed sex-differences in labour-market outcomes. Yet, and despite the centrality that preferences play in her model, to date Hakim has not provided an explanation of the sources that lead to preference heterogeneity amongst women. Preference theory is about the '*historical context in which core values become important predictors of behaviour*' (Hakim, 2003a, 355), not about the causes of core-value differentiation. If not the product of sex-role socialization, as in socio-cultural explanations, nor of economic specialization (i.e. rationality), as in standard economic arguments, what makes women differ in their work orientations? Women's preferences are seen by Hakim mainly as reflecting women's *agency*. That is why her approach must be differentiated from both socio-cultural and economic theories. Yet this is also why, in practice, it is often hard to identify which empirical predictions could differentiate Hakim's model from the predictions that stem from these two alternative approaches.

Data and Methodology

ESS data for the Spanish sample ($N = 1{,}663$) has been used to analyse the different economic and attitudinal factors that mediate the empirical correlation observed between occupational sex-composition and individual earnings. The analytical strategy adopted is based on nested equations within a model-building framework. Although this strategy also tackles allocation issues, the analytical stress is on valuation.

The basic approach followed here is akin to that of Tam (1997). The goal is to *explain away* the impact of occupational sex-segregation on wages by introducing in the earnings function theoretically driven indicators that could mediate between the two. In the case of Tam's influential paper, he used externally imputed measures of JSK based on the U.S. Dictionary of Occupational Titles (DOT) (Tam, 1997, 1670, 1675). I will be using both an indicator of JSK, an indicator of respondents' gender attitudes and an indicator of respondents' relative supply of domestic work and test their relative impact on wages, as well as their effect on the occupational sex-composition coefficient. Occupational sex-composition is calculated as the fraction of workers in respondents' occupation that are women, measured using the 4-digit ISCO classification. This baseline measure is complemented with 3-digit ISCO information for the 4-digit cells containing 0 number of women.[10]

The JSK indicator used here is based on respondents' own assessments of the time that it would be required for somebody with the right qualification to do respondents' job well. This self-reported estimate of JSK seems closer to the original theoretical concept than the externally imputed values used by Tam because it refers directly to respondents' jobs net of general human capital requirements, whereas DOT values are necessarily based on occupational-level information and do not distinguish clearly between the job-specific and the general human capital content of occupations.[11] It must, however, be noted that self-reporting introduces subjectivity, which could bias the results if incumbents' sex had a systematic impact on job evaluations. This possibility must be taken seriously as empirical research suggests that women could underestimate their capabilities (Correll, 2001). Note, however, that the direction of this hypothetical bias will presumably depend on the sex-composition of jobs. This is because of the very wording of the question used in the ESS. When asked how long it would take for *somebody* with the right qualification to learn to do respondent's job well, respondents will most probably think of that someone as a woman if they are employed in a female-dominated job and as a man if they are employed in a male-dominated job. If biased women consider men more capable than themselves, they will tend to report longer learning periods than actually required in the former case but shorter learning periods in the latter.[12] Self-reporting bias could thus reduce the capacity of JSK to absorb the association between sex-composition and earnings and hence produce lower-bound estimates of JSK.

Gender-role attitudes are measured using a battery of ESS questions from which an index of traditional sex-role attitudes has been constructed. The index combines the responses to the following 5 Likert-type items: (i) whether women should be prepared to cut down on their wages for the sake of their families, (ii) whether men should have equal domestic responsibilities as women, (iii) whether men should have preference over scarce jobs, (iv) whether parents should stick together for children even if they do not get along, and (v) whether a person's family should be his/her priority. The index shows a Cronbach's alpha of 0.7, it is normally distributed and ranges from 0 to 20, the latter value implying the highest score in 'traditional' gender attitudes. I take this index to represent general sex-role attitudes or ideologies acquired through socialization processes. The extent to which the index captures core individual preferences is admittedly open to discussion (Hakim, 2003b and below).

The relative amount of domestic work supplied by respondents (DS) is referred to domestic tasks such as cooking, washing, cleaning, shopping, property maintenance and the like, not including childcare nor leisure activities. The fact that childcare is not included allows to maximize the number of observations. DS is measured using respondents' self-reported estimates of the amount of time that they spent on such activities on a typical weekday relative to the total amount of time spent by all the people living in their households.[13] It consists of a 6-interval scale ranging from '*none or almost none*' to '*all or nearly all of the time*'.

