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The double role of ethnic heterogeneity in explaining welfare-state generosity

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Abstract

Based on theoretical models of budget-balanced social insurance and individual choice, we argue that in addition to the well-known *empathy mechanism* whereby ethnic heterogeneity undermines sentiments of solidarity among a citizenry to reduce welfare generosity, population heterogeneity affects the generosity of a polity’s social insurance programs through another distinct mechanism, *political conflict*. Ethnic heterogeneity likely intensifies political conflict and reduces welfare generosity because heterogeneity of unemployment risk makes it more difficult to achieve social consensus concerning tax-benefit programs. Utilizing two separate regression analyses covering highly diverse polities, the 50 U.S. states and District of Columbia (CPS data), and 13 OECD countries (LIS data), we find strong evidence that empirically distinct empathy and political conflict effects on unemployment insurance programs characterize contemporary politics. Our findings suggest existing analyses of the negative relationship between ethnic heterogeneity and the size of the welfare state likely over- or underestimate the empathy effect. For example, perhaps surprisingly, had our analysis of US data omitted a measure of unemployment dispersion, the negative effect of ethnic fractionalization would have been *underestimated.*

Key Words: political economy, welfare state, social insurance, ethnic fractionalization
JEL Classification: H53.

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1. Introduction

It is often held that the cause of the large welfare states in the Nordic countries is their homogeneous populations. The modern version of this view has been advanced by Alesina, Glaeser, and Sacerdote (2001) and Alesina and Glaeser (2004). The mechanism through which homogeneity is thought to work is through promoting solidaristic feelings among the population. If a society is culturally and ethnically homogeneous, then its members can empathize with each other, and so the genesis of generous welfare states is attributed to a kind of altruism that does not exist between groups who share little in culture, religion, or language.

There is, however, another mechanism through which homogeneity may induce generous welfare states, and this mechanism does not invoke empathy or altruism. If a society is culturally homogeneous, its members are likely to face similar risks – more similar, at any rate, than members of heterogeneous societies, if that heterogeneity is correlated with educational and life-style distributions. If all members of a society face the same risks – the same probability distribution of bad states occurring – then, assuming each member is risk averse, all will possess the same ideal budget-balanced insurance policy. This is so even if individuals possess different von Neumann – Morgenstern utility functions over income lotteries, as long as each is risk averse. It follows that it should be much easier to come to political agreement on what the optimal insurance policy is when the society is homogeneous, and this may produce larger welfare states. We repeat this mechanism does not rely on the empathy that homogeneity may induce, but rather on the consequences of homogeneity for the choice of ideal self-interested insurance policies.

Consider the following simple model to illustrate this point. Each member of a society faces the same probability of unemployment. A worker i is characterized by her wage w^i if employed, and her von Neumann – Morgenstern utility function u^i over money lotteries, which we assume is strictly concave. The distribution of wages is F with mean μ. Every worker faces the same probability p of being unable to earn income, in which case she collects a benefit that is a fraction of her wage, λw. Employed workers are taxed at a rate t on income. A policy (t, λ) is budget-balanced if:

$$t(1-p)\mu = \lambda p \mu \text{ or } t = \frac{\lambda p}{(1-p)}. \qquad (1.1)$$

Thus, worker i's ideal budget-balanced policy is given by:

$$\max_{(t,\lambda)} (1-p)u^i\left((1-t)w^i\right) + pu^i(\lambda w^i) \qquad (1.2)$$

$$\text{s.t. } t = \frac{\lambda p}{(1-p)}$$

Letting $r = \frac{p}{1-p}$, this is equivalent to solving:

$$\max_{\lambda} u^i((1-\lambda r)w^i) + ru^i(\lambda w^i)$$

whose first-order condition reduces to $u^{i'}\left((1-\lambda r)w^i\right) = u^{i'}(\lambda w^i)$. Strict concavity of the utility function implies $(1-\lambda r)w^i = \lambda w^i$, and hence $\lambda = \frac{1}{1+r} = 1-p$ and $t = p$. Thus every worker, regardless of her (risk averse) preferences or wage, possesses the same ideal policy, which is complete income-smoothing; she enjoys an income of $(1-p)w^i$ whether working or unemployed. If, however, workers face different risks, they will possess different ideal policies. No simple budget-balanced policy will satisfy everyone. It will be more difficult to reach a political agreement on social insurance.

Our conjecture is that there are two mechanisms operating in the genesis of social insurance, in both of which population homogeneity is important. The first we call the *empathy mechanism*. The second mechanism we call the *political-conflict mechanism*. The less homogeneous a polity, the more differentiated are the risks that its members face, causing political conflict over the level of social insurance, and hence less expansive social insurance.

We must justify the last clause in this claim. Why should the fact that polities face highly differentiated risks induce less expansive social insurance than otherwise? We suggest several reasons. The first is that the population type that suffers the smallest risks, which will tend to comprise its higher paid and more educated members, may have influence in the political process in excess of its numbers. Clearly, this would not happen with the median-voter model of politics, if the lower paid and less educated worker is the median worker. We are therefore assuming a political mechanism more like the one that Bartels (2008) and Gilens (2009) indicate exists in the United States, where richer voters have far more influence on policy than poorer ones.

A second reason may be that the actual political mechanism entails bargaining, rather than median-voter hegemony. Different population types are represented in political bargaining, and they reach a compromise. If the bargaining process is inefficient, then the social-insurance policy may be less comprehensive than any of the types would like. However, even if bargaining is efficient, then it will still represent a compromise between the low-tax-benefit policy that the type with the low probability of bad states would like and the high-tax-benefit policy that the type with high probability of bad states would like.

