Marriage market imbalances and labor force participation of Canadian women

J. C. Herbert Emery · Ana Ferrer

Abstract   Several studies find a negative correlation between the ratio of males to females and measures of female labor supply in the US. This negative correlation has been interpreted as empirical support for the hypothesis that marriage market conditions influence intra-household allocation decisions. Given the similarity of cultures and of labor supply behavior of women in Canada and the United States, and the fact that they both experienced baby-booms at roughly the same time, any explanation for changes in female labor supply would be expected to hold for both countries. We test the prediction that marriage market conditions have explanatory power for Canadian female labor force participation (LFP) rates over the period 1971–2001. We find smaller marriage market effects for Canada than those found for the US but similar in magnitude to those found for the US Midwest.

1 Introduction

Several empirical studies for the US establish a correlation between female labor supply and measures of marriage market conditions like sex ratios and cohort

J. C. H. Emery · A. Ferrer (✉)
Department of Economics, University of Calgary,
2500 University Drive NW, Calgary, Alberta, Canada T2N 1N4
e-mail: aferrer@ucalgary.ca

J. C. H. Emery
e-mail: hemery@ucalgary.ca

dummy variables.[1] This correlation has encouraged the development of a literature examining the relationship between marriage market conditions, intra-household allocation of resources and female labor supply. If marriage market conditions are useful for understanding female labor force participation (LFP), then we should expect that the correlation between measures of the population sex ratio and female LFP can be confirmed with studies of countries other than the US. An earlier sociology literature provided mixed evidence of a correlation between population sex ratios and female LFP using cross-sectional data for a set of countries from the 1960s and 1970s.[2] More recently, Fukuda (2006) finds evidence of cohort effects associated with sex ratio differences for female LFP in Japan, albeit they are weaker than for the US. Descriptive studies of Canadian female LFP suggest an absence of an influence of marriage market conditions on female labor supply.[3]

Canada provides a useful case to evaluate the marriage market hypothesis for female labor supply. Given the similarity of cultures and of labor supply behavior of women in Canada and the United States, and the fact that they both experienced baby-booms at roughly the same time, any explanation for changes in female LFP would be expected to hold for both countries. We follow the econometric strategies of Grossbard-Shechtman and Granger (1998)—GSG here after—and Grossbard and Amuedo-Dorantes (2007)—GAD here after—to test the prediction that marriage market conditions have explanatory power for changes in, and levels of, Canadian female LFP rates for the period 1971–2001. Following the latter paper's estimation strategy we find that the signs of the estimated sex ratios effects are consistent with what models of intrahousehold allocations would predict but they are statistically insignificant. The estimated coefficients for the sex ratio variable are smaller than those found by GAD for the US but similar in magnitude to what they find for the US Midwest in some specifications. Counter to expectations but consistent with their estimates for the US Midwest, sex ratios have large negative effects for university educated women in Canada.

The next section discusses the intuition of the marriage market explanations for female LFP patterns and empirical evidence related to that literature. Section 3 presents the empirical strategy that we apply to Canadian data. Section 4 presents the data and empirical results. Section 5 concludes.

---

[1] See Grossbard-Shechtman and Granger (1998), Angrist (2002), Chiappori, Fortin and Lacroix (2002), Fukuda (2006), Grossbard and Amuedo-Dorantes (2007) and Siow (2008).

[2] Ward and Pampel (1985a, b) find that the population sex ratio (female population to total population) is significant for explaining female LFP from 1955 to 1975 in 16 developed nations after controlling for a variety of factors. Pampel and Tanaka (1986) use pooled cross sectional data for 70 countries for the years 1965 and 1970 and find that for developed nations, more females in the population is correlated with higher female LFP. South and Trent (1988) find that the sex ratio was not correlated with female LFP in a cross-section of 58 developed nations.

[3] Ostry (1968) and Galarneau (1994) observe the dramatic rise in female labor force participation in Canada was general to all age groups and unrelated to year of birth. The absence of birth cohort effects suggests no correlation between labor supply and marriage market conditions.

2 Marriage market conditions and female labor force participation

A general conception for how the sex ratio influences female labor supply is through the standard income effects of married woman's non-labor income. Becker (1973, 1981) views marriage as a long term contract between two individuals that produces valuable, though partially intangible outputs such as children, love, companionship, security, income from market work, and household goods from home production. Within marriage, one partner often specializes in market work while the other specializes in work in the home and a share of household resources is allocated to the partner providing labor in the home. When marriageable women are abundant relative to the number of marriageable men, women would have lower bargaining power in the marriage market.[4] This lowers the "price" that a bread-winner spouse must pay for labor dedicated toward household production, shifting resources and family structures in favor of men.[5,6] Women with lower resources allocated to them within marriage, ceteris paribus, have a lower reservation wage for labor market participation resulting in a higher likelihood of working for pay.

