# Power Resources Theory and Inequality in the Canadian Provinces*

Michael Kellermann†

March 29, 2007

**Abstract**

In this paper, I demonstrate that higher levels of union membership and NDP provincial governments are associated with lower post-tax-and-transfer inequality in Canadian provinces. These results are consistent with the power resources theory of inequality and the welfare state first advanced by Korpi (1983) and Stephens (1979), which claims that differences in organizational resources such as unions and left political parties are responsible for differences in distributional outcomes. While many studies have found this association using cross-national data from rich democracies, the repeated use of data from the same set of countries raises the possibility that the relationship is due to unobserved country-specific characteristics. Using a pooled cross-sectional time series dataset from 1980 to 2003 and focusing on within-province variation in Canada, I find evidence consistent with the power resources model.

Over the past twenty-five years, power resources theory has provided one of the most influential accounts of variation in the size, characteristics, and outcomes of the welfare state. At its core, it asserts that working class power, achieved through organization by labor unions or left political parties, produces more egalitarian distributional outcomes (Korpi 1983, Stephens 1979). Relationships have been found between these variables and a number of measures of inequality and redistribution. In one recent contribution from the power resources school, Bradley et al. (2003) examine household income inequality before

*This paper prepared for the 2007 Annual Meeting of the Midwest Political Science Association, Chicago, Illinois. Thanks to Torben Iversen, Jim Alt, Federico Ferrara, and participants in workshops at Harvard for helpful comments on an earlier version of this paper.

†Ph.D. candidate, Department of Government, Harvard University (kellerm@fas.harvard.edu).

and after government redistribution, finding that union density decreases pre-tax inequality and left political parties achieve greater reductions in inequality through taxes and transfers. Unfortunately, these studies almost exclusively analyze data from the same set of rich industrial democracies, data used in a somewhat different form to generate the theory. New data sources are needed to determine whether the consistent relationship found between power resources variables and distributive outcomes within the OECD also applies in other contexts. In this paper, I take advantage of differences in inequality within a single country, Canada, to test the power resources approach. Following Bradley et al.(2003), I use market and post-tax-and-transfer household inequality as the outcomes of interest.

Canada provides an attractive setting in which to consider the relationship between power resources and inequality. Canadian provinces are both socially heterogeneous and have considerable political independence. Union density and support for the New Democratic Party (a left party with strong labor affiliations) vary across provinces and over time. While some studies have looked for differences in income inequality across provinces (MacPhail 2000, Gray et al. 2003), and others have examined the relationship between power resources variables and other dependent variables such as the growth of government spending (Petry et al. 2000), there has been little research on the political determinants of the level of inequality in Canadian provinces. In this paper, I use a cross-sectional time series dataset of Canadian provincial data from 1980 to 2003. The dependent variables are market and post-*fisc* income inequality among economic households headed by individuals under the age of 65.[1] I find that the power resources variables, union density and NDP government, have no effect on market income inequality, but that both reduce inequality after taxes and transfers conditional on the level of market inequality. These findings thus provide evidence consistent with the power resources model, evidence not obtained from the cross-national data traditionally used in these studies.

[1]Economic households are those with two or more individuals of any age.

The paper proceeds as follows. The first section briefly reviews the power resources model. This is followed by a critique of the repeated use of data from rich democracies and an argument in favor of using intra-country variation to test the implications of the model. The third section defends the choice of the Canadian provinces as an appropriate test case. The following two sections review the data and the estimation strategy used in the analysis. The central results, that union density and left party government are associated with lower post-tax-and-transfer inequality, are presented in the penultimate section. The paper concludes with a discussion of the implications of these results for power resources theory and the study of inequality more generally.

# 1 Review of Power Resources Theory

Power resources theory asserts that the level of inequality in a society and the accompanying degree of redistribution by the state are functions of the organizational resources of the working class (Korpi 1983, Stephens 1979). The model is firmly rooted in a class-based conception of politics in which capital and labor have opposite preferences over the level of inequality. Korpi (1983, 15) defines power resources as "characteristics which provide actors . . . with the ability to punish or reward other actors." In the absence of organization, through either labor unions or political parties representing their interests, individual workers have essentially no ability to reward or punish. This allows capital owners to control the vast majority of power resources in the economy. If workers organize, however, they can act collectively in order to pursue greater equality through economic or political action. While the working class can in theory organize in any sufficiently democratic society, "the degree of organization varies greatly across societies and through time within societies" (Bradley et al, 2003). In the power resources model, these variations in the capacity of the working class for collective action, and not differences in preferences, explain cross-national differences in distributive outcomes and the size and characteristics of government social policies. Korpi

(1998) argues that the power resources approach is essentially about strategic action, and one can think of it as implicitly positing a bargaining model between capital and labor in which the outcome is shaped by the outside options of the two parties.

Since the power of the working class cannot be observed directly, researchers must use proxies. Four variables have become standard measures of power resources in the literature: union density, union centralization, bargaining coordination, and the strength of left political parties (Korpi 1983, Robinson 1994). Union density measures union membership as a proportion of the labor force; this is the most basic measure of the degree to which labor has been organized for collective action. The next two variables, union centralization and bargaining coordination, capture variation in labor power due to organizational differences. When the labor movement is organized into encompassing union federations and when wages are negotiated at the national level (as opposed to negotiations at the level of the industry or the firm), its capacity for coherent action is greater. Finally, the strength of left political parties is thought to represent the degree to which labor can pursue its goals through political channels. These variables are highly correlated and causally interrelated; nevertheless, proponents of power resources theory suggest that the three measures of unionization should have a greater impact on the distributive outcomes generated in the labor market and left party strength should be more influential in determining variation in redistribution by the government (Bradley et al. 2003).

Of course, the power resources model is far from the only explanation for differences in the size, characteristics, and redistributive consequences of the welfare state. Other theories emphasize the interests of either capital or the state in creating redistributive institutions. The former suggests that industrial capitalism requires a certain degree of redistribution to maintain economic efficiency. The latter argues that differences in institutional structures mediate preferences over distributional policies. In countries where political power is diffused and many actors have the ability to hinder changes from the status quo, for example, the size and redistributive impact of state action is thought to be lower. Power resources arguments have also been criticized on their own terms. Orloff (1993) suggests that the traditional power resources model assumes an undifferentiated working class, overlooking differences in distributive consequences for men and women. Perhaps even more troubling, most early research in the power resources tradition implicitly assumed that the power resources of capital were comparable across countries and across time (Olsen 1992). If power resources theory is to be interpreted seriously as a bargaining model, focusing solely on the bargaining strength of only one side is problematic.

Despite these criticisms, researchers have produced a series of studies over the past twenty-five years suggesting that power resources, as defined above, explain variation in redistributive policies and outcomes even after accounting for alternate theories. Variables measuring the strength of unions and of left parties are associated with larger welfare states (Korpi 1983, Stephens 1979), more generous public pensions (Huber and Stephens 1993), lower market inequality and a greater proportional reduction in inequality through taxes and transfers (Hicks and Swank 1984, Bradley et al. 2003), and lower relative poverty (Moeller et al. 2003). These studies have increased in methodological sophistication, from the simple correlations used in early works (Korpi 1983, Stephens 1979) to more complex models designed to control for alternative explanations. Some recent studies suggest that the explanatory power of labor strength and left party governance has diminished as most industrial democracies have moved from a period of welfare state expansion to one of retrenchment (Huber and Stephens 2001), but others have found no evidence of decrease over time (Kwon and Pontusson 2003).