Consider the following simple Nash bargaining model. There are two worker types in the population, A and B, where type A constitutes fraction f of the population. A types have an average annual wage of w^A if employed, and face an unemployment probability p^A of being unemployed during the year in question. Type B workers earn an annual wage of w^B and face a probability p^B of unemployment. All workers have a strictly concave von Neumann-Morgenstern utility function u on money lotteries. Given these facts, the set of feasible tax-benefit policies (t,λ) of the type described earlier must satisfy the budget constraint:

$$\lambda \le t(K-1), \qquad (1.3)$$

where $K = \frac{fw^A + (1-f)w^B}{p^A f w^A + p^B (1-f) w^B}$.

The ideal policy for type $J = A$,B maximizes $p^J u(\lambda w^J) + (1-p^J)u((1-t)w^J)$ subject to (1.3). Let $u(0) = 0$. The Nash bargaining problem between these two groups of workers is to choose the policy (λ, t) that maximizes the Nash product

$$\begin{gathered}(p^A u(\lambda w^A) + (1-p^A)u((1-t)w^A) - (1-p^A)u(w^A)) \cdot \\ (p^B u(\lambda w^B) + (1-p^B)u((1-t)w^B) - (1-p^B)u(w^B))\end{gathered}$$

subject to (1.3). The impasse point in the bargaining problem is the expected utility of a worker, were there no insurance. Consider an example parameterized by the data $(f, w^A, w^B, p^A, p^B) = (0.4, 20, 60, p^A, p^B)$, let $u(x) = \ln x$, and suppose that an unemployed worker in the absence of insurance receives an income of one from some source, so that her utility is zero in this case. In the homogeneous situation, when $p^A = p^B = 0.1$, the ideal policy for both types is $t = 0.1,\ \lambda = 0.9$, and it is, of course, the solution to the Nash bargaining problem. Now suppose there is a mean-preserving spread of the probabilities of unemployment, so that $p^A = 0.2,\ p^B = 0.033$. Then the solution to the bargaining problem is $t = 0.054, \lambda = 0.867$: the social insurance is less generous than in the original homogeneous situation[1]. The tax rate is higher than type B would like, and lower than type A would like.

[1] A mean-preserving spread of the probabilities does not always decrease the tax rate that solves the Nash bargaining problem, but this appears to be so for changes in the probabilities that are not too small. The general characterization of when this occurs is complicated, and is not worth pursuing here.

Yet a third reason for the relatively small welfare states in heterogeneous societies is based upon the multi-dimensionality of politics. In heterogeneous societies, political competition over economic issues, here taken to be social insurance, is complicated by a second non-economic issue that often has to do with race, ethnicity, or religion. Roemer (1998) and Roemer, Lee and Van der Straeten (2007) show that when this second issue exists, the political equilibrium is likely to deliver a smaller welfare state than it would in the absence of the second issue. The last reference cited estimates that the income tax rate in the U.S. is as much as ten percentage points lower than it would be, were race not an important issue for the American polity. This reduction in the tax rate can be decomposed into an 'anti-solidarity effect' and a 'policy bundle effect'. The former is simply the absence of empathy and inter-ethnic altruism in an ethnically heterogeneous society; the latter has to do with the bundling together of policies. These authors estimate that about half of the reduction in tax rates referred to in the U.S. is due to the policy bundle effect. Intuitively, think of this as being manifest in the relatively large vote that the Republican Party receives from whites, because they prefer that party's policy on race to the Democrats' policy, even though the Republicans advocate a low-redistribution policy that is not in the economic interests of many of these white voters. Thus, the existence of the race issue in U.S. politics enables the Republican Party to achieve an approximately 50% vote share, while not moving their economic policy to the left, as competition would force them to do, were the race issue not so salient for voters.

Our reasoning suggests that we attempt to explain the differential size of welfare states not simply with an independent variable measuring ethnic heterogeneity, as Alesina and Glaeser do, but with a second independent variable as well, one that measures the degree of heterogeneity of income risks that a population faces. With these two variables in a regression, where the dependent variable is a measure of the generosity of social insurance, the risk-heterogeneity variable should pick up the political-conflict effect and the ethnic heterogeneity variable should pick up the empathy effect. We expect the coefficients on both variables to be negative.

The next section of the article generalizes our earlier simple model where workers were either employed or unemployed to a model allowing workers to experience spells of unemployment of varying length. We then illustrate that theoretical results from the latter model's more realistic portrayal of individual worker's unemployment risk remain consistent with our ideas and major hypotheses concerning the provision of social unemployment insurance. Afterwards, we subject the testable predictions of the theoretical models to two econometric analyses using data from the United States and several OECD nations, respectively. To explain varying generosity of jurisdictions' unemployment insurance benefits, we model the empathy and political conflict mechanisms using racial heterogeneity across the states (immigrant proportion of population in OECD nations), and the coefficient of variation of weeks unemployed across jurisdictions.

The econometric tests provide strong support for our theoretical predictions. We estimate robust statistically significant negative effects of ethnic/racial heterogeneity and risk heterogeneity on unemployment benefits across the states in the U.S. and across OECD nations. The main body of the paper concludes with a brief summary of our results. An appendix provides robustness tests for the U.S. regressions.

2. Risky length of unemployment spells

The measure of risk heterogeneity we use in the empirical analyses is the coefficient of variation of *weeks unemployed* experienced by a population. Therefore, in this section, we generalize the simple model presented in section 1 that assumed a worker was employed for the year or unemployed for the year to cover the desired degree of insurance as a function of the probability distribution of individual-unemployment-spell lengths.

Therefore, as before, let w^A and w^B be the full-employment wages for a worker in the respective groups, and let f^i be the group population proportions so $f^A w^A + f^B w^B = \mu$ is mean income at full employment. We suppose there is a discrete grid of unemployment-spell lengths, y^t, that varies from $y^1 = 0$ (full employment for the individual) to $y^J = 1$ (unemployed for the whole year). Generally, y is a fraction of the year. Let $q^y(w)$ be the probability that a worker of wage w will sustain an unemployment spell y during the year. By definition, $\sum_{y^1=0}^{y^J=1} q^y(w) = 1$, for every w.