North American women have typically married men who on average are 2 years older than themselves, so women born during periods of rising births are disadvantaged in the marriage market.[7] Marriage market conditions have varied over the century, most notably with the baby-boom (1946–1965) and subsequent baby bust (1966–1980). The predicted negative correlation between marriage market conditions as proxied by the sex ratio (males to females) and female LFP can be tested following the participation rates of women during these times of change in the marriage market. Women born during periods of increasing numbers of births faced tighter marriage market conditions and should have been more likely (*ceteris paribus*) to participate in the labor force than women born during periods of non-increasing numbers of births. For example, women born in the 1960s when the number of births was falling may have been less likely to participate in the labor force than women born during the period of rising numbers of births in the 1950s since they would have encountered higher intra-family transfers for spousal labor.

GSG (1998), Fukuda (2006) and GAD (2007) find empirical support for these predictions in their analyses of changes in female LFP since 1965.[8] Using different

---

[4] See Angrist (2002), Chiappori et al. (2002), Choo and Siow (2006) and Siow (2008).

[5] Grossbard-Shechtman (1984) develops a "demand and supply" model where the marriage equilibrium of a monogamous society is characterized by a quasi-wage received by the wife for spousal labor at which the amount of spousal labor a woman wants to supply equals the amount of spousal labor demanded by her mate.

[6] Angrist (2002: 998) also argues that "because sex ratios affect the *likelihood* of marriage, they may affect activities that complement or substitute for economic dependence on a spouse. For example, women who expect to marry need to worry less about developing an independent means of support."

[7] Grossbard-Schechtman (1984) highlights that American women have typically married men who on average are two years older than themselves This age pattern for first marriages is also true for Canadian women (Statistics Canada 2007). Foot and Stoffman (1996) and Beaujot and McQuillan (1982) suggest that while marriage squeezes may explain changing marriage patterns, marriage patterns are not exogenous to marriage market conditions. Neither of these studies link marriage market conditions to labor force participation.

[8] This negative correlation has been documented for other periods in the US as well. See Angrist (2002).

econometric approaches, the studies show that after controlling for measures of fertility, educational attainment, wages and incomes, women facing adverse marriage market conditions had significantly larger changes in LFP rates. In addition, women born in a period of declining births after the early 1960s had smaller changes in LFP than previous cohorts.

There are alternative models that could also explain cohort effects and sex ratio effects. Ward and Pampel (1985a, b) argue that the population sex ratio matters not through supply channels (e.g. reduced marriage opportunities pushes women into the labor market) but through labor market "competition" effects. When the sex ratio declines, there are insufficient males to fill the demand for male workers and employers resort to hiring women. Easterlin's (1980) relative income hypothesis provides another alternative to the marriage market explanation for why the LFP rates of women born into different cohorts have different patterns of LFP. The post World War II baby-boom generation grew up with relatively high levels of consumption, but faced extremely competitive labor markets due to the size of their generation, hence they faced low wages relative to their consumption expectations. This led to the rise of dual earner households and lower fertility levels. In a similar vein, Goldin (1990) ascribes changes in LFP across cohorts of American women over the last century to changes in fertility behavior. For example, young married women in the 1920s and 1930s were more likely to work early in their marriages than were women from earlier birth cohorts. Goldin suggests that these women were more involved in the labor force because they were having fewer children, and on average they had higher levels of education since many women working in the 1920s and 1930s completed high school. In the 1940s and 1950s these same women were the older married women who moved into the labor force (Rosett 1958).

GSG (1998) argue that cohort effects found after controls for income and fertility have been included cannot be explained by Easterlin's relative income hypothesis. Given that they find cohort and sex-ratio effects after controlling for income, education and fertility, the relative income hypothesis does not remain a credible rival to the marriage market model. Similarly for Goldin's (1990) cohort based arguments, it is not clear how she would explain cohort effects that remain after controlling for fertility and education.

3 Empirical strategy

We apply both of the GSG (1998) and GAD (2007) empirical methodologies to Canadian data to determine whether marriage markets are significant determinants of the changes in female LFP. We analyze individual labor market participation using micro data from the Family Census of Population (1971–1991). We are not able to use micro data from the 1996 or 2001 Family Censuses of Population since an individual's age is reported only by 10-year age groups so we cannot assign an individual to a 5-year birth cohort and corresponding sex-ratio. Given the limitations on our sample period for the micro data, we also perform the analysis using aggregate information from all available Census years (1971–2001). This aggregate approach enables us to assess the robustness of the micro data findings.

