Re-Thinking the Two-Body Problem:

The Segregation of Women into Geographically-Dispersed Occupations

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Abstract. Empirical research on the family cites the tendency for couples to relocate for husbands' careers as evidence against the gender-neutrality of household economic decisions. For these studies, occupational segregation is a concern because occupations are not random by sex and mobility is not random by occupation. I find the tendency for households to relocate for husbands' careers is better-explained by the segregation of women into geographically-dispersed occupations rather than by the direct prioritization of men's careers. Among never-married workers, women relocate for work less-often than men and the gender effect disappears after accounting for occupational segregation. While most two-earner families feature husbands in geographically-clustered jobs involving frequent relocation for work, families are no-less-likely

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to relocate for work when it belongs to the wife. I conclude future research in household mobility should treat occupational segregation occurring prior to marriage rather than gender bias within married couples as the primary explanation for the prioritization of husbands' careers in household mobility decisions.

Introduction

For young dual earner couples, work relocations often involve sacrificing one's career for the sake of the other. Sociological, economic, and demographic research consistently shows that work relocations strain marriages and impair careers of "trailing spouses" (see, for example, Cooke 2003; McKinnish 2008). Occupation-specific studies highlight the challenges of reconciling two careers among physical and life scientists (McNeil and Sher 1999), academics (Helppie and Close 2011), and military officers (Gill and Haurin 2002). Large employers often offer informal or formal spousal placement assistance to attract workers with highly-specialized skills, and rural universities typically refer to the issue of recruiting couples as the "two-body problem."

Just as dual-career couples have become increasingly-common, so too has research on the two-body problem. Early research offered gender-neutral hypotheses regarding the prioritization of one spouse's career in mobility decisions (Long 1974; Mincer 1978; Sandell 1977). Recent empirical work shows that relocation decisions tend to improve husbands' and impair wives' career development. Critics of neoclassical economic models of household mobility cite the prioritization of husband's careers in relocation decisions as evidence that mobility decisions are governed by normative gender roles (see, for example, Bielby and Bielby 1992; Cooke 2003; Shihadeh 1991; and Sorenson and Dahl 2011).

However, by on focusing on individual household mobility decisions, empirical work on household migration neglects the role of occupational segregation in the two-body problem. The standard regression models predicting who is a tied mover assume that the economic benefits of relocation are captured by earnings and educational controls, and that residual variation correlated with sex thereafter can be attributed to normative gender roles and power relations. However, occupations are not random by sex, and mobility is not random by occupation; this assumption is violated if men and women segregate into geographically constrained and flexible occupations, as would be the case if their career choices are made in anticipation of the two-body problem or if labor market processes otherwise channel women into jobs that can be performed anywhere. In other words, the tendency for families to relocate for men's careers may be a feature of men's and women's *ex ante* career choices rather than families' *ex post* prioritization of husband's careers, and therefore endogenous.

This study tests for the bias posed by occupational segregation in inducing families to prioritize husbands' careers in relocation decisions. I operationalize the constraint or flexibility of occupations by examining the observed geographic distribution of the occupation—particularly the degree to which an occupation is geographically "clustered" (like petroleum engineers) or "dispersed" (like elementary school teachers). I use a generalization of Duncan's dissimilarity index that indicates the share of workers in an occupation that would need to relocate for that occupation to employ the same number of workers per capita in every U.S. metropolitan area (Benson 2011). I calculate this index for each occupation in the Decennial Census 5% Public-Use Microdata Sample (which has many observations) and match it to household occupations in the 2003-2010 March CPS (which has work relocation data).

I use these data to produce precise estimates for the propensity to relocate for work by marital status and by sex, and before and after controlling for a measure of occupational clustering. Among never-married workers, I find that men are more likely than women to relocate for work. However, I find never-married men are more likely to have clustered jobs and clustered jobs are more likely to relocate for work, thereby signing direction of the omitted variable bias. After controlling for occupational clustering, the estimated coefficient of sex on relocation is precisely estimated and not significantly different than zero. Likewise, among married couples, I find that husbands tend to work in more-clustered occupations, that clustering is positively correlated with relocation, and that the coefficient of clustering on relocation is not statistically different between husbands and wives.

This study is agnostic to the particular reasons why this segregation occurs, and instead focuses on whether it confounds predictors of geographic mobility. One possibility is that nurturing and support occupations (such as nursing and secretarial work) attracts women, and in-person services tend to be ubiquitous. Alternatively, the tendency for relocation decisions to benefit men's careers may be endogenous if young men and women anticipate this when they choose their occupations.

These results contextualize prior empirical work on household mobility in a life-course setting. Specifically, the segregation of highly-educated women into geographically-dispersed skilled occupations (such as teachers, dentists, or general managers, and not nuclear engineers or naval architects) may explain why household relocations appear to be sensitive to the husband's college education but not the wife's (see, for example, Compton and Pollak 2007; McKinnish 2008). Results suggest that the mobility literature has mistakenly placed too much emphasis on direct intra-household discrimination as the explanation for why households tend to relocate for

husbands' jobs. Rather, the literature should emphasize on the segregation of women into geographically-flexible jobs, rather than direct intra-household discrimination, as a consequence of the two-body problem on inequality. This suggests that the two-body problem is not only an issue for household and mobility research, but also one for occupational segregation research. Lastly, results suggest that reducing the disproportionate impact of the two-body problem on women's careers may be more difficult than previously believed because it is embedded in occupational segregation occurring even among the never-married.

I. GENDER AND HOUSEHOLD MIGRATION

Beginning in the postwar period and civil rights movement, women have reversed the education gap and reduced the labor force participation and earnings gaps. Since 1982, women have constituted the majority of new bachelor's degree holders in the U.S. Women's rising educational attainment helped narrow the overall sex pay gap and increased the share of household income contributed by women (Brewster and Padavic 2000; Ciabattari 2001; Simon and Landis 1989).

Despite women's gains in educational attainment, occupational segregation and the specialization of women in household labor remains persistent. These patterns emerge early in the life cycle, potentially beginning with early childhood socialization (Friede Westring and Ryan 2011) or rational specialization (see, for example, Echevarria and Merlo 1999; Engineer and Welling 1999; Hadfield 1999).

Household research cites the tendency for women's earnings and employment to temporarily decline after household moves as evidence for the normative prioritization of men's participation in the external market. These analyses typically begin from neoclassical models of the household that hypothesize gendered domestic labor and household relocation decisions are attributable to

differences in men's and women's innate and accumulated human capital (such as physical strength and work skills), and that gender role ideology does not need to be invoked to explain the tendency for household economic decisions to advantage the husband's career. Mincer (1978) notes decisions to relocate for the economic well-being of the family may disadvantage the career of the "tied mover," while decisions not to relocate may disadvantage the career of the "tied stayer."