Let μ_1 be the average annual market income per capita of employed workers and let μ_0 be the average per capita income loss suffered by unemployed workers. By definition, $\mu = \mu_1 + \mu_0$. Note that $\frac{\mu_0}{\mu}$ is the fractional loss of GNP due to unemployment.

A *simple policy* is an ordered pair (λ, t) such that an unemployed worker of wage w receives the benefit $\lambda w y$ for a period y of unemployment, and a worker who is employed for fraction $1 - y$ of the year pays a tax of $tw(1-y)$. It follows that the balanced-budget equation for the society characterized by the probability distributions $\{q^y(w)\}$ is:

$$t\mu_1 = \lambda\mu_0 \,. \tag{2.1}$$

Explicitly we may write for $i = A, B$:

$$\mu_1 = \sum_i \sum_{y^1=0}^{y^J=1} q^y(w^i)(1-y) f^i w^i \tag{2.2a}$$

$$\mu_0 = \sum_i \sum_{y^1=0}^{y^J=1} q^y(w^i) w^i f^i y \tag{2.2b}$$

Let u be the concave von Neumann- Morgenstern utility function of an individual in the population over money lotteries. The ideal policy for an individual of capacity w is the solution to:

$$\begin{aligned}&\max_{\lambda,t}\sum_{y^1=0}^{y^J=1} q^y(w)\big(u(\lambda wy+(1-t)w(1-y)\big)\\&\text{subj. to}\\&t\mu_1=\lambda\mu_0\end{aligned} \quad . \tag{2.3}$$

Here, w is fixed, and the probability distribution is over unemployment-spell lengths, y.

Substituting $t=\frac{\lambda\mu_0}{\mu_1}$ into the objective function, we compute that the ideal 'replacement ratio' λ for the individual w – assuming that this ratio is interior in the interval $[0,\frac{\mu_1}{\mu_0}]$, is given by the first-order condition:

$$\sum_{y^1=0}^{y^J=1} q^y(w)u'(I(w,y))(wy-\frac{\mu_0}{\mu_1}w(1-y)))=0, \tag{2.4}$$

where $I(w,y)\equiv\lambda wy+(1-t)w(1-y)$ is the individual's annual income if he is unemployed for y fraction of the year.

To derive a relatively simple formula, we will at this point assume that

$$u(x)=\ln(x). \tag{2.5}$$

Then (2.4) can be written:

$$\sum_{y^1=0}^{y^J=1} q^y(w)\frac{y-\frac{\mu_0}{\mu_1}(1-y)}{\lambda y+\left(1-\lambda\frac{\mu_0}{\mu_1}\right)(1-y)}=0.$$

Letting λ^w denote the ideal replacement rate for agent w, this in turn can be written as:

$$\sum_{y^1=0}^{y^J=1}\frac{q^y(w)}{\varphi(y)+\lambda^w}=0, \tag{2.6}$$

where $\varphi(y) = \frac{1-y}{y-\frac{\mu_0}{\mu_1}(1-y)} = \frac{\mu_1(1-y)}{\mu_1 y-\mu_0(1-y)} = \frac{\mu_1(1-y)}{\mu y-\mu_0}$. The ideal replacement rate for agent *w*, assuming it is interior, is the value of λ^w satisfying equation (2.6).

Compute that $\varphi'(y) < 0$, so the function $\varphi(\cdot)$ is decreasing in the regions in which it is defined. Note that φ approaches an asymptote as $y \to \frac{\mu_0}{\mu}$, at which point it is undefined. Note that $\varphi(0) = -\frac{\mu_1}{\mu_0}$ and $\varphi(1) = 0$. It follows from these observations that φ has the shape illustrated in figure 1. Since the upper limit of λ is $\frac{\mu_1}{\mu_0}$, achieved when $t = 1$, it follows that $\varphi(0) + \lambda^w < 0$. Hence, from figure 1, it immediately follows that $\varphi(y) + \lambda^w < 0$ for $y < \frac{\mu_0}{\mu}$ and obviously $\varphi(y) + \lambda^w > 0$ for $> \frac{\mu_0}{\mu}$.

[figure 1 here]

We can therefore decompose (2.6) as:

$$\sum_{y<\frac{\mu_0}{\mu_1}} \frac{q^y(w)}{\varphi(y)+\lambda^w} + \sum_{y>\frac{\mu_0}{\mu_1}} \frac{q^y(w)}{\varphi(y)+\lambda^w} = 0 \tag{2.7}$$

where all the terms in the first sum are negative and all the terms in the second sum are positive.

As in our simpler model, in the homogeneous situation when $w^A = w^B$ or unemployment risk profiles are identical, the ideal policy for both types is identical. Now analogous to the mean-preserving spread of the earlier model where type *A*'s risk of unemployment became unambiguously greater, suppose there is a change in type A's wage and risk profile, so that all the fractions $q^y(w^A)$ for $y > \frac{\mu_0}{\mu}$ increase and all the fractions $q^y(w^A)$ for $y < \frac{\mu_0}{\mu}$ decrease. For any fixed value λ, all the terms in the first sum in (2.7) are greater for w^Athan for w^B, and all the terms in the second sum are similarly greater for w^A than for w^B. It follows that it must be the case that $\lambda^{w^A} > \lambda^{w^B}$ in order for (2.7) to hold for both groups.

We have shown the following:

Proposition: *Let* $u(w) = \ln(w)$ *for all w. Suppose that A and B are two groups such that, for*:

$$y < \frac{\mu_0}{\mu}, q^y(w^A) < q^y(w^B) \text{ and } y > \frac{\mu_0}{\mu}, q^y(w^A) > q^y(w^B)$$

Then $\lambda^{w^A} > \lambda^{w^B}$.

It is this proposition that supports our claim that the larger is the dispersion in unemployment spells in a population, the larger will be the dispersion in desired replacement ratios, and assuming that ethnic heterogeneity makes such dispersion more likely, political conflict over tax-benefit policy will intensify. Of course, the exact claim made by the proposition is not that general, which is why it is necessary to make the calculation leading to the proposition.

