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Marriage matching, risk sharing and spousal labor supplies

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Abstract

This paper develops the collective marriage matching model, a behavioral and empirically flexible framework that incorporates both marriage matching and intrahousehold allocations. The model shows how marriage market equilibrium and bargaining power within the family are simultaneously determined. The framework provides a solution to the problem of incorporating substitute sex ratios in empirical models of spousal labor supplies. Using data from the US 2000 census, the empirical results show that changes in marriage market tightness, the ratio of unmarried men to unmarried women, have large estimated effects on spousal labor force participation rates, and smaller effects on hours of work and hours in home production. Controlling for variation in labor market conditions across marriage markets has substantive implications for the parameter estimates.

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1 Introduction

A well known finding from recent empirical models of intrahousehold allocations is that marriage market conditions affect intrahousehold allocations.[1] Simultaneously, other researchers have investigated empirical models of marriage matching or choice of living arrangements more broadly.[2] As Lundberg and Pollak (1996) have observed, an open problem is to integrate these two lines of research.

The primary contribution of this paper is to develop the collective marriage matching model, a behavioral and empirically flexible framework that incorporates both marriage matching and intrahousehold allocations.[3] We use the collective model of intrahousehold allocation developed by Chiappori (1992, 1998) and his collaborators. The attractive feature of the collective model is that it is empirically tractable and it provides testable restrictions on behavior while minimizing strong functional form assumptions. We embed a collective model within the Choo Siow (Choo Siow 2006; hereafter CS) marriage matching model. CS does not impose any *a priori* restriction on the marriage matching distribution. Except for thin cells, it can fit any observed cross section marriage matching distribution.

Our version of the collective model allows for public goods following Blundell, Chiappori and Meghir (2005; hereafter BCM), efficient spousal risk sharing, and continuous/discrete spousal labor supply behavior. Our framework shows how the estimate of the effect of marriage market conditions on the labor force participation of married women is informative on the relative bargaining power of married women who work versus those who do not.

In the marriage market, individuals choose whether to remain unmarried or who to marry and their living arrangement. *The utility weights of husbands relative to their wives in the collective model are used to clear the marriage market.* We show the existence of marriage market equilibrium. The transferable utilities marriage market model, for example Becker (1991) and CS, is a special case of our collective model of marriage matching. In the collective tradition, Chiappori, Lacroix and Fortin (2002; hereafter CFL) and BCM are other special cases.

To keep our paper to a manageble length, we defer structural estimation

[1] See, for example, Angrist (2002), Chiappori, Fortin, and Lacroix (2002), Francis (2005), Grossbard-Schectman (1984, 1993), Lafortune (2008).

[2] E.g. Black, et. al. (2003); Baker, et. al. (2001), Lillard, et. al. (2006).

[3] An alternative formulation is Del Boca and Flinn (2006).

and testing of our collective marriage matching model to our companion paper, Choo, Seitz, and Siow, 2008; hereafter CSS. Our empirical contribution here is more limited. We exploit our theoretical structure to provide a solution to the problem of accounting for substitution effects when estimating the reduced form impact of marriage market conditions on spousal labor supplies. Currently, a researcher constructs an own sex ratio based on the type of husbands and wives whose labor supplies the researcher is studying. The researcher ignores the sex ratios of other types of men and women due to the problem of multicollinearity.[4] Of course ignoring substitute sex ratios makes it difficult to interpret existing empirical results. Our model provides a potential solution to this problem, a sufficient statistic for the marriage market condition facing any type i men and type j women. We call this statistic *marriage market tightness*, which is the ratio of unmarried men of type i to unmarried women of type j.[5] Substitution effects are embedded in marriage market tightness because the presence of substitutes will affect tightness. Unlike sex ratios, marriage market tightness is an endogenous variable and this endogeneity has to be dealt with.

We estimate the effect of market tightness on two elements of household behavior. First, we estimate models of household labor supply using the 2000 Census where each state is treated as a separate marriage market. Second, we estimate models of hours spent in home production using the 2003 American Time Use Survey (ATUS). Our estimation results are presented in Section 5. After controlling for labor market conditions, state effects, and individual characteristics, marriage market tightness is negatively (positively) correlated with wives' (husbands') labor supplies and hours in home production. A one standard deviation increase in marriage market tightness leads to more than a one quarter standard deviation decrease (increase) in wives' (husbands') log odds of participating in the labor market. The effects on usual hours of work per week and annual weeks worked are quantitatively smaller. Thus changes in marriage market tightness have quantitatively significant effects on intrahousehold reallocations in the direction predicted by our theory. Controlling for labor market conditions changes the parameters estimates substantially. To deal with the endogeneity issue, we instrument tightness for each $\{i, j\}$ match in each state with the birth rates and parental

[4]The problem of how many prices of substitutes or complements to include in the estimation of demand and supply functions are of course well known.

[5]Seitz (2004) was the first to discuss marriage market tightness.

educational attaintment at birth of the i and j individuals in that state. Controlling for endogeneity of market tightness is more important for the labor supplies of husbands than for the labor supplies of wives. Adding individual effects do not change the point estimates of tightness on spousal labor supplies.

In addition to spousal labor supplies, we also investigate the effect of tightness on other intrahousehold time use measures including leisure and hours spent on home production. Due to the small sample sizes, our empirical foray using time use data is at best suggestive. Most of the estimated signs are consistent with the theory. Thus the time use results adds to the evidence that marriage market conditions affect intrahousehold allocation of time use. But the large standard errors preclude any definite conclusion about the magnitudes of the effects.

In our companion paper CSS, we show that the parameters of the sharing rule in CFL are identified with marriage matching data. The substantive restriction in CFL is that it does not admit household public goods. As a consequence, CFL and CSS apply the empirical framework to couples without children. Without public goods, equilibrium household labor supplies can be decentralized into each spouse maximizing their own utility subject to their own wage and a share of non-labor assets (the "sharing rule"). This decentralization results in spousal indirect utilities which are functions of own wages and shares of non-labor assets. In the marriage market, individuals base their marital decisions on comparing the spousal indirect utilities that they can obtain from different marital choices. Changes in shares of non-labor assets via changes in the sharing rule will affect the level of spousal utility obtained in a particular marital match. From this paper, we know that the sharing rule adjusts to equilibrate the marriage market. CSS shows that the determinants of such equilibrium sharing rules are identified from marriage matching data. We exploit the simple indirect utilities from the CFL framework to estimate marital choice models.

CSS shows that the consideration of marriage matching introduces new determinants of sharing rule to the collective framework which has not be emphasized before. These new determinants are alternative wages and non-labor incomes from alternative marital choices. Unlike the standard collective model which is silent on how own wages and non-labor incomes affect the sharing rule, CSS imposes apriori restrictions on how own wages and non-labor incomes, as well as alternative wages and non-labor incomes should affect the sharing rule.

Also using data from the 2000 census, the empirical part of CSS shows that estimates from marriage matching data are more consistent with the collective marriage matching model than estimates using spousal labor supplies data following CFL.

This paper and CSS are first passes at integrating empirical marriage matching with intrahousehold allocations and we have left out many relevant complications. We will suggests some extensions in the conclusion of the paper. Siow (forthcoming) surveys the research program.

2 The Model

Consider a society in which there are I types of men, $i = 1, .., I$, and J types of women, $j = 1, .., J$. Let m_i be the number of type i men and f_j be the number of type j women, and $\mathbf{M}$ and $\mathbf{F}$ are the vectors of the numbers of each type of men and women, respectively. The type of an individual is defined by his or her preferences and ex-ante opportunities.

The model has two stages. In the first stage, individuals choose whether to marry and who to marry. An $\{i, j\}$ marriage is a marriage between a type i man and a type j woman. At the time of marriage, wages and non-labor incomes for each marital choice are random variables. In the second stage, wages and non-labor income for each household are realized and household labor supplies and public and private consumption allocations are chosen. The rationale for including public good consumption within marriage is to capture resources allocated to children in the marriage.

Let C_{ij} (c_{ij}) be the private consumption of the wife (husband), K_{ij} the household's expenditures on public goods and H_{ij} (h_{ij}) the wife's (husband's) labor supply. We normalize the total amount of time for each individual to 1. Preferences for the wife and husband are described by:

$$U_{ij}[\Omega_{ij}(C_{ij}, 1 - H_{ij}), K_{ij}] + \Gamma_{ij} + \varepsilon_{ij},$$

and

$$u_{ij}[\omega_{ij}(c_{ij}, 1 - h_{ij}), K_{ij}] + \gamma_{ij} + \epsilon_{ij},$$

respectively. The felicity functions, $U_{ij}(.)$ and $u_{ij}(.)$, depend on i and j to allow for differences in home production technologies across different types of marriages. We will assume that the felicity functions are weakly separable in public good expenditures. The rationale for weak separability will become apparent later.

The invariant gains to an $\{i, j\}$ marriage for the women and men, Γ_{ij} and γ_{ij} respectively, shift utility according to the type of marriage and allow the model to fit the observed marriage matching patterns in the data. Although invariant gains vary across different types of marriage, an important restriction is that Γ_{ij} and γ_{ij} are additively separable from consumption and leisure.

We further assume ε_{ij} and ϵ_{ij} are random variables realized before marriage decisions are made. The realizations of these random variables allow different individuals of the same type to make different marital choices. ε_{ij} is independent of C_{ij}, H_{ij}, K_{ij} and also ϵ_{ij}. We assume that ε_{ij} or ϵ_{ij} do not depend on the specific identity of the spouse so that an individual will treat all potential spouses of the same type as perfect substitutes. This will allow us to use competitive marriage market clearing as our equilibrium concept.

If a woman chooses not to marry, then $i = 0$ and if a man chooses not to marry, $j = 0$.

2.1 The collective model with efficient risk sharing[6]

The objective of this section is to derive two results. First, we will show how efficient risk sharing affects the expected felicities of the spouses as bargaining power within the household changes. Second, we will establish conditions under which the wife will on average work more and the husband will on average work less as the bargaining power of the husband increases.[7] We start first with the household consumption and labor supply decisions in the second stage. Wages and nonlabor incomes are assumed to be specific to the match. Consider a husband and wife in an $\{i, j\}$ marriage. In this section, we will hold the match, $\{i, j\}$, fixed and therefore dispense with the $\{i, j\}$ subscripts.

Before marriage, $\overline{W}$, $\overline{w}$ and $\overline{A}$, which are the components of wages and nonlabor incomes common to all individuals in $\{i, j\}$ matches, are observed. However, total non-labor family income (A), the wife's wage (W) and the husband's wage (w) have random components (ξ^A, ξ^W, and ξ^w, respectively)

[6] The collective model with public goods builds on BCM. Efficient intrahousehold risk sharing builds on Chiappori 1999.

[7] This section considers only interior solutions for hours of work. An extension to the labor force participation decision, and other extensions, will be discussed in Section 4. See also Blundell, Chiappori, Magnac, and Meghir (2007).

that are not realized until after the marriage decision is made:

$$A = \overline{A} + \xi^A$$
$$W = \overline{W} + \xi^W$$
$$w = \overline{w} + \xi^w.$$

Let Z be the vector of parameters which characterize the joint distribution of A, W and w. We pose the efficient risk sharing spousal arrangement as a planner solving:

$$\max_{\{C,c,H,h,K\}} \mathbf{E}_Z \Big[U[\Omega(C, 1-H), K]\Big] + p\mathbf{E}_Z \Big[u[\omega(c, 1-h), K]\Big] \tag{P1}$$

subject to the family budget constraint

$$c + C + K \leq A + WH + wh,$$

where $\mathbf{E}_Z$ is the expectations operator. In problem (P1), the planner chooses family consumption and labor supplies to maximize the weighted sum of the wife's and the husband's expected felicities subject to their family budget constraint. The weight allocated to the husband's expected felicity is denoted by p, $p \in R^+$, where $p > 1$ implies the husband has more weight than the wife and vice versa. We will refer to p as the husband's power in an $\{i, j\}$ marriage. The social planner takes p as given when solving the second stage household problem. The determination of p itself occurs in the marriage market and is a central focus of this paper. When the intrahousehold allocation is the solution to problem (P1), the allocation is efficient.