Another significant methodological difference with respect to Tam's approach is that I do not fit separate equations for men and women. This is because the Spanish ESS sample is rather small to begin with (N = 1,663) and gets much further reduced when restricted to wage earners for whom there is direct (N = 390) or imputable information on wages (N = 699). The analysis will therefore be based on wage equations that are fitted to a pooled sample of men and women, including obviously a sex-specific intercept, which is the standard approach for small samples (see e.g. Tomaskovic-Devey and Skaggs, 2002). Missing values have been dealt with using specific regression-based imputation methods. Self-selection bias has also been tackled using the Heckman method of estimation (Heckman, 1979). The use of these techniques, which are explained below, constitutes the final methodological departure from Tam's approach.

The dependent variable of our analyses, y, is the logarithm of gross hourly wages before deduction for tax and/or insurance in Euros. According to the ESS, the mean gross-pay for Spaniards in 2004 was 10€ per hour—i.e. 1,500€ per month. The crucial parameter in focus is the proportion of women in respondents' occupation (S). The analytical strategy implemented requires five different equations and is developed as follows. First, a linear regression is fitted to the ESS data in order to estimate the effect of respondents' sex (f) on wages (y), after controlling for standard demographic and worker characteristics, represented by vector (X). In accordance with standard practice, vector X includes total working experience, years with current employer, years of schooling completed, plus two other control variables, namely, firm's size, and region of residence (Equation 1). Secondly, I estimate the effect, both on y and on f, of the proportion of women employed in respondents' occupation (S), controlling for X (Equation 2). Once S is estimated, the goal is to explain away its effect. The hypothesis behind this goal is that, contrary to the expectations that follow from cultural-devaluation theory, occupational-sex segregation has no residual impact on wages, once the appropriate variables are controlled for. This hypothesis is tested by introducing both our JSK indicator (Equation 3) and, departing from Tam's strategy, also the index of sex-role attitudes (Equation 4) and respondents' self-reported relative supply of domestic work (Equation 5). The impact of sex-role attitudes on both S and y is tested through an interaction between gender attitudes and respondents' sex (ID*f), since it is expected that traditional gender values only depress wages in the case of women, but not of men. Similarly, it is expected that the relative supply of domestic work has an effect on the parameters in focus only in the case of people living in partnership. Hence an interaction between respondents' residential status (0 = living out of partnership; 1 = living in partnership) and their relative domestic supply is estimated (hh*DS). Practical analyses show that in order to properly test for the statistical effect of this interacted term, it is also necessary to control for a further interaction between residential status and respondents' sex (hh*f). This is because Spanish women living alone earn significantly less than men living alone and this must be accounted for in order to isolate the (conditional) effect of relative domestic supply on y. Formally, therefore, the following five models are estimated. The actual estimation technique applied is explained below:

$$y = f_1(\mathrm{f}, \mathrm{X}) \tag{1}$$

$$y = f_2(\mathrm{f}, \mathrm{X}, \mathrm{S}) \tag{2}$$

$$y = f_3(\mathrm{f}, \mathrm{X}, \mathrm{S}, \mathrm{JSK}) \tag{3}$$

$$y = f_4(\mathrm{X}, \mathrm{S}, \mathrm{JSK}, \mathrm{ID}^*\mathrm{f}) \tag{4}$$

$$y = f_5(\mathrm{X}, \mathrm{S}, \mathrm{JSK}, \mathrm{ID}^*\mathrm{f}, \mathrm{hh}^*\mathrm{DS}, \mathrm{hh}^*\mathrm{f}) \tag{5}$$

Expected results are: $|bs_{f2}| > |bs_{f3}| > |bs_{f4}| > |bs_{f5}|$, where bs are the unstandardized parameter coefficients of S and the subscripts refer to the number of the equation fitted to the ESS data. Standardized beta coefficients will also be reported for the best fit model. All the variables used in the statistical analyses are described in Table 1.

Missing Values and Self-Selection as Potential Sources of Bias

High levels of non-response to the earnings question have been dealt with as follows. First, a multivariate logit analysis of the relative probability of non-response has been carried out in order to evaluate the potential biasing impact of missing values.