3. Econometric Tests

This section features econometric testing of our hypothesis that population heterogeneity affects welfare state generosity through distinct empathy and political-conflict mechanisms. Controlling for other individual as well as jurisdictional characteristics, we regress measures of individuals' annual unemployment benefits on measures of the population heterogeneity and dispersion of unemployment risk within the political jurisdiction of the individual's place of residence. According to the theory (other relevant variables constant), greater levels of ethnic heterogeneity within a polity should reduce received unemployment benefits because of decreased population empathy and increased political conflict (due to greater dispersion of the risk of unemployment). We report two separate analyses of the effects of heterogeneity on the delivery of unemployment benefits. The first analysis examines heterogeneity and unemployment benefits using data from 51 (state and District of Columbia) political jurisdictions composing the United States. The second examines our hypotheses using data from 13 OECD countries available in the Luxembourg Income Study database (2014) (LIS), spanning the period 1984 to 2010. We first describe the U.S. data and samples, and then proceed to describe the data drawn from LIS.

3.1. U.S. data

Its governance system of federalism in which state governments with diverse ethnic populations, socioeconomic characteristics, and political orientations deliver publicly provided goods and services within parameters defined by a common federal government, makes the United States a virtual laboratory for empirical tests of hypotheses similar to ours. U.S. fiscal federalism's delivery of unemployment benefits is especially suitable for such an analysis. The U.S. Federal-State Unemployment Insurance Program delivers unemployment benefits to eligible workers under the auspices of the federal government. However, each of the local governments (the fifty states, District of Columbia, Puerto Rico, Guam, and the Virgin Islands) administers its own program subject to guidelines laid out in Federal law.

The program's main features are straightforward. Constrained by parameters set by federal law, the state where eligibility for benefits is established determines eligibility requirements, weekly benefit amounts (WBA), and the duration of benefit payments. To establish eligibility a claimant must apply, demonstrate a recent strong attachment to the labor force by satisfying state requirements for wages earned or time worked during a "base period", be determined unemployed through no fault of her own, and be ready, willing, and able to work. The base period is generally the first four of the five calendar quarters immediately preceding filing of a claim; no-fault eligibility rules out workers fired for cause or who quit; readiness requirements render sick or injured workers ineligible for unemployment benefits, although they may seek disability benefits.[2]

Variations in the method by which the WBA is determined differentiate benefits paid across the states. A large majority of states compute weekly benefits as a fraction (usually about one-half) of a claimant's usual weekly wages in one or more quarters of her base period. Typically, the WBA is a fraction of wages during the base period quarter with highest earnings; however, some states average two quarters. Two additional important sources of variation in the WBA paid by states are state determined maximum and minimum amounts. During 2006, among the fifty states and the District of Columbia, the maximum varied from a low of $235 per week in Mississippi to a high of $979 in Massachusetts, one of a minority of states supplementing a claimant's basic benefit with an allowance for dependents. The minimum WBA varies from a low of $5 in Hawaii to a high of $143 in Washington State. A final source of variation across jurisdictions is a state's limit on the duration of benefit payments. Although the limits varied from 20 weeks in South Carolina and Missouri to 30 weeks in Massachusetts, the vast majority of states allowed a maximum duration of 26 weeks.[3]

Our data set for the U.S. is the March 2007 CPS, which provides relevant data for the year 2006. Our model regresses the natural logarithm of annual benefits received (conditional on

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[2] Taxes on employers primarily fund the benefits in all but three states who also tax employees. Therefore, think of the benefit tax t as worker's common belief concerning the portion of the employer tax shifted to workers.

[3] In the 1970s, a permanent federal-state program of Extended Benefits was established for workers who exhaust their entitlement to regular state benefits during periods of high unemployment. The program is financed equally from federal and state funds.

receiving benefits) on ethnic fractionalization, the coefficient of variation of "covered" weeks unemployed, and a set of control variables described below. In addition to conditioning on the receipt of benefits, we apply data filters designed to replicate state program requirements to eliminate records that are inconsistent with eligibility requirements during the relevant year (e.g., individuals reporting positive benefits but zero weeks of unemployment, or are inconsistent with program benefit limitations such as maximums and minimums). The latter filters have the desired effect of eliminating obvious data reporting errors (e.g. six figure benefits, $1 of annual benefits) as well as reported benefits exceeding or falling below state specifications.

The filters replicate two policy instruments: 1. state limitations on the duration of benefit payments (covered weeks of unemployment); 2. state imposed maxima and minima weekly benefits.

A priori, we expect the recipient's weekly wage and her weeks unemployed to be the most powerful predictor of annual benefits. However, because many recipients reported more weeks of unemployment than the number for which a state would allow payment of benefits, using individuals' reported number of weeks unemployed would violate this important eligibility requirement, resulting in a serious miss-specification. We re-specify the relationship between annual benefits and weeks unemployed, by computing a new variable "covered weeks of unemployment" (covWksUn) equal to the minimum of each individual's reported weeks of unemployment and her state-of-residence's maximum duration of payment weeks.

Implementing policy instrument 2 could entail a straightforward requirement that no recipient's WBA lie outside the appropriate state's maximum and minimum WBA. Although we report the results from using this control during our discussion of regression robustness in the appendix, our primary model utilizes a different approach to placing bounds on a recipient's WBA. There are two primary reasons for this departure. First, complications derive from the fact that, although the great majority of recipient's state of residence (coded in the CPS) is also their state of benefit eligibility, this is not true for a substantial number of recipients. In addition to relocations across state boundaries during a spell of unemployment, many recipients' permanent employment is in jurisdictions other than their state of permanent residence. Thus, many large metropolitan centers (NYC, WDC, Philadelphia, Cincinnati, Boston, Jacksonville, Memphis) provide employment to large populations residing in multi-state regions.[4]

Secondly, the amount of a state's minimum WBA is not as obvious a signal of state generosity as might appear. For example, the smallest minimum WBA, Hawaii's $5, appears not to be a reflection of a lack of welfare generosity, but a liberal state's attempt to provide a benefit to even the minimal amount of work effort during a claimant's base period. Since Hawaii calculates a claimant's weekly benefit amount by dividing her high base quarter earnings by 21, a $5 WBA implies 13 weeks of earnings totaling just $105, approximately one hour of minimum wage earnings during a 13 week period. Payment of benefits this low would be extremely rare. For example, an eligible claimant working half-time (20 hours a week) at the federal minimum wage of $7.25 per hour would qualify for about a $90 WBA in Hawaii. For comparison, a minimum wage worker working one-half time in Mississippi would receive a WBA of $72.50, equivalent to 50 percent of her 13 week earnings.

