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THE CHANGING IMPACT OF MARRIAGE AND CHILDREN ON WOMEN'S LABOR FORCE PARTICIPATION

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THE CHANGING IMPACT OF MARRIAGE AND CHILDREN ON WOMEN'S LABOR FORCE PARTICIPATION

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Abstract: Cohany and Sok recently reported that the labor force participation rate of married women with children, and especially married women with very young children, declined between 1997 and 2005. In contrast, recent work by Boushey showed that the negative impact of children on work by women age 25-44 declined, rather than increased, in the two decades between 1984 and 2004. In this paper, I re-examine the effects of marriage and children on women's labor force participation rates between 1984 and 2004. I show that the presence of children has had a declining negative impact on work for single women, but an increasing negative impact for married women. Both of these changes occurred primarily in the 1993-2000 period and have been maintained through 2004, but not at the 1993-2000 rate of increase. Research that overlooks this differential effect of children by marital status draws an incorrect conclusion about the child penalty on women's labor force participation.

I. Introduction

Bureau of Labor Statistics tabulations show that the very steady increase in U.S. women's labor force participation that characterized the post-war period has largely subsided. For most groups of women (all women, married women, and women with children), the trend line in the labor force participation rate flattened out in the early-to-mid 1990s after nearly four decades of steady increases (Cohany and Sok; Mosisa and Hipple). But like many aggregate trends, substantial complexity and controversy lies just beneath the surface. Recent work by Boushey (2004) and by Cohany and Sok (2007) suggest two apparently inconsistent trends. Boushey, responding to anecdotal evidence in the popular press about a declining commitment to work of women with children, showed that the negative impact of children on work by women age 25-44 declined, rather than increased, in the two decades between 1984 and 2004. Cohany and Sok showed that the labor force participation rate of married women with children, and especially married women with very young children, declined between 1997 and 2005, which implies that the negative impact of children on work has increased, at least for this group of women.

Who is right? Actually, they both are. An analysis of data from the CPS-ORG samples from 1984 through 2004 shows that women age 25-44 with children are more likely to be working, not less likely, than in the past. But married women with children are indeed working less than a decade ago, although more than two decades ago. This is especially true for married women with young children. The difference between the findings is attributable to the behavior of single women with children, whose labor force participation jumped sharply in the 1990s. The labor force participation rate of single mothers age 25-44 increased nine percentage points between 1993 and 2000. The rate for single women, age 25-44 with children age five or less, jumped an astonishing 14 percentage points over that time period. In contrast, the labor

force participation rate for married women with children increased one percentage point and the rate for married women with children age five or less was flat over that same time period. More interestingly, the negative impact of children on the labor force participation of married women increased.

In this paper, I examine the changing impact of marriage and children on women's labor force participation between 1984 and 2004. I follow the general approach of Boushey in estimating a multivariate model using logit, but focus more carefully on interactions of marriage and children, an impact not revealed in Boushey's analysis. I also look more carefully at race and age of child effects. Data are from the CPS Outgoing Rotation Group samples for selected years from 1984 through 2004.

II. Background

The steady upward trend in the labor force activity of married women and of women with children in the post-war period is well-known. The labor force participation rate (LFPR) of married women, age 16 and older, rose from just 21% in 1950 to more than 60% in 1994, about where it now stands. The LFPR of married mothers followed a similar trend, rising from just 17% in 1948 to 70% in the mid-1990s (Cohany and Sok).[1] For all women with children two years old or younger, the rate increased from 34% in 1975 to 59% in 2005 (Mosisa and Hipple, Table 10). For all of these groups, the labor force participation rate rose quite steadily through the mid-1990s, but has been essentially unchanged since. For some groups, the LFPR peaked in 1997 and has fallen since.

The *New York Times* (Belkin, 2003 and Story, 2005) and *Time Magazine* (Wallis, 2004) both featured stories about what appeared to be a declining commitment to work among women with children, especially among married women and more educated mothers with young children. This evidence was almost exclusively anecdotal, but it clearly touched a nerve. "Opting out" became a catchphrase. It was suggested that the long march of married women

into the labor force was arguably nearing its end. “Off-ramps” and “on-ramps” have become a part of the jargon of discussing women’s labor force participation and the cycling that still characterizes life-cycle work patterns for many women.

Boushey responded to these accounts by examining data from the CPS Outgoing Rotation Groups (ORG) for selected years from 1984 to 2004.[2] She used a multivariate analysis to focus on the independent impact of children on the probability of labor force participation. Explanatory variables are primarily demographic, rather than economic: presence of a child, marital status, race/ethnicity, presence of a prime-age working male in the household, educational attainment, and year (to control for business cycle impacts). Her sample is limited to women age 25 to 44, but includes women of all marital statuses. Analysis is by logit. The key variables are interactions of the presence of a child with year, which measures what she calls the “child penalty.”

Boushey finds that the LFP penalty of having a child under age 18 declines from 20.7 percentage points in 1984 to 14.4 points in 1993 and further narrows to 9.9 points (2000) and 8.2 points (2004). The corresponding child penalties associated with having a child less than age six are 25.5 points (1984), 22.6 points (1993) and 21.1 and 19.7 points in 2000 and 2004, respectively. Both analyses thus show a narrowing difference in LFP between mothers and non-mothers. Having children has, according to these analyses, become less of a factor in LFP, not more of a factor.

