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Cohort-level sex ratio effects on women's labor force participation

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Abstract It follows from a number of theoretical models of marriage that the scarcer women are relative to men, i.e., the higher the sex ratio, the less married women are likely to participate in the labor force. Such sex ratio effects may be stronger among less educated women. These predictions are tested using individual data from Current Population Surveys for four regions of the U.S. (Northeast, Midwest, South and West), and for the U.S. as a whole, covering the period 1965–2005 at 5-year intervals. Within-region sex ratio variation results from variation in cohort size (due principally to large fluctuations in number of births) and limited fluctuations in the difference between male and female age at marriage. As hypothesized, we find that sex ratios are inversely related to women's labor force participation, reflecting that ceteris paribus women born in years of peak baby-boom are more likely to be in the labor force than women born in years of peak baby-bust. Additionally, weaker sex ratio effects are found among educated women in two of the four regions of the United States.

Keywords Female labor force participation · Sex ratios · Cohorts · Education · Marriage markets

JEL Classification J1 · J2

1 Introduction

Women's labor force participation (LFP) rates in the U.S. experienced a marked increase between 1965 and 1980, the improvement being most remarkable for

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married women. For instance, during these 15 years, the LFP rate of married women ages 25–29 rose from 33%–58% in the labor force, an increase of 25% points. In contrast, in the next 15 years—from 1980 to 1995—the LFP rate for this age group grew by only 11% points, to 69% of the labor force. Furthermore, between 1995 and 2005, the LFP rate of married women ages 25–29 decreased slightly to 66.5%. Likewise, during the same 10 years, the LFP of married women ages 30–34 decreased by 2.5% points.

Economic explanations of such historical trends in female labor supply have principally focused on the effects of wages, income, educational attainment, and the number of children born to potential labor force participants. In this paper, we argue that historical fluctuations in cohort size, due principally to the changing number of births in adjacent cohorts, cause fluctuations in sex ratios—the ratio of men to women—that affect marriage market conditions. We predict that such fluctuations in sex ratio help explain female LFP rates. Additionally, we predict that sex ratio effects (regional and at the cohort level) will be stronger among women with less education. Both predictions, consistent with both Demand and Supply and collective economic theories of marriage, are tested using individual data from Current Population Surveys for the U.S. as a whole and for each of its four regions (Northeast, Midwest, South and West) covering the 1965–2005 period at 5-year intervals.

Our analysis expands on previous analyses of changes in women's LFP over time by Pencavel (1998) and Grossbard-Shechtman and Granger (1998) that are also based on the analysis of Current Population Survey (CPS) data. Pencavel (1998) documented considerable variation in women's LFP rates across cohorts and the limited explanatory power of fluctuations in wage and income—both variables typically included in economic models of labor supply. Using time-series aggregated CPS data for the period 1965 to 1990, Grossbard-Shechtman and Granger (1998) showed that the women experiencing the most rapid increases in LFP had been born in a growing cohort, i.e. a baby-boom.[1] Using aggregate CPS data, and similar aggregate data for Japan, Fukuda (2006) utilizes an innovative Bayesian cohort model that allows simultaneous estimation of age, period, and cohort effects on female labor force participation. While cohort effects in both countries were substantial—exceeding the size of period effects in their magnitude—the U.S. had larger cohort effects on female LFP than Japan.

We add to these studies by including cohort-level sex ratios in our analysis, by allowing for regional variation, and by expanding the time span of the analysis to 2005. Additionally, we examine how sex ratio effects on women's LFP rates differ by educational attainment. Finally, we carry out the analysis for both married women and for all women regardless of marital status to address variations in the sex ratio effects according to women's marital status.

[1] Earlier, Heer and Grossbard-Shechtman (1981) had documented the large increase in LFP among female baby-boomers and explained it in terms of the low sex ratios characterizing that generation.

2 Predictions

2.1 Sex ratios

It follows from at least three economic theories of marriage that sex ratios of marriage eligibles will be inversely related to married women's LFP: Grossbard's demand and supply (D&S) model of marriage (Grossbard-Shechtman 1984, 1993), and collective models by Chiappori (1988, 1992) and Rapoport, Sofer, and Solaz (2006). That collective models and D&S models lead to similar predictions is not surprising given that the second step in these two-step decision–making models are very similar. In both kinds of models, Step 2 consists of individuals deciding on their labor force participation by maximizing their own utility subject to a budget constraint that includes pre-determined access to the spouse's income. In both kinds of models, access to the spouse's income is determined in Step 1 in some form of agreement between the spouses, and the agreement is likely to favor more those spouses who benefit from advantageous marriage market conditions such as high sex ratios (high ratios of male to female).

Step 1 is based on different assumptions in collective models and D&S models. In collective models, access to the spouse's income is determined by a sharing rule agreed upon by the couple. This sharing rule may not depend on work in household production, as is the case in Chiappori (1988) and the Chiappori model predicting sex ratio effects on female LFP (Chiappori, Fortin, & Lacroix, 2002), or it may depend on work in household production, as in Chiappori (1997) and Rapoport et al. (2006). In Grossbard's D&S model, work in household production plays a central role and a couple agrees on the terms of trade between work in household production and access to income (interpreted as a quasi-wage). Individuals have a supply of marital household labor, which implies that they set apart time resources that they are not willing to allocate towards any household production performed in couple, given possible levels of quasi-wage.[2] Likewise, individual demands for such work involves a separation between what one is willing to purchase at different prices, and what is not demanded.

Collective models and D&S models differ in the role marriage markets play in influencing individual bargaining. In collective models, there are no market level prices for work in household production that influence individual bargaining agreements, in a manner analogous to the role of market wages in bargaining between employer and worker. Marriage market conditions influence individual bargaining power but not market-level terms of trade between work and access to income. In D&S models, when marriage partners agree that one is willing to supply what the other is willing to demand at a given quasi-wage, they are influenced by the terms of trade established in informal collective bargaining between many men and women interacting in the same marriage markets. Quasi-wage earners are typically the primary workers in household production, while their spouses are typically the primary market wage earners. To the extent that there is a traditional division of

[2] The separation between time one is willing to supply and time that remains devoted to self is reminiscent of a similar separation found in Lundberg and Pollak's (1993) ''Separate Spheres'' model.

labor by gender, as assumed in Becker (1973, 1981), men are likely to pay women quasi-wages.

It follows from both kinds of models that marriage market conditions influence the individual's decision on how to allocate her/his own time to work in the labor force. When marriage markets are more favorable to women, women's quasi-wages are higher (according to Step 1 in a D&S model) and the sharing rule favors women more (according to Step 1 in a collective model). In both kinds of models, more favorable marriage market conditions for women translate into more material resources for women (in Step 1) and, consequently, as a result of an income effect, women are less likely to participate in the labor force (in Step 2).[3]

One advantage of the D&S models is that they are reminiscent of labor market models and ideas can be imported from labor market analysis to the analysis of marriage markets. For example, one can look at education as a factor that influences the aggregate demand for work in marital production by women or men of different educational levels. Likewise, education can affect aggregate supply of such work. As a result, education may affect market levels of quasi-wage and therefore individual opportunities for intra-household bargaining. According to Grossbard-Shechtman (1984, 1993), in Step 1, more educated people may be more likely to seek non-material quasi-wages for work in marital household production, leading to a smaller income effect in Step 2. It is also possible that more educated women may have more egalitarian marriages in which work in marital household production is more frequently replaced by contracts to outsiders, such as restaurants and child care workers. This would lead to lower material compensation for household production in Step 1 and a smaller income effect in Step 2. For both of these reasons, we predict a smaller correlation between marriage market conditions and women's labor force participation among college-educated couples than among less educated couples (see Grossbard-Shechtman & Neuman, 1988). However, it is also possible that education raises productivity in work in household production and that when sex ratios increase educated women's quasi-wages, established in markets for more educated women's work in marital production, benefit more than is the case with the quasi-wages of less educated women.

