Labor Market Dynamics and the Migration Behavior of Married Couples *

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Abstract

Between 1964 and 2000, the intercounty migration rate of married couples declined by 15%. Concurrently, female labor force participation and the relative wages of women increased. In 1964, 36% of married households had both spouses in the labor force and women earned only 50% of the wages of men. Over the following 36 years, the fraction of dual earner households increased to 75% and women's earnings rose to 64% of men's. Using a two location household level search model of the labor market, we show that both the increase in dual earner households and the rise in women's wages contributed significantly to the decline in the migration rate of married households, with each explaining 55% and 16% of the decline, respectively. In addition, we show that the co-location problem has important implications for estimates of lifetime earnings inequality.

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1 Introduction

Between 1964 and 2000 the increase in female labor force participation of married women has led to a more than doubling in the fraction of families with both spouses in the labor force, from 36% to 75%. Over the same period there has been a 30% increase in the female to male median wage ratio among married couples. How might such changes affect the willingness of married households to migrate across counties for new work opportunities?

In this paper we document that the intercounty migration rate between 1964 and 2000 of single households increased from 5.4% to 9.1% while the migration rate of married couples declined from 5.7% to 5.0%. After controlling for changes in other demographic characteristics such as age, education, and race, for example, we show that the downward trend for married couples persists whereas the upward migration trend for single households becomes flat. This suggests that the labor market experience of married women may be key in explaining the observed decline in couples' migration.

We estimate how much of the discrepancy in these migration patterns can be accounted for by the above mentioned forces. We find that, in total, the rise of dual labor force households can account for 55% of the decline in migration, whereas rising relative wages of women can account for 16% of the decline. These two mechanisms together can account for 65% of the total decline, implying negative complementarities between the two effects. Moreover, we find that approximately 10% of the rise in dual labor force households is induced by the rising job prospects of women. Accounting for this indirect effect of rising wages results in a wage effect that can account for 20% of the decline in migration. These results are both qualitatively and quantitatively similar in two extensions wherein we include an endogenous participation margin and exogenous, non-job related moves. Consistent with Cooke (2011) and Kaplan and Schulhofer-Wohl (2017) who suggest alternative mechanisms as the primary source of declines in mgiration–namely the Great Recession and increases in technology–for the post 2000 period, we show that female labor force participation among couples instead declined after 2000. Hence, our choice of focusing on the 1964-2000 time period for the effect on migration decisions.

In order to disentangle the composition effect, i.e. more dual searching households, from the wage effect, we construct a two location model with labor market frictions and allow both individuals to receive local and foreign offers while unemployed and employed, interpreting the acceptance of any foreign offer as a move to a new location. Once a move has taken place, only the spouse receiving the foreign offer remains employed. The interaction between mobility and on-the-job search has several implications on the reservation wages of individuals. First, if both spouses are employed, the foreign wage offers for which a household is willing to move is increasing in both spouses current wages, highlighting the fact that increased wages strengthens location ties. Second, if one spouse is employed, the local reservation wage of the unemployed spouse is everywhere decreasing in the employed spouse's wage. This results from increasing location ties as one spouse climbs the job ladder. That is, as the employed spouse's wage increases, both will be less likely to receive acceptable foreign offers, thereby making local offers more attractive and decreasing the reservation wage for the second spouse. The changing location ties of dual searching households, and in particular the strong location ties present when both spouses are employed, is the driving force of our mechanism.

We provide evidence of our mechanisms using household level data from the Current Population Survey (CPS) to test the effect of the joint decision process on migration. We find that households with both members in the labor force are 10% less likely to move. Furthermore, we show that among all households that moved, the relative probability that a dual searching household moves for job related reasons than for other reasons is 26% lower than for couples with only a single searcher. These results are consistent with work by Mincer (1978) and Costa and Kahn (2000) who show that the co-location problem faced by couples has a significant impact on migration decisions.

We also show that ignoring the co-location problem and these changing location ties has implications for estimates of lifetime earnings inequality. Ignoring this additional consideration of spouses within households can bias these estimates by as much as 20% for men, 36% for women, and 23% across all married individuals. This bias has increased for men and decreased for women as increasing wages for women cause fewer male-driven moves for high foreign wage offers and, therefore, more men to optimally enter the state of unemployment as a result of similar female-driven moves throughout their career. This results in dual searching men making choices that differ from their single searching counterparts to a greater degree, implying that the standard search model is no longer a good approximation for these men. The reverse is true for dual searching women.

Our model is similar to that developed in Guler et al. (2012). We add on-the-job search and gender specific offer distributions. Flabbi and Mabli (2018), Rendon and García-Pérez (2018), and Marcassa (2014) also extend the Guler et al. (2012) model to investigate the implications of dual searching households for estimates of lifetime earnings inequality, labor market policy reforms, and spousal unemployment duration, respectively. Karahan and Rhee (2017) argue that aging populations can explain almost half of the historical decline in aggregate migration. Kennan and Walker (2011) studies the effect of expected income on migration decisions. However, none of these papers address the dual-searcher household problem.

Molloy et al. (2017) argue that there is insufficient variation in the fraction of dual searching households to explain this long run decline, however, there is disagreement on this point. Cooke (2013), for example, instead shows that the fraction of dual worker households increased by roughly 15 percentage points from 1980-2010. We show empirically that the more than doubling of such households plays a major role in the migration decline in our sample. Taskin (2016) and Foged (2016) study how migration decisions of couples may change as relative wages change and find that migration is U-shaped in the wife's share of total family income. Neither, however, are able to decompose historical migration trends. Finally, a recent working paper by Guler and Taskin (2018) uses a similar model with marriage and divorce to investigate migration trends. Our paper differs in that we focus on married couples, as other demographic characteristics cannot explain the trend for this group. We allow search on and off the job to differ, and include an empirically representative mix of single and dual searching married couples.

The remainder of the paper proceeds as follows. Section 2 presents the CPS data used in our econometric analysis and highlights the key demographic trends underlying our mechanism. Section 3 outlines our model and derives the migration rates of dual and single searching households. Section 4 presents the results of our calibration and Section 5 conducts our quantitative experiment. Section 6 concludes.

2 Data

The goal of this paper is to understand how the increase in the number of households with both spouses in the labor force and the relative earnings of women have affected the mobility of married couples. We begin this section by documenting the decline in the migration rate of married couples and adjust the series for other demographic changes. Next we turn to the two channels we are interested in: the increase in dual searching households and the rise in the relative earnings of women. We show that both the fraction of dual searching households as well as the relative earnings of women increased rapidly until 2000.

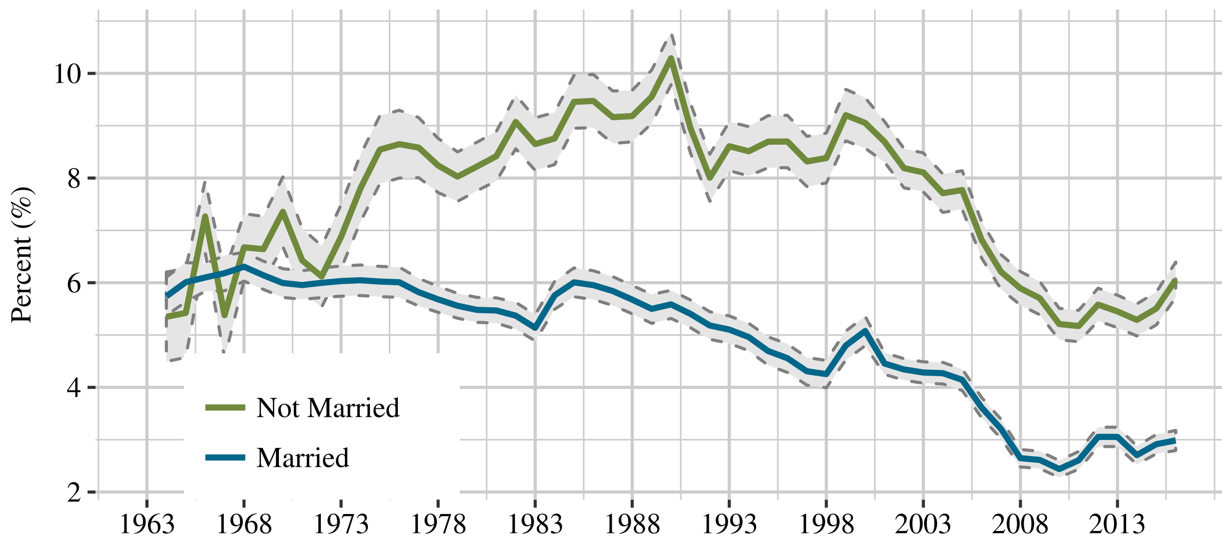
2.1 Mobility

In this section we document migration rates for married couples, the group of interest in this paper, discussed below. Our migration data come from the March sample of the Current Population Survey (CPS) through IPUMS (Flood et al., 2018). The variable of interest is the one year mobility question in which respondents were asked if they had changed residence since March of the previous year. Movers are divided into five categories: those who had moved within the same county (intracounty); those who had crossed county lines but stayed in the same state (intercounty-intrastate); those who had resided in a different state (interstate); and those who had migrated from abroad. Throughout the remainder of the paper, we refer to total intercounty migration (the sum of intercounty - intrastate and interstate migration) simply as intercounty migration.

We restrict our sample to civilian households in which the head of the household is at least 16 years old. We define married households as households in which the head of household state they are married with or without their spouse present, the remaining households are labeled as not-married. Further we define single searcher households as married households in which one individual is in the labor force and the other is out of the labor force and dual searcher households as married households in which both individuals are in the labor force.

Figure 1 shows the intercounty migration rate of civilian households between 1964 and 2015 by marital status of the head of household. The figure reveals that the trends in intercounty migration are different substantially by marital status. While the percent of not-married movers increased from 5.4% in 1964 to 9.1% in 2000, the percent of married movers decreased from 5.7% to 5.0% over the same time period. After 2000, both groups experienced a rapid decrease in their migration rates. As we will show in the following subsections, the trends in dual searcher households began to level out after the late 1990's. Moreover, Kaplan and Schulhofer-Wohl (2012) shows that a change

Figure 1: Intercounty Migration Rate by Marital Status



Notes: The 1-year geographic mobility question was not asked between 1972 to 1975, 1977 to 1980, 1985 and 1995 and a cubic spline was used. The figure shows the percent of civilian households that moved across county lines within the previous year by marital status. 95% confidence bands are plotted.

in the CPS imputation method in 2006 can explain almost half of the post 2000 decline and just over 60% of the sharp decline during the 2005-2006 period. As a result, our primary period of interest is 1964-2000 as our mechanism will be able to explain very little of the migration patterns following 2000.[1]

Others have identified that the composition of households had changed over this time period (e.g. Iyigun and Lafortune (2016)). As we discuss in Section 3, our model does not consider age, race, and education levels of household members. To remove these potential confounds from our migration data, we adjust the data to control for such characteristics. Figure 2 shows the adjusted percent of households that moved across county lines within the year. The adjusted series controls for changes in the age, sex, race, education of the head of household, total real family income, and the number of family members living in the household by estimating Eq. 1 in each year and

[1]CPS imputation flags are not available prior to 1995. Thus, we cannot drop imputed observations over the vast majority of our period of interest. As discussed in Kaplan and Schulhofer-Wohl (2012), our trends are likely robust to the imputation method used, so long as we use the same imputation method throughout our sample. We therefore also do not drop imputed values for 2000 in our main quantitative exericse.

subsequently adding the estimated time dummies to the 1964 migration rates.

$$move_{i,t} = \beta_0 + \beta_1 X_i + \eta_t + \varepsilon_{i,t} \tag{1}$$

where X_i contains the vector of controls discussed above. The figure shows that holding constant the composition of these factors at their 1964 levels removes any discernible trend for single searching households until 2000. However, holding the demographic characteristics of married households fixed at their 1964 values flattens the migration trend only slightly. These results indicate that demographic characteristics go a long way in explaining the migration trends of not-married households, but not married couples. Moreover, they provide a middle ground between past studies linking demographics and migration and Molloy et al. (2011), who suggest that demographic characteristics are likely to be unimportant for migration choices.

