Are Career Women Good for Marriage?*

Zvika Neeman† Andrew F. Newman‡ Claudia Olivetti§

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Abstract

We study US divorce rates, which despite the continuing rise in female labor force participation (FLFP), have been falling since the mid-1980s, reversing a two-decade trend. A cross section of U.S. states for the year 2000 displays a negative relationship between the divorce rate and FLFP.

We present theory and evidence in support of the view that these recent trends are the product of two distinct economic forces: relative to their non-career counterparts, career women display greater *selectivity* in the search for marriage partners and greater *flexibility* in sharing the benefits of a marriage with their partners.

Greater selectivity implies that career women will be older when they first marry and that their marriages will be of higher average "quality," possibly making them less prone to breakup. Greater flexibility implies that it is easier for two-earner families to re-adjust the intrahousehold allocation to compensate for changes in outside opportunities, making marriages more resistant to "shocks." Our evidence shows that both effects may be playing a role in generating the trends the trends.

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†Department of Economics, Boston University, 270 Bay State Road, Boston, MA 02215, and the Eitan Berglas School of Economics, Tel-Aviv University, Tel-Aviv, Israel 69978; Email `zvika@BU.edu`, URL http://people.bu.edu/zvika, http://www.tau.ac.il/~zvika/.

‡Department of Economics, Boston University, 270 Bay State Road, Boston, MA 02215, Email `afnewman@bu.edu`, URL http://people.bu.edu/afnewman.

§Department of Economics, Boston University, 270 Bay State Road, Boston, MA 02215, Email `olivetti@bu.edu`, URL http://people.bu.edu/olivetti/.

1. Introduction

From the early 1960s through the early 1980s, the divorce rate and the female labor force participation rate (FLFP) in the U.S. both trended upward. A large, multidisciplinary literature has offered many suggestions as to why this should be so. In the past two decades, however, the divorce rate has fallen, while FLFP continues to rise. Moreover, a cross section of U.S. states for the year 2000 displays a negative relationship between the divorce rate and FLFP. This evidence suggests that something is missing from existing theories that connect FLFP and divorce.

In our empirical analysis, the most important variables that are negatively correlated with divorce are the median age at first marriage (MAFM) and FLFP, even when both are present in the same specification. After exploring the evidence in greater depth, showing for instance that these two variables' contributions persist controlling for education, income inequality, and a number of demographic characteristics, we go on to suggest two economic forces that may be generating the trends. Relative to their non-career counterparts, career women display greater *selectivity* in the search for marriage partners and greater *flexibility* in sharing the benefits of a marriage with their partners.[1] The selectivity effect can be expected to be strongly associated with delaying the age at first marriage, but is less successful at explaining the additional contribution of FLFP.

One distinguishing feature of a career woman is that she values a marriage of given quality less than a non-career woman, because she has her own means of financial support. We show that this implies she will be choosier in selecting a partner; on average she will marry later and the quality of her marriage will be higher, leading to a lower chance of divorce. Countering this "ex-ante" *selectivity effect,* however, is an "ex-post" one: a career woman will be less tolerant of low-quality marriage; this could lead to higher divorce rates for career women relative to non-career women who marry at the same age.

The other distinguishing feature of the career woman is that the earnings she derives from the labor market facilitates surplus transfer between her and her partner. This is the origin of the *flexibility effect*: a marriage with two career partners is more stable than a marriage with only one career partner, simply because it is easier for the career partners to compensate each other for outside "temptations." We argue that this effect may have become stronger over the years with increasing marketization of formerly household-produced goods and with the narrowing of the gender earnings gap since the mid-1980s.

Existing explanations connecting divorce and FLFP are varied, but all suggest that FLFP and divorce rates should covary.[2] Most find causality running from FLFP to divorce rates:

[1] By "career woman" we mean one who works regardless of marital status; in 2000, over 85% of single women 25 to 34 were working, while only 70% of married women were. In our theoretical analysis, this is the distinction that is crucial for establishig the differential behavior of one-and two-earner fahouseholds. See also Goldin (1995).

[2] Rasul (2006) suggests that changes in divorce law would have led to temporary increases in divorce that would then have fallen back to trend levels, which have in fact been falling over the past twenty years; see also Wolfers (2006). It is not clear whether this "pipeline" effect can account for the whole trend over forty

career women are more independent and therefore willing to divorce, (Nock, 2001); the incomes of husbands and wives are substitutes, making marriage between equals less valuable (Becker et al., 1977); there is increased marital conflict within career couples (Mincer, 1985; Spitz and South, 1985), etc. An important strand finds causality running the other way: in the face of rising divorce rates, even married women have increased incentives to invest in careers, as a kind of self insurance – precautionary working, as it were (Greene and Quester, 1982; Johnson and Skinner, 1986) or because they spend less of their adult life in marriage, thus reaping fewer returns from specializing in the home, and having greater incentives to make larger investments in market work (Stevenson, 2007). Finally, some authors have suggested that the two trends reflect a spurious correlation: improvements in home production technology, which both lowers the opportunity cost of working and reduces the value of a marriage, have contributed to increased FLFP and to increases in divorce (Ogburn and Nimkoff, 1955; Greenwood and Gruner, 2004). In recent work, Stevenson and Wolfers (2007) suggest that other technological factors, such as the contraceptive pill, and changes in the wage structure, that have been found to be important determinant for the increase in labor force participation of married women might also be responsible for the concurrent increase in divorce rate.

The rest of the paper proceeds as follows: in the next section we present the empirical evidence about the relationship between FLFP and divorce. In Section 3 we discuss the selectivity effect, and in Section 4 the flexibility effect. In Section 5 we explain how our analysis can account for the shift from consent to unilateral divorce law that has swept the US during the 70s. Finally, Section 6 offer concluding remarks. Detailed descriptions of the data and some theoretical arguments are relegated to the appendix.

2. The Relationship between FLFP and Divorce

The trend reversal of the past two decades is illustrated in Figure 1. From 1980 to 2000, the rate of divorce in the US fell from 5.3 to 4.2 per 1000 people per year, undoing more than a third of the increase of the previous two decades. Meanwhile FLFP continued to rise, from 50% to 62%.

years, and in any case it makes no connection between divorce and FLFP.