According to both models, it is also possible for the sex ratio effects on women's LFP to vary with male income and education. Marriage market conditions may not simply be a matter of how many men per woman are available in a given marriage market, but also of the extent to which these men are willing to let women obtain access to their income. A higher sex ratio is more likely to induce women's withdrawal from the labor force if men have a more attractive financial position (on the importance of sex ratios incorporating men's earning power see Lichter, Anderson, & Hayward, 1995). It thus follows that there will be more of a negative correlation between married women's LFP and sex ratios at higher levels of male education and income.

[3] Many more of the same conclusions follow from both categories of models, including the prediction that married men will work more in the labor force if the sharing rule/quasi-wages are more in favor of women.

Sex ratio effects on women's LFP are also expected to vary according to their marital status. On the one hand, married women's access to men's income (Step 1) is likely to vary more with marriage market conditions (as measured by sex ratios) than that of women of any other marital status (single, cohabiting, separated, widowed or divorced). Accordingly, we expect stronger sex ratio effects on the LFP of married women than on the LFP of women of any marital status. On the other hand, the LFP of women of any marital status—married and unmarried—is expected to vary with the sex ratio more than that of married women for two reasons. First, sex ratio effects are likely to have an effect on marital status, the higher the sex ratio, the more women are likely to be married and married women work less in the labor force. Secondly, male willingness to give women access to their income will vary with their marital status. As such, analyses of the LFP of married women are more likely to provide evidence supporting the arguments found in D&S and collective models of marriage.

In the past, cross-city comparisons have provided evidence of a negative association between sex ratios and married women's labor supply. It has been found that married women are more likely to supply labor in cities where sex ratios are higher than average (Grossbard-Shechtman & Neideffer, 1997; Chiappori et al., 2002; Rapoport et al., 2006). However, a negative association between regional sex ratios and women's LFP is not necessarily caused by marriage market effects on LFP (income effects due to variation in quasi-wages or sharing rule). Good labor demand conditions for women may simultaneously lead women to participate more in the labor force and attract female migrants, thereby lowering sex ratios. This alternative explanation for a negative association between sex ratios and women's LFP only applies where labor migration is feasible. However, migration cannot occur across different birth cohorts and, therefore, evidence of sex ratio effects based on cohort comparisons is potentially more supportive of a marriage market theory of labor supply than evidence based on cross-regional or cross-city comparisons.

Sex ratios vary over time because of fluctuations in cohort size. Cohort size variations can cause fluctuations in sex ratio because, on average, men dating or marrying a particular group of women tend to be older than they are. For instance, if a particular cohort is larger than a preceding cohort, the sex ratio calculated for that particular group of women will be less than one.[4] Such is the case of cohorts of women born at the beginning of a baby boom and likely to marry men born prior to that baby boom who belong to smaller cohorts. Vice-versa, if a particular cohort of women is smaller than a preceding cohort, the sex ratio will be larger than one. For example, cohorts of women born at the beginning of a baby bust will typically marry men born prior to that baby bust who belong to larger cohorts.

Table 1 indicates substantial variation in this kind of sex ratio in the U.S.A. over the period 1965–2005 when we assume a difference of 2 years of age between men and women at marriage—an assumption in line with the overall evolution of that age difference. It follows from that table that women born during a baby boom

[4] Note that the more rigid age preferences, the more fluctuations in cohort size will cause marriage market imbalances.

Table 1 Generations of women, sex ratios, and changes in married women's labor force participation for four regions in the United States

Year of birth	Generation name	U.S. region	Sex ratio[1]	ΔLFP ages 25–29	ΔLFP ages 30–34	ΔLFP ages 35–39	ΔLFP ages 40–44
1926–1930	Pre-depression	**U.S.**	**.98**	**n.a.**	**n.a.**	**n.a.**	**4.96**
		NE	.95	n.a.	n.a.	n.a.	2.68
		Midwest	1.06	n.a.	n.a.	n.a.	10.05
		South	.96	n.a.	n.a.	n.a.	3.07
		West	.99	n.a.	n.a.	n.a.	3.31
1931–1935	Depression	**U.S.**	**1.00**	**n.a.**	**n.a.**	**6.69**	**5.65**
		NE	.96	n.a.	n.a.	4.80	3.87
		Midwest	1.06	n.a.	n.a.	8.87	5.63
		South	1.01	n.a.	n.a.	7.56	6.94
		West	1.01	n.a.	n.a.	3.87	5.05
1936–1940	New deal	**U.S.**	**.95**	**n.a.**	**6.84**	**5.09**	**9.23**
		NE	.95	n.a.	6.02	2.18	12.61
		Midwest	1.02	n.a.	5.85	6.03	9.75
		South	.92	n.a.	8.89	3.12	6.25
		West	.91	n.a.	6.22	9.75	10.35
1941–1945	World War II	**U.S.**	**.91**	**5.97**	**5.44**	**10.83**	**4.12**
		NE	.90	7.23	5.25	12.88	1.77
		Midwest	.93	6.69	6.28	12.33	3.96
		South	.88	1.74	3.34	9.34	4.52
		West	.92	9.86	7.17	9.47	5.68

Table 1 continued

Year of birth	Generation name	U.S. region	Sex ratio[1]	ΔLFP ages 25–29	ΔLFP ages 30–34	ΔLFP ages 35–39	ΔLFP ages 40–44
1946–1950	Post WW II	**U.S.**	**.87**	**11.84**	**13.51**	**6.61**	**7.05**
		NE	.89	11.05	15.26	8.61	11.11
		Midwest	.93	14.19	17.32	6.53	5.71
		South	.84	10.98	10.32	6.39	7.48
		West	.85	9.71	11.61	5.36	3.92
1951–1955	Korean War	**U.S.**	**.95**	**8.25**	**6.06**	**5.28**	**2.85**
		NE	.94	13.59	8.18	6.52	−.11
		Midwest	.99	8.80	5.55	4.82	6.84
		South	.93	6.70	6.45	3.86	.13
		West	.94	6.18	4.28	6.30	5.81
1956–1960	Sputnik	**U.S.**	**.97**	**7.40**	**3.82**	**1.40**	**1.23**
		NE	.97	5.44	2.46	.86	2.98
		Midwest	1.04	6.00	4.23	2.83	.27
		South	.93	7.26	5.59	2.47	3.29
		West	.96	9.84	2.81	−.78	−2.22
1961–1965	Kennedy	**U.S.**	**1.03**	**3.49**	**3.62**	**-1.84**	**n.a.**
		NE	1.01	2.54	4.54	−.77	n.a.
		Midwest	1.09	6.18	5.03	−2.67	n.a.
		South	1.01	6.90	2.98	−.58	n.a.
		West	1.01	−2.31	2.19	−2.97	n.a.

Table 1 continued

Year of birth	Generation name	U.S. region	Sex ratio[1]	ΔLFP ages 25–29	ΔLFP ages 30–34	ΔLFP ages 35–39	ΔLFP ages 40–44
1966–1970	Moon	**U.S.**	**1.06**	**−.14**	**−2.14**	**n.a.**	**n.a.**
		NE	1.05	2.59	−.80	n.a.	n.a.
		Midwest	1.16	4.28	−2.00	n.a.	n.a.
		South	1.03	−5.06	−1.22	n.a.	n.a.
		West	1.02	−.38	−4.45	n.a.	n.a.
1971–1975	Roe	**U.S.**	**1.07**	**.76**	**n.a.**	**n.a.**	**n.a.**
		NE	1.05	1.62	n.a.	n.a.	n.a.
		Midwest	1.11	−.11	n.a.	n.a.	n.a.
		South	1.06	4.06	n.a.	n.a.	n.a.
		West	1.06	−1.24	n.a.	n.a.	n.a.
1976–1980	First echo	**U.S.**	**1.01**	**n.a.**	**n.a.**	**n.a.**	**n.a.**
		NE	1.01	n.a.	n.a.	n.a.	n.a.
		Midwest	1.08	n.a.	n.a.	n.a.	n.a.
		South	.97	n.a.	n.a.	n.a.	n.a.
		West	.98	n.a.	n.a.	n.a.	n.a.