Table 1 Description of variables

Variable	Description	*N*	Mean or %	Standard deviation
Y	Is the log of the ratio of gross weakly earnings and usually weekly hours in €. Missing values have been restored by imputation.	699	2.02	0.53
tenure	Years with current employer	677	10.44	9.82
schooling	Years of schooling completed	1,629	11.15	5.52
experience	Total number of years in paid work	609	19.46	15.20
S	Proportion of female in respondent's occupation (ISCO-4 digits complemented with ISCO-3)	1,185	0.48	0.30
ID	Index of (traditional) gender role attitudes	1,663	9.58	3.63
f	Sex of active respondents	868		
	Male	520	59.91%	
	Female	348	40.09%	
JSK	Time that would be required for people with the right qualification to learn to do R's jobs well, measured using an interval scale ranging from 1 = less than a week, to 8 = more than 2 years	699	3.4	1.03
DS	Proportion of weekly housework typically provided by R, measured using an interval scale ranging from 1 = none to 6 = all	1,663	3.23	2.05
hh	Residential status of Rs	1,650		
	Not living in partnership	651	39.45%	
	Living with partner/spouse	999	60.55%	

Source: European Social Survey, Spanish sub-sample (2004).

The distribution of non-responses to the earning question shows virtually no interpretable structure and none of the main parameter estimators used in the five wage equations noted earlier seems to show any significant association with the probability of non-response.[14] This simple test already suggests that there is a very high random component in non-responses, from which a low potential biasing impact can be inferred. This finding lends support to using imputation techniques based on the restricted sample of actual responses to avoid the loss of an unacceptable number of observations (Goldstein, 1996). Imputation has been performed by best sub-set regression[15] (StataCorp., 2003, 120–125). Missing values have been filled in by using the predicted values and the standard errors estimated using the following equation:

$$\begin{aligned}\mathrm{Ln}(y)^{\mathrm{sex}} = {} & b_0^{\mathrm{sex}} + b_1^{\mathrm{sex}}\mathrm{Experience} + b_2^{\mathrm{sex}}\mathrm{Tenure} \\ & + b_3^{\mathrm{sex}}\mathrm{Tenure}^2 + b_4^{\mathrm{sex}}\mathrm{Schooling} + b_5^{\mathrm{sex}}\mathrm{JSK} \\ & + b_6^{\mathrm{sex}}\text{Size of establishment} + b_{Xj}^{\mathrm{sex}}\mathrm{EGP}_{Xj} \\ & + b_{Zj}^{\mathrm{sex}}\mathrm{Residence}_{Zj} + \mathrm{e}_i\end{aligned} \tag{6}$$

where JSK is the job-specialization indicator commented on above, subscript X_j refers to a set of parameter-estimates for the 5-value version of the Goldthorpe class schema ($j = 5$–1) (Erikson *et al.*, 1979) and subscript Z_j denotes the set of estimate dummies for region of residence ($j = 18$–1). The rest of the variables are self-explanatory. The sex superscript denotes that this equation has been fitted separately by sex.

The five earnings functions noted earlier have been fitted both to the restored sample as well as to the original sample restricted to actual respondents to the earnings question. Results are highly comparable across samples for all earnings functions and the parameter estimates of the best fit models are extremely similar regardless of the sample used. Confidence intervals are, however, narrower in the restored sample, which is presented in the findings section of the article. Results using the restricted sample are available on request.

A further source of potential bias is self-selection. Self-selection is known to bias estimates for wage equations because women's choices regarding whether or not to work are not made independently of the market wages offered. As such, we only observe wages for a particular self-selected group of women. Heckman (1979) proposed a method of estimation that deals with self-selection. The main thrust of this method, put in terms of our own research question, is to note that wages are jointly determined, not only by the variables that are captured in the standard wage models, but also by those affecting the decision to participate in paid work. Individuals compare market wages (y_i) with their reservation wage (y_{ri}) and only choose to work if $y_i > y_{ri}$. Market wages can be expressed by the regression equation:

$$y_i = X_i\beta + e_{1i} \text{ (regression equation)} \qquad (7)$$

where y_i is the hourly wage of individual i; X_i is a vector of variables affecting his/her wages; β is a vector of parameters to be estimated, and e_{1i} is the unexplained/unobserved component with $e_{1i} \sim N(0,\sigma)$. Yet y_i is only observed if:

$$Z_i\gamma + e_{2i} > 0 \text{ (selection equation)} \qquad (8)$$

where Z_i is a vector of variables affecting the decision to work (i.e. the reservation wage); γ is a vector of parameters to estimate; e_{2i} is a random variable with $e_{2i} \sim N(0,1)$ that captures unobserved characteristics affecting such decision.