Like research on the allocation of household chores, research on household relocation decisions is made difficult by the challenge of fully-controlling for the economic attributes of husbands and wives. These controls are essential, since men and women differ in education, experience, personal career goals, and other characteristics likely to affect the economic consequences of relocation decisions. Because empirical work estimating the likelihood of relocation necessarily interprets significant effects for sex net of the controls suggested by neoclassical theory as evidence of *direct* effect of sex, controls must capture all characteristics that are both correlated with sex and predictive of the economic benefits of relocation.

Consistent with Mincer, empirical work consistently finds that household relocations typically benefit the career of one spouse to the detriment of the other (Bailey and Cooke 1998; Battu, Harminder, and Sloane 1998; Boyle et al. 2001; Clark and Huang 2006; Jacobsen and Levin 1997; McKinnish 2008; Nivalainen 2005; Pixley and Moen 2003; Shauman and Noonan 2007). Critics of the neoclassical view interpret the prioritization of the husband's career in mobility decisions as an enactment of traditional gender roles (see, for example, Bielby and Bielby 1992; Nivalainen 2004; Cooke 2003; Swain and Garasky 2007). On the whole, studies find evidence that household relocation decisions are made with economic motives, but that sex effects after controlling for husbands' and wives' education and earnings imply non-economic

causes as well (Markham and Pleck 1986; McKinnish 2008; Noe et al. 1988; Ostroff and Clark 2001; Pixley and Moen 2003; Turban, Campion, and Eyring 1992; Zvonkovic et al. 1996).

While the immediate consequences of the two-body problem on the trailing-spouse's labor market outcomes are now well-evidenced, research has rarely and only recently shifted focus to the life-course consequences of the two-body problem. As Pixley (2008) notes, the emphasis on individual mobility decisions, rather than broader life-course implications, may be a result of a methodological convenience rather than analytical importance.

Since Pixley's critique, a few studies have explored how workers' occupations predict relocation. Using the Panel Survey of Income Dynamics and the Census, Shauman (2010) examines both individual and occupational predictors of relocation (for all reported reasons for relocating, unlike this study), and finds that controlling for occupational characteristics enhances individual-level predictors of mobility. Brandén and Ström (2011) find sex segregation around wage profile characteristics among Swedish couples and find evidence of both location coordination among working spouses and distinct patterns among men and women. While fully-controlling for the set of potential differences between occupations remains an ambition for this literature, on the whole, there appears to be evidence for both systematic sex segregation around occupational characteristics and significant residual variation in mobility explained by sex.

II. THEORY

The standard regressions estimating men's and women's economic outcomes upon household relocation assume that future economic rewards for relocation may be captured by the usual available controls, such as those for earnings and education. When predicting relocation for work by sex, occupational segregation is not a concern if this segregation is otherwise-random with respect to which men and women require calculated work relocations for career advancement.

However, if men segregate into geographically-clustered occupations involving frequent work relocation and women segregate into geographically-dispersed occupations that are relatively adaptable to spousal relocation, then occupational characteristics correlated with both sex and propensity to relocate remain in the residual term. As such, occupational segregation is a potential source of omitted variable bias.

There are several reasons to expect that occupational segregation would be non-random with respect to mobility. First, if men and women expect families to prioritize husbands' careers in relocation decisions, this may in turn compel women to sort into flexible occupations, thereby reproducing occupational segregation. Benson (2011) finds that women who enter constrained occupations disproportionately suffer lower earnings, later marriage, and higher divorce rates than men in those occupations or women in flexible occupations. In this case, the segregation of women into flexible jobs persists due to men's and women's aggregately self-fulfilling expectations and the tendency for families to relocate for husbands' careers may be endogenous to career choices that consider for their future expected mobility.

Second, the segregation of women into flexible occupations may be a byproduct of other occupational characteristics. For normative or preferential reasons, women may segregate into human service jobs (such as health, education, or administrative support) that can be performed anywhere (see Anker 1997 or Cohen and Huffman 2003, for reviews).

Third, organizational and labor market processes may also disadvantage women and segregate them into support occupations offering few opportunities for career advancement and few rewards for calculated relocations (see, for example, Acker 1990, Epstein 1990, Kanter 1977, Ressner 1987).

Although it is notoriously difficult to test the mechanisms behind theories of occupational segregation, empirical work has provided clear evidence that differences in men's and women's career and family expectations begin long before marriage, and that patterns in men's and women's job searches differ. Blau and Ferber (1991) find that men and women in college generally have different career and family ambitions. Daymont and Andrisani (1984) find that men are more likely to major in science, technology, engineering, and math, while women are more likely to major in education, the humanities, health, and biology; this is notable because the former typically lead to jobs in knowledge-work that are geographically-clustered, while the latter lead to jobs in human services that are generally flexible. Interviews by Hanson and Pratt (2005) find that women tend to search for jobs more-locally than do men, and that the geographic scope of women's job searches is less-likely to require household relocation. Becker and Moen (1999) find that dual-earner couples' relocation decisions are often governed by the explicit prioritization of one spouse's "career" over the other's "job," with most "careers" belonging to the husband, but not all.

Regardless of the cause, occupational segregation is a potential concern for empirical research on household mobility because occupations are not random by sex and geographic mobility is not random by occupation. These two conditions yield classic omitted variable bias (OVB), with the possibility that segregation anticipates the two-body problem further implying endogeneity. The results section replicates earlier empirical research on household mobility, signs the OVB posed by segregation (*i.e.*, mechanically shows whether it will be positive or negative), and then presents results after correcting for the bias.

Hypothesis 1 is that the tendency for never-married men to relocate for work more often than women is an artifact of the OVB posed by occupational segregation; regressions that do not

address this OVB will overestimate the independent effect of sex on the likelihood of relocating for work. Showing biases for one sex in relocation for work among never-married workers is striking since these men and women are, by definition, not tied to a spouse. This hypothesis is also important because it tests whether occupational sorting occurs in advance of marriage.