Notes: [1] Ratio of men age 22–26 to women age 20–24 or men age 27–31 to women age 25–29 calculated based on Census data from 1940 to 2000. The age group depends on the Census year. Sex ratios for last generation were calculated based on the 2000 Census using younger age groups. [2] Calculated from CPS years 1965–2005 (baby-boomers) would be more likely to participate in the labor force than baby-busters (Grossbard-Shechtman & Granger, 1998; Guttentag & Secord, 1983; Heer & Grossbard-Shechtman, 1981). This explains, for instance, why in the U.S. there was rapid growth in the LFP of married women ages 25–29 in the years 1965–1980. These are precisely the years during which baby-boomers were reaching these ages. In contrast, married women entering ages 25–29 in the period 1980–1995 were born during the baby-bust. Not surprisingly, relative to their baby-boom counterparts (such as the cohort born in 1946 that reached age 25 in 1971), these baby-bust women (such as the cohort born in 1964 that reached age 25 in 1989) have experienced a substantially slower growth in participation in the labor force.

Easterlin (1980) offers an alternative explanation for the inverse relationship between the LFP of women and fluctuations in cohort size: growing cohorts, such as baby-boomers, face worse income opportunities than the ones encountered by their parents when they were growing up. Baby-boom women thus may meet baby-boom men with relatively low incomes. This would also push married women into the labor force. Furthermore, according to Easterlin, baby-boom couples are also expected to have fewer children, which would also support the prediction that married baby-boom women have higher LFP rates. However, if sex ratio effects on married women's LFP persist after appropriately accounting for household income and fertility, this alternative explanation can be ruled out.

2.2 Other variables affecting female labor force participation

We also build on past research on women's LFP rates and include the following variables in our analyses.

2.2.1 Household income and female wages

According to Mincer (1962), higher wages had been a major reason why women were attracted to join the labor force prior to 1960. Mincer solved a puzzle that had confounded labor economists at the time: time series results showed that women's LFP and wages were growing in the same direction, in apparent contradiction to findings of a negative association between wages and women's LFP based on cross-sectional data. Mincer resolved this puzzle by separating the effects of male and female wages. What explained women's entry into the labor force in time series were increases in women's wages, whereas increases in male wages accounted for the negative association between wages and women's LFP in cross-sections studies. Mincer interpreted the effect of married women's own wages on their LFP rates as a substitution effect and the effect of husbands' wages primarily as an income effect.

While this wage/income explanation has held for earlier periods, its effectiveness in explaining recent trends in LFP seems limited. A number of studies have indicated that, in recent years, women's wages and their LFP have not been moving in the same direction. Rosen (1992) pointed out that the LFP rates of women increased greatly during the 1970's, when women's wages were stagnant or declining. It is possible that a low positive association between female wages and female labor supply reflects the fact that women entering the labor market for the first time had low levels of human capital.

Based on cross-sections from various years of the CPS, Leibowitz and Klerman (1995) found that, relative to women's wages, men's wages and unemployment explained more changes in married mothers' employment between 1971 and 1990. A possible explanation for the stronger effect exhibited by male wages on women's LFP relative to female wages is that female wages are endogenous in a study of female LFP. For any given demand for female labor, changes in the supply of that labor will cause fluctuations in wages. Therefore, we carry our analysis with and without controls for female non-labor income and for average wages earned by other women in the same age-education-region-year cell to assess the robustness of sex ratio effects in explaining female LFP to the inclusion of these two regressors.

2.2.2 Education

Previous studies have found that the rise in women's LFP was associated with increased levels of education.[5] This is consistent with the view that education improves the individual's preparation for the job market, raising the opportunity cost of leisure and home production, leading to increased participation in the labor market. Additionally, educational attainment may also impact women's LFP via its impact on marriage opportunities and on preferences. Educated couples tend to have a preference for a more egalitarian division of labor (Hersch, 2003), which in turn may facilitate married women's labor force participation. Therefore, we account for women's educational attainment and, as with female wages, we carry out the analysis with and without controls for educational attainment in order to assess the robustness of our sex ratio effect estimates to the inclusion of this potentially endogenous regressor.

2.2.3 Fertility

The growth in women's LFP rates over time has also been shown to coincide with decreases in fertility.[6] Causality here can go either way. Not only is it possible that lower fertility explained increases in women's labor supply, but higher LFP rates may have also lowered fertility. More generally, labor supply and fertility may also be spuriously related due to the effect of other variables on both labor supply and fertility (e.g. Deville, 1977; Lehrer & Nerlove, 1986). Acknowledging the importance of women's childbearing on their LFP choices, and despite the difficulties of isolating its true impact due to endogeneity reasons, we control for women's fertility in some specifications with a dummy variable indicative of the presence of children less than three years old living in the household.

[5] See, for example, Huet (1977), Shapiro and Shaw (1983), Smith and Ward (1984), Mincer (1985), Goldin (1990), and Leibowitz and Klerman (1995).

[6] See, for example, Mincer (1962), Deville (1977), Ekert (1983), Smith and Ward (1984), Mincer (1985), Goldin (1990), Rosen (1992), and Leibowitz and Klerman (1995).

2.2.4 Change in attitudes

Partially as a result of the decreasing explanatory power of wage/income variables,[7] scholars—especially those trained in sociology—have turned to cultural explanations focused on variables such as attitudes towards work and family. We include a time trend, region of residence dummies, and interaction terms between the time trend and the set of regional dummies to allow for differences in attitudes towards work and family across regions as well as over time.

In the next section, we examine the LFP effects of sex ratios over time (cohort size effects) and across regions while accounting for the aforementioned factors possibly influencing female LFP decisions.

3 Data and empirical methodology

3.1 Data

With the exception of the data on sex ratios, the data used in this study are extracted from Current Population Surveys (CPS, March files) for four regions (Northeast, Midwest, South and West) covering the period 1965–2005 at 5-year intervals. The CPS data are all individual-level data, except for average female wage, that is calculated for each region/age group/year cell. Non-labor household income is defined as total household income minus the respondent's labor income.

Our sample includes women 25–44, and we create age dummies for the following age groups: 25–29, 30–34, and 35–39. Since women with young children at home have a higher likelihood of receiving intra-marriage transfers of income, we only consider women younger than 44 years old.

In calculating sex ratios, we assume that the male/female age difference at marriage is, on average, equal to 2 years. This assumption fits the data on first marriages in the U.S. for most of the period that we cover:[8] in the period 1965–1995, despite a considerable rise in age at marriage, the age difference between male and female age at first marriage fluctuated between 2.2 and 2.7.[9] Sex ratios are calculated from Census data for 5-year age groups by dividing the number of all men 2 years older by the number of all women ages 20–24 or 25–29 (depending on the year), regardless of marital status, i.e. we compute:

$\textit{Sex Ratio} = (M_{t-2} + M_{t-1} + M_t + M_{t+1} + M_{t+2})/(F_t + F_{t+1} + F_{t+2} + F_{t+3} + F_{t+4})$, where M is the number of men and F the number of women. We expect that people are most likely to be influenced by sex ratios when they are in their twenties and are most likely to enter a first marriage. What moves cohort-level variation in sex ratios is the difference between the number of women born in years $(t + 3)$ and $(t + 4)$ and

[7] In this regard, Pencavel (1998) posits that variation in male and female wages accounts for less than half of the observed changes in women's LFP rates over time.

