PARTICIPATION IN HETEROGENEOUS COMMUNITIES*

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This paper studies what determines group formation and the degree of participation when the population is heterogeneous, both in terms of income and race or ethnicity. We are especially interested in whether and how much the degree of heterogeneity in communities influences the amount of participation in different types of groups. Using survey data on group membership and data on U. S. localities, we find that, after controlling for many individual characteristics, participation in social activities is significantly lower in more unequal and in more racially or ethnically fragmented localities. We also find that those individuals who express views against racial mixing are less prone to participate in groups the more racially heterogeneous their community is. These results are consistent with our model of group formation.

I. INTRODUCTION

Many observers, including economists, are convinced of the importance of the complex stock of social norms, trust, and networks of civic engagement that has been grouped under the term "social capital."[1] As forcefully argued by Putnam [1993, 1995a, 1995b], social capital may produce several positive socioeconomic effects which can spur economic success.

But, what determines social capital? Figure I displays the distribution across U. S. states of an index of social capital that embodies information on trust, membership in groups, and voting behavior.[2] Overall, this index is highest in the North/Northwest

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1. See Coleman [1990] for an extensive discussion of the foundations of "social capital" in social theory.

2. We constructed this index extracting the principal components from three variables obtained from the General Social Survey: the percentage of people in the state who belong to a group, the percentage who trust others, and the percentage who voted in the last presidential election. A detailed description of the data is given in the empirical part of the paper. Principal components analysis has been used to construct an index of social capital also by Putnam and Yonish [1998]. Our index is very highly correlated with theirs.

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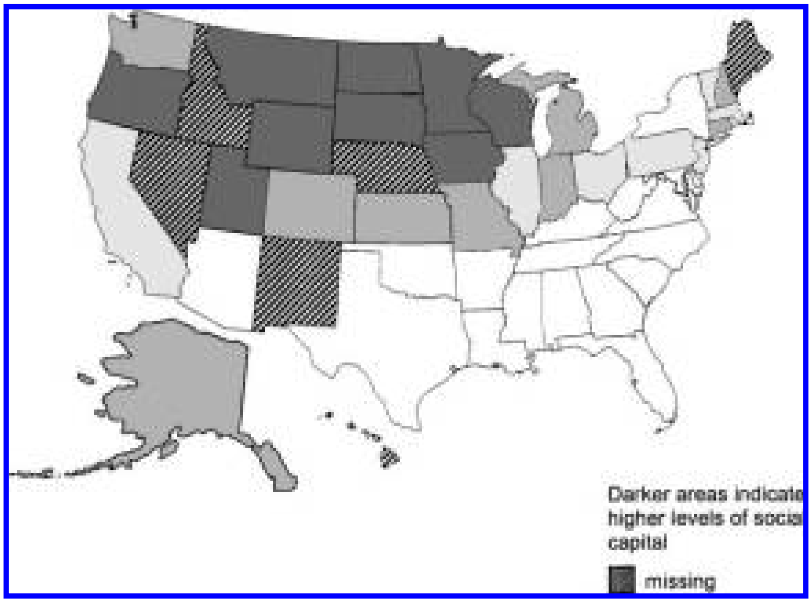

FIGURE I
Social Capital Index

and lowest in the South/Southeast. The former regions are characterized by a racially homogeneous population and relatively low income inequality, while the latter have the opposite features. In particular, the top five U. S. states in our social capital ranking are North Dakota, Utah, Minnesota, Wyoming, and Montana. All of them are very homogeneous: they are almost one standard deviation below the national mean for both racial fragmentation and income inequality.

The relationship between homogeneity and social capital may not be only a U. S. phenomenon. For instance, according to the international ranking reported by Knack and Keefer [1997], the five countries with highest levels of trust are Norway, Finland, Sweden, Denmark, and Canada; the same countries rank among the top ones for associational activity and norms of civic cooperation. These countries have an ethnically homogeneous population and very low levels of income inequality.

This paper shows that there is something systematic about the relationship between heterogeneity of communities (in terms of income and race) and the level of social capital: more homogeneous communities have a higher level of social interactions leading to more social capital. Even though this paper focuses on U. S. cities, our results have more general implications. Racial heterogeneity and income inequality vary greatly in different countries, and, through their effect on social capital, they may influence economic outcomes and public policies.

One of the common criticisms of the notion of social capital is that it is very hard to "measure," hence difficult to use in empirical analysis. Rather than focusing on a broad index of social capital, we study a critical component of it that can be measured fairly precisely, namely the participation in associational activities, such as religious groups, hobby clubs, youth groups, sport groups, etc. Our interest in these activities is motivated by Putnam [1993, 1995a, 1995b] who suggests that these types of social interactions are particularly conducive to generating the beneficial effects of social capital. More specifically, participation in social groups may lead to the transmission of knowledge and may increase aggregate human capital, and the development of "trust," which improve the functioning of markets.[3] In addition, social interactions and networks may influence individual outcomes, from criminal activities, to fertility, to the labor supply.[4] Also, participation in groups is known to be highly correlated with political participation, and the latter has critical implications for policy choices.[5] If the wealthy or more educated have a disproportionate propensity to join groups and engage in political action, then public policies may be tilted in their favor. This may lead to vicious circles, in which disadvantaged minorities participate less, have less "voice," and become even more disadvantaged, leading to a variety of social problems.[6]

3. On the effects of positive spillovers in the transmission of human capital, see Romer [1986], Lucas [1988], and Bénabou [1996]. On trust see La Porta et al. [1997].

4. For theoretical work on the effects of transmission of information in group and informational cascades, see, for instance, Banerjee [1992] and Ellison and Fudenberg [1995]. An early empirical contribution on the importance of networks is Case and Katz [1991].

5. See Verba and Nie [1987] and Verba, Schlozman, and Brady [1995].

6. Even though participation is typically associated with "positive" socioeconomic outcomes, social networks may also transmit "negative" norms. For example, the so-called "culture of poverty and welfare" may find its roots in social networks propagating incentives to search for welfare rather than work. See, in particular, Cutler and Glaeser [1997] and Bertrand, Luttmer, and Mullainathan [2000] for recent empirical work on this important question. A different theoretical perspective on groups and social norms is given by Berman [2000].

In our model individuals prefer to interact with others who are similar to themselves in terms of income, race, or ethnicity. If preferences are correlated with these characteristics, then our assumption is equivalent to saying that individuals prefer to join groups composed of individuals with preferences similar to their own.[7] Given this setup, one may expect that diffuse preferences for homogeneity may decrease total participation in a mixed group if fragmentation increases. However, individuals may choose to sort into homogeneous groups. Therefore, it is not clear a priori under which conditions more heterogeneity in the population would lead to more or less participation. In the theoretical part of the paper, we investigate this issue. Our model departs from standard club theory, since our groups do not require contributions, do not have congestion effects, and are based on free entry and exit of individuals. We make these assumptions because we want to focus on how the composition of the group affects individual choices of participation.

Our empirical results on U. S. localities suggest that income inequality and racial and ethnic heterogeneity reduce the propensity to participate in a variety of social activities including recreational, religious, civic, and educational groups.[8] Among the various forms of heterogeneity, racial fragmentation seems to have the strongest negative effect on participation. Furthermore, and consistent with our model, these results are more marked for the groups in which direct contact among members is important, like churches and youth clubs, while heterogeneity matters less or not at all in groups with a low degree of interaction, for instance professional associations. Finally, our model predicts that individuals relatively more averse to mixing with different types should be those more negatively influenced by heterogeneity in the community. We successfully test this more stringent implication of the model, by exploiting individual data on attitudes toward race

7. Theoretical results by Conley and Wooders [1996] are consistent with this assumption. They show that when agents can be crowded (positively or negatively) by the skills of other people in their jurisdictions, taste-homogeneous jurisdictions are optimal. To the extent that tastes are correlated with income and race, our assumption follows.

8. We define by "race" the census classification of black, white, Asian, American Indian, and other. We define by ethnicity the classification by ancestry, like Italian, Irish, etc. Throughout this paper we will use the terms "white" and "black" instead of Caucasian and African-American, for the sake of brevity.

relations. In summary, we find that social capital is lower in more unequal and heterogeneous communities.[9]

Recent research has highlighted a positive bivariate correlation between inequality and social capital measures at the state level.[10] Our multivariate analysis, conducted with individual level data on participation and community level measures of income inequality, sheds light on this issue. As for race, much empirical research has studied the effects on public policy of ethnic and racial heterogeneity. For instance, Alesina, Baqir, and Easterly [1999, 2000] show that the supply of "core" productive public goods is lower and measures of patronage are higher in more racially fragmented localities. Glaeser, Scheinkman, and Shleifer [1995], Cutler and Glaeser [1997], Poterba [1996], Luttmer [1997], and Goldin and Katz [1999] study the role of racial conflict as a determinant of education policies and several other characteristics of U. S. cities. Alesina, Baqir, and Hoxby [1999] link racial and ethnic fragmentation to the number of jurisdictions in the United States.

This paper is organized as follows. Section II presents a model that generates predictions linking heterogeneity of the population and the level of participation in social activities. Section III describes our empirical strategy and data. Section IV highlights some simple correlations at the state level between income inequality, racial and ethnic fragmentation, and measures of participation and social capital. Section V presents our econometric results. The last section concludes.

II. THE MODEL

II.1. Setup

Consider a community populated by two types, "blacks" and "whites," their size being $B > 0$ and $W > 0$, respectively. For ease of exposition, we shall think about "race" as what identifies the

9. To the extent that one views social capital as a public good, this result is consistent with the findings of Alesina, Baqir, and Easterly [1999]. That paper focuses on public goods provided by means of public policies (and taxes) chosen by a median voter who "decides" for the entire community. The present paper instead focuses on a public good (social capital) not generated by policies but by the interaction of private individuals in private groups. One may think of public policies that increase the incentive to participate and create social capital, but this issue is not explicitly addressed in this paper.

10. See Putnam's presentation at the Saguaro Seminar, October 1998, Harvard University.

types; below we discuss other discriminating characteristics, especially income. The population is uniformly distributed on a line, and both types have a uniform distribution on the interval [0,1].[11]

Each individual decides whether or not to participate in a group. We do not allow for participation in more than one group. There are no congestion costs and no economies of scale: thus, group size does not influence individual utility. Members of the group cannot exclude new members, and entry into (exit from) the group is free and costless. The reservation utility from nonparticipation is $\bar{u}$ for everybody.[12] The utility from participating in the group depends on the composition of the group and on the distance between the individual's and the group's location. Let P_B (P_W) be the proportion of blacks (whites) within the group. The utility from participation for an individual of type $j = W,B$ located at a distance l from the group's headquarters (HQ) is

$$(1) \quad U_j = u(\alpha,P_{-j}) + v(l) \qquad \text{for } j = B,W$$

$$u_\alpha(\cdot) < 0, \qquad u_P(\cdot) < 0, \qquad v_l(\cdot) < 0$$

$$(\text{A1}) \qquad u_{\alpha P}(\cdot) \leq 0$$

$$(\text{A2}) \qquad u_\alpha(\cdot)\big|_{P=0} = 0, \qquad u_P(\cdot)\big|_{\alpha=0} = 0,$$

where P_{-j} is the proportion of group members whose type is different from j's type, and α is a taste indicator that varies across individuals and captures the intensity of an individual's aversion to the opposite race. The functions $u(\cdot)$ and $v(\cdot)$ are continuously differentiable in their arguments; $u_\alpha(\cdot)$ and $u_P(\cdot)$ represent the partial derivatives of $u(\cdot)$ with respect to α and P_{-j}, respectively, and $v_l(\cdot)$ is the derivative of $v(\cdot)$ with respect to l. Underlying assumption $u_P(\cdot) < 0$ is the preference for participating in a social activity with members of one's own type. Given our assumption that the cross derivative of $u(\cdot)$ with respect to α and P_{-j} is (weakly) negative, an increase in the proportion of members of the opposite type decreases an individual's utility more, the higher is α. We label α the "degree of intolerance" or "aversion." Assumption

11. This assumption is obviously restrictive and does not address the presence of racial segregation in housing. The nature of our arguments remains valid as long as B and W have the *same* distribution (even if it is not uniform).

12. The reservation utility may of course be a function of individual characteristics. While this will be important for the empirical analysis that follows, we abstract from it in the theoretical discussion.

A2 states that utility, when evaluated at $P_{-j} = 0$, is independent of α, and when evaluated at $\alpha = 0$, is independent of P_{-j}. In other words, when the group is totally composed of one's own type, intolerance does not play any role, and when an individual has zero intolerance, the composition of the group does not affect her utility. In what follows, we will assume that $u(0,P_{-j}) = 0$ purely to simplify the notation.[13]

Each individual is therefore characterized by two parameters (in addition to the type B or W): aversion and distance. We assume that the aversion parameter α and individual location are independently distributed on [0,1]. We denote the density function for α and for the distance l with $f_\alpha(\cdot)$ and $f_l(\cdot)$, respectively, and assume that these densities are continuously differentiable on their support.

In summary, our model has three basic components. The first is that people prefer to interact with members of their own type. The second component is that people's utility from participation decreases the farther they have to travel to go to group meetings, etc. In the absence of traveling costs there is no reason why people should form mixed groups (everyone is better off by being with his own type, no matter how far away the group is located), nor is there a reason why more than two groups should form (or, in general, as many as the different types in the population), given that there is no congestion. We require the groups to have a minimum size in order to avoid the degenerate outcome in which every individual forms his or her own group. Finally, our model departs from club theory and from the local public goods literature in that no contributions are paid by the members and utility is derived directly from participation.

