Gender Dimensions of Changes in Earnings Inequality in Canada

by

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1. Introduction

The changes in the labour market outcomes of women are among the most important features of the changes in the Canadian labour market over the last two decades of the 20th century. During the 1980s, female labour force participation, which had climbed from below 30 percent in 1960 to 50 percent in 1980, continued to increase and appeared to reach a plateau around 60 percent in the early to mid-1990s.[1] During the 1990s, women's individual earnings made spectacular gains, increasing by more that 10 percent, while men's earnings were either stagnant or declining.[2] As a result, the female/male average earnings ratio of full-time full-year workers continued to climb in the mid-1990s, going from 67.6 percent in 1990 to 72.8 percent in 1996, when it seemed to reach a plateau.[3] In this chapter, we first document the economic progress of individual women: we report the impact of that progress on women's own earnings levels and dispersion, as well as on their earnings relative to men. Like Heisz et al. (2002), we find that over the last two decades male earnings inequality increased more than female earnings inequality did. Our numbers indicate that although the Canadian experience is quantitatively different from the American experience, it is not qualitatively different in these dimensions. But, we note that after a sharp increase in the early 1980s, family earnings inequality declined in the later part of the 1980s and increased again in the early 1990s.[4] By contrast, in the United States, there was no such decline in family earnings inequality in the mid to late 1980s (Daly and Valetta, 2000).

We thus next focus our analysis on the impact of the women's labour market gains on family market earnings inequality. Our aim is to offer an assessment as to whether the favorable changes in women's labour market outcomes actually translated into substantial changes in the welfare of women and their families. In the wake of increases in family

[1] After 1976, figures are from the CANSIM II Series - D980498 for all women 15 years of age and over. The participation rate of prime age women (25-44 years of age) was at a plateau of about 77 percent in the early to mid- 1990s. In the early 2000s, it had climbed again to the around 80 percent (CANSIM II - D980512).

[2] From 1980 to 2000, Heiz, Jackson and Picot (2002) report declines in the median earnings of full-time male workers, except among the 45-64 year olds. By contrast, they report increases in the median earnings of full-time female workers in all age groups except the 18-24 year olds.

[3] The ratios are from the CANSIM II Series - V1542068. The later 1990s, men's earnings began to climb again eroding women's relative progress and the ratio was still around. Drolet (2002) finds an higher number for the raw hourly wage ratio of 80 percent in 1997,which increases to 89 percent when accounting for a host of human capital, workplaces and other characteristics. We also find more favorable numbers from an hourly wage measure.

[4] Wolfson and Murphy (2000) also report substantial increases in family market earnings in Canada from 1985 to 1995.

dissolution (and decreases in family formation), have women's increased earnings merely compensated for their increased role in the financial support of the family? We note that one of the earliest Canadian family income inequality studies (Henderson and Rowley, 1977) had pointed out that changes in family size was one of the major reasons for increases in inequality over the 1965-1973 period. They argued that the decline in families with at least one male earner was an important contributing factor. Beach (1989) also classifies changing family composition as the dominant supply-side effect. In addition, it has been suggested that increases in assortative mating in the context of lower fertility rates could exacerbate family earnings polarization. For the United States, Juhn and Murphy's (1997) findings cast doubt on the notion that married women have increased their labour supply in recent decades to compensate for the disappointing earnings growth of their husbands. Instead Hyslop (2001) finds that the positively correlated labour market oucomes of working couples are due to permanent factors. Burtless (1999) estimates that 13 percent of the 1979-1996 increase in family equivalent income dispersion in the United States can be attributed to increases in assortative mating. Alternatively, family earnings inequality may still be largely driven by changes in male earnings inequality, as Daly and Valetta (2000) and Machin and Waldfogel (1994) have found for the United States and the United Kingdom, respectively.

On the demand side, changes in the wage structure stemming from skill-biased technological change remain the leading explanation for the important changes in U.S. wage inequality (Autor, Katz and Krueger, 1998; Katz and Autor, 1999). In Canada, these changes in the wage structure, in particular in the returns to education, are not as clearly identified and remain the subject of some controversy (Bar-Or et al., 1995; Beach and Slotsve, 1994; Beaudry and Green, 2000; Burbidge et al. 2002; Kapsalis et al., 1999). At best, they appear to be a phenomenon of the late 1990s and early 2000s following our period of study (Boudarbat et al., this volume). We nevertheless formally investigate the relative contribution to changes in family earnings inequality of the following factors: i) changes in the structure of male earnings; ii) changes in female labour force participation; iii) changes in the structure of female earnings; iv) changes in assortative mating; v) changes in family composition and living arrangements; and vi) changes in family characteristics, including the human capital variables. Our analysis therefore unfortunately neglects the impact of institutional changes, such as the decline in union density and the changes in the minimum wage, found to be important in Fortin and Lemieux (2000).

We use the distribution decomposition methodology pioneered by DiNardo, Fortin and Lemieux (DFL, 1996) and further expanded by Lemieux (2002) to allow for the separation of changes in the wage structure from changes in the productive characteristics of individuals. Our study thus parallels that of Daly and Valetta (2000) who also uses the DFL methodology to look at the impact of factors such as changes in male earnings, female labor force participation, family structure and family characteristics on family income inequality in the United States from 1968 to 1999. Daly and Valetta concludes that while changes in male earnings is the dominant factor in increasing family income dispersion, changes in family structure and rising female labour force participation also have substantial effects. One innovation in the present study is the ability to distinguish changes in the male and female wage structure from changes in human capital variables. This will enables us to better assess the respective impacts of supply-driven vs. demand-driven factors.

The remainder of the chapter is organized as follows. In section 2, we discuss some conceptual and practical issues in the measurement of individual earnings and family market earnings in Canada. In section 3, we describe the trends in women's labour market outcomes that we also compare to men's. We also document the changes in individual earnings inequality and in family earnings inequality across different family types. In section 4, we begin our analysis of the impact of the various explanatory factors identified above on family earnings inequality. Our main findings are summarized in section 5.

2. Data and Measurement Issues

Like most Canadian studies of individual and family earnings, we use Statistics Canada's Survey of Consumer Finances (SCF), the official source for income estimates prior to 1996, when much of the rise in women's labour force participation and labour earnings took place.[5] The publicly available Individuals Aged 15 and Over data files, used for the years 1981, 1982 and from 1984 to 1997, provide us with detailed information on individuals' characteristics, earnings, and labour market activity, but have no information on their union status. The publicly available Census Families data files, used for the years 1982, 1984, and from 1986 to 1997, provide information on the characteristics and earnings of the

[5] Other studies have used tax based data (Baker and Solon, 2003). See Frenette et al. (this volume) for more details on the impacts of using different data sources.

family unit, and the head and spouse in each family. The weights attached to each record in the files are created to reflect the non-random sampling design of the survey and are used throughout this study.[6]

We are concerned with the impact of changes in labour market outcomes, particularly those of women, on inequality in labour market earnings. Thus to characterize the labour market as accurately as possible we restrict the samples used throughout this paper to individuals aged 16-64, excluding the self-employed and those permanently unable to work[7]. We measure these individuals' annual earnings as income earned as wages and salaries from employment, expressed in 1992 dollars.[8] We note that Statistics Canada does not provide allocation flags to identify individuals with allocated wages. If the percentage of workers with allocated wages is important, this will bias downwards our measure of earnings inequality to the extent that there has been an increase in residual wage inequality over time (Baker and Solon, 2003). Comparisons with tax data (Frenette et al., this volume) indicate that discrepancies between the Gini coefficient from SCF data and from Tax data may be of the order of 10 percent. This caveat must be kept in mind with considering our Canada-U.S. comparisons.

To facilitate comparison with other studies (e.g. Morisette et al. (1994), Burbidge et al. (1997), Heisz et al. (2002)), we present first trends in individual's annual earnings using samples of full-time full-year workers. This sample selection was common in earlier studies when surveys did not collect information on the number of weeks worked a year or the number of hours worked. Given the sensitivity of inequality measures to outliers, we trim the sample of full-time full-year workers by eliminating observations reporting extremely low or high earnings, keeping individuals with annual earnings between $3000 and $200,000.[9]

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[6] The SCF employs a multi-stage stratified clustered probability sample. The weights are designed to reflect the sample design, differential response rates for households, among other things.

[7] We also exclude all observations with unclear labour force outcomes. That is, we exclude individuals who report being in the labour force but worked zero weeks, worked zero hours, or had zero annual earnings.

[8] The CPI is from the CANSIM II Series v737344. The questionnaire used in the survey asks the respondent "During the 12 months ending December 31, (year), what was your income from the following sources? 1. Wages and Salaries before deductions, including military pay and allowances." For non-respondents, income is imputed by Statistics Canada based on geographically-nearest respondent records possessing similar characteristics. However, Statistics Canada does not provide allocation flags to identify those individuals with allocated wages. If the percentage of workers with allocated wages were important, this would bias downwards our measure of earnings inequality to the extent that there has been an increase in residual wage inequality over time (Baker and Solon, 2003).

[9] One alternative to truncating the sample is to "winsorize" the data by, for example, setting the values of the observations in the bottom or top decile equal to the value of the observation at the 10^{th} or 90^{th} percentile. This techniques is more desirable than truncating if it is believed the extreme values are exaggerated versions of the

An important problem in using full-time full-year samples for gender comparisons is that among full-time workers, women's average number of hours worked a week are lower than men's are. Thus comparing the earnings of full-time full-year women and men confound labour supply issues with wage determination issues. To consider the latter issues, it is preferable to present trends in hourly wages. In the SCF, the usual hours of work are reported as the average hours worked in the three weeks prior to the individuals' interview, which takes place in the year following the year for which income is reported. By selecting individuals with at least on year of job tenure, we minimize the possibility of a job mismatch between earnings and hours of work. We construct an individual's hourly wage by dividing their annual earnings by annual hours (the product of reported weeks worked in a year and the usual hours worked per week). We also trim this sample to exclude individuals with hourly wages less than $1.50 or above $100 in $1992, a procedure commonly performed in U.S. studies to eliminate outliers (Gao, 2003). We can then present trends in hourly wages by using the sample of workers with at least one-year of job tenure.

The trends in inequality and earnings presented in the next section depend to some extent on the sample selection made. Our trimmed samples show less erratic changes in inequality measures, but also less amplitude especially in the early 1980s.[10] The trends presented appear to be robust to various age restrictions we investigated, although using a sample of men age 25-54 indicates larger increases in inequality among men in the 1990s than the broader sample used here.

Our principal focus of analysis is the distribution of family market earnings, where we use the SCF concept of family.[11] We restrict our sample to families whose heads are both aged 16-64, excluding families with self-employed family heads or family heads permanently unable to work.[12] Because families are of different size, we need to appeal to the concept of equivalent family earnings. A family's earnings are simply the sum of all family members'

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true values, but the true values lie in the tails. Truncating the sample is more desirable if the extremes are believed to simply be mistakes. See Angrist and Krueger (1999) for a thorough discussion of trimming techniques. Eliminating these extreme observations eliminates less than 1 percent of our sample.

[10] For example, for our female sample of workers with more than one year of job tenure, allowing for extremely low or high wages will result in measured increases in inequality among women in the early 1980s that our trimmed sample does not depict.

[11] The SCF defines a census family as either a husband and wife, or one parent, with their unmarried children. A single individual living alone or with unrelated individuals is a one-person census family. The analysis of inequality in more encompassing measures of family income can be found in Frenette et al. (this volume).

income earned as wages and salaries from employment, as reported in the SCF. Recognizing that economies of scale may be achieved in larger households, we adjust family earnings to compare families of different sizes. We follow Daly and Valetta (2000) in adjusting a family's earnings for the number of individuals in the economic family with the following formula,

$$Y = \frac{E}{F^{s}} \tag{1}$$

where Y is equivalent family earnings, E is the total family earnings, F is the number of individuals in the economic family. We set $\sigma=0.5$, the midpoint between the extreme values of 0 and 1.[13] Although this adjustment is widely used in the literature (e.g. Karoly and Burtless, 1995), there is no general consensus on what equivalence scale should be preferred when adjusting family earnings. Here we simply adjust family income for household size, thus abstracting from the fact that different types of families may have different needs. By contrast, Atkinson and Bourguigon (1987) suggest a two-stage procedure where the household's original income is first adjusted to accommodate differences in needs and then further adjusted for family size. Our emphasis on the comparison of living standards for different types of families follows the more recent literature (Ebert and Moyes, 2003, Davidson and Pendakur (1999)).

Other equivalence scales adjust family income according to the age structure of the family. For example, Statistics Canada uses an equivalence scale when deriving low income measures whereby the oldest individual in a family receives a weight of one, all other adults are weighted by 0.4 and all children (under age 16) are weighted by 0.3. The scale used by the OECD weighs the oldest individual by one, all other adults by 0.7, and children by 0.5. In Figure 1, we display trends in median log equivalent-family earnings and in the 90-10 differential that were calculated using various equivalence scales. We find that our simplified measure reproduces the same general trends of these other measures.

We generally consider four family types: 1) married with kids, 2) married without kids, 3) other families with kids and 4) other families without kids. Married with kids includes families with a head and spouse, married or living common law, who report the

[12] We also exclude all observations for which the labour force outcomes of either family head are unclear. That is, we exclude families whose head reports being in the labour force but worked zero weeks, worked zero hours, or had zero annual earnings.

presence of children under the age of 18 in the census family. Married without kids are similar to type 1 but do not report the presence of children. Other families with kids include those who are never married, separated, divorced, or widowed, who report the presence of children under the age of 18. Other families without kids include all other single individuals, never married, separated, divorced or widowed, who do not report the presence of children under the age of 18. In table 1, we report the relative shares of these different types of families from 1982 to 1997. There we separate never-married parents from the other types of unmarried parents given the importance of never-married mothers (87 percent of never-married parents) for policy purposes.

The most dramatic change in family composition is the large 7 percentage points drop in the proportion of families consisting of married couples with children and a similarly sizeable increase of 4 percentage points in the number of single individuals without children. Also significant, the proportion of families comprised of never-married parents has more than doubled over the 1980s and 1990s to reach 3 percent in 1997. Despite the fact that they represent only a small proportion of families, these families are more likely to have lower equivalent family earnings than married couples, hence an increase in this type of family places more families in the lower end of the earnings distribution.[14] Given these sizeable changes across the different family types, changes in inequality among all families will likely depend on the proportion of each family type in the total population of families

3. Trends in Women's Labour Market Outcomes and Family Earnings Inequality

Over the 1980s and 1990s, women in Canada have experienced significant changes in their labour market outcomes. One striking feature is that labour force participation rates among women increased substantially throughout the 1980s, although remaining roughly constant around 65 percent in the 1990s, as can be seen in Figure 2a. Beaudry and Lemieux (1999) who analyze both the rise and relative stagnation of female labor force participation from 1976 to 1994 in terms of macroeconomic effects, age effects and cohort effects find that the latter effects explain most of the phenomenon. While participation rates for men actually fell over this period, the change for men is not nearly as dramatic as for women. Between

[13] For example, setting σ=1 assumes that infinite economies of scale are achieved by a family, while setting σ=0 assumes that no economies of scale may be achieved

1981 and 1997, women aged 16-64 saw a 12 percentage point increase in participation rates while the comparable sample of men experienced a 3 percentage point decrease in participation rates.