[8] A more complete analysis would endogenize the gender difference in age at marriage, as it is also likely to vary with sex ratios.

[9] Since 1997 this gender difference in age at marriage has shrunk to an average of 1.7, fluctuating between 1.6 and 1.8 (www.census.gov).

the number of men born in years ($t-1$) and ($t-2$).[10] The resulting cohort-level sex ratios are merged into the CPS data according to women's birth year and region of residence.

Sex ratios for women born between 1926 and 1980 and for men born between 1924 and 1978 for the entire country and by U.S. regions are shown in Table 1. Each 5-year cohort was given a name related to historical events that occurred around their years of birth. For the U.S. as a whole (bold numbers), it can be seen that this sex ratio fluctuated dramatically from a minimum of .87 for the women born in the years 1946–1950, right after World War II, (men born 1944–1948), to a maximum of 1.07 for the women born in 1971–1975 (men born 1969–1973) around the passage of Roe versus Wade, a landmark ruling that led the number of abortions to increase in the United States.[11] The sex ratio for women born in 1966–1970 and men born in 1964–1968, the Moon generation, was also high at 1.06. It is also apparent from Table 1 that sex ratios were consistently higher in the Midwest than in other regions.

Table 1 also reports changes in LFP rates for married women of different ages, these changes being defined over the last five years.[12] The table indicates a negative correlation between sex ratio and changes in married women's LFP. For example, it can be seen that, at almost every age, the Post-World War II generation women—the women with the lowest sex ratio—experienced faster growth in LFP than the other 5-year cohorts of women. For example, on average, married women age 30–34 experienced an increase of 13.5% points in LFP between 1975 and 1980, corresponding to the entry of the women of the Post-World War II generation (born in 1946–1950) into this age group. In contrast, between 1995 and 2005, women of the baby-bust cohorts (the Moon and Roe generations characterized by unusually high sex ratios) entered this age group and the LFP of married women declined by 3.6% points: from 73% to 69.4%.[13]

We analyze sex ratio effects on the LFP of two samples of women: (1) about 200,000 women of varying marital status, and (2) around 140,000 married women. Appendix Table 11 in the appendix presents means, standard deviations, and definitions for the variables used in our analysis for the two samples for the U.S. as a whole. Descriptive statistics by U.S. region are provided in Appendix Table 12.

3.2 Empirical methodology

We first specify a parsimonious model of women's LFP that excludes any potentially endogenous regressors:

[10] The numbers of men and women born in years t, ($t + 1$), and ($t + 2$) are roughly equal and appear in both the numerator and the denominator.

[11] Links between abortion law changes and changes in fertility in the 1970s have been discussed, e.g. by Donohue and Levitt (2001) and Angrist and Evans (2002). Oreffice (2007) directly studies the association between abortion law changes and LFP.

[12] A similar table for women of all marital states is available from the authors upon request.

[13] Table 1 indicates that this decline covered all regions: the LFP rate of married women declined in every U.S. region during those 10 years.

$$\begin{aligned}\Pr(LFP_i = 1) = \Phi(&\beta_1 SR_{ir} + \beta_2 trend + \beta_3 A_i + \beta_4 NLI_i \\ &+ \beta_5 R_i + \beta_6 R_i * SR + \beta_7 R_i * trend)\end{aligned} \tag{1}$$

where Φ stands for the standard normal cumulative distribution, *LFP* equals 1 if women are in the labor force, *SR* is the sex ratio for women in a particular birth cohort and region, *A* indicates the particular age group to which the female respondent belongs to, *NLI* represents women's non-labor income, *R* stands for region dummies, and *trend* is a time trend. Interactions between the sex ratios and the region dummies recognize regional differences in changes in cohort size due principally to regional differences in fertility, inter-region as well as international migration. Interactions between the time trend and the region dummies address difficult to capture trends in work and family attitudes, as well as differences in the adoption timing of specific time-saving technologies across regions (see Greenwood & Guner, 2004)—all factors that could impact the LFP of women at different rates depending on the region where they live.

However, the model specified in Eq. (1) ignores the role played by other factors known to affect women's LFP discussed earlier, including their educational attainment, average wages earned by similar women, and the number of young children they may have. Therefore, we also estimate a model that includes women's educational attainment (E^f), the average market wage for women in their age/education group residing in a specific region in a particular year $\left(\overline{w}^f_{AERT}\right)$, and the number of young children (*F*), which results in our *second* specification:

$$\begin{aligned}\Pr(LFP_i = 1) = \Phi(&\beta_1 SR_{ir} + \beta_2 trend + \beta_3 A_i + \beta_4 NLI_i + \beta_5 R_i + \beta_6 R_i * SR \\ &+\beta_7 R_i * trend + \beta_8 E^f_i + \beta_9 \overline{w}^f_{AERT} + \beta_{10} F_i)\end{aligned} \tag{2}$$

Yet, if we want to further assess whether sex ratios have a differential effect on married women's LFP rates depending on their educational attainment, we need to interact the sex ratio (*SR*) with women's educational attainment (E^f). It follows from our earlier discussion that the coefficient of this interaction term is expected to be negative. We also include interactions between women's educational attainment and region of residence to account for regional differences in educational levels and policies. Finally, a three-way interaction term of sex ratios, women's educational attainment, and region dummies allows us to test whether there are regional differences in sex ratio effects by educational attainment. Hence, our *third* model specification is given by:

$$\begin{aligned}\Pr(LFP_i = 1) =\Phi(&\beta_1 SR_{ir} + \beta_2 trend + \beta_3 A_i + \beta_4 NLI_i + \beta_5 R_i + \beta_6 R_i * SR \\ &+ \beta_7 R_i * trend + \beta_8 E^f_i + \beta_9 \overline{w}^f_{AERT} + \beta_{10} F + \beta_{11} E^f_i * SR \\ &+ \beta_{12} E^f_i * R + \beta_{13} R_i * E^f_i * SR)\end{aligned} \tag{3}$$

All three models are estimated separately for: (a) women of any marital status and (b) for married women to gauge the robustness of the sex ratio effect on women's LFP depending on their marital status.

4 Results

Estimations of models 1 through 3 for the U.S. as a whole are displayed in Tables 2–4. The corresponding results by U.S. region are shown in Tables 5–7.[14] Each table presents results for women of any marital status and for married women. Models 1 and 2 involve a test of our main prediction: that sex ratios and women's LFP are inversely related. In accordance with our prediction, we find a negative association between sex ratio and women's LFP in all estimations of models 1 and 2, whether we just consider married women or whether we look at women of any marital status. For example, according to parsimonious model 1 (Table 2), an increase in the sex ratio from 1.00 to 1.10 is associated with a decrease in the LFP of married women of 2.15% points. The corresponding decrease in the LFP of women of any marital status is 3% points.

The regressions presented in Tables 3 and 6 test model 2. They show that, despite weakening, sex ratio effects on women's LFP remain robust to the inclusion of women's wages, education, and presence of young children in the analysis. For instance, the effect on LFP of an increase in sex ratio of .10 is comparable to that of an increase in education of a little more than 2 years. The lower sex ratio effects in Table 3 (after controlling for education, income and the presence of small children in the household) than in Table 2 (which lacks such controls) is suggestive of both: Easterlin's and the marriage market explanation of stronger sex ratio effects in the parsimonious model of Table 2. They also agree with Fukuda's (2006) cohort results using Japanese data.