Let $\overline{Q}(p, Z)$ and $\overline{q}(p, Z)$ be the expected indirect felicity functions of the wife and the husband, respectively. Appendix A shows that the solution to problem (P1) implies:

Proposition 1 *The changes in spousal expected utilities as the husband's power, p, increases, satisfy:*

$$\frac{\partial \overline{Q}(p, Z)}{\partial p} = -p \frac{\partial \overline{q}(p, Z)}{\partial p} < 0 \tag{1}$$

The wife's expected utility falls and the husband's expected utility increases as p increases. Equation (1) traces the redistribution of spousal expected utilities as the husband's power increases. We will now study how spousal labor supplies change as husband's power changes. A necessary condition for solving problem (P1) is that given realized wages and non-labor income the planner solves:

$$\max_{C,c,H,h,K} U[\Omega(C,1-H),K] + pu[\omega(c,1-h),K] \tag{P2}$$

subject to

$$c + C + K \leq A + WH + wh.$$

Denote $H^*(p,W,w,A)$ and $h^*(p,W,w,A)$ as the labor supplies for the wife and husband that result from solving (P2), for a given p,W,w,A. In general, it is difficult to determine analytically how spousal labor supplies respond to changes in p. Building on BCM, assume that an individual's felicity function is weakly separable in private goods (consumption and leisure) and the public good. Appendix B shows that by restricting leisure (with suitably defined individual private income) and the public good to be normal goods for each spouse:

Proposition 2 *The wife's labor supply is increasing in p whereas the husband's labor supply is decreasing in the husband's power, p:*

$$\frac{\partial H^*(p,W,w,A)}{\partial p} > 0 \quad \forall\, W,w,A, \tag{2}$$

$$\frac{\partial h^*(p,W,w,A)}{\partial p} < 0 \quad \forall\, W,w,A. \tag{3}$$

Equations (2) and (3) are expected. Problem (P2) is a unitary model of the family faced with nonlabor income A and wages W, w. Thus we cannot reject a unitary model of the family for $\{i,j\}$ couples in the same society if they share risk efficiently. For example, spousal labor supplies of different $\{i,j\}$ couples will satisfy Slutsky symmetry in response to different wage realizations. We will discuss the relationship between our risk sharing collective model and the deterministic collective model such as CFL in Section 4.2.

2.2 Marriage decisions in the first stage

Following CS, in the first stage agents decide whether to marry and who to marry if they choose to marry. We will use the additive random utility model to model this choice. Consider a particular woman of type j. Recall that she can choose between I types of men and whether or not to marry, $I+1$ alternatives in total. A particular man of type i can choose between J types of women and whether or not to marry, $J+1$ alternatives in total. The expected indirect utility from an $\{i,j\}$ marriage for a woman is:

$$\overline{V}_{ij}(p_{ij}, Z_{ij}, \varepsilon_{ij}) = \overline{Q}_{ij}(p_{ij}, Z_{ij}) + \Gamma_{ij} + \varepsilon_{ij},\ i = 0, ..I \tag{4}$$

where $i = 0$ means that she chooses to be unmarried.

The indirect utility from an $\{i,j\}$ marriage for a man is:

$$\overline{v}_{ij}(p_{ij}, Z_{ij}, \epsilon_{ij}) = q_{ij}(p_{ij}, Z_{ij}) + \gamma_{ij} + \epsilon_{ij},\ j = 0, .., J \tag{5}$$

and $j = 0$ means that he chooses to be unmarried.

Each individual draws an independent realization of the random component for each possible type of match he or she could enter and for remaining single. For a particular woman, let $\underline{\varepsilon_j} = [\varepsilon_{0j}, .., \varepsilon_{ij}, .., \varepsilon_{Ij}]$ denote the realizations of the shocks and $G(\underline{\varepsilon_j})$ denotes the joint density of $\underline{\varepsilon_j}$. Likewise for a man, let $\underline{\epsilon_i} = [\epsilon_{i0}, .., \epsilon_{ij}, .., \epsilon_{iJ}]$ denote the realizations of shocks and $g(\underline{\varepsilon_i})$ denotes the joint density of $\underline{\epsilon_i}$. The expected indirect utility from the female's optimal choice will satisfy:

$$\begin{aligned}\overline{V}^*(\underline{\varepsilon_j}) = \max[&\overline{V}_{0j}(p_{0j}, \varepsilon_{0j}, Z_{0j}), ..., \overline{V}_{ij}(p_{ij}, \varepsilon_{ij}, Z_{ij}), ..., \\ &\overline{V}_{Ij}(p_{Ij}, \varepsilon_{Ij}, Z_{Ij})]\end{aligned} \tag{6}$$

and the expected indirect utility from the male's optimal choice will satisfy:

$$\begin{aligned}\overline{v}^*(\underline{\epsilon_i}) = \max[&\overline{v}_{i0}(p_{i0}, \epsilon_{i0}, Z_{i0}), ..., \overline{v}_{ij}(p_{ij}, \epsilon_{ij}, Z_{ij}), ..., \\ &\overline{v}_{iJ}(p_{iJ}, \epsilon_{iJ}, Z_{iJ})]\end{aligned} \tag{7}$$

2.3 The Marriage Market

Let $\mathbf{p}$ be the matrix of husband's powers where a typical element is p_{ij} for $i,j \geq 1$. Under the assumption that there are many women and men of each type, the random vectors $\underline{\varepsilon_j}$ and $\underline{\epsilon_i}$ are independent of p. Assume further that $\underline{\varepsilon_j}$ and $\underline{\epsilon_i}$ are independent of all wages and assets in the society. Let $\Phi_{ij}(\mathbf{p})$ denote the probability that a woman of type j will choose a spouse of type i, $i = 0, ..I$. Since each woman of type j is solving the same spousal choice problem (6),

$$\Phi_{ij}(\mathbf{p}) = \Pr[\mathbf{R}_{ij}(i,i',j,\varepsilon_{ij}) > 0 \quad \forall i' \neq i] \tag{8}$$

$$= \int_{\varepsilon_{ij}=-\infty}^{\infty} \int_{\varepsilon_{0j}=-\infty}^{\mathbf{R}_{ij}(i,0,j,\varepsilon_{ij})} ... \int_{\varepsilon_{Ij}=-\infty}^{\mathbf{R}_{ij}(i,I,j,\varepsilon_{ij})} G(\underline{\varepsilon_j})\ d\varepsilon_{ij}\ d\underline{\varepsilon_{i'\neq ij}} \tag{9}$$

where

$$\mathbf{R}_{ij}(i,i',j,\varepsilon_{ij}) = \overline{Q}_{ij}(p_{ij}, Z_{ij}) + \Gamma_{ij} - \overline{Q}_{i'j}(p_{i'j}, Z_{i'j}) - \Gamma_{i'j} + \varepsilon_{ij}.$$

When there are f_j type j women, the number of type j women who want to choose type i spouses, $i = 0,..,I$ is approximated by

$$\overline{\mu}_{ij}(\mathbf{p}, f_j) = \Phi_{ij}(\mathbf{p}) f_j.$$

Using (8), for $i \geq 1$,

$$\frac{\partial \overline{\mu}_{ij}(\mathbf{p}, f_j)}{\partial p_{i'j}} = f_j \frac{\partial \Phi_{ij}(\mathbf{p})}{\partial p_{i'j}} = \begin{cases} \leq 0, & i' = i \\ \geq 0, & i' \neq i \end{cases} \tag{10}$$

where $\overline{\mu}_{ij}(\mathbf{p}, f_j)$ is the demand function by type j women for type i husbands. Equation (10) says that the demand function satisfies the weak gross substitute assumption. That is, the demand by type j women for type i husbands, $i \geq 1$, is weakly decreasing in p_{ij} and weakly increasing in $p_{i'j}$, $i' \neq i$. Such a result is expected. All other types of potential spouses, $i' \neq i$, are substitutes for type i spouses. When the bargaining power of type i spouses increase, demand for that type of spouse is expected to weakly fall and the demand for other types of spouses is expected to weakly increase.

Similarly, let $\phi_{ij}(\mathbf{p})$ denote the probability that a man of type i will choose a spouse of type j, $j = 0,..J$. Since each man of type i is solving the same spousal choice problem (7),

$$\phi_{ij}(\mathbf{p}) = \Pr(\mathbf{r}(j,j',i,\epsilon_{ij}) > 0 \ \ \forall j' \neq j) \tag{11}$$

$$= \int_{\epsilon_{ij}=-\infty}^{\infty} \int_{\epsilon_{i0}=-\infty}^{\mathbf{r}(j,0,i,\epsilon_{ij})} ... \int_{\varepsilon_{iJ}=-\infty}^{\mathbf{r}(j,J,i,\epsilon_{iJ})} g(\underline{\epsilon_i})\ d\epsilon_{ij}\ d\underline{\epsilon_{i,j'\neq j}}$$

where

$$\mathbf{r}(j,j',i,\epsilon_{ij}) = \overline{q}_{ij}(p_{ij}, Z_{ij}) + \gamma_{ij} - \overline{q}_{ij'}(p_{ij'}, Z_{ij'}) - \gamma_{ij'} + \epsilon_{ij}.$$

When there are m_i number of type i men, the number of type i men who want to choose type j spouses, $j = 0, .., J$ is approximated by

$$\underline{\mu}_{ij}(\mathbf{p}, m_i) = \phi_{ji}(\mathbf{p}) m_i.$$

Using (11), for $j \geq 1$,

$$\frac{\partial \underline{\mu}_{ij}(\mathbf{p}, f_j)}{\partial p_{ij'}} = m_i \frac{\partial \phi_{ij}(\mathbf{p})}{\partial p_{ij'}} = \begin{cases} \geq 0, & j' = j \\ \leq 0, & j' \neq j \end{cases} \tag{12}$$

where $\underline{\mu}_{ij}(\mathbf{p}, m_i)$ is the demand function by type i men for type j wives.

Marriage market clearing requires the supply of wives (husbands) to be equal to the demand (husbands) for wives for each type of marriage:

$$\underline{\mu}_{ij} = \overline{\mu}_{ij} = \mu_{ij} \; \forall \; \{i > 0, \; j > 0\} \tag{13}$$

There are feasibility constraints that the stocks of married and unmarried agents of each gender and type cannot exceed the aggregate stocks of agents of each gender in the society:

$$f_j = \mu_{0j} + \sum_i \mu_{ij} \tag{14}$$

$$m_i = \mu_{i0} + \sum_j \mu_{ij} \tag{15}$$

We can now define a rational expectations equilibrium. There are two parts to the equilibrium, corresponding to the two stages at which decisions are made by the agents. The first corresponds to decisions made in the marriage market; the second to the intra-household allocation. In equilibrium, agents make marital status decisions optimally, the sharing rules clear each marriage market, and conditional on the sharing rules, agents choose consumption and labor supply optimally. Formally:

Definition 3 *A rational expectations equilibrium consists of a distribution of males and females across individual type, marital status, and type of marriage* $\{\hat{\mu}_{0j}, \hat{\mu}_{i0}, \hat{\mu}_{ij}\}$*, a set of decision rules for marriage, a set of decision rules for spousal consumption, leisure and public goods* $\{C^*_{ij}(\cdot), c^*_{ij}(\cdot), L^*_{ij}(\cdot), l^*_{ij}(\cdot), K^*_{ij}(\cdot)\}$*, and a matrix of husbands' powers* $\mathbf{p}^*$ *such that:*

1. *Marriage decisions solve (6) and (7), obtaining* $\{V^*(\underline{\varepsilon_j}), v^*(\underline{\epsilon_i})\}$.

2. *All marriage markets clear implying (13), (14), (15) hold;*

3. *For a marriage between a type i man and a type j woman, the decision rules* $\{C^*_{ij}(\cdot), c^*_{ij}, L^*_{ij}(\cdot), l_{ij}(\cdot), K^*_{ij}(\cdot)\}$ *solve (P1).*

Theorem 4 *A rational expectations equilibrium exists.*

Sketch of proof: We have already demonstrated (1) and (3). So what needs to be done is to show that there is a matrix of husbands' powers, $\mathbf{p}^*$ which clears the marriage market. Consider a matrix of admissible husband's powers $\mathbf{p}$. For every marriage $\{i, j\}$ excluding $i = 0$ or $j = 0$, define the excess demand function for marriages by men:

$$D_{ij}(\mathbf{p}) = \underline{\mu}_{ij}(\mathbf{p}) - \overline{\mu}_{ij}(\mathbf{p}) \tag{16}$$

The demand and supply functions, $\underline{\mu}_{ij}(\mathbf{p})$ and $\overline{\mu}_{ij}(\mathbf{p})$, for every type of marriage ij, satisfy the weak gross substitute property, (10) and (12). So the excess demand functions also satisfy the weak gross substitute property. Mas-Colell, Winston and Green (1995: p. 646, exercise 17.F.16^C) provide a proof of existence of market equilibrium when the excess demand functions satisfy the weak gross substitute property. For convenience, we reproduce their proof in our context in Appendix 3.[8] Our collective model of marriage matching shows that the transferable utilities model of the marriage market can be generalized to non-transferable utilities where the marginal utilities of consumption are not constant.[9]

2.4 The logit spousal choice model

The rest of the paper concerns some empirical implications of the above model. From here on, we will assume the logit random utility model, that ε_{ij} and ϵ_{ij} are *iid* extreme value random variables. In this case, McFadden (1974) showed that for every type of woman j, the relative demand for type i husbands is:

$$\ln \overline{\mu}_{ij} - \ln \mu_{0j} = (\Gamma_{ij} - \Gamma_{0j}) + \overline{Q}_{ij}(p_{ij}, Z) - \overline{Q}_{0j}(Z_{0j}) \ , \quad i = 1, .., I \tag{17}$$

[8] Kelso and Crawford (1982) were the first to use the gross substitute property to demonstrate existence in matching models.