Indeed, Table 1 shows the results of the following linear regression for the 1964-2000 adjusted sample:

$$mrate_{t,g} = \beta_0 + \beta_1 \mathbb{1}(g = Married) + \beta_2 t + \beta_3 \mathbb{1}(g = Married)t + \varepsilon_{t,g} \tag{2}$$

where $mrate$ is the migration rate at time t for group g and g may take the values $Married$ or $Not\ Married$. The fact that the estimated coefficient on time, $\widehat{\beta}_2$, is not statistically significant suggests that demographic changes account for all of the trend in the migration rate of non married households. Moreover, the negative and significant coefficient for the married-time interaction term, $\widehat{\beta}_3$, indicates that a sizeable residual trend remains after controlling for changes in the demographic characteristics of married households.[2] Together, Figure 2 and Table 1 show that a residual trend only exists for married households. As a result, only a characteristic unique to married couples can explain this trend. One such aspect is the presence of a working spouse.

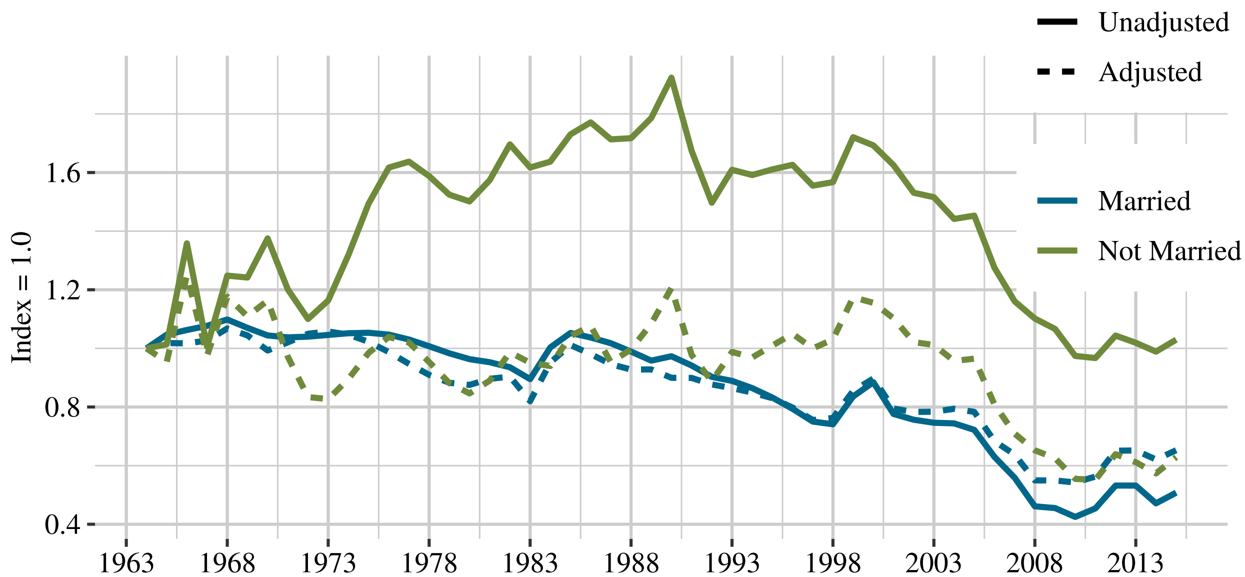
Table 1: Time Trend Estimation Results

	Not-Married Time Trend	Married-Time Interaction
Estimate	0.00093	-0.00741^{***}
Robust Standard Error	0.00127	0.00180
N	74	74

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}\ p < 0.001$

[2]An F-test also rejects the null hypothesis of $\beta_2 = \beta_3 = 0$ at the 1% level.

Figure 2: Adjusted and Unadjusted Intercounty Migration Rates



Notes: The figure shows the unadjusted and adjusted percent of households that moved within the previous year. The adjusted series are the sum of the percent of households that moved in 1964 for the group in question and the coefficients on the time dummies of a linear regression of the one year mobility status for that same group on time dummies, age, sex, race and education of the head of household , family size, and total real family income.

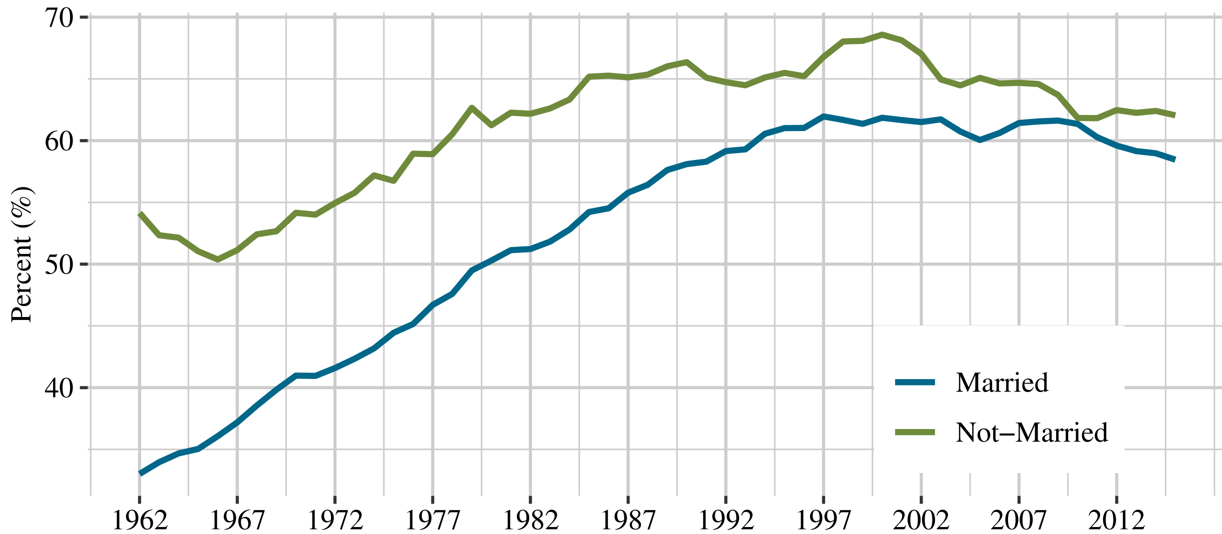
2.2 Female Labor Force Participation and the Relative Earnings of Women

It is well known that female labor force participation increased rapidly after the end of World War II, from 39% in 1964 to just over 60% in 2000. This trend is even more prevalent among married women. Figure 3 shows that both the number of married and not-married women entering the labor force over this time increased substantially and that the labor force participation rate of married women increased at a faster pace, rising by almost 30 percentage points from 1964 to 2000.[3] Importantly, Figure 4 shows that the percent of dual searcher households increased from 36% in 1964 to 75% in 2000. The trend in the percent of dual searcher households began to flatten in the late 1990's and has remained almost unchanged at roughly 75%.

The rise in the number of dual searching married households is our first channel of interest, we call this channel the composition effect. As shown in Appendix A, the migration rate of dual searcher households is less than that of single searcher households. This suggests that the increase

[3]Not shown is the labor force participation rate of married men that decreased from 85% in 1964 to 77% in 2000, implying that there may have been some crowding out of married men.

Figure 3: Female Labor Force Participation by Marital Status



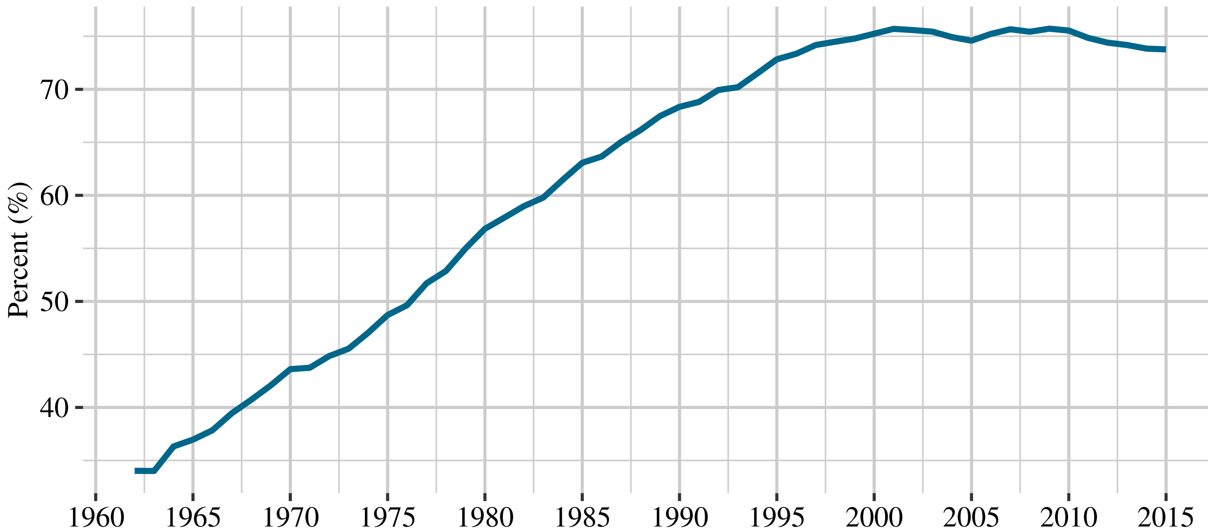
Notes: The figure shows the percent of women, age 16 or older, that are in the labor force by marital status. The data comes from the basic monthly files of the Current Population Survey.

in female labor force participation which gave rise to a compositional change of married households (the rise in dual searcher households) may have some power in accounting for the decrease in the migration rate of married couples.

The second channel we focus on that may drive down the mobility rate for married couples is the ratio of women's to men's earnings. An increase in the ratio of women's to men's earnings may drive down mobility rates of dual searcher households since outside offers must be larger when both individuals are earning more. Figure 5 shows the ratio of median yearly earnings of married women to married men from 1962 to 2017. The figure shows that women's earnings relative to men began to increase in the 1980's, rising from about 49% to 64% in 2000. This time period corresponds to the second half of our period of interest as well as the period that saw the largest decrease in the mobility of married couples. We call this channel the wage effect.