In contrast, Cohany and Sok document falling LFP by married mothers with young children, especially those with infants (children up to one year of age). 1997 appears to be the peak year for LFP for these groups. The LFPR for married mothers with children under age six fell from approximately 64% in 1997 to less than 60% in 2004, before rising slightly in 2005. The LFPR for married mothers with infants fell from 59.2% in 1997 to 51.7% in 2004, then rose to 53.5% in 2005 and 55% in 2006 (BLS, 2007).

Cohany and Sok's analysis is exclusively bivariate. They do show, however, that the downward trend from 1997 to 2004 holds for women 16-24 and 25-34, but not for older women; for non-Hispanic Blacks and for Hispanics more than for non-Hispanic Whites; for native born and foreign born women; and for women with all levels of education. None of these effects control for other variables.

One obvious complication in comparing these results is that the samples clearly differ – mothers age 25-44 and of any marital status with any children or with young children vs married mothers of all ages with very young children. Time frames differ as well. Additionally, Boushey's analysis is multivariate, while Cohany and Sok's is bivariate. And, more importantly, neither examines subtler interaction effects of marital status and children.

III. Data and Methods

I use data from the CPS Outgoing Rotation Group (CPS-ORG) samples for 1984, 1989, 1993, 2000, and 2004 – the same years used by Boushey.[3] The Outgoing Rotation Groups are a portion of the CPS monthly survey that are exiting the sample after either their initial four months or, following an eight-month absence from the sample, their final four months. Sample sizes are very large. In any month, one-fourth of the CPS sample is a member of one of the ORG samples. The annual CPS-ORG data files include all 12 months of ORG interviews, so the weighted total cumulates to three times the total population.

I sample all women age 25-44. For 1984, 1989, and 1993, sample sizes are approximately 70,000. For 2000 and 2004, sample sizes are 56,000 and 59,000, respectively. LFPR estimates from this data differ very slightly from official BLS reports, since the BLS analyses are based on the full CPS sample each month. For 2004, the BLS reports a LFPR for all women age 16 and older of 59.2% (U.S. Dept. of Labor, 2007); the corresponding CPS-ORG estimate is 59.1%. For women with children age 18 or younger, the corresponding BLS and CPS-ORG LFPRS are 70.7 and 70.2%, respectively. Similar comparisons by sample exist for

LFPRs by age of youngest child. These comparisons certainly suggest that the CPS-ORG panels are appropriate for studying trends in women's LFP.

I estimate a set of descriptive regressions of women's labor force participation, using both ordinary least squares (OLS) and logit. The OLS regressions are very easy to interpret: the estimated coefficients are simply the average effect of a particular variable on the labor force participation rate. The weakness of OLS is that resulting probabilities of participation can be less than 0 or greater than 1, something that is not possible. As a result, economists often use logit and probit analysis for variables like labor force participation; both methods appropriately constrain the impacts to be between 0 and 1. I use logit, which is generally easier to work with than probit. Logit coefficients do not, however, have a direct interpretation in terms of their impact on the labor force participation rate. Logit coefficients must be transformed into more interpretable probability effects.[4]

Explanatory variables include marital status, presence of children of various ages, year dummies, educational attainment, race/ethnicity, and age, all entered as dichotomous variables. I examine the impact of age of children in three specifications: any children less than age 18, less than age six, or less than age two.[5] The analyses of the impact of children less than age two are limited to 1989-2004, because this information is not available in the CPS-ORG file for 1984. I allow the impacts of marital status and presence of children of various ages to vary across the years to test for changing impacts. In addition, I test specifically for whether the child penalty varies across marital status.

IV. Analysis

Table 1 presents information on the characteristics of the CPS-ORG sample of women age 25-44. The figures shown are the means over all years (1984, 1989, 1993, 2000, and 2004), except for the presence of a child age 0-2 or 3-5 for which no 1984 data is available. All means are weighted and represent population estimates. Average age of these women is 34.4,

almost two-thirds are currently married and a similar proportion have a child age 18 or younger. One in six has a child age two or younger and one in five has a child age three to five.[6] 72 percent are non-Hispanic white, 13 percent are non-Hispanic black, and 11 percent are Hispanic. The average monthly labor force participation rate for these women over these years is 74 percent.

Figure 1 shows the overall trend in the labor force participation rate for all women 25-44 and separately by marital status. The LFPR for all 25-44 year old women (the middle line) rose sharply between 1984 and 1989, from 70.2% to 74.8%. Over the following five years, the rate increased just 0.4%, then rose just a point and a half in the next seven years (through 2000). Between 2000 and 2004, the proportion in the labor force fell by 2.1 percentage points, falling just below its 1989 level. The time trend for 25-44 year old married women essentially follows this same trend from a lower base. The trend for single women, however, is quite different. From a higher base (82.5% in 1984), their labor force participation rate declines steadily through 1993, then increases through 2000, more than making up for the earlier decline. Finally, between 2000 and 2004, their LFP profile declines, following the other two trend lines. These trends suggest a decline in the negative impact of marriage on labor force participation, from a gross (unadjusted) penalty of almost 17 percentage points in 1984 to 8-12 percentage points since 1989. In 2004, the difference was 11 percentage points. These differences do not, however, control for compositional effects.