II.2. Equilibrium with One Group

We begin by assuming that the minimum size of the group is greater than half of the population, so that at most one group can form in equilibrium. We relax this assumption below.[14] Consider an individual of type $j = B,W$ and aversion α who has the choice between joining a group with composition P_{-j} and distance l on the

13. If we remove this assumption, the only change is that the extreme of integration with respect to l in expression (3) below becomes $v^{-1}(\bar{u} - u(0,P_{-j}))$, which, anyway, is independent of P_{-j} given assumption A2.

14. For a simpler version of this model with one group and no travel costs, see Alesina and La Ferrara [1999].

one hand, and staying out on the other. This individual participates in the group if and only if

$$u(\alpha, P_{-j}) + v(l) \geq \overline{u}. \tag{2}$$

In what follows, we concentrate on symmetric equilibria in which the HQ is in the middle of the group.[15]

The probability that the participation constraint (2) is satisfied can be written as

$$\prod (P_{-j}, \overline{u}) \equiv 2 \int_0^{v^{-1}(\overline{u})} \int_0^{g(\overline{u}-v(l),P_{-j})} f_\alpha(\alpha)\, f_l(l)\; d\alpha\; dl, \tag{3}$$

where $g(\cdot)$ is obtained by inverting $u(\cdot)$. Our assumptions imply that the partial derivative of $g(\cdot)$ with respect to P_{-j} is negative and hence that $\partial\Pi(P_{-j},\overline{u})/\partial P_{-j} < 0$. Expression (3) states that, for given group composition P_{-j}, the individuals of type $j = B,W$ for whom the participation constraint holds are those who are located no farther than $v^{-1}(\overline{u})$ from the group headquarters *and* whose aversion parameter is no greater than $g(\overline{u} - v(l),P_{-j})$.

The total mass of individuals of type B and type W willing to participate in the social activity is then, respectively,

$$\tilde{B} = \Pi(P_W,\overline{u}) \cdot B \tag{4}$$

$$\tilde{W} = \Pi(P_B,\overline{u}) \cdot W. \tag{5}$$

DEFINITION 1. An *equilibrium* is a group composition (P_B^*,P_W^*) such that for both types none of the members wishes to leave the group and none of the nonmembers wishes to join.

In equilibrium the proportion of individuals of type B in the group, P_B, must be equal to the ratio of the mass of the participants $\tilde{B}$ to the total mass of participants $\tilde{B} + \tilde{W}$. The two conditions defining the equilibrium are therefore

$$P_B = \frac{\Pi(P_W,\overline{u}) \cdot B}{\Pi(P_W,\overline{u}) \cdot B + \Pi(P_B,\overline{u}) \cdot W} \tag{6}$$

$$P_W = 1 - P_B, \tag{7}$$

15. The symmetric case is the natural reference point, both from a normative and from a positive point of view. A utilitarian social planner would always locate the HQ in the middle of the group, to minimize total travel costs. Alternatively, if after the group were formed the members decided by majority vote where to locate the HQ, they would choose the middle of the group (which is the median, given our assumptions). We can also handle the nonsymmetric case. Details are available.

which together give us the "fixed point" equilibrium condition contained in the following.

PROPOSITION 1. There exists at least one equilibrium $P_B^* \in [0,1]$ which solves

$$P_B = \frac{\Pi(1 - P_B,\overline{u})}{\Pi(1 - P_B,\overline{u}) + \Pi(P_B,\overline{u})W/B}. \tag{8}$$

Proof. The proof of this proposition like all the other proofs is in Appendix 1.[16]

In what follows, for notational convenience we will suppress the term $\overline{u}$ from the arguments of $\Pi(\cdot)$. We will restrict our attention to locally stable equilibria. A formal definition and a necessary and sufficient condition for local stability are provided in Appendix 1.

Figure II provides the intuition. The equilibrium value(s) of P_B is (are) given by the intersection of the function in the right-hand side of (8) with the 45° line. The right-hand side of (8) represents the fraction of members type B in the group that is "generated by the reactions" of both types to a given composition $(P_B, 1 - P_B)$. The intersection(s) with the identity line give(s) the value(s) of P_B at which both reactions are consistent with the actual proportions. Our stability condition requires that the slope of the above function at the point of intersection with the 45° line be less than one.[17]

Figure II depicts various possible configurations of equilibria. In panel (a) we have a unique interior equilibrium, i.e., the group that forms is "mixed," with a proportion $P_B^* \in (0,1)$ of blacks and $(1 - P_B^*)$ of whites. This equilibrium is stable. Suppose that you add one black member to the group, so the composition becomes $P_B > P_B^*$. The shape of the curve in panel (a) suggests that this "more favorable" composition for types B does not trigger enough participation of new B members, nor does it induce enough W members to exit, so the group goes back to the initial equilibrium. In panel (d) the opposite occurs; that is, any slight increase

16. Underlying Proposition 1 is the condition, $\Pi(P_W^*,\overline{u}) \cdot B + \Pi(P_B^*,\overline{u}) \cdot W > .5(B + W)$, namely that the utility and distribution functions are parameterized so that the group that forms in equilibrium meets the minimum size requirement.

17. Notice that—given our assumptions on $u(\cdot)$, $v(\cdot)$, $f_\alpha(\cdot)$, and $f_l(\cdot)$—the slope of the "reaction function" defined by the right-hand side of (8) is always nonnegative.

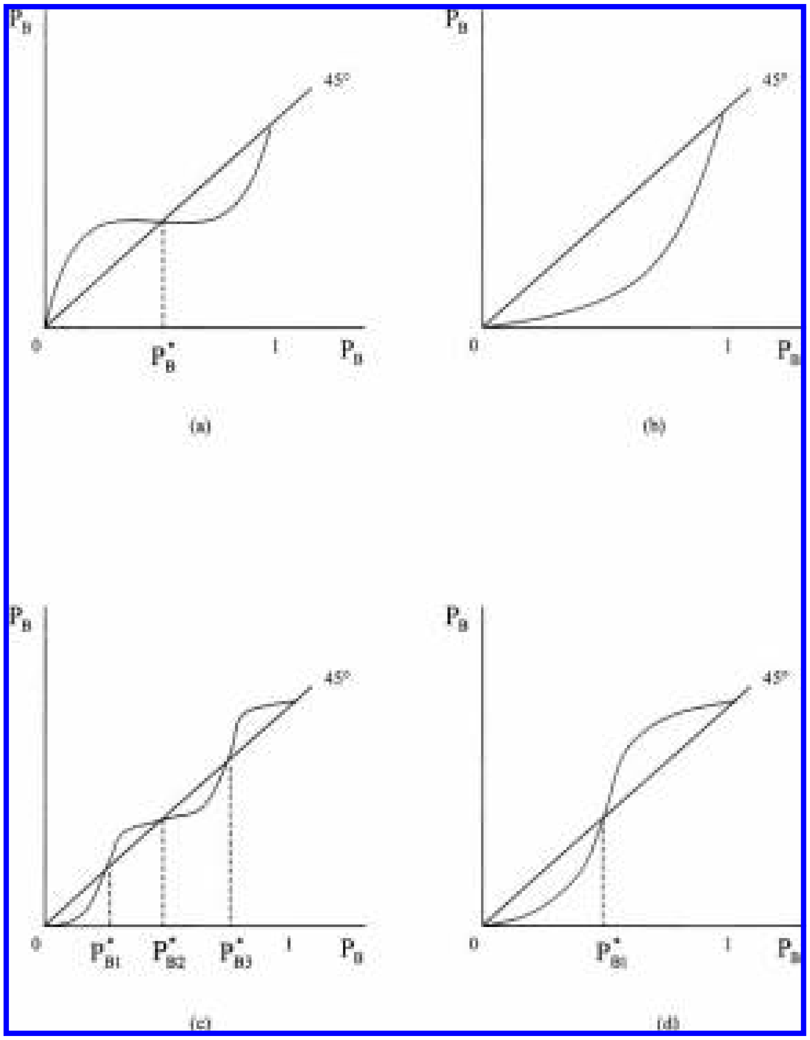

FIGURE II
Equilibrium Configurations

(reduction) in the fraction of B members from P_B^* will trigger an inflow (outflow) of B types and outflow (inflow) of W types, so that the composition of the group moves to complete homogeneity with $P_B^* = 1(P_W^* = 1)$. While in panel (d) either $P_B^* = 0$ or $P_B^* = 1$ can in principle be stable equilibria[18] in panel (b) only $P_B^* = 0$ is. Finally, panel (c) illustrates the case of multiple equilibria.[19]

We are interested in two features of the equilibrium. The first is how the composition of the group relates to the composition of the total population; the second is *who* among the heterogeneous individuals of a given type will choose to participate and who will stay out.

LEMMA 1. Let (P_B^*, P_W^*) be a unique stable equilibrium. Then $B < W \Leftrightarrow P_B^* < P_W^*$.

COROLLARY 1. If $B \neq W$ and (P_B^*, P_W^*) is a unique stable equilibrium, then either $(P_B^*/P_W^*) < (B/W) < 1$ or $1 < (B/W) < (P_B^*/P_W^*)$.

COROLLARY 2. Let (P_B^*, P_W^*) be a unique stable equilibrium. A necessary condition for an individual type $j = B,W$ to participate is that her α falls in the interval $[0,\bar{\alpha}^j]$, and her distance in the interval $[0,\bar{l}]$, where $\bar{\alpha}^j$ and $\bar{l}$ solve, respectively,

$$u(\bar{\alpha}^j, P_{-j}^*) + v(0) = \bar{u} \tag{9}$$

$$v(\bar{l}) = \bar{u}. \tag{10}$$

In particular, $B < W \Leftrightarrow \bar{\alpha}^B < \bar{\alpha}^W$.

Lemma 1 and Corollary 1 show that the unbalance between the types in the population is *magnified* within the group: not only is the minority type underrepresented in the social activity, but it is less than proportionately represented compared with its weight in the population. Thus, the social group is more homogeneous than the whole population.

Corollary 2 can help us understand this result. In equilibrium, the individuals participating in a mixed group are those who are "not too averse" to the opposite race, and "not too far" from the location of the group. Corollary 2 tells us exactly "how far" we can go in these two dimensions. The maximum geographical distance compatible with the participation constraint is such that

18. As a matter of fact, given our assumptions that $B < W$ and that the minimum size of the group is greater than ½, only $P_W^* = 1$ can be a stable equilibrium. As we will see, this will no longer be the case in the multiple-groups model.

19. Of the five equilibria depicted in the figure, 0, P_{B2}^*, and 1 are stable, while P_{B1}^* and P_{B3}^* are unstable. The issue of the selection among multiple equilibria will not be addressed in this paper.

an individual who receives the lowest possible disutility from heterogeneity (i.e., an individual with $\alpha = 0$) is indifferent between joining the group or not. Given A2, this distance is independent of the composition of the group and of the individual's type. The dimension we are most interested in is "aversion" to the opposite type. The maximum intolerance compatible with the participation constraint is that of the individual located exactly where the group's headquarters are, and is given by the $\overline{\alpha}^j$ solution of (9). We can refer to these individuals as the "most averse" members. Corollary 2 tells us that the most averse member of type B will be less "tolerant" than that of type W if and only if blacks are a minority in the population. In fact, although the two types have the same ex ante distribution of α, we cannot observe the same degree of "aversion" for B and W in the equilibrium composition of the group. If we did, the same α should be indifferent between participating as a majority or as a minority. Instead, the fact that B is underrepresented in the population induces even some relatively "moderate" B individuals (low α) to stay out, while W individuals manage to keep some relatively "participation averse" (high α) members in the group. This coupled with the sheer unbalance in the numbers B and W produces the "magnification" effect described above.

II.3. Heterogeneity and Participation

We are now ready to study how a change in the heterogeneity of the population influences the total mass of participants.

DEFINITION 2. The degree of *heterogeneity* is the probability that two randomly drawn individuals from the population belong to different types.

This is the same definition of our empirical analysis. Obviously, in our case of two types, a 50-50 split has the maximum level of heterogeneity. Denote by w the fraction of whites in the population; i.e.,

$$w \equiv W/(W + B). \tag{11}$$

An increase in w represents a *decrease* in heterogeneity if $w \geq \frac{1}{2}$ and an *increase* if $w < \frac{1}{2}$.

DEFINITION 3. The *aggregate level of participation S* is the share of the total population who belongs to a group:

$$S \equiv \frac{\tilde{B} + \tilde{W}}{B + W} = \prod (1 - P_B^*)(1 - w) + \prod (P_B^*)w. \tag{12}$$

Under very mild sufficient conditions on $\Pi(\cdot)$, described in Appendix 1, the following holds.