The observed increase in female labour force participation can be largely attributed to increased participation of more educated, married women. In terms of education attainment, labour force participation rates increased most for women with university degrees or high school education in the 1980s, as shown in Figure 2b. In the 1990s, participation rates for women with university degrees remained constant around 82 percent for the decade while women with high school or less actually experienced declines in participation. More remarkable is the 21 percentage points increase in participation rates of married women with children displayed in Figure 3. Participation rates of married women without children have also increased, in line with the full sample of women. In contrast, participation rates of single women have remained fairly constant. However, single women with children are the only group of women who experienced an increase in participation in the 1990s, following a sharp drop in the late 1980s. The increases in participation among single women with children may be related to significant changes to social assistance programs in the mid 1990s. In all provinces, single parents experienced real reductions to social assistance entitlements in the mid-1990s and some (most notably Ontario) have introduced work requirements.[15]

Despite the remarkable growth in women's labour force participation rates, the potential impact of higher participation on family earnings is diminished by an accompanying increase in the proportion of women working part-time. The proportion of women working part-time increased by an average of 7 percentage points from 1981 to 1997 and now stands at 30 percent (see Figure 3b). Notably however, the proportion of married women with children working part-time has actually decreased slightly while increases were most significant for single individuals, including lone parent families.

In the 1980s and 1990s women made spectacular gains in earnings and wages relative to a general stagnation in men's earnings. Figure 4 presents a clear picture of women's relative wage gains by displaying the kernel density estimates of the distributions of log

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[14] Yet, because they represent such a small proportion of families, this yield too small sample sizes to be considered separately in our decomposition analysis.

[15] See Crossley and Curtis (2000) for a discussion of changes to provincial social assistance programs. As an example, using information provided by the National Council of Welfare (2003), there was a 25 percent real decrease in benefits to single parents (with one child) in Ontario between 1994 and 1997. Only in New Brunswick did real benefits to single parents increase (by 5 percent) over this period.

earnings and log wages for men and women in 1982 and 1997. There is a clear rightward shift in the distributions of log earnings and log wages of women. In comparison, the distribution of log wages for men shows movement from the middle to the lower end of the distribution, with a substantial widening of the distribution of men's log earnings. An important feature of the change in the distribution of log hourly wages for men is a shift of mass around the mode of the 1982 distribution to the mid-lower tail of the distribution in 1997. That kind of shift has been identified in Fortin and Lemieux (2000) to be the result of the decline in union density over that period.

As a result of women's relative earnings gains, women's median earnings and wages have been steadily catching up to men's. Figure 5 indicates that the female/male median earnings ratio has increased steadily over the 1980s and 1990s. As expected, the female/male hourly wage ratio for the sample of workers with one year or more of job tenure is consistently larger, reaching the high seventies (as a percent) in the late 1990s, than the annual earnings ratio for a sample of full-time full-year workers, which is in the lower seventies. This is due to the fact that despite their increased participation, even when women work full-time they continue to work fewer hours than men. Furthermore, as demonstrated in Figure 6, the difference in hours between men and women has remained fairly constant over the two decades and across education groups or samples.[16]

Interestingly, the substantial earning gains of women relative to men's has not been achieved at the expense of relatively increasing inequality among women. Figures 7 and 8 display the trends in log earnings and log wage inequality among men and women from 1981 to 1997, plotting five different measures of inequality for two samples of men and women. Note that each of the inequality measures is sensitive to somewhat different shifts in the earnings distribution. For example, the Gini coefficient is most sensitive to changes in the middle of the distribution[17].

Among men employed full-time and full-year, earnings inequality, displayed in Figure 7, increased most substantially in the early 1980s, with only moderate increases in inequality for the rest of the sample period. For example, between 1981 and 1989 the Gini

[16] The male-female difference in hours for the sample of full-time full-year workers ranges from a low of 3.09 hours/week in 1981 to a high of 3.81 hours/week in 1994. The male-female difference in hours for a sample of workers with more than one year job tenure ranges from 6.35 81 hours/week in 1989 to 7.14 81 hours/week in 1991(**?**).

[17] See Kesselman and Cheung (this volume) for a thorough discussion of inequality measures.

index increased by 10 percent, representing an annualized 1.1 percent change. These results are consistent with the results presented by Morissette et al. (1995) who also reports a rise in the Gini index of 10 percent between 1981 and 1989 for men employed full-time full-year. The 90-10 differential in log annual earnings increased by 16 log points between 1981 and 1989, representing a 2 percent annualized log wage change. The Canadian trends the 1980s, while not as large as in the United States, are consistent with U.S. trends. Katz and Autor (1999) report increases of 20 log points in the 90-10 differential in log weekly wages of full-time full-year American male workers from 1979 to 1987, which represents an annualized 2.5 percent log wage change, and annual increases of 1.6 percent in the Gini coefficient.[18] In the 1990s, the increases in inequality among full-time full-year men in Canada were more moderate, inequality continued to grow more slowly than in the United States. Between 1990 and 1997, the Gini increased by only 3 percent representing an annualized 0.4 percent increase while the increase in the 90-10 log wage differential show an annualized 0.7 percent log wage change. Picot (1998) finds similar annualized increases of 0.6 percent in the Gini coefficient computed for male paid employees from 1990 to 1995. For the United States in the 1990s, Katz and Autor (1999) report increases in inequality representing an annualized 1.2 percent increase in the Gini index and annualized log wage increases of 0.9 percent in the 90-10 differential for male full-time full-year workers.[19]

The trend in inequality among men (controlling for changes in hours of work) is slightly different. For our sample of men with at least one-year of job tenure, log hourly wage inequality increased substantially in the early 1980s, dropping for the last half of the decade and then increasing significantly through the 1990s. For the overall decade, the 90-10 differential in log wages showed an annual rate of change of 1.3 percent and the Gini index increased by an annualized rate of 1 percent. Again, this is similar to trends in the United States where Katz and Autor (1999) report an annualized change of 1.7 percent in the 90-10 hourly log wage differential.[20] Inequality among men increased moderately throughout the 1990s, as the 90-10 differential show annualized 0.7 percent rate of change and the Gini coefficient increased by an annualized 0.6 percent.

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[18] Calculated from Katz and Autor (1999), using a sample of males working full-time full year between 1979 and 1987 (Table 1), or Katz and Autor (1999), sample of full time full year non-agricultural workers earnings more than ½ of the contemporaneous Federal Minimum wage between 1980 and 1990 (Table 8).

[19] Annualized changes calculated from Katz and Autor (1999) using a sample of males working full-time full-year between 1995 and 1987 (Table 1).

[20] Katz and Autor (1999) Table 4.B, the figure is from the CPS-MORG.

Some portion of the difference between the trends in Canadian hourly wage and annual earnings inequality among men may be accounted for by increased polarization in hours of work, as argued in Morissette et al. (1994). For example, in the mid-1980s average weekly hours of more and less educated men began to diverge, as indicated in Figure 6c. Although inequality in men's annual earnings stagnates after this time, it is likely that more educated individuals experienced cuts to their hourly wages and compensated for this by working relatively more hours, resulting in very little change in the distribution of annual earnings. Consistent with this notion, Morissette et al. (1994) find that the distribution of weekly hours among men had widened through the 1980s.[21]

Trends in earnings inequality among women, displayed in Figure 8, are substantially different from that of comparable men, as noted by others (e.g. Picot et al., 1998). For men, the widening of the earnings distribution occurs both at the bottom and top of the distribution, as the time trends in 90-50 and 50-10 differentials follow the trend in the 90-10 differentials. For women, inequality increases more at the bottom of the distribution with the 50-10 differential trending up while the 90-50 mostly declines, the result is a rather stable, if not declining, 90-10 differentials.

Inequality in annual earning among women working full-time and full-year increased moderately through the 1980s. For this group of women, the 90-10 differential showed annualized log wage changes of 1.3 percent and the Gini increased by an annualized 0.9 percent, notably less than the increase in inequality among their male counterparts. Women in the United States experienced much larger increases in inequality over this period than their Canadian counterparts and than American men. Katz and Autor (1999) report increases in inequality among women working full-time and full-year representing annualized 3.1 percent log wage changes in the 90-10 differential of log weekly wages and an annualized 2 percent increase in the Gini index over the 1980s.[22] DiNardo, Fortin and Lemieux (1996) have implicated the declining real value of the minimum wages in the United States as a powerful explanatory factor for the very large increases in inequality among American women. The 1990s also saw much different trends in inequality for men and women working

[21] Towards the end of the 1990s, we see a dramatic decline in the hours worked among more educated individuals. The source of that change is not clear. A reporting change is a possibility since there have been changes in the hours of work question series in the questionnaire of the LFS. Changes in the EI eligibility, which became hourly based, is another possibility.
[22] Annualized changes calculated from Katz and Autor (1999), using a sample of females working full-time and full-year between 1979 and 1987 (Table 1).

full-time full-year. Women experienced a slight increase in earnings inequality up to 1995 followed by a decline, so that the annualized change from 1990 to 1997 was basically nil at 0.1 percent while changes in the Gini stood at an annualized 0.4 percent. In contrast, women working full-time and full-year in the United States saw continued increases in inequality into the 1990s, as the 90-10 differential increased by an annualized 0.7 percent and the Gini increased by an annualized 1 percent.[23]

Inequality trends for our sample of women with at least one-year of job tenure are even flatter than trends for our sample of women working full-time and full-year. In both the 1980s and 1990s, inequality increases in hourly wages among women with at least one-year of job tenure fared at an annualized 0.6 percent log wage changes in the 90-10 differential in log hourly wages. For the entire period, there was a slight decrease in the Gini coefficient.

Given these trends in female earnings inequality, we would expect, as suggested by Picot (1998), the increased participation of women in the labour market to have moderating effect on trends in family earnings inequality, somewhat offsetting the increases in men's earnings inequality. This would explain the Canada-U.S difference in patterns of family earnings inequality in the 1980s. The substantial increases in female earnings inequality in the mid to late 1980s in the United States were accompanied by slight increases in family earnings inequality, increases not found in Canada.

Figure 9 presents these trends in our measure of equivalent family earnings for all families, and the four different family types: 1) married with children, 2) married without children, 3) others with children, and 4) others without children. First, in terms of levels of equivalent family earnings, for all families in panel A, we find a moderate increase of 9 percent in median family equivalent earnings over the 1980s, followed by a general stagnation of earnings in the 1990s. For families of married couples in panel B and C, we find similar patterns of increases in median family equivalent earnings. For other families with kids (headed at 87 percent by women) in panel D, the measure of median family equivalent earnings is somewhat noisy, but does not indicate a consistent pattern of increase. Similarly, for other families without kids (53 percent of which are headed by men), median earnings have barely kept up with inflation. To the extent that these last two types of families

[23] Calculated from Katz and Autor (1999) Table 1, using a sample of women working full time and full year between 1995 and 1987.

comprise younger heads, these trends are similar to earnings trends for younger individuals described in Picot (1998), Kapsalis et al. (1999) among others.[24]

Turning to inequality measures, for all families, we find that family equivalent earnings inequality increased in the early 1980s, peaking in 1984, and then declined for the rest of the 1980s. From 1984 to 1989, the 90-10 differential in log family equivalent earnings decline by an annualized –2 percent log earnings changes, while the Gini declined at an annualized rate of –0.6 percent. By contrast in the 1990s, earnings inequality among all families increase as the 90-10 differential increased by an annualized 0.8 percent log earnings change and the Gini by 0.4 percent through the 1990s. The increase in inequality among different types of families in the 1990s was most significant among married couples, which constitute roughly 55-60 percent of all families. It was most notable among married couples without children who saw an annualized 1.7 percent increase in the 90-10 differential. In contrast, inequality among other families (with and without children) was either constant or declining over the 1980s and 1990s.

In summary, the trends in family equivalent earnings inequality over the 1980s and 1990s appear to be dominated by trends among couples with trends among singles playing smaller offsetting effects.

3. Accounting for Changes in the Distribution of Equivalent Family Earnings.

a. Decomposition Methodology

We begin with a description of the procedure used to decompose changes in the density of log equivalent family earnings in terms of the six factors identified earlier. This description generally follows the notation of DiNardo, Fortin and Lemieux (1996). Taking the logarithm of equation (1), we get $y_t \equiv \log(Y_t) = \log(E_t) - 0.5\log(F_t)$, where family earnings, E_t, is the sum of the male earnings, E_{Mt}, and female earnings, E_{Ft}, when applicable,

$$E_t = E_t^M + E_t^F + u_t \qquad \text{where} \tag{2}$$

$$E_t^M = \exp(X_{Mt}b_{Mt} + e_{Mt} + w_{Mt})$$
$$E_t^M = \exp(X_{Ft}b_{Ft} + e_{Ft} + w_{Ft})$$

[24] Forty-six percent of other families are headed by (single) individuals under the age of 35. Seventeen percent are headed by individuals under the age of 25. This compares to only 25 percent of married couple families

where X_M is a vector of the male/husband characteristics, X_F is a vector of the female/wife characteristics, $\ _M$ and $\ _F$ represent the residual male and female earnings not explained by their characteristics, w_{Mt} and w_{Ft} represent the log of weeks worked by male/husband and female/wife, and   represents the residual family equivalent earnings, which may include children's earnings, that are not explained by the female and/or male characteristics.[25]

To understand the decomposition procedure, it is useful to view each family observation as a vector (y, $X_{MT}\ \ _M$, L, $X_{Ft}\ \ _F$, A, C, X, t) made up of the family's log equivalent family earnings (y), the men's predicted earnings ($X_{Mt}\ \ _M$), the female labour force participation choice (L), the woman's predicted earnings ($X_{Ft}\ \ _F$), the degree of assortative mating in the family (A), family composition (C), the vector X=[X_M , X_F] of family characteristics that may be partitioned as the male/husband characteristics (X_M) and the female/wife characteristics (X_F), and a date t. Note that we therefore abstract from changes in residual earnings inequality.[26] Each family observation belongs to the joint distribution F(y, $X_{Mt}\ \ _M$, L, $X_{Ft}\ \ _F$, A, C, X, t) and the joint distribution of log equivalent family earnings and characteristics at one point in time is the conditional distribution F(y, $X_{Mt}\ \ _M$, L, $X_{Ft}\ \ _F$, A, C, X | t). The density of log equivalent family earnings at one point in time, $f_t(y)$, can be written as the integral of the density of equivalent family earnings conditional on a set of family attributes and on a date t_Y, $f(y|$ $X_{Mt}\ \ _M$, L, $X_{Ft}\ \ _F$, A, C, X, $t_Y)$, over the distribution of family attributes at a date t:

$$
\begin{aligned}
f_t(y) = \iiint\iiint f\left(y \mid X_{Mt}b_M, L, X_{Ft}b_F, A, C, X, t_y = t\right) \\
dF\left(X_{Mt}b_M \mid L, X_{Ft}b_F, A, C, X, t_{X_{Mt}b_M|L,X_{Ft}b_F,A,C,X} = t\right) \\
dF\left(L \mid X_{Ft}b_F, A, C, X, t_{L|X_{Ft}b_F,A,C,X} = t\right) dF\left(F \mid A, C, X, t_{F|A,C,X} = t\right) \\
dF\left(A \mid C, X, t_{A|C,X} = t\right) dF\left(C \mid X, t_{C|X} = t\right) dF\left(X \mid t_{X_M,X_F} = t\right)
\end{aligned}
$$

$$
= f\left(y; t_y = t,\ t_{X_M b_M|L,X_F b_F,A,C,X} = t,\ t_{L|X_F b_F,A,C,X} = t,\ t_{F|A,C,X} = t,\ t_{A|C,X} = t,\ t_{C|X} = t, t_X = t\right). \qquad (3)
$$

The estimation of counterfactual densities for the decomposition involves the combination of different "datings", the last line introduces the notation that accounts for these. For example,

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headed by individuals under age 35 and 3 percent of married couples headed by individuals under age 25.

[25] Note that we add w in the above because we actually estimate the log weekly earnings of the head/spouse and then convert this back to log annual earnings.