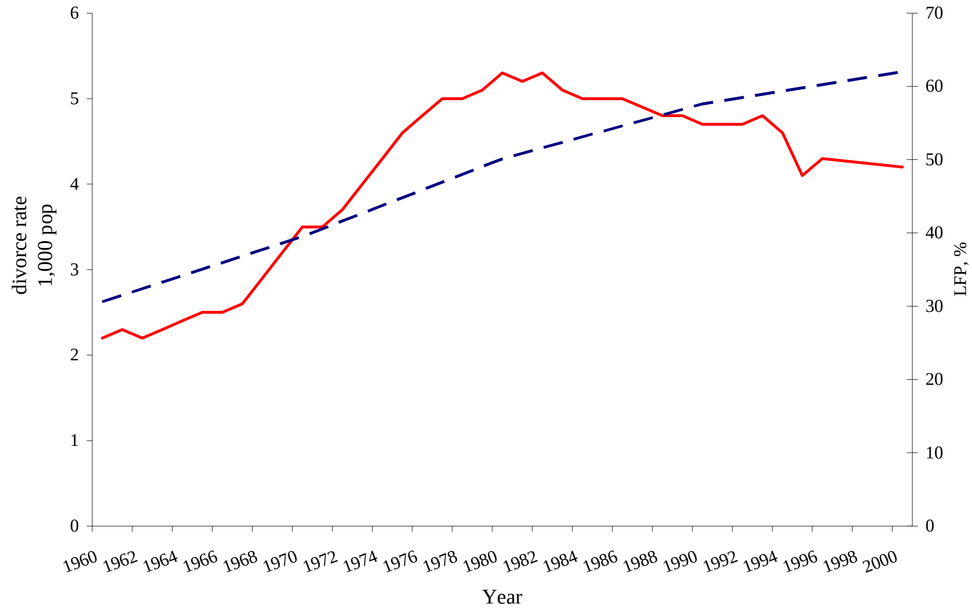

Figure 1: Divorce rate (per 1000 population) and Married Women's LFP: 1960-2000

What is more, if one looks at a cross section of US states, one finds a negative relationship between the divorce rate and FLFP. Figure 2 shows how the divorce rates, per 1,000 population, and labor force participation rates of married women vary across U.S. states in the year 2000. The divorce rate is high in states like Alabama, Kentucky and West Virginia where married women's labor force participation rates are relatively low - around 60%. Divorce rates tend to be lowest in states like Minnesota, Massachusetts, Vermont and Iowa where more than 70% of married women participate to the labor force. The (population-weighted) correlation coefficient between the two series is sizable, -0.5, and is statistically significant at the 1 percent level.

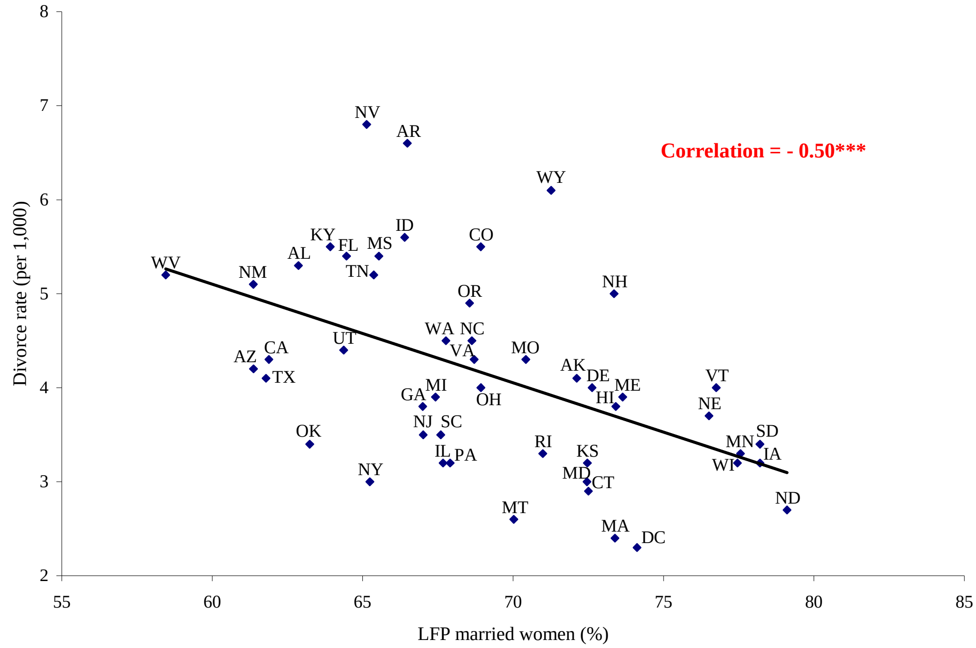

Figure 2: LFP rates of married women and divorce rates by state, 2000

Both the time-series and the cross-section evidence suggest that something is missing from the existing explanations for the relationship between FLFP and divorce. As explained in the introduction, this paper presents and discusses two economic forces that seem to have been overlooked by earlier analyses and that could account for the recent trends. However, before delving into a deeper explanation of these economic forces, we establish the robustness of the empirical findings described above.

We show that the cross section result holds even after controlling for a number of state-level characteristics that might account for a negative relationship between FLFP and divorce rates. For example, it has been shown (Martin, 2005) that more educated women are less likely to divorce than uneducated women, as well as more likely to work, so that as the average level of education in the female population increases, the divorce rate should fall. But controlling for the mean level of female education in the state, the negative correlation between FLFP and divorce persists. This correlation remains even after controlling for a number of other state-level demographic and economic characteristics such as income inequality and the median age at first marriage.

In Table 1 we present the results of state-level regressions of divorce rates on LFP of married women.[3] It also reports the results for specifications that progressively add factors, such as median age at first marriage, marriage rates, male income inequality, educational attainment, gender concentration by occupation and other socioeconomic and demographic variables (race and religion), that might be driving the negative cross-state correlation. In the basic regression, column 1, we find that a 10% difference in labor force participation rates

[3] See Data Appendix for a detailed discussion of data sources and variable definitions. In all the regressions the state level variables are population-weighted.