I test Hypothesis 1 using a sequence of four empirical tests, labeled Hypothesis 1A – 1D. Hypothesis 1A is that never-married men (women) sort into (away from) geographically-clustered occupations. This signs the direction of the correlation between the independent variable of interest with the omitted variable of concern (sex and clustering). Hypothesis 1B is that clustered occupations relocate for work more often than occupations that are dispersed. This signs the correlation between the omitted variable of concern with the dependent variable (clustering and relocation for work). Together, Hypothesis 1A and 1B mechanically sign the direction of OVB (men are more-likely to work in clustered occupations, and clustered occupations are more-likely to relocate for work among both sexes). Hypothesis 1C is that never-married men relocate more often than never-married women. This replicates earlier results by estimating the unconditional effect of sex on the propensity to relocate for work (sex and mobility). Lastly, Hypothesis 1D is that sex is not a significant predictor of mobility after controlling for the OVB posed by occupational clustering.[2] Note that Hypothesis 1D is a null prediction and should be interpreted to signify that we can achieve a reasonably-precise estimate of a small effect of sex after introducing the control.

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[2] The study of the segregation of women into low-paying jobs within firms is a familiar application of this approach; fine-grained controls reduce and virtually eliminate the estimated effect of sex (see, for example, Kilbourne et. al. 1994, Petersen and Morgan 1995, or Reskin and Bielby 2005). However, because relocations-for-work are rare, introducing occupational fixed effects makes tests very weak. Including a linear term for clustering instead maintains the statistical power of the test while tying results to the specific theoretical construct of provided by clustering.

Hypothesis 2 is that the tendency for households to relocate for the husband's job is explained by the segregation of women into dispersed occupations, and so is also thereby an artifact of the OVB posed by occupational segregation. I test Hypothesis 2 by executing the above empirical strategy for married workers.

Hypothesis 2A examines the correlation between occupational clustering and the propensity to relocate for work. Unfortunately, the CPS does not ask for whose occupation the household relocates for work, precluding straightforward estimation when both heads are in the labor force. To address this, Hypothesis 2A tests whether the effect of occupational clustering on job mobility among single-earner men and women is positive and whether the marginal effect of the maximal clustering score on mobility for dual-earner couples is positive. Hypothesis 2B examines whether dual-earner couples are more likely to feature a husband with a more-clustered occupation than the wife. As above, Hypotheses 2A and 2B sign the direction of the omitted variable bias. Finally, Hypothesis 2C estimates the independent effect of sex net of a control for the maximum degree of occupational clustering within the family. Once again, this is a null prediction and should be interpreted as a test that the predicted effect of sex is small and estimated with reasonably-high precision. This offers an alternative explanation for existing studies that use families' propensity to relocate for husbands' careers to infer a normative prioritization of husbands' careers in mobility decisions, suggesting they are endogenous to men's and women's occupations.

For each hypothesis, I analyze highly-educated workers separately. Primary and secondary education in the U.S. is highly general with specialization occurring largely through college and internships. For highly-skilled occupations, the penalties for being a "tied mover" may be higher, and the omitted variable bias is expected to be more-pronounced. For example, gaming cage

workers, who are generally employed by casinos to exchange money for chips, require little formal or informal training. Because they are concentrated in few cities (such as Las Vegas, Atlantic City, Reno, and Biloxi), gaming cage workers are a very geographically-clustered female-dominated occupation. While the cost of being a trailing spouse and switching occupations from a gaming cage worker to a cashier is likely to involve little loss in career prospects or professional identity, it is reasonable to expect greater costs of being a trailing spouse in clustered occupations that are more-highly educated (for example, a geographically-displaced nuclear engineer may become a math teacher).

III. DATA

I calculate an occupational clustering index using the 5% PUMS of the 1980, 1990, and 2000 Decennial Census.[3] The Decennial Census features the large number of observations (about eight million workers per Census year) needed to estimate the employment share of each occupation in each metropolitan statistical area. Occupations are standardized to the 1980 SOC codes using the Census Bureau's occupational crosswalk. The pre-normalized clustering index for a given occupation is taken from Benson's (2011) generalization of Duncan's dissimilarity index:

$$C_j^* = \frac{1}{2}\sum_{i=1}^{I}\left|\frac{n_{ij}}{n_j} - \frac{n_i - n_{ij}}{n - n_j}\right|$$

where *I* represents the set of metropolitan areas, *n* represents the counts of workers aged 18-65 in the labor force, and subscripts denote counts within metropolitan areas *i* and occupations *j*. I calculate this index for each occupation and Census year. The pre-normalized clustering index *C** is then normalized as follows. First, I take the log-transformation of *C**, yielding a roughly-normal distribution with values ranging from -3.22 for the most-dispersed occupation to -0.36 for the most-clustered. For simplicity, the index is normalized by addition so that the most-dispersed

[3] Available online. See Ruggles et al. (2010).

occupation receives a *C*-score of 0, with higher values corresponding to more-clustered occupations. The clustering indices are then merged to occupations in the CPS March Supplement, with the 2003-2010 March Supplements representing those with the six-digit 2000 Standard Occupational Classification (SOC) codes used in 2000 Decennial Census.[4]

While Duncan's D is conventionally used for dichotomous dissimilarity (e.g. sex), the generalized form is also appropriate for polytomous dissimilarity (e.g. U.S. metropolitan areas). The clustering index also retains the intuition interpretation that makes Duncan's D a desirable measure: it is the share of workers within an occupation that must relocate for the share of workers to be balanced in every metropolitan area. For example, in 1980, elementary school teachers constituted about 1% of the national labor force. The pre-normalized clustering index C^* for elementary school teachers is 0.053, signifying that in 1980, 5.3% of elementary school teachers would need to relocate for them to constitute 1% of the labor force in every U.S. metropolitan area. In contrast, the pre-normalized clustering index C^*_{1980} for mining and petroleum engineers is 0.731, signifying that 73.1% would need to relocate to equalize the labor force share of petroleum engineers. Within the 1980 sample, these occupations represent the most-dispersed and most-clustered non-military occupations, respectively, where the majority of workers have bachelor's degrees. In the full sample, textile winding setters are the most-clustered occupation and retail salespeople are the most-dispersed.

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[4] The 2000 PUMS distinguishes 337 occupations by (up to) a six-digit SOC code. By exploiting the hierarchical nature of SOC Codes, it is easy to show that aggregating occupations (for example, to three digit categories) reduces the magnitude of the effects, suggesting potential aggregation bias. This bias is also noted by Reskin (1993: 243): "`college and university teacher' includes someone teaching night classes on repairing office machines at Parkland Community College as well as a distinguished professor of mathematics at Harvard." Likewise, segregation into highly-mobile sub-specialties within occupations would not be captured by aggregate measures like SOC codes. The resulting measurement error (in the independent variable) is expected to cause attenuation bias, reducing the magnitude of the estimated coefficients and increasing the standard errors.