[26] As explained earlier, the empirical evaluation of residual wage inequality in the SCF may be difficult.

$$f\left(y;t_y=97,\ t_{X_M b_M|L,X_F b_F,A,C,X}=97,\ t_{L|X_F b_F,A,C,X}=97,\ t_{F|A,C,X}=97,\ t_{A|C,X}=97,\ t_{C|X}=97,\ t_X=97\right)$$

represents the actual density of equivalent family earnings in 1997, and

$$f\left(y;t_y=97,\ t_{X_M b_M|L,X_F b_F,A,C,X}=82,\ t_{L|X_F b_F,A,C,X}=82,\ t_{F|A,C,X}=82,\ t_{A|C,X}=82,\ t_{C|X}=82,\ t_X=82\right)$$

represents the density of equivalent family earnings that would have prevailed in 1997 had the distribution of all family attributes remained as it was in 1982.

We begin the primary order decomposition by adjusting the 1997 equivalent family earnings distribution for changes to men's wage structure. That is we want to find the counterfactual density

$$f\left(y;t_y=97,\ t_{X_M b_M|L,X_F b_F,A,C,X}=82,\ t_{L|X_F b_F,A,C,X}=97,\ t_{F|A,C,X}=97,\ t_{A|C,X}=97,\ t_{C|X}=97,\ t_X=97\right) \quad (4)$$

We first estimate male weekly earnings in 1982 and 1997. The econometric model used to describe male log weekly earnings, e_t^M, may be stated as

$$e_t^M = X_{Mt} b_{Mt} + e_{Mt} \quad (5)$$

where X_{Mt} is a vector of characteristics specific to the male head of the family. The vector of characteristics X_{Mt} includes a quadratic in age, and dummy variables indicating province of residence, education level, and full time or part time job status.

Male log weekly earnings are adjusted for wage structure by applying the 1982 parameter estimates ($_{82}$) to 1997 characteristics of men in the labour force and adding the residuals of the 1997 male log weekly earnings regression. That is,

$$e^M_{X97b82} = X_{M97} b_{M82} + e_{M97}. \quad (6)$$

These estimates are then used to adjust the equivalent family earnings of all families with a male head (in the labour force) present by replacing $E^M_{X97b82} = \exp(e^M_{X97b82} + w_{M97})$ in (2) with the adjusted value from (6).

Next, we adjust the density of log equivalent family earnings for changes in female labour force participation. That is we want to find

$$\begin{aligned}
&f\left(y;t_y=97,\ t_{X_M b_M|L,X_F b_F,A,C,X}=82, t_{L|X_F b_F,A,C,X}=82, t_{F|A,C,X}=97, t_{A|C,X}=97,\ t_{C|X}=97, t_X=97\right)\\
&=\iiint f\left(y\,|\,X_M b_M, L, X_F b_F, A, C, X, t_y=97\right) dF\left(X_M b_M\,|\,L, X_F b_F, A, C, X, t_{X_M b_M|L,X_F b_F,A,C,X}=82\right)\\
&y_{L|X_F b_F,A,C,X}(L, X_F b_F, A, C, X)\, dF\left(L\,|\,X_F b_F, A, C, X, t_{L|X_F b_F,A,C,X}=97\right) \qquad (7)\\
&dF(X_F b_F, A, C, X\,|\,t_{X_F b_F,A,C,X}=97)
\end{aligned}$$

where $y_{L|X_F b_F,A,C,X}(L, X_F b_F, A, C, X) = \dfrac{dF\left(L \mid X_F b_F, A, C, X, t_{L|X_F b_F,A,C,X} = 82\right)}{dF\left(L \mid X_F b_F, A, C, X, t_{L|X_F b_F,A,C,X} = 97\right)}$ is a reweighing function that represents the changes that have occurred between 1982 and 1997 in female labour force participation. Labour force status L takes on the values of 0 or 1, hence the reweighing function may be stated as

$$\begin{aligned} y_{L|X_F b_F,A,C,X}(L, X_F b_F, A, C, X) = L\, &\frac{\Pr\left(L = 1 \mid X_F b_F, A, C, X, t_{L|X_F b_F,A,C,X} = 82\right)}{\Pr\left(L = 1 \mid X_F b_F, A, C, X, t_{L|X_F b_F,A,C,X} = 97\right)} \\ &+ (1-L)\,\frac{\Pr\left(L = 0 \mid X_F b_F, A, C, X, t_{L|X_F b_F,A,C,X} = 82\right)}{\Pr\left(L = 0 \mid X_F b_F, A, C, X, t_{L|X_F b_F,A,C,X} = 97\right)} \end{aligned} \tag{8}$$

Note that the reweighing function $y_{L|X_F b_F,A,C,X}(L, X_F b_F, A, C, X)$ is set to one when a female head/wife is not present in the family.

We estimate the above probabilities using a probit model in which the latent variable determining a woman's labour force participation is a function of age, education, and province of residence. The predicted reweighing function is then multiplied by the weights of each observation for which a female head/wife is present.

Our third adjustment modifies this counterfactual density for changes to the wage structure of women. That is we want to estimate

$$f\left(y; t_y = 97,\ t_{b_M|L,X_F b_F,A,C,X} = 82,\ t_{L|X_F b_F,A,C,X} = 82,\ t_{X_F b_F|A,C,X} = 82, t_{A|C,X} = 97,\ t_{C|X} = 97, t_X = 97\right) \tag{9}$$

To estimate female weekly earnings, we use a Heckman two-step procedure to account for selection into the labour force. The econometric model for female log weekly earnings may be stated as

$$\begin{aligned} e_t^F &= \max\left\{0, X_{1Ft} b_{Ft} + e_{Ft}\right\} \\ I_t &= X_{2Ft} g_{Ft} + h_{Ft} \\ e_t^F &> 0 \text{ if } I_t > 0 \end{aligned} \tag{10}$$

Here, I_t is a latent variable determining entry into the labour force. X_{1Ft} includes a quadratic in age and dummy variables indicating province of residence, education level, and full time or part time job status. X_{2Ft} includes a quadratic in age, age-education interaction terms, and dummy variables indicating province of residence and education level. The estimates of $_{F82}$ are applied to the 1997 female characteristics to adjust female log weekly earnings

$$e^F_{X97b82} = X_{F97} b_{F82} + e_{F97}. \tag{11}$$

In turn, these estimates are used to adjust the family's equivalent family earnings in (2).

The fourth step in the decomposition is to adjust the density of equivalent family earnings for the degree of assortative mating within families. That is, we want to estimate

$$f\left(y;t_y=97,\ t_{X_M b_M|L,X_F b_F,A,C,X}=82, t_{L|X_F b_F,A,C,X}=82, t_{X_F b_F|A,C,X}=82, t_{A|C,X}=82,\ t_{C|X}=97, t_X=97\right) \tag{12}$$

To obtain this density we multiply the density represented by (9) by the reweighing function $y_{A|C,X}(A,C,X)=\dfrac{dF\left(A|C,X,t_{A|C,X}=82\right)}{dF\left(A|C,X,t_{A|C,X}=97\right)}$. This reweighing function represents the changes that have occurred between 1982 and 1997 in assortative mating and is similar in concept to the reweighing function for labour force participation. Assortative mating is described by the likelihood of a husband in the earnings decile *i* to be married to a wife in the earnings decile *j* (*i,j=1,2,…10*). The reweighting function may be written as

$$\begin{aligned} y_{A|C,X}(A,C,X) &= S\sum_{m=1}^{10} I_m \sum_{n=1}^{10} I_n \frac{\Pr\left(i=m,\ j=n\,|\,C,X,t_{A|C,X}=82\right)}{\Pr\left(i=m,\ j=n\,|\,C,X,t_{A|C,X}=97\right)} \\ &\quad +(1-S)\sum_{m=1}^{10} I_m \sum_{n=1}^{10} I_n \frac{\Pr\left(i=m,\ j=n\,|\,C,X,t_{A|C,X}=82\right)}{\Pr\left(i=m,\ j=n\,|\,C,X,t_{A|C,X}=97\right)} \end{aligned} \tag{13}$$

where S=1 if there are children present in the family and zero otherwise, I_m is an indicator function that takes on a value of 1 if $i=m$, and zero otherwise and similarly for I_n. The probabilities used to estimate this reweighing function are simple cross-tabulations of husbands' and wives' earnings deciles, and are estimated for married couples with and without children separately to recognize different degrees of assortative mating amongst couples with children and couples without children. Note that the reweighing function is set equal to 1 for single individuals. The counterfactual density is found by multiplying this reweighing function by the weights on each observation.

In the fifth stage of the decomposition we want to estimate the density

$$f\left(y;t_y=97,\ t_{X_M b_M|L,X_F b_F,A,C,X}=82, t_{L|X_F b_F,A,C,X}=82, t_{X_F b_F|A,C,X}=82, t_{A|C,X}=82, t_{C|X}=82, t_X=97\right) \tag{14}$$

which is found by multiplying the density represented by (12) by the reweighing function $y_{C|X}(C,X)=\dfrac{dF\left(C|X,t_{C|X}=82\right)}{dF\left(C|X,t_{C|X}=97\right)}$ which represents the changes that have occurred between 1982 and 1997 in family composition. The reweighing function may be written as

$$y_{C|X}(C,X)=\sum_{s=1}^{4} I_s \frac{\Pr\left(C=s\,|\,X,t_{C|X}=82\right)}{\Pr\left(C=s\,|\,X,t_{C|X}=97\right)} \tag{15}$$

where I_s is an indicator function taking on a value of 1 when C=s and zero otherwise. Recall that 4 categories for family composition are used here (married with kids, married without kids, others with kids, and others without kids). The probabilities used to estimate $y_{C|X}(C,X)$ are found using a multinomial logit model, the equation for which is

$$\Pr(C=s\,|\,X,t_{C|X}=t)=\frac{\exp(X_{st}b_{st})}{1+\sum_{j=1}^{4}\exp(X_{jt}b_{jt})} \tag{16}$$

where X_{st} represents a subset of X for family type s, which includes the age of the family head, age squared, and dummy variables indicating province of residence.

Finally, the densities are reweighed to account for changes to family characteristics. That is, we want to estimate

$$f\left(y;t_y=97,\ t_{X_M b_M|L,X_F b_F,A,C,X}=82,t_{L|X_F b_F,A,C,X}=82,t_{X_F b_F|A,C,X}=82,t_{A|C,X}=82,t_{C|X}=82,t_X=82\right)$$

$$\begin{aligned}
&=\iiint\iiiint f\left(y\,|\,X_M b_M,L,X_F b_F,A,C,X,t_y=97\right)dF\left(X_M b_M\,|\,L,X_F b_F,A,C,X,t_{X_M b_M|L,X_F b_F,A,C,X}=82\right)\\
&y_{L|X_F b_F,A,C,X}(L,X_F b_F,A,C,X)\,dF\left(L\,|\,X_F b_F,A,C,X,t_{L|X_F b_F,A,C,X}=97\right)\\
&dF\left(X_F b_F\,|\,A,C,X,t_{F|A,C,X}=82\right)y_{A|C,X}(A,C,X)\,dF\left(A\,|\,C,X,t_{A|C,X}=97\right)\\
&y_{C|X}(C,X)\,dF\left(C\,|\,X,t_{C|X}=97\right)y_X(X)\,dF\left(X\,|\,t_X=97\right)
\end{aligned} \tag{17}$$

where the reweighing function for family characteristics is

$$y_X(X)=\frac{dF\left(X\,|\,t_X=82\right)}{dF\left(X\,|\,t_X=97\right)}. \tag{18}$$

To estimate (18), we need to distinguish between married couples and families headed by single individuals. We therefore estimate the reweighing function as

$$\begin{aligned}
y_X(X)=I_C&\frac{\Pr(t_X=97\,|\,I_C=1)}{\Pr(t_X=82\,|\,I_C=1)}\frac{\Pr(t_X=82\,|\,X,I_C=1)}{\Pr(t_X=97\,|\,X,I_C=1)}\\
+(1-I_C)&\frac{\Pr(t_X=97\,|\,I_C=0)}{\Pr(t_X=82\,|\,I_C=0)}\frac{\Pr(t_X=82\,|\,X,I_C=0)}{\Pr(t_X=97\,|\,X,I_C=0)}
\end{aligned} \tag{19}$$

where I_C=1 if the family consists of a married couple and zero otherwise. Equation (19) found by applying Bayes rule. The conditional probabilities $\Pr(t_X=t\,|\,X,I_C=1)$ and $\Pr(t_X=t\,|\,X,I_C=0)$ are obtained by pooling the 1982 and 1997 samples and then using a probit model with the year as a binary dependent variable to estimate the probability of each observation falling into the 1997 sample, based on the characteristics *X* and given marital status. The unconditional probabilities $\Pr(t_X = t)$ are the weighted shares of the 1982 and 1997 samples in the pooled sample.

Implementation of the reverse order decomposition simply involves reversing the order in which the weights are applied or head's earnings are estimated. Table 2 summarizes the weights used in each stage of the primary and reverse order decomposition.

b. Results

In figure 10, we plot the estimated counterfactual densities that result from the sequential decomposition. We begin in panel A by presenting the 1982 and 1997 raw densities which show the widening of the distribution of family earnings over this period. There is a substantial shift of the upper tail of the family earnings distribution while a substantial portion of the lower tail remains pin at the 1982 density levels. There is also a sizeable reduction in mass around the mode of the distribution. The subsequent panels in Figure 10 represent counterfactual densities of the 1997 density of log family equivalent earnings holding each factor at 1982 levels. The inequality statistics corresponding to each density are presented in Table 3. Table 4 evaluates the effects of the counterfactual experiments on the changes in the distribution of log family equivalent earnings from 1982 to 1997. The magnitude of the changes in the different inequality measures corresponds to 6-8 percent of the 1982 values. For the United States, Daly and Valetta (2000) report changes in the inequality measures ranging from 6 percent (for the 90-50) to 16 percent (for the 90-10).

As seen in panel b of Figure 10, holding male wage structure at its 1982 level leads to a substantial rightward shift of the 1997 family earnings distribution. While there have been some mild increases in the returns to education for men (see appendix table A1), which should shift the counterfactual density to the left, the decrease in the constant in the 1997 wage structure actually cause the density to shift to the right. This decrease in the constant is consistent with evidence of declining wages for younger and less educated male workers (Picot, 1998, Beaudry and Green, 2000). The inequality statistics in Table 3 indicate that changes to male wage structure may explain around 13 percent of the increase in family earnings inequality that is seen between 1982 and 1997. Interestingly, as shown in Table 4, changes in the male wage structure account for similar increases in the 90-50 and in 50-10 log earnings differentials, reflecting the fact that the widening male wage distribution occurred both in the upper and lower tail as indicated earlier. Using a different methodology, Burtless (1999) found that changes in the distribution of male earnings could account for 28 percent of the increase in the Gini coefficient in the U.S between 1979 and 1996. Here we consider first only changes in the returns to the male individual characteristics, the changes in the characteristics themselves is considered last.

Second, it appears that the increase in female labour force participation may have acted to reduce inequality among families between 1982 and 1997. By adjusting the 1997 density to hold female labour force participation at 1982 levels, the measures of inequality (generally) increases, indicating that inequality in 1997 would have been higher had women not entered the labour force in greater numbers. This factor influences the 90-50 log differential and the Gini coefficient the most which would suggest that changes in women's participation were most important for pulling up the middle of the earnings distribution.

The third factor of the decomposition, female wage structure, explains a substantial portion (more than 20 percent) of the increase in family earnings inequality between 1982 and 1997. Given that there were substantial increases in the returns to experience and to education for women (as indicated in appendix table 1), it is not surprising that the effect of changes in the female wage structure on family inequality are greater than changes in the male wage structure. The effects of changes in the female wage structure on family equivalent earnings inequality are most important on 50-10 log earnings differential, consistent with our earlier observation that inequality for women increased more at the bottom of the wage distribution.[27] Furthermore, all of the inequality measures are affected by this factor indicating that female wage structure is important for explaining movement of families through the entire distribution.