PROPOSITION 2. If a unique stable equilibrium exists, an increase (decrease) in heterogeneity reduces (increases) total participation; i.e.,

$$w < \frac{1}{2} \Rightarrow \frac{dS}{dw} < 0$$

$$w \geq \frac{1}{2} \Rightarrow \frac{dS}{dw} > 0.$$

The intuition is simple. We can have two kinds of equilibria. In the first one the group is homogeneous: by the minimum size requirement this means that all the individuals of the majority type participate, and none of the minority type does. In the second kind of equilibrium the group is mixed, with proportion $0 < P_B^* < 1$ identified by Proposition 1. To examine the impact of heterogeneity, suppose that whites are the majority ($w > ½$), and consider a decrease in w leaving the size of the total population unchanged—an increase in heterogeneity. In the first kind of equilibrium, all that happens is that the total number of whites is lower, and since all W types were participating, the size of the group decreases, and so does aggregate participation.[20] More interesting is the case of the "mixed" group. We have established that if $B < W$ then $P_B^* < P_W^*$, and that $\partial P_W^*/\partial w > 0$. Consider what happens to S defined in (12) when whites are the majority and w decreases. First of all, since $\Pi(P_B^*) > \Pi(1 - P_B^*)$, the fall in w creates an absolute loss in participants greater than the gain created by the increase in $(1 - w)$. Furthermore, the fractions of the two types

20. One could devise a decrease in w such that B types become the majority. However, as long as the probability that two randomly drawn individuals belong to different types has increased (our definition of heterogeneity), then the total mass of B types must be lower than the total mass of W was before the change. This implies that the new (homogeneous) group will be smaller than before and hence confirms our conclusion.

who participate change with the new value of heterogeneity. The sufficient condition mentioned in the text ensures that the increase in the fraction of blacks is not so overwhelmingly larger than the decrease in the fraction of whites to overcome the first effect. Appendix 1 reports a simple example to highlight the critical features of our results in the one-group and in the multiple-group context, and provides an intuitive graphical illustration.

II.4. Multiple Groups

We now set a minimum size such that at most $N \geq 2$ groups can exist in equilibrium. If we denote by $\tilde{B}^k$ and $\tilde{W}^k$ the mass of B and W participants to group k, this amounts to requiring that

$$\tilde{B}^k + \tilde{W}^k > (B + W)/(N + 1). \tag{A3}$$

The main differences with the results obtained so far arise with respect to the possibility that multiple *homogeneous* groups form; i.e., that B and W types sort into perfectly segregated groups. For clarity of exposition, we will in turn analyze the case of *mixed* and that of *homogeneous* groups, although it should be clear that the same set of preferences can be consistent with some groups being mixed and some homogeneous if multiple equilibria for P_B^* exist, as depicted for example in Figure II, panel (c).

In equilibrium a member's utility from participating in a group must exceed not only the reservation utility $\bar{u}$, but also the utility that the same individual could get from joining another group. As we will see, this implies that if two adjacent groups exist, the individual located at the "border" between the two groups must be indifferent between them. Any group will therefore fall into one of two categories: one in which no neighboring group exists and the "geographic coverage" of the group (i.e., the distance between the two extreme members of the group) is maximum and is determined by a participation constraint like (2); and one in which there is at least one adjacent group and the border(s) of the group is (are) characterized by the indifference condition mentioned above. Before we turn to the analysis of these two configurations, we define our equilibrium concept in the multiple-group context.[21]

21. In the definition of an n-group equilibrium, we employ a slight abuse of notation in using the intersection operator to express that two groups cannot

DEFINITION 4. An n-group *equilibrium* consists of n couples (P_B^{k*}, L^k), $k = 1, \ldots, n \leq N$, with P_B^{k*} denoting the fraction of members type B in group k, and L^k denoting the distance between the two most distant members of group k, with $\sum_{i=1}^{n} L^i \leq 1$ and $\cap_{i=1}^{n} L^i = 0$, such that for each (P_B^{k*}, L^k) none of the members wishes to leave the group and none of the nonmembers wishes to join.

In what follows, we start by analyzing the case of multiple mixed groups. We consider the case in which the utility function admits a unique stable interior equilibrium, $P_B^* \in (0,1)$, and investigate how the conclusions of the one-group model are affected by the existence of multiple groups.[22] We start by considering an equilibrium with *n disjoint* mixed groups.

LEMMA 2. If an equilibrium with n disjoint mixed groups exists, all groups must be identical, and each of them is defined by the equilibrium conditions of the single-group model.

The fact that the groups are disjoint implies that the members at the "frontier" of the group are held down to their reservation utility $\bar{u}$. In this case the existence of other groups does not affect the conditions characterizing the borders of the group: the geographic coverage of the group is maximum and is determined by condition (10) in Corollary 2. Furthermore, the uniform distribution of B and W types on the line ensures that every segment of it is characterized by the same proportion B/W in the population. This, coupled with condition (10), implies that all groups will be identical n-replicas of the single group analyzed in the previous section. As a consequence, the following holds.[23]

"serve" the same segment of the population (i.e., no individual can belong to two groups at the same time).

22. When the utility function admits multiple interior equilibria in P_B^*, the analysis is complicated by the fact that the quantitative impacts of an increase in heterogeneity for two groups with different P_B^*'s will not be the same. An extensive treatment of this case would not add much to our analysis.

23. Notice that the existence conditions for the multiple-group case are a straightforward extension of the one-group model. In fact, as shown in the proof of Proposition 1, the existence of an equilibrium proportion P_B^* depends on the functional forms of $f_\alpha(\cdot)$, $f_l(\cdot)$, $u(\cdot)$, and $v(\cdot)$, which are the same in the single- and in the multiple-group case. We just have to modify the condition that the minimum size is met in equilibrium with the following one: $\Pi(P_W^*, \bar{u}) \cdot B^k + \Pi(P_B^*, \bar{u}) \cdot W^k > (B + W)/(N + 1)$, $k = 1, \ldots, n \leq N$, where $B^k(W^k)$ is the mass of $B(W)$ types in the segment $2l^k$ identified by condition $v(l^k) = \bar{u}$ for disjoint groups, and by $2l^k = 1/n$ for adjacent groups.

COROLLARY 3. An increase in heterogeneity decreases participation in an equilibrium with n mixed disjoint groups under the same conditions under which it does so in an equilibrium with one mixed group.

The consideration of multiple groups therefore does not alter the conclusions of the one-group model in the case in which all groups are disjoint and each of them behaves as an isolated community. What if the groups are "adjacent"? First of all, the "geographic coverage" of each group can be below the maximum identified by condition (10). Second, the members located at the "frontier" between two groups must be indifferent between joining one or the other. If we denote by $\hat{l}^k$ the distance of the farthest member of group k from its HQ, this implies that for any two neighboring groups k and $k + 1$ (with $k = 1, \ldots, n - 1$), the following holds for any member of group k with aversion parameter α^j, $j = B,W$, and distance $\hat{l}^k$:

$$u(\alpha^j, P_{-j}^{k*}) + v(\hat{l}^k) = u(\alpha^j, P_{-j}^{k+1*}) + v(\hat{l}^{k+1}). \tag{13}$$

Consider the case in which the utility function admits a unique stable interior equilibrium, $P_B^* \in (0,1)$, and let $(P_B^*, \hat{l}^k)$, $k = 1, \ldots, n$ be an equilibrium with n adjacent mixed groups. Then it must be $\hat{l}^k = \hat{l}^{k+1}$ for all $k = 1, \ldots, n - 1$. In other words, all groups are identical. For this case, we still have that an increase in heterogeneity decreases aggregate participation if it does so in the one-group model.[24] In the example presented in Appendix 1 we show analytically why this is the case.

We now turn to the case of multiple homogeneous groups, which yields the results most in contrast with the one-group model. We can have multiple homogeneous groups when the utility function gives corner equilibria like those depicted in Figure II, panel (d). In this case the impact of heterogeneity on participation depends on the relationship between the minimum size requirement and the mass of the minority type, as stated in the following proposition. Without additional conditions on the

24. Intuitively, what is different here is that the border members of each group are no longer identified by the condition $v(l^k) = \overline{u}$ (indeed, the utility of the border members with $\alpha = 0$ is now strictly greater than $\overline{u}$). This implies that the extremes of integration with respect to l in expression (3) will be different, although still independent of P_{-j}. However, the impact of heterogeneity on participation depends on how changes in w influence $\Pi(P^*_{-j}, \overline{u})$, via the impact on P^*_{-j}, and this basically depends on the shape of $u(\cdot)$. Due to the separability of the utility function, changes in the extreme distance do not affect the sign of dS/dw.

number of groups, Proposition 3 holds when there are an equal number of all-B and all-W groups. In Appendix 1, however, we develop the analysis with respect to a generic number N_B of all-B groups and N_W of all-W groups. Denote by $\Delta S/\Delta(1-w)$ the change in aggregate participation due to a change in the fraction of B types in the population.

PROPOSITION 3. Let $w > ½$. For any stable equilibrium in disjoint homogeneous groups,

$$\frac{\Delta S}{\Delta(1-w)} < 0 \qquad \text{for } 2v^{-1}(\bar{u})(1-w) < \frac{1}{n+1}$$

$$\frac{\Delta S}{\Delta(1-w)} > 0 \qquad \text{for } 2v^{-1}(\bar{u})(1-w) = \frac{1}{n+1}$$

$$\frac{\Delta S}{\Delta(1-w)} = 0 \qquad \text{for } 2v^{-1}(\bar{u})(1-w) > \frac{1}{n+1}.$$

The intuition underlying Proposition 3 is the following. In an equilibrium with disjoint homogeneous groups, all individuals within distance $v^{-1}(\bar{u})$ from the HQ of a group will join it, regardless of their aversion parameter α. Starting from a situation where type W is an overwhelming majority and homogeneous groups type B cannot meet the minimum size requirement, an increase in heterogeneity decreases aggregate participation because it reduces the mass of the all-W groups without inducing participation by B. At some point, however, the increase in $(1-w)$ makes the minimum size feasible for B types, so that the decrease in the size of existing W groups is more than compensated by the formation of new B groups, and there is a discrete increase in aggregate participation. Beyond this "switch" point, when equal numbers of all-B and all-W groups exist, any increase in heterogeneity leaves aggregate participation unchanged because the decrease in the size of W groups equals the increase in that of B groups.

In summary, many combinations are possible depending on the number of groups and on the adjacent versus disjoint nature of the groups. However, the main point is that compared with mixed-group equilibria in which the effect of heterogeneity on participation is "smooth" and (under mild sufficient conditions) negative, the case of multiple homogeneous groups yields a discontinuous effect, whereby increased heterogeneity can actually *increase* participation.

II.5. Discussion

We have shown, thus far, that an increase in heterogeneity can reduce participation, especially in the presence of mixed groups. Before moving to the empirical evidence, it is worth mentioning two points. First, the empirical literature on political participation suggests that, after controlling for socioeconomic status, blacks have a higher propensity to participate in groups and to vote than whites.[25] In our empirical analysis below, we also find the same result. The explanation offered by political scientists is that blacks are more conscious of being a minority and have an extra incentive to engage in political action to preserve their identity and foster their political and civil rights. In our model this could be accommodated by assuming that $\alpha^i = \alpha^i(W/B)$, with $i = B,W$. That is, the propensity to participate is a function of the distribution of types in the population, with $\partial\alpha^B/\partial(W/B) > 0$ and $\partial\alpha^W/\partial(W/B) < 0$. Therefore, if $B < W$, then we must have $\alpha^B > \alpha^W$. This implies that after controlling for all other determinants of participation, one should obtain the empirical finding that blacks participate more, since they are a minority.

Second, thus far we have focused on differences across types not based on income. In the empirical analysis we are also interested in the effect of an increase in income inequality on participation. A vast literature in local public finance addresses the issue of group formation and income levels.[26] Our model is not a contribution game. Differences in income matter only to the extent that they are correlated with preferences and culture. In this case our formalization could be reinterpreted in terms of income rather than race; i.e., individuals would prefer to participate in social activities with people from their own income bracket. A complication, however, is that income is a "continuous" variable, while race is much less so: in modeling income dispersion as a dispersion of "types," we are therefore simplifying the analysis.

25. See, for instance, Verba and Nie [1987].

26. See, for instance, Epple and Romer [1991]. Particularly related to our analysis is the work by Fernandez and Rogerson [1996] and La Ferrara [2000], since both examine the effects of changes in income inequality.

III. Empirical Strategy and Data

In the remaining part of the paper, our aim is to focus on the link between individual aversion to different types and the decision to join groups, and to estimate the impact of increased heterogeneity in the community on participation. For our basic specification, we assume that at any point in time the "latent variable" measuring the expected utility from participation in a group for individual i in community c can be modeled as

$$Y_{ic}^{*} = X_{ic}\beta + H_c\gamma + S_c\delta + T\lambda + \epsilon_{ic}, \tag{14}$$

where X_{ic} is a vector of individual characteristics; H_c is a vector of community variables (including heterogeneity), S_c is a dummy for the state where the individual lives, T is a year dummy, and ϵ_{ic} is an error term normally distributed with mean 0 and variance σ_c. The vectors β, γ, δ, and λ are parameters. We do not observe the latent variable Y_{ic}^{*} but only the choice made by the individual, which takes value 1 (participate in a group) if Y_{ic}^{*} is positive, and 0 (not participate) otherwise,

$$\begin{aligned} P_{ic} &= 1 \qquad \text{if } Y_{ic}^{*} > 0 \\ P_{ic} &= 0 \qquad \text{if } Y_{ic}^{*} \leq 0. \end{aligned} \tag{15}$$

We estimate the Probit model (14)–(15) using individual level data and taking Metropolitan Sampling Areas (MSA) and Primary Metropolitan Sampling Areas (PMSA) as "community" dimension. We are especially interested in the vector of coefficients γ, although many of the components of β will also be important to gain insights into the determinants of participation.