[9] See also Legros and Newman (2007).

where $\overline{\mu}_{ij}$ is the number of $\{i, j\}$ marriages demanded by j type females and μ_{0j} is the number of type j females who choose to remain unmarried. Similarly, for every type of man i, the relative demand for type j wives is:

$$\ln \underline{\mu}_{ij} - \ln \mu_{i0} = (\gamma_{ij} - \gamma_{i0}) + \overline{q}_{ij}(p_{ij}, Z_{ij}) - \overline{q}_{i0}(Z_{i0}), \quad j = 1, .., J, \tag{18}$$

where $\underline{\mu}_{ij}$ is the number of $\{i, j\}$ marriages supplied by j type males and μ_{i0} is the number of type i males who choose to remain unmarried.

When the marriage market clears, $\overline{\mu}_{ij} = \underline{\mu}_{ij} = \mu_{ij}$. We can add (17) and (18) to get:

$$\begin{aligned} \ln \frac{\mu_{ij}}{\sqrt{\mu_{i0}\mu_{0j}}} &= \frac{(\Gamma_{ij} - \Gamma_{0j}) + \overline{Q}_{ij}(p_{ij}, Z_{ij}) - \overline{Q}_{0J}(Z_{0j})}{2} \\ &+ \frac{(\gamma_{ij} - \gamma_{i0}) + \overline{q}_{ij}(p_{ij}, Z_{ij}) - \overline{q}_{i0}(Z_{i0})}{2} \end{aligned} \tag{19}$$

CS calls the left hand side of (19), which is observed, the total gains to an $\{i, j\}$ marriage. Total gains is the average of the log odds of type i males marrying type j females relative to them not marrying, and of type j females marrying type i males relative to them not marrying. The right hand side of (19) is equal to the average of the spouses' expected payoffs from an $\{i, j\}$ marriage relative to not marrying. If the average log odds of marrying are high, then total gains to marriage are large which should be expected.

In CS, which assumes transferable utilities, a change in p_{ij} have exactly equal and opposite impact on the husband's and wife's utilities. Thus in CS, changes in p_{ij} do not affect total gains. Unlike CS, the total gains to marriage in (19) depends on p_{ij}, husband's power. In the case considered here, with diminishing marginal utilities of consumption, total gains depends on p_{ij}.

3 Implications of the theory for reduced form labor supply estimation

In our companion paper (hereafter CSS) we establish formal identification of the structural parameters from observations on family labor supplies and marriage matching patterns in at least two marriage markets. Here, we will use the theory to provide a solution to the problem of accouting for substitute sex ratios in estimating the reduced form impact of sex ratios on labor supplies for married couples.

Using market clearing, and subtracting relative supply, (17), from relative demand, (18):

$$T_{ij} = \ln \frac{\mu_{i0}}{\mu_{0j}} = (\Gamma_{ij} - \Gamma_{0j}) + \overline{Q}_{ij}(p_{ij}, Z_{ij}) - \overline{Q}_{0J}(Z_{0j}) \tag{20}$$
$$- ((\gamma_{ij} - \gamma_{i0}) + \overline{q}_{ij}(p_{ij}, Z_{ij}) - \overline{q}_{i0}(Z_{i0}))$$

where T_{ij} is the log of the ratio of the number of unmarried type i men to unmarried type j women. This measure of marriage market tightness, or the net gain of the wife relative to her husband, is used in our empirical analysis in place of the aggregate ratio of men to women. Market tightness for an $\{i, j\}$ match in a market s is determined by two components of the matching environment. The first component is the invariant gains to entering an $\{i, j\}$relative to remaining unmarried. The higher the invariant gains to marriage, the greater is market tightness for $\{i,\ j\}$ matches. The second component depends on the indirect utility derived from an $\{i, j\}$ match in society s relative to remaining unmarried. As the relative indirect utility from an $\{i, j\}$ marriage increases, tightness is predicted to increase.

Equation (20) is a fundamental equilibrium relationship in the marriage market, a direct implication of marriage market clearing. It is the basis of the empirical content of marriage matching on the collective model in this paper. It is important to emphasize two points. First, market tightness (T_{ij}) is a function of the husband's power (p_{ij}) which is determined in marriage market equilibrium will thus depend on conditions in every other type of match an individual could have entered. Thus, market tightness depends on the labor market and marriage market conditions for all other types of matches. Second, T_{ij} and p_{ij} are both endogenous variables and simultaneously determined, thus equation (20) is not a statement about the causal effect of T_{ij} on p_{ij}.

Then invert equation (20) to derive an expression for the husband's power:

$$p_{ij} = g_{ij}(T_{ij}, \Gamma_{ij}, \Gamma_{0j}, \gamma_{ij}, \gamma_{i0}, Z_{ij}, , Z_{0j}, Z_{i0}). \tag{21}$$

Equation (21), which is useful here, is not a reduced form equation because T_{ij}, an endogenous variable, is on the right hand side. CSS derives the reduce form equation for p_{ij} which we will use for estimation there.

We now use equation (21) to derive empirical implications of our theory regarding the effect of marriage matching on spousal labor supplies. Let H^s_{ij} be the hours of work of a wife of type j in an ij marriage in society s.

Recall from Section 2.1 the labor supply of a wife in an $\{i, j\}$ match is:

$$H^*_{ij}(p_{ij}, W_{ij}, w_{ij}, A_{ij}) \tag{22}$$

In what follows, we derive an empirical counterpart to equation (22), as well as the labor supply equation for husbands. To start, notice that the female's labor supply depends on four components, the husband's power (p_{ij}), the wages of both spouses (W_{ij}, w_{ij}) and the couple's total nonlabor income (A_{ij}). The latter three are observed in the data, but p_{ij} is not observed.

Substituting equation (21) into the labor supply functions yields the wife's labor supplies:

$$H^*_{ij}(g_{ij}(T_{ij}, \Gamma_{ij}, \Gamma_{0j}, \gamma_{ij}, \gamma_{i0}, Z_{ij}, , Z_{0j}, Z_{i0}), W_{ij}, w_{ij}, A_{ij})$$

Labor market determinants as well as match specific wages and assets affects the wife's labor supply. Since we do not have instruments for the idiosyncratic componetns of individual level wages and assets, we will average the labor supplies in society s to get a mean labor supply equation for type j women in $\{i, j\}$ matches in that society:

$$\overline{H}^s_{ij}(T^s_{ij}, \Gamma_{ij}, \Gamma_{0j}, \gamma_{ij}, \gamma_{i0}, Z^s_{ij}, , Z^s_{0j}, Z^s_{i0})) \tag{23}$$

Similarly, the mean labor supply equation for all type i husbands in $\{i, j\}$ matches in society s is:

$$\overline{h}^s_{ij}(T^s_{ij}, \Gamma_{ij}, \Gamma_{0j}, \gamma_{ij}, \gamma_{i0}, Z^s_{ij}, , Z^s_{0j}, Z^s_{i0}) \tag{24}$$

We can use variations in mean labor supplies and T^s_{ij}, Z^s_{ij}, Z^s_{i0} and Z^s_{0j} across societies to estimate (23) and (24).

Equations (23) and (24) do not include individual spousal wages or non-labor incomes as covariates. The theory implies that the individual labor supply responses to spousal wages should satisfy Slutsky symmetry. However, this restriction cannot be tested with mean labor supply by $\{i, j, s\}$, which is used here. We do not use individual level data because wages and non-labor income are measured with error and we do not have instruments for the idiosyncratic component of individual wages and non-labor income.

Systematic components cannot be used as instruments to test Slutsky symmetry because the systematic components are known at the time of marriage, and therefore affect husband's power p_{ij}^s, and are also collinear with Z_{ij}^s.

It is also convenient at this point to discuss empirical tests of the static collective model using spousal labor supplies such as CFL. In their paper, they estimate restricted individual spousal labor supplies models where the restrictions are derived from a static collective model. They instrument spousal wages, children, and non-labor income with education, age, father's education, city size and religion. Different values of these instruments define different types of individuals in different regions. There is no instrument which captures the transitory component of wages. Our interpretation of their empirical results is that they provide evidence of (1) efficient bargaining between different types of spouses and (2), spousal bargaining power depends on the type of marriage matches as we assume in this paper. Their empirical results are not informative about whether there is efficient risk sharing with the household as we suppose, or whether there is not as they supposed. In order to empirically distinguish between whether there is efficient risk sharing or not, one would need an instrument for transitory wage shocks when one estimates individual spousal labor supplies equations. As mentioned in Section 4.2, the results in this paper also do not shed light on whether there is efficient risk sharing or not.

A caveat on unobserved heterogeneity is necessary. If our observed matches do not accord with the matches as perceived by market participants, then the marriages in each observed match may contain mixtures of different unobserved marital matches. As labor market conditions and other exogenous variables change, the mix of unobserved marital matches used to construct observed market tightness and other variables may change. How these unobserved resorting affects our results is unclear. This problem is not unique to our paper. To the extent that changes in exogenous variables change the composition of observed sex ratios, this problem affects all work this area.[10]

[10] For example, Angrist uses the sex ratio of immigrants as his measure of substitutes which he argues was driven by immigration policy. Differences in immigration policies will change the quality of mix of immigrants.

4 Model Extensions

4.1 Married women's labor force participation rate

In addition to matching by individual types, a potential couple also has to choose what kind of household to form. For example, they can marry or cohabit. The spouses can choose participate in the labor force or not. Our framework extends to choice of marital technologies and we will use it to study the married women's labor force participation rate here.

Let x denote specialized marriages where only the husband works and n denote non-specialized marriages where both spouse work. Now a marriage match is denoted by the tuple $\{i, j, t\}$ where $t = \{n, x\}$. Noticing that the number of unmarrieds in the same society, μ_{i0} and μ_{0j}, are the same for both types of marriages, use (19) to get:

$$\ln \mu_{ij}^x - \ln \mu_{ij}^n = \frac{(\Gamma_{ij}^x - \Gamma_{ij}^n) + \overline{q}_{ij}^x(p_{ij}^x, Z) - \overline{q}_{ij}^n(p_{ij}^n, Z) + (\gamma_{ij}^x - \gamma_{ij}^n) + \overline{q}_{ij}^x(p_{ij}^x, Z) - \overline{q}_{ij}^n(p_{ij}^n, Z)}{2} \tag{25}$$

Then:

$$\frac{\partial \ln \frac{\mu_{ij}^x}{\mu_{ij}^n}}{\partial \zeta} \propto \frac{\partial((\Gamma_{ij}^x - \Gamma_{ij}^n) + (\gamma_{ij}^x - \gamma_{ij}^n))}{\partial \zeta} + \frac{\partial(\overline{q}_{ij}^x - \overline{q}_{ij}^n + \overline{q}_{ij}^x - \overline{q}_{ij}^n)}{\partial Z}\frac{\partial Z}{\partial \zeta} \tag{26}$$
$$+ (1 - p_{ij}^x)\frac{\partial \overline{q}_{ij}^x}{\partial p_{ij}^x}\frac{\partial p_{ij}^x}{\partial \zeta} - (1 - p_{ij}^n)\frac{\partial \overline{q}_{ij}^n}{\partial p_{ij}^n}\frac{\partial p_{ij}^n}{\partial \zeta}$$

Using (21),

$$\frac{\partial p_{ij}^t}{\partial \zeta} = \rho_{ij}^t \frac{\partial((\Gamma_{ij}^t - \Gamma_{0j}) - (\gamma_{ij}^t - \gamma_{i0}))}{\partial \zeta} + \rho_{ij}^t \frac{\partial((\overline{q}_{ij}^t - \overline{q}_{0j}) - (\overline{q}_{ij}^t - \overline{q}_{i0}))}{\partial Z}\frac{\partial Z}{\partial \zeta} \tag{27}$$
$$- \rho_{ij}^t \frac{\partial T_{ij}^t}{\partial \zeta};\ \rho_{ij}^t \equiv [(1 + p_{ij}^t)\frac{\partial \overline{q}_{ij}^t}{\partial p_{ij}^t}]^{-1} > 0$$

Substitute (27) into (26):

$$
\begin{aligned}
\frac{\partial \ln \frac{\mu_{ij}^x}{\mu_{ij}^n}}{\partial \zeta} &= \frac{\partial(\Gamma_{ij}^x + \gamma_{ij}^x)}{(1+p_{ij}^x)\partial \zeta} - \frac{\partial(\Gamma_{ij}^n + \gamma_{ij}^n)}{(1+p_{ij}^n)\partial \zeta} + \frac{(p_{ij}^x - p_{ij}^n)}{(1+p_{ij}^x)(1+p_{ij}^n)} \frac{\partial(\Gamma_{0j} - \gamma_{i0})}{\partial \zeta} \\
&+ \Big(\frac{\partial(\overline{q}_{ij}^x + \overline{q}_{ij}^x)}{(1+p_{ij}^x)\partial Z} - \frac{\partial(\overline{q}_{ij}^n + \overline{q}_{ij}^n)}{(1+p_{ij}^n)\partial Z} + \frac{(p_{ij}^x - p_{ij}^n)}{(1+p_{ij}^x)(1+p_{ij}^n)} \frac{\partial(\overline{q}_{0j} - \overline{q}_{i0})}{\partial Z}\Big) \frac{\partial Z}{\partial \zeta} \\
&+ \frac{(p_{ij}^x - p_{ij}^n)}{(1+p_{ij}^x)(1+p_{ij}^n)} \frac{\partial T_{ij}}{\partial \zeta} \qquad (28)
\end{aligned}
$$

The economic content of (28) is as follows. After controlling for the marital match $\{i,j\}$, changes in invariant gains, changes in labor market conditions across two different societies, if the change in the log odds of specialized versus non-specialized marriages is positively correlated with the change in tightness, $p_{ij}^x > p_{ij}^n$ and vice versa. Put another way, *ceteris paribus,* if an increase in tightness reduces the labor force participation rate of married women, then $p_{ij}^x > p_{ij}^n$.