To some extent, the trends in Figure 1 conceal more than they reveal, since the real story turns on the interaction of marital status and the presence of children and, more specifically, on the change in that interaction over these years. For that, I turn to the regression analysis. Table 2 presents estimates from three OLS regression models and one logit model. Model (1) is similar to Boushey's model. It includes basic demographic information (race, education, and age, all entered as dummy variables), plus year dummies[7] and whether there is a child age 18 or younger in the household. The effect of a child on labor force participation is

allowed to vary by year; the coefficients in the table show the changing child penalty relative to 1984. Model (2) adds a variable for marital status; this provides another measure of the child penalty, this time controlling for marital status. Model (3) adds a variable combining marital status and presence of children. This enables me to examine whether the labor force participation of married women with children is changing over time relative to single women with children. Finally, Model (4) is a logit version of the specification used in Model (3). In this table, I focus on the impact of having a child age 18 or under. In Table 3, I examine the impact of younger children and also examine possible differences in responses by race and ethnicity.

Because the sample size is so large, almost all coefficients are statistically significant at the 10% level or better. Most are statistically significant at the 1% level. Estimates that are not statistically significant at the 10% level or better are shaded in the table.

Model (1) shows a very straightforward story about the impact of children on women's LFP. In 1984, the child penalty on LFP is 18.3 percentage points. The coefficients just below (Child-1989 to Child-2004) show the difference in the child effect in that year, relative to 1984; year-to-year changes can be obtained just by subtraction. The penalty falls in absolute value after 1984, by 3.4 percentage points by 1989, an additional 2.2 percentage points between 1989 and 1993 (the difference between the two year estimates), and then by 3.3 more percentage points by 2000. Between 2000 and 2004, no further change occurs; the two estimates of the child penalty are essentially unchanged. As of 2004, the child penalty was half its original 1984 level, down from 18 percentage points to nine. This is almost exactly what Boushey finds and thus concludes that the impact of children on labor force participation is falling. As far as she goes, she is entirely correct.

Model (2) adds control for marital status, interacted with year. The control slightly weakens the impact of children on participation, but the central story still holds. In this specification, the original negative impact of children is 14.4 percentage points and most of the change occurs between 1993 and 2000, rather than more steadily between 1989 and 2000.

The trend in the effect of marriage on labor force participation follows the child effect trend to some extent, but the timing differs. In 1984 (the base year) the LFP of married women was 11 percentage points lower, all else constant. This fell almost in half by 1989 and then fell further by 1993. But then the marital impact reverses course: between 1993 and 2000 and continuing into 2004, the negative impact of marriage on labor force participation increased. By 2004, the impact of marriage was nearly as large as it had been in 1984, -.086 vs -.110.

Models (3) and (4), OLS and logit respectively, add the marriage x child effect, allowed to vary by year. In this specification, the child coefficients are the impacts for single women, while the marriage x child variable measures the differential impact of children on the labor force participation of married women relative to single women. The marriage variable estimates are the impacts for married women without children. Using this model, it is possible to combine coefficients to compare the labor force participation of single women with children relative to single women without children, married women relative to single women, and married women with children relative to married women without children. For example, the effect of children on labor force participation for a married woman is the sum of three effects: the child effect, the marriage effect, and the child-marriage effect.

The OLS results in column (3) reveal entirely different trends for single and married women with children. In 1984, single women with children had a labor force participation rate 8.3 percentage points lower than single women without children (see the child < Age 18 coefficient), holding the other demographic factors in the model constant. The corresponding LFP rate for married women with children in that year was another 15.5 percentage points lower, the sum of the marriage estimate (-.065) and married x child effect (-.090). This is consistent with the estimated marriage coefficient of -0.11 in Model (2), which is roughly a weighted average of the marriage effect for women with children (-.155) and those without children (-.065).

Through 1993, the effect of children on the labor force participation rate of single women was essentially unchanged; the 1989 and 1993 child interactions are very small (-0.010 and -0.013) and not statistically significant. Over this same time period, however, the negative impact of children on the labor force participation rate for married women declined by two-thirds from 15.5 points to 5.5 points (based on the sum of the marriage and marriage x child interactions). By 1993, marriage had essentially no effect (-.008) on the labor force participation of women without children; this is shown by the difference between the marriage effect in 1984 (-0.065) and the change in the effect in 1993 (0.057). Then the trends changed course again. The labor force participation rate for single women with children jumped sharply (see the child-2000 and child-2004 coefficient estimates of 0.085 and 0.087), to the extent that by 2000 and through 2004, children no longer have a net marginal negative effect on work for single women. But married women did not follow that trend; for them, the child effect remained steady through 2000 and 2004.[8] This confirms that post-1993, the declining child penalty observed in Models 1 and 2 reflects the impact of single women with children.