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[4] Prospective claimants are advised "Generally, you should file your claim with the state where you worked. If you worked in a state other than the one where you now live or if you worked in multiple states, the state UI agency where you now live can provide information about how to file your claim with other states."

Our approach to these issues is to place state and individual specific lower and upper bounds on benefits received as follows. For the lower bound, we remove all records with annual benefits less than covered weeks of unemployment times the maximum of the state's minimum WBA or $72.50. For the upper bound, we remove all records with annual benefits above $716 times the state's maximum covered weeks of eligibility. The amount $716 is the maximum WBA allowed by the state of Maine, second to the $979 (including up to $326 for the beneficiary's dependents) allowed in Massachusetts. Because of Census Bureau limitations on the maximum weekly earnings reported (top coding), we considered any record with wage earnings above $2885 per week as an error and discarded it. Finally, we also eliminate individuals with weekly wages below $145 (20 hours at federal minimum wage) and those below age 18 or above age 65. Applying these filters to the data produced a data set quite comparable to the covered labor force reported by the Department of Labor (DOL) for the year 2006 (see Table 1).

Table 1
Comparison of Filtered Sample data with DOL reported "Covered Labor Force"

Sample	Mean Annual benefit	Mean duration of benefits (weeks)	Weekly benefit	Mean Weekly Wage	% With Benefits	n
DOL covered Lbr Force		15.3	$277	$797	1.9	
Replicated Covered Lbr Force	4087.70	14.84	$275.45	$798.91	1.8	80739

3.2. <u>LIS data</u>

We use data from the Luxembourg Income Study database (LIS, 2014) to examine the relationship between benefits, unemployment risk and ethnic heterogeneity across countries. There are two measures of length of time unemployed available in LIS, namely the *duration of a current* *spell of unemployment* and the *number of weeks of unemployment* during the income reference period, which in the vast majority of cases is the calendar year. We use this latter variable to measure unemployment risk. We are able to use data for 13 countries over multiple periods, with altogether 52 country-year pairs.[5] For all those observations where we can measure the number of weeks of unemployment, we also have access to a variable that indicates *immigrant status*, and we use immigrants' share of the country's population as our measure of ethnic or population heterogeneity. Finally, to measure *unemployment benefits*, we add means-tested, universal, and earnings-related unemployment benefits into a single variable, measured in terms of 2010 PPP-adjusted US dollars.

We are not able to reconstruct the covered labor force for each country-year pair as in the U.S. data. Our point of departure is the non-elderly adult (i.e., working-age) population. We measure unemployment risk by the coefficient of variation of weeks of unemployment among those who were unemployed for at least one week during the income reference period. We measure unemployment risk both including and excluding immigrants. We estimate the share of immigrants for the working-age population. Wages and unemployment benefits (for those with positive benefits) are estimated for those who are working-age and in the labor force.

Within each country-year pair in LIS, we divide the relevant samples into cells defined by age (4 groups, 10-year intervals), education (low, medium, high and "indistinguishable"; this last category often applies to immigrants), and gender, so there are 32 cells within each country-year. The regressions are estimated using cell-level estimates of the variables and weighted using cell size. All regressions include country-year fixed effects that control for overall level of GDP and other attributes shared within a country in a given year.

[5] The countries we include are Australia, Austria, Belgium, Canada, Czech Republic, Estonia, France, Germany, Greece, Ireland, the Netherlands, Spain, and the United States. See appendix for exact country years.

4. Regression results

4.1. U.S. regressions

In addition to main effects, we test if statewide variables such as ethnic fractionalization and unemployment dispersion interact to affect annual benefits, or if either moderates the increase in benefits due to an extra week of covered unemployment via interaction effects with the latter explanatory variable. We estimate the system:

1. $b^i = \alpha + \beta^I \cdot X_i + \beta^{is} \cdot X_{is} + \beta^{Int} \cdot X_{Int} + \epsilon$
2. $X_{icu} = \alpha_0 + \beta_u \cdot b^i + \beta^{I'} \cdot X_i' + \beta^{is'} \cdot X_{is}' + \beta_{wu} \cdot X_{iu} + \varepsilon_0.$

Here b^iequals individual i's annual benefits, X_i is a vector of i's individual characteristics (including X_{icu} covered weeks of unemployment); X_{is} is a vector of state characteristics in individual i's state of residence, X_{Int} is a vector of interaction variables, and $\beta^I, \beta^{is}, \beta^{Int}$ are vector-valued coefficients to be estimated. Primes in equation 2 indicate vectors including a subset of the components in the corresponding unprimed vector of equation 1. The primary individual predictor variables of interest include covered weeks of unemployment and the weekly wage; the primary state predictors include, the state's mean weekly wage, unemployment dispersion (the coefficient of variation of weeks unemployed (cv), ethnic fractionalization of the state's population), and policy mood, a measure of political liberalism in the state.[6] In addition to the interaction of fractionalization and unemployment

[6] Ethnic fractionalization within a state is computed as one minus the Herfindahl index of group population shares. The groups are African American, Asian American, Hispanic American, Native American, white non-Hispanic. Policy Mood is a measure of public support for government programs on the liberal-conservative continuum. The present authors obtained the policy mood data used in this paper from Peter K. Enns of Cornell University. See also Enns and Koch 2013).

dispersion, we tested other interaction variables multiplying various explanatory variables such as fractionalization or unemployment dispersion by individual covered weeks of unemployment.