## 2 The need for new data

The evidence regarding the importance of power resources is less convincing than the preceding litany of results might suggest. Almost all tests of the power resources model analyze data from the same set of industrial democracies.[2] While recent studies certainly reflect improvements in both statistical methods and data quality, it is difficult to argue that they constitute new evidence in favor of the power resources model. One cannot exclude the possibility that unobserved national characteristics, correlated with union and left party strength, account for the consistent pattern of distributive outcomes demonstrated by these studies. To put it concretely, Sweden's relatively low inequality (larger welfare transfers, low relative poverty, etc.) might not be due to its strong unions and history of Social Democratic government, but rather to the egalitarian preferences of Swedish voters. The inferential problems that arise from repeatedly estimating versions of the power resources model on the same data are compounded by the use of either balanced or unbalanced time series-cross sectional data. Researchers have turned to this data in order to allow for more complex models, but the resulting estimates are likely to be overconfident. This problem is particularly problematic when most of the variation is between rather than within countries.[3] In order to test adequately the power resources model, new data must be found from sources other than cross-national comparisons of the rich democracies.

Variations in the distributional outcomes observed in subnational political units offer a new source of data on which to test the power resources model. If the association between union strength, left party government and greater equality is not simply a consequence of unobserved national characteristics, then it would not be unreasonable to expect that the same relationship holds true within countries as well as between them. If, on the other

[2]Korpi (1983) considers Norway, Finland, Sweden, Denmark, Belgium, the Netherlands, Austria, Germany, France, Italy, Britain, Ireland, Japan, Australia, New Zealand, Canada, and the United States. Twenty years later, Bradley et al. (2003) drops Austria, Ireland, Japan, and New Zealand and adds Switzerland.

[3]See Goodrich (2004) for a discussion of these problems.

hand, the relationship is spurious, it is more likely to be discovered in data that was not used to generate the original model. There are some issues associated with the use of intra-country variation to evaluate this theory. Proponents recognize that external economic and social pressures affect the ability of working class organizations to achieve their desired distributional outcomes even at the national level. One would expect that these outside constraints are greater when looking at the balance of power resources in subnational regions. The level of inequality in a given subnational unit at a given time will be affected by national economic forces, not to mention the distributive policies of the national government, both of which would tend to reduce the impact of differences in working class organization at the subnational level.

States and provinces differ from countries in many ways, not the least of which is the degree of labor and capital mobility. These differences find a parallel in recent discussions about the implications of globalization for the power resources approach (Bradley et al. 2003). Some have argued that the increased capital mobility associated with globalization has allowed capital owners to resist the demands expressed through unions and left parties; capital mobility between subnational units is generally far greater, and so the same logic should apply. At the same time, labor mobility is far higher within countries than it is between them. The effects of labor mobility on the bargaining power of the working class are ambiguous. If mobility is beneficial for capital, it might be beneficial for labor as well by providing a higher threat point. At the same time, the possibility of exit might encourage workers to leave rather than press for redistributive policies, even if organizations designed to express the views of labor already exist. Taken together, these differences are likely to attenuate the effects of variation in the degree of working class organization on distributional outcomes; one would expect the reduction in inequality, for example, associated with socialist government at the state level to be lower than at the national level. Nevertheless, *if* the power resources model is right, it is difficult to imagine a scenario in which increased labor organizational power would produce greater inequality, less redistribution, and smaller welfare programs in subnational political units, holding all else equal.

Prospective countries must satisfy two criteria in order to allow a fair test of the power resources approach. First, subnational units must have sufficient political autonomy to achieve different distributive outcomes. Ideally, one would want to look for a country in which states or provinces could both impose income or capital taxes and could determine the generosity of transfer programs.[4] Particularly in regard to redistribution, one would not expect the extent of union or left party strength to greatly affect policies if subnational units primarily implement policies determined at the national level. Second, the indicators of labor power resources must vary across subnational units. It is unlikely that variation in union density, for example, would be as great within countries as it is across the industrial democracies typically studied; on the other hand, these differences are more likely to represent actual variation in the capacity of labor for collective action and less likely to be a consequence of unobserved national characteristics. For this study, I use data from Canada to test the power resources theory as it relates to pre and post-*fisc* inequality.

# 3 Why Canada?

Canada offers a promising context in which to study power resources within a single country, satisfying both of the requirements discussed above. Canadian provinces have extensive autonomy in areas affecting distributive outcomes. Canadian provinces are able to impose income taxes on their residents. Payroll taxes are generally dedicated to fund health care, but they function as general revenue taxes in some provinces (Kesselman and Cheung 2004). In the labor market, provinces determine the level of the minimum wage. Quebec operates an independent pension system, and other provinces have considered similar programs. Provinces also operate means-tested social assistance programs; when Ontario reduced the

[4]This militates against choosing Australia, for example, because the power to impose income taxes was taken away from state governments during World War II and has never been returned.

level of its benefits by 21% in 1995, the poverty rate increased from 4% to 6% (Osberg and Xu 1999). In short, without minimizing the importance of the federal government, Canadian provinces possess instruments with which to influence income inequality.

The key power resources variables also vary across the Canadian provinces. While the Liberal and Conservative parties, both of which fit the definition of cadre parties, have traditionally dominated Canadian politics, the New Democratic Party defined itself from the beginning as a programmatic Socialist party. The Commonwealth Cooperative Federation (an agrarian Socialist party) and the Canadian Labour Congress founded the NDP in 1961. It differs from European Socialist parties primarily in its political weakness. Despite the fact that over half of its members are from union households, it has never won a plurality of the union vote in a national election (Cross and Young 2004). The NDP has been more successful at the provincial level, forming the government once in Ontario (1990-1995) and on multiple occasions in Saskatchewan, Manitoba, and British Columbia. It has historically been very weak in Quebec and the Atlantic provinces.[5]

Turning to the other standard indicators of labor power resources, union density in the Canadian provinces varies significantly; mean density over the past thirty years was highest in Newfoundland at 48% and lowest in Alberta at 25.5%. For the other two union-related variables, union centralization and bargaining coordination, there is essentially no variation. Unions in Canada are closely linked to their counterparts in the United States; both feature low levels of centralization and weak federations, and negotiations are conducted at the firm level (Robinson 1994). Thus, union density is the primary sources of variation to be found in Canada.

The Canadian case also provides an attractive institutional environment in which to test the power resources model. Legislative institutions and electoral systems are homogenous across the provinces. Parliaments elected under majoritarian electoral rules tend to produce

[5]The Parti Quebecois, an advocate of separation from Canada, also has social democratic leanings, but is not organized as a left party in the same way as the NDP.

single-party governments with strong executives; governments that win office can generally enact their programs. While minority governments occur occasionally, formal coalitions are rare, reducing the need to worry about portfolio allocation within cabinets. The provinces also face common interest and exchange rates and (to a somewhat lesser extent) a common capital market regime. Taken together, these reduce the number of potentially confounding variables in the analysis. The main drawback of using Canadian data is the small number of provinces; with only ten, one must rely primarily on within-province variation to draw reasonable inferences.