Women participate in the labor force if their market wage exceeds their reservation wage (w*). Higher market wages for women increase the likelihood that reservation wages will be exceeded by the market wage rate, enticing women into the labor force. Higher shares of household resources allocated to the wife for spousal labor, on the other hand, will raise w*, reducing the likelihood that a woman will choose to participate in the labor force. The wife's share of household resources will depend on the relative supply of marriageable women, or more generally, marriage market conditions.

An empirical model for LFP can be specified as:

$$P_{it} = c_0 + c_1 w_{it} + c_2 \text{MM}_{it} + c_3 w_{it}^h + c_4 F_{it} + c_5 t + e_{it} \tag{1}$$

P_{it} equals one if female i participates in the labor market at time t, reflecting that the market wage exceeds the reservation wage, and 0 otherwise; w_{it} is female i's wage; MM_{it} is a measure of marriage market conditions for female i at time t; w_{it}^h is a measure of the female i's husband wages, F_{it} is a measure of fertility for female i at time t, and t is a time trend variable introduced to control for time-varying factors such as economic conditions. Depending on the assumption made for the functional form of the error term e_{it}, (1) can be estimated by OLS, Probit or Logit methods.

GAD analyze the effects of marriage market conditions on female LFP for the US using this empirical strategy and cross-sectional data at 5-year intervals between 1965 and 2005 for the US overall and for four regions of the US The authors use sex-ratios calculated at a regional level to measure marriage market conditions. The sex ratio variable is defined for 5-year birth cohorts and it is the ratio of men aged 22–26 to women aged 20–24 when the women of a given birth cohort were 20–24 years of age.[9] The sex ratio value is specific to birth cohorts and has the same value for that birth cohort in all time periods.[10]

The probability that a given woman will participate in the labor force described in (1) can be aggregated to generate the following participation model for age groups:

$$P_{jt} = c_0 + c_1 w_{jt} + c_2 \text{MM}_{jt} + c_3 Y_t + c_4 F_{jt} + c_5 t + e_t \tag{2}$$

where j represents a given age group. Therefore, P stands for the participation rate of age group j at time t, Y_t becomes a measure of aggregate income (such as the real GDP per capita) included to account for the general state of labor demand, and the remaining variables are the aggregate (over group j) counterpart of those defined in Eq. 1.

A number of variables on the right-hand side of Eq. 2 are not necessarily exogenous to female LFP. For example, higher GDP may reflect higher demand for

---

[9] For some birth cohorts GAD calculate the sex ratio for ages 25–29 since only decennial census data is available. For example, the sex ratio for the 1946–50 birth cohort is calculated with the number of women aged 20–24 in the 1970 Census while for the 1941–45 birth cohort, GAD use the number of women aged 25–29 in the 1970 Census.

[10] The idea that sex ratios specific to birth cohorts are constant over time is open to criticism. For example, as individuals marry at higher ages, or perhaps for second marriages, differences in the ages of husbands and wives may be greater. This would suggest that how sex ratios are calculated should change over the life-cycle.

labor in the economy, but also that more labor resources are devoted to measured market production rather than non-measured production in the home. To reduce the effect of some of the unmeasured factors that enter into the calculation of residual correlations, Eq. 2 can be differentiated with respect to time:[11]

$$\frac{\partial P_{jt}}{\partial t} = \tau_0 + \tau_1 \frac{\dot{w}_{jt}}{w_{fjt}} + \tau_2 \frac{\dot{sr}_{jt}}{\partial t} + \tau_3 \frac{\dot{Y}_t}{Y_t} + \tau_4 \frac{\dot{F}_{jt}}{F_{jt}} + \tau_5 + u_t \tag{3}$$

where $\frac{\dot{w}_{jt}}{w_{jt}}, \frac{\dot{F}_{jt}}{F_{jt}}, \frac{\dot{Y}_t}{Y_t}$, and $\frac{\dot{sr}_{jt}}{sr_{jt}}$, are the growth rates of the corresponding variables.

To estimate (1), following the micro data approach in GAD, we use information from the Family Census of Population (1971–1991) since it includes information about husband's wages and household fertility choices. Data for the aggregate approach used by GSG includes age group specific information on wages, education and LFP, obtained from various Censuses of Population (1971–2001). Other aggregate data regarding the Consumer Price Index, Gross Domestic Product, and fertility rates come from Statistics Canada. For both methodologies, our cohort level sex ratios are calculated using Census population counts, and for some birth cohorts prior to 1971, annual estimates of population by age from Statistics Canada. For this reason, we are able to use sex ratios calculated at age 20–24 for all 5-year birth cohorts.[12]

