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Christopher Kollmeyer

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FAMILY STRUCTURE, FEMALE EMPLOYMENT, AND NATIONAL INCOME INEQUALITY: A CROSS-NATIONAL STUDY OF 16 WESTERN COUNTRIES*

by

Dr Christopher Kollmeyer
Lecturer of Sociology
School of Social Science
University of Aberdeen
Aberdeen AB24 3QY
Scotland, United Kingdom
Email: c.kollmeyer@abdn.ac.uk

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ABSTRACT

This study assesses whether recent changes in family structure and female employment patterns have altered the distribution of income in some countries. Extant literature on this topic reaches inconsistent conclusions and overwhelmingly focuses on the United States. To address these shortcomings, the author draws on internationally comparable data for 16 Western countries to assess whether these social changes have distributional consequences. Specifically, the hypothesis is that increased female employment reduces income inequality, but that increased prevalence of single-mother families heightens income inequality. Results from two-way random effects regression models provide considerable support for this hypothesis. These effects are robust after controlling for variations in labour market institutions, social welfare provisions, and relevant social and economic structures. Limited evidence also suggests that educational homogamy between spouses and partners explains some of the differences in income inequality among countries. The study ends by discussing some of the implications of these findings.

Key Words: Income Inequality; Female Labour Market Participation; Single-Mother Families; Educational Homogamy

FAMILY STRUCTURE, FEMALE EMPLOYMENT, AND NATIONAL INCOME INEQUALITY: A CROSS-NATIONAL STUDY OF 16 WESTERN COUNTRIES

INTRODUCTION

After decades of slow decline, income inequality began to rise in many but not all Western countries in the early 1980s. Numerous studies have addressed this phenomenon, resulting in a body of literature that ranges widely across the social sciences. Within this broad literature, one prominent sociological perspective considers the possibility that changes in family structure may have altered the distribution of income in some countries (Blossfeld and Buchholz, 2009; Esping-Andersen, 2007; Martin, 2006; McLanahan and Percheski, 2008; Treas, 1987; Western, Bloome and Percheski, 2008). In this regard, two structural changes are thought to be most important. The first is the growing portion of families headed by single mothers. This change in family formation practices may increase inequality, largely because such families face entrenched barriers to maintaining income parity with other types of households—for example barriers arising from the gender pay gap or from families with dual incomes. Consequently, as the portion of single-mother families expands, the portion of households with incomes below the national median should expand as well, causing national income inequality to rise. This effect is well documented for the United States (e.g. Treas, 1987; Western, Bloome and Percheski, 2008), but studies from a cross-national perspective are limited and yield inconsistent results (cf. Bradley et al, 2003; Esping-Andersen, 2007).

The second change is the steady rise in female labour market participation. Over recent decades, the portion of women working outside the home has risen significantly in nearly all Western countries. For families and households with two or more adults, this phenomenon creates new opportunities to pool incomes and reduce risks associated with job loss and other traumatic life events. Yet despite the clear trajectory of this trend, its effect on income inequality is not fully understood. Most studies on this subject focus on the United States. Here some studies conclude that increased female employment actually heightens income inequality, largely because it can create dual-income households that combine two high-income earners or two low-income earners, thereby magnifying existing socioeconomic advantages or disadvantages (Burtless, 1999; Karoly and Burtless, 1995; Schwartz, 2010). However, most studies conclude the opposite—namely that increased numbers of women in the paid workforce moderate income inequality, ostensibly because women from low-income households are more likely to work than their counterparts in high-income households (Albrecht and Albrecht, 2007; Cancian and Reed, 1999; Chevan and Stokes, 2000; Nielsen and Alderson, 1997; Treas, 1987; Western, Bloome and Percheski, 2008). Unfortunately, studies on this topic using cross-national data are limited and often reach inconsistent results (cf. Gustafsson and Johansson, 1999; Alderson and Nielsen, 2002; Bradley et al, 2003; see also theoretical argument by Blossfeld and Buchholz, 2009).

In sum, the literature on how these social changes affect income inequality suffers from at least three interrelated deficiencies. (1) It overwhelmingly focuses on the United States, and when cross-national studies have been undertaken, family- and gender-related factors are treated as control variables rather than the primary focus of the study. (2) Comparing results among various studies is difficult due to methodological differences in how income inequality is measured (gross or net), how the data is analyzed (regression or decomposition), and how the underlying unit of analysis is defined (households or families). Finally, (3) the literature has not yielded a consensus on basic questions. For instance, do single-mother families increase income inequality only in the United States, due to its meagre welfare state, or is this phenomenon more generalized? Similarly, does growth in female employment intensify or mitigate income inequality, either in the United States or elsewhere?

In what follows, this study seeks to overcome these shortcomings by assessing how recent changes in family structure and female employment patterns have affected national income inequality. The empirical analysis is based on income inequality estimates from the Luxembourg Income Study (LIS) for a sample of 16 Western countries observed intermittently over recent decades. Results from two-way random-effects regression models generate support for the argument that female labour market participation reduces income inequality, but that the prevalence of single-mother families heightens income inequality. These findings are robust even after controlling for variations in labour market institutions, social welfare provisions, and relevant social and economic structures. It also appears that educational homogamy between spouses and partners explains some of the differences in income inequality among countries.