I use the Decennial Census to calculate the clustering scores and merge clustering scores to the 2000 SOC occupations in the 2003-2010 March CPS.[5] Using the distribution of occupations as the independent variable of interest, rather than observed differences in mobility rates by occupation, has several distinct advantages. First, the geographic distribution of an occupation is plausibly exogenous. For example, petroleum engineers are concentrated around oil fields, and primary school teachers are dispersed around children (see Ellison and Glaeser 1997 for a discussion of the geographic agglomeration of occupations). Realized relocation decisions are endogenous. Indeed, because relocation is the outcome variable, regressing relocation on average relocation by job threatens a mechanical relationship.

Second, the geographic distribution of an occupation may be a better measure of the geographic-flexibility of a trailing spouse than the incidence of observed relocations. It is not clear from the incidence of work relocations alone whether the geography of the relocating worker was constrained. For example, using mobility rates, McKinnish (2008) finds veterinarians, head cooks, and cashiers to relocate often, but intuitively these jobs would be relatively-robust to spousal relocations. However, some occupations appear exceptional; for example, clergy are dispersed, but any individual may be geographically constrained.

Third, an advantage of using the clustering index is that the Census does not report why individuals relocate for work. This is problematic because only 10% of recently-relocated households surveyed in the CPS report that the primary reason for doing so was for work or a job transfer, and workers may relocate precisely because their job is highly flexible (enabling them

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[5] The March CPS features 6-digit SOC codes beginning in 2003. Unlike the Census or PSID, it asks relocating households to report the primary reason for relocating in the prior year, including "for work or job transfer." Only 10% of relocations are for work, and geographic flexibility may lead workers to relocate for reasons other than work, confounding estimates that use observed relocations for all reasons. The March CPS also has much greater statistical power than the PSID. Data are accessed by the ipums.org (see Ruggles et al. 2010).

to relocate for family reasons, for personal preferences, or indeed for a spouse).[6] Taken together, it seems unlikely that the high mobility in certain occupations (such as cooks) necessarily implies that they are geographically constrained.

[FIGURE 1]

To provide an illustration of occupational clustering, Figure 1 presents the geographic distributions of the three SOC occupations: physicians and surgeons, medical scientists, and physicists and astronomers.[7] Physicians have a low clustering score, reflecting that they are spread relatively-evenly among US metropolitan areas. In contrast, physicists and astronomers have a high clustering score, with employment concentrated around the U.S. Department of Energy National Laboratories. A list of the fifteen most-clustered and most-dispersed occupations for which the majority of workers have graduate degrees is reported in Table 1.

[TABLE 1]

Table 1 shows that none of the twenty most-clustered occupations were majority-female in 1980, and only one became majority-female by 2000 (archivists and curators). In contrast, fifteen of the twenty most-dispersed occupations were majority female in 1980 or became majority female by 2000. The exceptions are "other financial specialists," "physicians," "clergy and other religious workers," and "dentists," which respectively rose in female share to 47%, 27%, 19%, and 30% in that period. "CEOs and public administrators" declined in female share in that period (from 25% to 19%); although categorized as a "dispersed occupation" in the Census, this occupation is conceptually problematic given that it includes everyone from owners of small restaurants to Fortune 500 CEOs. Table 1 also shows that many of the most highly-dispersed

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6 Unfortunately, while the CPS reports the principal reason for household relocations, it is far too small to estimate relocation likelihoods by occupation.

7 "Physicists and astronomers" are treated as distinct from "post-secondary teachers" by the Bureau of Labor Statistics.

high-skill occupations absorbed a larger share of highly-educated women than might otherwise be predicted by sex-typing or the segregation of women into "caring" occupations. For example, financial managers, accountants, auditors, advertisers, and clinical lab technicians are highly-dispersed occupations featuring few female-type work attributes, but from 1980 to 2000, each doubled in female-share to become majority-female.

While some highly-skilled occupations became more clustered or more dispersed, the mean extent of occupational clustering stayed roughly constant from 1980 to 2000. In particular, geologists, writers, therapists, and special education teachers became considerably more dispersed (the measure dropped by more than 20%) while only medical scientists became considerably more-clustered (the measure increased by more than 20%).

I define married couples as couples where both heads are married with a present spouse. I examine relocation incidence for workers younger than 35 to focus on those most at-risk of marriage and work relocations. Occupations are matched on the occupation in the previous year (i.e. for those relocating, I use occupations prior to relocation). The CPS asks households whether they relocated in the prior twelve months. Those that report doing so are asked for the primary reason for relocation. I define a "work relocation" as a change in residence across counties in which the respondent cites work or job transfer as the primary reason for relocation. Common reasons include relocation for family, upgrading housing, or change in marital status. Relocations "for work or job transfer" include about 10% of relocations.

IV. RESULTS

To test the hypothesis that women sort into geographically-dispersed occupations among the never-married (Hypothesis 1A), I perform a two-sample t-test for occupational clustering by sex among never-married men and women in the 2000 Decennial Census. Never-married men have

an estimated mean clustering score of 1.108, and never-married women have a mean clustering score of 0.891. Among bachelor's degree holders, these estimated means are 1.153 for men and 1.073 for women. All standard errors are less than 0.001, and two-sample t-tests reject that the mean clustering scores are equal among the population of men and women with $p<0.001$. I conclude never-married men have higher mean clustering scores than women. Figure 2 illustrates the relationship between occupational clustering and female share.

[FIGURE 2]

Next, I test the relationship between the occupational clustering index and the propensity to relocate for work among never-married men and women using the 2003-2010 CPS March Supplements (Hypotheses 1B-1D). Table 2 presents results.

[TABLE 2]

Table 2, Columns 1 and 2 show that both young never-married men and women in geographically-clustered occupations are more likely to relocate than those in dispersed occupations, lending support for Hypothesis 1B ($p<0.001$; two-tailed test). The estimated magnitude of the effect is quite stark; the range of the index is 2.45, and a one point increase is associated with a 38% increase in likelihood of relocating for work among all never-married men and a 57% increase among never-married women. Tests for Hypothesis 1A and 1B find men segregate into clustered occupations, and people in clustered occupations are more likely to relocate for work (for both sexes), thereby signing the OVB. [8] Column 3 reintroduces the OVB

[8] To sign the omitted variable bias posed by the correlation between sex and occupational clustering, I examine pairwise correlations between a female dummy, a worker's occupational clustering score, and the probability a worker relocates for work he correlation between female and the clustering score is negative (Hypothesis 1A) and the correlation between correlation between the clustering score and relocation for work is positive (Hypothesis 1B), both with $p < 0.01$. Therefore, without controlling for clustering, the raw correlation between female and relocation will be biased downward. These correlations are

by omitting occupational clustering. Consistent with prior research, I find women relocate for work less than men (even among never-married workers), lending support for Hypothesis 1C ($p<0.001$). Column 4 shows that, after including occupational clustering, the estimated correlation between the female indicator and relocation for work drops to 4%, which is no longer statistically significant ($p=0.30$). This precisely-estimated small effect is consistent with Hypothesis 1D.