It is assumed that e_{1i} and e_{2i} are jointly normally distributed and have correlation ρ. If $\rho \neq 0$, standard equation techniques applied to the first equation will yield biased results. Heckman (1979) shows that it is possible to obtain consistent estimates of β if the regression equation and the selection equation are estimated jointly. I have used Equations 1–5 above as the regression equations. For each and all of these models the following selection equation has been jointly estimated (variables are self-explanatory):

$$\begin{aligned}&\gamma_0 + \gamma_1\text{sex} + \gamma_2\text{married} + \gamma_3\text{married} * \text{sex}\\&\quad + \gamma_4\text{schooling} + \gamma_5\text{schooling} * \text{sex} + \gamma_6\text{age}\\&\quad + \gamma_7\text{age} * \text{sex} + \gamma_j\text{residence}\\&\quad + \gamma_j\text{residence} * \text{sex} + e_{2i} > 0 \qquad (9)\end{aligned}$$

Findings

Table 2 shows the results of fitting the five Heckman-estimation equations described above on the ESS imputed data for Spanish wage earners. Equation 1 is a standard human-capital earnings function that yields an 11 per cent wage penalty for women, which is a somewhat lower estimate than those usually reported in the literature (see e.g. Lago, 2002). Yet if the Heckman-selection estimation technique was not applied, the estimated coefficient would be higher (0.15) and in accordance with recent estimations (Polavieja, 2005; Amuedo-Dorantes and De la Rica, 2006). The ρ parameter is significant, which indicates that a standard human-capital model would indeed suffer from self-selection bias.

Equation 2 adds occupational sex-composition (S) to the previous model. Note that this takes up all the effect of respondents' sex on y, a finding that seems to suggest that within-job discrimination plays virtually no role in explaining sex-differences in earnings, which is in line with previous research (see e.g. Meyersson-Milgrom *et al.*, 2001). The unstandardized coefficient estimated for S, bs_{f2}, is –0.19. Since y is logged, this would mean that individuals employed in fully female occupations earn on average 19 per cent less than those employed in fully male ones. Large margin errors advise caution in the interpretation of this figure. Note, however, that this is the type of result that has conventionally been used in support for the cultural devaluation hypothesis. The argument is simple: if equally endowed individuals are rewarded differently depending on the female composition of their occupations is because there is a process of cultural undervaluation of female jobs, which is carried over into the wage-setting process. As explained earlier, this interpretation rests on the assumption that all possible factors affecting the relationship between individuals, jobs, and wages are controlled for. But is this really the case?

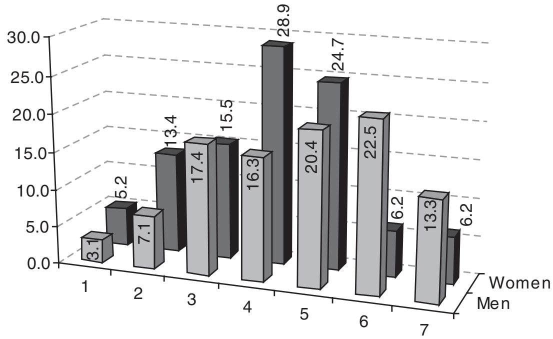
It seems not. The ESS shows very significant differences in the specific human capital requirements of jobs by sex. For instance, the proportion of employed respondents with tertiary education occupying jobs requiring learning periods longer than 1 year is 36 per cent amongst men, but drops to only 12 per cent in the case of women (Figure 1). These are indeed notable differences. Hence it is not surprising that when job-specialization (JSK) is introduced in the earnings function, a significant drop in both the coefficient and significant levels of S is observed, bs drops from –0.19 (Equation 2) to –0.09 (Equation 3) and becomes non-significant ($P > |t| = 0.15$). As in the case of Tam (1997), the introduction of JSK seems to explain away the effect of occupational sex-composition on earnings.