[Insert figure 1 about here.]

Family Structure and National Income Inequality

Given that families can pool incomes across multiple income earners and redistribute income from working to non-working members of society, numerous studies consider the possibility that changing family structures are affecting the distribution of income in some countries. As noted above, the vast majority of these studies focus on the United States (Albrecht and Albrecht, 2007; Breen and Salazar, 2011; Burtless, 1999; Chevan and Stokes, 2000; Daly and Valletta, 2006; Karoly and Burtless, 1995; Martin, 2006; Nielsen and Alderson, 1997; Western, Bloome and Percheski, 2008). During the mid-20th century, strong social and religious pressures tightly circumscribed the family structure in the United States and elsewhere, creating a situation in which the nuclear family (with the father as the sole breadwinner) was far and away the dominant family form. However, over recent decades, movements away from this family structure may have altered the distribution of income in some countries. The first change is the relative decline of the nuclear family and the concomitant rise of the single-parent family. For example, in the United States, the portion of children living with both parents has declined markedly over recent decades, falling from 90 percent of children in 1960 to only 68 percent of children in 2010 (US Census Bureau, 2011). The vast majority of these children live with their single mothers.

The studies cited above consistently find a positive association between the prevalence of single-mother families and income inequality. The common theoretical explanation for this finding is that families headed by single mothers face entrenched barriers to maintaining income parity with other types of households. These barriers can arise from three broad factors. (1) Gender pay gaps make it difficult for single mothers to earn equivalent incomes to comparably situated men. (2) The rise of dual-income families means that single-mother families often have one less income than other families and households. Finally, (3) at least in the United States, single mothers tend to be less educated than other women (McLanahan and Percheski, 2008). Combined, these three factors—lower pay for women, more dual-income families, and low educational attainments of single mothers—leave most single-mother families with household incomes well below the national median.

Is a similar phenomenon occurring in other Western countries? The answer is no according to Esping-Andersen (2007). He argues that the disequalizing effect of single motherhood is primarily limited to the United States, because of its meagre welfare state. However, given that all Western countries have noticeable pay gaps between men and women (Therborn, 2004: table 3.3), it seems unlikely that even the most generous welfare states could fully compensate for the barriers faced by single mothers. Recent cross-national research provides some tentative support for this alternative argument. Notably, in their broad analysis of income inequality, Bradley and his colleagues (2003) find a positive link between the prevalence of single-mother families and income inequality within a sample of 14 Western countries (see also Albertini, 2008; Kenworthy, 2008:45).

Furthermore, data from the LIS suggest that these factors may be causally linked across numerous countries. As shown in figure 2, the portion of children living in households headed by single mothers has increased (in some cases markedly) in all countries examined in this study except Austria, where it declined modestly. When compared to data on national income inequality (figure 1), it appears that some of the most unequal countries in the present study—Ireland, the United States, and the United Kingdom—also have some of the highest portions of children living with single mothers. Of course, this evidence is only anecdotal, but it does suggest that this issue should be investigated further.

[Insert figure 2 about here.]

Female Employment and National Income Inequality

The increased propensity of women to work outside the home is another family-related social change with possible implications for income inequality. During the post war era, when levels of income inequality were generally lower than present, most families in the West adhered to the male breadwinner model, in which the father worked full-time outside the home and the mother engaged in unpaid domestic work (Lewis, 1992). Over recent decades, however, this type of family has slowly declined in prominence as more women have taken jobs outside the home. Indeed, cross-national data suggests that, compared to the early 1970s, most countries in the present study have experienced almost a doubling of work-aged women in the workforce (see figure 2).

Although this trend is widespread, its effects on income inequality are disputed. According to several studies, increased levels of female labour market participation in the United States have contributed to rising income inequality in that country (Burtless, 1999; Karoly and Burtless, 1995; Schwartz, 2010). A central factor in this claim is “spousal or partner homogamy”—the tendency for spouses and partners to resemble one another in terms of their educational obtainments, class backgrounds, and career accomplishments. For instance, studies find that educational similarities among couples are ubiquitous in Western countries (Blossfeld and Timm, 2003), and becoming more common in the United States since the 1960s (Schwartz and Mare, 2005). When coupled with increased female employment, this phenomenon is thought to heighten income inequality by combining two high-income earners or two low-income earners into one household. This, of course, magnifies pre-existing socioeconomic advantages and disadvantages. However, the merits of this argument are disputed (Breen and Salazar, 2011; Cancian and Reed, 1999; Western, Bloome and Percheski, 2008). For instance, by decomposing household income inequality from the late 1970s and the early 2000s, Breen and Salazar (2011) find that changes in levels of educational homogamy had little effect on the distribution of income in the United States. However, given that their study assesses temporal effects within one country, their results leave open the possibility that spousal or partner homogamy can help explain variation in income inequality among countries.

Two points are worth noting about the possible distributional consequences of spousal or partner homogamy. (1) In order for this phenomenon to heighten income inequality, incremental increases in female employment need to be concentrated among households at the top of the income distribution. Otherwise, inequality will remain unchanged.[1] For the United States,

[1] The Gini coefficient is immune to equivalent changes occurring at the top and bottom of a distribution. For example, under a scenario in which low-income earners doubled their incomes and high-income earners doubled their incomes, the Gini coefficient of income inequality would remain unchanged.

