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Abstract

This paper considers the role of the entire marital history in labor market decisions. A distinction is made between married, remarried, single and divorced women in the estimation of standard participation and labor supply functions. In specifications controlling for unobserved individual heterogeneity, white remarried women are more likely to participate in the labor force and have higher labor supply than that of white married women. The results indicate that a substantial fraction of the total change in employment rates of all married women over time is due to the increase in the number of remarried women in the population.

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1. Introduction

The United States has witnessed dramatic changes in family structure and labor supply patterns over the last several decades. Since the 1960s, the U.S. labor market has been characterized by a striking rise in the labor force participation rates of women, especially married women. This increase in labor force participation rates has occurred alongside climbing divorce rates, raising interesting questions regarding the nature of the relationship between labor supply and marriage choices. According to economic theory, there are several reasons to suspect the decisions to form and dissolve a household are related to the labor supply decision. Many factors that influence the opportunity cost of market work, such as children and spousal income, tend to be associated with marriage. It is likely that unobserved attributes play a similar role: for example, women with unobserved preferences for "home work" relative to market work may be more likely to marry, as they may gain more from specialization of labor within the household. Alternatively, it may be the case that traits desired in the marriage market are also valued in the labor market (Cornwell and Rupert, 1997).

Labor supply is predicted to increase in response to divorce due to the large decline in non-labor income that tends to follow. For this reason, women may have incentives to accumulate experience and to increase their wages in anticipation of a future divorce (Johnson and Skinner, 1986). Noting a substantive increase in hours worked in the years prior and subsequent to divorce, Johnson and Skinner (1986) estimate a model where divorce probabilities and labor supply are simultaneously determined for a sample of initially intact families. The authors find labor supply increases in response to an increase in the probability of divorce, although labor force participation does not have a significant effect on divorce probabilities. In subsequent work, Johnson and Skinner (1988) attempt to determine the source of the large rise in participation rates for a sample of women who experienced a marital separation. Several hypotheses are considered, including declines in spousal income, higher after-tax wages and reductions in labor specialization within the home. After controlling for unobserved individual heterogeneity, changes in family income account for a only small fraction of the change in labor supply, while own wages and husband's home hours had no significant effects on labor supply. Interestingly, much of the change in labor supply is due to other characteristics related to the separation, as the marital status indicator accounted for most of the predicted change in hours.

Previous work therefore suggests divorce may have sizable effects on labor supply behavior. A natural question of interest is whether the effects of divorce on labor supply persist upon remarriage. This question is of importance as large numbers of divorced women remarry. Since remarriage likely results in a restoration of non-labor income and specialization within the household (Duncan and Hoffman, 1985),[1] the long-term effects of divorce on labor supply may be negligible. Alternatively, it may be the case that past divorce decisions continue to influence the current labor supply decisions of remarried women. Women who increase their labor supply as a consequence of divorce are likely to have higher wages as a result of increased labor market experience. In addition, since the probability of divorce tends to rise with the order of the marriage,[2] remarried women face higher divorce probabilities than women in their first marriages and, holding experience and other standard labor supply determinants constant, may therefore have higher labor supply. If so, part of the rise in the labor force participation rates of all currently married women since the 1960s may simply arise from changes in the composition of the sample over time. In particular, as divorce rates rise, the sample of currently married

[1] This issue is raised but not addressed by Johnson and Skinner (1988).

[2] Becker, Landes and Michael (1977) suggest remarried individuals have less marital-specific capital and larger variances in expected outcomes than married individuals and hence have higher divorce probabilities. They also suggest marital-specific capital from a first marriage may destabilize subsequent marriages.

women will contain a larger fraction of remarried women than in the past. If remarried women are more likely to work than women in their first marriages, the labor force participation rates for all married women will increase as a consequence of the increased number of remarried women in the population.

It may also be the case that the sample of currently single women pools two very heterogeneous groups due to the role of past marital status decisions. For example, divorced women may face higher opportunity costs of working than single women due to an accumulation of marital-specific capital enhancing productivity at home. Furthermore, if unobserved preferences for marriage and work are negatively correlated it is likely divorced women, having made an initial decision to marry, possess lower preferences for work and are less likely to participate in the labor force than never-married women.

The goal of this paper is to determine the extent to which the entire marital history influences current labor supply. Standard labor force participation and labor supply functions are estimated for black and white women separately, where a distinction is made between marriage and remarriage, as well as between single and divorced women. For ease of comparison, the Panel Study of Income Dynamics (PSID) is used to conduct the analysis, as related studies have utilized this data set in the past. The construction of the sample is discussed in section 2.

The specification and estimation of the model is outlined in section 3. The model is estimated under two assumptions. First, it is assumed that unobserved, time-invariant individual effects are uncorrelated with the determinants of the participation and labor supply decisions as is assumed in the majority of past studies. Second, to allow for the possibility that unobserved preferences over work are correlated with unobserved preferences over marriage, exogeneity assumptions are relaxed and the covariates are allowed to be correlated with unobserved individual heterogeneity in the labor supply and participation decisions.[3] Whereas most studies of female labor supply are limited to samples of continuously married women, or in the case of Johnson and Skinner (1986, 1988) to a sample of separated women, the data in this paper include women from all marital states.