Within a type of marriage, i.e. $\{i,j,t\}$, *ceteris paribus,* an increase in husband's power, p_{ij}^t, implies that the husband's expected utility is rising and the wife's expected utility is decreasing. But we cannot conclude that if $p_{ij}^x > p_{ij}^n$, wives in specialized marriages have less expected consumption or utility than their working counterparts.

Working without uncertainty, Blundell, et. al. (2007) also discusses the spousal labor force participation decision within the collective framework. Their model differs from ours in the following substantive way. Using our notation, they assume that an $\{i,j\}$ couple take p_{ij} as exogenous. Given $\{W_{ij}, w_{ij}, A_{ij}, p_{ij}(W_{ij}, w_{ij}, A_{ij})\}$, the couple efficiently chooses which of them should work and how much to work.[11]

Their participation model is nested in our setup. Section 2.1 assumes that optimal labor supplies are always interior. But depending on p_{ij}, the specification of preferences, and on the realizations of wages and non-labor income in the second period, some $\{i,j\}$ households will choose to be specialized and others will not. p_{ij} is the same for specialized versus non-specialized households under this scenerio. Without uncertainty, all $\{i,j\}$ households will choose one choice over the other.

[11] Since they ignore the marriage decision and there is no uncertainty in their model, p_{ij} is a function of W_{ij}, w_{ij}, A_{ij}.

However even without wages and assets income uncertainty, this section discussed an additional channel for specialization versus non-specialization which is not covered in their model. That is, some $\{i,j\}$ households will to choose to enter specialized marriages and others will not as part of their marital decisions. More importantly, p^x_{ij} and p^n_{ij}. are pinned down in the model as in equation (25).

4.2 One period marriage without uncertainty

Most of literature on the collective model deals with a static model of intrahousehold allocations without uncertainty. That is, wages and non-labor income are known as of the time the individuals enter into the marriage. Our marriage matching framework can accommodate this case and our structural labor supply paper, CSS, studies this case.

Let observed wages, non-labor income and labor supplies be equal to true wages, non-labor income and labor supplies plus measurement error:

$$\widetilde{W}_{ij} = W_{ij} + \varepsilon^{W\pi}_{ij} \tag{29}$$
$$\widetilde{w}_{ij} = w_{ij} + \varepsilon^{w\pi}_{ij} \tag{30}$$
$$\widetilde{A}_{ij} = A_{ij} + \varepsilon^{A\pi}_{ij} \tag{31}$$
$$\widetilde{H}_{ij} = H_{ij} + \varepsilon^{L\pi}_{ij} \tag{32}$$
$$\widetilde{h}_{ij} = h_{ij} + \varepsilon^{l\pi}_{ij} \tag{33}$$

where $\widetilde{X}_{ij}$ is the observed values of X_{ij}. $\varepsilon^{W\pi}_{ij}$, $\varepsilon^{w\pi}_{ij}$, $\varepsilon^{L\pi}_{ij}$, $\varepsilon^{l\pi}_{ij}$and $\varepsilon^{A\pi}_{ij}$ are measurment errors which are uncorrelated with the true values. Marriages are still identified by $\{i,j,\pi\}$. Thus we can still use p_{ij}, the husband's power, to clear the marriage market. Given p_{ij}, instead of problem P1, the planner will now solve:

$$\max_{\{C_{ij},c_{ij},H_{ij},h_{ij}\}} \widehat{Q}(C_{ij}, 1-H_{ij}, K_{ij}) + p_{ij}\widehat{q}(c_{ij}, 1-h_{ij}, K_{ij}) \tag{P1a}$$
$$\text{subject to } C_{ij} + c_{ij} + K_{ij} \leq A_{ij} + W_{ij}H_{ij} + w_{ij}h_{ij} \;\forall\; z_{ij}$$

(1), appropriately reinterpreted, continues to hold which is what is critical for marriage market clearing. Thus as long as we can identify the type of an individual and the marital matches that the individual can enter into, i.e. $\{i, j\}$, the empirical tests that we develop in this paper remain valid. Differences in observed spousal labor supplies across $\{i, j\}$ couples in the same society are interpreted as due to different realizations of measurement errors across these couples.

Thus the empirical results in this paper should be interpreted with care. *Even if our empirical results are consistent with our model predictions, they do not shed light on whether there is efficient risk sharing within the family or not.*

In our reduced form regressions, we do not include individual spousal wages as covariates. For every $\{i, j\}$ match, we observe labor income and labor supplies of multiple couples. Wages can be constructed by dividing labor income by hours of work. But measurement error in labor supplies and idiosyncratic labor supply shocks will induce variation in constructed wages as discussed above. Since risk sharing in marriage, measurement error in labor supplies, and idiosyncratic labor supply shocks are all salient factors in our data, and we do not have instruments for the idiosyncratic components of wages, we do not use constructed wages in our reduced form labor supply estimates. Consequently, we do not take a stand on how much risk sharing there is in our data.

Put in another light, the reduced form implications that we test in this paper are independent of whether there is risk sharing or not. Similarly, our results are also independent of whether there are public goods in marriage or not.

In CSS, we will estimate this one period model without uncertainty or public goods.

4.3 Home Production

In this section, we borrow results from BCM to extend our theoretical model to incorporate home production and a distinction between leisure time and time for work at home. Assume the household produces the public good K_{ij} at home using the production technology $G(\cdot)$ which is a function of purchased market inputs and time. The home production function is:

$$K_{ij} = G(X_{ij}, L_{ij}, l_{ij}),$$

where L_{ij} and l_{ij} are the hours of home production of the wife and husband, respectively. The efficient risk sharing arrangement in the collective model with home production is:

$$\max_{\{C,c,X_H,h,L,l\}} \mathbf{E}(\widehat{Q}(C_{ij}, 1 - H_{ij} - L_{ij}, K_{ij})|Z) \\ + p_{ij}\mathbf{E}(\widehat{q}(c_{ij}, 1 - h_{ij} - l_{ij}, K_{ij})|Z)$$

subject to the family budget constraint

$$X_{ij} + C_{ij} + c_{ij} \leq A_{ij} + W_{ij}H_{ij} + w_{ij}h_{ij}.$$

BCM establish identification of this model, including the function $G(X_{ij}, L_{ij}, l_{ij})$, under the assumption that data on time use is available. The reader is referred to BCM for full details. The solution to the model yields demand functions for private and public consumption expenditures, and labor supplies for home and market work, each a function of wages, nonlabor income and market tightness.

5 Empirical Analysis

In this section, we estimate the reduced form of our structural model relating marriage market tightness to spousal labor supplies in both market and home production. We investigate the empirical relevance of three issues highlighted by our theoretical model of marital matching and intra-household allocations: (i) the role of marriage market substitutes, (ii) the endogeneity of market tightness to labor market conditions, and (iii) heterogeneity in the marital production technology.

5.1 Data

The data used in our analysis comes from two sources: the 5% sample from the 2000 US Census and the 2003 American Time Use Survey (ATUS). The Census data is used to construct measures of sex ratios, marriage market tightness and labor market conditions in each marriage market and to estimate our reduced form labor supply regressions. The reduced form home production regressions are estimated using time use data from the ATUS.

We define an individual's type as a combination of race, age and education. For each gender, there are four contiguous age categories of 5 years each. The ages are slightly staggered across gender to reflect the fact that men tend to marry slightly younger women. The youngest female and male age categories, are 25-29 and 27-31 respectively. For each gender, we consider two schooling categories: high school graduates (at least 12 and up to and including 15 years of education) and college graduates (16 years of education and higher). For each race and gender, there are 8 potential types of individuals. Since we are only considering same race marriages, there are potentially $64 \times 3 = 254$ types of marital matches for each society.

We define each state as a separate society. With 50 states, there are potentially $254 \times 50 = 12,700$ cells across all marriage markets. However, the majority of these potential cells (marital match $\times$ state) have few or no marriages. To avoid thin cell problems, we delete a cell if the number of marriages in that cell is less than 5.[12] For most OLS regressions, we have about 2800 different cells (marital match $\times$ state), with about 180 distinct marital matches. For IV regressions, we have about 2400 observations. Most of the missing cells are due to non-white marriages, with large spousal age differences, in states with small populations. There are 750,000 same race couples in our Census sample before dropping the thin cell couples. After dropping the thin cell couples, about 3,000 couples, our base Census sample has approximately 747,000 couples. In other words, most of the thin cells that we dropped were empty cells. We also exclude mixed race couples to mitigate thin cells.[13]

Excluding thin cells from the empirical analysis should not affect the consistency of our estimates. Our labor supplies regressions, (23) and (24), have to hold for any subset of marital matches.

There is one selection criteria that is commonly imposed in the empirical collective labor supply literature that we do not impose here, at least for the labor supply regressions. Because we allow for public goods within marriage, we do not restrict our analysis of labor supply to childless individuals or couples. In contrast, we only consider childless couples in our home production model, as it is difficult to distinguish between home production and leisure for certain activities in households with children.

Market tightness for marital type $\{i, j\}$ in state s is defined as the log of the ratio of the number of unmarried type i males to the number of unmarried

[12] We have other minor selection rules.

[13] Also market tightness for mixed race couples which include white spouses are very different from own race couples because there are so many more whites than other races in the data. So we would need to have separate coefficients on tightness for each mixed race couples.

type j females in state s.[14]

Table 1 provides summary statistics for our sample. Average tightness across all marriages is -0.00278. Thus most couples choose marriage matches in which the numbers of unmarried men and women are approximately equal. The standard deviation is 0.542 which means that a significant number of marriage matches have quite uneven tightness.

We use five measures of sex ratios. The most refined measure is the sex ratio measured at the cell level (log of the ratio of the number of males of type i to the number of females of type j in state s). There are also sex ratios by education matches and state, age matches and state, and race and state. For all measures, across all individuals, mean sex ratios are approximately zero. The standard deviations are relatively large for marital matches narrowly defined and also by education.

The individual data shows that most couples choose partners in which mean tightness and sex ratio is approximately zero. However the standard deviation is significant. Thus there should be sufficient variation in marital matches to study their effect on spousal labor supplies.

Since we will analyze data at the cell level, Table 1 also shows, where an observation is a cell, mean T_{ij}^s is -0.129. The standard deviation has increased to 0.962. The increase in the standard deviation should be expected since each cell is an observation. Although most marriages occur around tightness equal to zero, a significant number of marriages occur elsewhere. When each cell is an observation, those marriages are responsible for the large standard deviation in tightness at the cell level. But there are relatively few marriages in the cells with tightness significantly different from zero. Similarly, the standard deviations of sex ratios by cell is also large relative to those by individuals.

At the cell level, about half the cells involve white marriages and the rest are black and hispanic marriages.[15] Mean ages by cell are between 35-39 for women and 37-41 for men. For both gender, there are slightly more high school graduates than college graduates.

To control for aggregate labor market conditions in an individual's local marriage market (i.e. Z_{ij}^s, Z_{i0}^s, Z_{0j}^s), we define the following three variables to characterize the earnings and non-labor income distributions. First, condi-

[14]An individual is unmarried if he or she is currently not married in the Census form (not code 1 or 2).

[15]Again this is the difference between individual observations, where 80% of the marriages are among white couples, and observations by cell.

tional on positive annual labor earnings for a type of unmarried individual, we construct the mean and standard deviation of log annual labor earnings for the distribution of unmarried individuals (wage and salary income). The second measure is the fraction of individuals with zero labor earnings for each match type in each marriage market. Finally, we construct the analogous variables for non-labor earnings, defined from the Census as total personal income minus wage and salary income.[16] We do not use wages or non-labor income for $\{i, j\}$ couples in state s to avoid issues of simultaneity. So we are assuming that the labor market and non-labor incomes of the unmarrieds in a state span the wages and non-labor income distributions for that state.

Table 1 also contains information on the labor supply behavior of married couples by cells in the 2000 Census. The labor force participation rates for husbands and wives are 92% and 75%, respectively. We consider three measures of labor supply and one measure of home production in our empirical analysis. Our first measure is the log odds of LFP (labor force participation). Conditional on participating in the labor force, our second measure of labor supply is the log of usual hours worked per week. Mean usual hours worked for men and women were 45 and 34 hours respectively. Conditional on being in the labor force, the third measure is log weeks worked per year. Mean weeks worked per year for men and women were 49 and 41 weeks respectively.