The rationale for testing interaction effects in equation 1 is to examine if any possible effects of ethnic fractionalization or dispersion of unemployment risk are affected by the level of the other variable. That is, is the presumed negative effect of ethnic fractionalization larger in a state with greater dispersion of unemployment durations? One plausible interpretation of such a result would be that ethnic fractionalization accentuates the political conflict mechanism. Similarly, we also examine if dispersion of weeks unemployed and ethnic fractionalization have moderating effects on benefit gains from an incremental week of covered unemployment.

Theory predicts explanatory variables 'covered weeks of unemployment' and an 'individual's weekly wage' should have strong positive effects, as should the control variables policy mood and state average wage. Alternatively, our hypothesis says increases in either ethnic fractionalization or dispersion of unemployment should decrease annual benefits through main effects, interactions, or both. Other theoretical considerations based on incentive effects suggest covered weeks unemployed are determined simultaneously with benefits. Therefore, we estimate it as an endogenous variable modeled in equation 2 as depending on the individual's benefits, weekly wage, total weeks unemployed (X_{iu}), and a vector of aggregate state variables including ethnic fractionalization and dispersion of weeks unemployed. Our final estimates of the explanatory variables in equation 1 uses 2sls with individual's total weeks of unemployment as an instrument for covered weeks of unemployment.[7]

[7] We used Hausman's (1978) test for endogeneity and determined that although the coefficient of variation in weeks unemployed appears not to be endogenous in equation 1, covered weeks of unemployment may be.

To perform satisfactorily as an IV, total weeks unemployed, must be correlated with covered weeks unemployed (correlation .97); satisfy the exclusion restriction (it does not appear in equation 1); and not be correlated with equation 1's error term after controlling for the explanatory variables appearing in equation 1. To see why we argue total-weeks- unemployed is exogenous in equation 1, observe that according to all state rules, given that equation 1 is conditional on receiving benefits, the only variables determining annual unemployment benefits subject to a claimant's decisions are the weekly wage and covered unemployment. Any other variable both correlated with benefits and individual total weeks unemployed (e.g. education, experience, etc.) could not affect the former directly, but only through weekly wages or covered weeks unemployed both of which are controlled for in equation 1, and would not be in the error term. In particular, any unemployment beyond the individual's covered unemployment could have no effect on benefits paid.

Results for the OLS and 2sls regressions appear in the first and second models of Table 2. There were no significant interaction effects, suggesting the effects of dispersion of unemployment risk on benefits paid do not depend on the level of ethnic fractionalization. The two sets of regression results are quite similar with the explanatory variables explaining about three-fifths of the variation in annual benefits received by recipients. As expected, the individual variables covered weeks unemployed and weekly wage both have statistically significant coefficients with positive signs indicating estimated elasticities of 0.82 and 0.22 respectively.

Let us put these estimates into perspective. A recipient with mean average wage and covered weeks of unemployment paid benefits by a state with approximate mean levels of policy mood, ethnic fractionalization, and dispersion of unemployment risk (e.g. Ct, Ohio, Pennsylvania, Washington State), would receive an increment to annual benefits of approximately \$218 from an additional week of covered unemployment (6.7% increase in covered unemployment). To duplicate this increase in benefits, with covered unemployment constant, the weekly wage would have to increase 25% ($8.72 per one percent increase in the weekly wage).

The regression results imply ethnic fractionalization, dispersion of weeks unemployed, and policy mood each have strong effects on a state's delivery of unemployment benefits. For example, holding constant all other explanatory variables, a beneficiary receives about 12 percent lower annual benefits for each one-point increase (two s.d.) in the dispersion of weeks unemployed. This suggests the average recipient in the previously mentioned states would receive from Michigan (with its nearly average state levels of policy mood and ethnic fractionalization but a cv of weeks unemployed about 1.33 s.d. below the mean) an increase in annual benefits of approximately $271 (about one full week of benefits). These are significant amounts, suggesting both the political conflict and empathy mechanisms are important contributors to the variation in unemployment benefits paid by states.

To underscore the importance to welfare generosity of <u>both</u> the empathy and political conflict mechanisms, we note that, as expected, fractionalization and unemployment dispersion were substantially associated across states, but unexpectedly, the association was negative (correlation -.51). Consequently, estimating the model with various specifications of the filters determining the covered labor force shows that omitting the unemployment dispersion variable results in a substantial *underestimate* of the effect of ethnic fractionalization on unemployment insurance generosity of U.S. states.[8]

[8] The negative coefficient on fractionalization was always reduced between 40 and 60 percent and sometimes became insignificant.

Table 2: Regression Models and Results

Model 1	constant	Ln Weekly wage	Ln covered weeks	Coefficient of variation	Fractionalization	Policy Mood	N=1421	Adj $R^2 = .616$	F=454.83
Coefficient	4.69**	.219**	.821**	-.122**	-.484**	.008**			
s.e.	.182	.021	.018	.032	.136	.002			
Mean				2.36	.30	42.56			
2SLS	4.66	.219	.832	-.121	-.488	.008			
	.182	.021	.018	.032	.136	.002	N=1421	.613	451.04

Dependent variable is ln (unemployment income), all models. **significant at 1% level. IV is total weeks unemployed for covered weeks unemployed.

4.2. LIS regressions

As discussed in Section 4.1, our two main explanatory variables representing the hypothesized two source effects of heterogeneity on unemployment insurance generosity are the coefficient of variation of weeks unemployed and immigrants' share of the population. Regression results for multiple countries based on LIS are estimated using age*education*gender cells. We include country-year fixed effects in the estimations, so the effect of both unemployment risk profile and ethnic fractionalization is identified off differential changes within countries across time affecting groups defined by age, gender, and education in unemployment risk, which in turn induces different benefit levels. This approach is similar to that first used by Blundell, Duncan and Meghir (1998) to examine the effect of tax reform on labor supply.