The logit estimates show an identical trend. As already noted, the logit coefficients in column 4 do not have a direct quantitative interpretation in terms of the probability of labor force participation, although sign and statistical significance can be readily assessed. The implied logit child estimates are shown in Figure 2, separately for single and married women. The figures shown are marital status-specific, i.e., they are relative to childless women of the same marital status. The different patterns are apparent. Through 1993, the child impacts are essentially constant, not as negative for single women (-11 percentage points) as for married women (-14 to -16 percentage points). Thereafter, the trends diverge, with the negative impact of children steady for married women and becoming less negative for single women. By 2000, the child effect is essentially zero for single women and 12 to 13 percentage points for married women. The net change in relative position from the 1980s and early 1990s to the 2000s is almost 10 percentage points.

Between 2000 and 2004, the labor force participation rate fell for single and married women, with and without children. But this decline is very similar for all of the groups; none of the 2004 marriage or children effects are statistically different from those in 2000.

The other variables in the regressions have very reasonable impacts that are consistent with other estimates of their effects. Controlling for marriage and children (Model 3 in Table 2), black and white women are both about six percentage points more likely to be participants than Hispanic women and Asian women (the omitted group). Without control for marriage (Model 1), black women are the most likely to be working, but this reflects their lower rates of marriage. The time dummies show an across-the-board negative effect between 1989 and 1993 and then another two point decline between 2000 and 2004. The impact of education is large. Women who have less than a high school degree have far lower rates of labor force participation –29 percentage points (models 2 and 3). High school graduates with no post-secondary education also have reduced LFP rates (11 percentage points). Logit estimates for these variables are quite similar.

Women with Younger Children. The analysis thus far has examined only the impact of having a child under age 18. Much of the focus, however, has been on women with younger children. In Table 3, I examine the impact on women's labor force participation of having a child less than age six (Column 1) or a child less than age two (Column 2).[9] The specification is the same as used in Table 2, Columns 3. For expositional ease, I use OLS, show only the core variables of interest, and do not include the standard errors. (Logit estimates are virtually identical and are available upon request.) Shaded entries are not statistically significant at the 10% level. The estimates in Column (2) are based only on 1989-2004 data, because information about the presence of very young children is not available earlier. In that model, 1989 is the omitted year and all year interaction effects are relative to that year.

As seen in Column (1), the impact of a child less than age six is very large and negative in 1984. The estimate (-.194) is more than twice as large as the corresponding estimate from

Table 2 (-.083). Through 1993, very little changes for single women, and then, exactly as before, the negative child effect diminishes sharply. By 2004, the negative impact is about six percentage points, less than one-third of its 1984 level. For married women with children young than six years, the effect of children on work barely changes over the twenty year period. In 1984, children reduced the labor force participation rate of married women by more than 20 percentage points (the sum of the child and married x child estimates). This effect diminished by two percentage points through 1993, but the 2004 effect is unchanged from the 1993 estimate. So again, the impact of children on the labor force participation of single women and married women diverged post 1993. In 1984, single women with young children had a labor force participation rate 11.6 points higher than married women. By 2004, this difference had increased by five percentage points.

The impact of very young children (model 2 of table 3) also follows the patterns seen, but with a more pronounced impact--as might be expected. In 1989 and 1993, a young child reduced the labor force participation of single women by about 24 percentage points. By 2000 and still in 2004, this effect attenuated, falling to less than half its previous value. For married women with very young children, the trends are similar to those for married mothers with older children but with a stronger post-1993 trend. These mothers increased their labor force participation slightly relative to married women without young children between 1984 and 1993, but thereafter the gap increased. The penalty of very young children for married women increased by three percentage points between 1993 and 2004. The net effect is that the penalty of very young children on the labor participation of married women is at the same level in 2004 as in 1984.

Models 3-5 of table 3 further disaggregate the sample by race and ethnicity, to examine whether the impacts are consistent across the various groups. The presence of a child less than 6 years is the child indicator in all of these analyses. Results for the presence of a child are similar for other ages. Again, only the key variables are shown. The general story here is that

the patterns hold across White, Black, and Hispanic women. For all three groups, a large negative impact of children on the labor force participation of single women persists through 1993 and then is sharply cut or even disappears (in the case of Black women) by 2000. Between 2000 and 2004, the child penalty rises 1-2 points for Whites and Blacks (see the change in the child estimates between those years), while it decreases slightly for Hispanics. For married women, the 1984 impact of children varies by race: the net effect, based on the sum of the married and married x child terms, is positive for Black women, zero for Hispanic women, and negative for White women. All three groups show a growing negative impact of children on participation between 1993 and 2000, extending into 2004.

V. Other Issues

My analysis focuses on women aged 25-44 (the sample range used by Boushey) and thus leaves out both younger and older mothers. In 2004, one-sixth of mothers with children age six or younger were less than age 25 and another 2.8% were over age 44. While women age 25-44 are an interesting and relevant age group, the younger ones may also be of interest. What is the effect of marriage and children on their labor force participation?