Until recently, income inequality in Canada, whether before or after taxes and transfers, has been relatively understudied (Mulé 2001, 74). Internationally, Bradley et al. (2003) estimates that Canada falls just above the median on both pre and post *fisc* inequality, measured by the Gini coefficient. Inequality in Canada appears to have fallen in the 1970's, in part due to expanded federal programs for redistribution. Market income inequality then rose during the 1980's and early 1990's (MacPhail 2000, Heisz et al. 2001). Several recent studies have examined differences in income inequality between the provinces. Gray et al. (2003) and Osberg and Xu (1999) find relatively small but significant differences across provinces in the Theil index and poverty incidence rates, respectively. Using regional data, Sharpe and Zyblock (1997) finds that higher unemployment is associated with higher levels of inequality, while Countryman (1999) finds that unemployment insurance (a federal program) has the greatest inequality-reducing effect in provinces where unemployment is high. Perhaps the most relevant to this study is MacPhail (2000), which shows that unemployment increases and union density decreases the degree of inequality in annual earnings and hourly wages.[6]

Cross-national studies in the power resources tradition generally consider Canada to be a liberal welfare state comparable to the United States, with low levels of unionization and a weak left party (O'Connor 1998). In addition to the findings on union density and wage

[6] This study uses sub-provincial regions rather than the provinces themselves, and does not consider post-tax-and-transfer inequality.

inequality, there is some qualitative evidence that labor power, either directly or mediated through the NDP, shaped the expansion of the Canadian welfare state. Mulé (2001) attributes the increase in federal redistributive programs, including unemployment insurance and family allowances, to the threat posed by the NDP to the governing Liberal party, which allowed welfarist Liberals to overcome opposition within their own ranks. The creation of a universal government-funded health care program by the NDP government in Saskatchewan led to pressure for a national system, which was created in 1966.[7] No systematic quantitative analysis, however, has examined the link between power resources variables and distributive outcomes. It is to such an analysis that this paper now turns.

# 4 Data

In addition to the substantive benefits described above, analyzing the effects of power resources variables on distributional outcomes within a single country allows for the use of more consistent and abundant data. Only recently, under the auspices of the Luxembourg Income Study, has data satisfying basic requirements of comparability been made available for cross-national research (Atkinson and Brandolini 2001). Even with this new data, Bradley et al. (2003) must resort to an unbalanced panel setup, with at most five observations per country. As a result, cross-national variance dominates the data. This study uses annual data, generally collected by Statistics Canada, covering the period from 1980 to 2003. Data is available for each of the ten provinces. Summary statistics by province are provided in Table 1; references to the data series used appear in the appendix.

As the dependent variables in this study, I measure household income inequality using Gini coefficients calculated from market income and from income after taxes and transfers.

[7] Olsen (1998) points out that, while Canada looks much like the United States in terms of income inequality, its health care system is quite redistributive and looks much more like those of the Social Democratic welfare states.

Table 1: Descriptive statistics: Mean values by province and correlations with inequality

| | pregini | postgini | unionden | ndpgovt | unemp | gdppercap | femlab | youth | netmigr | singpar |
|---|---|---|---|---|---|---|---|---|---|---|
| Canada | 38.44 | 30.00 | 34.14 | 0.00 | 9.17 | 27.90 | 66.10 | 22.08 | NA | 8.27 |
| | | | | | | | | | | |
| Newfoundland | 44.57 | 29.81 | 49.22 | 0.00 | 17.67 | 18.28 | 52.69 | 24.55 | -0.74 | 6.61 |
| Prince Edward Island | 37.92 | 27.03 | 30.60 | 0.00 | 13.81 | 18.88 | 67.37 | 23.90 | 0.03 | 8.39 |
| Nova Scotia | 38.97 | 29.12 | 31.49 | 0.00 | 11.59 | 21.04 | 61.36 | 21.96 | -0.06 | 8.57 |
| New Brunswick | 40.50 | 28.98 | 34.44 | 0.00 | 12.41 | 21.01 | 59.65 | 22.50 | -0.15 | 8.46 |
| Quebec | 39.08 | 28.70 | 40.08 | 0.00 | 11.09 | 24.61 | 61.20 | 21.06 | -0.19 | 8.63 |
| Ontario | 37.23 | 29.70 | 30.72 | 21.74 | 7.79 | 30.80 | 69.01 | 21.77 | 0.07 | 7.94 |
| Manitoba | 36.33 | 28.67 | 35.94 | 43.48 | 7.23 | 24.46 | 69.77 | 23.39 | -0.44 | 8.63 |
| Saskatchewan | 38.00 | 30.00 | 34.17 | 60.87 | 6.67 | 24.95 | 67.90 | 24.94 | -0.60 | 9.07 |
| Alberta | 36.57 | 29.91 | 25.96 | 0.00 | 7.44 | 33.24 | 71.24 | 24.38 | 0.28 | 8.10 |
| British Columbia | 37.32 | 30.04 | 38.74 | 43.48 | 9.84 | 28.46 | 67.07 | 21.15 | 0.47 | 8.57 |
| | | | | | | | | | | |
| Corr. w/ Market Inequality | 1.00 | 0.10 | 0.75 | -0.44 | 0.85 | -0.66 | -0.96 | 0.18 | -0.56 | -0.70 |
| Corr. w/ Post-*fisc* Inequality | 0.10 | 1.00 | 0.18 | 0.37 | -0.31 | 0.53 | -0.03 | 0.03 | 0.00 | -0.20 |

Notes: Means calculated using data from 1980 to 2003. Figures for Canada reflect the country as a whole, not the average of the provinces. Correlations are calculated using data from the ten provinces only. The variables are as follows: pregini - market inequality, measured by the Gini coefficient; postgini - post-tax-and-transfer inequality, measured by the Gini coefficient; unionden - union density (% of workers); ndpgovt - NDP government (years in office as % of years); unemp - unemployment rate (% of labor force); gdppercap - real GDP per capita (thousands of $CDN); femlab - female labor force participation (% of female population); youth - population aged 15 and younger (% of population); netmigr - net migration (% of population); singpar - % of economic families headed by lone parents.

Several measures of inequality exist; the Gini emphasizes differences in the middle of the income distribution, while others give more weight to differences at the upper or lower ends of the distribution (Mulé 2001). Gini coefficients can be calculated by plotting the proportion of aggregate income earned by the lowest $n$th percent of the population (the Lorenz curve); the Gini is calculated by doubling the area between this curve and a 45 degree line. Gini coefficients thus range between zero and one, with zero representing perfect income equality and one representing a society in which one household receives all income. Between these values, however, there is not a unique mapping between income distributions and Gini coefficients. While no single measure can capture all aspects of inequality, Gini coefficients are a reasonable choice and are widely used in the literature.

Using two measures of inequality allows distributive outcomes to be studied before and after government redistribution. As calculated, the pretax Gini includes earnings from wages, self-employment income, private pensions, and investment income (excluding capital gains). The post-tax-and-transfer Gini is calculated after adding transfers and subtracting income taxes (but not most payroll taxes) (Statistics Canada 2003).[8] Following Bradley et al. (2003), the Gini coefficients used in this study exclude households headed by individuals aged 65 and over; this focuses the study on distribution across classes rather than across age groups.[9]. The data used in this study is not directly comparable because it includes only economic families, which are defined by Statistics Canada to exclude unattached individuals. Thus, these measures capture income inequality in a subpopulation that includes single adults with children as well as couples or other household units with or without children.[10]

I now turn to the variables associated with power resources theory. Union density is

[8] Transfers include public pensions, unemployment and worker's compensation benefits, social assistance, and refundable tax credits for children and low-income individuals, among others.