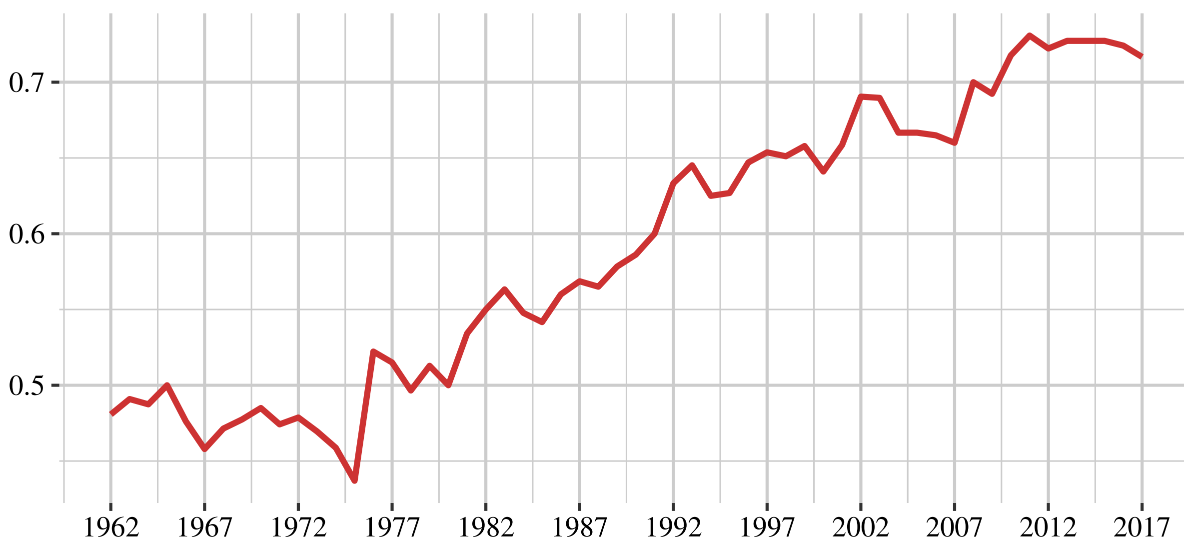
In Appendix A we test the effect of the dual searching households and relative wage differences on the migration decisions of married couples using household level data from the March CPS. We show that dual searcher households are 0.4 percentage points less likely to move across county lines and among all households that move, dual searcher households are 26% less likely to move for job related reasons. We also show that dual searcher households that live in states with higher

Figure 4: Dual Searcher Households



Notes: Plotted is the percent of households in which the head of household is in the labor force that have a spouse who is also in the labor force. Households in which the head of household is not married are included in the sample and the percent of dual earner households is calculated as: (# of married households with both spouses in the labor force)/(Total # of married households). The data comes from the basic monthly files of the Current Population Survey.

Figure 5: Ratio of Women's to Men's Median Earnings



Notes: The figure shows the ratio of women's median yearly earnings to men's median yearly earnings.

relative earnings of women are less likely to move and show suggestive evidence that among those households that do move, those that face higher relative earnings of women are less likely to move for job related reasons.

We have documented the fact that mobility rates have declined for married couples from 1960 to 2000 while female labor force participation and the fraction of dual searcher households concurrently increased. This evidence suggests that both spouses in the labor force has played a role in decreasing the fraction of married couples that choose to move. In the following section, we develop a model of dual labor search to quantify the extent to which the increase in the fraction of dual searcher households and the rise in the female-to-male median wage ratio have contributed to the observed decrease in mobility.

3 Model

We are interested in modeling the job search problem for married couples under two different scenarios. First, the single searcher household, in which only one spouse is actively searching and receiving job offers. Second, the dual searcher household, in which both spouses are actively searching and receiving job offers. The key features of our model are that searchers can receive either local job offers or foreign job offers and that men and women receive offers from gender specific wage offer distributions.

3.1 Environment

Risk neutral households search for jobs and enjoy utility over pooled income similar to Guler et al. (2012).[4] There are two types of households, single searcher households and dual searcher households. Within dual searcher households, both individuals receive job offers and are ex ante heterogeneous as they have gender specific job prospects and receive gender specific flow utility of unemployment. In the single searcher household, individuals differ as only one is searching for jobs. For now we take the increase in female labor force participation as given and do not model the household's decision of becoming either a single searching household or dual searching household

[4]Examples of alternatives to the unitary model of the household include Dey and Flinn (2005), Gemici (2016), and Lundberg and Pollak (1993).

as a benchmark. We relax this assumption and allow for an endogenous participation margin in Section 5.

Individuals who are out of the labor force receive flow utility b_O and do not receive job offers. Individuals who are in the labor force but unemployed receive flow utility b^i_U, where $i \in \{M, F\}$ indexes the gender of each household member, and can receive either local or foreign job offers. They receive local offers at wage w drawn from the c.d.f. $F^i(w)$ at exogenous poisson arrival rate α^u_l, and foreign offers at wage w drawn from the same c.d.f. at exogenous poisson rate α^u_f. Individuals also search for jobs while employed and can again receive local or foreign offers at wage w drawn from the same c.d.f. at rate α^e_l and α^e_f, respectively. All jobs separate at exogenous rate δ and households discount utility at rate r. If an individual accepts a foreign offer, the household must quit any locally held jobs and move locations.[5] Here, we abstract from moving costs since we are ultimately interested in the difference between moving rates for single searcher and dual searcher households. While Kennan and Walker (2011) discuss the importance of such costs in migration decisions, the moving costs will simply create a wedge between the reservation wage of local and foreign offers that is similar for both single and dual searching households so long as these costs do not differ across households.[6] As a result, and to simplify our calibration, we exclude these one time fixed costs. As we discuss below, dual searching households will instead incur a cost to migration in the form of lost spousal income.

3.2 Single Searcher Household

A single searcher household is composed of two individuals: one out of the labor force and the other of gender i in the labor force searching for jobs. Such a household can be in one of two states: employed-out of the labor force with value function $EO^i(w)$ or unemployed-out of the labor force with value function UO^i. The value functions are given by:

$$\begin{aligned} rUO^i = b_O + b^i_U &+ \alpha^u_l \int \max\{EO^i(w) - UO^i, 0\}\, dF^i(w) \\ &+ \alpha^u_f \int \max\{EO^i(w) - UO^i, 0\}\, dF^i(w) \end{aligned} \tag{3}$$

[5] An extension to allow for non-job related moves is discussed in Appendix B.

[6] Incorporating moving costs may differentially impact single and dual searching households when they are risk averse.

$$\begin{aligned} rEO^i(w) = b_O + w &+ \alpha_l^e \int_w EO^i(w') - EO^i(w)\, dF^i(w') \\ &+ \alpha_f^e \int_w EO^i(w') - EO^i(w)\, dF^i(w') \\ &+ \delta[UO^i - EO^i(w)] \end{aligned} \tag{4}$$

where b_U^i is the flow value of unemployment for the unemployed spouse.

In either state the household receives flow utility b_O from the spouse that is out of the labor force. Since there is no cost to moving for the single searching household as only one spouse participates in the labor force, the reservation wage for both local and foreign offers is the same. Let R_s^i be the reservation wage such that $EO^i(R_s^i) = UO^i$. This reservation wage is given by the implicit equation,

$$R_s^i - b_U^i = (\alpha_l^u + \alpha_f^u - \alpha_l^e - \alpha_f^e) \int_{R_s^i}^{\infty} \frac{1 - F^i(w)}{r + \delta + (\alpha_l^e + \alpha_f^e)[1 - F^i(w)]}\, dw. \tag{5}$$

The steady state unemployment rate, u_s^i, and steady state distribution of households employed at wage less than or equal to w, $G^i(w)$, are given by Eq. 6 and Eq. 7.

$$u_s^i = \frac{\delta}{\delta + (\alpha_l^u + \alpha_f^u)[1 - F^i(R_s^i)]} \tag{6}$$

$$G^i(w) = \begin{cases} \frac{\delta[F^i(w) - F^i(R_s^i)]}{\{\delta + (\alpha_f^e + \alpha_l^e)[1 - F^i(w)]\}[1 - F^i(R_s^i)]} & w \geq R_s^i \\ 0 & else \end{cases} \tag{7}$$

Each are derived as in Burdett and Mortensen (1998). The migration rate for single searcher households of type i is the weighted sum of the migration rate of the unemployed plus the migration rate of the employed, where the weights are given by the mass of households in each employment state. The migration rate of the unemployed is: $\alpha_f^u[1 - F^i(R_s^i)]$. The rate at which workers employed at wage w migrate is $\alpha_f^e[1 - F^i(w)]$. Therefore, the aggregate migration rate for single searcher households of type i, M_s^i, is:

$$M_s^i = u_s^i \cdot \alpha_f^u(1 - F^i(R_s^i)) + (1 - u_s^i) \cdot \alpha_f^e \int_{R_s^i}^{\infty} 1 - F^i(w)\, dG^i(w). \tag{8}$$

The aggregate migration rate for single searching households is then given by a simple weighted average over all household types as follows:

$$M_s = \xi_M \cdot M_s^M + (1 - \xi_M) \cdot M_s^F \tag{9}$$

where ξ_M denotes the fraction of single searching households with the husband in the labor force.

3.3 Dual Searcher Household

A dual searcher household is composed of two individuals both of whom are searching for jobs. Such a household can be in one of four states: employed-employed with value function $EE(w, w')$, husband employed-wife unemployed with value function $EU^M(w)$, wife employed-husband unemployed with value function $EU^F(w)$, and unemployed-unemployed with value function UU.

Just as in the single searcher household, the reservation wage for accepting jobs while in the unemployed-unemployed state is the same for both local and foreign offers as neither spouse must quit an existing job.[7] Let R_1^i be the reservation wage for spouse i when both members of the household are unemployed. Because both the offer distribution and flow utility of unemployment for each spouse is different, $EU^i(w)$ will differ by i. As a result, the reservation wage, R_1^i, is also indexed by i. The corresponding value function is:

$$\begin{aligned} rUU = b_U^M + b_U^F + (\alpha_l^u + \alpha_f^u) \int_{R_1^M}^{\infty} EU^M(w') - UU \, dF^M(w') \\ + (\alpha_l^u + \alpha_f^u) \int_{R_1^F}^{\infty} EU^F(w') - UU \, dF^F(w'). \end{aligned} \tag{10}$$

If one member of the household is employed, several decisions about accepting job offers need to be made. First, if the unemployed spouse receives a local job offer, they may take that offer if either the value of joint employment or the value of switching roles exceeds the current value of single employment. In the former case for example if the wife is employed at wage w' the husband will accept any local offer w such that $EE(w, w') \geq EU^F(w')$ and the household will enter a state of

[7] Ex post inspection of the steady state value functions reveals that a cutoff strategy is optimal for the dual searching household.

joint employment. Let $R_2^M(w)$ and $R_2^F(w)$ be the reservation wage for men and women to make this transition defined by $EE(w, R_2^F(w)) = EU^M(w)$ and $EE(R_2^M(w'), w') = EU^F(w')$, respectively. In the latter case, if the husband receives a wage offer sufficiently high to accept, but not high enough to enter joint employment, each spouse will switch roles and the household will remain in a state of single employment.

Related to this second transition is the fact that both the employed and unemployed spouse may receive a foreign offer. If the foreign offer is received by the spouse that is currently employed, the household will be willing to move for any wage greater than the one it is currently receiving. We do not allow for the possibility that members of the household can live in separate locations or that the household can split up. Therefore, if the unemployed spouse receives an acceptable foreign offer, the employed spouse must quit their job and transition into the unemployed state.

Since individuals are heterogeneous within the household, spouse i will accept the foreign wage offer, w', if and only if $EU^i(w') \geq EU^{-i}(w)$. Thus, the reservation wage to transition from employed-unemployed to unemployed-employed denoted R_3^i is given by $EU^i(w) = EU^{-i}(R_3^{-i}(w))$. Note that this reservation wage is identical for the local switching case discussed above and is not generally the 45° line. The value function for the employed-unemployed state is then given by:

$$
\begin{aligned}
rEU^i(w) = b_U^{-i} + w + (\alpha_l^e + \alpha_f^e) & \int_w^\infty EU^i(w') - EU^i(w)\, dF^i(w') \\
& + \alpha_l^u \int_{\phi^{-i}(w)}^\infty \max\left\{EE(w, w') - EU^i(w), EU^{-i}(w') - EU^i(w)\right\}\, dF^{-i}(w') \\
& + \alpha_f^u \int_{R_3^{-i}(w)}^\infty EU^{-i}(w') - EU^i(w)\, dF^{-i}(w') \\
& + \delta[UU - EU^i(w)]. \qquad (11)
\end{aligned}
$$

where $\phi^{-i}(w) = \min\left\{R_2^{-i}(w), R_3^{-i}(w)\right\}$.