Because marriage and fertility are endogenous and are atypical at these younger ages, I treat this issue cautiously. Figure 3 shows the labor force participation rate for women, age 16-24, with a child age six or less. Between 1984 and 1993, the rates are independent of marriage; approximately 50% of single and married women with a young child worked. Then, quite similar to the other analyses, the labor force participation rate for single women jumps, in this case by 19 percentage points between 1993 and 2000. The participation rate for married mothers also increases, by about five points. After 2000, the participation rate for both groups declined six to eight percentage points. This suggests that including these younger women in the analysis would not alter any of the conclusions drawn.

VI. Summary

The basic story revealed by the data on women's labor force participation between 1984 and 2004 is one in which the presence of children has had a smaller negative impact on work for all women age 25 to 44, a finding that confirms Boushey's report of a declining child penalty. But on closer inspection, this effect varies enormously by marital status. Single women with children sharply increased their labor force participation rate, while the declining impact of children on the labor force participation rate of married women stalled, beginning in 1993. Both of these changes occurred primarily in the 1993-2000 period and have been maintained through 2004, but not at the 1993- 2000 rate of increase. These child impacts are not much changed, whether I examine women with children under age 18, women with children less than age six, or women with children less than age two. The effects are also widespread across race and ethnicity. The negative impact of a child less than age six on the labor force participation of single black women literally disappears between 1984 and 2000. The key contribution of these analyses is to emphasize that focusing only on the effect of children on labor force participation provides a very incomplete picture of the very different effects on single and married women.

VII. Discussion

A full explanation of the changes documented here is a formidable and important challenge. At this point, I can identify potential candidates, but not assess them fully. The timing of the changes for single women tracks reasonably well with both welfare reform, including the pre-PRWORA state waivers, and the substantial increase in the generosity of the EITC. Between 1993 and 1996, 46 states received waivers for AFDC and Medicaid, including 33 that generally required work, set time limits for assistance, or increased work incentives (US GAO, 1997). Figure 4 shows the percentage point increase in the labor force participation rate by state for single mothers age 25-44 between 1993 and 2000; each bar represents a state, arrayed from biggest to smallest increase. Substantial variation across state is evident, which is

itself interesting and worth further consideration. The average increase is 9.9 percentage points and the median increase is 10.2 points. Seven states had decreases and another four had increases less than five percentage points. The largest increases are in Connecticut, Minnesota, Louisiana, and Massachusetts. All four states had waiver programs in place, but that is not by itself sufficient evidence of a causal impact.

A simple difference-in-difference (DID) calculation of changes in labor force participation rates for married and single women with children can crudely net out common within-state effects that are due to economic growth or other state-wide factors.[10] The range of DID estimates (single – married) is from 32.5 percentage points in Connecticut, where the LFPR for married mothers declined while the rate for single mothers increased sharply, to -6.3 percentage points in Kansas, where the rate for married mothers increased and the rate for single mothers fell. The top five states (CT, MN, IN, MA, and LA) all had waivers in place. Connecticut, Minnesota, and Indiana are particularly interesting in this computation, because their state LFPR for single women with children increased sharply, while the rate for married women with children fell.

Over the same time period, maximum EITC benefits more than doubled for women with two or more children and increased 50 percent for women with one children. For single women who are not in the labor force, the EITC labor supply incentives are unambiguously positive. The EITC acts as a wage subsidy, up to some earnings threshold, equal to 34% for women with one child and 40% for women with two children (Hoffman and Seidman, 2003). For married women, conflicting income and substitution effects may actually generate negative work incentives if family income, net of their own potential contribution, places them on the declining benefit portion of the EITC schedule; for evidence on this, see Eissa and Hoynes (2005).

Changes in fertility rates are a potential, although obviously endogenous, contributing factor for married women. Fertility rates rose for married women, especially older married women. Between 1993 and 2004, the fertility rate for married women, age 20-24, declined

3.3%, while the corresponding rates for 30-34 year olds and 35-39 year olds increased 20% and 44%, respectively.[11] More traditional economic analyses look to income effects via spouse incomes. That information is not available on the CPS-ORG file. It is also possible that the changes reflect different approach to the production of child services, with a substitution of own time for non-family caregiver time. These issues can more fruitfully be explored with data sets like the NLSY that combine detailed family income and employment information with employment, marriage, and fertility histories.

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Table 1. Sample Characteristics (Weighted), CPS-ORG, Women, Age 25-44, 1984, 1989, 1993, 200, and 2004

Variable	Mean	Std.Dev.
Age	34.41	5.681
Married	0.657	0.475
Has a Child Age 0-2[a]	0.169	0.375
Has a Child Age 3-5 [a]	0.202	0.402
Has a Child Age 6-13	0.412	0.492
Has a Child Age 14-17	0.193	0.395
Has a Child < Age 18	0.644	0.479
LFPR	0.744	0.436
White (Non-Hispanic)	0.718	0.450
Black (Non-Hispanic)	0.130	0.336
Hispanic	0.105	0.307
Educ < HS	0.113	0.317
Educ=HS Grad	0.350	0.477
Educ-Some College	0.275	0.447
Coll Grad	0.190	0.392
Adv Deg	0.071	0.257
Sample Size	326,624	

[a] Not available in 1984.