Equation 4 adds sex-role attitudes (interacted with respondents' sex) to the previous model and this reduces the effect of occupational sex-segregation on wages further ($bs_{f4} = -0.07$; $P > |t| = 0.25$). The interaction shows that traditional sex-role attitudes have no impact for men, but significantly reduce women's earnings. Moreover, when sex-role attitudes are

Table 2 Heckman regression models on the log of gross hourly wages. Imputed values for missing cases on output variable. Spain (2004)

	Model 1	Model 2	Model 3	Model 4	Model 5	
Variables	***b***	***b***	***b***	***b***	***b***	b^s
Female	−0.11**** (0.03)	−0.04 (0.05)	−0.03 (0.04)	(see f below)	(see f below)	
Experience	0.015**** (0.002)	0.015**** (0.002)	0.015**** (0.002)	0.014**** (0.002)	0.012**** (0.002)	0.25
Tenure	0.018*** (0.006)	0.017*** (0.006)	0.013*** (0.005)	0.013*** (0.005)	0.013** (0.01)	0.24
$(\text{Tenure}/100)^2$	−0.05**** (0.01)	−0.05**** (0.01)	−0.04**** (0.01)	−0.04*** (0.01)	−0.04**** (0.01)	0.25
Years education	0.05**** (0.003)	0.05**** (0.003)	0.046**** (0.003)	0.044**** (0.003)	0.044**** (0.003)	0.44
S (P female in R's occupation)		−0.19*** (0.07)	−0.09 (0.07)	−0.07 (0.06)	−0.06 (0.07)	−0.03
JSK (T required to learn R's job)			0.080**** (0.01)	0.078*** (0.01)	0.078*** (0.01)	0.078
*ID(sex-role attitudes)*f(female)*						
f				−0.08 (0.05)	(see f below)	
ID				−0.003 (0.005)	−0.004 (0.006)	−0.03
f*ID				−0.02*** (0.008)	−0.02*** (0.008)	−0.10
*hh(residential status)*DS*						
DS (Effect of P household work provided by R for Rs living out of partnership)					0.02 (0.011)	0.06
hh*DS (Difference in the effect of DS for Rs living in partnership)					−0.04** (0.017)	−0.10
*hh (residential status)*f (female)*						
f (Effect of f for non-partnered Rs)					−0.15** (0.058)	−0.14
hh (Effect of living in partnership for men)					−0.02 (0.01)	−0.005
f*hh (Difference in the effect of being in partnership for women)					0.14** (0.066)	0.11
Constant	1.03****	1.01****	0.74****	0.75****	0.78****	
Heckman's ρ	−0.23*** (0.09)	−0.21** (0.09)	−0.20** (0.10)	−0.16 (0.12)	−0.17 (0.11)	
N=	1,121	1,121	1,121	1,121	1,121	
Censored *N*=	540	540	540	540	540	
Uncensored *N*=	581	581	581	581	581	
Prob > F	0.0000	0.0000	0.0000	0.0000	0.0000	
Wald test of independent equations ($\rho = 0$): $\text{chi}^2(1)$=	0.0153	0.0292	0.0529	0.1895	0.1215	

Notes. All models control for size of establishment and autonomous community of residence. b = unstandardized coefficients. b^s *= Standardized coefficients. Heteroscedasticity-robust standard errors in parenthesis. Imputations on outcome variable made on the basis of Equation (6) above. Selection equation is Equation (9) above. ID has been centred to the mean so that it now ranges from −9 to 11.*
****Significance ≤0.001; ***Significance ≤0.01; **Significance ≤0.05; *Significance ≤0.1
Source. Calculated by the author from European Social Survey, Spanish sub-sample (2004).



Legend: 1= 1 day or less; 2= 2-6 days; 3=1-4 weeks; 4=1-3 months; 5= more than 3 months up to 1 year; 6= more than 1 year up to 2 years; 7= more than 2 years.