Table 1

	Dependent Variable is Divorce Rate, per 1,000 population								
Married Women LFP	-0.083***	-0.051**	-0.050**	-0.061**	-0.073*	-0.066*	-0.099**	-0.097**	-0.077*
	[0.027]	[0.024]	[0.022]	[0.023]	[0.037]	[0.038]	[0.040]	[0.036]	[0.041]
Female Age at First		-0.417***	-0.367***	-0.267**	-0.275**	-0.271**	-0.201	-0.360***	-0.289
Marriage		[0.091]	[0.087]	[0.113]	[0.125]	[0.126]	[0.139]	[0.131]	[0.205]
Marriage rate per			0.042***	0.041***	0.041**	0.042**	0.045***	0.034**	0.034**
1,000 population			[0.015]	[0.015]	[0.016]	[0.016]	[0.016]	[0.014]	[0.015]
Male Income Ineq.				0	0	0	-0.000*	0	0
				[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
% High School					0.015	0.015	-0.083	-0.067	-0.099
					[0.056]	[0.056]	[0.078]	[0.069]	[0.077]
% College					0.025	0.02	-0.021	-0.013	-0.058
					[0.055]	[0.055]	[0.065]	[0.058]	[0.073]
Gender Concentration						0.025	0.029	0.036	0.039
by Occupation						[0.034]	[0.035]	[0.031]	[0.032]
% White							0.076**	0.054	0.049
							[0.037]	[0.033]	[0.036]
% Black							0.062**	0.033	0.022
							[0.030]	[0.028]	[0.038]
% Asian							0.094*	0.079	0.057
							[0.055]	[0.049]	[0.057]
Number of Children								-6.281***	-6.690***
								[1.829]	[2.095]
% Protestant									-0.014
									[0.017]
% Catholic									-0.023
									[0.026]
Constant		17.888***	16.249***	15.762***	15.187***	14.458***	13.273***	22.423***	24.415***
		[2.357]	[2.284]	[2.291]	[2.792]	[2.985]	[3.050]	[3.788]	[5.501]
Observations	49	49	49	49	49	49	49	49	49
Adjusted R-squared	0.15	0.4	0.48	0.49	0.47	0.46	0.48	0.6	0.59

Robust standard errors in brackets, * significant at 10%; ** significant at 5%; *** significant at 1%.Missing observations on divorce rate for Indiana and Louisiana. For California and Colorado 1990 divorce rates. .Data Sources: Divorce and marriage rates are from the Vital Statistics of the United States, LFP & other variables are from Census IPUMS 2000, Median age at first marriage is from U.S. Census Bureau, American Community Survey 2002-2003, Census Supplementary Survey 2000-2001 Religion is from http://www.beliefnet.com/politics/religiousaffiliation.html for all states except Alaska, District of Columbia, and Hawaii for which we use www.thearda.com/mapsReports.

of married women across states translates into a lower divorce rate by 0.83. The estimate is significant at the 1 percent level. Since the average divorce rate is 4.2 per 1,000 population this is a sizeable number (roughly corresponding to a 20% lower divorce rate).

Column 2 in the Table reports the results of the regression when we also control for age at first marriage, the variable that captures our selectivity effect. Both the FLFP and age coefficients are negative, sizeable, and statistically significant at the one percent level. In this case, we find that a 10% difference in labor force participation rates of married women across states translates into a lower divorce rate by 0.51 (a 40% reduction relative to the estimate in column 1) and a 1 year difference in age at first marriage translates into a lower divorce rate by approximately 0.42. Married women labor force participation and age at first marriage are both important: taken together they can explain 40 percent of the overall cross-state variation in divorce rates (by adding a series of controls we can only increase the adjusted R-squared coefficient to 0.6). The importance of selectivity is not surprising given the strong negative association between age at first marriage and divorce rates, pointing to more stable marriages later in life.

There is also a strong positive association between median age at first marriage and FLFP (correlation coefficient = 0.37, significant at the 1% level, not shown), indicating that career women might be more selective in their choice of a partner and, as a consequence, have more stable marriages. Nevertheless, FLFP remains significant after including age at first marriage, suggesting that selectivity does not fully explain the trend.

Our results are robust to the inclusion of a string of additional explanatory variables, as shown by columns 3 to 9. For example, one could argue that the negative correlation between age at first marriage and divorce rates is driven by the fact that higher male income inequality increases the option value of a marriage thus increasing women's incentives to search longer for a partner (Gould and Paserman, 2003). This would decrease the marriage rate and, if waiting longer allows to form better matches, would also lower divorce rates thus generating a negative correlation between age at first marriage and divorce rates. However, this seems not to be the case (see column 4). In addition, higher male income inequality cannot account for the negative correlation between FLFP and divorce rate (and positive correlation between FLFP and age at first marriage) unless one is willing to assume that FLFP causes higher male inequality. Another possibility that has been discussed in the literature (McKinnish, 2004) is that lower occupational sex-segregation increased the meeting rate with opposite sex co-workers. This in turn would lead to higher marital instability. However, if this were the case, we should observe high FLFP state being characterized by higher, not lower, divorce rates. In any case controlling for a state-level measure of gender concentration by occupation (column 5) does not alter our results.

The negative correlation between divorce rates and FLFP of married women remains sizeable and significant in all specifications.[4] What could account for these findings? We

[4] We have experimented with alternative measures of married women labor force participation, such as full- and part-time participation, labor force participation of white women and labor force participation of 25-54 year old women. For all specifications we obtain results similar to the ones reported here.

argue that the fact that a woman works, thereby deriving monetary income, differentiates her from her noncareer counterpart in two important respects that affect the functioning of the "marriage market." We refer to the first such effect as selectivity, which pertains to behavior before marriage. The second effect is flexibility, which has to do with behavior inside the marriage. We hypothesize that career women will be more selective about whom they marry and more flexible in sharing surplus with their partners once they are married.

Greater selectivity implies that the career woman will search longer for a mate, waiting for indications of higher marital "quality" than an otherwise similar non-career woman. On average, then, we expect career women to be older than non-career women when they first marry and to have higher quality marriages, hence lower divorce rates. Our data support these conclusions: states with higher married women's LFP have higher median age at first marriage, and that variable in turn is negatively correlated with divorce rates.