The following conclusions can be drawn from the empirical analysis presented in section 4. First, there are no significant differences in participation or labor supply behavior for white women in the uncorrelated specification. However, in the specification controlling for unobserved individual heterogeneity, white remarried women are more likely to participate in the labor force and have higher labor supply than that of white married women. Second, single white women are more likely to participate and supply greater hours than divorced white women in the correlated specification. The results suggest the effects of the marital history are biased in the specification that assumes the covariates are uncorrelated with time-invariant unobserved heterogeneity. One interpretation of this finding, consistent with the literature on the earnings premium for married men, is that some unobserved traits may be attractive in both labor and marriage markets (Cornwell and Rupert, 1997). Surprisingly, the results for black women vary considerably from those for whites: there do not exist significant differences in employment behavior for black women in different marital states.

[3] A recent paper by van der Klaauw (1996) makes an important contribution to this literature by adopting a dynamic structural model where individuals choose their marital and labor force status each period. The results indicate the presence of strong interdependencies between labor supply and marital status decisions and suggest that the own wage effect on labor supply is underestimated when the endogeneity of the marital status choice is ignored. The paper provides evidence that marital status decisions have important consequences for labor supply behavior, motivating further investigation of this issue. However, van der Klaauw (1996) does not consider the effects of the entire marital history on current participation decisions.

2. Data

The data for this analysis come from the PSID. The original sample interviewed by the PSID in 1968 contains approximately 5,000 families and is composed of two separate sub-samples: a nationally representative sample of households and a sample of low-income families.[4] The original PSID sample, as well as any split-off families, is followed up to the present and data are currently available up to 1992.

The variables used in this analysis are defined as follows. Annual hours of work are the product of the number of weeks worked during the year and the average number of hours worked during the weeks the respondent worked. Participation is defined as working 775 hours or more per year. This measure of participation is a better indicator of significant attachment to the labor force than non-zero hours of work over the entire year. In terms of education, women with high school are defined as having exactly 12 years of education, while some college refers to more than 12 years of education. The wage measure is average hourly wages in 1983 dollars[5] and the measure of experience used in this paper is the total number of years the respondent worked since the age of 18. Experience is constructed using the measure of total experience reported in the PSID for all women in 1985 and information on employment status from all previous and subsequent years is used to update the experience measures accordingly.[6]

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[4] To adjust for unequal selection probabilities, sampling weights are used in the descriptive statistics and in estimation.

[5] Wages are constructed by dividing annual (weekly) earnings by 2000 (40). There are several outliers at the both ends of the wage distribution. Women in the 1st and 99th percentiles of the wage distribution are assigned missing values for the wage information but are not dropped from the sample.

[6] Some women are missing experience information in 1985. In these cases, total experience in 1987 and labor force status in previous and subsequent periods are used to construct experience.

The marital history is constructed using information on current and year-to-year changes in marital status, the relationship of the respondent to the head of the household and information on the starting and ending dates of the first marriage. Women are assigned one of four possible marital states: single, divorced, married and remarried. Single women are never-married women, while divorced women are currently single women who were married in the past.[7] Women in first marriages that are currently living with their spouses are defined as married. Remarried women are currently married women, living with their spouses, whose first marriages ended before the interview date. Cohabitors are defined as married or remarried for it is not possible to distinguish between legal marriage and cohabitation in the data.[8] Due to small sample sizes, widows are eliminated from the sample.

The sample used in estimation is restricted to prime working-age women with completed education.[9] Information on hours is missing for many women in 1970 and 1971, annual information on the formation and dissolution of cohabiting couples is not available before 1976, and information on student status for women over 25 is only available after 1978. Therefore, the sample is limited to the years 1979 to 1992. Non-sample members[10] and individuals with missing[11] and inconsistent responses are also eliminated. The current analysis is also limited to black and white women due to the small sample sizes or limited years of data available for

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[7] Women who are separated but not legally divorced are treated as divorced for the purposes of this analysis. With regards to currently single women who cohabited in the past, if the relationship ended before 1979, they are treated as single: if the relationship ended after 1979, they are treated as divorced.

[8] Information pertaining to the respondent's relationship to the head of the household does not distinguish between cohabitors and legal wives as of 1982.

[9] In particular, women who are 25 years of age or older in 1979 and 55 years of age or younger in 1992 are included in the sample.

[10] Non-sample members are defined as individuals who were not part of original PSID families and were not born to original PSID sample members. Further information is not collected for non-sample individuals if they leave their current PSID family. This is especially problematic for the current analysis, as attrition from the sample could result from divorce or separation.

[11] Attempts are made to fill in missing education, age and race information from information in other years where possible.

women of other ethnic backgrounds. Since experience is constructed using information on labor force status in each year, and for econometric purposes discussed in section 3, the sample is further restricted to women with a complete set of information in each year. The resulting balanced panel contains 957 women or 13,398 person-year observations.

Table 1 contains average characteristics of the sample for each year in the panel. It is of interest to examine how the characteristics of the sample evolve over time. First and foremost, participation rates and labor supply increase over time as consistent with the well-documented empirical trends. Participation rates rise from 72% in 1979 to 83% in 1992, while average annual hours for working women increase by 250 hours over the sample period. The upward trends in participation rates and hours occur in conjunction with increases in experience and a reduction in childbearing as the women in the sample age, but also alongside dramatic changes in marital status over the 14 years of the panel. In 1979, approximately two-thirds of the sample are currently in their first marriage. Of the 91% of women who had ever married, 24% have divorced. By 1992, the proportion of ever-married women increases to 95% and the proportion of women who have divorced to 43%. As a result, the proportion of currently married women who are remarried doubles over this period and 81% of currently single women have experienced a divorce by 1992.