[9] Bradley et al. (2003) use only data for households with heads between the ages of 25 and 59

[10] Whether this is good or bad is an open question. On the one hand, there is no theoretical reason to exclude single individuals. On the other hand, these Gini coefficients were calculated without adjusting for household size using some equivalence scale; excluding unattached individuals makes the sample considerably more homogenous.

generally measured as the proportion of employees who belong to trade unions. Data for the period 1980-1995 is available. To complete the time series, I follow McDonald and Myatt (2004) and divide the number of employees covered by unions by total employees. This data is available from 1997 forward; the missing year is interpolated by taking the mean of 1995 and 1997. These two series appear highly consistent with the exception of Newfoundland, where the change exceeds 10 percentage points.[11]

While one could use the share votes won by the NDP to measure the strength of labor-supported political parties, I create an indicator variable corresponding to the party holding the premier's office on the first of January of each year. All NDP provincial parties are institutionally attached to the federal NDP, so it is reasonable to code them together. It is not clear that this applies to the other parties; during the period considered in this study, two parties (Social Credit and Parti Quebecois) only held office in a single province, while the provincial branches of the Liberal and Conservative parties have weak to nonexistent institutional links and heterogeneous ideologies (Cross and Young 2004).[12] Using tenure in government emphasizes the influence of the NDP as an organization, and somewhat reduces the probability that a relationship between left party strength and distributional outcomes is due to differences in preferences. I use the party in office on January 1st, rather than use fractional values for election years, in recognition of the fact that it takes some time for incoming governments to change policies.

Given the small number of provinces, it would be exceedingly difficult to estimate the cross-sectional effects of union density and NDP governance while controlling for other variables. Examination of the bivariate relationships does not support the power resources model. Figure 1 plots the mean pre-tax (circles) and post-tax (triangles) Gini coefficients from 1980 to 2003 against mean union density and the proportion of time spent in office by

[11]The distinction between coverage and density is less important in Canada than it is in some European countries.

[12]For example, the current Liberal premier of Quebec used to be the leader of the federal Progressive Conservatives, while Conservative premiers in Alberta and Ontario were close to the federal Reform Party.

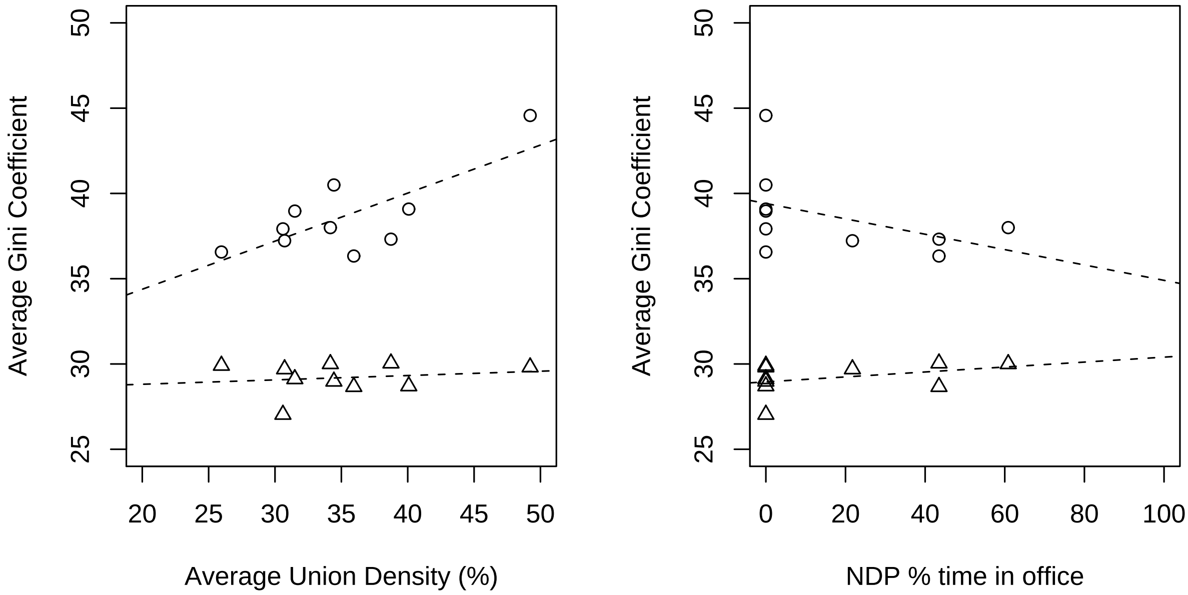


Figure 1: Provincial average market (circle) and post-*fisc* (triangle) Gini coefficients

---

the NDP. Contrary to expectations, the relationship between market inequality and union density is positive. Newfoundland, in the upper right, is clearly pulling the line up, but the relationship is (marginally) positive even after Newfoundland is dropped. The bivariate relationship between after-tax inequality and union density is essentially flat. It is difficult to judge the strength of the NDP relationships because six provinces have never elected NDP governments.

While there is not sufficient data to allow for more complex analyses of between-province variation, the use of time series data does allow for other factors to be controlled in the analysis of within-province variation. Two types of variables should be included as controls. The first are variables that provide proxies for rival explanations to the power resources model, while the second are those that control for asymmetric impacts of factors external to provinces that affect distributional outcomes.

The control variables used in this study parallel those of Bradley et al. (2003) as much as possible given the available data. Unemployment is included to control for variations in the business cycle; to the extent that unemployment benefits are available, one would expect that higher unemployment would produce lower post-tax-and-transfer inequality, holding market inequality constant. GDP per capita is typically included in inequality studies as a control variable, justified either as a proxy for the level of development or with the argument that redistribution is a luxury good. Since household income inequality is the basis for calculating the Gini coefficients, changes in female labor participation could affect distributional outcomes. The proportion of the population 15 years and under is included to control for the effects of federal redistributive programs targeted at children; seniors are not included since the dependent variables measure inequality among non-elderly households. The percentage of economic families that are lone-parent households taps in to changes in household composition over time, which as households become more heterogenous is likely to increase inequality. Finally, net interprovincial migration is included, both for its potential direct effect on inequality due changes in population structure and because of the concern that labor mobility may affect the power resources of the working classes.[13] Some variables traditionally included as controls in cross-national studies do not need to be used because there is no variation in the Canadian data. These include constitutional structure, capital market openness, and bargaining coordination, among others. Examining the bivariate correlations in Table 1, the strongest correlations are between market income inequality and both female labor force participation and unemployment. None of the correlations with post-*fisc* inequality are particularly strong. With only ten data points, the bivariate correlations are highly vulnerable to outlying cases; the negative correlation between market inequality and the percentage of lone-parent families is entirely due to Newfoundland, which has the highest average level of inequality and the lowest average rate of single parentage in the sample. The next section describes the statistical model to be used in parsing out the

[13] I looked for, but was unable to find, satisfactory measures of educational differences, which is also used by Bradley et al. (2003).

information contained in this data.

# 5 Estimation strategy

This study uses pooled time-series cross-sectional data, which typically presents several complications: temporal dependence of observations within a unit through time, correlation of errors across units by time period, and heteroscedasticity (Beck and Katz 1995, 1996). In recent years, the Beck and Katz approach to TSCS data has been standard in political science. This involves including a lagged dependent variable with the other explanatory variables to deal with autocorrelation, and then using panel-corrected standard errors to compensate for differences between units. Goodrich (2004) points out that this approach produces unreasonable estimates unless fixed effects are included for each of the cross-sectional units. I follow this amended strategy in this study, which has the effect of using only within-province variation to estimate the effects of the explanatory variables.