Information from the 1981 to 2001 Censuses regarding education, LFP and wages, is consistently defined and collected through the years. Data from the 1970s, however, require some additional consideration. The main obstacle is that the 1976 Canadian Census did not collect information on earnings or income. We were able to construct an aggregate measure of wages for 1976 from the *Survey of Consumer Finance (SCF),* which we use in the aggregate model. In addition, data on aggregated fertility for the year 1975 is actually from 1974, the last year that this information is provided in the Historical Statistics of Canada (catalogue 11-516-XIE). In the individual model, the lack of information regarding wages or household income is more problematic. We use an aggregate measure of male and female wages, by province and age group, obtained from the *SCF.* This is similar to of GAD (2007). We have checked that the results in level regressions are robust to the inclusion of an indicator for the year 1976. Further, in the aggregate case, we also checked the robustness of our results to the exclusion of data for 1971 and 1976. Since identification of marriage markets rests on recognition of cohort effects, any such effect present in our full range data should remain in a data set that restricts the sample to the years 1981–2001. Therefore, we are confident that our results in the aggregate model are not driven by the 1976 data.

---

[11] Given our limited degrees of freedom for this study, we do not include lagged variables to correct for this problem, but we do estimate a variety of reduced form specifications to assess the sensitivity/robustness of our results.

[12] Prior to 1971, Canada had decennial Censuses. Where GAD chose to calculate sex ratios for 2–5 year birth cohorts per Census at ages 20–24 and 25–29, for Canadian women born before 1946 we calculate the sex ratio based on population counts at ages 20–24 in all cases. In Census years, we use the actual Census population counts by age and for years between Censuses, we use Statistics Canada's estimates of population counts that are interpolations from actual Census counts.

Finally, sex ratios may measure true marriage market conditions with error due to factors like immigration.[13] In particular, if immigrants are more likely to marry other immigrants (Angrist 2002), and if immigrants are more likely to be men, it could be argued that we are overestimating the relevant sex ratio for native born women. An issue for our study in comparison to the US studies is that relatively higher immigration levels into Canada than to the US after World War II may mean that the cohort level sex ratios are a poorer proxy of marriage market conditions in Canada. Although there is no information to construct a sex ratio for immigrants prior to 1972, we have examined the immigrant sex ratio for which information is available and find that the immigrant sex ratio is relatively stable and generally less than one, even for more recent cohorts. Table 1 presents two alternative measures of the sex ratio; the first measures the sex ratio for birth cohorts of women at ages 0–4 which will primarily reflect the influence of numbers of births, while the second is the sex ratio for the same birth cohort at ages 20–24. The difference between the two measures of the sex ratio will reflect the influence of differential mortality between sexes and net migration. From the comparison of these two sex ratio measures, we feel that it is unlikely that we are overestimating the relevant sex-ratio in our analysis due to immigration. Between the two measures of the sex ratio in both countries, there are differences in levels of the measures but changes in the sex ratio across birth cohorts seems to be largely influenced by changes observed from the numbers of births as represented by the younger age sex ratio.

In the individual model, we estimate a Probit equation like that specified in (1) for pooled census years over the period 1971–1991.

$$P_{it} = c_0 + c_1 w_{it} + c_2 \text{MM}_{it} + c_3 w_{it}^h + c_4 F_{it} + c_5 t + c_6 X_{it} + e_{it}$$

where the variables are as defined in (1), with X_{it} representing a vector of additional controls such as region of residence at the time of the survey, and a polynomial specification of age to control for life cycle effects. We estimate regressions measuring marriage markets using the cohort level sex ratio.

In the aggregate model, we approximate the dependent variable in Eq. 3 with the first differences of LFP rates for each age group j ($\Delta \text{LFPR}_{jt} \approx \frac{\partial P_{jt}}{\partial t} \approx P_{jt} - P_{jt-1}$) and consider LFP rates at 5-year intervals over the period 1971–2001 for the following 5-year age categories: 20–24, 25–29, 30–34, 35–39, 40–44, 45–49, 50–54. We estimate the following regression:

$$\Delta \text{LFPR}_{jt} = b_0 + b_1 \Delta \log w_{jt} + b_2 \Delta \log w_{jt}^m + b_3 \Delta \log \text{GDP}_t + b_4 \Delta \text{HiEd}_{jt}$$
$$+ \sum_{k=1}^{7} \theta_k \Delta \text{MM}_{kt} + v_{jt}$$

where t runs from 1971 to 2001; $k = 1,\ldots,7$ represents the change of cohort dummy variables; $\Delta \log w_{jt}$ ($\Delta \log w_{jt}^m$) is the change in the logarithm of the real female (male) annual average wage per 5-year period, for 5-year age group j, between t and $t-1$; $\Delta \log \text{GDP}_t$ is the change in logarithm of the real per capita GDP between t and $t-1$

---

[13] See Angrist (2002) and Siow (2008) for discussions of the problems with identifying marriage market influences on female labor supply using indirect measures of marriage market conditions like sex ratios or birth cohort dummy variables.