5.2 Determinants of market tightness

Table 2 presents summary estimates of market tightness (The point estimates are in Appendix T1). This is our empirical estimate of $T(\Gamma^s, \gamma^s, Z^s, M^s, F^s)$. There are a total of 2395 cells (state $\times$ marital match) and each cell is one observation. Column 1 regresses market tightness on state effects alone. The p-value of the F-statistic shows that there is no evidence that state effects alone explain market tightness. Column 2 adds our proxies for sex ratio at birth data to the regression. The R^2 jumps to 0.876. In otherwords, our sex ratios at birth is a strong predictor of market tightness. This also means that we should not have a weak instrument problem when we do IV regressions. Column 3 adds labor market conditions to the regression. From the F tests, both proxies for sex ratio at birth and labor market conditions separately continue to affect market tightness. Finally in column 4, we add individual

[16] To be precise, we measure the fraction of individuals with non-positive non-labor income rather than zero non-labor income.

fixed effects: age, education and race of each spouse. The individual fixed effects, proxies for sex ratio at birth and labor market conditions all continue to affect market tightness.

6 Substitution in marriage markets

As discussed in the introduction, substitution or spillover effects in the marriage market affects the estimation of causal effects of sex ratios on spousal labor supplies. Table 3 shows that this concern is a significant issue in our data.

Different columns in Table 3 presents summary statistics for reduced form spousal labor supplies regressions on the own sex ratio, SR^s_{ij}, sex ratios by age and state, sex ratios by education and state, and sex ratios by race and state, and other covariates. The first three columns estimate the reduced forms by OLS. The last three columns assumes that the sex ratios by state are endogenous and instruments the sex ratios with the birth rate instruments from Table 1.

In addition to sex ratios, column 1 includes state effects as covariates. The dependent variables include, for an $\{i, j, s\}$ marriage, the log odds of the spouse being in the labor market (log odds LFPR), log usual hours of work per week (ln hrs/wk), and log weeks worked per year (ln wks/yr). For each dependent variable and covariates, Table 3 presents the p-value of the F test that the coefficients of all the other sex ratios other than SR^s_{ij} is different from zero. For the first dependent variable, log odds LFP of the wife, the p-values in columns (1) to (6) are smaller than one percent. In other words, all the other sex ratios also affect the labor force participation rate of the wife. Thus it becomes difficult to interpret what the estimated coefficient on SR^s_{ij} in such a regression imply, particularly when all the different sex ratios are quite collinear. The estimated coefficient on SR^s_{ij} depends on what other sex ratios are included.

The next dependent variable is ln hrs/wk of the wife. Here, the p-values show that all the other sex ratios also affect hours of work per week at the one percent significance level for columns (1), (2), (4) and (5). The p-values are 0.06 in columns (3) and (6). In those two specifications there is some weak evidence that the other sex ratios do not matter.

Looking through the Table, 50 out of 60 specifications show that, after controlling for SR^s_{ij}, the other sex ratios also affect spousal labor supplies at the one percent significance level. In most of the cases where we cannot reject the hypothesis that the other sex ratios do not matter are in columns (3) and (6) which include individual fixed effects for both spouses. In these specifications, the joint confidence intervals are large and there is little power against the null hypothesis of no effect.

Thus the evidence in Table 3 strong suggests that substitution in marriage markets affect spousal labor supplies, and that the estimated effect of SR_{ij}^s on spousal labor supplies depends critically on what other sex ratios are included in the estimation. Ignoring the other sex ratios is not a satisfactory solution. It is difficult to interpret the estimated effect of SR_{ij}^s alone on spousal labor supplies when we know that substitution effects in the marriage market are quantitatively important.

7 Wives' labor supplies

Table 4a presents the estimated effects of T_{ij}^s on log odds of LFP of wives. Columns (1) to (3) present OLS estimates. Columns (4) to (6) present IV estimates.

In addition to T_{ij}^s, column (1) adds state effects as covariates. The estimated coefficient on T_{ij}^s is 0.232. It is statistically different from zero at the 1% level and has the "wrong sign". Recall that the theory says that the variation in T_{ij}^s is a proxy for husband's power after controlling for invariant gains and labor market conditions. Column (1) does not control for labor market conditions or invariant gains, except via state effects, nor does it account for the endogeneity of T_{ij}^s.

In column (2), we add state effects and labor market controls. In this case, the estimated effect on T_{ij}^s is -0.196 and it is also statistically significant at the 1% level. So controlling for labor market conditions significantly changes our estimated effect of T_{ij}^s and in accord with direction predicted by out theory. This switch in estimated sign between column (2) and column (1) is consistent with our other estimates. In almost all cases, controlling for state effects alone result in a wrong estimated sign for T_{ij}^s. After controlling for labor market conditions, the estimated sign for T_{ij}^s is consistent with the theory. This is true whether we instrument T_{ij}^s or not. So, consistent with our model, T_{ij}^s is correlated with labor market conditions. Ignoring labor market conditions in estimation will result in inconsistent estimates of the effects of T_{ij}^s on spousal labor supplies.

The change in estimated sign between columns (1) and (2) suggests that after controlling for state effects, supply shocks to the labor market dominate. An exogenous increase in the supply of type j females will reduce T^s_{ij}, and reduce their wage and therefore their labor supplies throught a substitution effect. This will induce a positive correlation between female labor supplies and T^s_{ij} if we do not control for the reduction in the wage which is what we have in column (1). After we add labor market controls in column (2), the estimated "wrong" sign disappears.

Column (3) adds state effects, labor market controls and individual fixed effects. The point estimate on T^s_{ij} is -0.184 and again statistically different from zero at the 1% level. It is marginally different from that in column (2).

Columns (4) to (6) are IV estimates. T^s_{ij} is affected by variations in invariant gains. Since we at best control for these variations with state and individual effects, there are variations in invariant gains that we do not control for. Instrumenting T^s_{ij} with sex ratios at birth should mitigate the endogenity problem. The point estimate of 0.330 in column (4), which only includes state effects as the other covariates, shows that instrumenting T^s_{ij} is not sufficient to obtain a "right" estimated sign. Column (5) adds state effects and labor market conditions. In this case, the estimated coefficient on T^s_{ij} is -0.176 and it is statistically different from zero. The IV point estimate in column (5) is slightly different from the OLS estimate in column (3) but we will not reject the hypothesis that they are the same using a Hausman test. Put another way, the endogeneity of T^s_{ij} is not a quantitatively significant issue for estimation. This will be a consistent finding in the rest of our results.

Column (6) adds state effects, labor market conditions and individual effects. The point estimate on T^s_{ij} is -0.167 and the standard error is 0.065. So the point estimate is similar to that without individual effects but the standard error doubles. This fall in estimated precision is the same for the other specifications.

Using the point estimate of -0.176 in column (5) as a benchmark, a one standard deviation increase in T^s_{ij} will decrease the log LFP of wives by 0.26 standard deviation. Thus variations in tightness is quantitatively important for explaining variations in the log LFP of wives across matchs and or societies.

A behavioral interpretation of our results is that, *ceteris paribus,* women and men are more willing to enter into marriages where the wife does not work when T^s_{ij} increases. Using our discussion of female labor force participation rate in Section 25, the negative effect of T_{ij}^s on log LFP of wives says that *ceteris paribus,* husbands' power is larger in households where their wives do not work compared to households where their wives work.

By and large, the OLS estimates in columns (1) to (3) are similar to their IV counterparts in columns (4) to (6). This suggests that endogeneity of T_{ij}^s due to labor demand shocks are not of first order importance. Recall that we include state effects to control for state level labor demand shocks.

Table 4b presents estimates of the effect of T_{ij}^s on ln hrs/week of wives. Column (1) adds state effects. The point estimate on T_{ij}^s has the wrong sign and is statistically different from zero at the 5% level. Column (2) adds state effects and labor market conditions. The point estimate on T_{ij}^s is -0.026 and the estimated standard error is 0.004. So as in Table 4a, adding labor market conditions changes the estimated sign on T_{ij}^s. The estimated magnitude is smaller than for log odds LFP. Column (3) adds state effects, labor market conditions and individual effects. The point estimate on T_{ij}^s is -0.024 and the estimated standard error is 0.005. So adding individual effects do not changed the estimated effect on T_{ij}^s.

Columns (4) to (6) are IV estimates. As before, state effects alone in Column (4) results in a wrong estimated sign. Column (5) adds state effects and labor market conditions. The point estimate on T_{ij}^s is -0.028 and the estimated standard error is 0.005. This estimate is similar to the OLS estimate. Column (6) adds state effects, labor market conditions, individual effects. The point estimate on T_{ij}^s is -0.044 and the standard error is 0.010. Thus adding individual effects lower the estimated precision on T_{ij}^s. The estimates in Table 4b is qualitatively similar to their counterparts in Table 4a. The estimated magnitudes on ln hrs/wk are smaller than for participation.

Using the estimate in column (5) as a benchmark, a one standard deviation increase in T_{ij}^s results in 0.065 standard deviation decrease in ln hrs/wk. So variation in tightness explains less of the variation in mean log usual hours of work per week across matches and or societies.

Table 4c presents estimates of the effect of T_{ij}^s on ln wks/yr of wives. Column (1) adds state effects. The point estimate on T_{ij}^s has the wrong sign and is statistically different from zero at the 5% level. Column (2) adds state effects and labor market conditions. The point estimate on T_{ij}^s is -0.023 and the estimated standard error is 0.005. So as in Table 4a and 4b, adding labor market conditions changes the estimated sign on T_{ij}^s. Column (3) adds state effects, labor market conditions and individual effects. The point estimate on T_{ij}^s is -0.016 and the estimated standard error is 0.007.

Columns (4) to (6) are IV estimates. As before, state effects alone in Column (4) results in a wrong estimated sign. Column (5) adds state effects and labor market conditions. The point estimate on T^s_{ij} is -0.022 and the estimated standard error is 0.007. This estimate is similar to the OLS estimate. Column (6) adds state effects, labor market conditions, individual effects. The point estimate on T^s_{ij} is -0.014 and the standard error is 0.014. Thus adding individual effects lower the estimated precision on T^s_{ij} until it is no longer statistically significant at the 5% level. The point estimates in Table 4c are quantitatively similar to their counterparts in Table 4b. The estimated standard errors are larger.

Using the estimate in column (5) as a benchmark, a one standard deviation increase in T^s_{ij} results in 0.038 standard deviation decrease in ln wks/yr. So variation in tightness explains even less of the variation in mean log weeks worked per year of wives across matches and or societies.

For all three measures of labor supplies, it is important to control for labor market conditions to obtain consistent estimates of the effect of T^s_{ij} on the wives' labor supplies. The IV point estimates are not significantly different from OLS estimates. Although individual effects explain labor supplies, including them do not affect the estimated effects of T^s_{ij}.

Comparing results for the three different measures of wives' labor supplies, the qualitative impact of changes in tightness on spousal labor supplies are the same for all three measures. The quantitative magnitude is largest for participation and lowest for weeks worked per year, with hours per week in between. Due to changes in tightness, wives adjust their hours per week more than weeks worked per year. The relative large quantitative effect on participation suggests that not working is not simply more leisure. It is consistent with our view that participation is a choice between specialized and non-specialized home production and not just a redistribution of resources with a fixed marital technology.

If we use SR^s_{ij} instead of T^s_{ij} as a covariate in the regressions in Table 4, the point estimates on SR^s_{ij} are quantitatively similar to that for T^s_{ij} in their respective specifications. Thus empirically, the estimated effects of SR^s_{ij} on spousal labor supplies are similar to that of T^s_{ij}. However the interpretations are very different. When we use T^s_{ij} as a covariate and estimate its effect consistently, we recognize that it is an endogenous variable. We can calculate the effect of an exogenous variable on spousal labor supplies after calculating its effect on tightness. When SR^s_{ij} is used as a covariate, the standard practice is to use the point estimate as a causal estimate of the effect of SR^s_{ij} on spousal labor supplies which is incorrect. As Table 3 points out, the estimated effect depends on what other sex ratios are included.

8 Husbands' labor supplies

Table 5a presents the estimated effects of T^s_{ij} on log odds of LFP of husbands. Columns (1) to (3) present OLS estimates. Columns (4) to (6) present IV estimates.

In addition to T^s_{ij}, column (1) adds state effects as covariates. The estimated coefficient on T^s_{ij} is -0.018. It is not statistically different from zero at the 5% level.

In column (2), we add state effects and labor market controls. In this case, the estimated effect on T^s_{ij} is 0.067 and it is statistically significant at the 5% level. So controlling for labor market conditions significantly changes our estimated effect of T^s_{ij} in accord with direction predicted by out theory. This switch in estimated sign between column (2) and column (1) is consistent with our other estimates. In almost all cases, controlling for state effects alone result in a wrong estimated sign for T^s_{ij}. The argument is the same as in the wives' case. Supply shocks dominate in the labor market after we control for state level labor market conditions.

Column (3) adds state effects, labor market controls and individual fixed effects. The point estimate on T^s_{ij} is -0.015 and it is not statistically different from zero at the 5% level.