Results are similar among bachelor's degree holders. Never-married bachelor's degree holders are more likely to relocate when they work in clustered occupations and women are less-likely to relocate than men, consistent with Hypotheses 1B and 1C. Controlling for occupational clustering reduces the effect of being female among bachelor's degree holders, yielding small but significant estimates for effect of being female on mobility. Model 8 estimates that a one log point increase in clustering is correlated with an 52% increase in the likelihood of relocating for work among young, bachelor's degree-holding workers, and females are 15% less likely to relocate for work than men.

Results broadly corroborate Hypothesis 1. Together, results show that much of the statistical variation in relocation for work explained by sex in the regression in Column 3 (Hypothesis 1C) is captured by the OVB caused by occupational segregation (Hypothesis 1A and 1B). Omitting

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each robust and significant in all nine explored specifications: the full sample, by age less than 35, for never-married and married workers, and by college education (ie. $1 + 2^3$). While correlations were strongest among the young and college-educated, this check suggests the OVB very-broadly leads to downwardly-biased estimates of the independent effect of sex on relocation for work.

the geographic dispersion of an occupation from mobility regressions overestimates the independent effect of sex on likelihood of relocating for work.[9]

Next, I examine occupational clustering within families and its effect on relocation (Hypothesis 2). I use the CPS March Supplement 2003-2010 to test for the existence of clustering within the family and the marginal effect of clustering for both men and women for a family's mobility. Table 3 presents results.

[TABLE 3]

Consistent with Hypothesis 2A, men tend to work in occupations that are more-geographically clustered than women. For both dual-earner couples and the subset of "power couples" in which both heads have a bachelor's degree (or greater), there are about two times as many couples featuring a husband with the more-clustered occupation than the reverse.

Hypotheses 2B and 2C concern the correlation between occupational clustering and household mobility. I estimate the correlation between occupational clustering on mobility when the higher score belongs to the husband and also when it belongs to the wife. The full regressions take the form:

$$\text{logit}(p_i) = \hat{\beta}_0 + \hat{\beta}_1 C_{im} + \hat{\beta}_2 C_{if} + \mathbf{XB}$$

where p_i is the logit-estimated probability that a household makes an intercounty work relocation, C_{im} and C_{if} are the clustering indices of the male's/female's occupation if it is strictly

[9] Primary reason for relocation is a self-reported variable, and I also examine relocation reported for reasons other than work (results are available upon request). Treating all relocations as the dependent variable reduce the magnitude of coefficients, the effect of clustering remains significant in all regressions, female loses significance in columns (3) and (7). Treating relocations primarily "for family" as the dependent variable makes clustering lose significance, and females are more likely to relocate. I interpret results to suggest never-married women are more likely than men to cite family as the primary reason they relocate, and the geographic clustering of a job is a better predictor of relocations that households report are primarily "for work."

the more-clustered of the two occupations (and is otherwise zero), and XB includes controls.[10] The coefficients $\hat{\beta}_1$ and $\hat{\beta}_2$ can be interpreted as the correlation between maximal clustering and work relocation, and the regression allows this effect to differ by sex. For robustness, controls include two methods for accounting for the spouse's occupation. Table 4, Columns (1) – (3) includes a relative measure: the difference in clustering between the two heads' occupations. Column (4) reports an absolute measure: the clustering score of the less-clustered occupation. The direction of these coefficients are theoretically ambiguous; the spouse may benefit from relocation for their career, but may be less amenable to relocating for the other spouse.

[TABLE 4]

Table 4 shows that occupational clustering is highly correlated with work relocation, consistent with Hypothesis 2B. Across the four specifications and eight point estimates, the estimated rise in the likelihood a family relocates for work given a one log-point increase in the clustering index of the head's occupation ranges between 21% and 48%.[11] Seven estimates are statistically significant with $p<0.05$, and five are with $p<0.01$. The coefficient in Column (4), on the occupation of the wife when the wife has the more-clustered occupation is positive and not statistically significant. The standard error on this estimate is also particularly high, reflecting in part that households in which the wife has the more clustered occupation are a minority. Notwithstanding the final significance test, I conclude that households with more-clustered occupations are more likely to relocate for work. Together, Hypotheses 2A and 2B sign the

[10] Among dual earner couples, 5.09% of men and women have the same occupation. To avoid bias, the terms for maximal clustering is zero for both spouses when occupations match. These couples are significantly more likely to relocate for work than spouses where the occupations do not (23.9%, with a standard error of 9.3%). One possibility is that work relocations are easier when the husband and wife have the same occupation.

[11] That is, exp(0. 192) = 1.212, and exp(0.389) = 1.475.

omitted variable bias: households tend to exhibit men in more-clustered occupations, and households with more-clustered occupations tend to relocate for work.

Lastly, Hypothesis 2C examines whether the marginal effects are similar for men and women. Point estimates for a family's likelihood of relocating for work are slightly greater when the husband has the more-clustered occupation. However, the estimated difference when the wife has the more-clustered occupation is very small (between 1% and 6%) and not statistically significantly different.

Results are robust to the typical controls used in the family mobility literature, including children, education, and earnings. While children clearly reduce a families' mobility, the only other statistically-significant predictor of mobility (net of clustering and other controls) is the husband's education. This result is consistent with McKinnish's (2008) finding that family mobility is predicted by the husband's, but not the wife's, attainment of a bachelor's degree. It also offers a potential explanation: men's college education is more likely to lead to technical jobs that tend to be more geographically constrained and tend to involve relocation for work (even among the women in those occupations), whereas women's college education is more likely to lead to jobs in human services that tend to be ubiquitous and tend not to involve relocation for work (even among the men in those occupations).