Greater flexibility implies that once a "crisis" does occur, it is easier to find a compromise allocation of resources within the household when both partners work than when one of them doesn't. The reason is that with two earners in the household, surplus is more transferable. In this case, the distribution of resources within the household is irrelevant to the decision whether to continue a marriage: only the total surplus matters. In contrast, when one member of the household doesn't work, and therefore has only more costly instruments for surplus transfer, there are occasions when total surplus might be high enough to warrant continuation of the marriage, but the difficulties in arriving at a split of the surplus that compensates both partners for outside options results in marital breakup.

Both effects are likely to be important in explaining the trend: we have already mentioned the evidence for the selectivity effect; the fact that FLFP still negatively affects divorce even controlling for age of first marriage suggests that flexibility is also playing a role.

3. The Selectivity Effect

The data show that states with higher FLFP also have a higher median age at first marriage and lower overall divorce rate. In this section we present a simple search model that tries to capture these two effects as resulting from greater selectivity of career women when choosing a marriage partner.

As appealing as selectivity may be as an explanation, its effects are not as straightforward as might appear, because it actually operates in opposite directions before and after marriage. Assuming career women suffer less disutility from career than non-career women, they have lower search costs and will therefore be more selective by setting higher "reservation levels" of expected marriage quality. This will lead them to marry later on average and have higher average quality of marriages, which might be expected to reduce their divorce rate.

But this "ex-ante" selectivity effect has to confront an "ex-post" selectivity effect if it is to explain divorce trends, and this proves theoretically delicate. Career women will also be more selective about *remaining* married: they will divorce for a larger range of realized marriage qualities than their non-career counterparts. Thus, ex-ante effect must be sufficiently strong to skew the conditional distribution of marriage qualities enough to overcome the ex-post effect. In particular, while delaying marriage might lead women to make better choices, so that divorce rates may be lower overall when there are more career women, the ex-ante effect must also explain why divorce rates are apparently lower for career women who happen to marry early. While this is certainly possible, its theoretical likelihood does not appear overwhelming.

3.1. A Search Model

Consider the following model. There are two types of agents: males and females. Time is discrete and infinite, however males and females have only two chances to be matched and married with each other. In the first period, agents meet a potential partner and receive a signal σ of the match quality $\phi/2$ (value of public good for each). If the couple decide to marry, they learn the true quality in the next period. After learning quality, they either remain married forever (since quality is fixed for all time) or they divorce. We assume, for simplicity, that there is no remarriage.

If the agents do not marry in the first period, then they go back to the search market in the second period. This time, though, they are older and wiser, and we model this in the extreme way by assuming they can perfectly observe marriage quality before marrying in the second period.[5] Thus they will only marry if the actual quality of their match $\phi/2$ exceeds their utility when single, s.

The males all work and earn w. Females, both career and noncareer, also earn w. An agent's utility from private consumption is $c - \bar{c} - \mathbf{I}d$ for $c \geq \bar{c}$ and $-D - \mathbf{I}d$ for $c < \bar{c}$; here $\mathbf{I}$ is the indicator function assuming the value 1 if the agent works, zero otherwise. The difference between non career females and career females as well as males is that $d = 0$ for the latter while $d = d^N > w > 2\bar{c}$ for the former. Assume $D > d^N$.

We assume that utility is fully transferable. Under these assumptions, career females work whether married or not. Non career females work if single but not if married (the married couple's surplus in the latter case is $2w - 2\bar{c} - d^N$ if she works, which is less than the $w - 2\bar{c}$ obtained when she doesn't). Assume that in the latter case, the non career female produces an additional local public good $\varphi = \bar{c}$ that makes her male partner indifferent between marrying her and marrying a career female (this is a simple way of abstracting from general equilibrium effects on search strategies that would depend on the fraction of career women in the population).

[5] This is the main departure from standard analysis (Jovanovic, 1984; Bougheas and Georgellis, 1999; Rasul, 2006), which tends to assume stationarity of the signal structure. There are two problems here (aside from its empirical implausibility in this context). First, because of ex-post selectivity, it doesn't clearly result in lower divorce rates among two-earner than one-earner households. Moreover, it may not even lead to delayed marriage among working agents: nonworking agents in such a world have a stronger incentive to wait for good signals (if they don't discount the future too much), since they are likely to remain married longer than working agents.

The flow utility while single (either unmarried or divorced) is therefore $w - \bar{c}$ for career females, and $w - \bar{c} - d^N$ for non career females. Because of transferable utility, we can just as well consider the flow to a potential couple, which is obtained by adding $w - \bar{c}$ to these expressions. When married, the flow is $w - 2\bar{c} + \phi + \varphi$ for a couple with a noncareer agent and $2w - 2\bar{c} + \phi$ for one with two career agents. Thus we can normalize the flow in terms of the difference between the unmarried and the married states to be s when single (either before or after marriage) and ϕ when married, where $s = 0$ for career, and $s = -d^N < 0$ for noncareer.

It is important to note that greater selectivity need not imply that divorce is less probable overall for two-earner families, because their divorce probability need not be lower if they happened to marry in the first period: when s rises, one is more willing to divorce, so even if one's marriage quality is higher on average, the divorce probability might rise nonetheless. In fact, somewhat lengthy calculations show that the probability of divorce for those who marry in period 1, will be increasing in s for a broad range of parameters. (Therefore if we had assumed a stationary signal structure, we would be unable to show that selectivity robustly implies a lowering of divorce rates – the nonstationarity appears crucial.)

The important point is that empirically, the selectivity effect would manifest itself largely through the delay in marriage rather than through a reduction in divorce given age at marriage. Thus the fact that FLFP still enters negatively in our regressions when controlling for the age at marriage suggests that the flexibility effect is playing a role, and is not a residual of the selectivity effect uncaptured by controlling for the age at marriage.