If both members of the household are employed, each will accept local job offers above their current wage. If one receives a foreign offer, the household must decide whether or not to accept it and move. If the household chooses to move, the spouse who did not receive the offer transitions into the unemployed state and begins receiving flow utility b_U^{-i}. Moreover, if spouse i loses their job, their partner must decide whether or not to remain employed at their current wage or voluntarily quit and transition to the UU state rather than remain in the $EU^{-i}(w)$ state. Clearly iff $w \geq R_1^{-i}$, spouse $-i$ will remain employed rather than quit. Let $M^i(w, w')$ be the moving reservation wage for spouse i defined as $EE(w, w') = EU^i(M^i(w, w'))$ such that the household decides to move for all foreign offers above $M^i(w, w')$. Then the value function for the employed-employed state is:

$$
\begin{aligned}
rEE(w, w') = w + w' + \alpha_l^e \int_w^\infty & EE(w'', w') - EE(w, w') \, dF^M(w'') \\
+\alpha_l^e \int_{w'}^\infty & EE(w, w'') - EE(w, w') \, dF^F(w'') \\
+\alpha_f^e \int_{M^M(w,w')}^\infty & EU^M(w'') - EE(w, w') \, dF^M(w'') \\
+\alpha_f^e \int_{M^F(w,w')}^\infty & EU^F(w'') - EE(w, w') \, dF^F(w'') \\
& +\delta \left[\max\left\{EU^M(w), UU\right\} - EE(w, w')\right] + \delta \left[\max\left\{EU^F(w'), UU\right\} - EE(w, w')\right].
\end{aligned}
\tag{12}
$$

Again, notice that because the value of being in the employed-unemployed state differs by the gender of the employed spouse, $M^i(w, w')$ is indexed by gender.

The spillovers between spouses detailed above can be seen as one particular way to micro-found the reduced form location match shocks detailed in Coen-Pirani (2010) needed to match population flows across locations. For example, suppose the male in a $EU^M(w)$ household separates from their job and so the household transitions into the UU state. For the case when the on the job arrival rate is less than that when unemployed, this household will now be more likely to migrate as the previously employed spouse now receives foreign offers at a higher rate. Additionally, the spouse that was previously unemployed is more likely to cause a move as the household is not tied to any particular location by two employed spouses. On the other hand, suppose the unemployed spouse in the same $EU^M(w)$ household instead accepted a local offer and so the household transitions into the $EE(w, w')$ state. The household is now less likely to move as both spouses receive foreign offers at a lower rate and are tied to their current location by their spouse. These two situations can be interpreted as a bad and good location match shock, respectively. Moreover, this example further illustrates that dual searching households who have recently moved to a particular location are also more likely to out migrate as recent movers are least likely to be in the $EE(w, w')$ state. This is again consistent with Coen-Pirani (2010).

A steady state among dual searcher households consists of a set of four value functions, eight reservations wages, four measures of households, and three steady state distributions of households across jobs. As in the case of single searcher households, the reservation wages, measure of households in each state, and steady state distributions are sufficient to derive the aggregate migration rate. Let uu_d, eu_d^M, eu_d^F, and ee_d be the measure of households in the the unemployed-unemployed state, husband employed-wife unemployed state, wife employed-husband unemployed, and employed-employed state. Moreover, let $T^i(w)$ be the measure of households in the respective employed-unemployed state that are employed at wage less than or equal to w; and, let $H(w, w')$ be the measure of households in the employed-employed state in which one member is employed at wage less than or equal to w and the other is employed at wage less than or equal to w'.

The migration rate for dual searcher households is the weighted sum of the migration rates of all four states. uu_d households can move if either spouse receives an acceptable foreign offer while unemployed whereas ee_d households employed at wages (w, w') move if either spouse receives a foreign offer while employed in excess of $M^M(w, w')$ and $M^F(w, w')$, respectively. eu_d^i households employed at wage w can move for two reasons: if spouse i receives a foreign offer above w while employed or if spouse $-i$ receives a foreign offer while unemployed above $R_3^{-i}(w)$. Thus, the aggregate migration rate for dual searchers, M_d, is given by

$$\begin{aligned} M_d &= \alpha_f^u(2 - F_m(R_1^m) - F_f(R_1^f)) \cdot uu \\ &+ \alpha_f^e\left(eu_f \cdot \int_{R_1^F}^{\infty} 1 - F_f(w)dT_f(w) + eu_m \cdot \int_{R_1^M}^{\infty} 1 - F_m(w)dT_m(w)\right) \\ &+ \alpha_f^u\left(eu_f \cdot \int_{R_1^F}^{\infty} 1 - F_m(R_3^m(w))dT_f(w) + eu_m \cdot \int_{R_1^M}^{\infty} 1 - F_f(R_3^f(w))dT_m(w)\right) \\ &+ ee \cdot \alpha_f^e\left(\int_{R_1^M}^{\infty}\int_{R_2^f(w)}^{\infty} 1 - F_m(M_m(w, w'))d^2H(w, w') + \int_{R_1^F}^{\infty}\int_{R_2^m(w')}^{\infty} 1 - F_f(M_f(w, w'))d^2H(w, w')\right) \end{aligned} \tag{13}$$

Finally, the aggregate migration rate of married couples is then a weighted average of that for single and dual searching households:

$$M_{agg} = \zeta_d \cdot M_d + (1 - \zeta_d) \cdot M_s \tag{14}$$

where ζ_d is the fraction of dual searching households among married couples.

4 Calibration

To carry out our quantitative experiment and decompose the contribution of the increase in female labor force participation and the increase in relative wages of women to the decline in the aggregate inter-county migration rate of married households, we first calibrate the model economy at an annual frequency. We fix a number of parameters and functional forms and use simulated method of moments to calibrate the remaining parameters. We use a time period of one year to calculate our moments. Note that our model is a continuous time model and so calibrating to annual targets still allows for multiple transitions to take place within a year. The discount rate is set to 0.04 to match an annual discount factor of 0.96. We normalize the flow value of being out of the labor force, b_O, to 0. The wage offer distribution, $F^i(\cdot)$, is assumed log-normal. The separation rate, δ, is set to 0.15, which matches closely estimates of involuntary separations detailed in Hall (2005).[8]

This leaves the local and foreign arrival rates of job offers both on and off the job, $\{\alpha_l^e, \alpha_f^e, \alpha_l^u, \alpha_f^u\}$ the flow values of unemployment for both men and women, $\{b_U^M, b_U^F\}$, and the location and shape parameters of the male and female offer distributions, $\{\mu^M, \sigma^M, \mu^F, \sigma^F\}$. These parameters are calibrated by matching key moments in the data. To maintain consistency with our econometric sample in Section 2 and Appendix A, we calibrate to the year 2000. In our quantitative exercise in Section 5, we then adjust $\{\mu^F, b_U^F\}$ to match the median wage gap in 1964 while holding the ratio of b_U^F to the mean offer fixed.

We use the March Annual Social and Economic Supplement (ASEC) supplement to the CPS to calculate our targeted moments. The median and 90^{th} percentile to median wage ratio of the observed wage of married men and women measured in 1999 dollars pin down the parameters of

[8]The separation rate is the weighted annual separation rate for each type of household within our model, namely households of type $EE \to EU^i$, $EU^i \to UU$, and $EO^i \to UO^i$. Each separation rate is calculated as described by Shimer (2012) and weighted by the relative fraction of each type of household in our sample.

Table 2: Calibrated Moments

Moment	Model	Data	Targeted
Single Searcher Mig. Rate	0.057	0.057	✓
Dual Searcher Mig. Rate	0.047	0.047	✓
Mass in EE	0.82	0.79	✓
Mass in EU^M	0.12	0.13	✓
Mass in EU^F	0.056	0.047	✓
Mass in EO^M	0.95	0.89	✓
Mass in EO^F	0.90	0.82	✓
Male Median Wage ($)	37,342	38,000	✓
Female Median Wage ($)	23,636	23,000	✓
Male 90-50 Wage Ratio	1.38	2.15	✓
Female 90 – 50 Wage Ratio	1.43	2.17	✓
Flow from UU to EU_m	0.98	0.99	
Flow from UU to EU_f	0.84	0.94	
Flow from EU_m to EE	0.80	0.99	
Flow from EU_f to EE	0.98	0.99	
Flow from UO_m to EO_m	0.93	0.99	
Flow from UO_f to EO_f	0.74	0.99	

the offer distribution while the inter-county migration rates for single and dual searchers pin down the foreign arrival rates. Finally, the mass of female and male single searchers, and the mass of dual searching households within each state jointly pin down the local arrival rates and the flow value of unemployment for men and women. The results of the calibration are shown in Table 6 and Table 3.[9] The model moments closely match the data in all respects except for the 90^{th} percentile to median wage ratio, which we underestimate. Although the model is only calibrated to match the mass of households in each labor market state, the model also does relatively well in matching the flows between states.

The flow value of unemployment relative to the mean offer for men matches closely that of Shimer (2005). The flow value of unemployment for women is substantially closer to the mean wage offer, more similar to the estimates of Hagedorn and Manovskii (2008). This may reflect a number of factors including the division of homework within the household (Coltrane, 2000), gender stigmas such as those discussed in Evertsson and Nermo (2004), and child care considerations (e.g. Lundberg and Rose, 2000) that we do not model in this paper. Moreover, the fact that our estimated local arrival rates are substantially above their foreign counterparts suggest that households send

[9]We classify individuals as employed if they report usual weekly time spent working of at least 20 hours.

the majority of job applications to local labor markets. This is consistent with the findings of Marinescu and Rathelot (2018).

Table 3: Calibrated Parameters

Parameter	Value	Description
α_l^u	28.591	Local unemp. arrival rate
α_l^e	14.878	Local emp. arrival rate
α_f^u	13.850	Foreign unemp. arrival rate
α_f^e	1.025	Foreign emp. arrival rate
μ_M	9.351	Male location parameter
σ_M	0.490	Male shape parameter
μ_F	8.730	Female location parameter
σ_F	0.546	Female shape parameter
b_U^M	4,600	Male flow utility of unemp.
b_U^F	6,996	Female flow utility of unemp.

The resulting reservation wages implied by the model are shown in Figure 6 and Table 4. Panel (a) shows the reservation wages for a spouse in the uu, eu_M, and eu_F states. Noticeably, $R_2^F(w)$ and $R_2^M(w)$ begin above R_1^F and R_1^M, respectively, and decrease below these two reservation wages as the wage of the employed spouse increases. When both spouses are unemployed, and one spouse receives an employment opportunity, they have the opportunity to provide higher consumption for the entire household thereby increasing the reservation wage of the still unemployed spouse; the same reason the reservation wage in a standard job search model increases with an increase in unemployment benefits. On the other hand, however, the now employed spouse also ties the household to their current labor market, decreasing the option value to search of the unemployed spouse and therefore their reservation wage. This is similar in spirit to how the reservation wage of an unemployed individual responds to a decrease in the arrival rate of job offers in the standard search model. When the employed spouse garners a low wage, the former effect dominates whereas the latter effect dominates as the wage of the employed spouse increases.

Panel (b) displays $R_3^M(w)$, $R_3^F(w)$, and the 45-degree line. Note that both $R_3^M(w)$ and $R_3^F(w)$ are bounded below by R_1^M and R_1^F, respectively. Here, the reservation wage for an unemployed male to begin working and their employed spouse to quit either voluntarily or due to a move is everywhere above the 45-degree line, whereas the opposite is true for an unemployed female and employed male. Because men have better job prospects than their wives, the option value for a searching male is higher than that for a searching female. As a result, couples are willing to sacrifice a small amount of consumption today with the wife working so that the unemployed male spouse can search for an even better job. On the other hand, it is costly for couples to allow the female spouse to search while the husband works because he is less able to take advantage of his superior offer distribution. Also of note is that both $R_3^M(w)$ and $R_3^F(w)$ are increasing in the wage of the employed spouse. This indicates that there are some foreign offers that dual searching households reject that their single searching counterparts would accept.