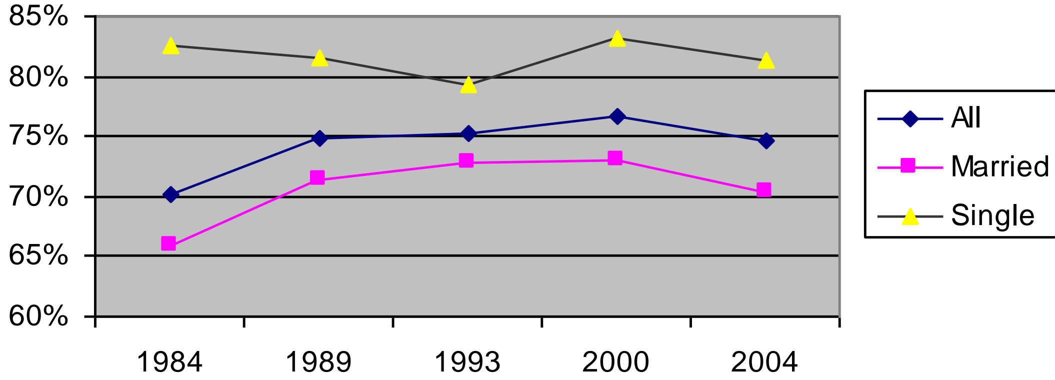
Fig 1. LFP Rate, Women, Age 25-44, Selected Years, 1984-2004, By Marital Status
85%
80%
75%
70%
65%
60%
1984
1989
1993
2000
2004
All
Married
Single

Table 2 OLS and Logit Estimates of Effect of Children and Marriage on LFP of Women, 25-44, Selected Years, 1984-2004

Variable	(1)		(2)		(3)		Logit	
	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.
Constant	0.891	0.005	0.952	0.005	0.931	0.006	2.388	0.039
Child-1984	-0.183	0.003	-0.144	0.004	-0.083	0.006	-0.635	0.041
Child-1989	0.034	0.005	0.016	0.005	-0.010	0.009	-0.034	0.054
Child-1993	0.056	0.005	0.026	0.005	-0.013	0.008	0.035	0.052
Child-2000	0.089	0.005	0.079	0.005	0.085	0.009	0.612	0.056
Child-2004	0.090	0.005	0.080	0.005	0.087	0.009	0.644	0.054
Year=1989	0.012	0.004	-0.016	0.005	-0.007	0.005	-0.057	0.038
Year=1993	-0.006	0.004	-0.049	0.005	-0.034	0.005	-0.286	0.037
Year=2000	-0.012	0.004	-0.034	0.005	-0.038	0.006	-0.309	0.039
Year=2004	-0.031	0.004	-0.051	0.005	-0.057	0.005	-0.444	0.037
Black	0.082	0.004	0.064	0.004	0.059	0.004	0.317	0.022
White	0.057	0.003	0.058	0.003	0.059	0.003	0.323	0.018
Hispanic	0.018	0.004	0.015	0.004	0.015	0.004	0.111	0.022
Educ < HS	-0.280	0.004	-0.286	0.004	-0.290	0.004	-1.530	0.022
Educ = HS	-0.101	0.003	-0.102	0.003	-0.106	0.003	-0.688	0.020
Some College	-0.052	0.003	-0.054	0.003	-0.057	0.003	-0.416	0.021
Adv Degr	-0.035	0.003	-0.033	0.003	-0.034	0.003	-0.268	0.021
Age 25-32	-0.036	0.002	-0.042	0.002	-0.041	0.002	-0.219	0.011
Age 33-39	-0.010	0.002	-0.014	0.002	-0.012	0.002	-0.065	0.011
Married-1984			-0.110	0.004	-0.065	0.006	-0.516	0.038
Married-1989			0.050	0.005	0.034	0.008	0.265	0.053
Married-1993			0.083	0.005	0.057	0.008	0.465	0.051
Married-2000			0.028	0.006	0.049	0.008	0.408	0.055
Married-2004			0.024	0.005	0.048	0.008	0.405	0.053
Marr x Child-1984					-0.090	0.008	-0.256	0.048
Marr x Child-1989					0.034	0.011	0.051	0.067
Marr x Child-1993					0.053	0.011	0.065	0.065
Marr x Child-2000					-0.029	0.011	-0.414	0.070
Marr x Child-2004					-0.032	0.011	-0.412	0.068
R^2 (adj.)	.063		.068		.070			

Models (1)–(3) are estimated by OLS. Presence of child refers to children age 18 and younger.
Note: Sample size for all models is 326,664
Shaded cell = not statistically significant at 10% level or less.