3.2. Example.

Suppose that:

single career females obtain a payoff of $s = 0$ in every period in which they are single;

single non career females obtain a payoff of $s = -1$ in every period in which they are single;

marriage is worth X, Y, and Z each period with probabilities x, y, and z respectively. Suppose that $X < -1 < Y < 0 < Z$, and $xX + yY + zZ > 0$.

agents discount their future utility by β.

Suppose that women observe a signal about the quality of the match in the first matching period. Suppose that the signal, $\sigma \leq 1 - y$, is the probability that the marriage would be of high quality. The probability that the marriage is of the intermediate quality is known to be equal to y.

A non career woman (with $s < Y$) who observes a signal σ obtains the following expected payoff from marrying in the first match period:

$$(1 - y - \sigma) X + yY + \sigma Z + \frac{\beta}{1 - \beta} [(1 - y - \sigma) s + yY + \sigma Z]$$

She obtains the following expected payoff from delaying her marriage to the second matching period:

$$s+\frac{\beta}{1-\beta}\left[xs+yY+zZ\right]$$

Such a woman would therefore marry in the first period IFF

$$(1-y-\sigma)X+yY+\sigma Z+\frac{\beta}{1-\beta}\left[(1-y-\sigma)s+yY+\sigma Z\right]\geq s+\frac{\beta}{1-\beta}\left[xs+yY+zZ\right]$$

IFF

$$\sigma\geq\frac{(1-\beta)s-(1-\beta)\left((1-y)X+yY\right)+\beta z\left(Z-s\right)}{Z-(1-\beta)X-\beta s}$$

A career woman (with $s>Y$) who observes a signal σ obtains the following expected payoff from marrying in the first match period:

$$(1-y-\sigma)X+yY+\sigma Z+\frac{\beta}{1-\beta}\left[(1-y-\sigma)s+ys+\sigma Z\right]$$

She obtains the following expected payoff from delaying her marriage to the second matching period:

$$s+\frac{\beta}{1-\beta}\left[xs+ys+zZ\right]$$

Such a woman would therefore marry in the first period IFF

$$(1-y-\sigma)X+yY+\sigma Z+\frac{\beta}{1-\beta}\left[(1-y-\sigma)s+ys+\sigma Z\right]\geq s+\frac{\beta}{1-\beta}\left[xs+ys+zZ\right]$$

IFF

$$\sigma\geq\frac{(1-\beta)s-(1-\beta)\left((1-y)X+yY\right)+\beta z\left(Z-s\right)}{Z-(1-\beta)X-\beta s}$$

Observe that the threshold signal above which such a woman marries in the first period is increasing in s, which implies that a career woman is more likely to delay her marriage than a non career woman.[6]

If we assume that σ is uniformly distributed on the longest possible interval on which it can be distributed, then since it must be that $E[\sigma]=z$, it must be that if $2z<1-y$ then $\sigma\sim U[0,2z]$, and if $2z>1-y$ then $\sigma\sim U[2z-(1-y),1-y]$. The fact that more than half of the couples never divorce is consistent with $z\geq 1/2$, which is consistent with the latter uniform distribution. We therefore assume that $\sigma\sim U[2z-(1-y),1-y]$ and proceed to calculate the likelihood of divorce conditional on marriage in the first period and overall probability of divorce for career and non career women.

[6]

$$\begin{aligned}\frac{d}{ds}\left(\frac{(1-\beta)s-(1-\beta)\left((1-y)X+yY\right)+\beta z\left(Z-s\right)}{Z-(1-\beta)X-\beta s}\right) &= \frac{(1-\beta)\left(Z(1-z\beta)-Yy\beta-X(1-(z+y)\beta)\right)}{(X-Z-X\beta+s\beta)^2}\\ &> 0\end{aligned}$$

A career woman who marries in the first period divorces with probability $1-\sigma$, where $\sigma \geq \sigma_C \equiv \sigma(s=0)$. Taking an expectation of signals, the expected rate of divorce for such a woman is:

$$d_C \equiv 1 - \frac{\sigma_C}{2} - \frac{1-y}{2}$$

A non career woman who marries in the first period divorces with probability $1-y-\sigma$, where $\sigma \geq \sigma_N \equiv \sigma(s=-1)$. Taking an expectation of signals, the expected rate of divorce for such a woman is:

$$d_N \equiv 1 - y - \frac{\sigma_N}{2} - \frac{1-y}{2}$$

A career woman is therefore more likely to divorce than a non career woman conditional on having married in the first period IFF

$$1 - \frac{\sigma_C}{2} - \frac{1-y}{2} \geq 1 - y - \frac{\sigma_N}{2} - \frac{1-y}{2}$$

IFF

$$2y \geq \sigma_C - \sigma_N$$

Because $\sigma(s)$ is increasing in y (at a rate of $Y-X$ for a fixed s) this inequality is satisfied as long as y is "large enough." For example, if we assume that

$$\begin{aligned}
s_C &= 0 \\
s_N &= -1 \\
\beta &= .9 \text{ (close to 1)} \\
z &= .5 \\
Y &= -\varepsilon \text{ (close to 0)} \\
Z &= -X+1
\end{aligned}$$

then it is satisfied approximately for all values of $y \geq 0$.

Observe that under this parametrization,

$$\begin{aligned}
\sigma_C &\equiv \sigma(s=0) \\
&\approx \frac{-.1(1-y)X + .45(1-X)}{1-X-.1X}
\end{aligned}$$

and

$$\begin{aligned}
\sigma_N &\equiv \sigma(s=-1) \\
&\approx \frac{-.1 - .1(1-y)X + .45(1-X+1)}{1-X-.1X+.9}
\end{aligned}$$

And

$$\begin{aligned}
2y - (\sigma_C - \sigma_N) &\approx 2y - \frac{-.1(1-y)X + .45(1-X)}{1-X-.1X} + \frac{-.1 - .1(1-y)X + .45(1-X+1)}{1-X-.1X+.9} \\
&\underset{X=-10}{\geq} 0
\end{aligned}$$

$$2y - (\sigma_C - \sigma_N) \approx 2y - \frac{-.1(1-y)X + .45(1-X)}{1-X-.1X} + \frac{-.1 - .1(1-y)X + .45(1-X+1)}{1-X-.1X+.9} \underset{X=-3}{\geq} 0$$