Columns (4) to (6) are IV estimates. The point estimate of T^s_{ij} is -0.034 in column (4) which only includes state effects. Column (5) adds state effects and labor market conditions. In this case, the estimated coefficient on T^s_{ij} is 0.145 and it is statistically different from zero at the 1% level. The IV point estimate in column (5) is significantly different from the OLS estimate in column (3). So there is some evidence that the endogeneity of T^s_{ij} is a quantitatively significant issue for estimation for husbands' labor supplies. This will be a consistent finding in the rest of our results for husbands' labor supplies.

Column (6) adds state effects, labor market conditions and individual effects. The point estimate on T^s_{ij} is 0.015 and the standard error is 0.072. The point estimate is smaller than in column (5) and the standard error is significantly larger. This fall in estimated precision is the same for the other specifications.

Using the point estimate of 0.145 in column (5) as a benchmark, a one standard deviation increase in T_{ij}^s will decrease the log LFP of husbands by 0.53 standard deviation. Thus variations in tightness has a larger impact on the log LFP of husbands than wives. This larger impact by standard deviation is driven by the smaller standard deviation in log LFP of husbands across cells. The quantitative estimate for husbands is actually smaller than that for wives.

Table 5b presents estimates of the effect of T_{ij}^s on ln hrs/week of husbands. Column (1) adds state effects. The point estimate on T_{ij}^s is not statistically different from zero at the 5% level. Column (2) adds state effects and labor market conditions. The point estimate on T_{ij}^s is -0.012 and the estimated standard error is 0.003. So as in Table 5a, adding labor market conditions changes the estimated sign on T_{ij}^s. Column (3) adds state effects, labor market conditions and individual effects. The point estimate on T_{ij}^s is 0.004 and it is not statistically different from zero at the 5% level.

Columns (4) to (6) are IV estimates. As before, state effects alone in Column (4) results in a estimate that is statistically not different from zero. Column (5) adds state effects and labor market conditions. The point estimate on T_{ij}^s is 0.018 and the estimated standard error is 0.003. This estimate is larger than the OLS estimate. Column (6) adds state effects, labor market conditions, individual effects. The point estimate on T_{ij}^s is not statistically different from zero at the 5% level. Thus adding individual effects lower the estimated precision on T_{ij}^s. The estimated elasticities on hrs/wk are smaller than for participation.

Using the estimate in column (5) as a benchmark, a one standard deviation increase in T_{ij}^s results in 0.068 standard deviation increase in ln hrs/wk. So variation in tightness explains less of the variation in mean log usual hours of work per week of husbands than participation.

Table 5c presents estimates of the effect of T_{ij}^s on ln wks/yr of husbands. Column (1) adds state effects. The point estimate on T_{ij}^s is not statistically different from zero at the 5% level. Column (2) adds state effects and labor market conditions. The point estimate on T_{ij}^s is again not statistically different from zero at the 5% level. Column (3) adds state effects, labor market conditions and individual effects. The point estimate on T_{ij}^s is also not statistically different from zero at the 5% level.

Columns (4) to (6) are IV estimates. As before, state effects alone in Column (4) generates an estimate that is not statistically different from zero at the 5% level. Column (5) adds state effects and labor market conditions.

The point estimate on T_{ij}^s is 0.011 and the estimated standard error is 0.005. This is the only specification which is statistically different from zero at the 5% level. Column (6) adds state effects, labor market conditions, individual effects. The point estimate on T_{ij}^s is not different from zero at the 5% level.

Using the estimate in column (5) as a benchmark, a one standard deviation increase in T_{ij}^s results in 0.038 standard deviation increase in ln wks/yr. So variation in tightness explains less of the variation in mean log weeks worked per year of husbands than wks/yr or participation.

For all three measures of labor supplies, it is important to control for labor market conditions to obtain consistent estimates of the effect of T_{ij}^s on the husbands' labor supplies. Unlike the estimates for wives, the IV point estimates for husbands' are significantly different from the OLS estimates. Thus controlling for endogeneity is important to obtain consistent estimates for husbands. With T_{ij}^s as a covariate, there is no gain to adding SR_{ij}^s as a covariate.

Similar to that for wives, the quantitative magnitude is largest for participation and lowest for weeks worked per year, with hours per week in between.

In general, estimates for husbands are less precise than that for wives which may be due to the smaller variation in husbands' labor supplies. Still for all three labor supplies measures, we obtain the correct sign and point estimates that are statistically different from zero at the 1% and 5% level in the estimates in column (5). This finding compares favorably with other studies on the effect of marriage market conditions on males' labor supplies. In general, they obtain imprecise results.

8.1 Robustness

To deal with the variable number of individual observations used to construct mean values for different cells, we also estimated the models by weighing the data by marriage counts for each cell. The weighted results are similar to that without weighting and therefore are not presented. There was no systematic increase in the precision of the weighted estimates.

We also experimented with allowing the effects of tightness on spousal labor supplies to differ by $\{i, j\}$. Unfortunately, the proliferation of interaction effects led to imprecise and uninterpretable empirical estimates. So although the theory and empirical framework allows for heterogenous treatment responses, we are unable to usefully estimate these effects with the current data set.

We also experimented by dividing the sample into two samples, marriages with kids and marriages without kids. The number of observerations for each cell fell substantially, we lost more cells in the estimation, and the estimates because substantially more imprecise for both samples.

In earlier drafts of this paper, we estimated the model using the individual observations and clustering the standard errors by $\{i, j, s\}$. The estimates were substantially more imprecise.

9 Time use results

Our time use data comes from the 2003 American Time Use Survey. While the population surveyed is consistent with the 2000 census, the time use survey has much less observations. Compared with our census sample of 746,963 individual observations, we only have 408 individual observations in the time use survey. Table 6 compares the two surveys.

Our measure of home production, presented in Table 6 is the "total" non-market work definition of Aguiar and Hurst (2007), minus shopping activities (obtaining goods and services). In particular, an individual's hours supplied to home production is defined as the total time spent on meals (preparation, presentation, and cleanup), housework (interior cleaning, laundry, sewing, repairing and maintaining textiles, storing interior household items including food) and interior and exterior maintenance, repair, and decoration, vehicle repair and maintenance, and appliance and tool set-up, repair, and maintenance, household management (except mail and email), and lawn, garden, houseplant and pet care. Our measure of home production does not include time spent obtaining goods as services, as it is arguably more difficult to disentangle home production time and leisure time for both of these categories. For similar reasons, we limit our analysis to couples with no children in the household to abstract from decisions regarding time spent with children decisions. On average, husbands supply 10 hours per week to home production while wives supply 16 hours.

Table 7 contains regression results with the time use data. Due to the small number of observations, we did not include state fixed effects. We include labor market controls and race fixed effects. Consider first the estimated effect of tightness on log hours of leisure per week of wives. Whether we use sex ratio as a covariate, OLS or IV with tightness, the estimated signs are positive but statistically insignificant due to the small number of observations in the regressions. The positive signs are consistent with our theory. The point estimates are implausibly large relative to those from the equivalent labor supplies estimates.[17]

Row 2 presents estimates of the effect of tightness on the log hours of home production per week of wives. The three estimates have negative signs but again are not statistically significant. Again the negative signs are consistent with our theory. The point estimates in this case are more plausible.

Turning to the results for husbands, row 3 presents estimates of the effect of tightness on the log hours of leisure per week of husbands. Two estimates have negative signs and one is a small positive estimate. The large standard errors preclude any definite conclusion.

Row 4 presents estimates of the effect of tightness on the log hours of home production per week for husbands. All three estimates are positive. The standard errors remain large but smaller than in row 3. These positive estimates are consistent with the theory.

Due to the small sample sizes, our empirical foray using time use data is at best suggestive. Most of the estimated signs are consistent with the theory. Thus the time use results adds to the evidence that marriage market conditions affect intrahousehold allocation of time use. But the large standard errors preclude any definite conclusion about the magnitudes of the effects.

10 Conclusion

This paper presents a flexible empirical framework which intergrates the collective model of intrahousehold allocation with the CS model of marriage matching. We use this framework to show how one can deal with the problem of including substitute sex ratios in estimating the effects of marriage market conditions on spousal labor supplies. Finally, using data from the 2000 US census, the paper presents empirical evidence that substitute sex ratios do affect spousal labor supplies, and how marriage market tightness affects intrahousehold allocation of spousal labor supplies and time use which are consistent with the theory.

[17] Even ignoring the small sample size, the results are not exactly comparable because these time use estimates do not control for state effects.

There is a lot of work which remains to be done. CSS estimates and tests the CFL version of the model discussed here. How to empirically estimate and test more general collective models in our framework needs to be done. Extending the model to incorportate pre-marital investments is also important (E.g. Chiappori, Iygun and Weiss (forthcoming); Peters and Siow (2002)). Lifecycle considerations are also important (E.g. Choo and Siow (2005)). The collective model is a complete contracting model of intrahousehold allocation. There is substantial evidence that in certain contexts, the complete contracting assumption is implausible. How we should extend this framework to those contexts is an important avenue for further research.

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A Proof of Proposition 1

Abstracting from i, j, g, G, the social planner solves:

$$\max_{\{C,c,H,h,K\}} \mathbf{E}(\widehat{Q}(C, 1-H, K|Z) + p\mathbf{E}(\widehat{q}(c, 1-h, K)|Z)$$

subject to, for each state S,

$$c + C + K \leq A + WH + wh \tag{34}$$

Let Z^* be the value of Z evaluated at the optimum. The first order conditions with respect to c, C, H, h, K and the multiplier λ for each state S are:

$$\widehat{Q}^*_C = \lambda \tag{35}$$
$$p\widehat{q}^*_c = \lambda \tag{36}$$
$$\widehat{Q}^*_{1-H} = \lambda W \tag{37}$$
$$p\widehat{q}^*_{1-h} = \lambda w \tag{38}$$
$$\widehat{Q}^*_K + p\widehat{q}^*_K = \lambda \tag{39}$$

Using the first order conditions, as p changes, for each state S,

$$\frac{\partial \widehat{Q}^*}{\partial p} = \lambda(C^*_p - WH^*_p + K^*_p) - p\widehat{q}^*_K K^*_p \tag{40}$$
$$\frac{\partial \widehat{q}^*}{\partial p} = \frac{\lambda}{p}(c^*_p - wh^*_p) + \widehat{q}^*_K K^*_p \tag{41}$$

which imply:

$$\frac{1}{p}\frac{\partial \widehat{Q}^*}{\partial p} + \frac{\partial \widehat{q}^*}{\partial p} = \frac{\lambda}{p}(c^*_p - wh^*_p + K^*_p + C^*_p - WH^*_p) \tag{42}$$

Since the budget constraint has to hold for every S,

$$c^*_p + C^*_p + K^*_p - wh^*_p - WH^*_p = 0$$
$$\Rightarrow \frac{\partial \widehat{Q}^*}{\partial p} = -p\frac{\partial \widehat{q}^*}{\partial p} \tag{43}$$

Since (43) holds for every state S, (1) obtains.

B Proof of $\frac{\partial \mathbf{E} H^*}{\partial p_{ij}} > 0$ and $\frac{\partial \mathbf{E} h^*}{\partial p_{ij}} < 0$

For an $\{i, j\}$ family, given realizations of wages and asset income, and taking p as given, the planner solves a one period household maximization problem, P2. The objective of this appendix is to show that for any admissible realization of wages and asset income, and taking p as fixed, labor supply of the wife will increase and labor supply of the husband will decrease as p increases. Assuming that realized wages and asset income are W, w and A, the planner's problem is:

$$\max_{C,L,c,l,K} \widehat{Q}(\Omega(C,L),K) + p\widehat{q}(\omega(c,l),K) \tag{44}$$

$$s.t.\ c + C + K + WL + wl \leq A + W + w = I \tag{45}$$

Given the weak separability between private goods and the public good in each spouse's utility function, let Y and y be the expenditure on the wife's and husband's private goods respectively. Then the wife will solve:

$$\max_{C,L} \widehat{Q}(\Omega(C,L),K) \tag{46}$$

$$s.t.\ C + WL \leq Y \tag{47}$$

Due to the weak separability, the optimal levels of private goods, C and L, only depend on W and Y, and are independent of K. We will assume that the optimal level of L is increasing in Y. The standard restriction on $\Omega(C,L)$, i.e. concavity and $\Omega_{LL} - \Omega_{CL} < 0$, that is leisure increases as Y increases, is sufficient. Solving (46) will result in an indirect utility:

$$\widetilde{Q}(Y,K) \tag{48}$$

The husband will solve:

$$\max_{c,l} \widehat{q}(\omega(c,l),K) \tag{49}$$

$$s.t.\ c + wl \leq y \tag{50}$$

Again, the optimal levels of private goods, c and l, only depend on w and y, and are independent of K. We will assume that the optimal level of l is increasing in y. The standard restriction on $\omega(c,l)$, i.e. concavity and