Table 1 Sex ratios by birth cohort at ages 0–4 and 20–24, Canada and the US

U.S. birth cohort	Canada birth cohort	(1) US Males 2–6/ Females 0–4		(2) Canada Males 2–6/ Females 0–4		(3) US Males 22–26/ Females 20–24		(4) Canada Males 22–26/ Females 20–24		(5) US midwest Males 22–26/ Females 20–24
		Level	Change	Level	Change	Level	Change	Level	Change	Level
1936–1940	1937–1941	1.04		1.014		0.95		0.979		1.02
1941–1945	1942–1946	0.988	−0.052	0.954	−0.060	0.91	−0.04	0.911	−0.068	0.93
1946–1950	1947–1951	0.975	−0.013	0.966	+0.012	0.87	−0.04	0.939	+0.028	0.93
1951–1955	1952–1956	1.000	+0.025	1.024	+0.058	0.95	+0.08	0.967	+0.028	0.99
1956–1960	1957–1961	1.014	+0.014	1.015	+0.009	0.97	+0.02	0.986	+0.019	1.04
1961–1965	1962–1966	1.062	+0.048	1.100	+0.085	1.03	+0.05	1.077	+0.091	1.09
1966–1970	1967–1971	1.087	+0.025	1.120	+0.020	1.06	+0.03	1.092	+0.025	1.16
1971–1975	1972–1976	1.099	+0.012	1.088	−0.032	1.07	+0.01	1.056	−0.036	1.11

Sources: The sex ratios in columns (1) and (2) are the number of males aged 2–6 to the number of females aged 0–4 when females of a given birth cohort were aged 0–4. The sex ratios in columns (3), (4) and (5) are the number of males aged 22–26 to the number of females aged 20–24 when the females of the birth cohort were aged 20–24. The US ratios in (1) are constructed from data from US Census Bureau, National Estimates by Age, Sex, Race: 1900–1979 (PE 11) http://www.census.gov/popest/archives/pre-1980/PE-11.html. The ratios for the U.S. in (3) and (5) are from GAD (2007, Table 1).The Canadian ratios in (2) and (4) are constructed from data from CANSIM Matrix 510026 which provides estimates of population by age for 1921–1971, and CANSIM Matrix 510001 which provides estimates of population by age for 1971–2007

(we also use the change in real GDP); ΔMM_{it} indicates a change in measured marriage market conditions from birth cohort $t - 1$ to birth cohort t. As with the micro data approach, we estimate regressions measuring marriage markets using the change in the sex ratio observed for age group j between t and $t - 1$. $\Delta HiEd_{jt}$ is the change in the percentage of "highly educated women" (defined as those with post-secondary education), for each 5-year age group, between t and $t - 1$. v_{jt} is Gaussian white noise term.

4 Marriage market effects in Canada

Table 2 shows that Canadian female participation rates have been steadily increasing for all age groups from 1971 to 2001. This trend has been well-documented. In addition, large increases have been observed for all age groups, with the exception of a decline in LFP rates of 20–24 year old women between 1991 and 1996. This decrease could be picking up a cohort effect related to favorable marriage market conditions for this group, or it could reflect the recession from 1989 to 1993 which impacted hard on men and women aged 20–24.[14] It may also reflect the rising participation in post-secondary education on the part of women in this age group (Emery 2005).

The common increasing tendencies of participation rate for all age groups between census years are period effects. Cohort effects, on the other hand, are identified by "above (or below) average" magnitudes of change in LFP rates between years as the cohort under observation in that age group changes. Cohort effects would show up as differences in behavior across birth cohort LFP rate-age profiles. Table 2 shows the LFP rates by 5-year age group for Canada and the US. In the US, there are significant differences between the LFP rate-age profiles of women born in 1941–1945 and 1946–1950, again consistent with a cohort effect. GSG (1998) highlight the change in LFP rates for 25–29 year old women (between 1970 and 1975) and for 30–34 year old women (between 1975 and 1980), which was around 12 percentage points in each case (See lower panel of Table 2). The cohort trailing the early baby-boom women had a lower increase in LFP rate for the 30–34 year age group. The cohort prior to the pre-baby-boom did not have as dramatic an increase over its preceding cohort. In short, for the US the behavior of women born in 1946–1950 appears to be different from women of preceding and trailing birth cohorts. This is not apparent in our data. The change in LFP rates for 25–29 year old Canadian women between 1970 and 1975 is around 9%, which is similar to that of previous cohorts for the same time period. Similarly, the change in LFP rates for 30–34 year old women between 1975 and 1980 are around 13%, and are in line with the changes in participation rates for other cohorts between 1975 and 1980 (See upper panel of Table 2). In contrast to the US, the large increases in participation rates are observed not only for women in the early baby-boom cohort (born 1947–1951). Preceding and trailing cohorts also display large increases in

---

[14] See Beaudry et al. (2000). In addition, the stagnation of female LFP rates for recent cohorts of Canadian women has also been documented by Beaudry and Lemieux (1999).