The main source of data for our regressions is the General Social Survey (from now on, GSS) for the years 1974–1994. This survey interviews approximately 1500 individuals every year from a nationally representative sample, and contains information on a variety of sociopolitical indicators, as well as on demographic and income characteristics of the respondents.[27] In particular, the questionnaire includes questions regarding the

27. Note that the survey was not conducted in 1979, 1981, and 1992. Moreover, in 1982 and 1987 black individuals were oversampled; therefore, in our regressions we will use the weights provided by the cumulative GSS file 1972–1994 to correct for this oversampling. The original 1972–1994 cumulative file contains nominal income data for all years and real income up to 1993. In order to maximize the number of observations, we have constructed real income figures for 1994 following the same procedure that had been used for the previous years,

respondents' membership in organizations such as political groups, religious groups, unions, school associations, service groups, fraternities, sports and hobby clubs, etc. We use the answers to these questions to construct our dependent variables. The GSS also contains information about individual attitudes toward race relations and racial mixing. This will allow us to construct proxies for our parameter of "participation aversion" (α) in order to test the implications of our model.[28]

Among the explanatory variables we include individual controls taken from the GSS, as well as community variables capturing heterogeneity in race, ethnicity, and income in the place where the individual lives. All variables are described in Appendix 2. The remainder of this section illustrates our procedure for constructing community level variables.

It is possible to match approximately two-thirds of the respondents from the GSS 1972–1994 with the MSA/PMSA where they live.[29] We have used Census data to build community level variables, adopting the MSA/PMSA as our geographic notion of "community." Our measure of income inequality is the Gini coefficient for the MSA/PMSA computed using family income figures from the 1970, 1980, and 1990 Censuses. The values of *Gini* for the remaining years were obtained by linear interpolation and extrapolation. Our results are not sensitive to the interpolation procedure. Moreover, we computed Gini coefficients at the state level from the Current Population Survey (CPS) every year between 1974 and 1994, and the correlation between the CPS and the Census interpolated Gini's was .65. The state level Gini's from the CPS are those we use in Figures IV, VII, and VIII.

Our racial fragmentation index (*Race*) is constructed from the Census 1990 according to the following formula:

$$Race_i = 1 - \sum_k s_{ki}^2, \tag{16}$$

where i represents a given MSA/PMSA and k the following races: i) White; ii) Black; iii) American Indian, Eskimo, Aleutian; iv)

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which is described in Ligon [1989]. For more detailed information about the GSS, the reader is referred to Davis and Smith [1994].

28. In our model, traveling costs are also important. The GSS data, however, do not allow us to know the distance between individuals' residence and the location of the group's headquarters.

29. For about 40 percent of the individuals who *can* be matched with their MSA/PMSA, however, the membership data are missing. This is why the number of observations in our regressions will be smaller than the full GSS sample.

Asian, Pacific Islander; v) other. Each term s_{ki} is the share of race k in the population of MSA/PMSA i. The index (16) measures the probability that two randomly drawn individuals in area i belong to different races. Therefore, higher values of the index represent more racial fragmentation.[30]

The ethnic fragmentation index (*Ethnic*) is computed by a formula analogous to (16), using *ancestry* instead of race. In other words, s_{ki} in the formula now represents the fraction of people in area i whose first ancestry is type k. The original ancestries reported by the 1990 Census (35 categories) have been aggregated into 10 different groups on the basis of common language, culture, and geographic proximity (see Appendix 2 for a precise definition). We have chosen to aggregate these data in order not to give the same "weight" in the definition of *Ethnic* to very similar countries of origin, say Norway and Sweden, and two very different ones, say India and Ireland. Our results are not unduly sensitive to reasonable changes in our aggregation rules.

Note that we use the values of *Race* and *Ethnic* in 1990 for the whole sample. Our reasons for not interpolating are twofold. First, we believe that racial and ethnic fragmentation within MSAs are sensibly more stable over time than, say, income inequality. Second, and most importantly, in order to get variation over time, we should have resorted to the 1970 and 1980 Censuses, which contained fewer categories. For example, all Censuses before 1990 distinguished only three races: white, black, and other. Relying on years earlier than 1990 would thus have meant sacrificing the precision of our heterogeneity measures to a considerable extent. We felt that the loss in explanatory power due to this oversimplification outweighed the potential gain from time variation of the above indexes. Hence we chose to adopt the 1990 measures as our best proxies for racial and ethnic fragmentation.

Before proceeding, we need to justify our choice of MSA/PMSA as our geographical units. To begin with, we cannot use a smaller unit such as the PUMA because of a lack of respondent identifiers at this level in the GSS. However, we feel comfortable with MSA/PMSA for several reasons. First, one may argue that by

30. The Census did not identify "Hispanic" as a separate racial category. However, Alesina, Baqir, and Easterly [1999], who use the same measure of racial fragmentation, note that the category "Hispanic" (which they obtain from a different source) has a correlation of more than 0.9 with the category "other" in the Census data. Thus, for all practical purposes, the category "other" in the Census is virtually a measure of the Hispanic population.

using a relatively large geographical unit one may bias our regressions against finding an effect of fragmentation on participation. Second for several of the groups that we consider, it is reasonable to assume that direct interaction among group members occurs at the MSA level (think of unions, sport groups, and boy scouts, for instance). Third, and most importantly, we checked the correlation between our measures of fragmentation at the MSA/PMSA and at the PUMA level. The results are comforting. The correlation for the Gini coefficient is 0.65; the correlation for *Race* is 0.75; and the one for *Ethnic* is 0.75.[31] Finally, we have checked that a large proportion of the respondents in the GSS lives in nonsegregated areas, and therefore potentially interacts with the opposite race. It turns out that 71 percent of white respondents say that "there are black families living close" to them, and 48 percent of white respondents say that black families live "on this block, a few doors/houses away."

IV. Descriptive Results

We begin by presenting summary statistics and a few simple correlations among membership rates and our measures of heterogeneity. Summary statistics and definitions of the full set of variables are in Appendix 2.

The top panel of Table I shows the sample characteristics of some of our data. Participation rates are on average very high: overall, 71 percent of the respondents are members of at least one group, the average number of group memberships being 1.8 per person. Also, there is considerable variation in participation rates, both across individuals and across groups: the standard deviation of our basic membership variable is 0.45. The fraction of participants in the various groups ranges from 0.02 for farmers' associations to 0.34 for religious groups. Sport groups are the second most popular category, with a participation rate of 0.20, followed by professional associations (0.17), unions (0.15), and school service groups (0.14). Literary groups, hobby clubs, fraternities, and service groups (Rotary, Lions, etc.) have participation rates of 0.09–0.10. Most notable is the low enrollment in political associations: only 5 percent of the respondents are members of a political

31. These correlations are weighted by the population share of each PUMA in the MSA/PMSA.

TABLE I
DESCRIPTIVE STATISTICS

Panel A: Summary statistics

	Mean	Std. dev.
Member of any group	.71	.45
Member of church group	.34	.47
Member of fraternity	.09	.29
Member of service group	.10	.30
Member of hobby club	.10	.29
Member of sport club	.20	.40
Member of youth group	.10	.30
Member of literary group	.10	.30
Member of school service group	.14	.35
Member of school fraternity	.05	.23
Member of veterans' group	.06	.24
Member of political group	.05	.21
Member of nationality group	.04	.20
Member of union	.15	.36
Member of professional association	.17	.37
Member of farmers' group	.02	.14
Member of other group	.11	.31
Gini	.41	.03
Racial fragmentation (*Race*)	.36	.15
Ethnic fragmentation (*Ethnic*)	.67	.07

Panel B: Correlations

	Membership	*Gini*	*Race*
Gini	−.06**		
Race	−.05**	.34**	
Ethnic	−.04**	.09	.56**

*Denotes significance at the 5 percent level; **at the 1 percent level.

group. Nationality groups, which we will consider in more detail below, are joined by about 4 percent of the respondents.

The last three variables in Panel A of the table are measures of heterogeneity in income, race, and ethnicity in the MSAs where the respondents live. The mean of the Gini coefficient is 0.41, with a standard deviation of 0.03. Our racial fragmentation index has a mean of 0.36 (standard deviation 0.15), while ethnic fragmentation is higher at 0.67 (standard deviation 0.07). The correlation among these three measures of heterogeneity is quite high, as shown in the bottom panel of Table I: *Gini* is correlated 0.34 with

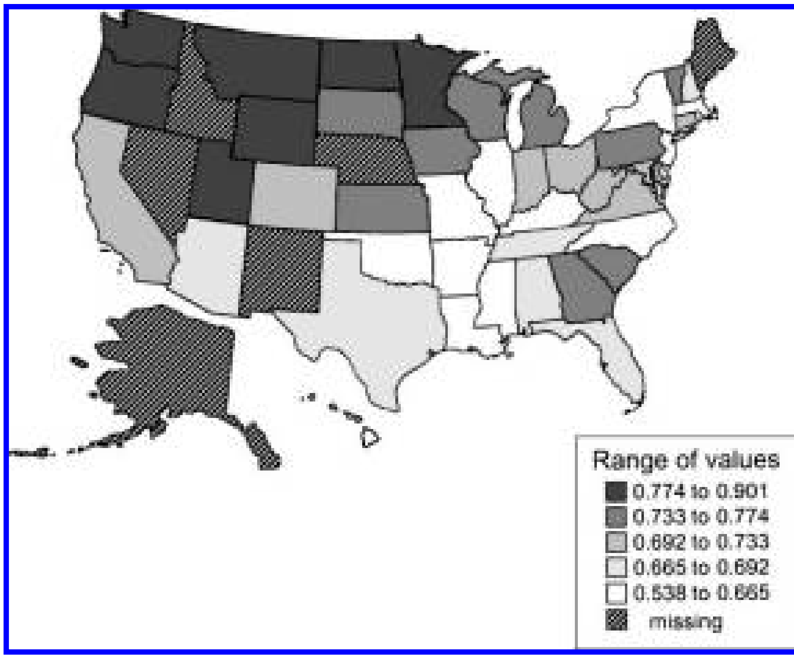

FIGURE III
Average Membership Rate, 1973–1994

Race and 0.09 with *Ethnic;* the correlation between *Race* and *Ethnic* is 0.56.[32] On the other hand, the simple correlation between average membership in the MSA and the Gini index is not quantitatively significant; this will no longer be true when we turn to multivariate analysis.

Figures III, IV, V, and VI illustrate the geographic distribution of our variables of interest by reporting sample averages at the state level.

Figure III shows the distribution of participation rates from the GSS data set, that is, the percentage of respondents in each

32. We also explored the correlation among our heterogeneity variables and the measures of racial or income segregation of Cutler, Glaeser, and Vigdor [1999]. It turns out that segregation by income is positively correlated with income inequality, and that the various measures of racial segregation are positively correlated with racial and ethnic fragmentation. This may suggest that segregation is "valued" relatively highly in places where heterogeneity is higher, an observation consistent with the argument that individuals prefer contact with people similar to themselves.

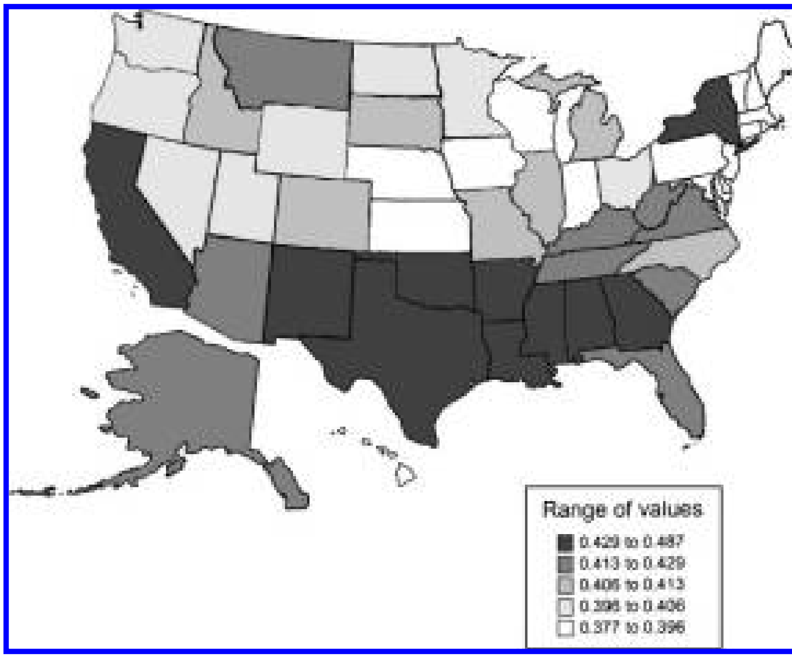

FIGURE IV
Gini Coefficient, 1972–1994

state who are members of at least one group (average from 1974 to 1994). As one can see, this percentage is highest in states of the north central and northwest regions, and lowest in the south and southeast. Figures IV, V, and VI show the distribution of the Gini coefficient (average 1974–1994 from CPS data), of Racial fragmentation, and of Ethnic fragmentation (both measured in 1990), all calculated at the state level. These maps show a rather striking pattern, when compared with Figure III: racial and ethnic fragmentation, as well as income inequality, are highest in the southeast and lowest in the northeast; i.e., those regions where participation is, respectively, lowest and highest.

Similar implications can be gathered from the three panels of Figure VII, which plot state level participation rates against the Gini index, Race fragmentation, and Ethnic fragmentation. In all three panels a negative correlation between membership in groups and heterogeneity clearly emerges.