In the regressions in Tables 2 and 3, sex ratio effects are the result of both regional and cohort-level variation in sex ratios. The regional models in Table 5 estimate the parsimonious model for each region separately. Within each region, the sex ratio effect originates from differences in cohort size due to fluctuations in the number of births around women's birth year. Sex ratio effects on women's LFP are smaller in the Midwest than in any of the other major three regions of the U.S., while they are strongest in the Northeast and the West. For example, in the West, an increase in sex ratio from 1.00 to 1.10 is associated with a decrease in the LFP of married women of up to 6.5% points. One can only speculate as to why there is a greater effect of sex ratio on women's LFP in these regions. Could there be more variation in women's bargaining power in marriage (an implication of Step 1 in both the D&S and collective models) in certain regions? Perhaps such bargaining power is related to sex ratios, with the Midwest having a substantially higher mean sex ratio over the period under study. For instance, the average sex ratio for the period under analysis was 1.03 in the Midwest compared to .96 in the Northeast and West, and .95 in the South. This means that marriage market conditions in the Midwest are more favorable to women, which could possibly stem from selective gender migration.

[14] We have also estimated models 2 and 3 excluding fertility to gauge the robustness of the sex ratio effects to the exclusion of this variable. Results were very similar to the ones we report and are available from the authors upon request.

Table 2 Probit regression estimates of women's labor force participation: Model (1)

Variables	Married women			All women		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−.572***	.118	−.215	−.873***	.104	−.301
Age 25–29	−.197***	.010	−.075	−.113***	.009	−.040
Age 30–34	−.184***	.010	−.070	−.120***	.009	−.043
Age 35–39	−.069***	.010	−.026	−.053***	.009	−.019
NLI	−5.14E-6***	1.76E-7	−1.93E-6	−7.53E-6***	1.56E-7	−2.7E-6
NE	.968***	.211	.310	.331*	.108	.111
South	.406**	.153	.147	.120	.113	.042
West	1.207***	.170	.373	.651***	.148	.207
Time trend	.033***	.001	.012	.030***	.001	.011
NE*Time trend	−.002	.001	−.001	−.002**	.001	−.001
South*Time trend	−.001***	.001	−.004	−.007***	.001	−.003
West*Time trend	−.011***	.001	−.004	−.010***	.001	−.004
NE*SR	−1.102***	.228	−.414	−.463**	.195	−.164
South*SR	−.295*	.163	−.111	−.074	.143	−.026
West*SR	−1.159***	.183	−.435	−.608***	.160	−.216
Number of observations	140592			197991		
Wald Chi-square	8677.45			10784.86		
Prob > Chi-square	.000			.000		

Notes: S.E. stands for 'standard error' and M.E. for 'marginal effect.' * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 and Midwest are used as reference categories

We also predicted that the inverse association between the sex ratio and female LFP is less likely to hold for more educated women than for women with less education, i.e. the interaction term between the sex ratio and women's years of schooling in Eq. (3) is expected to be positive. Model 3, containing an interaction term between the sex ratio and female education and the corresponding regional dummies, is estimated for the U.S. as a whole (Table 4) and for each major region (Table 7). The results from Table 4 indicate that the interaction between education and sex ratio has a negative sign among married women, whereas the simple effect of sex ratio is positive in the case of the reference region (the Midwest). The net effect of sex ratio in the Midwest and the Northeast is negative (obtained after adding the negative marginal effect −.08 to the positive marginal effect .13), while the interaction between education and sex ratio has a positive sign. Among women of any marital status, the interaction between education and sex ratio is not significant, except in the West, where it takes a positive sign.

In light of the regional differences pointed out earlier, we also display the results by region. It is also apparent from the results in Table 7 that sex ratio effects vary by education in opposite ways in the Midwest relative to the Northeast and West of the U.S. Our prediction was that material quasi-wages

Table 3 Probit regression estimates of women's labor force participation: Model (2)

Variables	Married women			All women		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−.219*	.120	−.082	−.558***	.108	−.195
Age 25–29	.015	.012	.005	.064***	.010	.022
Age 30–34	.045***	.011	−.017	.002	.010	.001
Age 35–39	−.019*	.011		−.009	.009	−.003
NLI	−8.0E-6***	2.0E-7	−2.9E-06	−9.95E-6***	1.8E-7	−3.5E-6
NE	.692***	.216	.232	.173	.185	.059
South	.212	.157	.077	−.095	.139	−.033
West	.577***	.178	.199	−.017	.157	−.006
Time trend	.016**	.001	.006	.012***	.001	.004
NE*Time trend	−.003***	.001	−.001	−.003***	.001	−.001
South*Time trend	−.010***	.001	−.004	−.009***	.001	−.003
West*Time trend	−.005***	.001	−.002	−.004***	.001	−.001
NE*SR	−.801***	.233	−.299	−.293	.200	−.102
South*SR	−.037	.168	−.014	.224	.148	.078
West*SR	−.583***	.190	−.217	.004	.168	.001
Young children	−.425***	.007	−.158	−.463***	.006	−.161
Education	.093***	.001	.035	.106***	.001	.037
Average wages	7.2E-5***	5.6E-6	2.7E-05	7.1E-5***	4.7E-6	2.5E-5
Number of observations	140592			197189		
Wald Chi-square	15007.38			22117.23		
Prob > Chi-square	.000			.000		

Notes: S.E. stands for 'standard error' and M.E. for 'marginal effect.' * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 and Midwest are used as reference categories

matter more in marriages involving less educated women, among whom traditional gender roles are likely to play a more important role than among marriages involving more educated women. This prediction seems to apply in the West and the Northeast, but not in the Midwest. The possibility exists that economic conditions in the Midwest induced more educated married women in this region to step into the labor force at a higher rate than their counterparts in the Northeast and West of the U.S.

Next, we investigate whether the sign on the interaction terms between female education and the sex ratio possibly originated from a positive correlation between male and female education. Specifically, for married women, we estimate a fourth model similar to model (3) that, in addition to the sex ratio interacted with the wife's education, includes the husband's education and interactions between the sex ratio and the husband's education. This model was first estimated for the U.S. as a whole. Results are displayed in Table 8. According to our estimates in Table 8, adding husbands' educational attainment and the corresponding interaction terms does not

Table 4 Probit regression estimates of women's labor force participation: Model (3)

Variables	Married Women			All Women		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	2.623***	.560	.978	−.836	.513	−.292
Age 25–29	.017***	.012	.006	.064***	.010	.022
Age 30–34	−.043	.011	−.016	.002	.010	−.001
Age 35–39	−.020	.011	−.007	−.010	.009	−.003
NLI	−7.9E-6***	2.0E-7	−3.0E-6	−10.0E-6***	1.8E-7	−3.5E-6
NE	4.784***	1.006	.749	1.103	.832	.308
South	3.050***	.761	.699	−.731	.657	.267
West	5.162***	.801	.802	1.273*	.699	.346
Time trend	.016***	.001	.006	.012***	.001	.004
NE*Time trend	−.002**	.001	−.001	−.003***	.001	−.001
South*Time trend	−.010***	.001	−.004	−.009***	.001	−.003
West*Time trend	−.005***	.001	−.002	−.003***	.001	−.001
NE*SR	−4.510***	1.028	−1.714	−1.261	.849	−.440
South*SR	−2.850***	.775	−1.029	.857	.662	.230
West*SR	−5.170***	.815	−1.928	−1.288*	.704	−.450
Young children	−.424***	.007	−.159	−.463***	.006	−.162
Education	.323***	.046	.121	.086**	.040	.030
Average wages	7.2E-5***	5.6E-6	2.7E-5	7.1E-5***	4.7E-6	2.5E-5
Education*SR	−.211**	.045	−.082	.022	.039	.008
NE*Education	−.315	.074	−.118	−.071***	.062	−.024
South* Education	−.220***	.059	−.082	.050	.502	.018
West*Education	−.356**	.061	−.133	−.102*	.053	−.036
NE*Education*SR	.290	.076	.108	.074	.063	.026
South*Education*SR	.211	.051	.079	−.050	.051	−.017
West*Education*SR	.357**	.061	.133	.102*	.053	.035
Number of observations	140952			197189		
Wald Chi-square	15137.15			22102.14		
Prob > Chi-square	.000			.000		

Notes: S.E. stands for 'standard error' and M.E. for 'marginal effect.' * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 and Midwest are used as reference categories

increase the explanatory power of our model. The new terms are statistically not different from zero.