Assortative mating demonstrates to be an important factor for explaining the increase in family earnings inequality. Cross-tabulations of the husbands and wives earnings deciles indicate a substantial increase in assortative mating from 1982 to 1997. For example, among married couples without children (see appendix table 2), in 1997 wives in the three upper deciles are more likely than in 1982 to have husbands whose earnings are in those same deciles than in the first three lower deciles. Thus women in the lower three deciles of female

[27] Burtless (1999) also consider the impact of changes to the distribution of female earnings in the U.S. and finds that 5 percent of the increase in the Gini coefficient could be attributed to changes in female earnings inequality.

earnings were less likely to be married to men in the three upper deciles. Although appearing visually insignificant in panel d of Figure 10, the increased likelihood of couples to have similar incomes has taken many families out of the lower-middle portion of the family earnings distribution and moved them into the upper and (predominantly) lower ends of the distribution. This is supported by the large changes (from 27 to 32 percent) in the 90-10 and 50-10 log differentials relative to the other measures of inequality.

As earlier studies predicted, changes in family composition are found to be a dominant factor accounting for increases in inequality among families. From panel f of Figure 10, it is apparent that many families moved from the middle to the lower end of the family earnings distribution as a result of changes to family composition reflecting, for example, the significant increase in lone parents.[28] More precisely, the changes to family composition account for 47 percent of the changes in the 90-10, 28 percent of the changes in the 90-50, and 56 percent of the changes in the 50-10 log earnings differentials. For the United States, Daly and Valetta (2000) find that changes in family structures are a factor explaining from 22 to 53 percent of the changes in measures in family equivalent income inequality from 1979 to 1989.

Finally, we adjust the 1997 family earnings distribution to account for changes in family characteristics. Since both male and female heads and/or spouses became older and more educated from 1997 to 1982, reverting these characteristics to the 1982 levels will naturally generates a substantial leftward shift of the density. However, because we do not have information on union status, we cannot generate a union lump, that is the mass present near the mode in 1982. Instead we generate a lump around the 25th centile of the log family equivalent earnings distribution, which actually leads to an increase in our measures of earnings inequality. Fortin and Lemieux (2000) show that the decline in union density from 1995 to 1998 generates such a transfer of mass from the mode to the mid-lower part of the wage distribution.

Because we cannot account for changes in union density, our end results seem to have less explanatory power in terms of the inequality measures than in terms of the densities once we visually transpose the counterfactual mass around the 25th percentile to the 50th percentile. In panel h of Figure 10, we superimpose the counterfactual density that accounts for all of the above six factors and the actual 1997 and 1982 densities. Discrepancies between the 1982 density and the counterfactual density reflect unexplained changes. Aside from the effect of unionization, our explanatory factors seem to account for most of the changes in the distribution of log family equivalent earnings, as shown by the 50-10 differential for example.

c. Robustness

In the bottom panel of Table 3 and in Table 5 we repeat the analysis in reverse order. We thus perform our sequential analysis beginning with the effect of family characteristics, followed by the effects of family composition, assortative mating, female wage structure, female labour force participation and finally male wage structure. The decomposition appears relatively robust to the order of the decomposition. However, changing the order of the decomposition changes the impact (however small) of adjusting the 1997 density for female labour force participation, in that the 90-10 and 50-10 differentials demonstrate that increasing female labour force participation has contributed to increasing family earnings inequality rather than offsetting the effects of other factors. The reverse order decomposition also changes the explanatory power of different factors. For example, in the primary order decomposition assortative mating explained 27 percent and 31 percent of the increase in the 90-10 and 50-10 log differential respectively. In the reverse order decomposition, assortative mating explains 37 percent and 47 percent of these measures. In contrast, the explanatory power of family composition is reduced using the reverse order decomposition. Overall, the factors presented appear slightly less able to explain increasing family earnings inequality when using the reverse order decomposition although qualitatively the results are very similar using either order.

4. Conclusion

In this chapter, we begin by describing the important economic progress of women in the last two decades of the 20th century. This progress, in terms of labour force participation, is

[28] Reweighting the 1997 density to reflect 1982 family composition reduced the proportion of single parent families (as a proportion of all families with at least one head in the labour force) from 6.2 percent to 5.2

found to be more important in the 1980s, and in terms of labour market earnings in the early 1990s. Similarly, the relative earnings of women measured by the female/male earnings ratio increased more importantly in the early 1990s. There are some concerns regarding the stagnation of female labour force participation since the early 1990s, of female median earnings and of the female/male earnings ratio since the mid-1990s. The analysis of these troubling trends will be an important topic of future research since it will likely need to include data up to 2005. Here we simply investigate the impact of the 1980s and early 1990s improvements in the labour market outcomes of women on the levels and dispersion of family earnings.

From 1982 to 1997, the increase in median log family equivalent earnings corresponds to a change of 6 percentage log points while the increase in the 90-10 differential was of 12 percentage log points with most of the increase (8 log points) occurring in the 50-10 differential. We find that changes in the wage structure dwarf socio-demographic factors in accounting for the changes in family equivalent earnings dispersion. Over that period, the workforce aged and reached substantially higher levels of educational attainment, which caused a substantial shift in the upper part of the family equivalent earnings distribution. Notwithstanding an unaccounted for union effect, these demographic changes explain most of the increase in the 90-50 differential. Another notable socio-demographic change over that period was the substantial decrease in the proportion of married (including co-habitating) couples among all families, this factor accounts for more than half of the increase in the 50-10 differential. We find that other gender specific factors, such the increase in female labour force participation and changes in the female wage structure, had very significant impacts on family earnings dispersion. The changes in the female labour participation to some extent contributed to offset the increase in the 50-10 differential, consistent with the idea that female sole heads of household are more likely to work than when a male earner is present. Changes in the female wage structure increased wage dispersion more evenly across the family earnings distribution. With increased assortative mating, many married working women found themselves in the upper deciles of the family earnings distribution, but many married women in the lower deciles of the female earnings distribution found that their husband's earnings had slipped into lower deciles of the male earnings distribution. Also, many more women in 1997 than in 1982 were finding themselves sole heads of households.

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percent.

We thus find that the implications of women's economic progress for the welfare of families have been uneven over the past two decades. Women and their families in the upper deciles of the family earnings distribution have been enjoying rising incomes, while families in the middle of the family earnings distribution have seen their family earnings eroded by the stagnation of men's earnings in the 1990s. There has been relatively little change for those families at the bottom of the family earnings distribution.

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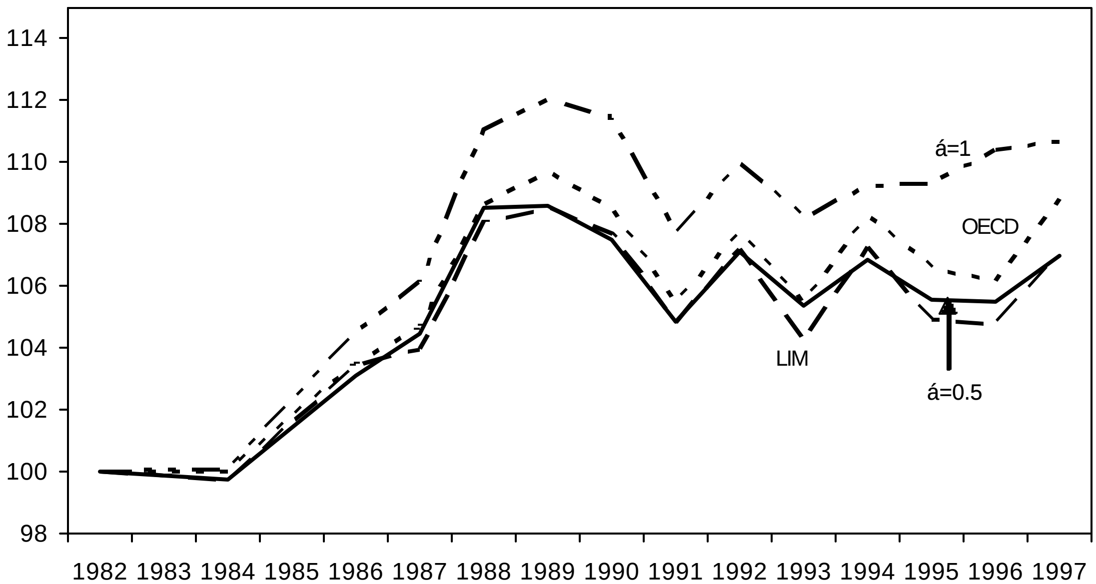
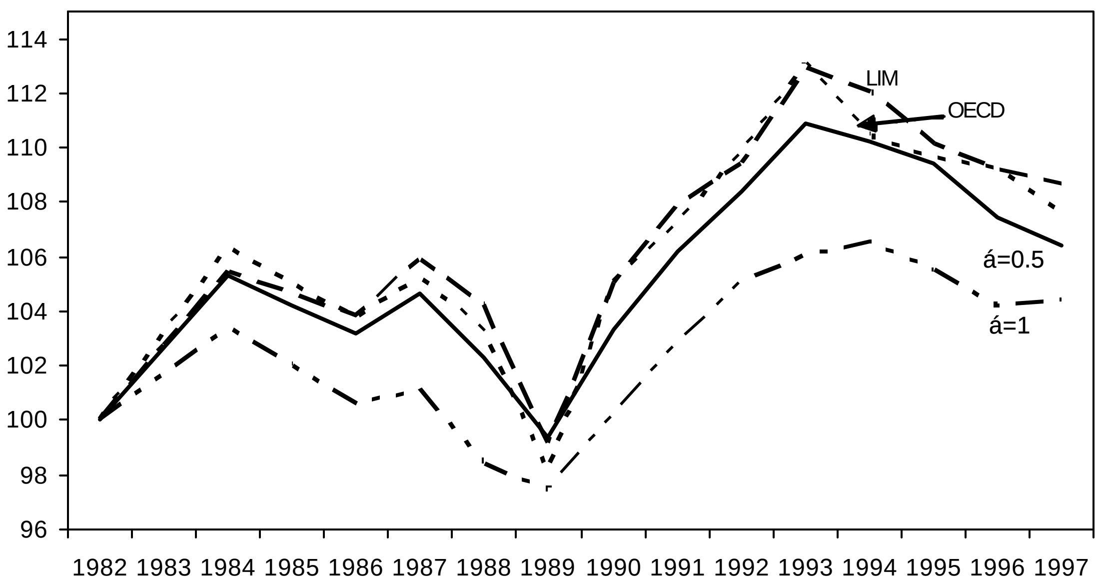
Figure 1
Log Equivalent Family Earnings Using Various Equivalence Scales

A. Median Log Equivalent Family Earnings

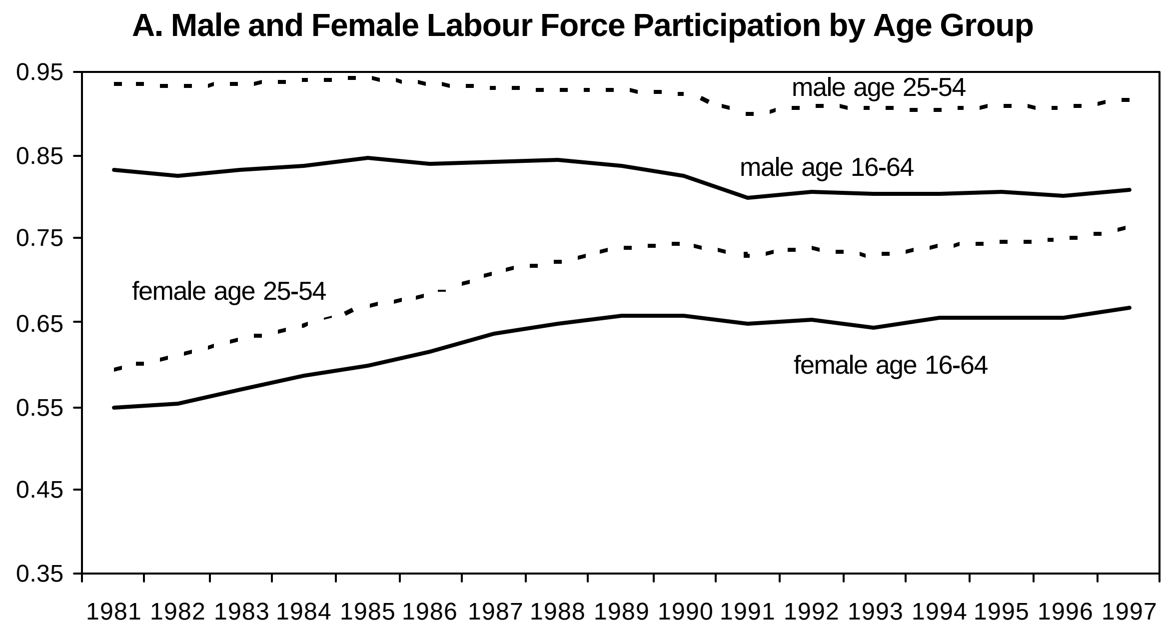


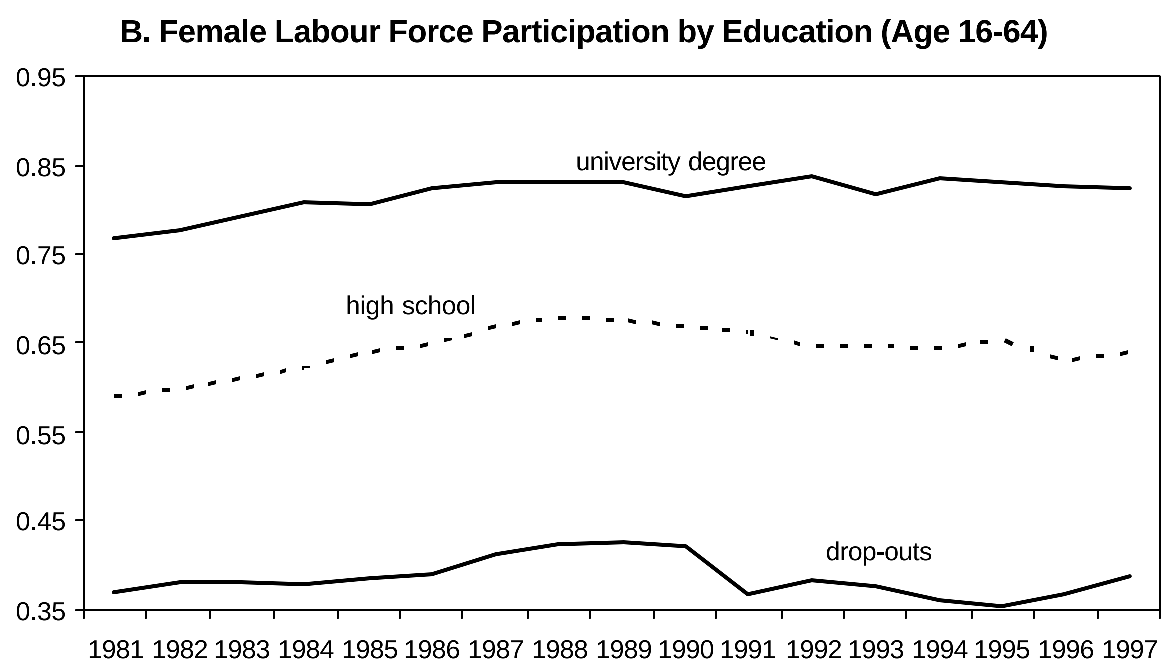
B. 90-10 Differentials



Note: Calculated using the SCF, Census Families, 1982, 1984, and 1986-1997. The data points presented for 1983 and 1985 are averages. Sample restricted to families whose head and spouse are aged 16-64, who are not self-employed, with at least one head in the labour force. Equivalent family earnings are measured as (i) E=Y/F where =1, (ii) E=Y/ F where =0.5, (iii) E=Y/(1+0.4(number of adults)+0.3(number of children)) for LIM, and (iv) E=Y/(1+0.7(number of adults)+0.5(number of children)) for OECD.