Panels (c) and (d) show the moving reservation wage for dual searching households in the *ee* state. The reservation wage is increasing in both spouses wages, again illustrating the increasing location effect as your spouse's wage increases. Note that this again implies that there are foreign jobs that a dual searcher in the employed state will reject that their single searching counterparts would accept and move. Furthermore, the moving reservation wage is everywhere below the sum of both spouses' wages. Upon moving, spouses enter the *eu* state, increasing the value of search for both spouses. This move frees the household to climb the job ladder quicker by reducing a couple's location ties.

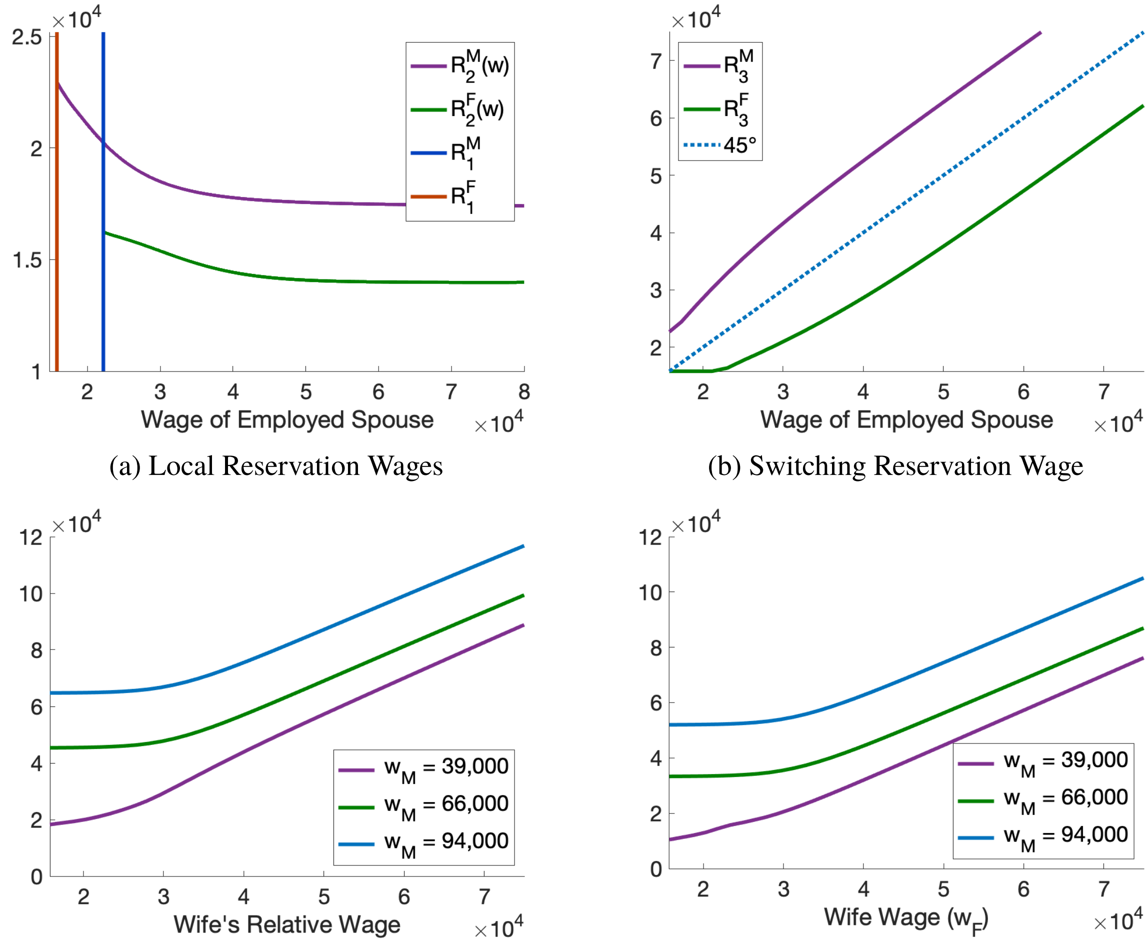
Table 4 shows the reservation wages of single searching households in the *uo* state and dual searching households in the *uu* state. The reservation wage of dual searchers in the *uu* state is lower than that of their single searcher counterparts in the *uo* state. This arises due to the fact that accepting a mediocre job offer as a dual searching couple is less harmful than for a single searcher. That is, unemployed spouses are willing to accept comparatively worse offers today to boost household consumption and later quit either because their unemployed counterpart accepts a foreign offer or because they enter the breadwinner cycle.[10] Neither of these types of quits are available to single searching couples.

Table 4: Reservation Wages in the *uu* and *uo* States

	Men ($)	Women ($)
Dual Searcher	22,195	15,815
Single Searcher	24,475	17,020

[10]This result is analogous to a single searchers problem with a higher job separation rate.

Figure 6: Model Reservation Wages



(a) Local Reservation Wages

(b) Switching Reservation Wage

(c) $M^M(w_M, w_F)$

(d) $M^F(w_M, w_F)$

5 Quantitative Experiment

5.1 Composition and Wage Effects

We now turn to our main quantitative experiment. In particular, we are interested in the effect of the increase in female labor force participation among married households and the increase in the relative wage of wives on the aggregate migration rate. To find the composition effect, we hold all calibrated parameters constant and adjust the share of dual searching households to match that in 1964. To estimate the wage effect, we fix the wage offer distribution of men and adjust the μ_F to match the 1964 female to male median wage ratio and adjust b_U^F such that the ratio of b_U^F to the mean wage offer remains constant. Moreover, we fix all other parameters including the fraction of dual searchers during this counterfactual.

The findings are given in Table 5 where we have normalized the migration rates in the calibration year to 100. We first describe our estimate of the composition effect. The model implies that the aggregate migration rate increased from 4.92% in 2000 to 5.33% in 1964. The composition effect therefore implies an overall change of 0.41 percentage points (8.3%). In the data we see that the migration rate instead increased from 5.0% in 2000 to 5.74% in 1964, a total change of 0.74 percentage points (14.8%). Thus, we conclude that the composition effect accounts for approximately 55.4% of the total decline in intercounty migration of married households seen in the data. If, in addition to changing the fraction fo dual searchers, we also change the fraction of male single searchers among all single searchers to match its 1964 value of 95.8%–up from 78.3% in 2000–then the composition effect accounts for 64.9% of the observed decline.

Next, we investigate the contribution of the wage structure to the decline in the migration rate. In our sample, the median real female wage increased from $\$15,243$ in 1964 to $\$23,000$ in 2000. The median real wage for men increased from $\$31,273$ to $\$38,000$ over this same period. These changes correspond to an increase in the female to male median wage ratio from 0.48 to 0.64. Decreasing the relative wage of married women to target a median wage ratio of 0.48 results in an increase in the migration rate of dual searching households from 4.66% to 4.82%, and in a negligible change in the migration rate of single searching households–it remained at 5.71%. Overall, the model implies an increase in the aggregate migration from 4.92% in 2000 to 5.04% in 1964. The wage effect thus implies a change of 0.12 percentage points (2.4%). Thus, we conclude that the wage effect accounts for 16.2% of the change in intercounty migration of married households. Moreover, the model implies that the composition effect is stronger overall than the wage effect. This indicates that the fact that women are entering the labor force is more important than the fact that they are becoming more similar to men in their roles within the household in explaining these long term migration trends within the United States.

To measure the degree to which complementarities exist between the two effects, we also conduct a joint counterfactual. In particular, we adjust the female offer distribution to match the female to male median wage ratio in 1964 after fixing the fraction of dual searching households to equal that in 1964. This joint counterfactual implies an aggregate migration rate of 5.39%, a

Table 5: Quantitative Results

	1964	2000	Change
Composition Effect			
Model	108.2	100.0	8.2
Data	114.8	100.0	14.8
Contribution	–	–	55.4%
Wage Effect			
Model	102.4	100.0	2.4
Data	114.8	100.0	14.8
Contribution	–	–	16.2%
Combined Effect			
Model	109.6	100.0	9.6
Data	114.8	100.0	14.8
Contribution	–	–	64.8%

rise of 0.47 percentage points (9.6%). The combination of the forces can therefore account for approximately 65% of the decline in aggregate migration over this time period. Notice that the rise in dual searcher's migration rate resulting from the wage effect is mitigated by the fact that there are simply fewer dual searchers in the counterfactual. Hence the total effect of decreasing women's relative wages and decreasing the fraction of dual searching households is less than the sum of these two individual effects. If we again also change the fraction of male single searchers among all single searchers, then the total effect increases to $\sim$ 72.9%.

Our results indicate that our mechanism was a key force reducing the number of moves by married couples before the 21st century. Consistent with Kaplan and Schulhofer-Wohl (2017) and Cooke (2011) the demographic characteristics herein studied, namely the rise in dual searcher households and relative wages of women, can explain very little of the decline in intercounty migration during the 21st century.

5.2 Endogenous Labor Force Participation

To estimate the indirect effect of wages on migration through labor force participation decisions, we include a participation margin. In the model, we allow each member of the household to be endowed with a flow value of non-participation, b^i_O, that has c.d.f. $G^i(b^i_O)$, where i indexes men and women. For simplicity, we assume that $b_O^M \perp b_O^F$. Conditional on a draw of b_O^i, the household decides whether to be a single searching household with the male in the labor force, a single searching household with the female in the labor force, or a dual searching household. Once this decision is made, the participating spouse(s) enter the labor force in the unemployed state and begin searching.[11] We maintain our assumption that only employed and unemployed individuals receive job offers. Several characteristics of our model with a participation decision are worth noting. First, because we have assumed that household preferences are linear, the option value to search of the participating member of single searching households is unaffected by the level of b_O^i received by the nonparticipating household member. Second, since the value of non-participation is independent of labor market parameters, conditional on their participation decision, households solve the same problem described in section 3 of the paper.

To calibrate the parameters of G^i and make the participation decision quantitative, we choose a parametric form. We follow Flinn (2006) and set G^i to be the c.d.f. of the exponential distribution with parameter λ^i, which adds two parameters that need to be calibrated: one scale parameter for each distribution. We use the fraction of dual searching households and the fraction of male single searchers among all single searching households as our calibration targets for λ^M and λ^F.

Table 6: Calibrated Moments: Participation Decision

Moment	Model	Data
Mass of Dual Searchers	0.752	0.752
Mass of Male Single Searchers	0.783	0.783

Table 7: Calibrated Parameters: Participation Decision

Parameter	Value	Description
λ^M	20,224	Mean Value of Male Non-participation
λ^F	34,640	Mean Value of Female Non-participation

Table 6 reports the data and model moments and Table 7 reports the calibrated parameters. We begin by estimating the total effect by recalibrating μ_F and b_U^F to match the median wage ratio of women and men in 1964 and λ^M and λ^F to match the fraction of dual searching households and

[11] Pissarides (2000) and Flinn (2006) model the participation margin analogously in standard search models.

the fraction of male single searchers among all single searching households in 1964. We estimate the wage effect by solving for the model implied migration rates using the participation parameters, λ^M and λ^F, calibrated to 2000 and the wage parameters, μ_F and b_U^F, calibrated to 1964. Similarly we estimate the composition effect by solving for the model implied migration rates using the wage parameters calibrated to 2000 and the participation parameters calibrated to 1964.