Schwartz's (2010: 1542) study of married couples between 1967 and 2005 finds that this has indeed occurred—namely that employment rates among women with high-income husbands rose faster than employment rates among women with low-income husbands, even though the latter still work outside the home more than the former (see also Karoly and Burtless, 1995). Nonetheless, changes in the homogomy of spouses and partners would need to be significant to produce discernible effects on income inequality. (2) Available cross-national data suggest that educational homogamy between couples is indeed a prominent characteristic of all Western countries, but whether this trend is intensifying over time is unclear (Blossfeld and Timm, 2003).[2]

Although a consensus is lacking, the preponderance of evidence now supports the idea that increased female employment generally reduces income inequality. Numerous studies on income inequality in the United States reach this conclusion, albeit noting that the equalizing effect of female employment can vary from decade to decade, depending upon the types of women drawn into the workforce (Albrecht and Albrecht, 2007; Cancian and Reed, 1999; Chevan and Stokes, 2000; Daly and Valletta, 2006; Nielsen and Alderson, 1997; Treas, 1987; Western, Bloome and Percheski, 2008). While the theoretical rationale for these empirical findings is not always clear, it seems likely that the disequalizing effect of female employment (arising from homogamy among affluent couples) is usually more than offset by its equalizing effect (arising from lower-income households gaining additional sources of income).

To confuse matters further, cross-national studies on the determinants of income inequality reach inconsistent conclusions about the relationship between female employment and income inequality. For example, Gustafsson and Johansson's (1999) study of 16 OECD countries between 1966 and 1994 finds no link between female labour market participation and income inequality—a finding that is reasonably consistent with Breen and Salazar's (2010) study of the United Kingdom. However, despite analyzing the same 16 countries over nearly the same period, Alderson and Nielsen (2002) find that female labour market participation heightens income inequality—a finding that contrasts with their earlier study of the United States (Nielsen and Alderson, 1997). Finally, based on analysis of 14 OECD countries over similar decades, Bradley and his colleagues (2003) find that female labour market participation generally reduces income inequality.

How can these contradictory results be explained? To begin, it should be noted that three of the studies mentioned in the preceding paragraph do not control for changes in family structure (i.e. Gustafsson and Johansson, 1999; Breen and Salazar, 2010; Alderson and Nielsen, 2002). This omission may be decisive. It is my contention that part of the confusion over whether female labour market participation increases or decreases income inequality stems from the fact that it is

[2] As calculated from the LIS micro-data, the correlations between the educational obtainments of spouses and partners range from a low of 0.48 (Canada 2004) to a high of 0.89 (Italy in 1987).

bound up with the rise of single-mother families. These trends are bound together because single motherhood necessitates that, in the absence of full-support from the state or other sources, affected women must enter the labour market to support their families. When prodded by these circumstances, increased female employment should elevate income inequality, since it represents households with only one source of income, which is being earned by persons economically disadvantaged by gender and perhaps other factors (e.g. low educational levels).

Conversely, the remaining portion of increased female employment should help to reduce income inequality. This should occur because it represents either (1) married or partnered women entering the workforce to provide additional incomes for their families or households, or (2) single women without children earning incomes for themselves. Under the first scenario, as long as women from lower-income families and households continue to work in large numbers, rising female employment should moderate income inequality. This idea is consistent with Kenworthy's (2008:45) general contention that widespread labour market participation (either by males or females) helps to distribute income more equally across society.

DATA AND RESEARCH METHODS

Measurement and Data

The study's dependent variable measures inequality in the distribution of income across households. The data come from the LIS (2010), which is generally regarded as the best source of data for cross-national comparisons (Kenworthy, 2008:39). The LIS gathers detailed data from nationally representative household surveys, and then harmonizes these data to yield measures of income inequality that are consistent across countries and across years. The national surveys are conducted in "waves," occurring approximately every five years. Waves begin for some countries in the late-1970s, but for others in the mid-1980s. For select countries—including Sweden, the United Kingdom, and the United States—LIS also provides historical data starting in the late 1960s.

The LIS inequality measure used in this study has several properties that should be highlighted. (1) It measures disposable rather than gross income, meaning that it accounts for the moderating effects of progressive taxation and redistributive social programmes. This is preferable on sociological grounds, because net income captures the way people actually experience income and income inequality. (2) The LIS weights household income by the square root of the number of household members. This accounts for the economies of scale enjoyed by larger households, and results in what the LIS calls "equivalent income." (3) Inequality in the national distribution of equivalent income is expressed as a Gini coefficient, which theoretically ranges from 0 (each household has the same equivalent income) to 1 (a single household as all of the equivalent income and the others have none). To make the regression output more readable, I multiple the Gini coefficient by 100.