Figure 2

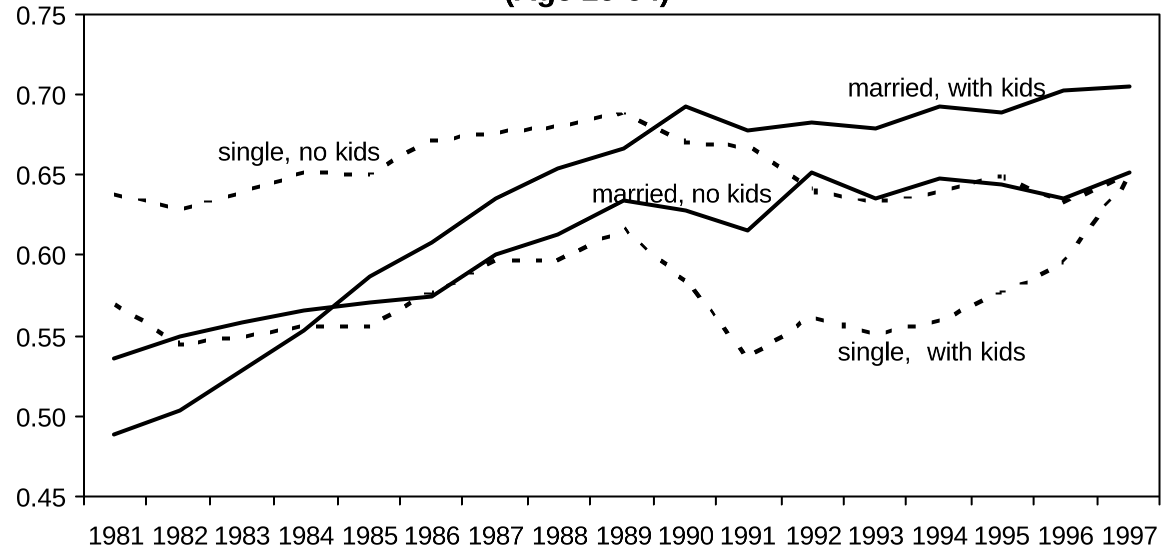




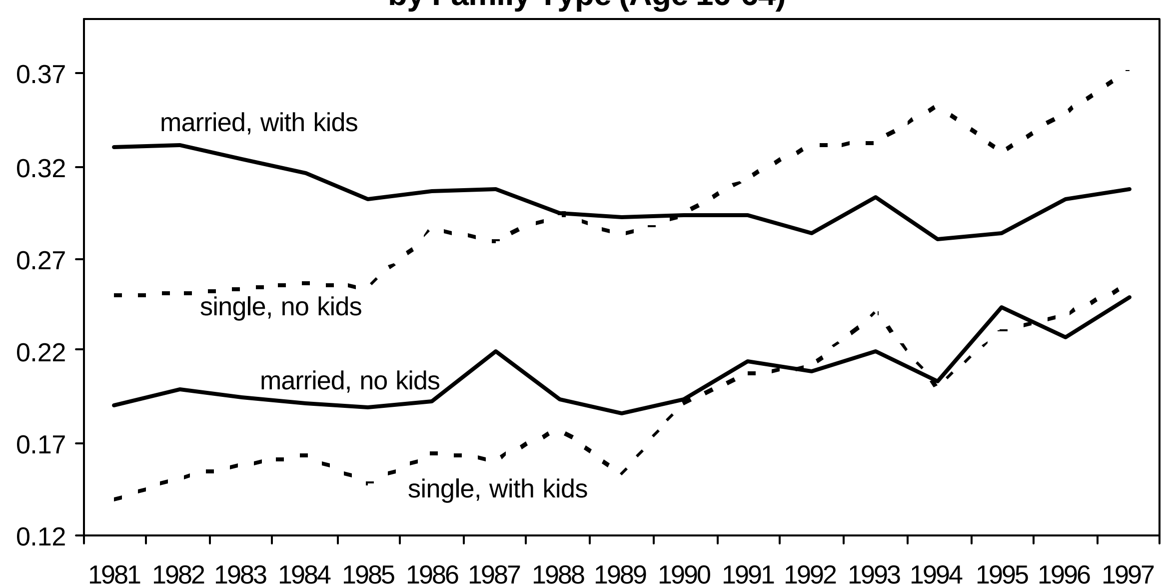
Note: Calculated using the SCF, Individuals Aged 15 and Over, 1981-1982 and 1984-1997. The data point presented for 1983 is an average of 1982 and 1984. Sample excludes the self-employed and observations with non-sensible labour force status. Drop-outs are those individuals reporting to obtain less than grade 10. High school includes individuals with grade 11-13 completed. Individuals classified as having a university degree does not include those with other post-secondary education.

Figure 3

A. Female Labour Force Participation by Family Type (Age 16-64)

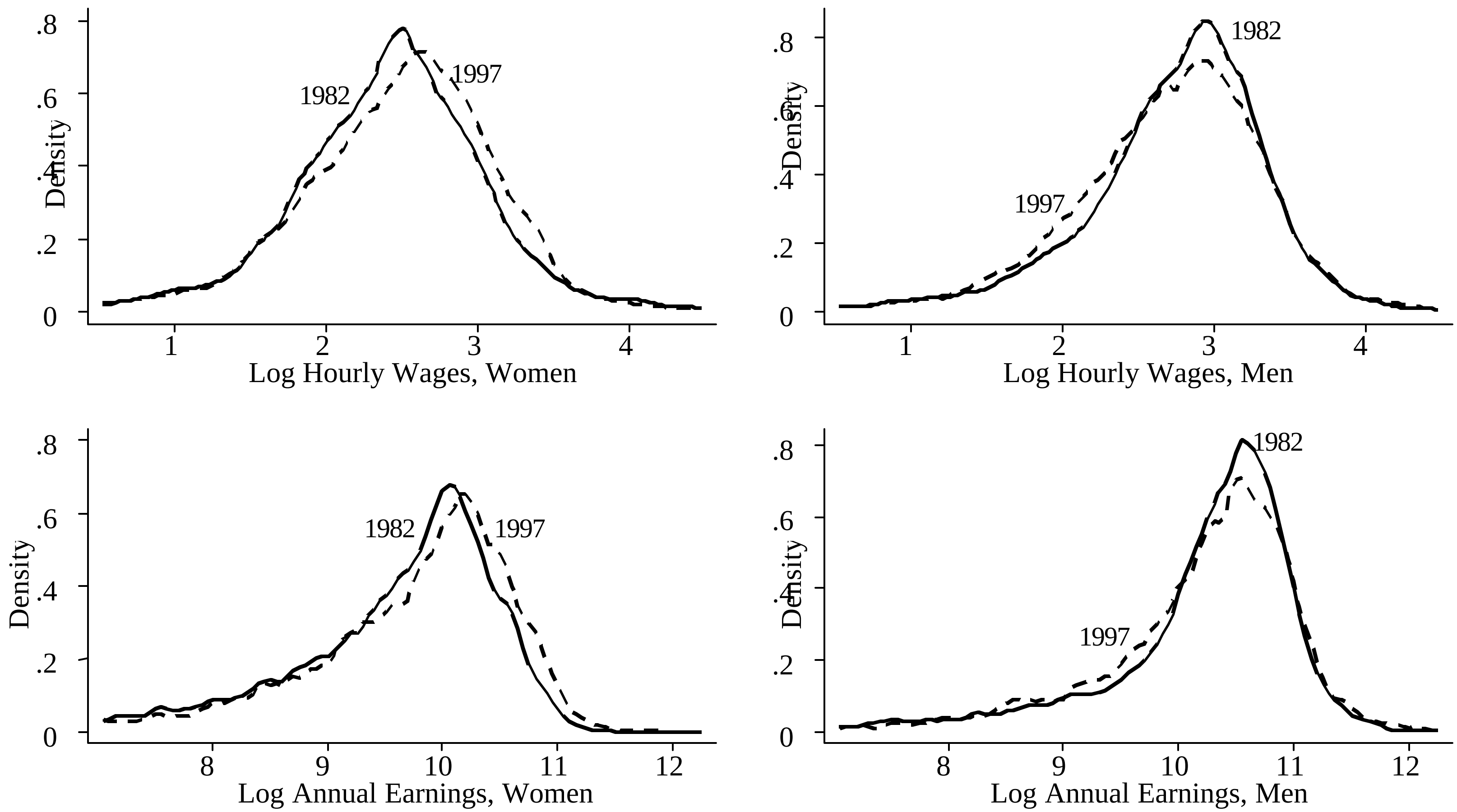


B. Proportion of Participating Women Working Part Time by Family Type (Age 16-64)



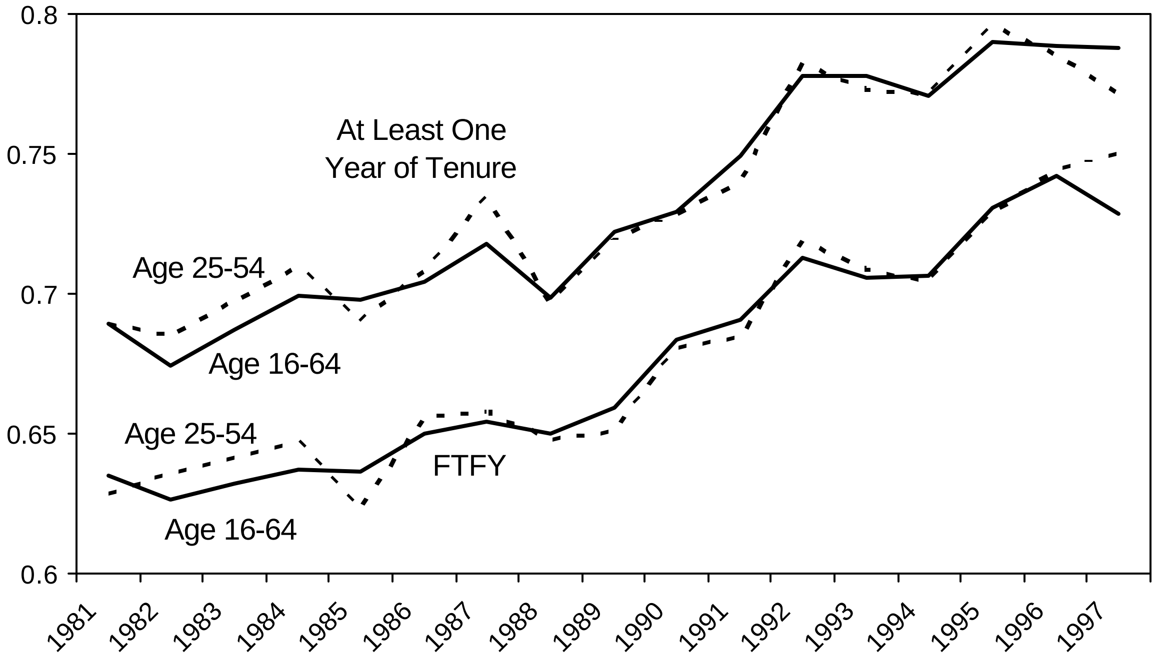
Note: Calculated using the SCF, Individuals Aged 15 and Over, 1981-1982 and 1984-1997. The data point presented for 1983 is an average of 1982 and 1984. Sample excludes the self-employed and observations with non-sensible labour force status. Married individuals include those living common law. Single individuals include those who have never married and others who may be divorced, separated or widowed. An individual with kids reports that there are children under the age of 18 in the census family. The sample of participating women excludes the self-employed and those not in the labour force. A women is working part time if she reports working mostly part time in the year.

Figure 4
Densities of Log Hourly Wages and Log Earnings for Women and Men



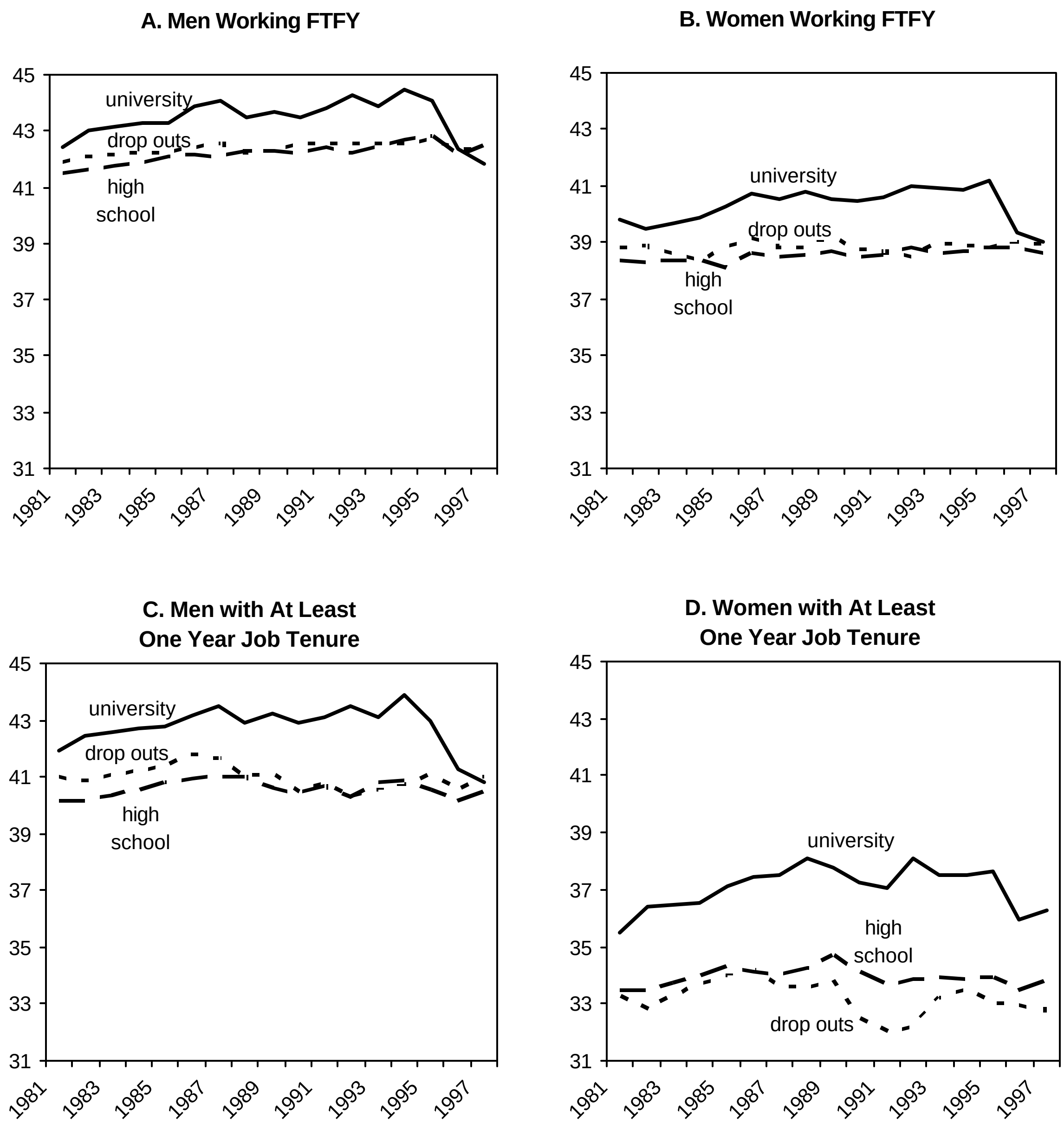
Note: Densities obtained using the SCF, Individuals Aged 15 and Over. Sample restricted to individuals aged 16-64, who are not self-employed, and are in the labour force.