The nature of the dependent variable makes it particularly important to allow for temporal dependence and errors correlated across units during a given year. Prior research has shown that market income inequality, at the very least, has strong temporal dependence in Canadian provincial data (Wang and Ogwang 2004). One would also expect that common shocks would affect inequality in the provinces. If the federal government, for example, doubled the level of unemployment benefits in 1985, the fitted values for post-*fisc* inequality in that year would almost certainly be too high in every province. Rather than attempting to include variables that capture federal distributive policy directly, I include time dummies to correct for the common shocks. This, combined with the provincial fixed effects, has the effect of de-meaning the data on both the cross-sectional and time dimensions. Finally, I use panel-corrected standard errors to compensate for remaining panel heteroskedasticity. This estimation strategy is quite conservative; it allows for different intercepts by province and by year, but assumes that the slopes of the explanatory variables are the same across provinces and across time.

I estimate two models, the first with market income inequality as the dependent variable, and the second for post-tax-and-transfer income inequality. The models are not symmetric; in a given year, market income inequality is assumed to precede post-*fisc* inequality. In the short run, this seems like a defensible assumption, since the amount of most taxes and transfers are contingent on market income.[14] Thus, the model for market income inequality has the form:

$$\begin{aligned} marketgini_t &= \alpha + \phi * marketgini_{t-1} + \beta_1 * uniondensity_t + \beta_2 * ndp_t \\ &\quad + controls_t + provinces + years + \varepsilon. \end{aligned} \tag{1}$$

The model for post-tax-and-transfer income inequality is similar, but with the addition of contemporaneous market income inequality on the right-hand side.

$$\begin{aligned} posttaxgini_t &= \alpha + \phi * posttaxgini_{t-1} + \gamma * marketgini_t + \beta_1 * uniondensity_t + \beta_2 * ndp_t \\ &\quad + controls_t + provinces + years + \varepsilon. \end{aligned} \tag{2}$$

For each dependent variable, I begin by estimating the model with all control variables included. I then remove variables with $t$-statistics less than 1.65 and re-estimate the model. The goal of this exercise is not to search for the "best" model, but to consider the robustness of the estimates for the power resources variables.

[14] This does not rule out the possibility that, in the very long run, post-*fisc* inequality affects market inequality, by changing educational outcomes, for example.

# 6 Results

Table 2 presents the results for models with market income inequality as the dependent variable. Model (1) includes all of the control variables. Model (2) drops net migration and population 15 and under, both of which have standard errors larger than the coefficient estimates, but retains the two power resources variables. Model (3) eliminates the power resources variables. The results of these models suggest that market inequality is characterized by temporal dependence, but the estimated coefficient for lagged inequality is approximately 0.33, which indicates that much of the variation is due to other factors (including the year fixed effects). There is no evidence in these results to indicate that changes in power resources within provinces over time affect the level of market inequality in those provinces. The estimated coefficients on union density and NDP governments have the expected signs, but are statistically indistinguishable from zero. The lack of a stronger effect for union density, in particular, is somewhat surprising given the results of MacPhail (2000). That study examined wage inequality in particular, while the measure of inequality used in this paper includes non-wage market income and aggregates by household. It may be the case that whatever egalitarian effects union density has on wage dispersion are overwhelmed by these aggregation effects.[15]

Within provinces, higher levels of unemployment and higher percentages of lone-parent families increase market inequality, while higher female labor force participation and GDP per capita decrease inequality. That unemployment should increase market inequality is intuitive; when unemployment is high, more households have members who receive no wages. Since average female labor force participation is above 50%, increases in female labor participation holding unemployment constant are likely to result in more two-earner households, thus reducing household inequality. Conversely, increasing the number of lone-parent families

[15] It is also the case that MacPhail (2000) relies primarily on cross-sectional variance; the statistical model includes time dummies but no unit fixed effects, since only three time periods are observed for each unit.

Table 2: Determinants of market income inequality

| | Dependent variable: Market inequality$_t$ | | |
|---|---|---|---|
| | (1) | (2) | (3) |
| Market inequality$_{t-1}$ | 0.322 | 0.329 | 0.334 |
| | (0.074) | (0.074) | (0.074) |
| Union density (%) | -0.049 | -0.024 | – |
| | (0.047) | (0.042) | |
| NDP government | -0.181 | -0.159 | – |
| | (0.261) | (0.262) | |
| Unemployment rate (%) | 0.159 | 0.127 | 0.134 |
| | (0.082) | (0.074) | (0.075) |
| GDP per capita (thousands) | -0.146 | -0.144 | -0.142 |
| | (0.085) | (0.076) | (0.076) |
| Female labor force participation (%) | -0.296 | -0.311 | -0.301 |
| | (0.067) | (0.064) | (0.064) |
| Population 15 and under (%) | 0.100 | – | – |
| | (0.117) | | |
| Net migration (%) | 0.089 | – | – |
| | (0.265) | | |
| Lone-parent families (%) | 0.270 | 0.255 | 0.223 |
| | (0.105) | (0.105) | (0.099) |
| N | 230 | 230 | 230 |
| $R^2$ | 0.892 | 0.892 | 0.891 |

Notes: All models estimated using OLS regression. Numbers in parentheses are panel-corrected standard errors. Each model includes a full set of province and time dummies (not shown).

should increase the number of single-earner households, increasing household inequality. The logic behind the GDP effect is less obvious; one could resort to a Kuznets curve argument, but its relevance for within-province variation in Canada is not immediately obvious.

Turning to household inequality after taxes and transfer, Table 3 presents the results of three models. Model (1) again includes all of the control variables. Model (2) drops GDP per capita, net migration, and female labor force participation. Model (3) eliminates the percentage of lone-parent households and population 15 and under, both of which have estimated coefficients between one and two times the magnitude of their standard errors. The estimated coefficients for the remaining variables are quite similar across models, providing

Table 3: Determinants of post-tax-and-transfer income inequality

| | Dependent variable: Post-*fisc* inequality$_t$ | | |
|---|---|---|---|
| | (1) | (2) | (3) |
| Post-*fisc* inequality$_{t-1}$ | 0.022 | 0.017 | 0.037 |
| | (0.036) | (0.035) | (0.035) |
| Market inequality$_t$ | 0.674 | 0.661 | 0.660 |
| | (0.027) | (0.025) | (0.025) |
| Union density (%) | -0.039 | -0.038 | -0.027 |
| | (0.016) | (0.016) | (0.013) |
| NDP government | -0.210 | -0.217 | -0.252 |
| | (0.089) | (0.089) | (0.087) |
| Unemployment rate (%) | -0.150 | -0.154 | -0.175 |
| | (0.034) | (0.030) | (0.029) |
| GDP per capita (thousands) | -0.004 | – | – |
| | (0.028) | | |
| Female labor force participation (%) | 0.032 | 0.044 | – |
| | (0.026) | (0.026) | |
| Population 15 and under (%) | 0.079 | 0.067 | – |
| | (0.045) | (0.042) | |
| Net migration (%) | 0.039 | – | – |
| | (0.088) | | |
| Lone-parent families (%) | -0.067 | -0.067 | – |
| | (0.039) | (0.039) | |
| N | 230 | 230 | 230 |
| $R^2$ | 0.942 | 0.941 | 0.939 |

Notes: All models estimated using OLS regression. Numbers in parentheses are panel-corrected standard errors. Each model includes a full set of province and time dummies (not shown).

some comfort that the results are not highly model dependent. The magnitude of the coefficient on contemporaneous market inequality is unsurprising; the two variables are clearly linked.[16] Somewhat more surprising is the estimated coefficient on lagged post-*fisc* inequality, which is essentially zero. The immediate implication of this estimate is that, controlling for contemporaneous market inequality, there is not a lot of temporal dependence in the dependent variable. Given the rate at which the economy and public policy change, such a result is somewhat implausible, and bears further consideration.