The study focuses on two independent variables. The first of these variables, *female labour market participation*, equals the number of women in the paid workforce as a percentage of all women aged 15 to 64. Data come from the Organization for Economic Cooperation and Development (OECD) (2010a). The other variable attempts to measure the prevalence of single motherhood, but is less straightforward. Using the LIS micro-data (LIS 2011), it is possible to calculate the percentage of households headed by single mothers living with their children.[3] This is the most direct measure of this phenomenon, but unfortunately it yields only 81 observations from 15 countries. This compares unfavourably to the 110 observations from 16 countries for the dependent variable. To help remedy this problem, the LIS (2010) offers a proxy variable that contains all 110 observations. This variable, the *percentage of children living with single mothers,* indirectly captures the prevalence of single motherhood and its potential effect on income inequality. For this reason, it will be used in the regression analyses. This choice does not affect the substantive conclusions of the study.[4]

The study also develops a measure of homogamy between spouses and partners. Using the LIS micro-data, *educational homogamy* is measured as the correlation between the educational obtainments of married or unmarried partners living in the same household.[5] Unfortunately, this measure suffers from two complications. (1) Educational obtainment is measured in idiosyncratic ways across the national surveys comprising the LIS micro-data. For example, the United States uses a 16-point scale to measure educational obtainment whereas the United Kingdom uses a four-point scale. To help remedy this situation, the LIS offers a recoded variable in which respondents are labelled as having either primary, secondary, or tertiary educations. This recoded measure is used in this study, but the standardization comes at the expense of eliminating some of the finer but possibly important distinctions in educational obtainment. (2) The LIS's standardized measure of educational obtainment is missing significant amounts of data. Specifically, it has only 68 of the possible 110 observations. Another measure of homogamy, one based on income rather than

[3] This measure requires the researcher to merge the individual- and household-level micro data, so that the variable "household type" can be broken out by the sex of the householder.

[4] The full model was also estimated with the *percentage of households headed by single mothers* instead of the proxy variable. This causes the number of observations to fall from 94 to 69. Nonetheless, the *percentage of households headed by single mothers* is statistically significant with the expected positive sign (b =.508*). Furthermore, two parameter estimates change in substantively meaningful ways: *union density* becomes statistically insignificant, and *population under 15* becomes statistically significant.

[5] This measure requires the researcher to change the format of the individual-level data from long to wide, so each partner's highest level of educational obtainment is located on the same row of the dataset. The correlation is based on Goodman and Kruskal's gamma, a measure of correlation designed for instances in which both variables are ordinal and neither constitutes a dependent variable.

education, was considered but not adopted.[6] Due to these limitations, the educational homogamy variable is not used in the full regression models.

Since the forces affecting the distribution of income are manifold and complex, the model controls for other factors that may influence income inequality. These factors are derived from the cross-national studies of income inequality cited above. Two variables account for variations in the composition of welfare states. *Welfare expenditures* measures total public outlays for social welfare programmes and transfers across all levels of government expressed as a percentage of GDP. Data come from the OECD (2010a). *Welfare generosity* gauges the replacement rates, duration, and eligibility criteria for three prominent welfare programmes: unemployment, disability, and old-age insurance. Higher scores indicate more generous benefits. Data come from Scruggs (2005). Two variables account for important differences in labour market institutions—*union density* (trade union membership as a percentage of the workforce) and *corporatism* (a five-point scale measuring the centralization and coordination of wage bargaining procedures). Data for both variables come from Visser (2011).

Several variables account for economic and social structures that may affect the distribution of income. *Industrial employment* accounts for deindustrialization, and *agriculture employment* accounts for lingering aspects of sector dualism as described by Alderson and Nielsen (2002). Both variables are measured as the number of workers in that sector as a percentage of the civilian workforce. Data come from the OECD (2010b). *Manufacturing trade with the South* measures economic linkages with low-wage countries. The South is defined as non-OECD countries, plus Mexico and Turkey, and manufactured goods are defined as categories five through eight of the international standard industrial classification scheme, revision two. Trade flows are expressed as a percentage of GDP. Trade data come from OECD (2011) and GDP data come from Penn World Tables (Heston, Summers and Aten, 2011). Following Alderson and Nielsen (2002), *gross domestic product* and its squared term are used as controls for broad social and economic changes linked to economic development. Data come from the Penn World Tables. Following Gustafsson and Johansson (1999), the size of economically inactive populations are captured by three variables. *Population under 15* measures children as a percentage of the population; *population over 65* measures pensioners as a percentage of the population; and *unemployment* measures the jobless as a percentage of the civilian workforce. Data come from the OECD (2010b).

[6] At least two problems beset this income-based measure of homogamy. (1) Some countries report individual earnings before taxes but others after taxes. This makes the measure inconsistent. (2) Only 88 out of 110 observations are present.

Statistical Estimation

The study analyzes 16 Western countries observed at unequally spaced intervals between 1967 and 2005. The 16 countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Sweden, Switzerland, the United Kingdom, and the United States. Japan and New Zealand are excluded, because the LIS has no data for these countries. The data set contains a maximum of 110 observations per variable.

Due to the panel structure of the data, standard applications of ordinary least squares (OLS) regression are inappropriate for this study. The main problem is that the country-year observations comprising the data are not independent of one another, as OLS regression requires, but rather linked in highly structured ways. In panel data, linkages occur in both the cross-sectional and temporal dimensions of the data. For example, all observations from the United Kingdom are linked cross-sectionally, whereas all observations from the year 2000 are linked temporally. When pooling such observation together, standard applications of OLS will yield misleading results if "unmeasured effects" are unintentionally omitted from the list of regressors.