Figure 1 Self-reported job-specialization by gender. Employed respondents with tertiary education, Spain (2004). *Source:* ESS, Spanish sub-sample (2004)

introduced in the earnings function, the Heckman self-selection coefficient also loses its statistical significance ($\rho = -0.16$; $P > |t| = 0.19$). This indicates that sex-role attitudes are affecting (negatively) both the probability that women enter paid employment and the wages they obtain if they choose to do so and this is why, once attitudes are controlled for, self-selection bias disappears for the analysed sample. These findings point in the direction of some sort of self-selection process whereby working women holding traditional sex-role views are more likely to (i) stay at home and (ii) choose particular jobs that entail both a higher proportion of women and lower monetary rewards.[16] Yet both the direct earning consequences of sex-role attitudes (for women) and their effect on the relationship between sex-composition and wages seem rather small in comparison to sex-differences in job-specialization (see standardized coefficients in Equation 5).

Finally, Equation 5 introduces respondents' share of total housework hours typically provided during a week interacted with residential status (hh^*DS), plus an interaction between residential status and respondents' sex (hh^*f). As explained above, the latter interaction is a rather technical requisite that stems from the data structure. It allows us to isolate the effect of hh^*DS on wages, given the significantly lower earnings displayed by women living out of partnership. Equation 5 shows that the distribution of domestic work is significantly associated with the earnings of respondents' living in partnership: the more unequal this distribution, the lower the earnings.[17] Note also that Equation 5 shows that the relative supply of domestic work seems to affect wages *irrespectively* of gender attitudes. In other words, the hourly earnings of people with different sex-role views seem equally affected by their relative input in the domestic production function. Equation 5, therefore, predicts that all individuals living in partnership, regardless of their sex-role attitudes/preferences, would see their wages depressed if they had to bear with a disproportional share of domestic work.[18]

It is well-known that it is invariably women who do bear with most of the housework burden. According to the ESS, half of all full-time employed Spanish women living in partnership report doing more than 3/4 of all the housework, whereas nearly 70 per cent of all employed married or cohabiting men admit doing less than 1/4 of it. Moreover, the housework burden seems unalleviated by sex-role attitudes as 44 per cent of full-time employed married or cohabiting women holding non-traditional sex-role views still report doing more than 3/4 of domestic work.[19] The Pearson correlation coefficient between sex-role attitudes and the relative supply of domestic work for employed women living in partnership is only 0.12. So it seems that the impact of domestic supply on earnings has very little to do with gender values, as captured by the sex-role attitudinal scale. This finding casts doubts on socio-cultural arguments linking general attitudes to the actual household burden.

The striking sex-differences in domestic-work supply found amongst Spanish wage-earners also cast doubts on economic-specialization arguments *a la* Becker. According to the Spanish sub-sample of the ESS, about 40 per cent of full-time employed married or cohabiting women with tertiary education still report doing more than 3/4 of domestic work at their households (Figure 2). Moreover, nearly 70 per cent of them consider that there are so many things to do at their homes that they often run out of time before they get them all done. So it seems that this observed domestic-supply imbalance might not only have negative consequences for women's earnings, as suggested by Equation 5, but also seems quite inefficient in terms of the domestic production function.

Discussion

The study of sex-segregation and the gender wage gap has been typically hampered by the paucity of survey indicators. The 2004 round of the ESS overcomes such limitations as it includes a very exhaustive list of theoretically relevant survey questions. But this is unfortunately done at the price of small country sample sizes. Observations shrink further when analysing earnings. As a result, and despite the plethora of

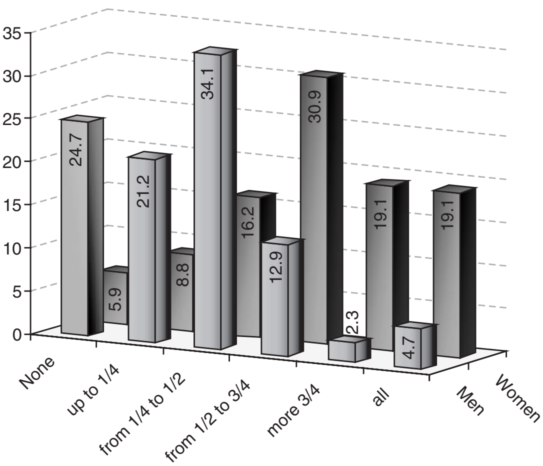


Figure 2 Proportion of domestic work supplied at the household. Full-time employed married or cohabiting respondents with tertiary education, Spain (2004). *Source:* ESS, Spanish sub-sample (2004)

indicators at hand, empirical findings based on the ESS country samples yield large standard errors and hence very wide confidence intervals. Empirical results should, therefore, be interpreted cautiously. It is with caution that the following conclusions can be drawn.