[16] In addition, one would assume that measurement error in the two measures of inequality is correlated by year. Including pre-tax-and-transfer inequality thus should control for some of this error.

The results of these models are consistent with the implications of the power resources model. Both left party government and union density decrease the expected level of post-*fisc* income inequality, conditional on the other variables. Using panel-corrected standard errors, the estimated coefficient on both variables is statistically significant at .05 in all three models. More importantly, however, is the consistency of the magnitudes on both variables across models. NDP control of the provincial government is associated with an expected decrease of .2 to .25 points in the Gini coefficient, while an increase of one percentage point in union density is estimated to decrease the post-tax-and-transfer Gini by approximately .02 to .04 points. These coefficients are undoubtedly small, but not substantively insignificant. The average standard deviation of post-*fisc* inequality within provinces is just over one Gini point, so one fifth of a point is not trivial. Nevertheless, the magnitude of the coefficient on NDP government is smaller than those on the time and year dummies. This is consistent with the attenuation expected because within-province variation in distributional outcomes takes place within a federal system that both redistributes itself and produces correlation in the economic shocks facing its constituent units.

A clearer sense of the magnitude of the change in post-*fisc* inequality associated with the two power resources variables can be obtained by assuming a particular income distribution and then calculating changes in income associated with a given change in Gini coefficients. The log-normal density is often used to model income distributions; in addition, Gini coefficients can be calculated easily from the parameters of this distribution.[17] The mean post-*fisc* Gini coefficient for Canada as a whole is 30.0; decreasing this to 29.75 implies that the variance of the log-normal distribution changes from .771 to .764. Given an average income for economic families of about $60,000 CDN in 2003, this implies that the gap between the income at the 90th percentile and the 10th percentile of the distribution would shrink by approximately $1600 CDN if the post-*fisc* Gini were reduced by .25.

[17]In particular, the relationship is $G = 2\Phi(\sigma/2) - 1$.

Table 4: Robustness of estimated coefficients, post-tax-and-transfer inequality

| | Dependent variable: Post-*fisc* inequality$_t$ | | | | | | |
|---|---|---|---|---|---|---|---|
| | posttax$_{t-1}$ | pretax$_t$ | union | unemp | ndp | youthper | singper |
| Ex Newfoundland | 0.004 | 0.713 | -0.040 | -0.156 | -0.232 | -0.064 | -0.046 |
| Ex PEI | 0.016 | 0.689 | -0.048 | -0.230 | -0.241 | 0.059 | -0.051 |
| Ex Nova Scotia | 0.016 | 0.652 | -0.041 | -0.147 | -0.228 | 0.076 | -0.070 |
| Ex New Brunswick | 0.028 | 0.667 | -0.040 | -0.141 | -0.196 | 0.079 | -0.071 |
| Ex Quebec | 0.008 | 0.658 | -0.045 | -0.148 | -0.242 | 0.081 | -0.055 |
| Ex Ontario | 0.028 | 0.663 | -0.038 | -0.137 | -0.193 | 0.075 | -0.078 |
| Ex Manitoba | 0.006 | 0.665 | -0.035 | -0.156 | -0.414 | 0.080 | -0.068 |
| Ex Saskatchewan | 0.014 | 0.631 | -0.027 | -0.139 | -0.107 | 0.079 | -0.016 |
| Ex Alberta | 0.020 | 0.649 | -0.030 | -0.145 | -0.168 | 0.047 | -0.078 |
| Ex British Columbia | 0.025 | 0.650 | -0.034 | -0.163 | -0.177 | 0.061 | -0.087 |

Notes: All models estimated using OLS regression. Standard errors not shown. Each model includes a full set of province and time dummies (not shown).

To evaluate the robustness of these results, I re-estimate model (2) from Table 3 several times, each time dropping data from one province. The assumption underpinning the statistical model, that the coefficients on explanatory variables be constant across provinces, may be violated if dropping a province causes the results to change dramatically. The coefficients estimated for each of these models are presented in Table 4. This table reveals the effects that a single province can have, even though the fixed characteristics of the provinces are removed from the model. The estimated effects of unemployment, union density, and percentage of lone-parent families are fairly stable, retaining negative and substantively similar coefficients.[18] The estimate for youth population is less stable: the sign flips on youth population when Newfoundland is removed. The consequences of dropping provinces for the NDP variable are somewhere in between; the estimated coefficient is always negative, but its absolute magnitude is greater when Manitoba is dropped and smaller when Saskatchewan is dropped.

The effect of dropping provinces on the estimated effect of the NDP bears further ex-

[18]The absence of a change in union density when Newfoundland is reassuring, given the issue mentioned earlier about the 1996 series break.

amination, given the centrality of this variable to the power resources model. In one sense, the variability of the estimates is unsurprising; unlike all of the other variables, information about within-province variation due to the NDP comes from only four provinces. The natural intuition from these results suggests that NDP governments were more effective in reducing in equality in Saskatchewan and less effective in Manitoba, with British Columbia and Ontario falling in the middle. While it is always dangerous to make judgments about multivariate relationships from bivariate data, this conclusion seems to be borne out by the data. Figure 2 plots the difference between market and post-*fisc* inequality in the four provinces where the NDP governed at some point between 1980 and 2003; periods of NDP government are shaded. The effect of the NDP in Ontario is quite striking, and reflects two different processes. Market inequality grew during the early 1990's, but post-tax-and-transfer inequality did not increase significantly, thus causing a widening of the gap between the two measures. When the Conservatives replaced the NDP government in the middle of 1995, post-tax-and-transfer inequality also increased, thus narrowing the gap again at higher levels of inequality. This post-1995 increase is consistent with policy changes made by the Harris government, such as the 21% cut in social assistance mentioned earlier (Osberg and Xu 1999). Turning to Manitoba and Saskatchewan, one observes that the gap is quite small in Manitoba during the NDP governments of the early 1980s, while it is large during the tenure of the NDP in Saskatchewan in the 1990s. The variability in the estimated coefficients is likely due to this difference, even after controlling for time effects.[19] The degree to which this is a problem depends on whether one believes that the data generating process is the same across provinces. If so, the variability provides greater identification of the effects of NDP governments; if not, one might be concerned that Saskatchewan is driving the result. I tend toward the former, largely due to the stability that results when the other two provinces with NDP governments are removed.

[19]With 1981 as the reference year, the estimated coefficients on the time fixed effects are negative throughout the 1990s.