However, because regional differences could mask cohort-level sex ratio effects, similar regressions are also estimated by region in Table 9. We find evidence of a larger negative sex ratio effect on the LFP of women married to more educated men in the Northeast and the South. This result confirms the importance of incorporating men's earning power when calculating how sex ratios affect marriage conditions, as argued by Lichter et al. (1995).

Table 5 Regression estimates of women's labor force participation by region: Model (1)

Variables	Married: Northeast			All: Northeast			Married: Midwest			All: Midwest		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−1.536***	.158	−.562	−1.601***	.193	−.614	−.794***	.097	−.278	−.860***	.105	−.300
Age 25–29	−.044***	.019	−.016	−.186***	.023	−.072	−.125***	.018	−.045	−.131***	.018	−.047
Age 30–34	−.131***	.018	−.049	−.205***	.021	−.080	−.139***	.018	−.049	−.143***	.018	−.051
Age 35–39	−.064***	.018	−.023	−.093***	.021	−.036	−.039**	.018	−.014	−.040***	.018	−.014
NLI	−5.1E-6***	2.7E-7	−1.9E-6	−5.5E-6***	3.8E-7	−2.1E-6	−7.6E-6***	2.9E-7	−2.6E-6	−8.5e-06***	3.36e-07	−2.9E-6
Time trend	.027***	.001	.010	.031***	.001	.012	.029***	.001	.010	.030***	.0006154	.010
Observations	53743			33529			57881			47052		
Wald Chi-square	2602.41			2363.07			3769.42			3528.79		
Prob > Chi-square	.000			.000			.000			.000		
Variables	Married: South			All: South			Married: West			All: West		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−.904***	.091	−.320	−.843***	.108	−.318	−1.219***	.114	−.438	−1.557***	.134	−.593
Age 25–29	−.089***	.016	−.032	−.168***	.019	−.064	−.169***	.018	−.062	−.210***	.021	−.081
Age 30–34	−.070***	.016	−.025	−.110***	.018	−.042	−.137***	.018	−.050	−.180***	.020	−.070
Age 35–39	−.030*	.016	−.011	−.043**	.018	−.016	−.079***	.018	−.029	−.097***	.021	−.037
NLI	−7.3E-6***	2.6E-7	−2.6E-6	−5.0E-6***	3.1E-7	−1.9E-6	−6.4E-6***	3.3E-7	−2.3E-6	−3.2E-6***	3.6E-7	−1.2E-6
Time trend	.022***	.001	.008	.024***	.001	.009	.019***	.001	−.438	.021***	.001	.008
Observations	73387			48395			57056			37650		
Wald Chi-square	3048.15			2363.26			1654.14			1348.52		
Prob > Chi-square	.000			.000			.000			.000		

Notes: S.E. stands for 'standard error' and M.E. for 'marginal effect.' * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 is used as reference categories

Table 6 Regression estimates of women's labor force participation by region: Model (2)

Variables	Married: Northeast			All: Northeast			Married: Midwest			All: Midwest		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−.609***	.170	−.220	−.875***	.205	−.334	−.402***	.103	−.138	−.058	.118	−.021
Age 25–29	.086***	.020	.031	−4.4E-4	.024	.000	.036*	.020	.012	−.055**	.024	−.020
Age 30–34	−.011	.019	−.004	−.072***	.023	−.027	−.022	.020	−.008	−.153***	.023	−.057
Age 35–39	−.017	.019	−.006	−.041*	.022	−.016	−.013	.019	−.004	−.035*	.021	−.013
NLI	−7.7E-6***	3.1E-7	−2.8E-6	−7.6E-6***	4.3E-7	−2.9E-6	−9.7E-6***	3.3E-7	−3.3E-6	−9.1E-6***	4.2E-7	−3.4E-6
Time trend	.012***	.001	.004	.024***	.002	.009	.020***	.001	.007	.026***	.002	.010
Young children	−.550***	.012	−.198	−.497***	.014	−.190	−.484***	.011	−.166	−.433***	.013	−.160
Education	.100***	.002	.036	.078***	.003	.030	.107***	.003	.037	.105***	.003	.039
Average wages	6.1E-5***	5.7E-6	2.2E-5	2.8E-5***	8.4E-6	1.1E-5	3.4E-5***	6.2E-6	1.2E-5	2.1E-5***	8.4E-6	7.6E-6
Observations	53551			33449			57722			39868		
Wald Chi-square	6089.54			3829.60			6810.15			5054.68		
Prob > Chi-square	.000			.000			.000			.000		
Variables	Married: South			All: South			Married: West			All: West		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−.361***	.102	−.126	−.484***	.120	−.181	−.376***	.129	−.133	−.580***	.155	−.219
Age 25–29	.012	.018	.004	−.066***	.021	−.025	.087***	.022	.030	.076***	.027	.029
Age 30–34	−.018	.017	−.006	−.064***	.020	−.024	.019	.020	.007	−.011	.023	−.004
Age 35–39	−.016	.017	−.006	−.037**	.019	−.014	−.016	.019	−.006	−.026	.022	−.010
NLI	−1.1E-5***	3.1E-7	−3.8E-6	−8.4E-6***	3.8E-7	−3.2E-6	−8.7E-6***	4.1E-7	−3.1E-6	−6.4E-6***	4.8E-7	−2.4E-6
Time trend	.012***	.001	.004	.021***	.002	.008	.010***	.001	.003	.011***	.002	.004
Young children	−.442***	.010	−.154	−.400***	.012	−.150	−.483***	.011	−.171	−.443***	.012	−.168

Table 6 continued

Variables	Married: South			All: South			Married: West			All: West		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
Education	.109***	.002	.038	.098***	.003	.037	.099***	.002	.035	.096***	.003	.036
Average wages	3.2E-5***	6.1E-6	1.1E-5	8.0E-8	8.0E-6	3.0E-8	7.0E-5***	7.7E-6	2.5E-5	7.8E-5***	9.9E-6	2.9E-5
Observations	73073			48245			56740			37489		
Wald Chi-square	7250.65			4567.44			5350.18			6305.48		
Prob > Chi-square	.000			.000			.000			.000		

Notes: S.E. stands for 'standard error' and M.E. for 'marginal effect.' * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 is used as reference categories

Table 7 Regression estimates of women's labor force participation by region: Model (3)