Table 2 Labor force participation rates of married females by 5-year age group Canada and the United States,[a] 1970–2001

	20–24		25–29		30–34		35–39		40–44	
	LFPR	Change	LFPR	Change	LFPR	Change	LFPR	Change	LFPR	Change
Canada										
1971	62.86		46.35		41.57		42.96		45.80	
1976	66.57	3.71	55.46	9.11	50.80	9.24	53.45	10.49	54.36	8.56
1981	77.16	10.59	68.13	12.68	63.47	12.67	63.31	9.86	64.38	10.02
1986	80.69	3.53	75.57	7.43	71.41	7.94	72.81	9.50	72.18	7.80
1991	81.11	0.41	79.39	3.82	77.63	6.23	79.30	6.49	80.56	8.38
1996	75.66	−5.45	78.24	−1.15	77.25	−0.38	78.11	−1.19	79.96	−0.60
2001	78.32	2.66	80.16	1.92	79.09	1.83	80.60	2.50	81.86	1.90
United States										
1970	57.70		45.2		44.7		49.2		52.8	
1975	64.10	6.40	57.3	12.10	51.9	7.20	55	5.80	56.7	3.90
1980	68.90	4.80	66.7	9.40	64.1	12.20	64.9	9.90	66.1	9.40
1985	71.80	2.90	71.4	4.70	70.3	6.20	71.7	6.80	71.9	5.80
1990	71.60	−0.20	73.8	2.40	73.4	3.10	75.5	3.80	77.6	5.70

[a] The U.S. rates are from Grossbard-Shechtman and Granger (1998), Table 1

LFP rates over the immediately preceding birth cohort. No birth cohort stands out in the Canadian data, except for the most recent cohorts.

4.1 The individual data approach

Table 3 shows the results from the individual Probit model using available Census data. The age of the individuals ranges from 25 to 44. Columns (1) and (2) show the effect of the sex ratio in a parsimonious specification. We find that as predicted by the marriage squeeze hypothesis, the sign of the sex ratio is negative and that non-married women have smaller sex ratio effects. However, the estimated coefficients for the sex ratio variable are not statistically significant. Columns (3) and (4) show that adding measures of fertility and female education levels has little effect on the estimated sex ratio coefficients for married females but result in a larger negative estimated effect for non-married women. The estimated coefficients for the rest of the covariates have the expected signs. Husband's wages are negatively associated with the LFP of married women in all cases. Age has a negative effect on participation in the simplest specifications since no control for fertility is included. The signs are reversed once the number of children in the household is taken into account (columns (3) and (4)). Finally, educated females participate more in the labor market.[15]

---

[15] Alternatively, we have tried a specification closer to that in GAD (2007) that omits husband's wages but includes husband educational levels. The results are similar to those shown in Table 3, with the sex ratio variable being small and positive but not significant. Having a husband with higher education reduces female LFP.

Table 3 Marginal effects on female LFP micro data for Canadian women aged 25–44, 1971–1991 (Standard errors in parentheses)

	(1) Married	(2) Non-married	(3) Married	(4) Non-married	(5) Married
Sex ratio[a]	−0.049 (0.031)	−0.018 (0.044)	−0.045 (0.032)	−0.057 (0.037)	0.023 (0.039)
Sex ratio * Post-secondary Education	–	–	–	–	−0.111 (0.068)
Sex ratio * University education	–	–	–	–	−0.338 (0.096)**
Husband wage ($Th.)	−0.001 (0.000)**	–	−0.002 (0.000)**	–	−0.002 (0.000)**
Age	−0.035 (0.003)**	−0.013 (0.004)	0.038 (0.003)**	0.024 (0.003)**	0.038 (0.003)**
Age 2 * 100	0.050 (0.004)**	0.020 (0.005)	−0.058 (0.004)**	−0.037 (0.005)**	−0.057 (0.004)**
Children under 14	–	–	−0.098 (0.001)**	−0.079 (0.001)**	−0.098 (0.001)**
Female education level					
Post-secondary	–	–	0.117 (0.002)**	0.085 (0.002)**	0.209 (0.053)**
University degree	–	–	0.170 (0.003)**	0.115 (0.002)**	0.341 (0.033)**
Region[b]	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes
Region * Year	Yes	Yes	Yes	Yes	Yes
Observations	179,211	58,181	179,211	58,181	179,211
R-Squared	0.09	0.06	0.15	0.23	0.15

[a] Sex ratio is the number of males aged 22–26 relative to women aged 20–24 in a birth cohort

[b] All regressions include controls for region of residence at the time of the survey (Ontario, Quebec, Maritimes, Prairies and BC), survey year and an interaction of the two to control for time varying regional macroeconomic effects

* indicates significant at 5%; ** significant at 1%

The estimated coefficients on the sex ratio variable indicate that between the birth cohort with the lowest measured sex ratio in column (4) of Table 1 (0.911 for 1941–1945) and the birth cohort with the highest measured sex ratio (1.092 for 1966–1970), marriage market conditions can explain a one percentage point difference in LFP. This magnitude of effect is smaller than GAD's estimated effect for the US but similar in magnitude to the estimated effect of the sex ratio in the US Midwest (See Table 6 in GAD 2007). As Table 1 shows, based on the cohort level sex ratios, Canada shows greater demographic similarity to the US Midwest than to the US overall.