We estimate two sets of regressions: one regresses the level of benefits in each cell on unemployment risk, earnings, and immigrant status, and the other regresses the benefit replacement rate (benefits/earnings) on those same variables. We estimate each of the regressions twice; first measuring unemployment risks and benefits (and replacement rates) for natives only and again including the immigrant populations. We show results for explanatory variables unemployment risk and immigrant share separately in Table 3 (Columns 1 and 2 respectively) and then together in

Column 3. All regressions control for the level of earnings. Panel A in Table 3 exhibits the results for the case with the dependent variable equal to benefit levels, and panel B shows results for benefit replacement rates.

Table 3. Regression results – benefits and benefit-replacement rates regressed on unemployment risk and immigrant share

A. Dependent variable: unemployment benefits (2010 international USD)

rbind	Natives Only			Also Immigrants		
	1	2	3	4	5	6
cvweeksue	-3776.8 (397.4)		-3776.8 (397.4)	-4225.6 (380.8)		-4225.6 (380.8)
Earnings	58.8 (7.8)	63.1 (7.9)	58.8 (7.8)	59.8 (7.0)	65.5 (7.2)	59.8 (7.0)
Immigrant		-49074.7 (6638.1)	-48467.6 (6422.0)		-47786.4 (5963.3)	-46997.0 (5786.8)
N	1266	1353	1267	1331	1407	1332
K	61	60	61	61	60	61
σ	4e+04	4e+04	4e+04	4e+04	4e+04	4e+04
Adj R^2	0.7	0.7	0.7	0.7	0.7	0.7
Countries	13	13	13	13	13	13
Countries - years	52	52	52	52	52	52

B. Dependent variable: unemployment benefits replacement level (benefits/earnings)

rbind	Natives Only			Also Immigrants		
	1	2	3	4	5	6
Cvweeksue	-12.29 (1.10)		-12.29 (1.10)	-14.14 (1.09)		-14.14 (1.09)
Earnings	-0.03 (0.02)	-0.03 (0.02)	-0.03 (0.02)	-0.04 (0.02)	-0.04 (0.02)	-0.04 (0.02)
Immigrant		-163.28 (19.40)	-164.73 (17.70)		-157.42 (19.84)	-158.16 (16.62)
N	1266	1353	1267	1331	1407	1332

K	61	60	61	61	60	61
σ	1.1e+02	1.2e+02	1.1e+02	1.1e+02	1.3e+02	1.1e+02
Adj R^2	0.75	0.69	0.75	0.76	0.66	0.76
Countries	13	13	13	13	13	13
Countries - years	52	52	52	52	52	52

Source: Authors' estimates using grouped data from the Luxembourg Income Study. Regressions include cell as well as country-year fixed effects.

The general impression from these regressions is consistent with our theoretical hypotheses, and the results from the U.S. data. As expected because of the reduction in population empathy, benefit levels decline significantly as the immigrant share of the population rises. Moreover, across country variation in unemployment risk profiles shows dispersion of risk (as measured by the coefficient of variation of weeks unemployed) also has a large negative effect on benefits, and that effect is unchanged when we include immigrants' population share in the regression.

5. Conclusion

Basing our argument on theoretical models, we claim population heterogeneity should affect the generosity of a polity's social insurance programs through two distinct channels. First, population heterogeneity likely reduces the well-known positive empathy effect thereby reducing welfare generosity because ethnic diversity undermines sentiments of solidarity among a citizenry. Second, ethnic heterogeneity likely increases dispersion of incomes thereby intensifying political conflict because heterogeneity of individual income risks renders it more difficult to achieve social consensus concerning tax-benefit programs. We utilized regression analysis on two data sets covering highly diverse polities, and found that distinct empathy and political conflict effects on unemployment insurance programs do appear to characterize contemporary politics.

Depending on the specific cross-jurisdiction relationship between ethnic and unemployment heterogeneity within a data sample, estimates of the empathy and political conflict effects require modeling both mechanisms. To date, analyses of the negative relationship between ethnic heterogeneity and the size of the welfare state probably over- or underestimate the empathy effect. For example, perhaps surprisingly, within the U.S. the cross-state relationship between ethnic fractionalization and unemployment dispersion is negative (correlation -0.51). Had our analysis omitted a measure of unemployment dispersion, the negative effect of ethnic fractionalization would have been underestimated! Future research should investigate these effects and relationships in more historical terms, especially with respect to the welfare states of Europe. Unfortunately, at this time, we are unaware of data that would enable us to measure risk heterogeneity in the European countries around 1960, when welfare-state legislation took off.

Appendix: Robustness Tests of U.S. Data and Description of LIS Data

To examine the robustness of the regression results, we subjected the regression model to several tests. First, we estimated the model using alternative data filters corresponding to different programmatic unemployment insurance policies. Secondly, we subjected the model to robust heteroskedasticity estimation. Third, we examined the 2sls model using a different, but weaker, instrumental variable for covered weeks of unemployment. Finally, we examined how robust our results are to different specifications of the aggregate state variables policy mood, ethnic fractionalization, and dispersion of unemployment risk.

Table 4 exhibits the regression results (equation 1 only) from several alternative filters of the CPS data. Model 3 in row 1 establishes an initial baseline indicative of the robustness of the hypothesized effects. The reported results are from a regression that applies no filters to the CPS data. The results appear quite similar to those of our original model in Table 2. All estimated coefficients retain the correct signs and remain significant, and the estimated coefficients for log of the weekly wage and policy mood are virtually unchanged. Differences in the estimated coefficients for the other variables are primarily due to lower correlations between these variables (especially log of covered weeks unemployed) with the dependent variable, a result indicated by a nearly two-thirds reduction in R^2.