Table 8: Counterfactual Parameters: Participation Decision

Parameter	Value	Description
λ^M	17,912	Mean Value of Male Non-participation
λ^F	105,384	Mean Value of Female Non-participation

Table 9 reports the results of the three counterfactual exercises. The composition effect results in a rise of the aggregate migration rate from 4.92% to 5.37%, or 62% of the total change. This results from the fact that there are both fewer dual searchers and more male single searchers as a fraction of all single searchers. The former decreases from 75.2% to 39.4% whereas the latter rises from 78.3% to 96.5%. Moreover, the migration rate of single searchers rises from 5.71% to 5.83% due to the increase in the fraction of single searching couples that have men in the labor force. Of note is Table 8, which shows the counterfactual scale parameters for the distribution of non-participation values. Relative to 2000, the mean value of non-participation for men declined by 11% whereas the mean value of non-participation for women rose by factor of ~ 3. This captures a number of features, including increased stigma, home production expectations, and child rearing, for women in the year 1964 as compared to 2000. In contrast, the outside option for men changes only slightly, implying that stigma and expectations in the home have increased to a far lesser degree than the associated changes for women.

The wage effect instead results in a rise of the migration rate of dual searchers from 4.66% to 4.82% and a negligible change in the migration rate of single searchers. In addition, the fraction of dual searching households among all households decreases from 75.2% to 71.3% as a result of fewer women participating in the labor force. This change accounts for roughly 10% of the rise in dual searching households since 1964, implying the other factors such as stigma were a more important driver than rising female job prospects for the rise of female labor force participation among married women. The total model implied aggregate migration rate is 4.92% in 2000 and

5.07% in 1964; thus the total wage effect accounts for 23.7% of the decline in aggregate migration. The combination of these forces, i.e. the wage and composition effects, results in a rise of the aggregate migration rate from 4.92% to 5.46%. Thus, the total effect is approximately 73%. As in our benchmark model, we see that there are negative complementarities between the wage and composition effects.

Table 9: Quantitative Results: Endogenous Participation

	1964	2000	Change
Composition Effect			
Model	109.2	100.0	9.2
Data	114.8	100.0	14.8
Contribution	–	–	62.2%
Wage Effect			
Model	103.5	100.0	3.5
Data	114.8	100.0	14.8
Contribution	–	–	23.7%
Combined Effect			
Model	110.8	100.0	10.8
Data	114.8	100.0	14.8
Contribution	–	–	73.0%

5.3 Exogenous Moves

We introduce exogenous moving shocks into the benchmark model to account for the fact that not all moves are a result of job changes. Let η_d and η_s be an exogenous poisson arrival rate of non-job related moves for dual searching households and single searching households, respectively. Further, we assume that the arrival rate of such moves is not optimal, i.e. all employed members of the household become unemployed after an exogenous move. We continue to allow for endogenous participation. The value functions, migration rates, and calibrated parameters for the model with exogenous moves can be found in Appendix B. We calibrate $\eta_d = 0.0287$ and $\eta_s = 0.0314$ to match the fraction of non-job related moves shown in Table A.4.[12] Table B.1 reports the parameter values and Table B.2 reports the data and model moments.

[12]The fraction of non related moves is the sum of Family and Other in column Dual-TOT for dual searchers and Single-TOT for single searchers.

We conduct the same counterfactual exercises as before. The results are both qualitatively and quantitatively similar to those in Section 5.2 and reported in Table 10. In particular, the combination of changing female labor force participation and rising female wages results in a decline in the migration rate of 7.8% (0.39 percentage points), from 5.40% to 5.01%. Relative to the decline of 14.8% (0.71 percentage points) observed in the data, these two forces can explain roughly 52.7% of the decline. The wage effect, i.e. the change in the migration rate from only adjusting female wage offers, can explain roughly 21.6% of the observed decline. This results from the fact that the migration rate of dual searchers declines from 5.05% to 4.84% and the fraction of dual searching couples rises from 73.3% to 75.2%. Thus, the wage effect causes more households to become location constrained and existing location constraints to become stronger. The migration rate of single searchers is unchanged from the wage effect.

Table 10: Quantitative Results: Exogenous Moves

	1964	2000	Change
Composition Effect			
Model	106.2	100.0	6.2
Data	114.8	100.0	14.8
Contribution	–	–	41.9%
Wage Effect			
Model	103.2	100.0	3.2
Data	114.8	100.0	14.8
Contribution	–	–	21.6%
Combined Effect			
Model	107.8	100.0	7.8
Data	114.8	100.0	14.8
Contribution	–	–	52.7%

The composition effect, i.e. the change in the migration rate resulting from the change in the mix of dual versus single searching households, can explain approximately 41.9% of the observed decline. The composition effect results from the fact that there are both more dual searching couples and fewer male single searchers, who tend to have higher migration rates than their female counterparts. In this case, the migration rate of single searchers declines from 5.60% to 5.51% and the fraction of dual searchers rises from 37.5% to 75.2%. The dual searcher's migration rate is unchanged. Evidently, changes in the outside option of women have been more important for rise in dual searchers than the increase in their wage offers when participating in the labor force. In fact, the counterfactual outside option of women shown in Table B.3 is 3.2 times higher than that in 2000. Finally, there are negative complementarities between the composition and wage effects. This results from the fact that, the large increase in the migration rate of dual searchers in our counterfactual year is mitigated by the fact that there are fewer such households in 1964.

5.4 Lifetime Wage Inequality

Flinn (2002) argues that lifetime inequality measures should be estimated using a structural approach. In particular, many large data sets available to researchers are repeated cross-sections rather than long panels of individual earnings. Those panels in which the entire career history is observable are only available for some populations, thereby making generalization difficult. Moreover, estimates obtained in the absence of a structural model do not allow researchers to investigate the effects of potential labor market reforms or of different existing institutional frameworks across labor markets. Our model has important implications for these types of structural exercises. Flabbi and Mabli (2018) and Bowlus and Robin (2004) have since estimated lifetime earnings inequality in exercises similar to Flinn (2002), however, all three have ignored the co-location problem. Here, we add to their results by estimating the bias of lifetime earnings inequality resulting from ignoring the co-location problem. We further show how the importance of explicitly modeling it has evolved over time. We use our previous calibrations and simulate the career paths of 100,000 households of each type. Each household begins in the unemployment state and ends their career with the first labor market spell, either employment or unemployment, that ends after a total work history of 40 years. We then calculate lifetime wage earnings for spouse j as

$$\omega(j) = \sum_{i=1}^{N} e^{-r\tau_i} \int_{0}^{t_i} w_{i,j} e^{-rv} dv \tag{15}$$

where t_i is the duration of labor market spell i for the household, $\tau_{i+1} = \tau_i + t_i$ is the starting time of the $i + 1$ labor market spell for the household, and N denotes the number of labor market spells that begin prior to the 40^{th} year. Moreover, we set $w_{i,j} = 0$ when spouse j is unemployed.

Table 11 displays the coefficient of variation for men and women when assuming they are single searchers as opposed to dual searchers in both our calibration and counterfactual year. The bottom panel further shows that the coefficient of variation of lifetime earnings for all married individuals when simulating the appropriate mix of dual and single searchers in 1964 and 2000. Our model shows that ignoring the dual search problem can substantially bias estimates of lifetime earnings inequality. For men, ignoring the dual search problem biases estimates of lifetime earnings inequality upwards. This results from the fact that dual searching men optimally choose to enter the state of unemployment as a result of their wives accepting foreign offers and fewer dual searching men accepting higher, but foreign wage offers that would satisfy their single searching counterparts. In 1964, however, this fact was relatively unimportant as female job prospects within the household were relatively unimportant. As female job prospects have improved over time, the co-location problem and therefore the bias in these estimates becomes more severe for men.

Analagous mechanisms are at play for estimates of female lifetime earnings inequality. When women's job prospects were relatively unimportant within the household, more dual searching women optimally choose to enter the state of unemployment as a result of their husbands accepting foreign offers and fewer dual searching women accepting higher, but foreign wage offers that would satisfy their single searching counterparts. This results in very few high wage women relative to low wage women causing the single searcher assumptions to overestimate earnings inequality. Conversely, as female job prospect become more important within the household, a larger fraction–but still not all–of women decline to quit their jobs so that their husbands may accept foreign offers and instead accept high paying foreign offers themselves. As a result, bias in estimates of lifetime earnings inequality decrease, and in this case, become negative.

Our model suggests that lifetime earnings inequality across all married people has declined, with the coefficient of variation falling from 0.46 to 0.38. Much of this decline is due to the fact that men and women are becoming more equal in terms of labor market outcomes. Even so, the model implies that the bias in this measure has increased in magnitude from 3.3% to 23.1%. This results directly from the fact that dual searching households look less and less like their single searching counterparts and that there is a larger fraction of these types of households in 2000 than in 1964. This again illustrates the importance of explicitly modeling the co-location problem when studying inequality.

Despite abstracting from a number of important features detailed in Huggett et al. (2011) that account for earnings inequality, our results suggest that explicitly modeling the co-location problem

Table 11: Lifetime Earnings Inequality

	1964	2000
Men		
Single Searchers	0.1316	0.1715
Dual Searchers	0.1307	0.1426
Bias (%)	0.7	20.3
Women		
Single Searchers	0.1979	0.1473
Dual Searchers	0.1454	0.1548
Bias (%)	36.1	−4.8
All		
Without Dual Searchers	0.4479	0.2919
With Dual Searchers	0.4634	0.3796
Bias (%)	−3.3	−23.1

is also important for these estimates. Not doing so has the potential to produce severely biased results, particularly among men during the post 2000 period and women prior to the rise in their relative wages. A potential extension of the mechanisms presented here would include heterogeneity in initial human capital and ability, asset accumulation, and life-cycle skill accumulation in the face of borrowing constraints as in Huggett et al. (2011), Rendon and García-Pérez (2018), and Griffy (2017), respectively. Adding these features would allow us to compare the relative importance of the co-location problem to these other channels in a more detailed way and allow us to assess the impact of potential policy reforms on lifetime inequality for all family types. We leave such extensions to future work.

6 Conclusion

Between 1964 and 2000, the fraction of couples in which both spouses are in the labor force nearly doubled while the female to male wage ratio increased by 30%. Contemporaneously, the intercounty migration rate of married couples decreased from 5.74% to 5.0%, whereas the migration rate of single people increased from 5.4% to 9.1%. These differential trends suggest important differences in the decision making process of singles and married couples. Using the March CPS supplement from 1999-2015, we show that dual searching couples are 10% less likely to move than their single searching counterparts. Moreover, among those couples who did move, dual searching couples are 26% less likely to move for job related reasons.

Using a two location job search model with both single and dual searching households we then decompose the contribution of increasing female labor force participation rates and increasing female wages relative to men to the historical decline in migration. We find that the increase in the fraction of dual searching households can account for roughly 55.4% of the decline whereas the rise in relative female wages can account for approximately 16.2% of this decline. Moreover, we show that ignoring the co-location problem biases estimates of lifetime earnings inequality for married individuals downward by up to 23%, and that explicitly modeling this decision within the household has become more important as female job prospects within the household have become more equal.

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A Household-level Analysis

We use the household level data from the March CPS to show the effect that both channels have on the migration decisions of households. First, we show that dual searcher households are less likely to move across county lines and among all households that move, dual searcher households are less likely to move for job related reasons. Second, we show that dual searcher households that live in states with higher relative earnings of women are less likely to move and show suggestive evidence that among those households that do move, those that face higher relative earnings of women are less likely to move for job related reasons.

Since the primary reason for moving was only asked post 1999, we restrict the data from 1999 through 2015. Although this is not our primary time period of interest we use this time period to test cross-sectionally the hypothesis that an increase in dual searching households and in the relative earnings of women is correlated with a decrease in intercounty mobility.