$\omega_{ll} - \omega_{cl} < 0$, that is leisure increases as y increases, is sufficient. Solving (49) will result in an indirect utility:

$$\widetilde{q}(y, K) \tag{51}$$

All the above implications of (44) and (45) are known from BCM. Assume that $\widetilde{q}(y, K)$ is increasing and quasi-concave, and $\widetilde{q}_{yK} > 0$. So we can rewrite the planner's problem as:

$$\max_{Y,y,K} \widetilde{Q}(Y, K) + p\widetilde{q}(y, K) \tag{52}$$

$$s.t.\ Y + y + K \leq I \tag{53}$$

Let $\mathbf{Y} = -Y$. Then the planner's problem, (52) and (53), can be rewritten as:

$$\max_{\mathbf{Y},y} R(\mathbf{Y}, y, p) = \widetilde{Q}(-\mathbf{Y}, I - y + \mathbf{Y}) + p\widetilde{q}(y, I - y + \mathbf{Y}) \tag{54}$$

$R(\mathbf{Y}, y, p)$ is supermodular in $\mathbf{Y}, y, K$ and p if:

$$R_{\mathbf{Y}y} = \widetilde{Q}_{YK} - \widetilde{Q}_{KK} + p(\widetilde{q}_{Ky} - \widetilde{q}_{KK}) > 0 \tag{55}$$

$$R_{\mathbf{Y}p} = \widetilde{q}_K > 0 \tag{56}$$

$$R_{yp} = \widetilde{q}_y - \widetilde{q}_K > 0 \tag{57}$$

The first order condition to the planner's problem is:

$$-\widetilde{Q}_Y + \widetilde{Q}_K + p\widetilde{q}_K = 0 \tag{58}$$

$$-\widetilde{Q}_K + p(\widetilde{q}_y - \widetilde{q}_K) = 0 \tag{59}$$

(59) implies (57).
(55) and (56) are implied by the assumption that $\widetilde{Q}(Y, K)$ is increasing in both arguments and quasi-concave in K, and $\widetilde{Q}_{YK} > 0$. An economically meaningful interpretation is that K is a normal good. In terms of the planner's primitive objective function (44), a sufficient condition is $\widehat{Q}(\Omega(C, L), K) + p\widehat{q}(\omega(c, l), K) = \Omega(C, L)\widehat{\Omega}(K) + p\omega(c, l)\widehat{\omega}(K)$ for increasing concave functions $\widehat{\Omega}$ and $\widehat{\omega}$. Since $R(\mathbf{Y}, y, p)$ is supermodular, using the monotone theorem of Milgrom and Shannon (1994), $\mathbf{Y}$ and y are both increasing in p, and thus Y is decreasing in p. Since L and l are increasing in Y and y respectively, L will decrease and l will increase as p increases. Thus H and h are increasing and decreasing in p respectively. See BCM for other implications of the weakly separable collective model of spousal labor supplies with public goods.

B.1 Cobb-Douglas preferences

Let the preferences of the husband and the wife be:

$$\widehat{q}(c,l,K) = l^{\alpha_h} c^{1-\alpha_h} K^{\delta_h} \tag{60}$$

$$\widehat{Q}(C,L,K) = L^{\alpha_f} C^{1-\alpha_f} K^{\delta_f} \tag{61}$$

Then:

$$\omega(c,l) = l^{\alpha_h} c^{1-\alpha_h} \tag{62}$$

$$\widehat{\omega}(K) = K^{\delta_h} \tag{63}$$

$$\Omega(C,L) = L^{\alpha_f} C^{1-\alpha_f} \tag{64}$$

$$\widehat{\Omega}(K) = K^{\delta_f} \tag{65}$$

Given y and Y, optimal leisure will satisfy:

$$l^* = \frac{\alpha_h y}{w} \tag{66}$$

$$L^* = \frac{\alpha_f Y}{W} \tag{67}$$

l^* and L^* are increasing in y and Y respectively as required. The indirect utilities are:

$$\widetilde{Q}(Y,K) = \alpha_f Y K^{\delta_f} \tag{68}$$

$$\widetilde{q}(y,K) = \alpha_h y K^{\delta_h} \tag{69}$$

for positive constants α_f and α_h. $R(\mathbf{Y}, y, p)$ is supermodular as required. Thus l^* will increase and L^* will decrease as p increases.

C Proof of Existence of Equilibrium

In the proof, we need:

$$E_{ij}(\underline{p}) > 0 \text{ as } \underline{p} \to \infty \qquad \text{(Condition A1)}$$

$$E_{ij}(\underline{p}) < 0 \text{ as } \underline{p} \to 0 \qquad \text{(Condition A2)}$$

That is, the utility functions q and Q must be such that as $\underline{p}$ approaches 0, men will not want to marry. And as $\underline{p}$ approaches ∞, women will not want to marry. Let $\beta_{ij} = (1 + p_{ij})^{-1}$ where $\beta_{ij} \in [0, 1]$ is the utility weight of the wife in an $\{i, j\}$ marriage and $(1 - \beta_{ij})$ is the utility weight of the husband. We know:

$$\frac{\partial \underline{\mu}_{ij}}{\partial p_{ij}} > 0 \tag{70}$$

$$\frac{\partial \underline{\mu}_{ij}}{\partial p_{ik}} < 0,\ k \neq j \tag{71}$$

$$\frac{\partial \underline{\mu}_{kl}(\beta)}{\partial p_{ij}} = 0;\ k \neq i, l \neq j \tag{72}$$

$$\frac{\partial \overline{\mu}_{ij}}{\partial p_{ij}} < 0 \tag{73}$$

$$\frac{\partial \overline{\mu}_{ij}}{\partial p_{kj}} > 0,\ k \neq i \tag{74}$$

$$\frac{\partial \overline{\mu}_{kl}(\beta)}{\partial p_{ij}} = 0;\ k \neq i, l \neq j \tag{75}$$

Let β be a matrix with typical element β_{ij} and the IxJ matrix function $E(\beta)$ be:

$$E(\beta) = \underline{\mu}(\beta) - \overline{\mu}(\beta) \tag{76}$$

An element of $E(\beta)$, $E_{ij}(\beta)$, is the excess demand for j type wives by i type men given β. An equilibrium exists if there is a β^* such that $E(\beta^*) = 0$. Assume that there exists a function $f(\beta) = \alpha E(\beta) + \beta$, $\alpha > 0$ which maps $[0, 1]^{I*J} \to [0, 1]^{I*J}$ and is non-decreasing in β. Tarsky's fixed point theorem says if a function $f(\beta)$ maps $[0, k]^N \to [0, k]^N$, $k > 0$, and is non-decreasing in β, there exists $\beta^* \in [0, k]^N$ such that $\beta^* = f(\beta^*)$. Let $f(\beta) = \alpha E(\beta) + \beta$, $k = 1$ and $N = I * J$, and apply Tarsky's theorem to get $\beta^* = \alpha E(\beta^*) + \beta^* \Rightarrow E(\beta^*) = 0$. Thus the proof of existence reduces to showing $f(\beta)$ which has the required properties. We know from (70) to (75) that:

$$\frac{\partial E_{ij}(\beta)}{\partial \beta_{ij}} < 0 \tag{77}$$

$$\frac{\partial E_{ik}(\beta)}{\partial \beta_{ij}} > 0 \tag{78}$$

$$\frac{\partial E_{kj}(\beta)}{\partial \beta_{ij}} > 0 \tag{79}$$

$$\frac{\partial E_{kl}(\beta)}{\partial \beta_{ij}} = 0;\ k \neq i, l \neq j \tag{80}$$

(77) to (80) imply that $E(\beta)$ satisfies the Weak Gross Substitutability (WGS) assumption. We now show that the WGS property of $E(\beta)$ implies that we can construct $f(\beta)$, such that $f(\beta)$ maps $[0,1]^{I*J} \to [0,1]^{I*J}$ and is non-decreasing in β. The proof follows the solution to exercise 17.F.16^C of Mas-Colell, Whinston and Green given in their solution manual (Hara, Segal and Tadelis, 1996). N.B. Unlike them, we do not start with Gross Substitution, we begin from WGS, but it turns out to be sufficient for Tarsky's conditions. For notational convenience, now onwards we'll treat the matrix function $E(\beta)$, as a vector function. Let $N = I * J$ and 1_N be a $N \times 1$ vector of ones. $E(\beta) : [0,1]^N \to R^N$ is continuously differentiable and satisfies $E(0_N) >> 0_N$ and $E(1_N) << 0_N$ (Conditions A1 and A2). For every $\beta \in [0,1]^N$ and any n, if $\beta_n = 0$, then $E_n(\beta) > 0$. For every $\beta \in [0,1]^N$ and any n, if $\beta_n = 1$, then $E_n(\beta) < 0$. If $\beta = \{0_N, 1_N\}$, the facts follow from Conditions A1 and A2. Otherwise, they are due to Conditions A1 and A2, and (77) to (80), i.e. WGS. For each n, define $C_n = \{\beta \in [0,1]^N : E_n(\beta) \geq 0\}$ and $D_n = \{\beta \in [0,1]^N : E_n(\beta) \leq 0\}$. Then $C_n \subset \{\beta \in [0,1]^N : \beta_n < 1\}$ and $D_n \subset \{\beta \in [0,1]^N : \beta_n > 0\}$. Then by continuity, the following two minima, ${}_{ij}((1-\beta_n)/E_n(\beta) : \beta \in C_n)$ and ${}_{ij}(-\beta_n/E_n(\beta) : \beta \in D_n)$, exist and are positive. Let $\underline{\beta}_n > 0$ be smaller than those two minima. Then, for all $\alpha \in (0, \underline{\beta}_n)$ and any $\beta \in [0,1]^N$, we have $0 \leq \alpha E_n(\beta) + \beta_n \leq 1$. For each n, define $L_n =_{ij} \{|\partial E_n(\beta)/\partial \beta_n| : \beta \in [0,1]^N\}$. Then, for all $\alpha \in (0, 1/L_n)$,

$$\frac{\partial(\alpha E_n(\beta) + \beta_n)}{\partial \beta_n} = \alpha \frac{\partial E_n(\beta)}{\partial \beta_n} + 1 \geq -\alpha L_n + 1 > 0$$

$$\frac{\partial(\alpha E_n(\beta) + \beta_n)}{\partial \beta_m} = \alpha \frac{\partial E_n(\beta)}{\partial \beta_m} \geq 0; n \neq m, \text{ follows from (77) to (80).}$$

Now let $K = {}_{ij}\{\underline{\beta}_1, .., \underline{\beta}_N, 1/L_1, .., 1/L_N\}$, choose $\alpha \in (0, K)$, then $f(\beta) = \alpha E(\beta) + \beta \in [0,1]^N$ and $\partial f(\beta)/\partial \beta_n \geq 0$ for every $\beta \in [0,1]^N$, and any n. Hence Tarsky's conditions are satisfied.

Table 1: Summary statistics (2000 US census)

Variable	Obs	Mean	Std. Dev.	Min	Max
INDIVIDUAL					
T(i,j,s)	746943	-.0027757	.5425194	-3.60296	3.03591
SR(i,j,s)	746943	-.003512	.4543587	-3.017524	2.338186
SR(age,s)	746943	-.0277066	.0823402	-.4369596	.3772942
SR(race,s)	746943	-.0037428	.0799888	-.4177496	.5132542
SR(educ,s)	746943	-.0237409	.4472689	-1.593234	1.491079
SR(s)	746943	-.0246046	.0460074	-.1440098	.173324
CELL AVERAGE					
T(i,j,s)	2998	-.1293465	.9615514	-3.416714	2.48425
SR(i,j,s)	2998	-.0779628	.8426416	-2.847724	2.157963
SR(educ,s)	2998	-.0332998	.6369095	-1.593234	1.491079
SR(race,s)	2998	-.0558353	.134684	-.4177496	.5132542
SR(age,s)	2998	-.0211743	.1003443	-.4369596	.3426507
SR(s)	2998	-.0253774	.0542566	-.1440098	.173324
Black	2998	.2731821	.4456679	0	1
White	2998	.5376918	.4986605	0	1
Hispanic	2998	.1891261	.3916741	0	1
Age M	2998	2.528686	1.095542	1	4
Age F	2998	2.531354	1.043363	1	4
High school M	2998	.554036	.4971545	0	1
High school F	2998	.5483656	.4977383	0	1
LFPR M	2998	.918946	.2729636	0	1
LFPR F	2998	.7461641	.4352774	0	1
Ln odds LFP F	2894	1.109142	.6913693	1.791759	3.540959
Ln odds LFP M	2518	2.601714	.9234381	.4054651	5.451038
Ln hrs/wk M	2868	3.787856	.2626683	1.386294	4.59512
Ln hrs/wk F	2336	3.557191	.4330918	0	4.564348
Ln wks/yr M	2868	3.870028	.2888853	1.098612	3.951244
Ln wks/yr F	2336	3.69982	.5736629	0	3.951244
Mean ln earn M	2998	10.16656	.3120247	9.081255	11.16829
Mean ln earn F	2998	9.937329	.3475555	8.290825	10.89547
Frac earn zero M	2998	.1150971	.075068	0	.4883721
Frac earn zero F	2998	.1114505	.0693188	0	.3673469
SD ln earn M	2998	.8329061	.1378765	.299653	1.857046
SD ln earn F	2998	.8318596	.1413075	.1315346	1.901179
Mean ln asset M	2994	7.46694	.5156904	4.493598	10.10028
Mean ln asset F	2994	7.459356	.5812731	4.877075	11.11245
SD ln asset M	2982	2.001771	.3480703	.0330005	4.720074
SD ln asset F	2992	1.823201	.340452	.1246987	3.480812
Frac asset zero M	2998	.6944602	.1316078	.2222222	1
Frac asset zero F	2998	.611422	.1078636	.183908	1

Table 2: Determinants of market tightness (P value of F test)

Observations	2998	2395	2395	2395
R-squared	0.013	0.876	0.891	0.913
State	0.57	0.00	0.00	0.00
Instruments		0.00	0.00	0.00
Labor market			0.00	0.00
M ages, edu				0.00
F ages, edu				0.00
Race				0.00

Table 3: Controlling for SR(i,j,s), P value of F test of coefficients of other sex ratio measures, SR(age,s), SR(race,s), SR(edu,s), being different from zero.