Figure 5
Female/Male Median Earnings Ratio



Note: Calculated using the SCF, Individuals Aged 15 and Over, 1981-1982 and 1984-1997. The data point presented for 1983 is an average of 1982 and 1984. Sample excludes the self-employed and those not in the labour force. The FTFY sample is further restricted to those who report working full time full year, whose annual earnings are between $3000 and $200,000. For this sample, the earnings ratio is measured as median female annual earnings as a ratio of the median male annual earnings. The One year of tenure sample is further restricted to those who report being with the current employer for more than one year, whose hourly wages are between $1.50 and $100. The earnings ratio is measured as median female hourly wages as a ratio of median male hourly wages.

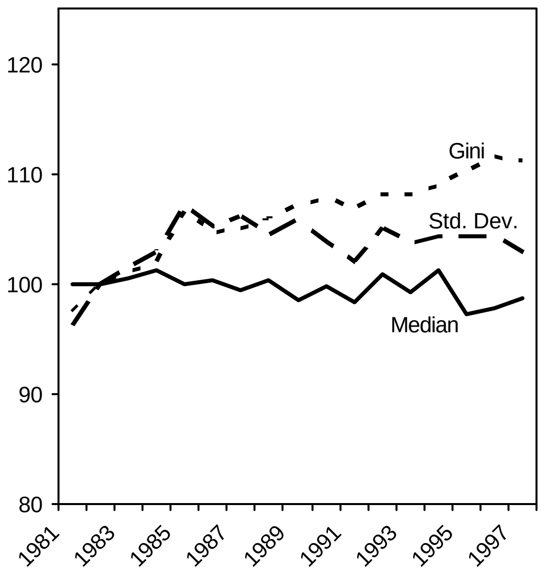
Figure 6
Average Weekly Hours by Education Level



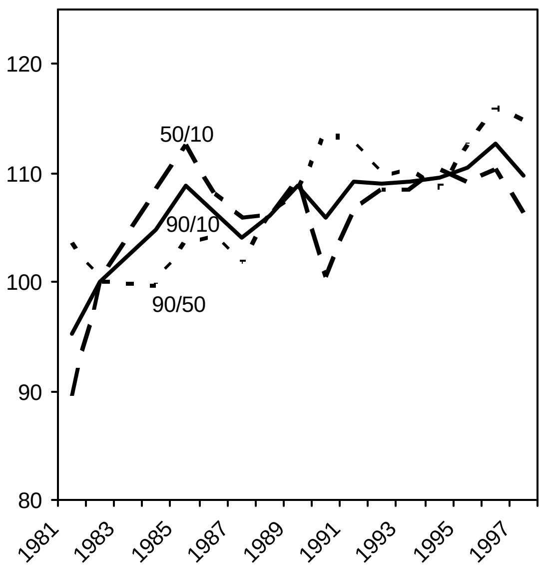
Note: Calculated using the SCF, Individuals Aged 15 and Over, 1981-1982 and 1984-1997. The data point presented for 1983 is an average of 1982 and 1984. Sample restricted to individuals aged 16-64 who are not self-employed and are in the labour force. High school drop-outs are those individuals reporting to have less than grade 10. High school refers to those individuals with grade 11-13 completed. Individuals classified as having a university degree do not include those with other post-secondary education.

Figure 7
Summary Measures for Male Log Wages and Salaries (Age 16-64), 1982=100

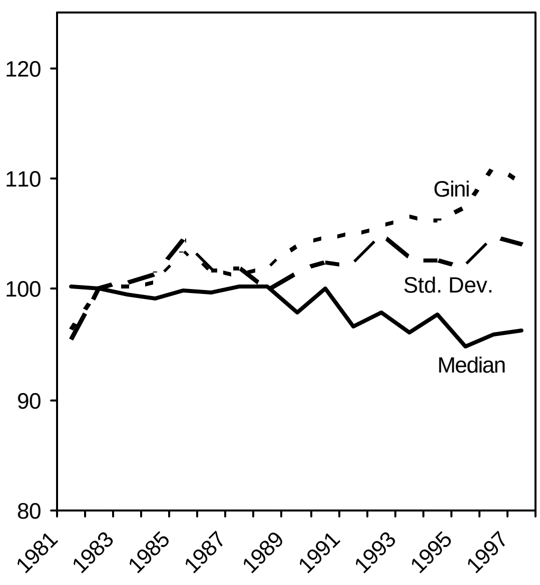
A. FTFY Workers, Annual Earnings



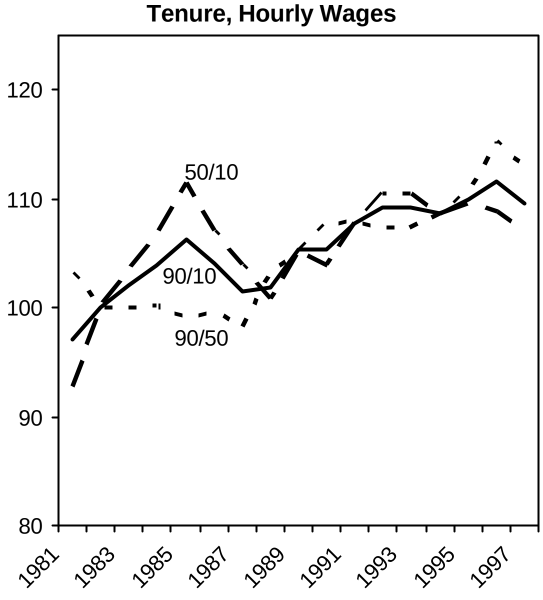
B. FTFY Workers, Annual Earnings



C. Workers with At Least One Year Job Tenure, Hourly Wages



D. Workers with At Least One Year Job Tenure, Hourly Wages

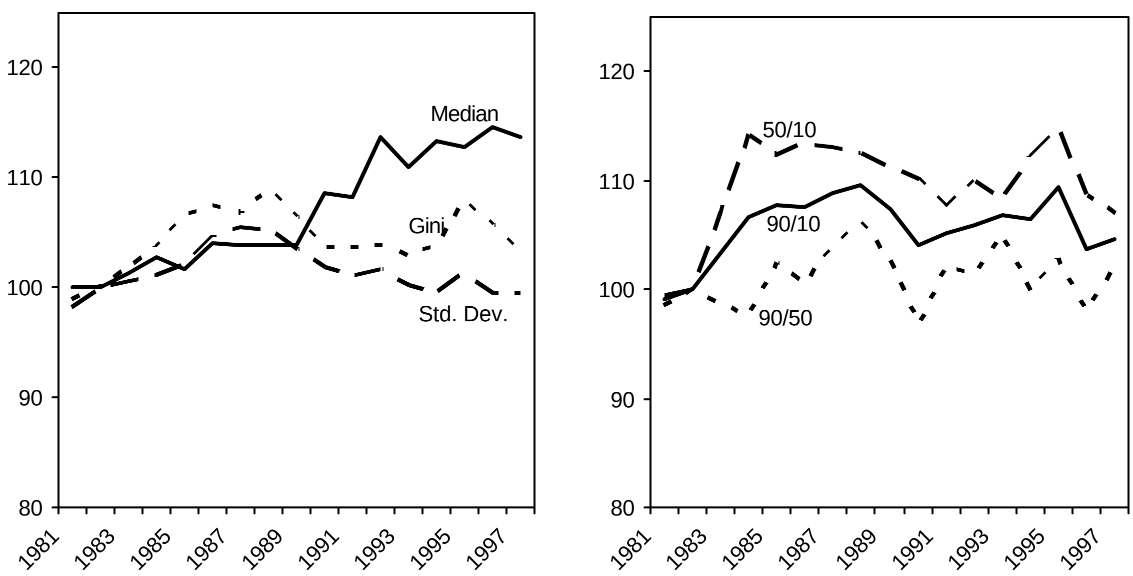


Note: Calculated using the SCF, Individuals Aged 15 and Over, 1981-1982 and 1984-1997. The data point presented for 1983 is an average of 1982 and 1984. Sample restricted to men aged 16-64 who are not self-employed, are in the labour force, with annual earnings between $3000 and $200,000 (1992 constant dollars) or hourly wages between $1.50 and $100.

Figure 8
Summary Measures for Female Log Wages and Salaries (Age 16-64), 1982=100

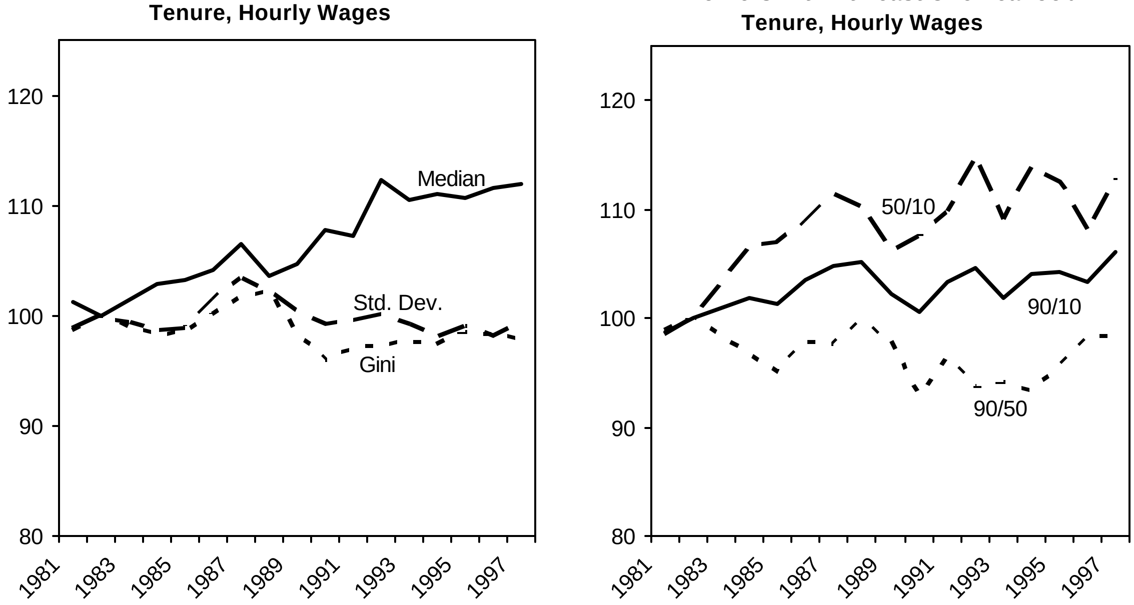
A. FTFY Workers, Annual Earnings

B. FTFY Workers, Annual Earnings



C. Workers with At Least One Year Job Tenure, Hourly Wages

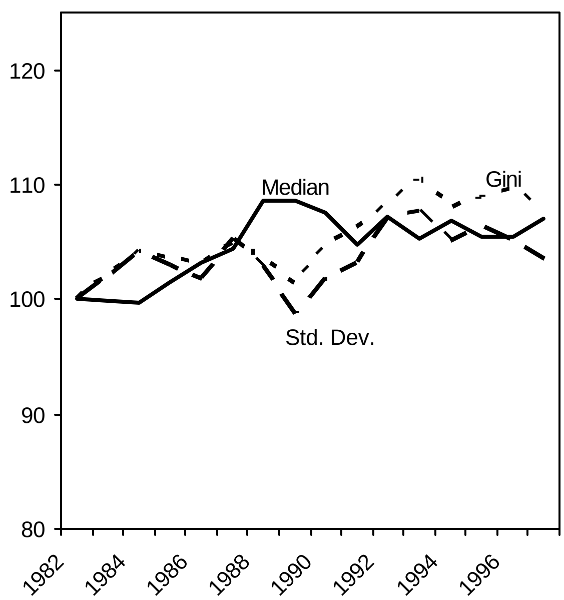
D. Workers with At Least One Year Job Tenure, Hourly Wages

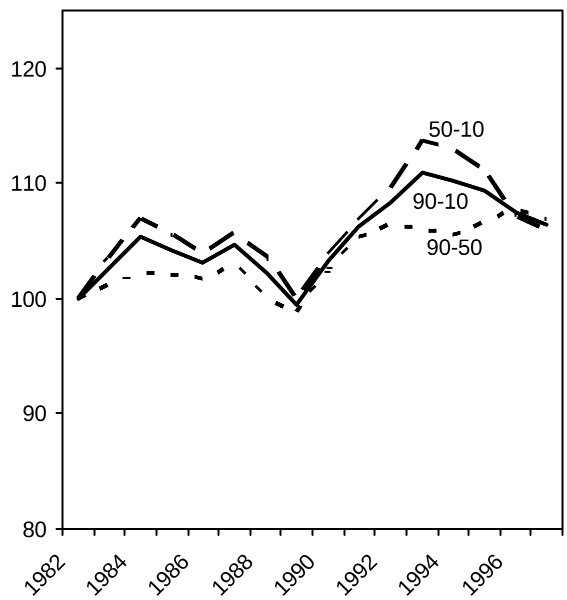


Note: Calculated using the SCF, Individuals Aged 15 and Over, 1981-1982 and 1984-1997. The data point presented for 1983 is an average of 1982 and 1984. Sample restricted to women aged 16-64 who are not self-employed, are in the labour force, with annual earnings between \$3000 and \$200,000 (1992 constant dollars) or hourly wages between \$1.50 and \$100.

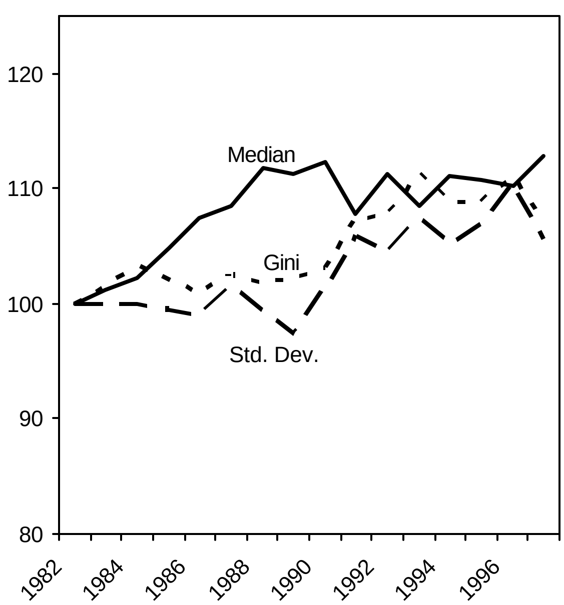
Figure 9
Summary Measures for Log Equivalent Family Earnings, 1982=100
(Head and Spouse Age 16-64)