More educated women are potentially less affected by variables measuring the tightening of marriage markets. GAD (2007: 252) expect that the strength of influence of marriage market conditions on female LFP diminishes with the level of a female's education. In particular, more educated women may have more egalitarian marriages which means that they have greater bargaining power within marriage and access to their spouse's income. We test this implication in Column (5), which reports estimated coefficients for a specification that includes an interaction of the sex ratio variable with an indicator for university educated and for non-university post-secondary educated. The estimated coefficient on the sex ratio variable in column (5) now refers to the effect of the sex ratio on the LFP of women with high school or less as their highest level of education. Contrary to expectations, females with the lowest level of education do not reduce their LFP when marriage markets are less competitive, whereas females with post-secondary education and university do. The estimated effects of the sex ratio for women with higher education levels are large, particularly for university educated women. For university educated women in Canada, a 10% change in the sex ratio decreases LFP by 3.2 percentage points.[16] This finding that the importance of sex ratios for explaining LFP increases with female education levels is consistent with GAD's results for the US Midwest region. They find that in the US West and Northeast, negative sex ratio effects are stronger for lower years of education, whereas in the US Midwest, the negative sex ratio effects are strongest for women with higher levels of education.[17]

4.2 The aggregate data approach

Table 4 reports the estimated coefficients of the regressions with changes in 5-year age group LFP rates as the dependent variable for all women, married women and university educated women. The results in Table 4 show that for all women, increases in female wages are associated with higher participation rates, while increases in male wages reduce female participation rates. The effect of GDP is

---

[16] The total effect for college educated women is −0.9 (0.023 − 0.111 = −0.088). For university educated it is −0.32 (0.023 − 0.338 = −0.315).

[17] This puzzling result suggests that there is a cohort effect negatively associated with the LFP of university educated females in Canada (and in these regions of the US). One plausible explanation is that education also enhances the productivity of household work. In areas where the difference between the value of market work and household work is small, an increase in the fraction of available men may reduce the LFP of (relatively scarce) educated females.

Table 4 OLS regression estimates of changes in female LFP 5-year age groups of Canadian women aged 20–54, 1971–2001 (Standard errors in parentheses)

Change in:	(1) All women	(2) Married women	(3) University educated women
Log of female wage	0.935***	0.959**	1.220***
	(0.344)	(0.379)	(0.390)
Log of male wage	−0.877***	−0.983***	−1.386***
	(0.309)	(0.341)	(0.350)
Log of per capita GDP	0.255	0.312	0.608**
	(0.247)	(0.272)	(0.280)
Fertility	−0.0004	−0.0002	−0.0004
	(0.0003)	(0.0004)	(0.0004)
% of women with post-secondary education	0.095	0.190	–
	(0.107)	(0.116)	
Sex ratio[a]	−0.160	−0.111	−0.076
	(0.146)	(0.161)	(0.166)
Constant	0.013	0.008	−0.014
	(0.017)	(0.019)	(0.02)
R-squared	0.26	0.26	0.37
N	42	42	42

[a] Sex ratio denotes the fraction of males aged 22–26 relative to females 20–24 in a birth cohort

* significant at 10%, ** significance at 5%, *** indicates significance at 1%

positive but not statistically significant, as are increasing education levels of Canadian females. Changes in fertility are insignificant throughout all models.

The results of the aggregate estimation follow those obtained with the micro specification. In this case, however, the influence of the sex ratio on LFP is weaker for higher educated women. The estimated coefficient suggests a sizeable influence of the sex ratio on female LFP. For example, between the 1956–1960 and 1961–1965 birth cohorts, the sex ratio increased from 0.986 to 1.077 (column (4) Table 2). This change in the sex ratio would result in a decrease in female LFP of between 0.5 percentage points for university educated women, 1 percentage point for married women and 1.5 percentage points for all women. While the negative sign of the estimated sex ratio coefficient supports the marriage market hypotheses for female labor supply, none of the estimated coefficients are statistically significant. This may reflect the difficulties of achieving statistical significance with such a small sample size.