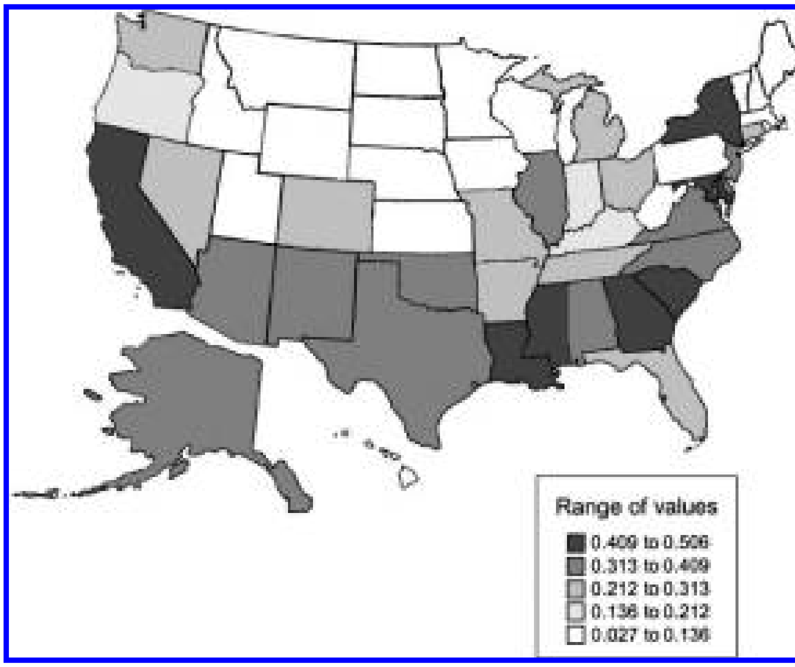

FIGURE V
Racial Fragmentation, 1990

V. THE ECONOMETRIC EVIDENCE

V.1. Basic Regressions

Table II displays our basic probit regression using the GSS data set and including only individual controls. The dependent variable takes the value 1 if the respondent belongs to at least one group, and 0 otherwise. The regressors include a set of individual characteristics that, in our model, may influence either the individual's reservation utility if not participating, $\bar{u}$, or the preference for participation captured by the parameter α.[33]

33. The political science literature has generally looked at these individual determinants of participation in isolation, i.e., correlating one or two variables at a time with membership rates. See, for instance, Verba and Nie [1987] and Verba, Schlozman, and Brady [1995]. A multivariate analysis with demographic controls similar to those we include is in Glaeser and Glendon [1997], but their dependent variable is church attendance rather than group participation.

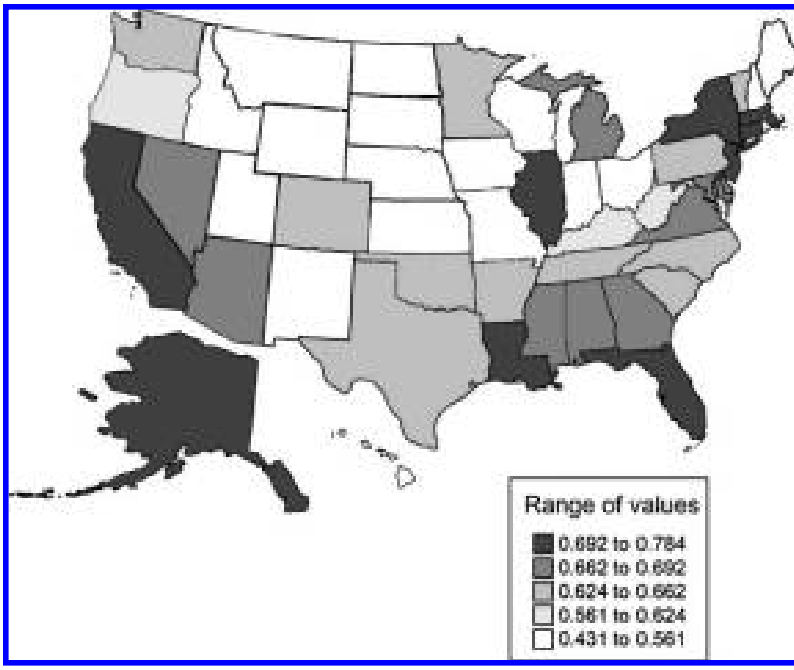

FIGURE VI
Ethnic Fragmentation, 1990

The estimates in the first column of Table II are marginal probit coefficients evaluated at the means; in the second column we report heteroskedasticity-corrected standard errors adjusted for intra-MSA clustering of the residuals. First of all, the cohort variable suggests a decline in participation by younger cohorts. Second, the age distribution variables show a dip in participation for individuals in their thirties. Child-rearing activities reduce the time available for participation: in fact, the coefficient on the variable that captures whether the respondent has children below the age of five is negative and significant. Note that the dummy for age group 30–39 and that for children below age 5 are highly correlated. More generally, both variables capture a period of individual lifetime that is particularly "busy" because of marriage, having children, setting up new households, etc. There is some weak evidence that older people participate more, probably because they have more time if they are retired, although health

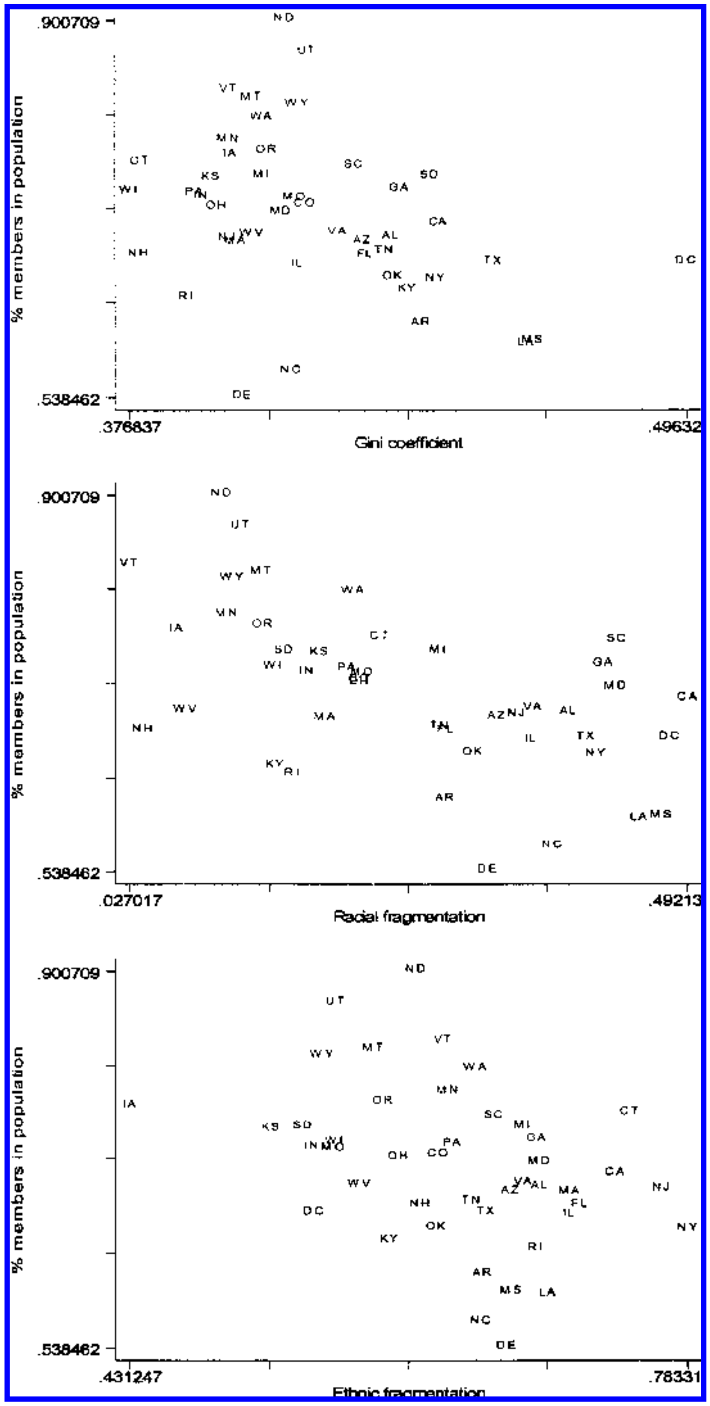

FIGURE VII
Heterogeneity and Participation in Groups

TABLE II
INDIVIDUAL DETERMINANTS OF PARTICIPATION

	Marg. Probit coeff.[a]	Std. error[b]
Cohort	−.002*	(.001)
Age <30	−.035	(.028)
Age 30–39	−.029*	(.018)
Age 50–59	.004	(.016)
Age ≥60	.035	(.027)
Married	−.001	(.011)
Female	−.046**	(.011)
Black	.045**	(.011)
Educ <12 yrs	−.122**	(.013)
Educ >16 yrs	.144**	(.012)
Children ≤5 yrs	−.035**	(.014)
Children 6–12	.071**	(.012)
Children 13–17	−.005	(.014)
ln (real income)	.074**	(.006)
Full-time	.025**	(.011)
Part-time	.067**	(.016)
STATES	Yes	
YEARS	Yes	
No. obs.	10534	
Pseudo R^2	.09	
Observed P	.72	
Predicted P	.74	

*Denotes significance at the 10 percent level; **at the 5 percent level.
a. Marginal probit coefficients are calculated at the means.
b. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.

considerations may work the other way. This result, together with the cohort effect, accounts for the notion of "older civic generation" emphasized by Putnam [1995a, 1995b].

Years of schooling are positively associated with participation: high school dropouts participate significantly less, while college graduates significantly more. The coefficients on the education variables remain highly significant and stable throughout all specifications. Among the possible explanations for this strong association, Verba and Nie [1987] suggest that more education is generally combined with a higher evaluation of one's own ability to influence sociopolitical outcomes, and with a higher level of social interaction. In Table II we also see that women participate significantly less than men. Our interpretation is that they often carry the weight of a job plus a preponderant share of household chores and child-rearing activities: this heavy load leaves women with less time for leisure and participation. We checked, in fact, that women do *not* participate less in voting, an act of participation which does not require a significant amount of time. The probability of being a member of a group is increasing in family income of the respondent, suggesting that participation is a "normal good." We investigate in more detail the nonlinear effects of income on participation in Section V.

Consider now the time spent at work. The omitted category captures people who are not working, including homemakers, retirees, students, and unemployed. After controlling for the level of income, the effect of time spent at work could be twofold. On the one hand, a constraint on time may decrease participation; on the other hand, socialization in the workplace may increase social interaction, incentives, and ability to participate. Our results on this point are consistent with basic economic principles. The coefficient on part-time workers is larger (and significantly so) than the one on full-time workers. This suggests that, even though socialization in the workplace helps (in fact, full-time workers participate more than those out of the labor force), the time constraint is binding for people who work full time. It is less binding for part-time workers who, on the other hand, still get the benefits of social interactions in the workplace.[34] When we control for all these variables, marital status does not seem to affect participation (contrary to the common notion in sociological and political analyses based on partial correlations, which indicate that married people participate significantly more).

For our purposes, a particularly interesting variable is "black." As we can see from Table II, ceteris paribus, members of this racial group participate significantly more, a result which emerges clearly when we control for the other individual determinants of participation but which is obscured if we only look at partial correlations. Note that this result is not driven by the higher church attendance of blacks in the south: in fact, it survives if churches are left out of the definition of groups and if the south is omitted from the regression. As discussed above, our model could be extended to incorporate a feature of group consciousness, in which the minority type participates more to preserve identity

34. This interpretation is confirmed by additional sensitivity analysis. Interestingly, the unemployed participate less even after controlling for income.

and to defend its role in the community. More importantly, since blacks are a minority in virtually all MSAs, the percentage of black residents is positively associated with racial fragmentation. The result on black propensity to participate implies that if we find that participation is *lower* in more racially fragmented communities, this result is not due to the positive correlation between percentage of blacks and racial fragmentation. On the contrary, the fact that blacks participate more works *against* finding a significant effect of heterogeneity on participation.

Regressors not shown include year dummies and state dummies. The pattern of year dummies is broadly consistent with the declining trend in participation rates, already partly captured by the variable cohort. Many of the state dummies are statistically significant, indicating a need to include them.

The coefficients on individual controls are very stable and robust to different specifications.[35] Therefore, to economize on space from now on, we will not report them, although it should be kept in mind that they have always been included in the regressions, together with the state and year dummies. We next extend our analysis by incorporating variables that capture the characteristics of the community where the respondent lives.

In Table III we include the size of the place where the individual lives, the median income level in the MSA and its square (all in logs), together with our measures of heterogeneity. Size has a negative but not significant coefficient, while the coefficients on the income variables indicate that richer communities participate more but at a decreasing rate. Finally, we move to the characteristics of communities which are the focus of the present paper.

The first measure of heterogeneity included in column (1) is income inequality. The coefficient on *Gini* is negative and significant at the 1 percent level, indicating that *people living in more unequal communities are less likely to join groups.* Column (2) includes our measure of racial fragmentation, which also has a negative and significant coefficient: *individuals living in more*

35. We do not include home ownership among our regressors because this would restrict our sample to 3101 observations only. When included in the regression, home ownership has a positive and significant coefficient, consistent with the findings of DiPasquale and Glaeser [1999].