A. All Families





B. Married With Kids



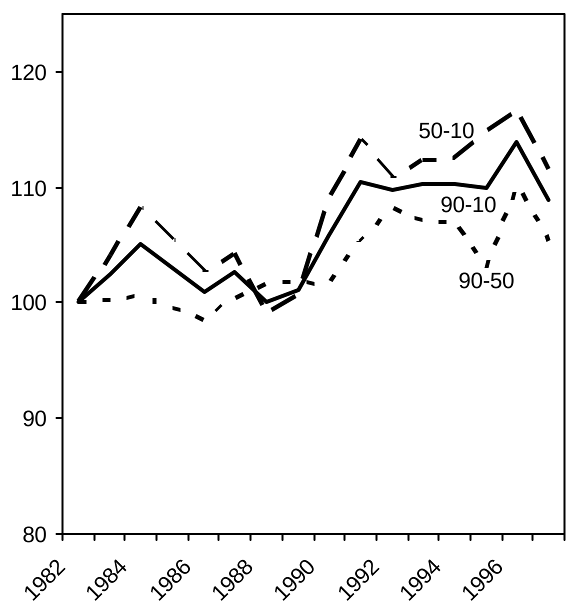
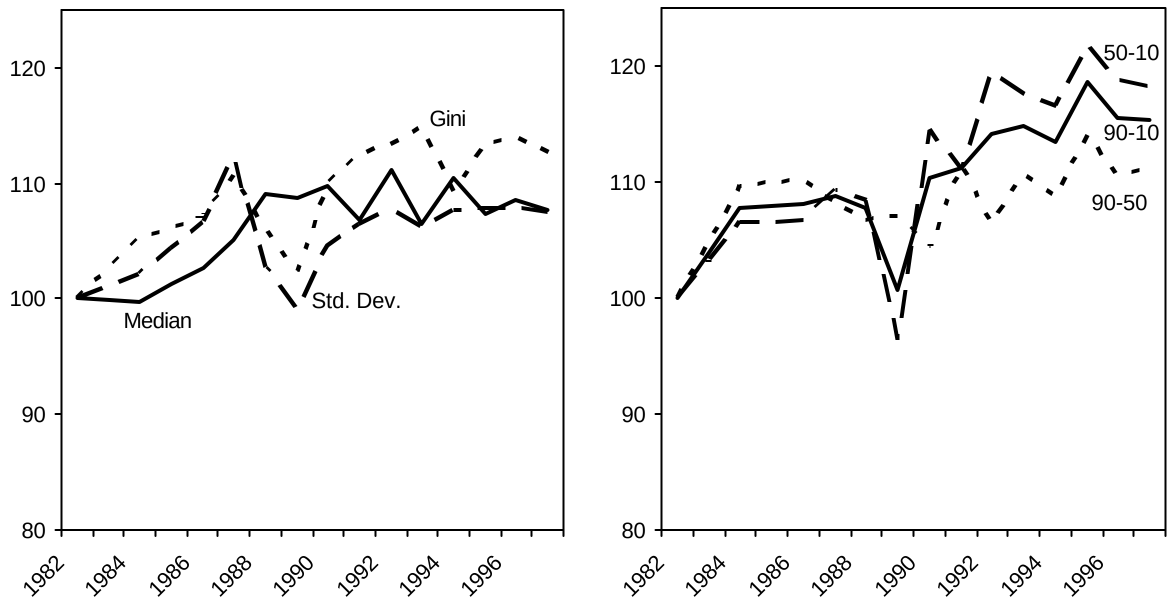


Figure 9 (continued)
Summary Measures for Log Equivalent Family Earnings, 1982=100
(Head and Spouse Age 16-64)

C. Married Without Kids



D. Single with Kids

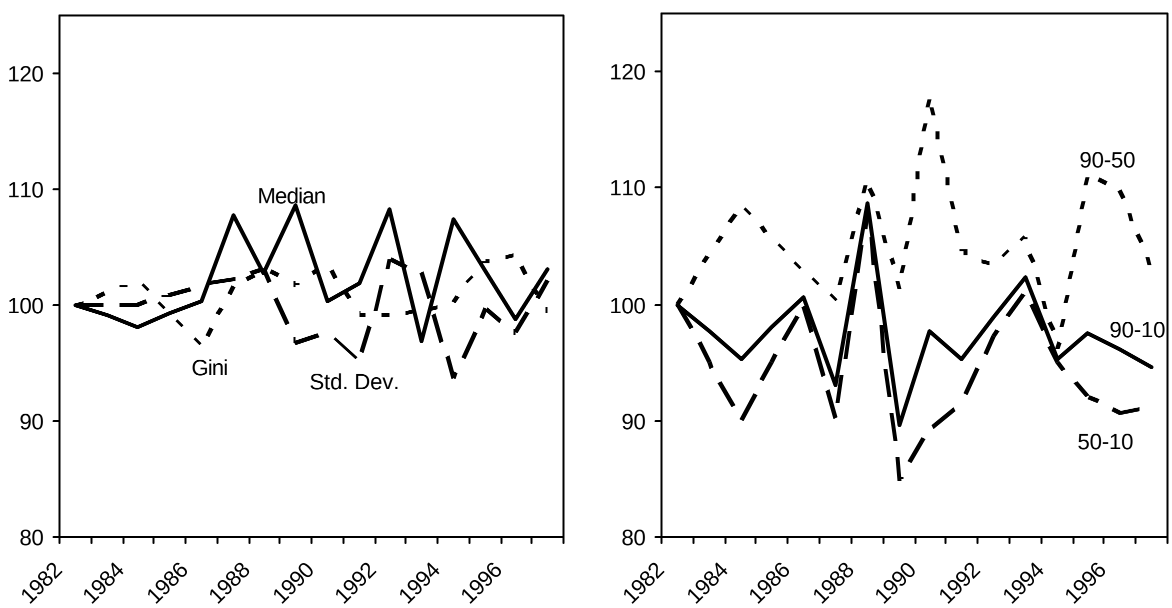
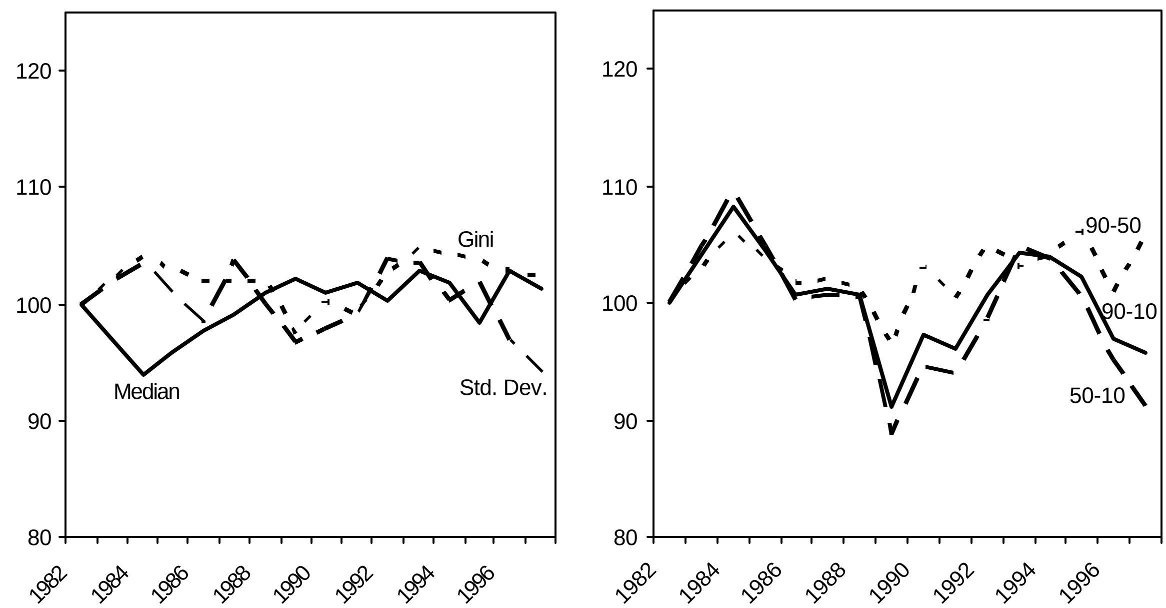


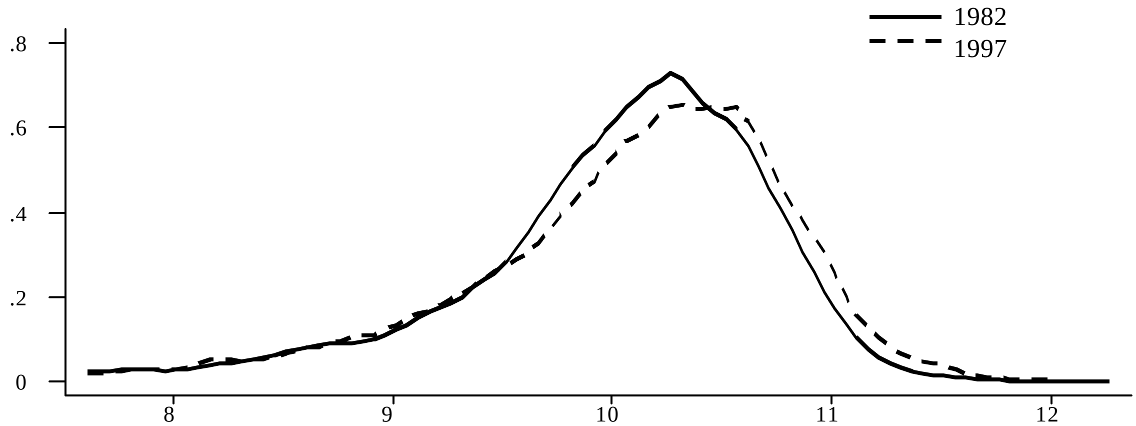
Figure 9 (continued)
Summary Measures for Log Equivalent Family Earnings, 1982=100
(Head and Spouse Age 16-64)

E. Single without Kids

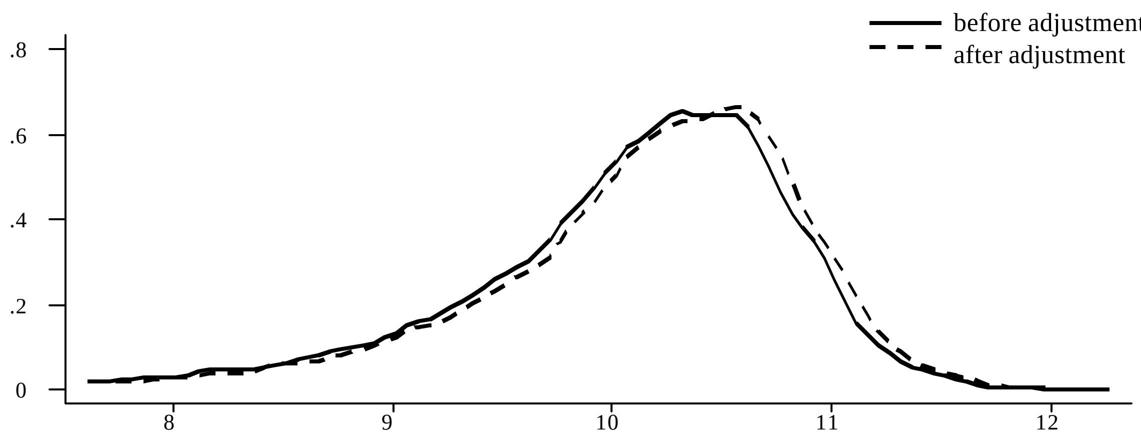


Note: Calculated using the SCF, Census Families, 1982, 1984, and 1986-1997. The data points presented for 1983 and 1985 are averages. Sample restricted to families whose head and spouse are aged 16-64, who are not self-employed, with at least one head in the labour force. Married individuals include those living common law. Single individuals include those who have never married and others who may be divorced, separated or widowed. An family with kids reports that there are children under the age of 18 in the census family.

Figure 10. 1997 Densities of Log Equivalent Family Earnings ($1992) Adjusted for the Indicated Factors



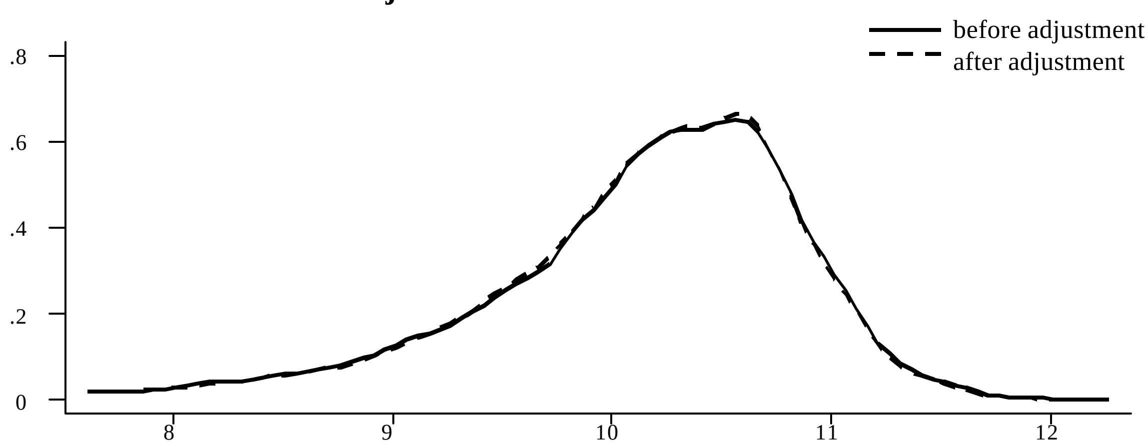
a. Actual Densities



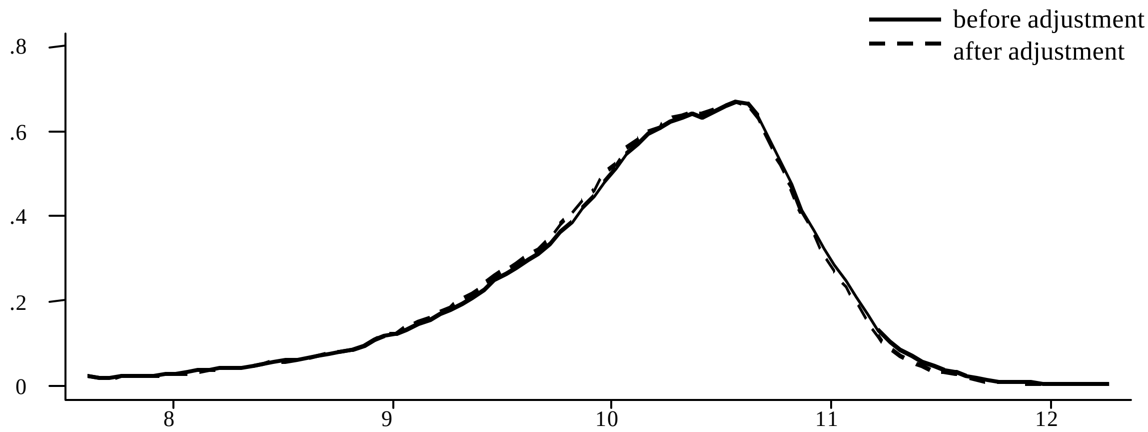
b. Male Wage Structure

c. Female Labour Force Participation

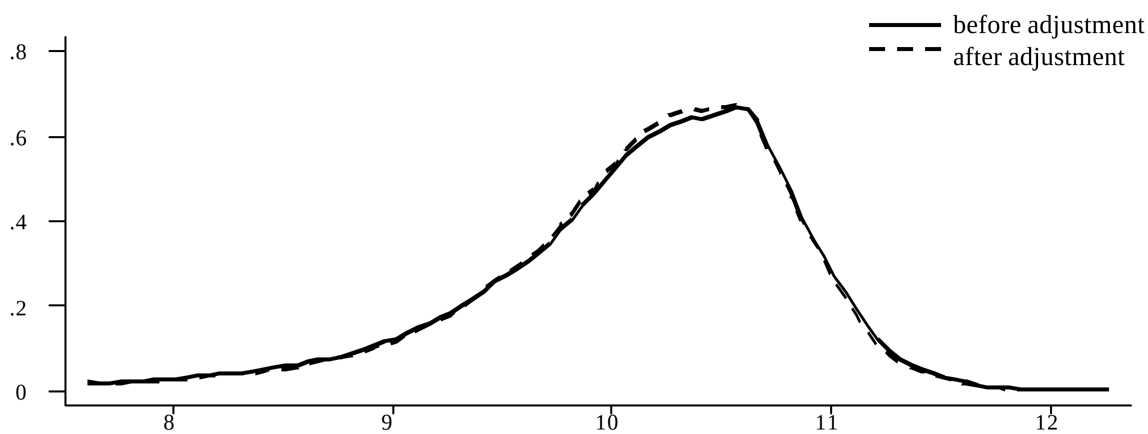
Figure 10 (continued). 1997 Densities of Log Equivalent Family Earnings ($1992) Adjusted for the Indicated Factors