[TABLE 5]

Table 5 presents results for "power couples," or couples in which both the husband and the wife have bachelor's degrees. Because power couples tend to have occupationally-specific skills, high earnings potential, and high labor force attachment, they are often the subject of research on couples' colocation problem. Results for power couples are substantively the same, except all eight tests corresponding to Hypothesis 2B are now positive and statistically significant. Point

estimates for the effect of marginal occupational clustering is greater for women in all four regressions, although the difference is not statistically significant.

V. CONCLUSIONS

Prior research has used the tendency for families to prioritize husbands' careers in relocation decisions as evidence for normative family relations and as a mechanism for continued labor market inequality. As evidence, this research has cited how sex remains a predictor for why families relocate, even after including several controls. This study departs from prior research by distinguishing a variable—the geographic constraint of an occupation—that is not random by sex, not adequately controlled for by education and earning covariates, and that predicts relocation. Using the Census to measure the dispersion of occupations, and data on households and work relocations from the March CPS, I replicate prior research finding men tend to relocate for work more than women, both among the never-married and married. I sign the direction of the omitted variable bias posed by the segregation of women into geographically flexible jobs, and provide evidence that this segregation is a principal driver of the finding that men tend to relocate for work more than women. I also find that results are particularly pronounced among bachelor-degree holders.

This study emphasizes the problem posed by omitted variable bias posed by occupational segregation, and is does not make claims as to why this segregation occurs. This segregation may occur specifically in anticipation of the prioritization of husband's careers in mobility decisions, or it may be due to other features of geographically flexible occupations.

This form of occupational segregation may help explain several outstanding empirical puzzles in the study of household mobility and in inequality. One of the most robust findings in the household mobility literature is that family mobility is sensitive to whether the husband is

college-educated, but not the wife (Compton and Pollak 2007; McKinnish 2008). The tendency for men to pursue college majors leading to technical careers and for women to pursue majors leading to jobs in human services may explain why families' mobility tends to be more sensitive to the husband's higher education. The segregation of women into geographically-dispersed occupations may also explain why women tend to work closer to home than men, particularly when women work in female-segregated jobs (see Hanson and Pratt 1995 or Fernandez and Su 2004 for reviews).

For the occupational segregation literature, the two-body problem may offer an explanation for why the highly-educated, and traditionally male-dominated occupations that were geographically-dispersed (such as physicians, dentists, general managers) absorbed so much of the growth of highly-educated women in recent decades, while more–clustered occupations (typically more-technical occupations, such as specialized engineers) have not.[12]

An advantage of using the geographic distribution of the occupation is that it is plausibly-exogenous, while using actual relocation propensities may be endogenous or mechanical. However, occupational clustering remains an imperfect proxy of the degree to which an occupation truly enjoys geographic flexibility. For example, post-secondary teachers are relatively-dispersed, but the careers of young academics are likely to benefit from the ability to make calculated relocation decisions. Future research may operationalize other measures of how robust occupations are to exogenous relocations. Future research may also examine whether skills can be transferred across occupations or correlation in the geographic overlap between

[12] There are also exceptions to the generalization that engineering and technical occupations are geographically-clustered; for example, civil engineers, accountants, and auditors are technical occupations and they are dispersed. They also have rapidly absorbed highly-educated women.

different geographically-constrained occupations. Indeed, the relatively high mobility among couples with matching occupations suggests this mitigates the two-body problem.[13]

For the literature on gender and the family, future research may consider *ex ante* segregation of women into ubiquitous jobs rather than the *ex post* normative prioritization of men's careers in mobility decisions as the primary explanation for why families tend to relocate for husbands. Future work may also consider the two-body problem's role in perpetuating occupational segregation. More broadly, research should consider how the prioritization of husbands' careers in mobility decisions affects young men and women's career paths and how the two-body problem promotes the intergenerational reproduction of occupational segregation.

[13] This approach would also build upon research on the migration of college-educated couples into large metropolitan areas that may be amenable to supporting two careers (see, for example, Costa and Kahn 2000).

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TABLES AND FIGURES

Table 1. Most-Clustered and Most-Dispersed High-Skill Occupations in 2000

Fifteen Most-Clustered	C	Fifteen Most-Dispersed	C
Petrol & Mining Engineers	2.45	Primary School Teachers*	0.00
Aerospace Engineers	2.09	Secondary School Teachers*	0.29
Atmosphere & Space Scientists	2.05	Managers (Education)**	0.33
Physicists & Astronomers	2.02	Managers (Health Occupations)**	0.33
Actuaries	1.93	Registered Nurses*	0.36
Legislators	1.86	Other Financial Specialists	0.43
Chemical engineers	1.83	Financial Managers**	0.56
Medical scientists	1.81	Accountants & Auditors**	0.58
Metallurgical & Materials Eng.	1.75	Advertising & Related Sales**	0.59
Podiatrists	1.74	CEOs & Public Administrators	0.59
Mathematicians & Math Sci.	1.73	Vocational & Educ. Counselors*	0.60
Other Physical Scientists	1.73	Welfare Service Aides*	0.61
Actors, Directors, & Producers	1.71	Clinical Lab Technicians*	0.65
Geologists	1.66	Physicians	0.69
Business & Promotion Agents	1.63	Physical therapists*	0.69
Agricultural & Food Scientists	1.62	Clergy & Religious Workers	0.7
Biological Scientists	1.6	Pharmacists**	0.71
Airplane Pilots	1.55	Social Workers*	0.74
Mechanical Engineers	1.55	Marketing Managers**	0.78
Archivists & Curators**	1.51	Dentists	0.81

Note: *: Majority female in 1980 and 2000. **: Became majority female within 1980-2000 "C" is the normalized clustering index (otherwise 0.04 to 0.73, un-normalized), calculated in the 2000 Census. Occupations are taken from 1980 SOC Codes, excluding agriculture and military occupations. "High Skill" occupations are the seventy-seven in which where the majority of workers have bachelor's degrees.

Table 2. Logistic Regression for Probability Young Never-Married Men and Women Relocate for Work, 2003-2010

	All Workers				Bachelor's Degree Holders			
Sex:	Men	Women	Both	Both	Men	Women	Both	Both
Hypothesis:	H1B	H1B	H1C	H1D	H1B	H1B	H1C	H1D
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Clustering "C"	0.361***	0.385***		0.370***	0.411***	0.309***		0.364***
	(0.044)	(0.053)		(0.034)	(0.0730)	(0.080)		(0.054)
Female			-0.118**	-0.0407			-0.22***	-0.156**
			(0.037)	(0.038)			(0.055)	(0.056)
Constant	-4.38***	-4.45***	-3.95***	-4.39***	-3.34***	-3.37***	-2.82***	-3.27***
	(0.060)	(0.062)	(0.025)	(0.049)	(0.103)	(0.097)	(0.039)	(0.080)
LR Chi-Squared	65.4***	49.9***	10.1***	125.3***	31.3***	14.8***	15.3***	60.5***
Observations	87,943	79,736	167,679	167,679	12,501	14,737	27,238	27,238

Note: Standard errors in parentheses. Data from CPS March Supplement 2003-2010, and includes labor force participants under age of 35.