We restrict the sample of households to civilian households in which the head of household is between the ages of 16 and 65. We create 3 samples of households: (1) all households in which the head of household is married (TOT), (2) all households in which the head of household is married and the spouse is present in the home (LT), and (3) all households in the TOT sample plus households that include unmarried partners (COH). Our variables of interest are the labor force status of both spouses. Thus, we divide the households into 2 subgroups: (1) both spouses are in the labor force (Dual) and (2) one spouse is in the labor force and the other is not (Single).

Table A.1 gives summary statistics for the characteristics of the head of household and their spouse, real family income, and home ownership rates for each subsample. Households across the samples differ in several respects. First, dual searcher households tend to have higher homeownership rates than single searcher households. Second, single searcher households tend to have a lower real family income. Households also differ slightly in educational attainment. In particular, dual searcher households tend to be slightly more educated than single searcher households. Both dual and single searcher households are roughly the same age and demographic makeup. Here we use all intercounty moves, below, section A.1 shows all estimates using only interstate moves.

To test the hypothesis that dual searcher households are less likely to move across county lines

Table A.1: Summary Statistics: Married Households

	Dual LT	Single LT	Dual TOT	Single TOT	Dual COH	Single COH
County Move	0.03	0.04	0.03	0.04	0.04	0.04
Real Family Income	83348.05	63294.73	83256.44	62064.38	80620.84	61534.13
Own Home	0.84	0.75	0.84	0.73	0.82	0.73
Head of Household Characteristics						
Age	43.09	44.05	43.09	43.87	42.69	43.76
White	0.85	0.84	0.85	0.83	0.85	0.84
Black	0.07	0.07	0.07	0.07	0.07	0.07
One race - Other	0.06	0.08	0.06	0.08	0.06	0.08
Multiple races	0.01	0.01	0.01	0.01	0.01	0.01
Less than High School	0.03	0.07	0.03	0.07	0.03	0.07
High School	0.32	0.39	0.32	0.39	0.33	0.40
Some College	0.29	0.24	0.29	0.24	0.29	0.25
College	0.24	0.19	0.24	0.19	0.23	0.19
Advanced Degree	0.13	0.10	0.13	0.10	0.13	0.10
Spouse Characteristics						
Age	43.20	44.46	43.19	44.45	42.76	44.12
White	0.85	0.84	0.85	0.79	0.85	0.84
Black	0.07	0.07	0.07	0.06	0.07	0.07
One race - Other	0.06	0.08	0.06	0.07	0.06	0.08
Multiple races	0.01	0.01	0.01	0.01	0.01	0.01
Less than High School	0.03	0.07	0.03	0.07	0.03	0.07
High School	0.32	0.39	0.32	0.37	0.33	0.40
Some College	0.28	0.24	0.28	0.23	0.29	0.24
College	0.24	0.19	0.24	0.18	0.23	0.18
Advanced Degree	0.13	0.10	0.13	0.15	0.13	0.10
Observations	786,805	362,946	788,243	385,329	832,232	378,948

we estimate a probit model with the following specification:

$$\text{P}(\text{move}_i = 1) = \Phi(\beta_0 + \beta_1 dual_i + X_i\gamma + \eta_t + \varepsilon_i) \tag{16}$$

where dual_i is an indicator for both spouses in household i being in the labor force, X_i is a set of household covariates that include: age, age squared, race, education for both the head of household and the spouse, an indicator for homeownership, real total family income, and an indicator that takes on the value 1 if a child is present in the home. η_t is a year fixed effect and $\Phi(\cdot)$ is the c.d.f. of the normal distribution. We define single searching households to be our base case so that our coefficient of interest is β_1, which we expect to be negative.

Table A.2 gives the estimated coefficients on the labor market indicators. The sign of β_1 indicates that the probability of moving when both spouses are in the labor force is in fact less than that when only one spouse is in the labor force. Table A.3 gives the marginal effects of the labor market indicator for a household in which both spouses are white, 40 years old, have a college degree, own a home, and have a child present in the home in the year 2000. Focusing on the "Total" column of Table A.3 shows that the probability of moving when both spouses are in the labor force is 0.399 percentage points lower than when only one spouse is in the labor force. Given that the average probability of moving across county lines across the entire sample of married households is 3.1%, this effect is quite large.

Table A.2: Probit Estimation Results: Dual Searching

	Total	Living Together	Cohab
dual	-0.0521***	-0.0510***	-0.0255**
	(0.00865)	(0.00867)	(0.00799)
N	508,371	507,457	561,633

Robust Standard errors in parentheses
$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table A.3: Probit Marginal Effects: Dual Searching

	Total	Living Together	Cohab
dual	-0.00399***	-0.00389***	-0.00202**
	(0.000682)	(0.000680)	(0.000639)
N	508,371	507,457	561,633

Marginal effects evaluated for a household in which both spouse are white, 40 years old, with a college degree, own a home and have a child present in the home in the year 2000. Robust standard errors in parentheses.
$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Next we use the reason for moving response to test if households with both spouses in the labor force are less likely to move for job related reasons. There are 19 categories for the reason for moving variable which we regroup into 4 broader categories. Our main category of interest is "New job or transfer," which includes only households that indicated that a new job or job transfer was their primary reason for moving. Our second category is "Other job reasons", which includes all households that indicated they moved to look for work or lost job, for an easier commute, retired, or other job-related reason as their primary reason for moving. Our third category is "Family," which includes all households that indicated a change in marital status, to establish own household, or other family reason as their primary reason for moving. Our fourth category is "Other," which includes all remaining reasons for moving.[13] Table A.4 gives a summary of the reasons for moving for all of our subgroups. For all subgroups, "Other" is the largest reason for moving and "New job or transfer" is the second largest for all subgroups. For the full sample of single searcher households, the percent of households that moved for "New job or transfer" reasons, 32.3%, is almost the same as the percent of households that moved for other reasons, 34.5%.

Table A.4: Reasons for Moving: Married Households

	Dual-TOT	Single-TOT	Dual-LT	Single-LT	Dual-COH	Single-COH
New job or transfer	27.0	32.3	27.1	32.4	24.5	29.6
Other job reasons	12.1	12.8	12.1	12.9	12.3	12.7
Family	22.4	20.4	22.3	20.3	25.0	22.3
Other	38.5	34.5	38.5	34.4	38.2	35.4
Total	100.0	100.0	100.0	100.0	100.0	100.0
Observations	11071	6775	11028	6757	15032	8035

We use the four broader categories to estimate the probability that a household with both spouses in the labor force will move for job related reasons. Specifically we model the probability that a move occurred for reason $j \in \{\text{New job or transfer, Other job reasons, Family, Other}\} = K$ as,

$$\text{P}(\text{whymove}_i = j) = \frac{e^{(\beta_{0j} + \beta_{1j} dual_i + X_i \gamma_j + \eta_{tj} + \varepsilon_i)}}{1 + \sum_{k \in K} e^{(\beta_{0k} + \beta_{1k} dual_i + X_i \gamma_j + \eta_{tk} + \varepsilon_i)}} \tag{17}$$

where the variables are defined as in the probit estimation. Here again, our base case is single searching households so our coefficient of interest is β_{11}. If dual searching households are in fact less likely to move for new jobs than single searching households, then we should expect β_{11} to be negative.

Table A.5 gives the estimated coefficients on the labor market indicators where all probabilities are relative to moving for other reasons. Again focusing on the "Total" column, we find that the coefficient on *dual* is negative and statistically significant for the new job or transfer reasons

[13]The remaining reasons for moving are: wanted own home - not rent, wanted new or better housing, wanted better neighborhood, for cheaper housing, other housing reason, attend/leave college, change of climate, health reasons, other reasons, natural disaster, and foreclosure or eviction.

Table A.5: Multinomial Logit Results: Dual Searching

	Total	Living Together	Cohabitating
New job or transfer			
dual	-0.300***	-0.303***	-0.273***
	(0.0479)	(0.0479)	(0.0448)
Other job reasons			
dual	-0.0734	-0.0786	-0.0290
	(0.0621)	(0.0622)	(0.0570)
Family			
dual	0.0900	0.0929	0.0948*
	(0.0524)	(0.0526)	(0.0469)
N	14,793	14,745	19,152

Robust standard errors in parentheses. Base case is other reasons for moving.
$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

for moving. The estimated coefficient implies that the relative probability a household with both spouses in the labor force moves for a new job or transfer is 25.9% lower than for a household with one spouse in the labor force[14]. The coefficients for all other reasons for moving across all subsamples are not statistically different from zero. Thus, we conclude that the strongest evidence of the difference between single and dual searcher household migration patterns points towards differences in the job search process.

Next we turn to the second channel of interest: the relative earnings of women. We construct the dependent variable, earnings ratio (ER), as the ratio of the total wage and salary income of married women to men in 1999 dollars at the state level. Since we do not know the exact location of individuals a year before the interview (we only know whether the individuals have moved and whether the move occurred across state or county lines), we now restrict our sample to households that only moved within a state in order to identify the relative earnings ratio the household faced last year. Since we are interested in the effect of the relative earnings of women on dual searcher households' migration decisions, we further restrict the sample to only dual searcher households.

To test the effect of the relative earnings of women on dual searcher household's migration decisions, we estimate a probit model with the following specification:

$$P(\text{move}_i = 1) = \Phi(\beta_0 + \beta_1 ER_s + X_i\gamma + \eta_t + \varepsilon_i) \tag{18}$$

[14]The relative probabilities of the multinomial logit are $1 - exp(\beta_j)$.

where ER_s is the relative earnings of women in the state in which the household lives. The rest of the covariates are the same as above. The coefficient of interest is β_1, which we expect to be negative.

Table A.6 gives the estimated coefficients on the relative earnings of women. The sign indicates that the probability of moving is decreasing in the earning ratio. Table A.7 gives the marginal effect of the earnings ratio for a household in which both spouses are white, 40 years old, have a college degree, own a home, and have a child present in the home in the year 2000. Again focusing on the "Total" column, Table A.7 shows that a 10 percentage point increase in the relative earnings of women decreases the probability that a dual searcher household moves by 0.07 percentage points.

Table A.6: Probit Estimation Results: Relative Earnings of Women

	Total	Living Together	Cohab
Earnings Ratio	-0.0160	-0.0156	-0.0328***
	(0.0102)	(0.0102)	(0.00885)
N	371,497	370,816	414,451

Robust Standard errors in parentheses
$^{*}p < 0.05$, $^{**}p < 0.01$, *** $p < 0.001$

Table A.7: Probit Marginal Effects: Relative Earnings of Women

	Total	Living Together	Cohab
Earnings Ratio	-0.000669	-0.000648	-0.00143***
	(0.000423)	(0.000420)	(0.000381)
N	371,497	370,816	414,451

Marginal effects evaluated for a household in which both spouse are white, 40 years old, with a college degree, own a home and have a child present in the home in the year 2000. Robust standard errors in parentheses.
$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Finally, we test whether those dual searching households that moved were less likely to move for job related reasons if they faced a higher earnings ratio. We use a multinomial logit with the same categories as in Eq. 17 and the same covariates used in Eq. 18. Again our coefficient of interest is that corresponding to the earnings ratio. Table A.8 gives the estimated coefficients for the multinomial logit. Although the estimated coefficients are negative, they are no longer statistically significant. This is most likely due to the fact that we have a very small sample with all the restrictions in place. However, we take the fact that the estimated coefficients have a negative sign as suggestive evidence that the earnings ratio decreases their probability of moving for job related reasons relative to other reasons for dual searcher households.