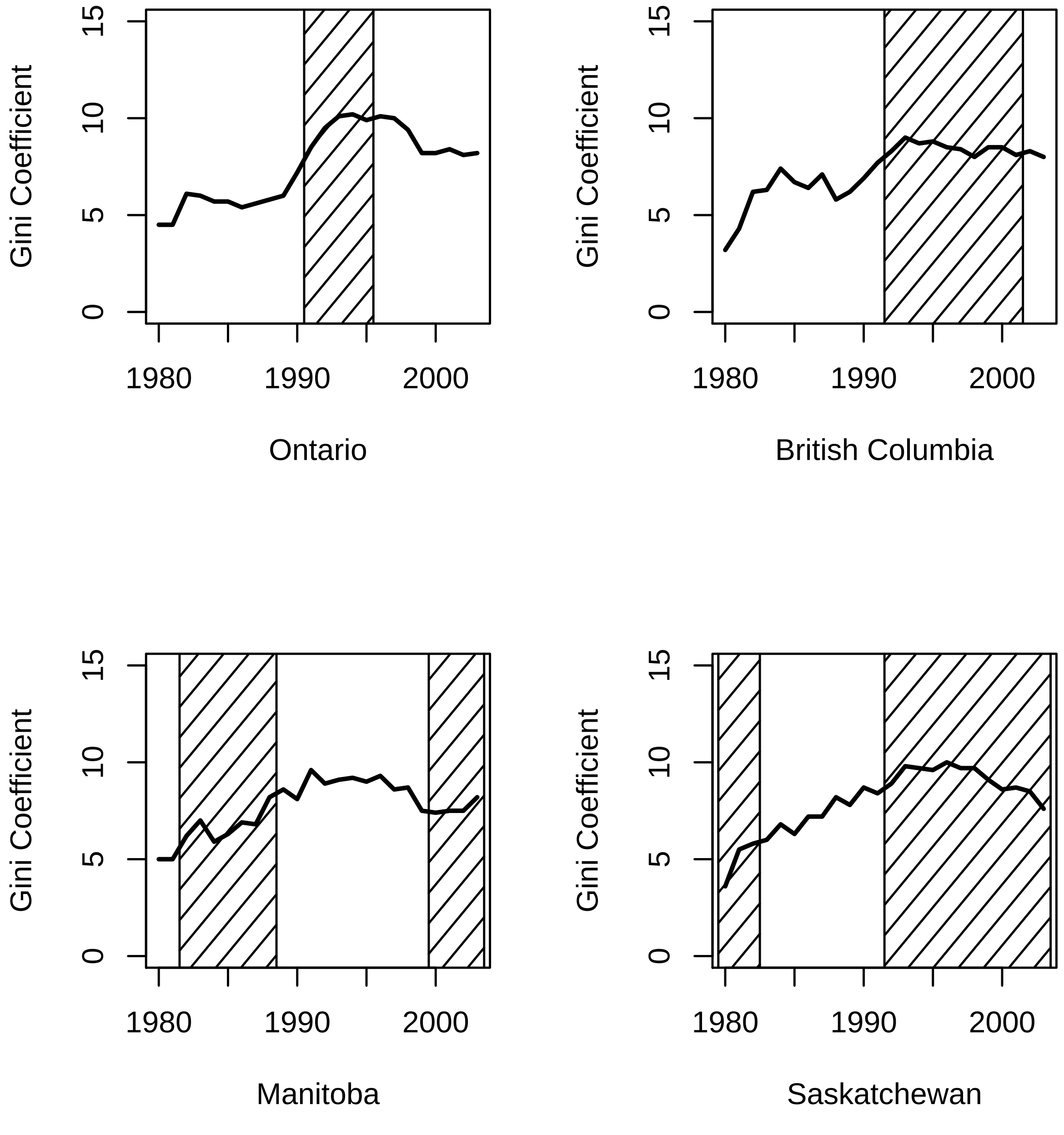


Figure 2: Market Gini minus post-*fisc* Gini for four provinces

# 7 Discussion

This paper provides evidence that NDP provincial governments and higher levels of union membership are associated with lower income inequality after taxes and transfers, conditional on market inequality. It finds no statistically reliable evidence that these variables affect the degree of market inequality. The substantive effects of left party government and union density are admittedly small, but they are not trivial. The effects are estimated using variation within provinces over time, and inequality does not change dramatically within provinces from year to year. To strengthen the results of this study, it would be advisable to extend the data both forward and backward in time to determine whether the relationships identified hold in other periods, particularly during the period of welfare state expansion in Canada during 1960s and 1970s. More effort is needed on the determinants of market inequality, which is likely to be related to education and to the composition of the economy. I do not expect that either of these extensions would change the substantive conclusions of this paper.

The results presented in this paper are, on balance, quite supportive of the power resources approach. Both of the variables have the relationship with post-*fisc* inequality predicted by the theory. One advantage of the Canadian data is that union density and left party government are not nearly as correlated as they are across the OECD, allowing both to be entered in the same statistical model. A greater advantage, however, is that this paper constitutes a new test of the theory: Korpi (1983), Stephens (1979), and their many followers did not develop the power resources model with the Canadian provinces in mind. While the magnitudes of these effects are small, attenuation is to be expected because of the many external determinants of distributional outcomes in subnational political units. This paper contributes most significantly to the literature by broadening its horizons beyond the standard OECD data.

In demonstrating that it is possible to test theories about inequality and redistribution using variation within countries, this study suggests an avenue for expanded research. It may only be possible to study some variables hypothesized to affect inequality in a national context; it is unlikely, for example, that labor bargaining will occur at the firm level in one province and at the provincial level in another. When sufficient variation exists, however, the implications of various theories should be tested within as well as between countries.

# Data appendix

This appendix provides references to the data used in the paper. For data from CANSIM II, the series numbers are in the following order: Canada, Newfoundland, Prince Edward Island, Nova Scotia, New Brunswick, Quebec, Ontario, Manitoba, Saskatchewan, Alberta, and British Columbia.

**Pre and post-tax-and-transfer inequality**: This data measures the Gini coefficient for economic families headed by individuals under the age of 65. Economic families, in the Canadian data, consist of households with two or more individuals. Data from CANSIM II table 2020705. Market income Gini series: V21151590 , V21151806, V21151914, V21152022, V21152130, V21152238, V21152346, V21152562, V21152670, V21152778, V21152886. After-tax Gini series: V21151662, V21151878, V21151986, V21152094, V21152202, V21152310, V21152418, V21152634, V21152742, V21152850, V21152958.

**Union density**: Union density measures the share of employees who belong to a trade union. Data on union density from 1976 to 1995 is calculated from data collected under the Corporations and Labor Unions Returns Act and the Labor Force Survey, CANSIM II table 2790025, series: V810365, V810368, V810371, V810374, V810377, V810380, V810383, V810386, V810389, V810392, V810395. Following McDonald and Myatt (2004), data for the period 1997-2003 is obtained by dividing the number of employees covered by union agreements by total employees. Data from CANSIM II table 2820073. The series for employees with union coverage are: V3075071, V3075076, V3075081, V3075086, V3075091, V3075096, V3075101, V3075106, V3075111, V3075116, V3075121. The series for total employees are: V3075016, V3075021, V3075026, V3075031, V3075036, V3075041, V3075046, V3075051, V3075056, V3075061, V3075066. The missing year (1996) is filled in by taking the mean of 1995 and 1997.

**Party of government**: This set of dummy variables codes the party that held the premiership of the province on January 1 of the year in question. Data from Canadian Parliamentary Guide, various years.

**Unemployment rate**: This variable measures the annual unemployment rate for both sexes aged 15 and over. The source of the data is the Labor Force Survey. Data from CANSIM II table 2820002. Data series: V2461224, V2461854, V2462484, V2463114, V2463744, V2464374, V2465004, V2465634, V2466264, V2466894, V2467524.