TABLE III
HETEROGENEITY AND PARTICIPATION

	(1)	(2)	(3)	(4)	(5)	(6)
Size of place	−.003	−.002	−.003	−.002	−.002	−.002
	(.002)	(.002)	(.002)	(.002)	(.002)	(.002)
Med HH income	4.379**	5.733**	4.587**	5.230**	4.259**	4.994**
	(1.491)	(1.688)	(1.620)	(1.688)	(1.577)	(1.740)
Med HH inc. ∧2	−.220**	−.282**	−.227**	−.259**	−.213**	−.248**
	(.071)	(.080)	(.077)	(.080)	(.075)	(.082)
Gini	−.916**			−.443**	−.773**	−.468*
	(.238)			(.249)	(.233)	(.261)
Race		−.201**		−.141**		−.112
		(.057)		(.068)		(.071)
Ethnic			−.253**		−.161*	−.082
			(.106)		(.092)	(.089)
INDIV CONTROLS[a]	Yes	Yes	Yes	Yes	Yes	Yes
STATES	Yes	Yes	Yes	Yes	Yes	Yes
YEARS	Yes	Yes	Yes	Yes	Yes	Yes
No. obs.	10534	10534	10534	10534	10534	10534
Pseudo R^2	.09	.09	.09	.09	.09	.09
Observed P	.72	.72	.72	.72	.72	.72
Predicted P	.74	.74	.74	.74	.74	.74

*Denotes significance at the 10 percent level; **at the 5 percent level.

Marginal probit coefficients are calculated at the means. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.

a. Individual controls: all those listed in Table II.

racially fragmented areas participate less. In column (3) "racial" fragmentation is replaced by "ethnic" fragmentation as measured through the ancestry data. Again, the negative and significant coefficient on this variable suggests that *participation is lower in more ethnically fragmented communities.*

In the last four columns of Table III we introduce in the same regression both inequality and our measures of racial or ethnic heterogeneity. *Gini* and racial fragmentation remain significant when introduced jointly; however, the absolute values of their coefficients fall due to the positive correlation among the two variables, highlighted in Table I. Similar results obtain in column (5), when we introduce inequality and ethnic fragmentation together. Finally, both *Race* and *Ethnic* lose significance in the last column, when they are included in the same regression with *Gini*. Given the high degree of correlation between the three indexes of

TABLE IV
SENSITIVITY ANALYSIS

	Dependent variable: Member Excl. influential observations[b]			Dependent variable: Member (excl. nationality groups)	
	(1)	(2)	(3)	(4)	(5)
Gini	−1.947**				
	(.742)				
Race		−.481**		−.212**	
		(.145)		(.057)	
Ethnic			−.747**		−.285**
			.286		(.104)
INDIV CONTROLS[a]	Yes	Yes	Yes	Yes	Yes
STATES	Yes	Yes	Yes	Yes	Yes
YEARS	Yes	Yes	Yes	Yes	Yes
No. obs.	9935	9888	9983	10425	10425
Pseudo R^2	.11	.13	.12	.09	.09
Observed P	.75	.75	.74	.72	.72
Predicted P	.77	.79	.77	.74	.74

*Denotes significance at the 10 percent level; **at the 5 percent level.

Marginal probit coefficients are calculated at the means. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.

a. Individual controls: all those listed in Table II.

b. Influential observations identified by predicting DFbetas for the relevant variable from the full sample regression and then dropping those observations for which $abs\,(DFbeta) > 2/\sqrt{\#obs}$.

heterogeneity, in what follows, we present results from regressions where the above measures are introduced one at a time.[36]

V.2. Sensitivity Analysis and Causality

In Table IV we conduct a sensitivity analysis by controlling for influential observations whose presence would sensibly bias our estimates. We do this by calculating the DFbetas from each original regression, and dropping those observations that lead to major changes in the coefficients of our heterogeneity measures.[37]

The results in Table IV are even stronger than those in Table III, in the sense that the coefficients on *Gini*, *Race*, and *Ethnic* are larger in absolute value. Using the estimated coefficient in column

36. When we introduced them jointly, the results we obtained were broadly consistent; but in some instances some of the three would lose statistical significance at standard levels of confidence.

37. Specifically, we dropped those observations for which $abs(DFbeta) > 2/\sqrt{\#obs}$ (see, e.g., Belsley, Kuh, and Welsch [1980, p. 28]).

(2), we calculate that, starting from the sample mean, an increase in *Race* by one standard deviation leads to a reduction in the probability of participation of eight percentage points. This is quite a sizable effect, if we compare it with the impact of other significant determinants of participation. For example, having a child below the age of five reduces the propensity to participate by about 3.5 percentage points: living in a community that is one standard deviation above the mean in racial fragmentation reduces the probability of participating by more than twice as much. Moving from a full-time to a part-time job increases the propensity to participate by four percentage points, that is half the impact of an increase of one standard deviation in *Race*. Take as another example, education. Ceteris paribus, going from the status of high school dropout to high school graduate or higher increases the probability of being a member of a group by about thirteen percentage points. An increase of one standard deviation in racial fragmentation has about two-thirds of this effect.

Inequality and ethnic fragmentation also have sizable coefficients. From column (1), starting at the sample mean, an increase in *Gini* by one standard deviation leads to a reduction in the probability of participation of six percentage points. This is almost twice as much as the effect of having a small child. Compared with the effect of education (an increase in probability of thirteen percentage points when going from high school dropout to high school graduate), moving from a community that is one standard deviation *above* the mean for inequality to one that is one standard deviation *below*, increases the probability of participation by about the same amount. Similarly for ethnic fragmentation (column (3)): an increase of one standard deviation above the mean reduces the probability of participating in a group by six percentage points.

Finally, it is worth noting that one of the categories included in our dependent variable is "nationality groups." One should expect that for this type of group racial and especially ethnic fragmentation should not have a negative effect, but rather a *positive* effect (see the next section). We ran the same regressions of Table III excluding nationality groups. The estimated coefficients on *Race* and *Ethnic* when the dependent variable is membership in any group other than a nationality group are reported in columns (4) and (5) of Table IV. Compared with the estimates of Table III (columns (2) and (3), respectively), the impact of racial and ethnic heterogeneity is now quantitatively more important, as expected.[38]

To explore other dimensions of heterogeneity, we have also considered the effect of heterogeneity in the age of the population living in the community. We built an age fragmentation index similar to that used for racial and ethnic fragmentation, using a variety of age breakdowns. None of the age fragmentation variables was significant when added to our basic specification (the sign was generally negative as expected), while *Gini*, *Race*, and *Ethnic* remained unaffected.[39]

Next we considered nonlinear effects of individual income on participation. This is important in order to check that our results on inequality are not a statistical artifact. In fact, if the relationship between individual income and participation were concave and we omitted nonlinear income terms from the participation regression, moving from a more equal to a less equal distribution of income would automatically reduce participation even if inequality per se were not a determinant of participation.[40] Appendix 4 reports our results on nonlinearities in income when we add to the basic specification a quadratic and a cubic term in income (possibly with a poverty dummy) together with our three measures of heterogeneity. The coefficients of *Gini*, as well as those of *Race* and *Ethnic*, remain negative and highly significant. The values of the coefficients on individual income suggest that the relationship between participation and income is increasing and is convex for low levels of income, and concave at high levels.

An additional test we performed was with regard to the variable *Black*. As discussed above, ceteris paribus, blacks participate more. We have tested whether when racial heterogeneity increases white participation falls more than that of blacks. When we introduce an interaction term of black and racial fragmentation, the coefficient has the expected sign, namely positive, but it

38. Analogous results were obtained on the sample purged of influential observations (i.e., on the sample comparable to columns (2) and (3) of Table IV).

39. Results on this point are available. The fragmentation index was constructed according to (16), where each term s_{ki} was the share of people in age category k in MSA/PMSA i, computed from Census data. For the age categories we experimented with different degrees of aggregation, e.g., years 0–14, 15–24, 25–34, 35–49, 50–64, 65 and above (relatively disaggregate), as well as 0–14, 15–34, 35–64, 65 and above (relatively aggregate).

40. This point was raised in the context of the effects of income inequality on health outcomes, and it is addressed in Gravelle [1998].

is not statistically significant at conventional levels.[41] We performed a variant of this test interacting the race of the respondent (black, white, or other, in the GSS) with the share of population in the MSA/PMSA belonging to that race (from Census data). The interaction term for blacks was positive and significant.[42]

We next consider the issue of potential endogeneity of *Gini*. A high degree of participation may reduce income inequality by increasing availability of information, options, and opportunities. Also, communities prone to social activities may be more favorable to redistributive policies. These problems are much less important for measures of ethnic or racial fragmentation: therefore, in Table V we concentrate on instrumenting for Gini.

We consider three instruments: the number of municipal and township governments in 1962, the percentage of fiscal revenues from intergovernmental transfers in 1962, and the share of the labor force employed in manufacturing. The number of governments in 1962 can safely be considered exogenous to participation in 1974–1994, and it can have influenced the degree of income inequality in the MSA. Within a metropolitan area that was fragmented into many smaller jurisdictions, it is more likely that significant differences in policies, local public good provision, and income levels persist among those jurisdictions. The amount of fiscal resources obtained from higher levels of government may have influenced inequality in the MSA/PMSA. However, fiscal transfers may target unequal MSAs, and, to the extent that inequality is serially correlated, this instrument may be imperfect. The share of manufacturing is certainly not exogenous to union participation, so when we use this instrument, we exclude participation in unions from our dependent variable. To a lesser extent, the share of labor force employed in manufacturing may not be exogenous to participation in other groups as well. In

41. According to the model, the majority type (whites) should drop out proportionally more as heterogeneity increases. The lack of significance of the interaction term may be related to the fact that the variable black alone is highly significant. Note that if the variable black is dropped the interaction is highly significant and positive. Also, in evaluating the empirical test for the "magnification effect," it should be kept in mind that black is not the only minority type. All these results are available.

42. Extending the same logic to religious categories, we regressed membership in church groups on the interaction terms between the respondent's denomination and the share of people belonging to that denomination in the state. Our categories were Protestant, Catholic, Jew, Other religion, and nonreligious (omitted). We found positive and significant coefficients on most categories. A similar test for religious attendance is reported by Glaeser and Glendon [1997].

TABLE V
INSTRUMENTING GINI

	Dependent variable: Member			Dependent variable: Member (excl. unions)	
		2SLS Instrument set:			2SLS Instrument set:
	OLS	NGOV62	NGOV62 REVIG62	OLS	NGOV62 MANSHR
	(1)	(2)	(3)	(4)	(5)
Gini	−.858**	−2.130**	−2.252**	−1.027**	−1.730*
	(.240)	(.742)	(.730)	(.241)	(.752)
INDIV CONTROLS[a]	Yes	Yes	Yes	Yes	Yes
STATES	Yes	Yes	Yes	Yes	Yes
YEARS	Yes	Yes	Yes	Yes	Yes
No. obs.	10333	10333	10333	10243	10243
R^2	.11	.10	.10	.11	.11
Hausman (p-value)		.01	.03		.01
Sargan (p-value)			.81		.70

*Denotes significance at the 10 percent level; **at the 5 percent level.
Marginal probit coefficients are calculated at the means. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.
a. Individual controls: all those listed in Table II.

summary, we are most comfortable with our first instrument (number of governments in 1962) relative to the other two, but we present results using various combinations of the three.

In column (1) of Table V we report estimates of the linear probability model for the sake of comparison. The coefficients of interest and goodness of fit measures from the first-stage regressions are reported in Appendix 5.[43] In columns (2) and (3) we report 2SLS estimates when we instrument *Gini* with the number of governments in 1962, and with the same variable plus the share of transfers received by higher levels of governments. *Gini* remains highly significant in both cases. In column (5) we use as instruments the employment share in manufacturing together with the number of governments in 1962. Since union participation is certainly not exogenous to the share of manufacturing, in columns (4) and (5) we exclude unions from our membership

43. The predicted value of Gini from the regressions in Appendix 5 was substituted in the linear probability model to obtain the estimates in Table V (correcting the Standard errors).

dependent variable. Once again, the 2SLS coefficient on *Gini* remains significant and is higher in absolute value than the OLS one (column (4)).

We should pause to analyze the fact that the coefficient on inequality in the 2SLS regressions is larger in absolute value than in the linear probability model. Suppose that those individuals who are more likely to participate in social activities are also more favorable to redistribution, thus reducing inequality. This would imply an upward bias (in absolute value) of the estimated coefficient on Gini in the participation regression. An alternative argument is that individuals who are more prone to participate are those who are less averse to mixing with individuals with a different income level. This is in fact the basic idea of our model. But then individuals less averse to income heterogeneity may also be more prone to live in communities with more income heterogeneity. This would imply a downward bias (in absolute value) of the OLS coefficient on *Gini*. The patterns of the coefficients in Table V seem to support the second interpretation.[44]

V.3. Types of Groups

The groups included in the GSS questionnaire are quite diverse, ranging from unions to literary clubs to church groups. It is therefore instructive to analyze participation in each of them separately to see whether heterogeneity plays a different role in different types of groups. This is done in Table VI.

We have run the same regressions of Table III, using as dependent variable individual membership in a given type of group. Each cell in the table refers to a separate regression and shows the marginal probit coefficient on the variable listed by column (*Gini*, *Race*, or *Ethnic*) for the type of group listed by row.[45]

44. Finally, we have explored the effects of our heterogeneity variables using both a fixed-effect and a random-effect model. There are several difficulties in pursuing this strategy. First of all, the GSS is not a panel. Second, we cannot construct time series of our racial and ethnic fragmentation variables. Thus, we can test within-MSA effects only for the Gini coefficient, for which we have a time dimension (even though some interpolation is involved, as discussed above). For all these reasons the results on within-MSA variations have to be taken *cum grano salis*. Our results are mixed. In the fixed effect model *Gini* retains a negative coefficient, although not statistically significant (t-statistic -0.81). With random effects the coefficient on *Gini* is negative (-1.17) and statistically significant (t-statistic -5.66). The Breusch-Pagan test for random effects rejects the null hypothesis with a p-value of .005.