Table A.8: Multinomial Logit Results: Relative Earnings of Women

	Total	Living Together	Cohabitating
New job or transfer			
Earnings Ratio	-0.0502	-0.0518	-0.00514
	(0.0844)	(0.0845)	(0.0729)
Other job reasons			
Earnings Ratio	0.0558	0.0519	0.0450
	(0.0906)	(0.0908)	(0.0758)
Family			
Earnings Ratio	-0.0994	-0.107	-0.108
	(0.0700)	(0.0701)	(0.0577)
N	5,675	5,651	7,927

Robust standard errors in parentheses. Base case is other reasons for moving.
$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

The evidence presented here runs counter with that given in Molloy et al. (2017) and Kaplan and Schulhofer-Wohl (2017) who argue that the co-location problem did not contribute significantly to the historical decline in migration rates. The key differences arise because we condition on married households and do not restricting attention to couples employed in technical occupations as they do. Moreover, we control for self-selection into marriage by conditioning on married couples throughout our analysis and add to our empirics a comparison of moves for job related reasons between dual and single searchers.

A.1 Interstate Moves

Table A.9 gives the summary statistics for the different sample groups for interstate moves. Interstate moves are less, however, single searching household are observed to move more than dual searching household for the living together and total married samples. All other summary states are identical to those presented in Table A.1. Table A.10 gives the probit results for interstate moves. The results confirm those presented in Table A.2 and are of larger magnitude. The probability of moving across state lines when both individuals in the labor force of moving across states is 0.45 less than when only one spouse is in the labor force. Table A.12 shows the reasons for moving across state lines. Across all samples "New job or transfer" is the most common reason for moving. Consistent with intercounty moves, single searching households are more likely to move across state lines for job related reason than dual searching households across all samples. Table A.13 gives the results for the multinomial logit on interstate moves. Again the results are consistent with than those presented in Table A.5 and are of larger magnitude.

Table A.9: Summary Statistics: Interstate Moves

	(1) Dual LT	(2) Single LT	(3) Dual TOT	(4) Single TOT	(5) Dual COH	(6) Single COH
State Move	0.01	0.02	0.01	0.02	0.02	0.02
Total Real Family Income	83348.05	63294.73	83256.44	62064.38	80620.84	61534.13
Own Home	0.84	0.75	0.84	0.73	0.82	0.73
Head of Household Characteristics						
Age	43.09	44.05	43.09	43.87	42.69	43.76
White	0.85	0.84	0.85	0.83	0.85	0.84
Black	0.07	0.07	0.07	0.07	0.07	0.07
One race - Other	0.06	0.08	0.06	0.08	0.06	0.08
Multiple races	0.01	0.01	0.01	0.01	0.01	0.01
Less than High School	0.03	0.07	0.03	0.07	0.03	0.07
High School	0.32	0.39	0.32	0.39	0.33	0.40
Some College	0.29	0.24	0.29	0.24	0.29	0.25
College	0.24	0.19	0.24	0.19	0.23	0.19
Advanced Degree	0.13	0.10	0.13	0.10	0.13	0.10
Spouse Characteristics						
Age	43.20	44.46	43.19	44.45	42.76	44.12
White	0.85	0.84	0.85	0.79	0.85	0.84
Black sp	0.07	0.07	0.07	0.06	0.07	0.07
One race - Other	0.06	0.08	0.06	0.07	0.06	0.08
Multiple races	0.01	0.01	0.01	0.01	0.01	0.01
Less than High School	0.03	0.07	0.03	0.07	0.03	0.07
High School	0.32	0.39	0.32	0.37	0.33	0.40
Some College	0.28	0.24	0.28	0.23	0.29	0.24
College	0.24	0.19	0.24	0.18	0.23	0.18
Advanced Degree	0.13	0.10	0.13	0.15	0.13	0.10
Observations	786,805	362,946	788,243	385,329	832,232	378,948

Table A.10: Probit Estimation Results: Interstate Moves

	Total	Living Together	Cohab
dual	-0.107***	-0.106***	-0.0879***
	(0.0111)	(0.0111)	(0.0102)
N	508,371	507,457	561,633

Robust Standard errors in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.11: Probit Marginal Effects

	Total	Living Together	Cohab
dual	-0.00454***	-0.00449***	-0.00386***
	(0.000532)	(0.000532)	(0.000490)
N	508,371	507,457	561,633

Marginal effects evaluated for a household in which both spouse are white, 40 years old, with a college degree, own a home and have a child present in the home in the year 2000. Robust standard errors in parentheses.
$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$

Table A.12: Reasons for Insterstate Moves: Married Households

	Dual-TOT	Single-TOT	Dual-LT	Single-LT	Dual-COH	Single-COH
New job or transfer	38.6	43.7	38.6	43.9	34.8	40.4
Other job reasons	11.3	11.6	11.3	11.6	11.6	11.5
Family	21.1	19.8	21.1	19.7	23.5	21.6
Other	29.0	24.9	29.0	24.9	30.1	26.5
Total	100.0	100.0	100.0	100.0	100.0	100.0
Observations	5,154	3,644	5,154	3,633	6,896	4,232

Table A.13: Multinomial Logit Results: Interstate Moves

	Total	Living Together	Cohabitating
New job or transfer			
dual	-0.385***	-0.392***	-0.356***
	(0.0681)	(0.0683)	(0.0630)
Other job reasons			
dual	-0.120	-0.129	-0.0758
	(0.0952)	(0.0954)	(0.0865)
Family			
dual	-0.0222	-0.0184	0.00108
	(0.0784)	(0.0786)	(0.0703)
N	5,208	5,189	6,477

Robust standard errors in parentheses. Base case is other reasons for moving.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

B Exogenous Moves

B.1 Model

The value function for single searchers in a model with exogenous moves that arrive at Poisson rate η are:

$$\begin{aligned} rUO^i = b_O + b_U^i + \alpha_l^u \int \max\{EO^i(w) - UO^i, 0\}\ dF^i(w) \\ + \alpha_f^u \int \max\{EO^i(w) - UO^i, 0\}\ dF^i(w), \end{aligned} \tag{19}$$

$$\begin{aligned} rEO^i(w) = b_O + w + \alpha_l^e \int_w EO^i(w') - EO^i(w)\ dF^i(w') \\ + \alpha_f^e \int_w EO^i(w') - EO^i(w)\ dF^i(w') \\ + \delta[UO^i - EO^i(w)] + \eta[UO^i - EO^i(w)]. \end{aligned} \tag{20}$$

The migration rate for a type i singe searcher is:

$$M_s^i = u_s^i \cdot \alpha_f^u(1 - F^i(R_s^i)) + (1 - u_s^i) \cdot \alpha_f^e \int_{R_s^i}^{\infty} 1 - F^i(w)\ dG^i(w) + \eta. \tag{21}$$

The value functions for dual searchers in a model with exogenous moves are:

$$
\begin{aligned}
rUU = b_U^M + b_U^F &+ (\alpha_l^u + \alpha_f^u) \int_{R_1^M}^{\infty} EU^M(w') - UU \; dF^M(w') \\
&+ (\alpha_l^u + \alpha_f^u) \int_{R_1^F}^{\infty} EU^F(w') - UU \; dF^F(w'),
\end{aligned} \tag{22}
$$

$$
\begin{aligned}
rEU^i(w) = b_U^{-i} + w + (\alpha_l^e + \alpha_f^e) &\int_{w}^{\infty} EU^i(w') - EU^i(w) \; dF^i(w') \\
&+\alpha_l^u \int_{\phi^{-i}(w)}^{\infty} \max \left\{ EE(w, w') - EU^i(w), EU^{-i}(w') - EU^i(w) \right\} \; dF^{-i}(w') \\
&+\alpha_f^u \int_{R_3^{-i}(w)}^{\infty} EU^{-i}(w') - EU^i(w) \; dF^{-i}(w') \\
&+\delta[UU - EU^i(w)] + \eta[UU - EU^i(w)],
\end{aligned} \tag{23}
$$

$$
\begin{aligned}
rEE(w, w') = w + w' + \alpha_l^e &\int_{w}^{\infty} EE(w'', w') - EE(w, w') \; dF^M(w'') \\
&+\alpha_l^e \int_{w'}^{\infty} EE(w, w'') - EE(w, w') \; dF^F(w'') \\
&+\alpha_f^e \int_{M^M(w,w')}^{\infty} EU^M(w'') - EE(w, w') \; dF^M(w'') \\
&+\alpha_f^e \int_{M^F(w,w')}^{\infty} EU^F(w'') - EE(w, w') \; dF^F(w'') \\
&+\delta \left[\max \left\{ EU^M(w), UU \right\} - EE(w, w') \right] + \delta \left[\max \left\{ EU^F(w'), UU \right\} - EE(w, w') \right] \\
&+\eta[UU - EE(w, w')].
\end{aligned} \tag{24}
$$

The migration rate for a dual searching household is:

$$
\begin{aligned}
M_d = & \left(\alpha_f^u(2 - F_m(R_1^m) - F_f(R_1^f)\right) \cdot uu \\
& + \alpha_f^e \left(eu_f \cdot \int_{R_1^F}^{\infty} 1 - F_f(w) dT_f(w) + eu_m \cdot \int_{R_1^M}^{\infty} 1 - F_m(w) dT_m(w) \right) \\
& + \alpha_f^u \left(eu_f \cdot \int_{R_1^F}^{\infty} 1 - F_m(R_3^m(w)) dT_f(w) + eu_m \cdot \int_{R_1^M}^{\infty} 1 - F_f(R_3^f(w)) dT_m(w) \right) \\
& + ee \cdot \alpha_f^e \left(\int_{R_1^M}^{\infty} \int_{R_2^f(w)}^{\infty} 1 - F_m(M_m(w,w')) d^2H(w,w') + \int_{R_1^F}^{\infty} \int_{R_2^m(w')}^{\infty} 1 - F_f(M_f(w,w')) d^2H(w,w') \right) \; + \eta.
\end{aligned}
$$

B.2 Calibration

Table B.1: Calibrated Parameters: Model with Exogenous Moves

Parameter	Value	Description
α_l^u	27.770	Local unemp. arrival rate
α_l^e	15.508	Local emp. arrival rate
α_f^u	1.076	Foreign unemp. arrival rate
α_f^e	0.866	Foreign emp. arrival rate
μ_M	9.385	Male location parameter
σ_M	0.518	Male shape parameter
μ_F	8.687	Female location parameter
σ_F	0.550	Female shape parameter
b_U^M	5,005	Male flow utility of unemp.
b_U^F	6,004	Female flow utility of unemp.
η_s	0.0314	Single Searcher Exogenous move arrival rate
η_d	0.0287	Dual Searcher Exogenous move arrival rate
λ_M	20,400	Mean Value of Male Non-participation
λ_F	34,608	Mean Value of Female Non-participation

Table B.2: Calibrated Moments: Model with Exogenous Moves

Moment	Model	Data
Single Searcher Mig. Rate	0.0551	0.057
Dual Searcher Mig. Rate	0.0484	0.047
Mass in EE	0.80	0.79
Mass in EU^M	0.13	0.13
Mass in EU^F	0.068	0.047
Mass in EO^M	0.96	0.89
Mass in EO^F	0.93	0.82
Male Median Wage ($)	$39,591$	$38,000$
Female Median Wage ($)	$21,507$	$23,000$
Male 90-50 Wage Ratio	1.43	2.15
Female $90-50$ Wage Ratio	1.45	2.17
Fraction of Dual Searchers	0.752	0.752
Fraction of Male Single Searchers	0.784	0.783

Table B.3: Counterfactual Parameters: Participation Decision with Exogenous Moves

Parameter	Value	Description
λ^M	19,088	Mean Value of Male Non-participation
λ^F	110,672	Mean Value of Female Non-participation