* $p < 0.05$ (two-tailed test); ** $p < 0.01$; *** $p < 0.001$

Table 3. Relative Geographic Constraint of Occupations within the Family

	All Workers		"Power Couples" Two Bachelor's Degrees	
	Share	SE	Share	SE
Two-Earner Couples				
Husband's Occupation is More-Clustered	42.7%	(0.164%)	42.9%	(0.164%)
Wife's Occupation is More-Clustered	20.9%	(0.097%)	22.9%	(0.103%)
Same Occupation	3.4%	(0.035%)	6.3%	(0.049%)
One Earner Couples				
Husband Works, Wife Does Not	23.9%	(0.106%)	23.2%	(0.104%)
Wife Works, Husband Does Not	5.2%	(0.044%)	2.9%	(0.033%)
Zero-Earner Couples	3.8%	(0.037%)	1.9%	(0.026%)

Note: Percents denote shares of households featuring the corresponding labor force circumstances. For example, 42.7% of households exhibited two earners, of which the husband worked in the more-clustered of the two occupations. Data from CPS March Supplements 2003-2010.

Table 4. Probability of Relocating for Work among Young Dual-Earner Couples with Both Spouses Present

	(1)	(2)	(3)	(4)
Higher C Occupation				
C of Husband's Occ	0.389***	0.343**	0.265*	0.251**
x (Husb Higher C)	(0.114)	(0.114)	(0.117)	(0.0927)
C of Wife's Occ	0.360**	0.317**	0.242*	0.192
x (Wife Higher C)	(0.117)	(0.117)	(0.121)	(0.177)
Spouse's Occupation				
Difference in C	-0.228	-0.179	-0.0826	
x (Wife Higher C)	(0.128)	(0.127)	(0.131)	
Difference in C	-0.146	-0.111	-0.0520	
x (Husb Higher C)	(0.164)	(0.164)	(0.168)	
C of Husband's Occ				0.0607
x (Wife Higher C)				(0.157)
C of Wife's Occ				0.0430
x (Husb Higher C)				(0.0578)
Controls				
Family Has Child		-0.921***	-0.730***	-0.736***
		(0.0835)	(0.0869)	(0.0837)
Husband has Bachelor's			0.849***	0.846***
			(0.103)	(0.101)
Wife has Bachelor's			0.109	0.122
			(0.103)	(0.101)
Husband's Log-Wage			0.0782	0.0839
			(0.107)	(0.105)
Wife's Log-Wage			-0.0565	-0.0573
			(0.0823)	(0.0813)
Constant	-5.499***	-4.714***	-5.079***	-4.880***
	(0.438)	(0.444)	(0.625)	(0.712)
LR Chi-Squared	13.02***	235.73***	261.82***	262.13***
Observations	34392	34392	32614	34570
$\beta_1 - \beta_2$	0.030	0.026	0.023	0.059
	(0.059)	(0.059)	(0.061)	(0.194)

Note: Standard errors in parentheses. From March CPS 2003-2010. $\beta_1 - \beta_2$ is the estimated difference in the first two coefficients. Includes couples where the mean age is at most thirty-five, with present spouses.

* $p < 0.05$ (two-tailed test); ** $p < 0.01$; *** $p < 0.001$

Table 5. Logistic Regression Estimating the Probability of Relocating for Work among Young Dual-Earner Couples, among "Power" Couples

	(1)	(2)	(3)	(4)
Relative Geo. Constraint				
C of Husband's Occ	0.515**	0.446**	0.455**	0.341*
x (Husb Higher C)	(0.172)	(0.172)	(0.173)	(0.138)
C of Wife's Occ	0.616***	0.552**	0.569**	0.612*
x (Wife Higher C)	(0.177)	(0.178)	(0.179)	(0.260)
Spouse's Occupation				
Difference in C	0.0167	0.0682	0.0599	
x (Wife Higher C)	(0.237)	(0.237)	(0.238)	
Difference in C	-0.310	-0.270	-0.317	
x (Husb Higher C)	(0.197)	(0.197)	(0.201)	
C of Husband's Occ				-0.120
x (Wife Higher C)				(0.223)
C of Wife's Occ				0.0956
x (Husb Higher C)				(0.0728)
Controls				
Family Has Child		-0.740***	-0.736***	-0.742***
		(0.123)	(0.124)	(0.118)
Husband's Log-Wage			0.195	0.188
			(0.244)	(0.234)
Wife's Log-Wage			0.0678	0.0591
			(0.174)	(0.168)
Constant	-5.726***	-5.055***	-6.285***	-6.169***
	(0.681)	(0.689)	(1.373)	(1.411)
LR Chi-Squared		49.5***	49.7***	56.4***
Observations	8438	8438	8215	9107
$\beta_1 - \beta_2$	-0.100	-0.107	-0.113	-0.271
	(0.094)	(0.095)	(0.096)	(0.284)

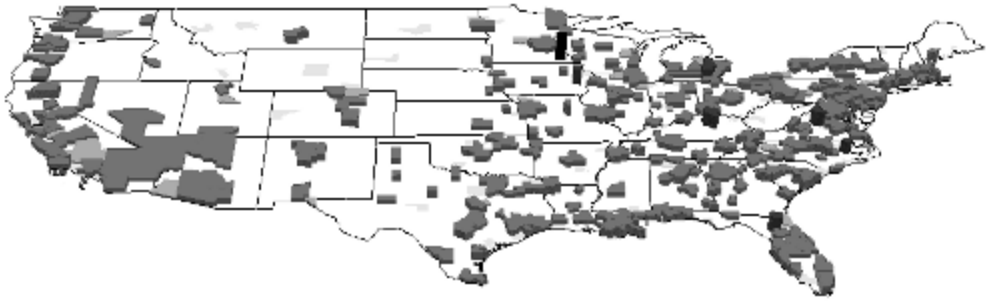
Note: Standard errors in parentheses. From March CPS 2003-2010. $\beta_1 - \beta_2$ is the estimated difference in the first two coefficients. Includes couples where the mean age is at most thirty-five, both spouses are present, and both spouses have a bachelor's degree (or higher).