45. All regressions include the individual and community controls listed in previous tables. We have also run the group-by-group regressions using the DFbeta method employed in Table IV. Results are very similar to those reported in Table VI.

TABLE VI
PARTICIPATION BY TYPE OF GROUP[a]

Dependent variable is membership in	Marginal Probit coefficients on:		
	Gini	*Race*	*Ethnic*
Church groups	−1.027**	−.156**	−.278**
	(.208)	(.044)	(.104)
Fraternities	−.323*	−.031	−.014
	(.185)	(.030)	(.081)
Service groups[b]	−.348**	−.057**	−.103**
	(.139)	(.028)	(.053)
Hobby clubs	−.520**	−.078**	.007
	(.131)	(.028)	(.057)
Sport clubs	−.789**	−.123**	−.038
	(.291)	(.064)	(.078)
Youth groups[c]	−.430*	−.060	−.007
	(.235)	(.043)	(.064)
Literary groups[b]	−.019	−.022	−.025
	(.193)	(.047)	(.070)
School service groups[d]	−.957**	−.112	−.136
	(.433)	(.072)	(.146)
School fraternities[b]	−.176	.003	−.026
	(.117)	(.032)	(.046)
Veterans' groups[e]	−.147	−.024	−.050
	(.114)	(.023)	(.039)
Political groups	−.271**	−.028	−.023
	(.128)	(.024)	(.048)
Nationality groups	.033	.018	.077**
	(.090)	(.017)	(.036)
Unions[f]	−.021	−.015	.107
	(.243)	(.050)	(.071)
Professional associations[g]	.129	.113	.198
	(.447)	(.104)	(.182)
Farmers' groups[h]	2.965	.465	.338
	(4.184)	(.503)	(.967)
Other groups	.032	−.050	−.077
	(.210)	(.037)	(.072)

*Denotes significance at the 10 percent level, **at the 5 percent level.

Marginal probit coefficients are calculated at the means. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.

Each cell reports the marginal probit coefficient on the variable listed in the column heading from a regression in which the dependent variable is membership in the type of group described in the row heading. All regressions include the individual controls listed in Table II, state, and year dummies.

a. The sample for each regression is restricted to those individuals who can potentially be members of that particular group.

b. Sample includes individuals with at least twelve years of education.

c. Sample includes individuals younger than 50.

d. Sample includes individuals with children age six to seventeen.

e. Sample includes cohorts 1920 to 1955.

f. Sample includes production, clerical, sales, and service workers.

g. Sample includes professional workers.

h. Sample includes only workers whose occupation code corresponds to Agriculture. Due to the small size of the sample, state and year dummies are omitted from these regressions.

For each group we have excluded from the sample those respondents who for one reasons or another *cannot* be members of a particular group. For instance, individuals below a certain age cannot be members of a veteran group, people who are not farmers cannot be members of a farmers' group, retirees cannot be members of a union, etc. The exact exclusion rules from the regression for each group are reported at the bottom of the table (the qualitative nature of our results is robust to the specification of these exclusions).

What patterns are we looking for? According to the basic ideas underlying our model, measures of heterogeneity should be less important for groups with a relatively *high degree of excludability or a low degree of close interaction* among members. The results of Table VI are broadly supportive of these hypotheses. Church groups are those with the strongest effect of all three types of heterogeneity. These are groups with very little excludability and a high degree of interaction. At the opposite extreme we have professional associations and farmers' groups, which have a very low level of personal interaction, although low excludability; the coefficients on *Gini*, *Race*, and *Ethnic* are in fact insignificant in these regressions. Service groups, hobby clubs, sports clubs, and youth clubs have a high degree of interaction and a less than perfect degree of excludability. Participation in these groups tends to be negatively influenced by heterogeneity, in particular by inequality and racial fragmentation (the coefficients on ethnic fragmentation are not statistically significant in every regression).

Interestingly, and consistent with what one would expect, Ethnic fragmentation is positively associated with membership in nationality groups. This observation suggests that in fragmented communities, individuals may feel more of a need to preserve and actively promote their own cultural identity and values, an observation broadly consistent with the spirit of our model.

Finally, notice that, aside from church groups, school service groups are the ones for which income inequality has the strongest negative impact. This is not surprising, given that high degrees of inequality are likely to be associated with marked heterogeneity in preferences for the type of education and services that schools should provide.

V.4. Individual Preferences and Participation

Our model implies that in heterogeneous communities those who choose not to participate should be individuals who are relatively more averse to mixing with different types. This implication can be investigated, since the GSS asks several questions aimed at directly identifying individual preferences and attitudes toward racial mixing.

We modify the specification in (14) as follows:

$$Y^*_{ic} = X_{ic}\beta + H_c I^A \gamma^A + H_c I^N \gamma^N + S_c \delta + T\lambda + \epsilon_{ic}, \tag{17}$$

where all variables are defined as in Section III, except that H_c is racial fragmentation in the community, I^A is a dummy equal to 1 if individual i is "averse to the opposite race" (in a way that will be defined below), and $I^N = 1 - I^A$. The coefficient γ^A therefore captures the impact of racial heterogeneity on participation for those individuals who explicitly declare their aversion to "mixed interactions," while γ^N captures the impact of heterogeneity for the respondents who are indifferent or have mild preferences on racial mixing. In Table VII we estimate this modified probit model and test the statistical difference between the two coefficients. Our theoretical framework implies that *the coefficient on Race should be larger in absolute value for the subsample of individuals relatively more averse to racial mixing.* A test along these lines is especially valuable in that it captures a variation across individuals with varying degrees of aversion *within* the same MSA/PMSA—hence it helps address potential concerns of spurious correlation between heterogeneity and participation in the cross section.

In Table VII the dependent variable is membership in a subset of groups for which heterogeneity is important. More specifically, we choose all the groups for which at least one of the three measures of heterogeneity is negative and significant at the 5 percent level in Table VI. These groups are church, hobby, sport, youth, service, political, and school service groups.[46]

The interesting thing for the measurement of individual "aversion" is that the GSS includes many questions explicitly concerning attitudes toward race relations. Some of them are "yes" or "no" questions, others range on a scale of 1 to 3 or 1 to 4. In all cases, we have created a binary variable distinguishing

46. The results of Table VII are not unduly sensitive to the choice of groups. They are robust to focusing on a smaller or larger subset of groups.

TABLE VII
PARTICIPATION AND AVERSION TO RACIAL MIXING

Dependent variable: Member	Probit coeff. on Race for those who answer: Yes (γ_A)	No (γ_N)	Test $\gamma_A = \gamma_N$ (p-value)	Fraction of Yes
(1) Have *not* had opp. race home for dinner in last few years (NOBLKDINNER)	−.436** (.078)	−.254** (.076)	.00	.65
(2) Would *not* change racist rules in club (NORACCHNG)	−.266** (.111)	−.174 (.113)	.04	.34
(3) Whites have right to segregated neighborhood (RACSEGR)	−.449** (.110)	−.279** (.077)	.04	.09
(4) Oppose your children going to school with most opposite race (NOMOSTSCHOOL)	−.325** (.093)	−.230** (.085)	.03	.41
(5) Racist has right to teach (RACTEACH)	−.345** (.080)	−.246** (.078)	.01	.43
(6) Would oppose black president[a] (NOBLKPRESID)	−.455** (.098)	−.308** (.080)	.01	.14
(7) Would favor a law against mixed marriages (NOMIXMARRIAGE)	−.419** (.092)	−.409** (.076)	.86	.20
(8) Oppose busing (NOBUSING)	−.273** (.074)	−.320** (.079)	.29	.77
(9) Think that blacks should not push[a] (BLKNOPUSH)	−.385** (.091)	−.322** (.090)	.27	.33

*Denotes significance at the 10 percent level; **at the 5 percent level.
Marginal probit coefficients calculated at the means. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.
All regressions include the individual controls listed in Table II, state, and year dummies.
a. Sample includes nonblacks only.

"averse" from "nonaverse" people, and we have reported a rough description of the criterion (or question) in rows (1) to (9) of Table VII. Details on the definition of each binary variable, as well as the exact wording of the question in the GSS, are provided in Appendix 2. Out of the many questions on race relations contained in the GSS, we have chosen those for which we had a sufficiently high number of respondents on both sides.[47] Table VII reports the marginal probit coefficients on the racial fragmentation variable from a participation regression and a test that the two coefficients

47. In practice, this means that we have excluded several questions for which the yes answer had only about 200 observations. For these cases, the sizes of the coefficients on *Race* were consistent with our hypothesis, even though the scarcity of observations made the estimation unreliable.

of interest are significantly different from each other. The last column reports the fraction of respondents who answered "yes" to the specific question.

The estimates in Table VII provide considerable support for this implication of our model. For eight out of nine questions concerning attitudes toward race relations, the effect of racial heterogeneity is strongest for the individuals more averse to racial mixing, namely the coefficient in the first column is larger in absolute value than the one in the second column. In six of these eight cases the difference in the coefficients is statistically significant, with a p-value of less than 0.05.

A particularly interesting question is the one about whether one has had a black person home for dinner in the past few years.[48] In fact, this is a question concerning individuals' actual *behavior,* as opposed to a test of a generic *attitude* toward race in an abstract sense. Therefore, it may be a better measure of individuals' true attitude toward racial mixing and interracial direct contacts, which is the essence of our model. In fact, recent results by Glaeser et al. [2000] are quite supportive of this interpretation. They find that when individuals are asked in the abstract they claim to have a large amount of "trust" toward others, but their behavior in actual experiments is much less prone to trusting others. Our results based upon this behavioral question are very strongly in favor of our hypothesis, as shown in line (1). Also, especially interesting is question (2), which explicitly addresses issues of racial mixing in groups by asking people if they would try to change the rules in a club that would not let a member of the opposite race join. For this question as well we obtain strong results. Question (3) addresses the right to live in segregated neighborhoods; again we find that those who strongly support segregation are more negatively affected by heterogeneity. Similar considerations apply to question (4), which concerns mixing of children in schools.[49] We also obtain favorable results on the question of whether racists should be allowed to teach (row

48. Until 1978 this question was asked of nonblacks only; since 1978 it was asked of all respondents in terms of "opposite race." In Table VII we report the results when the sample is restricted to nonblacks for all years. Analogous results obtain if we include all races using the years from 1978 on.

49. The GSS also includes question concerning sending children to school with "a few" children of the opposite race. Most people answered "no" making this question hard to use. There is also a question about "half" children of the opposite race: results on this one are similar to those reported in the text for the almost identical question.

(5)).[50] The only question for which the two coefficients have the reverse pattern concerns "busing." However, we suspect that the answer to this question may have more to do with an individual's stand on mandatory busing as "reverse discrimination" versus "affirmative action," than with the actual willingness to interact with different races.

Unfortunately, we cannot perform a test similar to that reported in Table VII on our two other measures of heterogeneity. The GSS does not include questions that can proxy for attitudes toward ethnicity. As for income heterogeneity, the survey asks a few questions on attitudes toward redistributive fiscal policy (e.g., whether the government should actively help the poor, or whether the federal income is too high), but nothing that may capture attitudes toward *interactions* with individuals of a different income or social level.

An additional set of variables that we considered is the measures of racial and income segregation constructed by Cutler, Glaeser, and Vigdor [1999]. Our model does not deal directly with segregation, but one may argue that in more segregated communities racial and income mixing is lower, and therefore even people relatively averse to heterogeneity may be willing to participate in groups.[51] If this interpretation is correct, and if heterogeneity and segregation were positively correlated (as indeed they seem to be), omitting segregation from our regressions would bias our results *against* finding an effect of heterogeneity on participation. We have tested the effect of segregation and found results that are only weakly supportive of this interpretation. In particular, the various measures of segregation do not always appear to have significant coefficients, neither alone nor interacted with the heterogeneity variables. A possible explanation is that several of the groups in the GSS questionnaire are generally city based: thus, segregation may not be sufficient to prevent mixing, particularly in small cities. Also, recall that—as mentioned at the end of Section III—the degree of housing segregation in the GSS sample seems limited.

50. A "racist" is defined in the GSS as someone who believes that blacks are genetically inferior to whites. The GSS also includes a question on the right of racists to speak. The answers are extremely highly correlated with those to the question about the right to teach.

51. In terms of the model this may imply that in more segregated communities the case of multiple homogeneous groups is more likely to occur.

VI. Conclusions

Participation in social activities is positively associated with several valuable phenomena, like trust and human capital externalities. The propensity to participate is of course influenced to a large extent by individual characteristics, but it also depends on the composition and degree of heterogeneity of the community. In the theoretical part of this paper, we show under which conditions more heterogeneity in the population leads to less social interaction. We then explore the evidence on U. S. cities and find that income inequality and racial fragmentation are strongly inversely related to participation. Ethnic fragmentation also negatively influences participation, but less than racial fragmentation. The groups that are more affected by heterogeneity are those in which members directly interact to a significant extent, and in which excludability is low. Also, in accordance with our model, we find that the individuals who choose to participate less in racially mixed communities are those who most vocally oppose racial mixing.