This problem can be addressed in a variety of ways (Halaby, 2004). Perhaps the simplest approach uses country-dummy variables to capture unmeasured cross-sectional effects and year-dummy variables to capture unmeasured temporal effects. This type of model—known as a two-way, fixed effects (FE) model—can be estimated with OLS and the resulting parameter estimates will be unbiased and consistent. However, compared to the FEM, the random effects model (REM) is equally effective but yields more accurate p-values since its estimator is more efficient. Under this approach, unmeasured effects that vary across countries but not across time are modeled with country-specific error terms, which accompany the general error term for the whole model. An advantage of this approach is that it can estimate slow moving or time-invariant explanatory variables, such as the wage bargaining variable used in this study. For these reasons, I adopt the REM rather than the FEM approach, but augment the former by including a dummy variable for each year in the data set. This yields a two-way REM, in which unmeasured cross-sectional effects are captured with country-specific error terms and unmeasured temporal effects are captured with a battery of year-dummy variables.

A few additional steps are taken to improve the accuracy of the model. First, to control for the likely presence of groupwise heteroscedasticity, robust standard errors are employed. Second, to indentify potentially problematic outliers, the Hadi robust outlier detection algorithm (p .05) is used on the full model. This step, however, identifies no observations as outliers. Finally, the presence of serial-correlation is considered but not addressed for the following reasons: (1) AR(1) estimation cannot be undertaken with unbalanced data sets; (2) the REM with temporal dummy variables should already mitigate the presence of serial-correlation; and (3) given that serial correlation affects the estimates of the standard errors, not the estimates of the parameters, the use of robust standard errors should already mitigate possible distortions arising from serial correlation.

RESULTS

Table 1 shows result from five regression models, each capturing some portion of the effects of family structure and female employment on income inequality. Model 1 begins with a simple bivariate model of the distributional consequences of female labour market participation. As anticipated, the parameter estimate for this variable is statistically significant and exhibits the expected negative sign. Substantively, this result suggests that the growing presence of women in the paid workforce over recent decades has helped to reduce income inequality among households. Similarly, model 2 examines the bivariate relationship between my proxy for single-mother families and income inequality. The estimated parameter for this variable is found to be statistically insignificant, but it does exhibit the expected positive sign. This result suggests that the prevalence of single motherhood does not affect income inequality, a finding that is consistent with Esping-Andersen's argument that the disequalizing effect of single motherhood is confined to the United States. However, when examined in a multivariate context, this interpretation is no longer supported.

[Insert table 1 about here]

Model 3 is the last bivariate model. It examines the distributional effects of educational homogamy among spouses and partners. Recall that this variable is missing considerable amounts of data, and hence results from this model should be interpreted accordingly. Despite the missing data, the resulting parameter estimate supports the argument that the propensity of spouses and partners to resemble one another in terms of their educational obtainments does exacerbate income inequality. Like all models presented in table 1, the portion of the variance explained between countries is significantly larger than the portion of the variance explained within countries—meaning that the models explain more of the cross-national than temporal variation in income inequality. Since this is especially true of the educational homogamy variable, the findings from model 3 do not necessarily conflict with Breen and Salazar's (2011) recent finding that educational homogamy is unrelated to rising income inequality in the United States.

The remaining two models in table 1 examine how family structure and female employment affect income inequality in multivariate contexts. Recall that one contention being put forward in this study is that female labour market participation and single motherhood are interlinked with each other and income inequality, which means that their distributional effects must be considered simultaneously. Model 4 sheds light on this contention. Compared to the bivariate models discussed above, this multivariate model yields two interesting changes: (1) the parameter estimate for female labour market participation becomes larger, and (2) the parameter estimate for the percentage of children living with single mother also becomes larger (in absolute terms) and statistically significant. Importantly, these changes are consistent with the effect of female employment on income inequality being confounded by single motherhood.[7] Indeed, it seems likely that some women, over recent decades, entered the paid workforce because they became single parents and could not rely on financial support from spouses or partners. Hence, once the model accounts for the disequalizing effect of single motherhood, the remaining equalizing effect of female employment becomes stronger.

Model 5 introduces a control for educational homogamy. Recall that this variable has limitations, including missing data. The inclusion of this variable, however, does not alter my general argument that female employment and single motherhood exert meaningful but countervailing effects on the distribution of income across a range of Western countries. More specifically, under this model specification, all three variables are statistically significant and exhibit the expected signs. Furthermore, it is found that the inclusion of the educational homogamy variable reduces the equalizing effect of female labour market participation, a finding that is consistent with the general claim that educational similarities between spouses and partners can alter the distributional effects of female employment patterns.

[Insert table 2 about here]

To minimize possible distortions arising from omitted variable bias, table 2 introduces controls for labour market institutions, welfare state characteristics, and numerous social and economic structures. For easy of comparison, model 3 from table 1 is reproduced as model 6 in table 2. (Due to missing data, the educational homogamy variable is not use in the comprehensive models shown in table 2.) Model 7 builds on model 6 by adding controls for two basic labour market institutions. As expected, the results support the contention that trade union density and centralized wage bargaining are inversely related to income inequality. This is evidenced by both variables being statistically significant and exhibiting negative signs. More importantly, however, the introduction of these control variables does not materially change the estimated effects of the study's primary variables.