**Female labor force participation**: This variable measures percentage of women aged 15 to 64 who are in the labor force. The source of the data is the Labor Force Survey. Data from CANSIM II table 2820002. Data series: V2461672, V2462302, V2462932, V2463562, V2464192, V2464822, V2465452, V2466082, V2466712, V2467342, V2467972.

**GDP per capita**: This variable measures average real income per person for each province, calculated by dividing GDP by population. Provincial GDP is measured in constant 1997 dollars; the data comes from Provincial Economic Accounts. Data from CANSIM

II table 3840002, series: V3839887, V3839933, V3839979, V3840025, V3840071, V3840117, V3840163, V3840209, V3840255, V3840301, V3840347. Population data from Estimates of Population by Age and Sex for Canada, the Provinces and Territories, CANSIM II table 510001, series: V466668, V466983, V467298, V467613, V467928, V468243, V468558, V468873, V469188, V469503, V469818.

**Net migration**: This variable measures net interprovincial migration as a percentage of population; positive numbers indicate a net inflow of individuals. Calculated by dividing total net migrants by the population (see above). Data on number of net migrants from CANSIM table 510012, series: V445531, V446476, V447421, V448366, V449311, V450256, V451201, V452146, V453091, V454036, V454981.

**Youth**: This variable measures the population aged 15 and under as a percentage of the total population. Calculated by dividing the population, both sexes, 0-15 years by the population (see above). Data from CANSIM table 510001, series: V466959, V467274, V467589, V467904, V468219, V468534, V468849, V469164, V469479, V469794, V470109.

**Lone-parent families**: This variable measures economic families headed by a lone parent as a percentage of all non-elderly economic families. Calculated by dividing the number of lone-parent economic families by the total numbe rof non-elderly economic families. Data for 2002-2003 extrapolated from available data. Lone-parent family data from CANSIM table 2020901, series: V21158865, V21159009, V21159081, V21159153, V21159225, V21159297, V21159369, V21159513, V21159585, V21159657, V21159729. Total non-elderly family data taken from CANSIM table 2020901, series: V21158854, V21158998, V21159070, V21159142, V21159214, V21159286, V21159358, V21159502, V21159574, V21159646, V21159718.

# Bibliography

Atkinson, Anthony B. and Andrea Brandolini. 2001. “Promise and pitfalls in the use of ‘secondary’ data-sets: Income inequality in OECD countries as a case study,” *Journal of Economic Literature* 39(3): 771-799.

Beck, Nathaniel and Jonathan N. Katz. 1995. “What do do (and not to do) with time-series cross-section data,” *American Political Science Review* 89 (3): 634-647.

Bradley, David, Evelyne Huber, Stephanie Moller, Francois Nielsen, and John D. Stephens. 2003. “Distribution and redistribution in postindustrial democracies,” *World Politics* 55: 193-228.

Countryman, Galen J. 1999. “The effect of unemployment insurance benefits on income inequality in the Canadian provinces,” *Canadian Public Policy* 25 (4): 539-556.

Cross, William and Lisa Young. 2004. “The contours of political party membership in Canada,” *Party Politics* 10 (4): 427-444.

Goodrich, Ben. 2004. “Problems with and solutions for two-dimensional models of continuous dependent variables,” Manuscript, Harvard University.

Gray, David, Jeffrey A. Mills, and Souruche Zandvakili. 2003. “An analysis of differential provincial income inequality trends in Canada.” Working paper, University of Cinncinnati.

Heisz, Andrew, Andrew Jackson, and Garnett Picot. 2001. “Distributional outcomes in Canada in the 1990s,” *Review of Economic and Social Progress* 2001: 247-272.

Huber, Evelyne and John D. Stephens. 2001. *Development and Crisis of the Welfare State: Parties and Politics in Global Markets.* Chicago: University of Chicago Press.

Huber, Evelyne and John D. Stephens. 1993. “Political parties and public pensions: A quantitative analysis,” *Acta Sociologica* 36: 309-325.

Kesselman, Jonathan R. and Ron Cherung. 2004. “Tax incidence, progressivity, and inequality in Canada.” Working paper, Simon Frasier University.

Korpi, Walter. 1998. “Power resources approach vs action and conflict: On causal and intentional explanations in the study of power.” In *Power Resources Theory and the Welfare State: A Critical Approach,* edited by Julia S. O’Connor and Gregg M. Olsen. Toronto: University of Toronto Press.

Korpi, Walter. 1983. *The Democratic Class Struggle.* London: Routledge & Kegan Paul.

Kwon, Hyeok Yong and Jonas Pontusson. 2003. ”Welfare spending in OECD countries revisited: Has the salience of government partisanship really declined?” Paper presented at the 2003 MPSA conference, Chicago, Ill.

MacPhail, Fiona. 2000. “What caused earnings inequality to increase in Canada during the 1980’s,” *Cambridge Journal of Economics* 24: 153-175.

McDonald, James Ted and Anthony E. Myatt. 2004. “The minimum wage effect on youth employment in Canada: Testing the robustness of cross-province panel studies.” Working paper, University of New Brunswick.

Moller, Stephanie, David Bradley, Evelyne Huber, Francois Nielsen, and John D. Stephens. “Determinants of relative poverty in advanced capitalist democracies,” *American Sociological Review* 68 (1): 22-51.

Mulé, Rosa. 2001. *Political Parties, Games, and Redistribution.* Cambridge, UK: Cambridge University Press.

O’Connor, Julia S. 1998. “Welfare expenditure and policy orientation in Canada, 1960-1990, in comparative perspective.” In *Power Resources Theory and the Welfare State: A Critical Approach,* edited by Julia S. O’Connor and Gregg M. Olsen. Toronto: University of Toronto Press.

Olsen, Gregg M. 1998. “Locating the Canadian welfare state: Family policy and health care in Canada, Sweden, and the United States.” In *Power Resources Theory and the Welfare State: A Critical Approach,* edited by Julia S. O’Connor and Gregg M. Olsen. Toronto: University of Toronto Press.

Olsen, Gregg M. 1992. *The Struggle for Economic Democracy in Sweden.* Aldershot, UK: Avebury/Gower.

Osberg, Lars and Kuan Xu. 1999. “Poverty intensity: How well do the Canadian provinces compare?” *Canadian Public Policy Journal* 25: 179-195.

Petry, Francois, Louis Imbeau, Jean Crete, and Michel Clavet. 2000. “Explaining the evolution of government size in the Canadian provinces,” *Public Finance Review* 28(1): 26-47.

Robinson, Ian. 1994. “NAFTA, social unionism, and labor movement power in Canada and the United States,” *Relations industrielles/Industrial Relations* 49(4): 657-695.

Sharpe, Andrew and Myles Zyblock. 1997. “Macroeconomic performance and income distribution in Canada,” *North American Journal of Economics and Finance* 8(2): 167-199.

Statistics Canada. 2005. “CANSIM II (Canadian Socio-Economic Information Management System) [electronic database],” accessed through http://dc1.chass.utoronto.ca.ezp1.harvard.edu/cansim2/English/index.html.

Statistics Canada. Income Statistics Division. 2003. *Income Trends in Canada 1980 to 2001 - User’s Guide.* Ottawa: Ministry of Industry.

Stephens, John D. 1979. *The Transition from Socialism to Capitalism.* Urbana, Ill.: University of Illinois Press.

Wang, Baotai and Tomson Ogwang. 2004. “Is the size distribution of income in Canada a random walk?” *Economics Bulletin* 3(29): 1-9.