	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	OLS	IV	IV	IV
SR(i,j,s)	X	X	X	X	X	X
Other SR (race, age, edu)	X	X	X	X	X	X
States	X	X	X	X	X	X
Labor market conditions		X	X		X	X
M effects			X			X
F effects			X			X
Observations	2894	2873	2873	2330	2330	2330
R-squared	0.363	0.445	0.509	0.418	0.497	0.527
Log odds LFP F	0.00	0.00	0.00	0.00	0.00	0.00
Observations	2849	2838	2838	2322	2322	2322
R-squared	0.350	0.417	0.450	0.428	0.496	0.518
Ln hrs/wk F	0.00	0.00	0.06	0.00	0.00	0.06
Observations	2518	2503	2503	2083	2083	2083
R-squared	0.176	0.563	0.657	0.330	0.573	0.634
Ln wks/yr F	0.00	0.00	0.00	0.00	0.00	0.23
Observations	2518	2503	2503	2083	2083	2083
R-squared	0.176	0.563	0.657	0.330	0.573	0.634
Log odds LFP M	0.00	0.00	0.00	0.00	0.00	0.23
Observations	2970	2951	2951	2380	2380	2380
R-squared	0.162	0.273	0.288	0.251	0.331	0.294
Ln hrs/wks M	0.00	0.01	0.42	0.00	0.00	0.42
Observations	2970	2951	2951	2380	2380	2380
R-squared	0.073	0.159	0.181	0.098	0.166	0.163
Ln wks/yr M	0.00	0.40	0.10	0.00	0.12	0.43

Table 4a: Effects of market tightness on log odds of labor force participation of wives

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS	OLS	OLS	IV	IV	IV	IV
T(i,j,s)	0.232	-0.196	-0.184	0.330	-0.176	-0.665	-0.167
	(0.014)**	(0.024)**	(0.030)**	(0.016)**	(0.032)**	(0.209)**	(0.065)**
SR(i,j,s)						0.537	
						(0.229)*	
Observations	2894	2873	2873	2330	2330	2330	2330
R-squared	0.170	0.391	0.485	0.229	0.480	0.473	0.545
States	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Labor mark		0.00	0.00		0.00	0.00	0.00
M type			0.00				0.00
F type			0.00				0.00

Table 4b: Effects of market tightness on log hours per weeks of wives

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS	OLS	OLS	IV	IV	IV	IV
T(i,j,s)	0.027	-0.026	-0.024	0.037	-0.028	-0.089	-0.044
	(0.002)**	(0.004)**	(0.005)**	(0.002)**	(0.005)**	(0.031)**	(0.010)**
SR(i,j,s)						0.068	
						(0.034)*	
Observations	2849	2838	2838	2322	2322	2322	2322
R-squared	0.173	0.405	0.445	0.248	0.499	0.489	0.525
States	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Labor mark		0.00	0.00		0.00	0.00	0.00
M type			0.00				0.00
F type			0.00				0.00

Table 4c: Effects of market tightness on log annual weeks of wives

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS	OLS	OLS	IV	IV	IV	IV
T(i,j,s)	0.018	-0.023	-0.016	0.022	-0.022	0.011	-0.014
	(0.003)**	(0.005)**	(0.007)*	(0.003)**	(0.007)**	(0.042)	(0.014)
SR(i,j,s)						-0.037	
						(0.046)	
Observations	2849	2838	2838	2322	2322	2322	2322
R-squared	0.071	0.145	0.185	0.097	0.170	0.172	0.205
States	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Labor mark		0.00	0.00		0.00	0.00	0.00
M type			0.25				0.26
F type			0.00				0.00

Robust standard errors in parentheses; * significant at 5%; ** significant at 1%

Table 5a: Effects of market tightness on log odds labor force participation of husbands

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS	OLS	OLS	IV	IV	IV	IV
T(i,j,s)	-0.018	0.067	-0.015	-0.034	0.145	1.680	0.015
	(0.018)	(0.029)*	(0.037)	(0.019)	(0.041)**	(0.260)**	(0.072)
SR(i,j,s)						-1.695	
						(0.285)**	
Observations	2518	2503	2503	2083	2083	2083	2083
R-squared	0.067	0.561	0.652	0.138	0.582	0.536	0.645
States	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Labor mark		0.00	0.00		0.00	0.00	0.00
M type			0.00				0.00
F type			0.07				0.20

Table 5b: Effects of market tightness on log hours per weeks of husbands

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS	OLS	OLS	IV	IV	IV	IV
T(i,j,s)	-0.001	0.012	0.004	-0.003	0.018	0.102	-0.002
	(0.001)	(0.003)**	(0.004)	(0.002)	(0.003)**	(0.026)**	(0.008)
SR(i,j,s)						-0.092	
						(0.029)**	
Observations	2970	2951	2951	2380	2380	2380	2380
R-squared	0.064	0.269	0.286	0.095	0.330	0.289	0.343
States	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Labor mark		0.00	0.00		0.00	0.00	0.01
M type			0.11				0.02
F type			0.51				0.80

Table 5c: Effects of market tightness on log annual weeks of husbands

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	OLS	OLS	OLS	IV	IV	IV	IV
T(i,j,s)	0.000	0.003	-0.003	-0.001	0.011	0.071	-0.004
	(0.002)	(0.004)	(0.006)	(0.002)	(0.005)*	(0.031)*	(0.010)
SR(i,j,s)						-0.066	
						(0.034)	
Observations	2970	2951	2951	2380	2380	2380	2380
R-squared	0.045	0.158	0.176	0.061	0.175	0.155	0.194
States	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Labor mark		0.00	0.00		0.00	0.00	0.23
M type			0.92				0.30
F type			0.05				0.53

Robust standard errors in parentheses; * significant at 5%; ** significant at 1%

Table 6: Sample Statistics for the 2000 US Census and the 2003 American Time Use Survey

	2000 Census		2003 ATUS	
	Racial Composition			
White	85.6		84.8	
Black	8.1		4.6	
Hispanic	6.2		10.6	
	Marriage Characteristics			
	Husbands	Wives	Husbands	Wives
College graduate	0.6634	0.6653	0.4398	0.4812
Participation rate	0.9412	0.7269		
Usual weekly hours	44.9972	34.1112		
Weeks per year	48.6386	41.4128		
Weekly housework			10.3	15.5
Observations	746,908		408	

Table 7: Time use results for men and women

	Sex Ratio	Tightness	Instrument Tightness	Observations
		Women		
Leisure	3.3036	0.9074	2.1089	94
	(14.2688)	(15.9733)	(15.8214)	
Home Production	-0.7615	-0.8015	-0.8055	81
	(0.5968)	(0.5237)	(0.5647)	
		Men		
Leisure	-25.8676	-33.4539	1.3333	72
	(19.8472)	(20.8519)	(23.4958)	
Home Production	3.1593	2.6557	2.6557	48
	(2.4134)	(2.9029)	(2.9029)	
State controls	No	No	No	
Labor market controls	Yes	Yes	Yes	
Match-specific controls	No	No	No	
Race controls	Yes	Yes	Yes	
Age and education controls	Yes	Yes	Yes	

Determinants of market tightness

	(1)	(2)	(3)	(4)
	OLS	OLS	OLS	OLS
Lcolpw_f		0.204	0.213	0.076
		(0.031)**	(0.033)**	(0.036)*
lcolpw_m		-0.062	-0.033	-0.078
		(0.057)	(0.055)	(0.051)
lcolmw_f		0.042	0.028	0.093
		(0.033)	(0.033)	(0.029)**
lcolmw_m		-0.131	-0.112	-0.111
		(0.071)	(0.071)	(0.098)
dlcolpw_f		-0.245	-0.203	-0.218
		(0.039)**	(0.039)**	(0.035)**
dlcolpw_m		0.563	0.472	0.398
		(0.094)**	(0.092)**	(0.084)**
Dlcolmw_f		-0.200	-0.144	-0.218
		(0.037)**	(0.037)**	(0.038)**
Dlcolmw_m		0.205	0.196	0.350
		(0.074)**	(0.075)**	(0.127)**
lhspw_f		0.207	0.253	0.106
		(0.089)*	(0.085)**	(0.084)
lhspw_m		0.153	0.254	0.303
		(0.175)	(0.172)	(0.159)
lhsmw_f		-0.199	-0.233	-0.112
		(0.091)*	(0.089)**	(0.094)
lhsmw_m		-0.536	-0.509	-0.540
		(0.407)	(0.420)	(0.522)
dlhspw_f		0.465	0.425	0.142
		(0.195)*	(0.208)*	(0.186)
dlhspw_m		-0.556	-0.662	-0.449
		(0.325)	(0.317)*	(0.274)
Dlhsmw_f		-0.614	-0.340	-0.349
		(0.410)	(0.409)	(0.403)
Dlhsmw_m		2.052	1.895	2.429
		(0.597)**	(0.591)**	(0.846)**
Lmweight		0.143	0.118	0.039
		(0.022)**	(0.021)**	(0.027)
Lfweight		-0.113	-0.102	-0.092
		(0.025)**	(0.026)**	(0.031)**
M avg learn			-0.708	0.289
			(0.133)**	(0.143)*
F avg learn			0.545	0.277
			(0.109)**	(0.156)
M zero earn			-0.580	1.798
			(0.361)	(0.339)**
F zero earn			-0.879	-0.968
			(0.412)*	(0.390)*
M sd learn			-0.194	0.161
			(0.175)	(0.184)
F sd learn			0.263	0.115
			(0.152)	(0.147)
M avg lasset			-0.166	-0.013
			(0.034)**	(0.032)
F avg lasset			0.150	0.095

			(0.041)**	(0.044)*
M sd lasset			-0.176	-0.104
			(0.056)**	(0.049)*
F sd lasset			0.108	0.084
			(0.088)	(0.080)
m_asset_zero			-0.330	-0.729
			(0.215)	(0.230)**
f_asset_zero			0.370	0.554
			(0.212)	(0.196)**
Observations	2998	2395	2395	2395
R-squared	0.013	0.876	0.891	0.913
State	0.57	0.00	0.00	0.00
Instru		0.00	0.00	0.00
Unmar earn			0.00	0.00
M type				0.00
F type				0.00

Robust standard errors in parentheses
* significant at 5%; ** significant at 1%

Lcolpw_f : log fraction college educated father for j type women.
Lcolpw_m : log fraction college educated father for i type men.
lcolmw_f : log fraction college educated mother for j type women.
lcolmw_m : log fraction college educated mother for i type men.
dLcolpw_f = Lcolpw_f*college educated type j women.
dLcolpw_m = Lcolpw_m*college educated type i men.
dLcolmw_f = Lcolpw_f*college educated type j women.
dLcolmw_m = Lcolpw_m*college educated type i men.
lhspw_f : log fraction high school graduate father for j type women.
lhspw_m : log fraction high school graduate father for i type men.
lhsmw_f : log fraction high school graduate mother for j type women.
lhsw_m : log fraction high school graduate mother for i type men.
d lhspw_f = lhspw_f *college educated type j women.
dLhspw_m = Lhspw_m*college educated type i men.
dLhsmw_f = Lhspw_f*college educated type j women.
dLhsmw_m = Lhsmw_m*college educated type i men.
Lmweight: log number of type I men (by race and age) in 1960 or 1970.
Lfweight: log number of type j women (by race and age) in 1960 or 1970.
M avg learn: average type I unmarried log labor earnings.
F avg learn: average type j unmarried log labor earnings.
M zero earn: Fraction type I umarried with zero labor earnings.
F zero earn: Fraction type j umarried with zero labor earnings.
M avg lassets: average type I unmarried log non labor earnings.
F avg lassets: average type j unmarried log non labor earnings.
M zero assets: Fraction type I umarried with zero non labor earnings.
F zero assets: Fraction type j umarried with zero non labor earnings.
M SD learn: Standard deviation type I unmarried log labor earnings.
F SD learn: Standard deviation type j unmarried log labor earnings.
M SD lassets: Standard deviation type I unmarried log non labor earnings.
F SD lassets: Standard deviation type j unmarried log non labor earnings.