Denote the rate of divorce of a career and non career woman conditional on having married in the first period by d_C and d_N, respectively. Then, overall divorce rate of career women is lower than that of non career women IFF

$$\frac{1-y-\sigma_C}{2(1-y-z)} \cdot d_C \leq \frac{1-y-\sigma_N}{2(1-y-z)} \cdot d_N$$

Under the parametrization above, this inequality holds approximately IFF

$$\frac{1-y-\frac{(1-y)+.45(11)}{11+1}}{2(1-y-.5)} \left(\frac{1}{2} - \frac{\frac{(1-y)+.45(11)}{11+1}}{2} + \frac{y}{2} \right)$$
$$\leq \frac{1-y-\frac{-.1+(1-y)+.45(11+1)}{1+10+1+.9}}{2(1-y-.5)} \left(\frac{1}{2} - \frac{\frac{-.1+(1-y)+.45(11+1)}{1+10+1+.9}}{2} - \frac{y}{2} \right)$$

IFF y is sufficiently close to zero.

4. The Flexibility Effect

Compare a marriage in which both partners work with one in which only one partner works. Assume that other than work status, all characteristics of the partners are identical across marriages. Assume that upon marriage, a household produces a local public good yielding utility ϕ to each partner.[7]

Upon marriage, the partners will settle on an allocation of goods on the household utility possibility frontier, e.g. via Nash bargaining relative to the disagreement point of consuming their respective wages (consistent with the data, it is reasonable to assume that both partners work prior to marriage – below we model this explicitly when considering the selectivity effect).

Now suppose one partner experiences a marriage-specific shock: perhaps it receives an "outside offer" worth more than the utility it derives in the current marriage.[8] The career spouse of the shocked partner has something that the non-career spouse doesn't have: money,

[7] We assume away increasing returns to scale in market purchased goods: allowing it might favor two-earner households in local public good production, but the greater time budget available for home production in one-earner households could wash this effect out.

[8] Similar effects could be obtained from a preference shock, e.g. a change to the utility derived by one partner from the local public good.

the instrument par excellence for transferring utility from one partner to the other.[9] More often than not, cash would not be transferred directly; rather cash allows the compensating partner to purchase (close to) an optimal market basket of goods (plasma TV, fur coat) for the shocked partner; a non-career spouse can only supply household-produced goods that are likely to be imperfect substitutes.

In short, the two-earner household well approximates the case of transferable utility. If the total surplus generated in the current household exceeds that of the outside offer, a new intra-household allocation can always be found that will keep the shocked partner indifferent between the marriage and the outside offer and still leave the spouse better off than being single. The marriage stays together. (If the outside offer exceeds the total surplus of the household, the marriage breaks apart. Either way, the outcome is efficient.)

In the one-earner household, the spouse who doesn't work has no money to make a counteroffer, only less efficient instruments for transferring utility such as payment in household-produced goods. In some cases, the spouse will not be able to compete with the new offer and the marriage dissolves. The one-earner household is a case of non- (or imperfectly) transferable utility, and in some cases the marriage will break up inefficiently. We illustrate the difference in Figure 3.

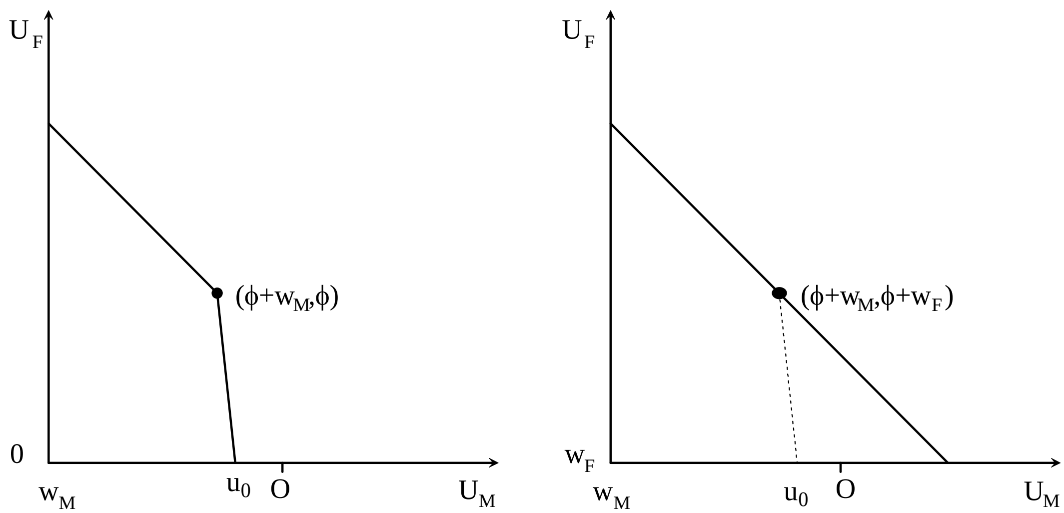

Figure 3: The Flexibility Effect

In both graphs, U_M and U_F denote the utility of the male and female, respectively, w_M and w_F denote the income, or wage, of the male and female, respectively, and ϕ denotes the value that the male and female each derive from their marriage. The left graph describes the utility possibility frontier of a married couple where the woman is not working, and the

[9] Partly this follows from a basic implication of standard consumer theory analysis, which leads to one of the favorite provocative lessons of intermediate micro teachers everywhere that it is better to give gifts (or government subsidies) in cash than in kind. One way of seeing this is spelled out in the Appendix. But also, the immediacy of cash transfers also reduces the incentive problems inherent in (often, delayed) in-kind transfers such as effort around the house, favors, etc. For more on the underpinnings and implications of nontransferability in matching environments, see Legros and Newman (2007).

right graph describes the utility possibility frontier of a married couple where the woman is working.

A working spouse, having greater means to transfer utility to the partner than a non-working one, can do so in case the partner receives an "outside offer." If partner M receives an offer such as O that is larger than $\phi + w_M$ (but less than $2\phi + w_M + w_F$) then the one-earner household will have a greater difficulty finding an allocation that yields M at least O, whereas the two-earner household could match it. Thus the two-earner household will more often (i.e. for a fixed distribution of outside offers) remain married than a one-earner household. Thus, the greater flexibility offered by the cash instrument actually contributes to the relative stability of the working women's marriages.