Considering the change in the marital histories of women over time, it is of interest to examine the average characteristics of women in different marital states. To avoid confounding differences across individuals with differences across time, statistics for the 1992 cross-section are provided. The wide variation in marriage and employment patterns across race is illustrated in table 2. On average, white women have higher levels of education and are less likely to have young children than black women. Whites also tend to earn higher wages and are more likely to participate in the labor force than blacks. Finally, table 2 illustrates striking differences across race with respect to marital status: white women are twice as likely to be married and black women are twice as likely to be single or divorced.

Comparisons of married and remarried women, as well as comparisons of single and divorced women, may prove instructive, as it is these groups that are generally pooled together in studies of female labor supply. Considering the differences across race, statistics for black and white women are presented separately. The statistics in table 3 support the hypothesis that divorce may have persistent effects on labor market behavior after remarriage. For whites, remarried women supply significantly greater hours to the labor market than married women, even though remarried women are less educated on average and do not differ significantly in terms of other observed characteristics. Interestingly, while differences in labor market behavior across married and remarried women stem from differences in hours for whites, variation for blacks is through participation rates. Although there are no significant differences in average characteristics, black remarried women are significantly more likely to work than their married counterparts.

Substantial differences in individual characteristics are evident upon comparison of single and divorced white women. Single whites are younger, have higher wages and are more educated on average than divorced whites. Despite these differences, participation rates and hours worked for single and divorced women are quite similar in the sample of white women. For blacks, there are no significant differences in terms of observed characteristics or labor market behavior. As expected, single and divorced white women tend to work more and have higher annual hours than married and remarried whites; surprisingly, the converse holds for black women. The sample statistics suggest labor force participation rates and annual hours worked vary substantially across women in different marital states. The extent to which such differences are explained by differences in observed characteristics is the focus of the empirical analysis below.

3. Econometric Specification

For ease of comparison with previous studies, a simple version of the standard labor supply model is adopted. In this model, women choose their hours of market work to maximize utility subject to a budget constraint. The full marital history, as compared to an indicator of whether a woman is currently married, is included in the specification to capture differences in labor market behavior among women who have divorced in the past. It is assumed the resulting schedule for desired hours of work can be expressed as

$$h_{it}^{*} = \alpha_{it}^{h} X_{it}^{h} + \varepsilon_{it}, \tag{1}$$

where desired hours are only observed when $h^{*}>0$. The set of characteristics assumed to determine hours (X^{h}) includes the woman's wage, her age, education, the presence of children under the age of 6, spousal labor income and education if married and her marital history.

The market wage is a function of a set of standard characteristics (X^{w}) including education, age, experience, region of residence, as well as the marital history

$$w_{it} = \alpha_{it}^{w} X_{it}^{w} + u_{it}. \tag{2}$$

The marital history is included in the wage equation, as previous studies on male earnings have found marital status to be an important determinant of wages.[12]

Women decide to participate in the labor market as long as the market wage exceeds the reservation wage (r). It is assumed the reservation wage is a function of characteristics (X^r) including education, the marital history, spousal income and education if married or remarried and the presence of young children in the household

$$r_{it} = \alpha_{it}^{r} X_{it}^{r} + \upsilon_{it}. \tag{3}$$

Before proceeding with estimation, an important issue must be addressed. All of the elements of X^h, X^w and X^r are generally assumed to be exogenous. It is very likely that this assumption is violated in practice. For one, experience may be correlated with the unobserved components determining the wage and hence the participation and hours decisions. When used as an instrument for the market wage, previous studies have found experience induces an upward bias on the wage parameter in the labor supply function (Mroz, 1987). Children are an oft-cited example of another variable which is likely endogenous in labor supply decisions (Browning, 1992).

An issue receiving less attention in studies of labor supply is the endogeneity of marital status. The substantial literature on female labor supply has ignored the potential endogeneity of marital status in the labor supply decision to a large extent.[13] In general, labor supply functions have been estimated using a sample of married women or controls for current marital status only have been included in the labor supply function. Both approaches rely on the assumption that unobserved factors determining marital status decisions do not also determine labor supply.

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[12] See for example Cornwell and Rupert (1997), Nakosteen and Zimmer (1987) and Korenman and Neumark (1991).
[13] Important exceptions include Johnson and Skinner (1986, 1988) and van der Klaauw (1996).

Violating this assumption in the first approach induces selection bias in the labor supply and participation estimates, as the sample of currently married women will not represent a random sample of women. In terms of the latter approach, including endogenous controls for marital status in the labor supply equation will introduce bias in the parameter estimates of interest. Furthermore, if individuals sort in the marriage market on unobserved characteristics correlated with labor supply, then the exogeneity of spousal characteristics in the labor supply function is also called into question. Therefore, it is of interest to compare the parameter estimates from a specification that assumes marital status and other covariates are exogenous to one that allows such variables to be correlated with unobserved components determining the hours and participation decisions.