d. Female Wage Structure

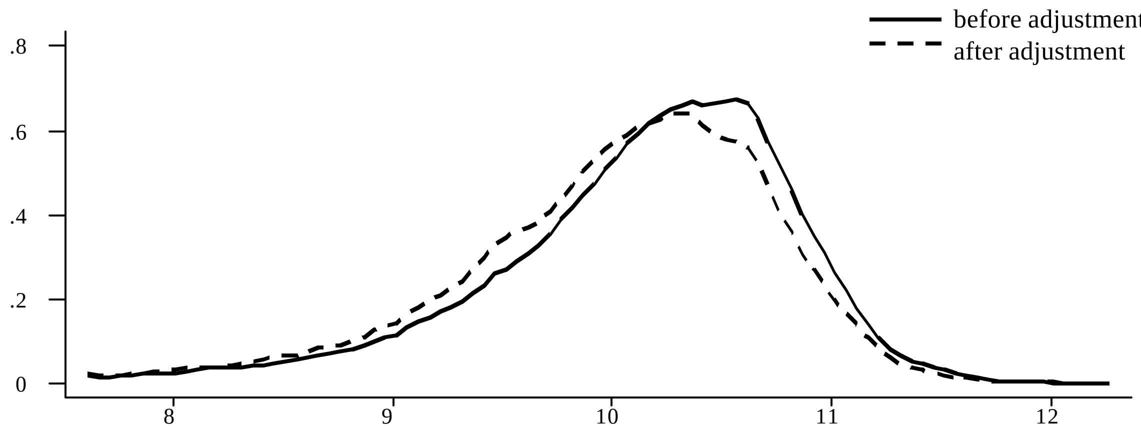


e. Assortative Mating

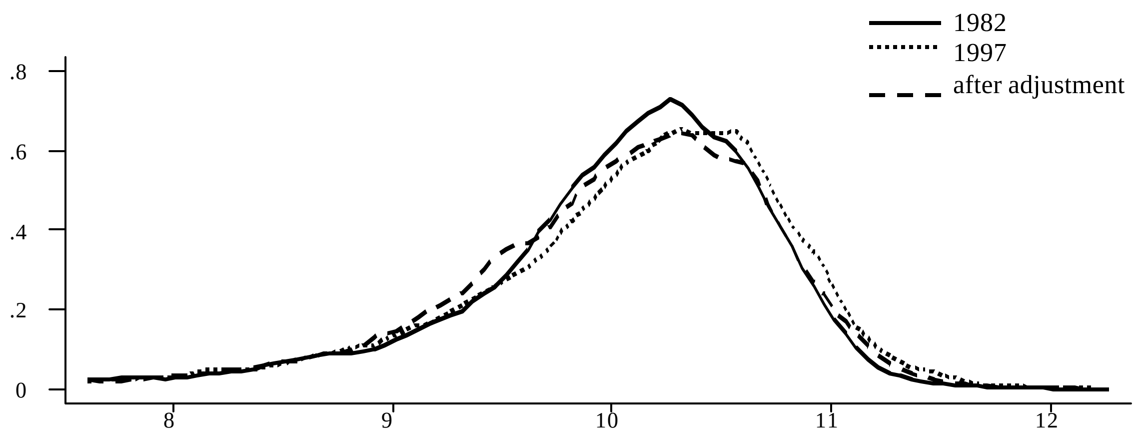


f. Family Composition

Figure 10 (continued). 1997 Densities of Log Equivalent Family Earnings ($1992) Adjusted for the Indicated Factors



g. Family Characteristics



h. All Factors

Table 1

Proportion of Families in Indicated Categories

Year	Married, with kids	Married, without kids	Never Married, with kids	Others, with kids	Others, without kids
1982	0.357	0.249	0.012	0.056	0.326
1984	0.338	0.251	0.015	0.055	0.341
1986	0.336	0.252	0.015	0.050	0.347
1987	0.327	0.252	0.016	0.051	0.355
1988	0.329	0.258	0.019	0.049	0.345
1989	0.324	0.258	0.020	0.050	0.348
1990	0.317	0.256	0.019	0.055	0.353
1991	0.311	0.256	0.022	0.051	0.360
1992	0.308	0.256	0.026	0.058	0.351
1993	0.303	0.252	0.025	0.059	0.361
1994	0.302	0.260	0.026	0.052	0.360
1995	0.302	0.257	0.024	0.052	0.364
1996	0.302	0.262	0.032	0.051	0.353
1997	0.293	0.257	0.031	0.051	0.368

Note: Calculated using the SCF, Census Families, 1982, 1984, and 1986-1997. Sample restricted to families whose head and spouse are aged 16-64, who are not self-employed with sensible labour force status. . Married with kids includes families with a head and spouse, married or living common law, who report the presence of children under the age of 18 in the census family. Married without kids includes families with a head and spouse, married or living common law, who do not report the presence of children. Never married with kids include single individuals never married who report the presence of children under the age of 18 in the census family. Others with kids include those who are separated, divorced, or widowed, who report the presence of children under the age of 18. Others without kids include all other single individuals, never married, separated, divorced or widowed, who do not report the presence of children under the age of 18.

Table 2

Weights Used in the Decomposition

	Primary Order Decomposition		Reverse Order Decomposition	
(1)	Male Wage Structure	y	Female Characteristics	$y\,y_{X_F}$
(2)	Female Labour Force Participation	$y\,y_{L}$	Male Characteristics	$y\,y_{X_F} y_{X_M}$
(3)	Female Wage Structure	$y\,y_{L}$	Family Composition	$y\,y_{X_F} y_{X_M} y_{C}$
(4)	Assortative Mating	$y\,y_{L} y_{A}$	Assortative Mating	$y\,y_{X_F} y_{X_M} y_{C} y_{A}$
(5)	Family Composition	$y\,y_{L} y_{A} y_{C}$	Female Wage Structure	$y\,y_{X_F} y_{X_M} y_{C} y_{A}$
(6)	Male Characteristics	$y\,y_{L} y_{A} y_{C} y_{X}$	Female Labour Force Participation	$y\,y_{X_F} y_{X_M} y_{C} y_{A} y_{L}$
(7)	Female Characteristics	$y\,y_{L} y_{A} y_{C} y_{X_M} y_{X_F}$	Male Wage Structure	$y\,y_{X_F} y_{X_M} y_{C} y_{A} y_{L}$

The functions $y_L, y_A, y_C, y_{X_M}, y_{X_F}$ are the conditional reweighting functions defined by equations (7), (12), (14), and (18). represents the original raw weights contained in the SCF microdata files. Notation for conditional weights has been simplified here for the sake of exposition.

Table 3 Inequality Statistics

A. Initial Estimates	90-10	90-50	50-10	Standard Deviation	Gini
1982	1.7509	0.6445	1.1065	0.8056	0.3222
1997	1.8733	0.6864	1.1869	0.8726	0.3433

B. Primary Order Decomposition	90-10	90-50	50-10	Standard Deviation	Gini
Male Wage Structure	1.8577	0.6809	1.1768	0.8654	0.3406
Female Labour Force Participation	1.8659	0.6879	1.1780	0.8653	0.3432
Female Wage Structure	1.8337	0.6789	1.1547	0.8500	0.3401
Assortative Mating	1.8007	0.6716	1.1291	0.8381	0.3360
Family Composition	1.7438	0.6599	1.0839	0.8160	0.3294
Family Characteristics	1.8136	0.7020	1.1116	0.8560	0.3429

C. Reverse Order Decomposition	90-10	90-50	50-10	Standard Deviation	Gini
Family Characteristics	1.9590	0.7258	1.2331	0.9313	0.3575
Family Composition	1.9047	0.7171	1.1876	0.9088	0.3520
Assortative Mating	1.8559	0.7092	1.1468	0.8949	0.3465
Female Wage Structure	1.8226	0.7056	1.1170	0.8326	0.3432
Female Labour Force Participation	1.8216	0.7103	1.1113	0.8291	0.3448
Male Wage Structure	1.8050	0.7062	1.0988	0.8270	0.3431

Note: Calculated using the SCF, Census Families, 1982, 1997. See methodology section for details. Sample restricted to families whose head and spouse are aged 16-64, who are not self-employed, with at least one head in the labour force.

Table 4 Primary Order Decomposition Results

	90-10	90-50	50-10	Standard Deviation	Gini
Total Change	0.1224	0.0419	0.0805	0.0671	0.0212
Effect of					
Male Wage Structure	0.0156 (12.77)	0.0055 (13.05)	0.0102 (12.62)	0.0073 (10.85)	0.0027 (12.81)
Female Labour Force Participation	-0.0082 (-6.73)	-0.0070 (-16.57)	-0.0013 (-1.60)	0.00002 (0.04)	-0.0025 (-11.98)
Female Wage Structure	0.0323 (26.36)	0.0090 (21.34)	0.0233 (28.98)	0.0154 (22.93)	0.0031 (14.66)
Assortative Mating	0.0330 (26.93)	0.0074 (17.58)	0.0256 (31.81)	0.0119 (17.69)	0.0041 (19.37)
Family Composition	0.0569 (46.51)	0.0117 (27.90)	0.0452 (56.21)	0.0221 (32.94)	0.0066 (31.14)
Family Characteristics	-0.0698 (-57.05)	-0.0421 (-100.46)	-0.0277 (-34.42)	-0.0400 (-59.67)	-0.0135 (-63.86)
Unexplained	0.0627 (51.21)	0.0575 (137.16)	0.0052 (6.41)	0.0504 (75.22)	0.0207 (97.87)

Note: Percent of total variation explained in parentheses. "Unexplained" is the residual not accounted for by all other factors. The effect of an explanatory factor indicates how much of the divergence between the 1997 and 1982 density is explained by replacing the 1997 density by the corresponding counterfactual density.

Table 5 Reverse Order Decomposition Results

	90-10	90-50	50-10	Standard Deviation	Gini
Total Change	0.1224	0.0419	0.0805	0.0671	0.0212
Effect of					
Family Characteristics	-0.0856 (-69.97)	-0.0394 (-93.99)	-0.0462 (-57.45)	-0.0586 (-87.40)	-0.0142 (-66.91)
Family Composition	0.0543 (44.36)	0.0088 (20.87)	0.0455 (56.60)	0.0225 (33.49)	0.0055 (25.86)
Assortative Mating	0.0487 (39.81)	0.0079 (18.83)	0.0408 (50.75)	0.0139 (20.68)	0.0056 (26.29)
Female Wage Structure	0.0334 (27.26)	0.0036 (8.48)	0.0298 (37.04)	0.0623 (92.97)	0.0032 (15.35)
Female Labour Force Participation	0.0010 (0.81)	-0.0047 (-11.14)	0.0057 (7.04)	0.0035 (5.21)	-0.0016 (-7.48)
Male Wage Structure	0.0166 (13.58)	0.0041 (9.75)	0.0125 (15.58)	0.0021 (3.17)	0.0017 (8.10)
Unexplained	0.0541 (44.16)	0.0617 (147.21)	-0.0077 (-9.55)	0.0214 (31.88)	0.0209 (98.78)

Note: Percent of total variation explained in parentheses. "Unexplained" is the residual not accounted for by all other factors. The effect of an explanatory factor indicates how much of the divergence between the 1997 and 1982 density is explained by replacing the 1997 density by the corresponding counterfactual density.

Appendix Table 1
Results of Log Weekly Earnings Regressions.

	1982				1997			
	Men		Women		Men		Women	
	Mean	Coefficient	Mean	Coefficient	Mean	Coefficient	Mean	Coefficient
Variable								
Age	38.57	.0701	35.83	.0414	39.98	.0692	38.31	.0754
		(.0025)		(.0054)		(.0034)		(.0063)
Age Squared /100	16.21	-.0736	14.12	-.0445	17.04	-.0705	15.77	-.0863
		(.0031)		(.00827)		(.0041)		(.0097)
Education Dummy Variables:								
High School	.32	.1661	.39	.1774	.25	.0769	.27	.2915
		(.0105)		(.0374)		(.0164)		(.0429)
Post-secondary	.22	.2221	.25	.3551	.42	.1757	.45	.4549
		(.0117)		(.0536)		(.0153)		(.0619)
University degree	.15	.4245	.13	.6351	.20	.4139	.19	.8901
		(.0129)		(.0805)		(.0171)		(.0914)
Constant	-	4.0702	-	4.1185	-	3.7911	-	3.5385
		(.0611)		(.0834)		(.0771)		(.0834)

Note: Standard errors are in parentheses. Dummy variables for province of residence and full-time/part-time status were also included in the regressions. See the methodology section for more details.

Appendix Table 2
Cross-Tabulation of Male and Female Earnings Deciles For Married Couples Without Children

1982

Male Earning Decile	Female Earning Decile 1	2	3	4	5	6	7	8	9	10	Total
1	0.96	1.25	0.8	0.82	0.96	0.54	0.5	0.34	0.52	0.29	6.98
2	1.04	1.37	1.18	1.04	1.37	0.66	0.54	0.67	0.46	0.4	8.73
3	1.59	1.21	0.95	1.45	1.41	1.09	0.74	0.38	0.56	0.53	9.91
4	1.27	1.21	1.21	1.18	1.25	1.15	0.91	0.93	0.57	0.49	10.17
5	1.32	1.17	0.99	1.19	0.98	0.84	1.04	0.86	0.59	0.56	9.55
6	1.26	1.36	1.16	1.21	1.06	0.95	0.76	0.95	0.96	0.61	10.29
7	1.26	0.98	1.66	1.36	0.85	1.03	1.63	0.83	0.83	0.83	11.25
8	1.24	1.07	1.52	0.92	1.1	0.97	0.77	1.16	0.84	0.92	10.51
9	1.38	1.44	1.22	1.25	0.88	0.99	0.99	0.98	0.98	1.35	11.46
10	1.61	1.1	1.23	1.13	0.93	0.67	0.97	0.83	1.22	1.46	11.15
Total	12.92	12.18	11.9	11.55	10.8	8.89	8.85	7.92	7.54	7.44	100

1997

Male Earning Decile	Female Earning Decile 1	2	3	4	5	6	7	8	9	10	Total
1	1.63	0.88	0.88	0.5	0.51	0.36	0.4	0.27	0.38	0.18	5.99
2	0.99	1.3	1.25	0.98	0.63	0.58	0.71	0.27	0.41	0.31	7.43
3	1.07	0.78	1.19	0.89	1.2	0.93	0.6	0.87	0.39	0.35	8.28
4	1.18	1.33	1.22	1.15	1	1.71	1.32	0.94	0.96	0.59	11.42
5	0.78	0.92	1.14	1.18	1.08	0.84	1.17	0.9	0.71	0.48	9.19
6	0.93	1.01	1.09	0.99	1.55	1.34	0.8	1.11	0.88	0.78	10.49
7	1.51	1.19	1.06	1.05	0.97	0.95	1.09	1.46	1.46	0.76	11.51
8	1.59	1.34	0.83	0.94	0.98	1.12	1.64	1.26	1.15	1.54	12.4
9	1.22	1.07	0.89	1.16	1.13	1.23	0.87	1.19	1.41	1.58	11.77
10	1.02	1.2	0.85	0.92	0.81	1.1	1.14	1.04	1.4	2.04	11.52
Total	11.93	11.03	10.4	9.78	9.89	10.15	9.75	9.31	9.15	8.61	100

Note: The degree of assortative mating may be characterized by the sum of the diagonal. The degree of assortative for couples without children was 11.89% in 1982 and 14.23% in 1997. The degree of assortative mating for couples with children was 11.64% in 1982 and 13.49% in 1997.