Model 8 continues in a step-wise fashion by introducing controls for the size and generosity of the welfare state. Here the expectation is that the amount of financial resources devoted to social welfare programmes and the generosity of these programmes (independent of whether they are used) will be inversely related to income inequality. The results from model 8 support this expectation, as both welfare state variables are statistically significant and exhibit

[7] Using the same regression techniques, a simple bivariate model finds that the percentage of children living with single mothers has a positive and statistically significant effect on female labour market participation (b= 1.18**).

negative signs. Again, the addition of new control variables has little substantive bearing on the estimated effects of the other variables in the model, which means that the net effect of female labour market participation and the percentage of children living with single mothers remain robust net of the other explanatory variables in the model. Impressively, this relatively parsimonious model explains 87 percent of the overall variance in income inequality found in this sample of 16 countries.

Finally, model 9 represents the full model of income inequality. It builds on model 8 by introducing controls for various economic and social structures that may affect income inequality. Under this comprehensive model, neither the size nor statistical significance of the parameter estimates of theoretical interest change in substantively meaningful ways. Most of the newly added variables prove to be statistically insignificant. Yet it is found that the size of the economically inactive population, especially those of retirement age, does exert some upward pressure on income inequality. Overall, the results from the final model provide strong support for the notion that family structure and female employment (along with the generosity of welfare states and the specificity of labour market institutions) are important determinants of income inequality.

CONCLUSION

Many sociologists maintain that a thorough understanding of income inequality and its dynamics requires a consideration of how incomes are pooled by families and households. Recently, two social changes have affected this pooling process, and consequently may have altered the distribution of income in some countries. (1) In nearly all Western countries, the portion of women working outside of the home has risen substantially over recent decades. This has given many low-income families and households new and valuable sources of income, but the tendency for couples to resemble one another in terms of their educational achievements and earning potentials may be offsetting some of this equalizing effect. Hence, the overall effect of female employment on income inequality is not fully understood. (2) In many but not all Western countries, the portion of households headed by single mothers has increased over recent decades. This is thought to heighten income inequality, because single-mother households usually have only one source of income, earned by persons economically disadvantaged by gender. Although this phenomenon is well documented for the United States, some scholars believe that generous welfare states prevent other Western countries from experiencing similar outcomes. On the whole, sociologists are well aware of these important social changes, but their distributional consequences are not fully understood.

Motivated to fill these gaps in the literature, this study has attempted to clarify whether these gender-related social changes have had distributional consequences for a sample of 16

Western countries. To make this assessment, the study uses income inequality estimates from the LIS, which start as early as the late 1960s for some countries and reoccur about every five years until mid-2000s. Since the forces that affect the distribution of income are manifold and complex, the regression models control for temporal and cross-national differences in labour market institutions, the provision of social welfare, and numerous economic and social structures. Results from two-way random effects regression models indicate that growth in female employment moderates income inequality, but that growth in single-mother families exacerbates income inequality. Limited evidence also suggests that the tendency for spouses and partners to have similar educational levels explains some of the cross-national variation in income inequality.

These findings have important substantive and theoretical implications. Substantively, the results confirm that these two social changes have real distributional consequences. On one hand, the near universal expansion of female labour market participation over recent decades has helped to cushion some of the disequalizing effects arising from other societal changes, such as the decline of organized labour and the retrenchment of social welfare provisions. For the sample as a whole, the percentage of working-aged women in the paid workforce increased from 48 percent in 1970 to 69 percent in 2005. Undoubtedly, without this pronounced social change, income inequality would be markedly higher in these 16 countries. On the other hand, the equalizing effect of rising female employment was undermined in some countries by growing portions of households headed by single mothers. This was particularly true of the United Kingdom and the United States, the two countries in this study with the highest levels of income inequality during the mid-2000s. Presently, in both of these countries, more than 20 percent of children live in households headed by single mothers. For the United Kingdom, this amounts to almost a fourfold increase from the early 1970s. Importantly, the overall results of this study imply that governments can reduce income inequality by encouraging the formation and maintenance of two-parent families, and by encouraging women to work outside the home. Notably, in countries where this has been accomplished—such as France, the Netherlands, and Switzerland—income inequality has not risen much over recent decades.

These finding also have implications for debates in the literature. Including the results presented here, there is now significant evidence suggesting that increased female labour market participation reduces income inequality, but that increased percentages of households headed by single mothers heightens income inequality. This has been demonstrated not only in the present study, but also in several studies of the United States (e.g. Albrecht and Albrecht, 2007; Western et al., 2008) and by Bradley and his colleagues' (2003) study of 14 Western countries. In each of these studies, scholars simultaneously consider how female employment and single-mother families affect income inequality. When these two factors are considered together, the results are consistent. When only female employment is considered, the results have been inconsistent (cf.

Alderson and Nielsen, 2002; Gustafson and Johansson, 1999; Nielsen and Alderson, 1997). It is my contention that these two trends must be considered together because single motherhood is itself a factor in expanding female labour market participation. Hence, if one factor is considered but the other is not, statistical models will conflate the two countervailing effects, leading to incorrect conclusions about the distributional consequences of these social changes. Finally, when better data become available, the present study can be improved by considering more fully how spousal and partner homogamy affects income inequality.