One of the primary advantages of using panel versus cross-sectional data is that one can attempt to control for the correlation between unobserved individual heterogeneity and the regressors in the equation of interest. It is this advantage that is exploited where possible to assess the consequences of ignoring the potential endogeneity of marital status. Two alternative assumptions are imposed in estimation. Consider the desired hours equation from (1) in the case of panel data, with N individuals observed over T periods

$$h_{it}^{*} = \alpha_{it}^{h} X_{it}^{h} + \varepsilon_{it}$$

where i=1,2,...N and t=1,2,...T. Decomposing ε_{it} into two components yields

$$\varepsilon_{it} = \eta_i + e_{it}. \tag{4}$$

The first component (η_i) represents individual heterogeneity, capturing any time-invariant components that are specific to the individual and not included in the set of regressors (Baltagi, 1995). The second component (e_{it}) varies over time and across individuals and by assumption is not correlated with the regressors or with η_i. Under the assumption that η_i is uncorrelated with the regressors, the random effects estimator produces consistent and efficient estimates. Unlike random effects, the fixed effects estimator does not require η_i to be uncorrelated with the regressors. Therefore, if the fixed effects estimator produces results that are significantly different than those from the random effects estimator, it is likely the regressors are endogenous and the parameter estimates from the latter are inconsistent.[14] Estimates of participation and labor supply equations from the random and fixed effects specifications are compared below to assess the extent to which some of the regressors may be endogenous.[15]

For the linear regression case, estimation of the fixed effects specification is straightforward in most instances. Generally, it is possible to apply the within-groups estimator to remove η_i from the equation of interest. Unfortunately, it is difficult to control for sample selection on e_{it} in this framework. The difficulty arises due to the fact that individual heterogeneity in the sample selection rule enters the labor supply function in a non-linear fashion and therefore cannot be removed by differencing (Kyriazidou, 1997). For similar reasons, controlling for individual heterogeneity in non-linear labor force participation functions is equally difficult.

Direct estimation of the fixed individual effects eliminates both of the above problems. In general, it is not practical to estimate so many individual coefficients. More importantly,

[14] It is also possible that the random effects results differ from the fixed effects results because the distributional assumption imposed on η_i is incorrect.

[15] If changes in marital status are correlated with transitory shocks, then fixed effects estimation will not solve the endogeneity problem. The common approach to deal with this issue is to instrument marital status. This approach is not taken here because it is difficult to find appropriate instruments in the data. However, it is expected that unobserved preferences over work and marriage are likely constant over time and will be captured by controlling for time-invariant unobserved heterogeneity in estimation.

estimation of the fixed effects for individuals with very few time periods is inconsistent.[16] Moving from the unbalanced to the 14 year balanced panel mitigates both problems, as it reduces the sample size such that estimation of the individual fixed effects is feasible in the participation, hours and wage functions and, considering the length of the panel, is likely to produce consistent estimates.[17,18]

4. Results

4.1 Participation

To assess the potential endogeneity of marital status as well as other determinants of labor supply, two specifications of the labor force participation decision are estimated. Results are presented in table 4. To estimate the fixed effects version of the model, individuals who do not change their labor force status during the 14 years of the panel must be eliminated. This restriction reduces the number of white women in the sample from 631 to 297. Column 1 contains labor force participation estimates for white women, derived under the assumption that the regressors are uncorrelated with the unobserved determinants of participation. The results

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[16] To deal with this problem, several approaches, for example Chamberlain (1982), Wooldridge (1995) and Bover and Arellano (1997), assume the fixed effects are linear functions of the exogenous variables in the model. This restrictive assumption may also result in serious inconsistencies in the parameter estimates if the fixed effects are not correctly specified.

[17] There are disadvantages to moving from the unbalanced to the balanced panel. If observations are randomly missing from the data then using a balanced panel, with fewer observations, results in a loss of efficiency. A more serious problem arises when the balanced sample represents a self-selected group of individuals: in particular, if the probability of discontinuing in the sample is correlated with labor supply behavior or marital status, inconsistency in the parameter estimates of the labor supply function will result. Sample statistics for the balanced and unbalanced panels are presented in table B1 of the Appendix to assess the representativeness of the restricted sample. The average characteristics are similar across the samples, suggesting the restricted sample is quite representative of the full sample. Table B2 contains a comparison of parameter estimates from a random effects employment probit on the balanced and unbalanced panels. With few exceptions, the estimates are comparable in sign, significance and magnitude across the two samples. Considering the evidence, it is likely that any inconsistencies introduced by moving to the balanced panel are outweighed by the availability of more robust estimation methods.

[18] Heckman (1981) provides Monte Carlo evidence that the fixed effects probit performs well on panels as short as 8 years in the case where no lagged dummy variables are included in the model.

from the fixed effects specification for whites are presented in column 2. The analogous results for black women are presented in columns 3 and 4, where the restriction of the sample to women whose participation status changes over the 14 years of the panel reduces the sample from 326 to 142 women.[19]

Of particular interest are the coefficients on the marital history across specifications. In general, the results tend to vary widely across specifications and race. For whites, there is no significant difference in employment probabilities across single and divorced women, or across married and remarried women in the random effects case. Moving to the fixed effects estimates, the marital history parameters increase in magnitude and significance: the estimated coefficient for remarried women increases by a factor of 4 and for single women by a factor of 20. Under this specification, there now exist significant differences in participation probabilities among women in different marital histories. The probability of participating in the labor market is higher for remarried women than for married women, consistent with the hypothesis that remarried women work more than married women in response to a higher probability of future divorce. As expected, divorced whites are less likely to work than their single counterparts. One possible interpretation of this finding is that marital-specific capital, accumulated while married, may increase the opportunity costs of working.