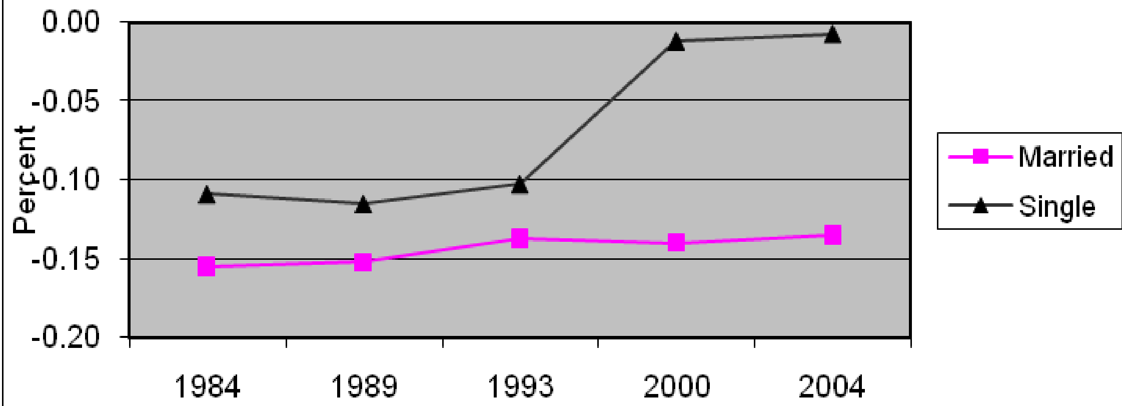
Fig 2. Logit Estimates of Impact of Children, Age 18 and Under, on LFP of Women Age 25-44, by Marital Status, Selected Years, 1984-2004
Percent
0.00
-0.05
-0.10
-0.15
-0.20
1984
1989
1993
2000
2004
Married
Single

Table 3 Effect of Children and Marriage on LFP of Women, 25-44, Selected Years, 1984-2004, By Age of Child, Race, and Ethnicity, OLS Estimates

Variable	(1) Child Age 0-5	(2) Child Age 0-2	(3) White, Child Age 0-5	(4) Black, Child Age 0-5	(5) Hispanic, Child Age 0-5
Constant	0.897	0.895	0.965	0.924	0.878
Child-1984	-0.194	xx	-0.162	-0.150	-0.297
Child-1989	0.005	0.240	-0.005	-0.018	0.079
Child-1993	0.010	0.004	-0.007	0.014	0.080
Child-2000	0.144	0.146	0.104	0.152	0.212
Child-2004	0.136	0.144	0.090	0.130	0.225
Year=1989	-0.004	xx	-0.001	0.000	-0.014
Year=1993	-0.030	-0.026	-0.019	-0.048	-0.047
Year=2000	-0.015	-0.002	-0.019	-0.002	0.007
Year=2004	-0.031	-0.019	-0.039	-0.017	0.011
Married-1984	-0.104	xx	-0.119	-0.003	-0.114
Married-1989	0.050	-0.077	0.048	0.040	0.052
Married-1993	0.079	0.029	0.076	0.067	0.071
Married-2000	0.067	0.008	0.078	0.009	0.058
Married-2004	0.064	0.007	0.079	0.014	0.028
Marr x Child-1984	-0.012	xx	-0.063	0.065	0.118
Marr x Child-1989	0.003	0.041	0.017	0.011	-0.041
Marr x Child-1993	0.016	0.022	0.037	-0.021	-0.044
Marr x Child-2000	-0.122	-0.131	-0.081	-0.138	-0.218
Marr x Child-2004	-0.111	-0.129	-0.055	-0.144	-0.234
Sample size	326,664	255,979	245,517	36,255	28,255
R^2 (adj.)	.088	.075	.088	.085	.102

Note: In model 2, all year interactions are relative to 1989.
Shaded = not statistically significant at 10% level or less.

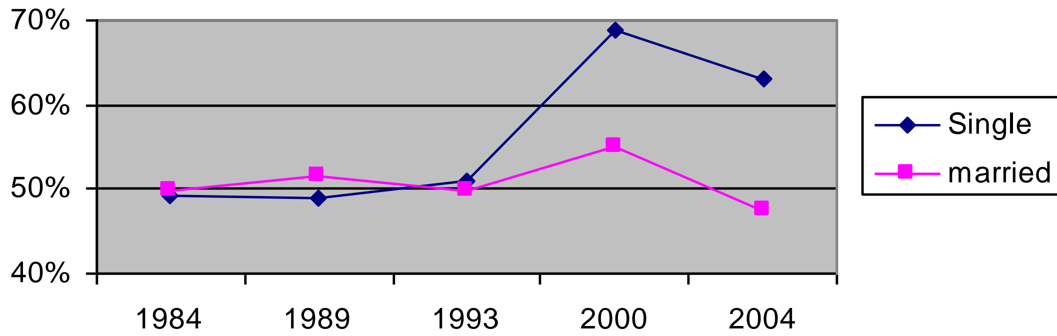
Fig 3. LFP Rate, Women, Age 16-24 with Young Children, by Marital Status, Selected Years, 1984-2004
70%
60%
50%
40%
1984
1989
1993
2000
2004
Single
married