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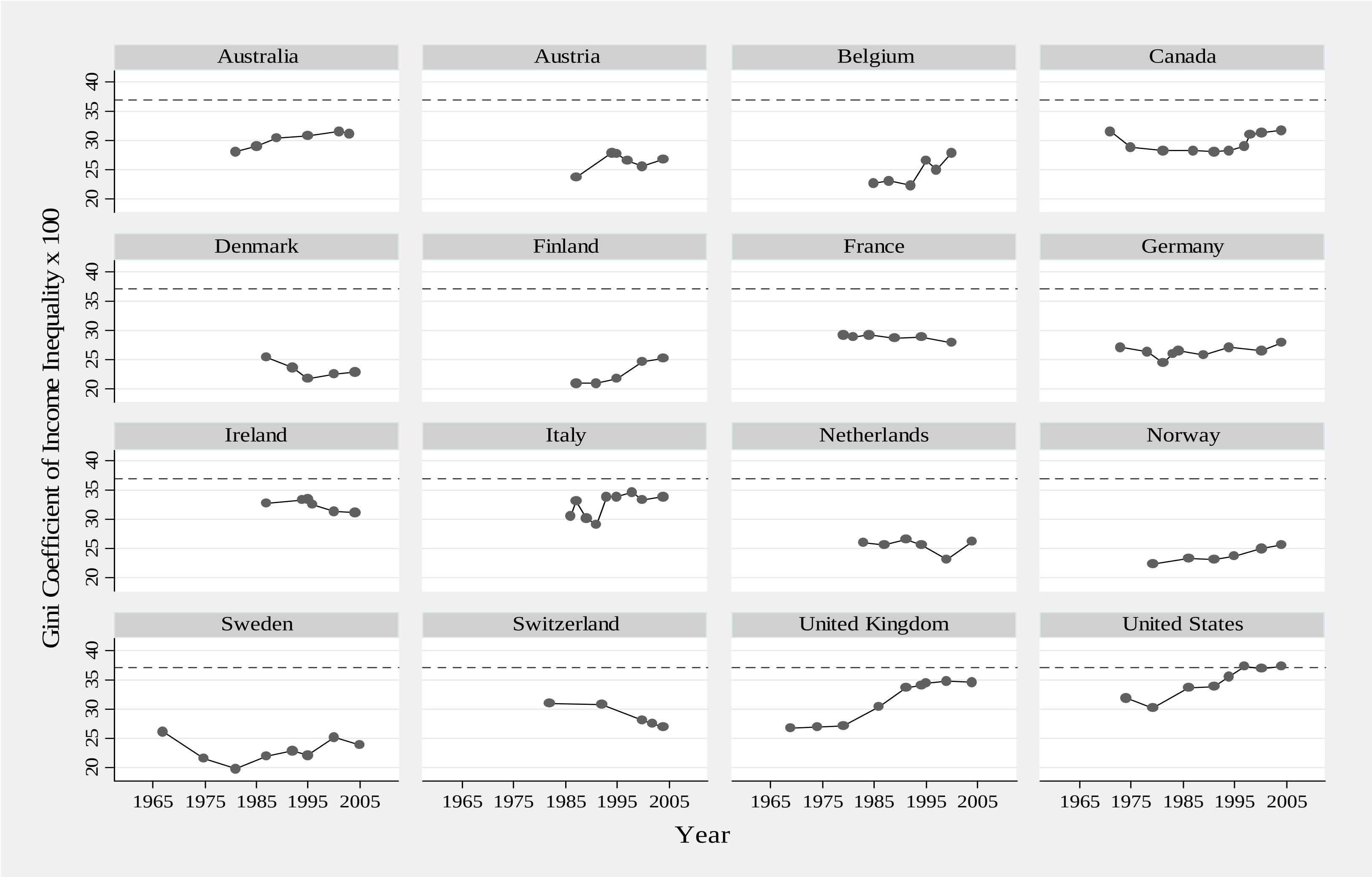
FIGURE 1. Gini Coefficient of Income Inequality for 16 Western Countries, 1967 to 2005



Note: Data from the Luxembourg Income Study (LIS 2010). Dashed lines represent highest level of income inequality in the sample—US 2004.

FIGURE 2. Changes in Family Structure and Female Employment in 16 Western Countries.