The ESS data for Spain lends little support to the cultural devaluation hypothesis. As Equations 3–5 show, job and supply-side characteristics absorb all the effect of occupational sex-composition on earnings. This suggests that *valuative* processes—i.e. the processes that link individuals (in jobs) to rewards—are not driven by employers' discriminatory tastes. The extent to which employers' discriminatory tastes play a role in *allocative* processes—i.e. the processes that link individuals to jobs—remains, however, an open question that the ESS data cannot answer.

Although the main analytical focus of this article has been on *valuation* processes, the joint-estimation techniques applied can also shed some light on the supply-side mechanisms of allocation. Results reported in Equation 4 suggest that sex-role attitudes are associated with women's self-selection into paid employment, as well as with both the occupational sex-composition of their jobs (when they work) and their earnings. These results seem to suggest that there is a *certain* degree of *attitudinal consonance* involved in *allocative* processes, from which indirect *valuative* consequences could follow. Attitudinal consonance seems consistent with socio-cultural theories of allocation.

Yet there is very little empirical association between the sex-role attitudes of employed women and their domestic workload. Non-traditional women doing paid work also end up doing most of the work at home. Moreover, the unbalanced division of domestic work by sex observed in Spain could have negative consequences for women's earnings, as suggested by the estimates reported in Equation 5. When it comes to housework, what we observe is, therefore, a high degree of attitudinal 'dissonance' amongst Spanish working women. This finding seems less consistent with socio-cultural explanations—unless it is assumed that sex-role attitudes are only weakly linked to personal preferences.[20]

The evidence on relative domestic supply and its effects on earnings also cast doubts on pure economic-specialization arguments. The ESS data shows very high levels of domestic supply unbalance by sex amongst Spaniards with tertiary education. It is hard to see why it should be economically rational that highly educated working women end up bearing with such a disproportional share of the domestic workload, particularly when it has been observed that such division could have negative consequences for their earnings. Empirical findings seem, therefore, more consistent with the idea that the unequal distribution of housework by sex acts as a *structural constraint* that could hinder women's career progression and from which earning consequences could follow. This idea was already defended in Polavieja (2005) but could not be tested then due to data limitations.

Yet the most important variable mediating the association between occupational sex-composition and earnings seems to be job specialization. Of the three variables tested, job specialization has the largest impact on earnings and is the only one that can absorb by itself all the statistical effect of occupational sex-segregation.[21] This finding is in line with those reported by Tam (1997, 2000) for the U.S., whilst it complements those reported in Polavieja (2005) for Spain in a very consequential way, as it provides the *direct* test that was hitherto lacking and which was then called for (Polavieja, 2005, 176).

There seems to be now sufficient accumulated evidence to argue that job specialization plays a *crucial* role in the gender wage gap.[22] Investigating sex-differences in access to job specialization seems, therefore, an obvious and promising avenue for further research on this topic.

Notes

1. Admittedly, reducing all the theoretical controversies that populate the literature on gender

segregation and the gender wage gap to a single debate between socialization and economic rationality is a simplification. I believe, however, that this simplification points indeed to a real debate that lies at the heart of the most consequential theoretical controversies.

2. Research currently under way is exploiting the comparative potential of the ESS using the whole pool of countries (Polavieja, 2007).

3. Pooling together all the country samples of the ESS does very little in the way of reducing uncertainty levels around crucial parameter estimates as all country samples are small for the purposes of labour market analysis, in particular when it comes to measuring earnings.

4. Gender-role socialization could explain why both young men and women tend to aspire to sex-typed occupations, even long before domestic/market work specialization takes place (Shu and Marini, 1998). It could also explain why women appear as disproportionately satisfied with their jobs, despite the fact that job segregation concentrates them in the least rewarded positions (Hakim, 1991). Sex-typed occupational choices at the time of labour-market entry have important implications for subsequent occupational careers (see e.g. Blossfeld, 1987).

5. The allocation process can also be affected by consumers' *gendered* preferences (Blau *et al.*, 2001[1986], 226; Reskin, 1993, 253); the segregation of recruitment channels (Braddock and McPartland, 1987; see, however: Petersen *et al.*, 2000); and the gender attitudes of co-workers (Jacobs, 1989, 181–182; Tomaskovic-Devey and Skaggs, 2002).