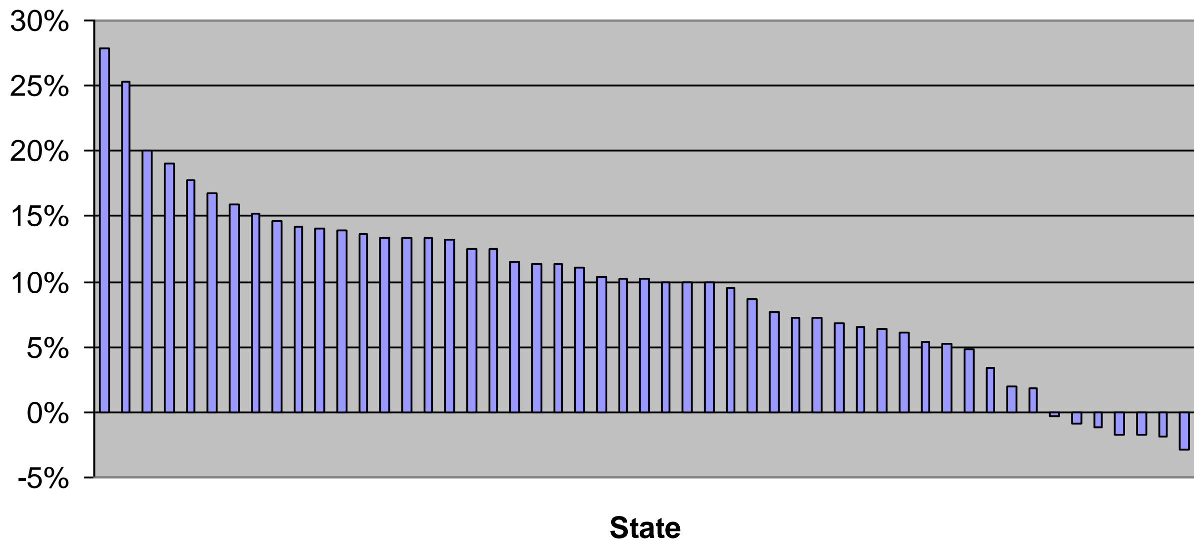
Fig. 4. Percentage Point Change in LFPR, 1993-2000, Single Women with Children, by State
30%
25%
20%
15%
10%
5%
0%
-5%
State

APPENDIX

This table shows the effect of converting missing data on the presence of children in various age groups to zero in the CPS-ORG data. Without the conversion, the distribution of women by age of children is widely different from BLS tabulations; compare the first and third columns. With the conversion, the distributions are nearly identical.

Table A1. Percentage Distribution of Women by Age of Children, 2004

	BLS Estimates[a]	CPS-ORG Estimates	
	% of All	% of All After Conversion of Missing Data to Zeros	% of All with No Conversion of Missing Data
With Children under 18 years old	31.8%	31.6%	48.7%
With Children 6 to 17, none younger	17.7%	17.6%	27.1%
With Children under 6 years	14.1%	14.0%	21.6%
With Children under 2 years	8.2%	8.1%	12.4%
With no children under 18 years old	68.2%	68.4%	51.3%

[a] BLS Estimates from *Women in the Labor Force: A Databook*, Table 6, Report 985, May, 2005.

[1] The LFPR for this group is higher than for all married women, because mothers of children age 18 or less are younger than the population of all married women, a group with no upper age limit.

[2] The years included in the analysis are dictated by the unavailability of information on presence of children by age in the CPS-ORG data between 1993 and 1999.

[3] Data files were obtained from the CEPR data archive at http://www.ceprdata.org/cps/org_index.php .

[4] The logit probability is exp (XB)/[1+exp(XB)], where the B's are the estimated coefficients. The marginal effect in a logit model is B x P x (1-P), where P is the mean sample proportion.

[5] I cannot replicate the sample analyzed by Cohany and Sok, because it is not possible to identify mothers of infants (age 0-1) in the CPS-ORG data. This information is not available on any publicly-released CPS data set (Sok, personal correspondence).

[6] These proportions are based on information on the presence of a child in given age ranges. Thirty five percent of the observations have missing data for all child age variables. It is clear that the missing data are actually substantive 0's. With this conversion, I exactly replicate BLS distributions of women by age of child (BLS, Table 6, 2006); without it, the distributions are widely different. This information is shown in Appendix Table 1. It appears that some skip sequence triggered the missing data, but the details of this are not obvious in the CPS-ORG data.

[7] Boushey interprets year dummies as business cycle variables. In modeling women's labor force participation, it is problematic to interpret time trends or year dummies as solely business cycle effects. She finds that control for year has a large effect on the estimated impact of having a child on labor force participation.

[8] This calculation reflects the changing estimates of the effects of children, marriage, and marriage x children between 1993 and 2004.

[9] Just under half of the women with children age 18 or younger have a child younger than six.

[10] The DID calculation is ($LFP_{M,2000}$ -$LFP_{M,1993}$) - ($LFP_{S,2000}$ –$LFP_{S,1993}$).

[11] The underlying fertility rates are 205.2 and 198.4 (age 20-24), 98.5 and 118.0 (age 30-34), and 37.8 and 54.5 (age 35-39) (US Vital Statistics, 2002; Martin et al 2005, 2006). Over the entire 1984-2004 period, the fertility rate for 30-34 year old married women increased 43% and the rate for 35-39 year olds increased 107%.