The large increase in the marital history coefficients indicates the presence of a positive correlation between unobserved factors correlated with the married state and the participation decision. This finding is consistent with the literature on the earnings premium for married men,

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[19] Table C1 in the Appendix compares the average characteristics of the balanced relative to the unbalanced panel. Table C2 compares parameter estimates from random effects labor force participation probits on both samples. Although the sample statistics are quite different across the samples, the estimated coefficients are similar in sign and significance across specifications, suggesting that the differences across the samples do not result in marked differences for inference.

where it is suggested that some individual traits may be desirable in both the marriage and labor markets (Cornwell and Rupert, 1997),[20] inducing a positive correlation between marriage and work.

Interestingly, the results for blacks vary considerably as compared to those for whites: in both the random and fixed effects specifications, there do not exist significant differences in the employment probabilities of black women in different marital states. This result contributes to the large body of evidence documenting black-white differences in employment and marriage behavior. To determine whether systematic differences exist in the parameter estimates from the fixed and random effects estimations, the test proposed by Durbin (1954), Wu (1973) and Hausman (1978) is applied. Under the null hypothesis, if the covariates are not correlated with the individual fixed effects, the estimates under the fixed and random effects specifications are both consistent although the latter are inefficient. If the covariates are correlated with the individual unobserved heterogeneity, only the fixed effects estimator will yield consistent estimates. Durbin-Wu-Hausman tests reject the random effects specification for both white and black women. Considering the large change in the marital history coefficients across specifications for whites, the results suggest marital status decisions are endogenous in the participation decision.[21]

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[20] Cornwell and Rupert (1997) find evidence of a large decrease in the marriage premium for men after controlling for fixed individual effects.

[21] The estimates for other standard determinants of participation from the random and fixed effects specifications are quite similar in many respects for white and black women. Spousal income and the presence of young children in the household have the expected negative effect on participation and in most cases do not change in terms of sign and significance across specifications. In general, spousal education is not a significant determinant of the participation decision for whites, as consistent with Johnson and Skinner (1986). However, black women with more educated spouses are more likely to work in the fixed effects specification.

4.2. Labor Supply

Estimates of the labor supply function under alternative assumptions regarding the endogeneity of the covariates are presented in table 5. A fixed effects wage specification, including an inverse Mill's ratio term to correct for sample selection on the idiosyncratic component of the error term (Heckman, 1979), is estimated and the predicted wage is included in the labor supply function.[22] Column 1 contains labor supply estimates under the random effects specification and column 2 contains estimates under the fixed effects specification for whites. Column 3 contains a selection-corrected version of the specification in column 2. The analogous results for black women are presented in columns 4 to 6.

The effects of the marital history on labor supply generally support the findings from the participation estimates. For whites, the random effects coefficients appear to be biased when compared to the parameters from specifications that allow for the time-invariant unobserved heterogeneity to be correlated with the regressors. In the latter specifications, single white women supply significantly higher hours to the labor market than divorced white women and likewise for remarried and married women. In contrast, black women in different marital histories do not differ significantly in terms of their labor supply behavior. One exception is that divorced black women tend to work fewer hours than married women; however this effect is not significantly different from zero once controls for sample selection are included in the labor supply function.

As consistent with the participation estimates, Durbin-Wu-Hausman tests reject random effects in favor of the fixed effects specification for whites and blacks. For whites in particular,

[22] The random effects wage specification was rejected by Durbin-Wu-Hausman tests for both whites and blacks. Wage estimates are presented in table A1 of the Appendix.

the results from the preferred specification indicate the entire marital history is important in the labor supply decision.[23]

5. Conclusion

Previous research suggests marriage and divorce are important determinants of labor market decisions for women. This paper presents evidence that the entire marital history may play a role in current labor supply decisions. After controlling for a set of standard characteristics and for fixed individual effects, remarried white women are more likely to participate and supply greater hours to the labor market than married white women. The results also indicate single white women have significantly higher participation rates and hours than divorced women.

A large literature has been concerned with explaining the rise in participation rates of married women since the 1960s. As mentioned in the introduction, if remarried women have higher labor force participation than married women, then the labor force participation rates of all married women may have increased in part as a consequence of an increase in the number of remarried women over time. The proportion of remarried white women in the sample of currently married women and the participation rate of currently married white women increased by 14 and 15 percentage points, respectively over the period 1979 to 1992. According to the estimates in table 4, the effect of remarriage on the probability of participating in the labor

[23] In terms of standard labor supply determinants, the estimated wage elasticity falls considerably when moving from the random to the fixed effects specification and after correcting for selection bias is no longer significantly different from zero for both blacks and whites. Interestingly, after controlling for unobserved individual heterogeneity and sample selection, spousal income has no effect on the labor supply decision for all women. Children have an insignificant effect on the labor supply of whites, as consistent with Johnson and Skinner (1986). However, children have a positive effect on the labor supply of black women. Spousal education does not appear to influence the labor supply of white women, consistent with the results for participation and with Johnson and Skinner (1986). However, black women with university-educated spouses tend to supply fewer hours to the labor market than women with less educated spouses.

market is approximately 12 percentage points.[24] As a result, the increased number of remarried women in the sample of married women over time raised the total participation rate of married women by 1.7%. In other words, 11.2% of the total change in the labor force participation rates of married women is due to the increase in the number of remarried women in the population.

The relationship between the marital history and labor supply tends to vary markedly by race. In contrast to whites, there are no significant differences in employment behavior for black women in different marital states. The wide variation in marriage and employment patterns is well known. However, the extent to which differences in underlying marriage or labor market conditions contribute to the differences in observed behavior across race warrants further study.