6. Note that the higher the firm-specific component of job specialization, the steeper the tenure-earnings profile will be.

7. Service-class occupations are also characterized by high monitoring costs, which are the second crucial dimension in contractual-hazard models of the employment contract (Goldthorpe, 2000, ch. X). Unfortunately, monitoring costs cannot be measured directly in most standard surveys, including the ESS.

8. In Polavieja (2005), job-specialization was imputed indirectly using a condensed version of Goldthorpe's class schema (Goldthorpe, 2000, ch. X).

9. Beckers' assumption that families have a single utility function has been challenged by bargaining models of the family, which see the unequal distribution of domestic work as the result of long-term coordination between spouses/partners (see e.g. Lundberg and Pollack, 1993; Ermisch, 2003).

10. Calculating the proportion of women in respondents' occupation using 4-digit information yields a 0 value for as much as 23 per cent of the cases. This is an important source of bias due to small sample size. In order to mitigate this distorting effect, calculations for these cases have been made using a 3-digit classification. 7 per cent of the 4-digit ISCO cells are occupied only by women. No changes have been made to these latter cells. All the models presented in this article have also been fitted using an alternative measure of occupational segregation calculated using 4-digit information drawn from the whole ESS pool. Results do not change (available on request).

11. Yet it must be noted that our sex-composition measure is based on 4-digit ISCO information for occupations, not for jobs, so there is a certain degree of discrepancy of measurement levels that Tam did not face. This discrepancy could reduce the explanatory potential of the JSK estimates.

12. The possibility that women underreported job-learning periods in female-dominated jobs (or over-reported them in male-dominated ones) seems unlikely as it is inconsistent with the idea that women underestimate their own capabilities. Upward bias in JSK estimation is thus considered improbable.

13. Subjectivity could also bias self-reported domestic supply if estimations were dependent on respondents' sex.

14. Results are available on request.

15. Let y_j be an observation for which wages are missing. A regressor list is formed from all the dependent variables $(x_1, x_2 \ldots, x_k)$ containing all xs for which x_{ij} is not missing. If this regressor list is not empty, a regression of y on the list is fitted. The imputed value $\hat{y}_j$ is defined as the predicted value of y_j, whereas the square of the standard error of the prediction yields an imputed variance value $(\hat{s}_j^2)$.

16. Note, however, that the possibility that sex-role attitudes are a *post hoc* rationalization of women's own occupational situation cannot be ruled-out, given the cross-sectional nature of the data.

17. Again it must be noted that the cross-sectional nature of this analysis precludes any interpretation as to the direction of causality.

18. An interaction effect between relative housework supply and respondents' sex has been tested and rejected.

19. Non-traditional women are defined as those scoring below 7 in the sex-role attitudinal scale.

20. This latter possibility has been defended by Hakim (2003b), who argues that survey attitudes do not measure individuals' own sex-role preferences but capture instead general societal norms. Applying this line of reasoning to our subject matter, one could argue that, behind the pro-egalitarian discourse reflected in their answers to the ESS, many Spanish women are actually hiding their 'true' core preferences for a more traditional sexual division of labour. Note that, however provoking, this argument is impossible to falsify using survey research, as it is precisely based on the assumption that surveys cannot capture core preferences (Hakim, 2003b, 340).

21. Neither f*ID nor hh*DS can by themselves absorb all the statistical effect of S on y, although both reduce it (results available on request).

22. See also: Tåhlin (2007, *forthcoming*).

Acknowledgements

The author wishes to thank the Norwegian Social Science Data Services (NSD) as the data archive and distributor of the European Social Survey Data (ESS). The ESS Central Co-ordinating Team (CCT) and the producers bear no responsibility for the uses of the data, or for interpretations or inferences based on these uses. Research for this paper was supported by a grant by the Fundación BBVA and has been developed within the Employment and the Labour Market Research Group of the EQUALSOC Network of Excellence. I am also grateful to Sunnee Billingsley, Duncan Gallie, Francesc Ortega, Luncinda Platt, Michael Tåhlin, and the two anonymous referees of the *European Sociological Review* for their helpful comments and criticisms. All errors are my own.

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Manuscript received: February 2007