Variables	Married: Northeast			All: Northeast			Married: Midwest			All: Midwest		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−1.243**	.614	−.448	−.628	.798	−.240	−.140	.469	−.048	2.907***	.567	1.077
Age 25–29	.086***	.020	.031	.000	.024	.000	.037*	.020	.012	−.045*	.024	−.017
Age 30–34	−.011	.019	−.004	−.071***	.023	−.027	−.021	.020	−.007	−.140***	.023	−.053
Age 35–39	−.018	.019	−.006	−.041*	.022	−.016	−.012	.019	−.004	−.030	.022	−.011
NLI	−7.8E-6***	3.1E-7	−2.8E-6	−7.6E-6***	4.3E-7	−2.9E-6	−9.7E-6***	3.3E-7	−3.3E-6	−9.0E-6***	4.2E-7	−3.3E-6
Time trend	.012***	.001	.004	.024***	.002	.009	.020***	.001	.007	.026***	.002	.009
Young children	−.550***	.012	−.198	−.497***	.014	−.190	−.484***	.011	−.166	−.432***	.013	−.160
Education	.053	.043	.019	.096*	.056	.037	.128***	.037	.044	.341***	.044	.127
Average wages	6.1E-5***	5.7E-6	2.2E-5	2.8E-5***	8.4E-6	1.1E-5	3.4E-5***	6.3E-6	1.2E-5	2.5E-5***	8.5E-6	9.2E-6
Education*SR	.049	.045	.018	−.019	.058	−.007	−.020	.036	−.007	−.231***	.043	−.086
Observations	53551			33449			57722			39868		
Wald Chi-square	6080.56			3833.92			6825.03			5161.73		
Prob > Chi-square	.000			.000			.000			.000		
Variables	Married: South			All: South			Married: West			All: West		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	−.091	.384	−.032	−.502	.471	−.188	−1.295***	.455	−.458	−1.740***	.539	−.659
Age 25–29	.013	.018	.005	−.066***	.021	−.025	.087***	.022	.030	.077***	.027	.029
Age 30–34	−.017	.017	−.006	−.064***	.020	−.024	.018	.020	.006	−.012	.023	−.005
Age 35–39	−.016	.017	−.005	−.037**	.019	−.014	−.016	.019	−.006	−.026	.022	−.010
NLI	−1.1E-5***	3.1E-7	−3.8E-6	−8.4E-6***	3.8E-7	−3.2E-6	−8.7E-6***	4.1E-7	−3.1E-6	−6.5E-6***	4.8E-7	−2.4E-6
Time trend	.012***	.001	.004	.021***	.002	.008	.010***	.001	.003	.011***	.002	.004

Table 7 continued

Variables	Married: South			All: South			Married: West			All: West		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
Young children	−.442***	.010	−.154	−.400***	.012	−.150	−.482***	.011	−.171	.012	.009	.003
Average wages	3.2E-5***	6.2E-6	1.1E-5	5.6E-8	8.1E-6	2.1E-8	7.0E-5***	7.7E-6	2.5E-5	9.9E-6	.091**	.034
Observations	73073			48245			56740			37489		
Wald Chi-square	7268.07			4568.79			5332.11			3591.66		
Prob > Chi-square	.000			.000			.000			.000		

Notes: S.E. stands for ‘standard error’ and M.E. for ‘marginal effect.’ * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 is used as reference categories

Table 8 Regression Estimates of married women's labor force participation

Variables	Model (3) with husband information		
	Coeff.	S.E.	M.E.
SR	2.891***	.639	1.078
Age 25–29	.025**	.012	.009
Age 30–34	−.035***	.011	−.013
Age 35–39	−.014	.011	−.005
NLI	−7.4E-06***	2.1E-07	−2.8E-06
NE	4.115***	1.08	.691
South	2.806***	.815	.661
West	5.481***	.866	.821
Time trend	.016***	.001	.006
NE*Time trend	−.001	.001	−.001
South*Time trend	−.010***	.001	−.004
West*Time trend	−.005***	.001	−.002
NE*SR	−3.839***	1.106	−1.431
South*SR	−2.492***	.822	−.929
West*SR	−5.521***	.874	−2.058
Young children	−.425***	.007	−.159
Education	.338***	.057	.126
Average wages	7.4E-05***	5.6E-06	2.7–05
Education*SR	−.211***	.056	−.082
NE*Education	−.380***	.094	−.142
South* Education	−.271***	.074	−.101
West*Education	−.335***	.077	−.125
NE*Education*SR	.365***	.962	.136
South*Education*SR	.265***	.075	.099
West*Education*SR	.327***	.077	.122
Husband's education	.004	.049	.002
Husband's education*SR	−.025	.041	−.009
NE*Husband's education	.116	.049	.043
South* Husband's education	.071	.084	.026
West*Husband's education	−.047	.064	−.018
NE*Husband's education*SR	−.133	.068	−.050
South*Husband's education*SR	−.076	.087	−.028
West*Husband's education*SR	.059	.074	.022
Number of observations	140133		
Wald Chi-square	15325.43		
Prob > Chi-square	.000		

Notes: S.E. stands for 'standard error' and M.E. for 'marginal effect.' * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 and Midwest are used as reference categories

Table 9 Regression estimates of married women's labor force participation by region: Model (3) with husband information

Variables	Northeast			Midwest			South			West		
	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.	Coeff.	S.E.	M.E.
SR	.442	.858	.155	3.232***	.606	1.181	−.076	.491	−.021	−1.735***	.577	−.652
Age 25–29	.010	.024	.004	−.039*	.024	−.014	−.060***	.021	−.023	.080***	.027	.030
Age 30–34	−.060***	.023	−.023	−.134***	.023	−.050	−.060***	.020	−.023	−.008	.023	−.003
Age 35–39	−.037*	.022	−.014	−.028	.022	−.010	−.035*	.019	−.013	−.025	.022	−.009
NLI	−6.8E-6***	4.4E-7	−2.6E-6	−8.7E-6***	4.3E-7	−3.2E-6	−8.2E-6***	3.9E-7	−3.1E-6	−6.2E-6***	4.8E-7	−2.4E-6
Time trend	.025***	.002	.009	.026***	.002	.010	.021***	.002	.008	.011***	.002	.004
Young children	−.500***	.014	−.191	−.433***	.013	−.161	−.400***	.012	−.150	−.447***	.012	−.169
Education	.034	.072	.016	.333***	.055	.126	.050	.044	.017	.027	.049	.011
Average wages	3.0E-5***	8.4E-6	1.1E-5	2.6E-5***	8.5E-6	9.4E-6	7.9E-7	8.1E-6	3.2E-7	7.8E-5***	9.9E-6	3.0E-5
Education*SR	.066	.075	.023	−.213***	.053	−.081	.058	.047	.023	.080	.052	.029
Husband's education	.134**	.065	.047	.031	.047	.008	.077**	.038	.031	−.020	.044	−.008
Husband's education*SR	−.171**	.068	−.061	−.046	.047	−.013	−.092**	.040	−.037	.009	.046	.004
Observations	33341			39792			48104			37318		
Wald Chi-square	3910.05			5175.73			4582.85			3633.43		
Prob > Chi-square	.000			.000			.000			.000		

Notes: S.E. stands for 'standard error' and M.E. for 'marginal effect.' * Significant at $p > .10$; ** significant at $p > .05$; *** significant at $p > .01$. All regressions include a constant term. Age 40–44 is used as reference categories

5 Discussion and conclusions

Using individual CPS data for selected years spanning 1965–2005, we test the prediction of a negative association between sex ratios and women's LFP for married women as well as for all women regardless of their marital status. For the U.S. as a whole, variation in the sex ratio—the ratio of men 2 years older to women 2 years younger regardless of marital status—is obtained from regional differences and cohort differences. A negative association between sex ratios and women's LFP could be the result of differences in labor market opportunities across regions which, in turn, could lead to migrating patterns that end up changing sex ratios in the aforementioned direction. Therefore, we also carry the analysis by U.S. regions so that sex ratio effects only reflect cohort-level effects.