The flexibility effect therefore helps to explain the trend in US divorce rates and the cross sectional variation in divorce and FLFP we observe at the state level. As more households become two-earner, there should be greater stability of the marriages that form, all else the same. More specifically, the continued significance of FLFP in our regressions after controlling for MAFM (the positive relationship between FLFP and MAFM being the most robust implication of our selectivity model) supports the view that flexibility is also playing a role.

4.1. "General Equilibrium" Effects and Law and Economics

So far we have made a purely compositional argument for the flexibility effect's role in explaining divorce trends: if we are talking about one couple in a sea of couples, then the analysis might stop there. But when we are talking about trends, and whether the presence of more career women accounts for the greater stability of marriage, we are talking about the whole population, and so the analysis ought not stop here. All else isn't the same if we compare two populations, one with few two-earner households, one with many. For a working spouse might be able to match a given outside offer more easily than a nonworking spouse, but what if the outside offers are better when there are many two-earner households than when there are not?

To address this question, we need to allow some mechanism by which the presence of two-earner households could affect the distribution of offers, and for this we have to allow for the possibility of remarriage. Elsewhere (Neeman, Newman and Olivetti, 2007) (NNO), we have studied such a model, which adopts a simple search-and-bargaining framework in the spirit of Becker et al. (1977). Individuals may encounter potential partners whether or not they are married (offers by such potential partners constitute the "outside options" that affect intra-household bargaining). Individuals who work are now potentially able to offer higher net surplus to potential partners than individuals who don't. Finally, we take the degree of female labor force participation to be the exogenous parameter.[10]

Using these ingredients, we compute aggregate rates of divorce as a function of FLFP

[10] Of course, LFP is endogenous; but in light of the literature, it's hard to see how endogenizing it would explain its negative correlation with divorce, so we feel justified taking it as exogenous.

under two forms of divorce law, *unilateral* and *consent*. The former allows a divorce if one partner decides to leave the marriage. The other requires mutual consent of both partners. The model is therefore useful as well for testing some of the conjectures about divorce law that have appeared in the literature.

In the fully transferable world of Becker et al., it doesn't matter which law is in force. Divorce occurs if and only if the spouse with an outside opportunity generates higher surplus outside the marriage than with the current partner. Thus when all women work, the divorce law is irrelevant.

Not so when there are some non-career women: when there is no longer transferable utility, in particular for the couple with a nonworking wife, the classical Coasian logic breaks down. The logic of whether or not a divorce occurs is similar to the logic of auctions with liquidity constrained bidders (Che and Gale, 1998). For instance, a married woman and a single woman might "compete" for the first woman's husband. A nonworking woman is financially constrained relative to a working woman and is therefore disadvantaged relative to a working woman. Consequently, a married working woman may be able to inefficiently retain her husband in the face of a proposal from a nonworking woman, and a married nonworking woman may not be able to match the offer made to her husband by a single working woman, which would result in inefficient divorce.

Under unilateral divorce, there will be both inefficient divorces, as husbands of nonworking women are easily wooed away by working women who can easily outbid the nonworking wife, even if the total surplus of the original marriage is high, and inefficient failures to divorce, as working married women outbid nonworking single women who propose to their husbands, even though the total surplus that would be created if husbands were to leave their working wives and marry nonworking women may be large. The latter effect (but not the former) is present as well under consent law, so there will be too little divorce, and in particular there will less divorce than there is under the unilateral law. The difference in divorce rate is largest when the level of FLFP is zero percent and declines as the level of FLFP approaches 100%.

As others (Wolfers, 2006; Rasul, 2006) have noted, when a state switches from consent to unilateral divorce (as most have, beginning in the early 1970s) there ought to be a sudden increase in divorces as the economy moves from the lower consent-divorce FLFP-divorce rate curve to the higher unilateral-divorce curve; over time, as FLFP increases, divorce rates will eventually tend to decline

This suggests a possible efficiency explanation for the nearly universal switch in US state divorce laws from consent to unilateral. The point has been made that consent divorce is transactionally very costly, while unilateral divorce is much less so (e.g. Hakim, 2006). The above discussion implies that when FLFP rises enough, divorce rates are little affected by which form of divorce law is in place, and it become efficient to avoid the transaction costs of consent divorce and impose unilateral divorce law instead. Thus for instance the recent push in New York (a high FLFP state) to adopt unilateral divorce will likely lead only to a small initial increase in divorce, with little long-run impact on divorce rates there.

Under either form of law, the model in NNO predicts an inverted-U relationship between FLFP and the rate of divorce. When initially FLFP is low, increases in FLFP lead to increases in the divorce rate, just as earlier studies have suggested. But there is a turning point, above which further increases in FLFP lead to *declines* in divorce. The model therefore accords well with the evidence presented above.

The intuition is that at low levels of FLFP, introducing a small number of working women destabilizes the nonworking women's marriages (the working women are unlikely to meet men already married to working women, so the fact their marriages are more stable has little effect), thereby raising the divorce rate. At high levels of FLFP, adding more working women increases the average stability of marriages, while the destabilizing effect on nonworking women's marriages is second order (since there are so few of them), and the divorce rate falls.

The flexibility effect could then account for the entire non-monotonic time series relationship between FLFP and divorce rate. But this places considerable import on the remarriage market, and there are reasons to believe other factors (including the many that have been suggested in the literature) may have been responsible for the rise in divorce from the 1960's to the 1980's. We will discuss some of them in the conclusion.

5. Conclusion

Why then has there been a decline in divorce rates since the 1980s? One possibility is that the effects identified elsewhere in the literature simply ceased to be operative after the mid 1980s. For instance, technological advances (machines to do household work, contraception) that might have made marriage relatively less valuable (and more fragile) might have diffused more or less completely by the mid-1980s. This by itself doesn't explain the decline, though, so we are led to invoke other effects (such as the two we have identified).

The selectivity effect – which as we have pointed out, is really two effects working in opposite directions – would have become more important simply as more as households become two-earner. However, this assumes that the net effect is indeed to reduce divorce rates. Moreover, it has a harder time explaining reduced divorce rates for two-earner households that happen to marry early.