* $p < 0.05$ (two-tailed test); ** $p < 0.01$; *** $p < 0.001$

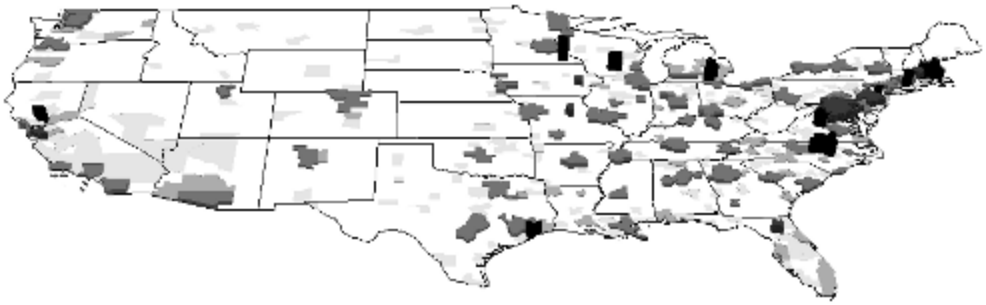
FIGURE 1

Metropolitan Workers Per Capita in Selected Occupations in 2000

(i) Physicians and Surgeons, C = 0.69 (un-normalized, C* = 0.12)



(ii) Medical Scientists, C = 1.81 (un-normalized, C* = 0.36)



(iii) Physicists and Astronomers, C = 2.02 (un-normalized, C* = 0.45)

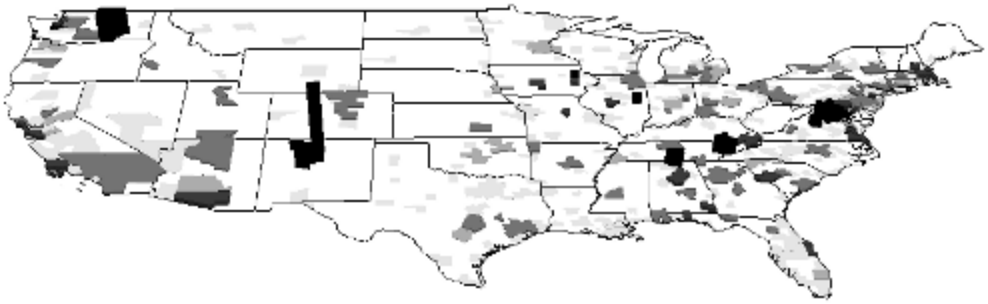
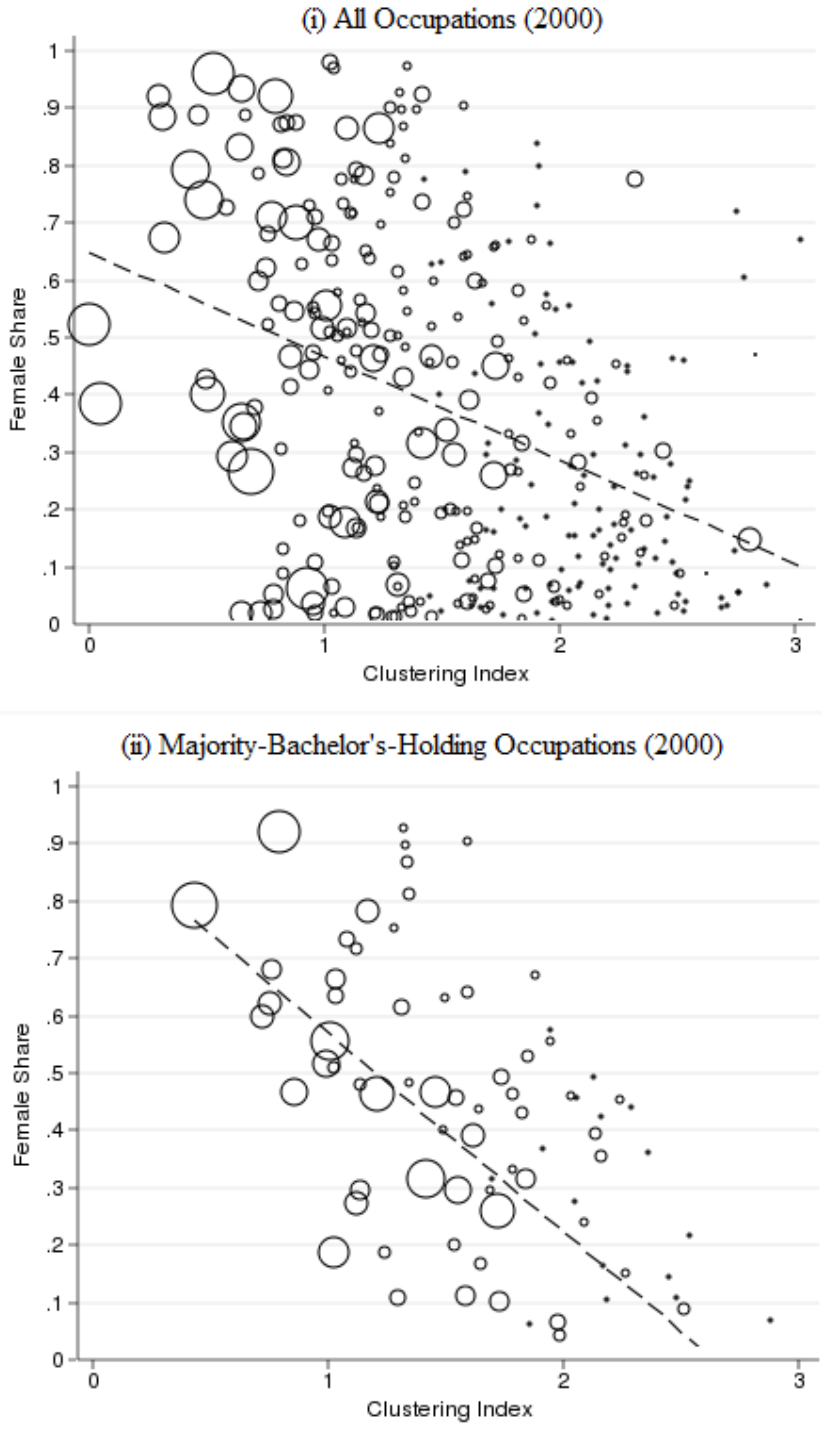


FIGURE 2

Scatterplots of Occupational Clustering versus Female Share,
By Year and Skill Level



*Note***.** Size denotes employment