[24] Average derivates for the parameter estimates in table 4 are available upon request.

Table 1. Sample Statistics for the Balanced Panel, 1979-1992

Year	Participation	Hours	Single	Married	Divorced	Remarried
1979	.72	1449	.09	.67	.12	.12
1980	.74	1428	.08	.66	.13	.13
1981	.75	1507	.08	.63	.15	.14
1982	.78	1477	.07	.62	.17	.14
1983	.77	1478	.07	.61	.17	.15
1984	.79	1532	.07	.59	.18	.17
1985	.82	1572	.06	.58	.17	.18
1986	.81	1596	.06	.58	.16	.19
1987	.81	1645	.06	.57	.17	.20
1988	.81	1666	.06	.56	.18	.20
1989	.82	1683	.06	.55	.19	.21
1990	.84	1649	.06	.53	.20	.21
1991	.85	1661	.06	.53	.21	.21
1992	.83	1701	.05	.52	.22	.21
Number of Observations						957

Table 2: Sample Statistics by Race, 1992 Cross-Section			
Variable	Whites	Blacks	Equality of Means (t-Statistic)
Age	45.543	45.145	1.098
	(0.202)	(0.301)	
Children	0.060	0.188	-5.587
under 6	(0.009)	(0.021)	
High School	0.414	0.393	0.908
	(0.019)	(0.026)	
Some College	0.489	0.362	3.931
	(0.020)	(0.026)	
Experience	18.676	19.445	-2.195
	(0.310)	(0.504)	
Single	0.049	0.087	-2.152
	(0.009)	(0.015)	
Married	0.526	0.263	8.557
	(0.020)	(0.024)	
Divorced	0.187	0.419	-7.566
	(0.015)	(0.027)	
Remarried	0.214	0.137	3.137
	(0.016)	(0.019)	
Wage	8.809	6.434	5.812
	(0.266)	(0.311)	
Hours	1697.367	1822.754	-2.712
	(30.962)	(34.345)	
Participation	0.730	0.679	1.653
	(0.017)	(0.025)	
N	631	326	

Notes: Standard Errors in Parentheses

Table 3: Sample Statistics by Marital Status, 1992 Cross-Section

Variable	Whites M	Whites $R^{\ddagger}$	Whites S	Whites $D^{\dagger}$	Blacks M	Blacks $R^{\ddagger}$	Blacks S	Blacks $D^{\dagger}$
Age	45.86	45.08	42.84	45.55	45.41	44.16	43.78	45.05
	(0.28)	(0.42)	(0.86)	(0.48)	(0.58)	(0.84)	(0.80)	(0.47)
		[1.55]		[2.75]		[1.37]		[1.37]
Children	0.07	0.08	0.04	0.02	0.21	0.12	0.11	0.18
under 6	(0.01)	(0.02)	(0.03)	(0.01)	(0.04)	(0.05)	(0.05)	(0.03)
		[-0.45]		[0.63]		[1.41]		[1.20]
High School	0.41	0.49	0.16	0.41	0.43	0.44	0.45	0.39
	(0.03)	(0.04)	(0.07)	(0.04)	(0.05)	(0.08)	(0.08)	(0.04)
		[-1.60]		[-3.10]		[-0.11]		[0.67]
Some	0.50	0.40	0.81	0.49	0.35	0.41	0.23	0.36
College	(0.03)	(0.04)	(0.07)	(0.05)	(0.05)	(0.08)	(0.07)	(0.04)
		[2.00]		[3.72]		[-0.64]		[-1.61]
Experience	17.66	18.75	22.50	20.61	21.05	20.16	18.99	19.71
	(0.43)	(0.60)	(1.19)	(0.74)	(0.89)	(0.84)	(1.58)	(0.74)
		[1.48]		[1.35]		[0.73]		[-0.41]
Wage	10.13	9.89	14.45	11.51	9.73	10.04	7.62	8.51
	(0.37)	(0.46)	(1.17)	(0.58)	(0.54)	(0.90)	(0.80)	(0.45)
		[0.41]		[2.25]		[-0.27]		[0.97]
Participation	0.80	0.84	0.94	0.93	0.73	0.91	0.66	0.72
	(0.02)	(0.03)	(0.04)	(0.02)	(0.05)	(0.05)	(0.07)	(0.04)
		[1.11]		[0.22]		[-2.55]		[0.74]
Annual	1510.05	1751.50	1999.36	1988.30	1830.18	1827.16	1884.99	1788.78
Hours	(42.26)	(69.42)	(119.18)	(61.92)	(47.21)	(123.10)	(74.58)	(50.61)
		[-2.97]		[0.08]		[0.02]		[1.07]
Spousal	38836.31	34787.67			22085.92	23213.87		
Labor	(1602.01)	(2075.38)			(1542.21)	(2404.58)		
Income		[1.54]				[-0.39]		
Spouse –	0.29	0.40			0.49	0.38		
High School	(0.02)	(0.04)			(0.05)	(0.08)		
		[2.46]				[1.17]		
Spouse –	0.57	0.47			0.29	0.48		
University	(0.03)	(0.04)			(0.05)	(0.08)		
		[2.00]				[2.01]		
Number of observations	338	139	32	122	93	41	43	150

Notes: Standard errors in parentheses. Hours are conditional on participation.
† t-statistic for equality of means between single and divorced women in brackets.
‡ t-statistic for equality of means between married and remarried women in brackets.