The differences between models 1 and 3 are largely due to two facts. The completely unfiltered data behind model 3 contains many records with either large reported benefits and small (most often zero) reported weeks of unemployment or large durations of unemployment with small annual benefits. These records undercut the relationship between benefits and covered weeks of unemployment. Many are simply data outliers created by reporting and coding errors, and the fact that, due to administrative lags in processing claims, many recipients (especially those with zero reported unemployment) likely qualified for benefits during the latter part of 2005, did not receive payments until 2006, and found employment early in 2006. These issues are not likely to affect correlations between benefits and weekly wages, fractionalization, or policy mood, but will certainly affect the correlation between benefits and weeks unemployed and the latter's dispersion. As suggested, we addressed these issues by applying data filters based on allowed minimum and maximum WBA restrictions on the duration of benefits.

We subjected the model to several alternative specifications of the data filters simulating policy variables governing the measured relationships between benefits and the explanatory variables. Each of these produced results similar to those of model one with estimated coefficients varying to some degree, but always statistically significant and with the hypothesized signs. To illustrate with a particularly relevant alternative filter, Table 4's model 4 is the result of applying the actual state maximum and minimum WBAs to each record.

Table 4

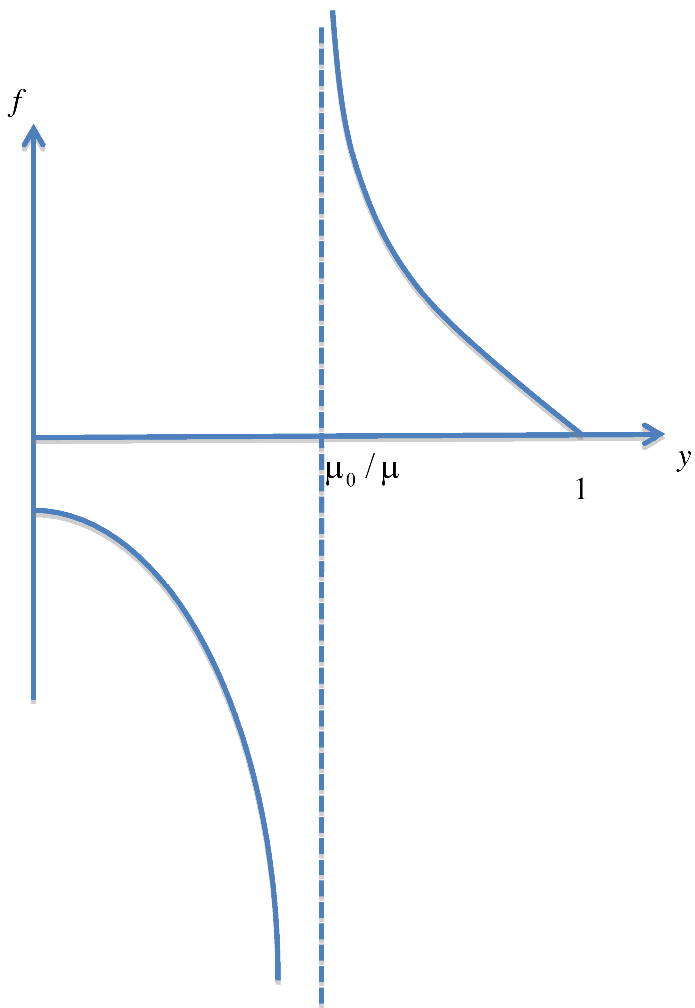
Model 3	constant	Ln Weekly wage	Ln covered weeks	Coefficient of variation	Fractionalization	Policy Mood	N=1874	Adj R^2 .22	F=108.83
Coefficient	5.30	.218	.565	-.177	-.549	.009			
s.e.	.262	.025	.027	.052	.219	.004			
t value	20.29**	8.61**	21.12**	-3.42**	-2.51**	2.61**			
Mean				2.36	.30	42.75			
Model 4	4.76**	.188**	.841**	-.131**	-.717**	.009**	N=1393	.530	315.35
	.198	.020	.022	.037	.157	.003			
Mean				2.37	.299	42.68			
Model 5	5.06	.215	.682	-.129	-.443	.008			
	.375	.021	.122	.033	.143	.002	N=1421	.099	32.25

Our final two tests involved correcting for possible heteroskedasticity in the error terms of our primary equation, and estimating the model via 2sls with an alternative instrument for covered weeks of unemployment. Examining the scatter plot between the dependent variable and the regression's standardized residuals suggested the possibility of heteroskedasticity. However, robust estimation produced virtually identical standard errors and identical coefficients significant at one percent levels, and the results are not reported. Our alternative instrument for covered weeks of unemployment was a respondent's "work status" during March of 2007 the year following the year under discussion. Work status encompassed possibilities ranging from employed full-time or part-time to unemployed or not in labor force. Coding a categorical variable 1 if respondent was full-time and 2 otherwise produced weak negative correlations between covered weeks of unemployment (-0.05) and logs of these variables (-0.15). We argue work status during March of 2007 does not affect unemployment benefits received during 2006 (hence the error term of equation 1) after a recipient's covered unemployment and wage are controlled. Model 5 of Table 3 shows the much weaker instrumental variable resulted in a reduced coefficient effect from covered weeks of unemployment, however the effect remained highly significant, and the effects of other explanatory variables changed little.

Table 5: Countries and Years Covered in LIS Data

Country	Years
Australia	1985, 1989
Austria	1994, 1997, 2000, 2004
Belgium	1995, 2000
Canada	1987, 1991, 1994, 1997, 2000, 2004
Estonia	2007, 2010
France	2005
Germany	1984, 1989, 1994, 2000, 2004, 2007, 2010
Greece	1995, 2000, 2004, 2007, 2010
Ireland	1994, 1995, 1996, 2000, 2004, 2007, 2010
Netherlands	2004, 2007, 2010
Slovak Republic	2004, 2007, 2010
Spain	1995, 2000, 2004, 2007, 2010
United States	1997, 2000, 2004, 2007, 2010

Figure 1   Graph of the function φ



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