Finally, there are reasons to believe that the flexibility effect has strengthened recently in the face of two further trends. Observe first that as partners in a marriage have equal incomes, there is greater transferability over a range of surplus levels that is proportional to the total income (it appears reasonable that the level of public good might vary positively with total income). Thus as gender earnings gaps have closed (this trend also began in the mid 1980s), the amount of transferability within two-earner households would have increased. Our preliminary explorations indicate that gender earnings gaps, which vary across US states, are positively correlated with divorce rates, as the flexibility model would suggest.

A second trend, apparent since the 1960s, and one that has been particularly strong in the US compared to Europe, is the increased availability in the market of goods that were formerly produced in the home (Freeman and Schettkat, 2005). In the Appendix, we illustrate the simple economics of monetary versus in-kind transfers in the context of a simple household production model. One implication of such a model is that the greater efficiency of monetary transfers depends in part on all (or enough) commodities being available for purchase in the market. Otherwise, the two-earner household will have a smaller transferability advantage relative to the one–earner household because while the second earner generates monetary income, she forgoes household production, and those goods are not available for purchase in the market.

So in the early 1960's when the market did not effectively supply many household-producible goods, transferability would not have differed so much across household types. With "marketization" of household goods, the advantage of money as a means of surplus transfer has increased, and career women have become better able to make these transfers – thus the flexibility effect has become increasingly powerful over time. This in turn may have contributed to the downturn in divorce. We hope to investigate this channel empirically in future drafts.

Appendix 1. Data

We use the following data sources:

Marriage and divorce rates: Vital Statistics of the United States (several years). Marriage and divorce rates used for most states are for 2001. For California and Colorado, 1990 divorce rates (the last year for which information is available) are used instead. Divorce rates for Indiana and Colorado are not available after 1980.

Labor force participation, education, race, and income: 2000 Census (IPUMS, 1% sample). The sample is restricted to working-age population (16-64 years old), not living in group quarters (GK=1). All state-levels averages and medians (for income) are population-weighted.

Age at first marriage: U.S. Census Bureau's American Community Survey 2002-2003.

Religion: For all states except Alaska, District of Columbia, and Hawaii, the data is taken from Beliefnet.com (http://www.beliefnet.com/politics/religiousaffiliation.html). Data was compiled by Professor John Green, Ray C. Bliss Institute of Applied Politics, University of Akron, Beliefnet contributing editor. The cited sources are 2004 National Surveys of Religion and Politics, 2001 American Religious Identification Survey, 2000 Glenmary Religious Congregations and Membership in the U.S., United States Census, 2002 Gallup Polls." For Alaska, District of Columbia, and Hawaii, data is produced by the Association of Religion Data Archives (http://www.thearda.com/mapsReports). According to the site, "data [was] collected by representatives of the Association of Statisticians of American Religious Bodies (ASARB)" in 2000. Note that "While quite comprehensive, this data excludes most of the historically African-American denominations and some other major groups."

"Gender concentration" in industries/occupations: Percentage of working women in industries, occupations, and industry-occupation cells where the state-level ratio of women to men is less than 50%. Based on Census data (IPUMS, same samples as above), using 1950-adjusted industry and occupation codes.

Appendix 2. The Flexibility Effect – Transferability with Working and Non-working Partners

Suppose the two partners have Cobb-Douglass utilities over private goods of $h^{1/3}m^{1/3}l^{1/3}$ where h is a household-producible good that can also be purchased on the market, m is a market produced good, and l is leisure. Each partner is endowed with a unit of leisure, the market wage is w, and p_h and p_m are the market prices of h and m, which the partners take as given.

If both partners work, it is straightforward to see that a household's Pareto efficient consumption has them both consuming multiples of the vector $q = (\frac{w}{3p_h}, \frac{w}{3p_m}, \frac{1}{3})$; utility is then proportional to q for each partner, and a transfer of utility from one to the other occurs along a frontier of negative unit slope. The maximum private good utility one partner could have is $u^2_{\max} = 2\frac{w^{2/3}}{3p_h^{1/3}p_m^{1/3}}$.

If one partner does not work in the market, she can produce h at home; assume, to put the two kinds of households on an equal footing, that she produces w/p_h units of h for every unit of l she gives up. A feasible point on the household utility possibility frontier is $u^* \equiv \left(\frac{w^{2/3}}{3p_h^{1/3}p_m^{1/3}}, \frac{w^{2/3}}{3p_h^{1/3}p_m^{1/3}}\right)$; this is obtained with one partner working 2/3 of the time in the market and using the income to purchase m, while the other partner works 2/3 of the time at home to produce h (this point can also be achieved in the two-earner household with, for example each partner working 2/3 and buying the market basket $\left(\frac{w}{3p_h}, \frac{w}{3p_m}\right)$).

Now, if the working partner wants to transfer starting at any point on the frontier above u^*, he does this as in the two-earner household. But if the non-working partner wishes to transfer starting at u^*, for instance if she tries to give all of her utility to her husband, this would be accomplished by having her set $l = 0$ and giving all w/p_h of the h that she produces to her husband. He would then maximize $(h' + w/p_h)^{1/3} m^{1/3} l^{1/3}$ subject to $p_h h' + p_m m + wl = w$ and $h' \geq 0$ (if h' can be negative, then the his partner is effectively selling l at price w and therefore working!). The solution is $\left(\frac{w}{p_h}, \frac{w}{2p_m}, \frac{1}{2}\right)$, yielding $u^1_{\max} = 2^{-2/3}\frac{w^{2/3}}{p_h^{1/3}p_m^{1/3}}$, which is less than $u^2_{\max}$. The frontier is linear between $(u^1_{\max}, 0)$ and u^*. This resembles Figure 3.

Observe that if h has only imperfect substitutes on the market (for instance, is not at all available there), then in fact career women may be at no greater advantage in transferring to husbands than non-career women.

Since utility transfers are accomplished in different ways by the two household types, a model such as this would in principle enable one to investigate the effects of labor market participation on intra household allocations as they respond to shocks.

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