After accounting for a time trend and other relevant factors, we find that women born during a baby-boom and, therefore, facing a low sex ratio when they enter dating and marriage markets 20 years later, had significantly higher LFP than women born in a baby-bust or during times of constant population growth. Vice-versa, women who were born during a baby-bust and, therefore, experiencing a high sex ratio in marriage markets, had significantly lower LFP than women born in a baby-boom or during times of constant population growth. These results, which hold for all women, regardless of their marital status, as well as for the U.S. as a whole and for most U.S. regions, are based on the assumption that the difference in average age at marriage between men and women stays equal to 2 years over the entire period studied—a reasonable assumption given the actual evolution of that age difference.

Our results help explain why in recent years we have observed some drops in women's LFP, such as a decrease of 3.5% points between 1995 and 2005 in the LFP rates for married women ages 30–34 (see Table 1). During those years, the women who entered these age groups were women of the Moon and Roe generations characterized by high sex ratios. In contrast, when women from cohorts characterized by low sex ratios entered this age group, we observe rapid increases in their LFP. For instance, married women of the Post World War II generation, born in the years 1946–1950, experienced an increase in LFP of 13.5% points when they entered the same age category between 1970 and 1975.

We also predicted that the negative association between sex ratios and women's LFP would be less valid for educated women than for women with low education. Two regions offer evidence for this prediction: the West and the Northeast. In those regions, sex ratio effects on women's LFP are stronger for women who are less educated. In the Midwest, we consistently get the opposite result, i.e. sex ratio effects on women's LFP are stronger for more educated women. Additionally, the analysis by women's marital status reveals that our prediction of a smaller sex ratio effect on women's LFP among more educated women—a prediction that makes sense within the context of an economic analysis of bargaining in marriage, is corroborated for married women.

To the extent that it applies to women with partners, our basic finding of an inverse association between sex ratio and women's LFP is consistent with both D&S and collective models of decision-making in marriage or cohabitation.

According to the collective models, higher sex ratios imply that women will obtain a more advantageous sharing rule when they and their partners (or partners to be) bargain over the distribution of resources that were pooled together. According to the Grossbard's D&S model, higher sex ratios imply that marriage markets have established more advantageous quasi-wages for female household production workers. Our second finding—a positive association between women's LFP and the interaction between female education and sex ratio is interpreted in terms of less educated women being more interested in working in remunerated household production. In the Midwest, we find a negative association between women's LFP and the interaction between female education and sex ratio. This result could be indicative of educated women living in the Midwest encountering a greater opportunity cost for not working.

In future work, it would be of interest to use alternative measures of sex ratio, including measures that only take singles into account—either calibrated to every single-year birth cohort or considering a range of possible age differences as in Goldman, Westoff, and Hammerslough (1984). Likewise, it would be helpful to gain a better understanding of how sex ratios affect other dimensions of labor supply, such as the labor supply of men, and the attachment of men and women to the labor force in an international context.

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Appendix

Table 10 Variable definitions, means and standard deviations

Variables	Definitions	Mean	S.D.
All women			
LFP	Dummy equal to 1 if in the labor force	.68	.47
SR[1]	Sex ratio	.98	.07
Age 25–29	Age group dummy	.21	.41
Age 30–34	Age group dummy	.21	.41
Age 35–39	Age group dummy	.20	.40
Age 40–44	Age group dummy	.19	.39
Young children	Number of children three years old and younger	.27	.54
Education	Years of schooling	12.84	2.70
Average wages	Average female wages for that age/education/region/year cell	7714.44	2584.04
NLI	Household non-labor income (including public assistance)	20409.33	−20801.74
NE	Region dummy	.22	.42
Midwest	Region dummy	.24	.43
South	Region dummy	.30	.46

Table 10 continued

Variables	Definitions	Mean	S.D.
West	Region dummy	.24	.42
Trend	Time trend	21.91	11.91
Married women			
LFP	Dummy equal to 1 if in the labor force	.63	.48
SR	Sex ratio	.97	.07
Age 25–29	Age group dummy	.20	.40
Age 30–34	Age group dummy	.23	.42
Age 35–39	Age group dummy	.23	.42
Age 40–44	Age group dummy	.22	.42
Young children	Number of children three years old and younger	.33	.58
Education	Years of schooling	12.82	2.68
Husband's education	Years of schooling	12.93	3.08
Average wages	Average female wages for that age/education/region/year cell	7781.38	2634.59
NLI	Household non-labor income (including public assistance)	25749.14	20302.99
NE	Region dummy	.21	.41
Midwest	Region dummy	.25	.43
South	Region dummy	.30	.46
West	Region dummy	.24	.43
Trend [2]	Time trend	21.16	12.31

Sources: March CPS 1965–2000. [1] U.S. Bureau of the Census. See the notes at the bottom of Table 1. [2] The time trend takes the following values: 0 for 1965, 5 for 1970, 10 for 1975, and so on until 40 for 2005

Table 11 Variable definitions, means and standard deviations by region

Variables	Northeast				Midwest			
	Married		All		Married		All	
	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.
LFP	.66	.47	.60	.49	.69	.46	.63	.48
SR[1]	.96	.05	.96	.05	1.03	.07	1.03	.07
Age 25–29	.20	.40	.19	.39	.21	.41	.20	.40
Age 30–34	.21	.41	.24	.43	.21	.41	.24	.43
Age 35–39	.21	.40	.24	.43	.20	.40	.23	.42
Age 40–44	.20	.40	.24	.43	.19	.39	.22	.41
Young children	.25	.52	.33	.58	.28	.55	.33	.59
Education	13.01	2.71	13.03	2.66	12.95	2.42	12.91	2.39
Husband's education	–	–	13.16	2.99	–	–	13.04	2.79
Average wages	7965.09	2716.98	8039.81	2805.39	7643.42	2749.32	7663.47	2842.56

Table 11 continued

Variables	Northeast				Midwest			
	Married		All		Married		All	
	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.
NLI	21668.23	2133.90	27109.43	20229.05	21369.21	20641.06	26258.41	19793.04
NE	1.00	.00	1.00	.00	.00	.00	.00	.00
Midwest	.00	.00	.00	.00	1.00	.00	1.00	.00
South	.00	.00	.00	.00	.00	.00	.00	.00
West	.00	.00	.00	.00	.00	.00	.00	.00
Trend	21.44	11.93	20.63	12.44	21.21	12.18	20.38	12.55

Variables	South				West			
	Married		All		Married		All	
	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.
LFP	.68	.47	.63	.48	.67	.47	.62	.49
SR[1]	.95	.07	.95	.06	.96	.06	.96	.06
Age 25–29	.21	.41	.20	.40	.22	.41	.22	.41
Age 30–34	.21	.41	.23	.42	.22	.41	.23	.42
Age 35–39	.20	.40	.22	.42	.20	.40	.23	.42
Age 40–44	.19	.39	.21	.41	.18	.39	.21	.40
Young children	.26	.53	.30	.56	.30	.57	.37	.61
Education	12.63	2.79	12.58	2.77	12.85	2.82	12.81	2.86
Husband's education	–	–	12.58	3.27	–	–	13.06	3.15
Average wages	7679.10	2679.83	7703.38	2715.46	7595.84	2095.45	7756.92	2090.30
NLI	18851.96	20167.25	24008.70	20116.01	20252.88	21127.08	25976.47	20820.87
NE	.00	.00	.00	.00	.00	.00	.00	.00
Midwest	.00	.00	.00	.00	.00	.00	.00	.00
South	1.00	.00	1.00	.00	.00	.00	.00	.00
West	.00	.00	.00	.00	1.00	.00	1.00	.00
Trend [2]	21.92	11.96	21.01	12.28	23.03	11.48	22.67	11.85

Sources: March CPS 1965–2000. [1] U.S. Bureau of the Census. See the notes at the bottom of Table 1 . [2] The time trend takes the following values: 0 for 1965, 5 for 1970, 10 for 1975, and so on until 40 for 2005

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