Table 4: Labor Force Participation Estimates from Random and Fixed Effects Probit Specifications

Variable	Whites: Random Effects	Whites: Fixed Effects	Blacks: Random Effects	Blacks: Fixed Effects
Spousal Income/1000	-0.010***	-0.015***	-0.016	-0.023***
	(0.002)	(0.002)	(0.011)	(0.007)
Children under 6	-0.660***	-0.784***	-0.570***	-0.825***
	(0.086)	(0.077)	(0.135)	(0.104)
Single	0.272	5.84***	-0.038	0.266
	(0.433)	(1.50)	(0.517)	(0.297)
Divorced	0.422**	0.773***	-0.239	-0.218
	(0.168)	(0.245)	(0.307)	(0.273)
Remarried	0.090	0.364**	0.263	0.162
	(0.097)	(0.181)	(0.322)	(0.287)
Spouse – High School	0.034	-0.020	0.586	0.936***
	(0.124)	(0.180)	(0.370)	(0.273)
Spouse – Some College	0.211	0.267	0.031	0.478*
	(0.134)	(0.200)	(0.306)	(0.275)
Durbin-Wu-Hausman Test Statistic: χ^2 (27)	221.459		109.669	
Number of individuals	297		142	
Number of observations	4158		1988	

Notes: Standard Errors in Parentheses. The dependent variable, participation, is equal to one if the respondent works 774 hours or more in the survey year and zero otherwise. All specifications also contain controls for region of residence, urbanicity, time effects, age and experience. Random effect specification also includes controls for education. Where applicable, education is interacted with the time effects. Full regression results are available upon request.

*coefficient significant at 10% level

**coefficient significant at 5% level

***coefficient significant at 1% level

Table 5: Labor Supply Estimates

Variable	White Women: Random Effects	White Women: Fixed Effects	White Women: Fixed Effects Selection Corrected	Black Women: Random Effects	Black Women: Fixed Effects	Black Women: Fixed Effects Selection Corrected
ln(Wage)	0.396***	0.019	-0.019	0.551***	0.417***	0.228
	(0.045)	(0.042)	(0.068)	(0.076)	(0.095)	(0.212)
Children	-0.300***	-0.230***	-0.054	-0.127**	-0.037	0.099**
under 6	(0.049)	(0.026)	(0.034)	(0.057)	(0.029)	(0.045)
Spousal	-0.007***	-0.004***	-0.0005	-0.008***	-0.005***	-0.0002
Income/1000	(0.001)	(0.001)	(0.0009)	(0.003)	(0.002)	(0.003)
Spouse –	-0.161	0.048	0.076	0.051	-0.285***	-0.250**
High School	(0.080)	(0.069)	(0.073)	(0.130)	(0.070)	(0.0128)
Spouse – Some	-0.011	0.026	-0.013	-0.023	-0.196***	-0.106
College	(0.079)	(0.073)	(0.078)	(0.108)	(0.070)	(0.109)
Single	0.135	0.505***	0.248**	-0.032	-0.045	-0.041
	(0.097)	(0.125)	(0.120)	(0.136)	(0.096)	(0.154)
Divorced	0.129*	0.226**	0.066	-0.169	-0.199**	-0.053
	(0.077)	(0.080)	(0.088)	(0.136)	(0.083)	(0.125)
Remarried	0.063	0.166***	0.109*	-0.044	-0.070	-0.031
	(0.051)	(0.051)	(0.058)	(0.130)	(0.075)	(0.129)
Durbin-Wu-Hausman Test Statistic: χ^2 (23)		131.457			139.705	
Selection Correction Term			0.538***			0.500***
			(0.040)			(0.111)
Number of Individuals		605			307	
Number of Observations		7022			3574	

Notes: Standard errors in parentheses. The dependent variable is the log of annual hours. All specifications also contain controls for age and time effects. Random effect specifications also include controls for education, interacted with time effects. The selection correction term is constructed using a fixed effect probit of non-zero hours of work in the survey year. Full regression results are available upon request.

*coefficient significant at 10% level

**coefficient significant at 5% level

***coefficient significant at 1% level

Appendix
A: Estimated Wage Regressions

Table A1: Wage Regression Estimates

	White Women			Black Women		
Variable	Random Effects	Fixed Effects	Fixed Effects Selection Corrected	Random Effects	Fixed Effects	Fixed Effects Selection Corrected
Experience	0.034***	0.079***	0.076***	0.008	0.043***	0.034
	(0.003)	(0.008)	(0.013)	(0.007)	(0.014)	(0.039)
Urban	0.185***	0.024	0.026	-0.052	-0.127***	-0.141
	(0.036)	(0.021)	(0.040)	(0.090)	(0.035)	(0.124)
Northeast	0.108**	0.192***	0.124	0.194	0.251***	0.163
	(0.054)	(0.069)	(0.106)	(0.138)	(0.090)	(0.134)
West	0.031	0.241***	0.232**	0.149	0.345***	0.371***
	(0.056)	(0.048)	(0.108)	(0.137)	(0.074)	(0.128)
South	0.111**	0.161***	0.154*	-0.020	-0.072	-0.149
	(0.044)	(0.058)	(0.092)	(0.092)	(0.067)	(0.102)
Single	0.022	0.314***	0.236*	-0.066	0.270***	0.210
	(0.068)	(0.065)	(0.131)	(0.089)	(0.053)	(0.147)
Divorced	0.028	0.059**	0.017	-0.068	-0.019	-0.023
	(0.041)	(0.029)	(0.051)	(0.085)	(0.040)	(0.093)
Remarried	-0.015	-0.021	-0.029	0.102	-0.002	-0.004
	(0.045)	(0.032)	(0.059)	(0.140)	(0.051)	(0.106)
Age	0.008	0.050***	-0.008	-0.025	0.024**	-0.051
	(0.017)	(0.007)	(0.023)	(0.030)	(0.009)	(0.036)
Square of Age	-0.0003	-0.0003***	-0.0002	-0.0003	0.00003	0.0002
	(0.0002)	(0.0001)	(0.0002)	(0.0004)	(0.0001)	(0.0003)
Durbin-Wu-Hausman Test Statistic: χ^2 (22)	238.779			102.54		
Selection Correction Term			0.086***			0.100
			(0.038)			(0.091)
Number of Individuals		6787			3524	
Number of Observations		604			306	