5 Conclusions

Our analysis of Canadian female LFP from 1971 to 2001 provides evidence in support of the observation that marriage market conditions add to our understanding

of female labor supply in Canada. Our micro data estimations provide evidence consistent with that in GAD (2007) for the US Midwest, a region that is demographically and economically similar to Canada. Our aggregate data estimations indicate that sex ratio effects are consistent with theoretical predictions but they are not statistically significant. The results of our micro data estimations suggest that university educated females in Canada are more responsive to marriage market conditions than females with lower education levels. We do not find this result in the aggregate data GSG estimation approach, nor do we have an explanation for why university educated females in Canada and the US Midwest behave differently from University educated females in other regions of the US. GAD (2007: 264) do not have an explanation for this outcome either but they suggest that it could reflect that economic conditions in the Midwest meant that educated women faced a higher opportunity cost of not working than educated women living in other regions of the US.

References

Angrist, J. (2002). How do sex ratios affect marriage and labor markets. Evidence from America's second generation. *Quarterly Journal of Economics, 117*, 997–1038.

Beaudry, P., & Lemieux, T. (1999). Evolution of the female labour force participation rate in Canada, 1976–1994: A cohort analysis. Applied Research Branch HRDC W-99-4E.

Beaudry, P., Lemieux, T., & Parent, D. (2000). What is happening in the youth labour market in Canada? *Canadian Public Policy, 26*(s1), 59–83.

Beaujot, R., & McQuillan, K. (1982). *Growth and dualism: The demographic development of Canadian society*. Toronto: Gage Publishing.

Becker, G. S. (1973). A Theory of marriage: Part 1. *Journal of Political Economy, 81*, 813–846.

Becker, G. S. (1981) *A treatise on the family*. London: Harvard University Press.

Chiappori, P., Fortin, B., & Lacroix, G. (2002). Marriage market, divorce legislation and household labor supply. *Journal of Political Economy, 110*(1), 37–72.

Choo, E., & Siow, A. (2006). Who marries whom and why. *Journal of Political Economy, 114*(1), 175–201.

Easterlin, R. A. (1980). *Birth and fortune: The effects of generation size on personal welfare*. New York: Basic Books.

Emery, J. C. H. (2005). Total and private returns to university education in Canada: 1960 to 2000 and in comparison to other postsecondary training. In C. M. Beach, R. W. Boadway, & R. Marvin McInnis (Eds.), *Higher education in Canada*. Kingston: John Deutsch Institute, Queen's University.

Foot, D. K., & Stoffman, D. (1996). *Boom, bust & echo: How to profit from the coming demographic shift*. Macfarlane: Walter & Ross.

Fukuda, K. (2006). A cohort analysis of female labor participation rates in the U.S. and Japan. *Review of Economics of the Household, 4*(4), 379–393.

Galarneau, D. (1994). *Female baby-boomers: A generation at work. Focus on Canada. Catalogue no. 96-315E*. Ottawa: Statistics Canada.

Goldin, C. (1990). *Understanding the gender gap*. New York: Oxford University Press.

Grossbard, S., & Amuedo-Dorantes, C. (2007). Cohort-level sex ratio effects on women's labor force participation. *Review of Economics of the Household, 5*(3), 249–278.

Grossbard-Shechtman, A. (1984). A theory of allocation of time in markets for labor and marriage. *Economic Journal, 94*, 863–882.

Grossbard-Shechtman, S., & Granger, C. W. J. (1998). Women's jobs and marriage—baby-boom versus baby-bust. (English Translation). *Population, 53*, 731–752.

Ostry, S. (1968). *The female worker in Canada*. Ottawa: Dominion Bureau of Statistics.

Pampel, F. C., & Tanaka, K. (1986). Economic development and female labor force participation: A reconsideration. *Social Forces, 64*(3), 599–619.

Rosett, R. N. (1958). Working wives: An economic study. In T. F. Dernburg, R. N. Rosett, & H. W. Watts (Eds.), *Studies in household economic behavior*. New Haven: Yale University Press.

Siow, A. (2008). How does the marriage market clear? An empirical framework. *Canadian Journal of Economics 41*(4).

South, S. J., & Trent, K. (1988). Sex ratios and women's role: A cross-national analysis. *American Journal of Sociology, 93*(5), 1096–1115.

Statistics Canada. (2007). *Marriages 2003*. The Daily, January 17, 2007 http://www.statcan.ca/Daily/English/070117/td070117.htm, accessed February 2007.

Ward, K. B., & Pampel, F. C. (1985a). Structural determinants of female labor force participation in developed nations, 1955–75. *Social Science Quarterly, 66*(3), 654–667.

Ward, K. B., & Pampel, F. C. (1985b). More on the meaning of the effect of the sex ratio on female labor force participation. *Social Science Quarterly, 66*(3), 675–679.