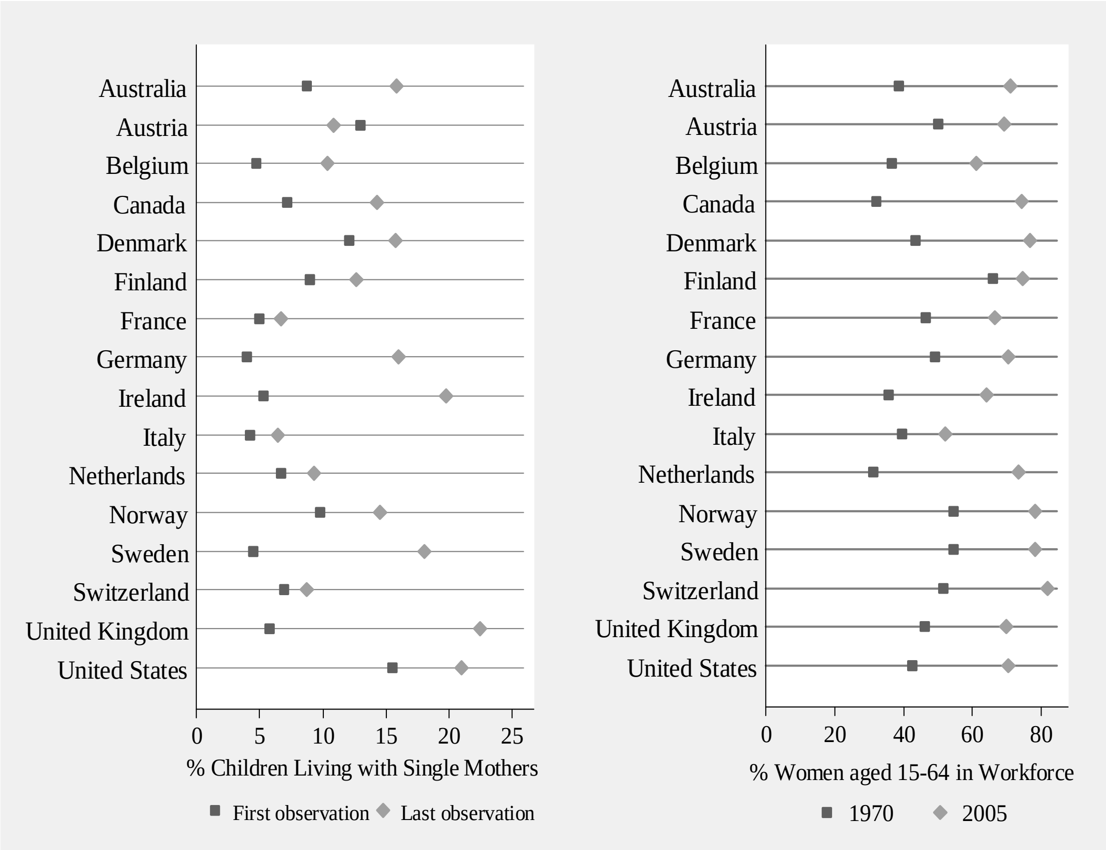


Table 1. Random-Effects Regression Estimates of Income Inequality (Gini Cofficient x 100): 16 Western Countries, 1967 to 2005

	(1)	(2)	(3)	(4)	(5)
Gender and Family					
Female Labour Market Participation	-0.184**			-0.320***	-.184*
	(.066)			(.079)	(.071)
Children with Single Mothers		0.169		0.547**	.443**
		(.307)		(.211)	(.164)
Educational Homogamy			.256**		.193*
			(.074)		(.073)
Constant	36.12***	25.19***	9.56	41.04***	15.81
	(3.63)	(1.39)	(5.96)	(3.91)	(10.67)
Number of Observations	110	110	68	110	68
Dummy Variable for Each Year?	Yes	Yes	Yes	Yes	Yes
R2 (within)	.123	.143	.091	.200	.120
R2 (between)	.513	.545	.547	.726	.708
R2 (overall)	.355	.229	.361	.546	.666

Note: Unstandardized parameter estimates with standard errors in parentheses. * = p < .05; ** = p < .01; *** = p < .001. Number of observations drop in models 4 and 5 due to missing data of educational homogamy.

Table 2. Random-Effects Regression Estimates of Income Inequality (Gini Cofficient x 100): 16 Western Countries, 1967 to 2005

	Models			
	(6)	(7)	(8)	(9)
Gender and Family				
Female Labor Market Participation	-0.320***	-.214***	-.162***	-.238***
	(.079)	(.059)	(.038)	(.048)
Children with Single Parents	0.547**	.407***	.320***	.299***
	(.211)	(.096)	(.067)	(.079)
Labour Market Institutions				
Union Density		-.082**	-.052***	-.037*
		(.026)	(.013)	(.183)
Centralized Wage Bargaining		-7.881**	-4.189*	-4.421*
		(2.69)	(1.866)	(2.115)
Welfare State				
Welfare Expenditures			-.161*	-.289**
			(.078)	(.096)
Welfare Generosity			-.250***	-.229**
			(.065)	(.078)
Social and Economic Structure				
Trade with the South				-8.991
				(13.006)
Industrial Employment				-.119
				(.093)
Agricultural Employment				.166
				(.106)
GDP				-.623e-3
				(.000e-3)
GDP-squared				.235e-8
				(.609e-8)
Population under 15				.212
				(.159)
Population over 65				.808***
				(.184)
Unemployment				-.131
				(.162)
Constant	41.04***	46.82***	46.68***	42.11**
	(3.91)	(3.01)	(2.62)	(18.69)
Number of Observations	110	110	94	94
Dummy Variable for Each Year?	Yes	Yes	Yes	Yes
R2 (within)	.200	.356	.341	.391
R2 (between)	.726	.891	.957	.976
R2 (overall)	.546	.756	.872	.932

Note: Numbers in parentheses are standard errors. * = p < .05; ** = p < .01; *** = p < .001. Number of observations drop in models 8 and 9 due to missing data on welfare generosity.