Notes: Standard errors in parentheses. The dependent variable is the log of annual hours. All specifications also contain controls for time effects. Random effects specifications also include controls for education. The selection correction term is constructed using a fixed effects probit of non-zero hours of work in the survey year. Full regression results are available upon request.
*coefficient significant at 10% level
**coefficient significant at 5% level
***coefficient significant at 1% level

B. Comparison of Balanced and Unbalanced Samples

Table B1: Comparison of Sample Statistics for the Unbalanced and Balanced Panels, 1979-1992

Variable	Unbalanced Panel	Balanced Panel
Age	37.983	38.896
	(0.05)	(0.057)
Children under Age 6	0.260	0.222
	(0.002)	(0.004)
High School	0.421	0.413
	(0.003)	(0.004)
Some College	0.459	0.480
	(0.003)	(0.004)
Black	0.112	0.090
	(0.002)	(0.002)
Single	0.097	0.065
	(0.002)	(0.002)
Married	0.574	0.586
	(0.003)	(0.004)
Divorced	0.163	0.173
	(0.002)	(0.003)
Remarried	0.167	0.175
	(0.002)	(0.003)
Wage	7.919	8.227
	(0.030)	(0.049)
Participation	0.790	0.797
	(0.002)	(0.003)
Hours	1588.287	1578.606
	(4.532)	(6.920)
Number of individuals	3739	957
Number of observations	33822	13398

Notes: Standard errors in parentheses.

Table B2: Labor Force Participation Estimates from Random Effects Probits

Variable	Unbalanced Panel	Balanced Panel
Spousal Income/1000	-0.005*	-0.007*
	(0.001)	(0.001)
Children under 6	-0.474*	-0.473*
	(0.019)	(0.029)
Single	0.246*	0.444*
	(0.061)	(0.133)
Divorced	0.252*	0.259*
	(0.045)	(0.077)
Remarried	0.070*	0.146*
	(0.033)	(0.055)
High School	0.503*	0.548*
	(0.053)	(0.108)
Some College	0.951*	0.975*
	(0.058)	(0.113)
Number of individuals	3739	957
Number of observations	33822	13398

Notes: Standard errors in parentheses. The dependent variable, participation, is equal to one if the respondent works 774 hours or more in the survey year and zero otherwise. All specifications also contain controls for region of residence, urbanicity, age, spousal education and time effects. Full regression results are available upon request.
*coefficient significant at 5% level

C. Comparison of Balanced and Restricted Samples

Table C1: Comparison of Sample Statistics for the Balanced and Restricted Panels, 1979-1992

Variable	Balanced Panel	Restricted Panel
Age	38.896	38.378
	(0.057)	(0.083)
Children under Age 6	0.222	0.279
	(0.004)	(0.006)
High School	0.413	0.454
	(0.004)	(0.006)
Some College	0.480	0.400
	(0.004)	(0.006)
Black	0.090	0.084
	(0.002)	(0.004)
Single	0.065	0.0219
	(0.002)	(0.002)
Married	0.586	0.638
	(0.004)	(0.006)
Divorced	0.173	0.128
	(0.003)	(0.004)
Remarried	0.175	0.212
	(0.003)	(0.005)
Wage	8.227	6.407
	(0.049)	(0.068)
Participation	0.797	0.652
	(0.003)	(0.006)
Hours	1578.606	1279.317
	(6.920)	(12.182)
Number of individuals	957	439
Number of observations	13398	6146

Notes: Standard errors in parentheses.

Table C2: Labor Force Participation Estimates from Random Effects Probits

Variable	Balanced Panel	Restricted Panel
Spousal Income/1000	-0.007*	-0.009*
	(0.001)	(0.001)
Children under 6	-0.473*	-0.619*
	(0.029)	(0.043)
Single	0.444*	0.414*
	(0.133)	(0.201)
Divorced	0.259*	0.332*
	(0.077)	(0.103)
Remarried	0.146*	0.151*
	(0.055)	(0.070)
High School	0.548*	0.418*
	(0.108)	(0.107)
Some College	0.081	0.158
	(0.068)	(0.087)
Number of individuals	957	439
Number of observations	13398	6146

Notes: Standard errors in parentheses. The dependent variable, participation, is equal to one if the respondent works 774 hours or more in the survey year and zero otherwise. All specifications also contain controls for region of residence, urbanicity, age, spousal education and time effects. Full regression results are available upon request.
*coefficient significant at 5% level

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