Appendix 1

A. Proof of Proposition 1

Proposition 1 can be proved by applying Brouwer's fixed point theorem to the function $f(P_B)$ defined by the right-hand side of (8). Notice first of all that $f(P_B)$ maps the interval [0,1] into itself, and that [0,1] is clearly a nonempty, compact, and convex set. All we need to show therefore is that the right-hand side of (8) is a function, and that it is continuous. This follows from our assumptions that $f_\alpha(\cdot)$ and $f_l(\cdot)$ are continuously differentiable, $u(\cdot)$ and $v(\cdot)$ are well-behaved, and $B,W > 0$. $\square$

B. Stability

We define an equilibrium $(P_B^*,1 - P_B^*)$ as *locally stable* if for given W and B a small perturbation, say to $(P_B^* + \epsilon,1 - P_B^* - \epsilon)$ with $\epsilon \lesseqgtr 0$ sufficiently small, reverts to the original $(P_B^*,1 - P_B^*)$. In other words, a group is "stable" if when we add (remove) one member of either type, so that the composition of the group changes, this individual will choose to exit (reenter) the group. Applying Leibnitz's rule to the function $\Pi(\cdot)$ defined in (3), we get the following necessary and sufficient condition for $(P_B^*,1 - P_B^*)$ to be locally stable:

$$\text{(A.1)}\quad \frac{\Pi(P_B^*)\cdot\partial\Pi(1-P_B^*)/\partial P_B^* - \Pi(1-P_B^*)\cdot\partial\Pi(P_B^*)/\partial P_B^*}{[\Pi(1-P_B^*)+\Pi(P_B^*)(W/B)]^2}\cdot\frac{W}{B}<1.$$

C. Proof of Lemma 1

The proof of Lemma 1 can be divided into two parts.

(a) $B < W \Rightarrow P_B^* < P_W^*$.

Apply the implicit function theorem to (8) to get

$$\text{(A.2)}\quad \frac{\partial P_B^*}{\partial(W/B)} = \frac{\Pi(P_B^*)\cdot\partial\Pi(1-P_B^*)/\partial P_B^* - \Pi(1-P_B^*)\cdot\partial\Pi(P_B^*)/\partial P_B^*}{[\Pi(1-P_B^*)+\Pi(P_B^*)(W/B)]^2}\frac{W}{B}-1.$$

Under the stability condition (A.1) this derivative is unambiguously negative. Note that by symmetry when $B = W$ the unique equilibrium must be $P_B^* = \frac{1}{2} = P_W^*$. Condition (a) follows from these two facts.

(b) $P_B^* < P_W^* \Rightarrow B < W$.

By contradiction. Suppose that $B > W$. Then by the same arguments as in part (a) we should have $P_B^* > P_W^*$, which contradicts the hypothesis. □

D. Proof of Corollary 1

Let us start by showing that $B < W \Rightarrow (P_B^*/P_W^*) < (B/W)$. From (6) we can write the ratio of the two proportions as

$$\frac{P_B^*}{P_W^*} = \frac{\Pi(P_W^*)}{\Pi(P_B^*)}\frac{B}{W}.$$

From Lemma 1, $B < W$ implies that $P_W^* > P_B^*$. Given that $\partial\Pi(P_{-j})/\partial P_{-j} < 0$, this in turn implies that $\Pi(P_W^*) < \Pi(P_B^*)$, which proves the first part of the corollary. The second part, namely $B > W \Rightarrow (P_B^*/P_W^*) > (B/W)$, can be proved with the same arguments. □

E. Proof of Corollary 2

Starting from the participation constraint (2) and observing the monotonicity of its left-hand side, it is straightforward to obtain (9) and (10)—for the latter, remember that $u(0,P_{-j}^*) = 0$.

The last part of the corollary follows from the fact that $\bar{\alpha}_j = g(\bar{u} - v(0), P^*_{-j})$, where $\partial g/\partial P^*_{-j} < 0$, coupled with Lemma 1. $\square$

F. Proof of Proposition 2

From (12) we obtain

$$\text{(A.3)} \quad \frac{dS}{dw} = \left[\Pi(P^*_B) - \Pi[1 - P^*_B)\right] + \frac{\partial P^*_B}{\partial w} \cdot \left[w \frac{\partial \Pi(P^*_B)}{\partial P^*_B} + (1 - w) \frac{\partial \Pi(1 - P^*_B)}{\partial P^*_B}\right].$$

We need to find the conditions under which (A.3) is negative. Let us start by showing that $w < ½$ implies that $dS/dw < 0$. As proved in Lemma 1, when $w < ½$, we have $P^*_B > P^*_W$. Given that $\partial \Pi(P_{-j})/\partial P_{-j} < 0$, the expression in the first square brackets in (A.3) is thus negative. As for the second part of the derivative, we know that in a stable equilibrium $\partial P^*_B/\partial w < 0$, so it is sufficient (but not necessary) to show that the expression in the second square brackets is positive to prove our result. This amounts to requiring that

$$\text{(A.4)} \qquad \frac{w}{1 - w} < -\frac{\partial \Pi(1 - P^*_B)/\partial P^*_B}{\partial \Pi(P^*_B)/\partial P^*_B}.$$

Notice that the left-hand side of (A.4) is less than one by assumption.[52] Therefore, a sufficient, though not necessary, condition for the above inequality to hold is that $\partial^2 \Pi(P^*_B)/\partial P^*_B \geq 0$. Intuitively, $\partial^2 \Pi(P^*_B)/\partial P^*_B \geq 0$ says that the negative effect of heterogeneity on participation will likely be observed when the distribution of the α's is uniform or skewed to the right (i.e., when there is a significant part of the population who dislikes interaction with the opposite type) and when the fraction of people whose utility exceeds the reservation level decreases relatively more at low levels of the proportion of the opposite type in the group.[53] However, note that even when $\Pi''(\cdot) < 0$, i.e., when most individuals have mild preferences on racial relations, it is still possible

52. Notice also that $\partial \Pi(P^*_B)/\partial P^*_B$ and $\partial \Pi(1 - P^*_B)/\partial P^*_B$ have opposite signs. In fact, the former denotes how the α of W types changes when P^*_B changes (hence it is negative), while the latter denotes how the α of B types changes when P^*_B changes, namely $\partial \Pi(P^*_W)/\partial P^*_B > 0$.

53. The latter condition is much less restrictive than it looks. In fact it is *always* satisfied when α and P enter multiplicatively in the utility function, whatever the exact functional form.

that $dS/dw < 0$ because all other effects work in this direction. The condition $\Pi''(\cdot) \geq 0$ is in fact "twice sufficient": the first time because it is sufficient but not necessary for (A.4) to hold; the second time because (A.4) is sufficient but not necessary for (A.3) to be satisfied.

The second half of the proposition, namely the fact that $w \geq 1/2$ implies $dS/dw > 0$, can be proved along the same lines. □

G. An Example

Suppose that α has a uniform distribution on [0,1], individuals are uniformly distributed on [0,1], and that the utility function for type $j = B, W$ is

$$U_j = -\alpha\sqrt{P_{-j}} - l. \tag{A.5}$$

The reservation utility from nonparticipation is $\overline{u} \in (-.2,0)$ for everyone.[54] Notice that the functional form (A.5) has $\partial U_j/\partial P_{-j} < 0$ and $\partial^2 U_j/\partial P_{-j}^2 > 0$. A positive second derivative implies that increasing the proportion of whites decreases the marginal utility of blacks by more if there are very few whites. Suppose that a group is completely homogeneous; the first few participants of different types may require the adoption of different procedures, a different language etc. These costs would be declining as the minority becomes larger. This specification leads to a solution like the one that is represented graphically as in panel (a) of Figure I.[55] For a ratio $W/B = 2$, for instance, we have a unique stable equilibrium in which $P_W^* = 0.8$ and $P_B^* = 0.2$. The aggregate participation rate is

$$S = \overline{u}^2\left[\frac{1 - w}{\sqrt{1 - P_B^*}} + \frac{w}{\sqrt{P_B^*}}\right]. \tag{A.6}$$

It is easy to verify that the stability condition is always satisfied in this example and the derivative of S with respect to w is unambiguously positive for $w > 1/2$; i.e., reduced heterogeneity

54. The specific interval chosen for $\overline{u}$ ensures that in our example with $W/B = 2$ the domain of integration for α and l is a simple triangle as depicted in Figure VIII. Our results carry over to different parameter values: it is just a matter of splitting the integral to have the correct extremes of integration.

55. A different example can be constructed starting from the functional form $U_j = -\alpha(P_{-j})^2 - l$, which has both the first and second derivatives with respect to $P_{-j}(j = B,W)$ negative. The interpretation would be that of groups where majority voting matters for certain decisions, so that the marginal utility of losing members of your own type may be increasing as you are approaching a half and half split. This example generates homogeneous groups in equilibrium; i.e., $P_B^* = 1$ or $P_W^* = 1$.

increases participation. Figure VIII can help understand why this is the case.

On the horizontal axis we measure individual location, and on the vertical axis individual aversion. The members from the two types who participate in the group can be represented by "triangles." In fact, the frontier of (α, l) combinations satisfying the participation constraint is

$$\alpha < (-\overline{u} - l)/\sqrt{P_{-j}}. \tag{A.7}$$

The basis of the triangles in Figure VIII is $2\overline{l} = -2\overline{u}$ and is the same for both types, given that the maximum distance from the group's location is independent of W/B. The height of the triangles, given by $\overline{\alpha}_j = -\overline{u}/\sqrt{P_{-j}}$, is lower for the minority type,

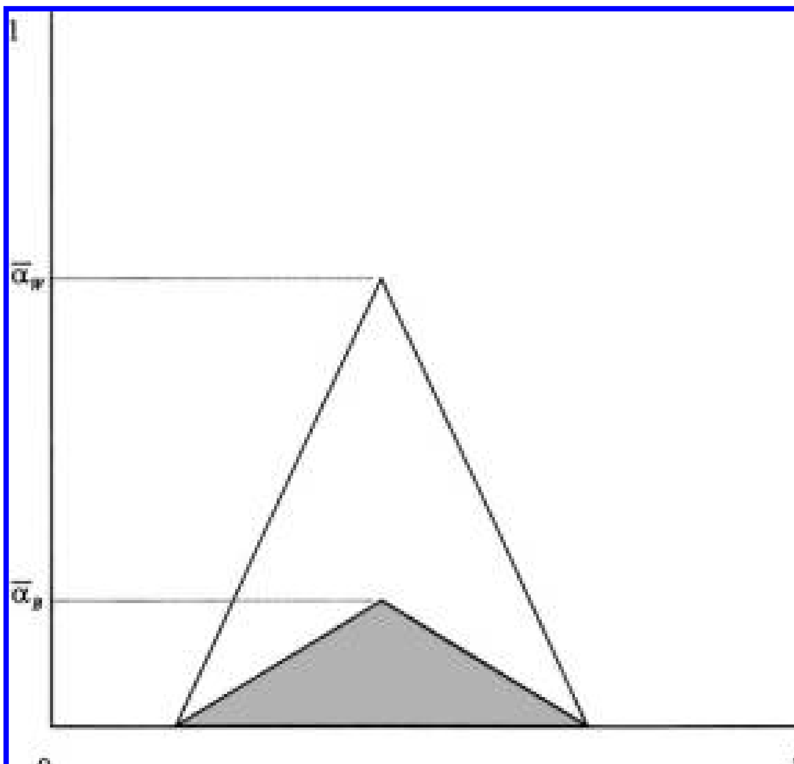

FIGURE VIII
Example of a Mixed Group

following Corollary 2. The "mass" of participants from the two types, $\tilde{B}$ and $\tilde{W}$, is obtained multiplying the areas of the two triangles by a third dimension, B and W, respectively. In graphical terms, an increase in heterogeneity translates into two changes. On the one hand, the height of the B triangle ($\overline{\alpha}_B$) increases, and that of the W triangle ($\overline{\alpha}_W$) decreases. On the other hand, the depth B by which we multiply the smaller triangle increases (and W decreases by the same amount). Notice that from (A.7) we have $\partial^2\overline{\alpha}_j/\partial P_{-j}^2 > 0$, which implies that $\Delta\overline{\alpha}_B < -\Delta\overline{\alpha}_W$ (where Δ denotes the absolute change). If the changes in $\overline{\alpha}_j$ were applied to the same "depth," by themselves they would reduce aggregate participation; the fact that $B < W$ reinforces this decrease.

Turning to the case of multiple groups, in subsection II.4 we prove that in the case of *disjoint* mixed groups the results of the single-group model carry over without alterations. Hence, in this example we focus on multiple *adjacent* groups. When the geographic coverage of each group is $1/n < 2|\overline{u}|$, the fraction of members type $j = B, W$ belonging to any group is $\Pi(P_{-j},\overline{u}) = [-(\overline{u}/n) - (1/4n^2)](P_{-j})^{-1/2}$. The aggregate rate of participation is now

$$S = \left(-\frac{\overline{u}}{n} - \frac{1}{4n^2}\right)\left[\frac{1-w}{\sqrt{1-P_B^*}} + \frac{w}{\sqrt{P_B^*}}\right], \tag{A.8}$$

which coincides with (A.6) except for the multiplicative factor in front of the square brackets. The sign of dS/dw, i.e., the impact of heterogeneity on participation, will therefore be the same as in the one-group model.[56]

H. Proof of Proposition 3

In any equilibrium with disjoint homogeneous groups, all and only the individuals type j located within $v^{-1}(\overline{u})$ from the headquarters will join a group with $P_j^* = 1, \forall\alpha \in [0,1]$. The aggregate participation rate is therefore

$$S = 2v^{-1}(\overline{u})w \cdot N_W + 2v^{-1}(\overline{u})(1-w) \cdot N_B I_B, \tag{A.9}$$

where N_j is the number of groups with $P_j^* = 1$, $j = B,W$, and the

56. This is not surprising, given that the linearity of utility in distance implies that when we "cut the sides" of a group we leave out the same proportion of B and W types. However, our results hinge on the separability of the utility function in P_{-j} and l, not on the specific way l enters utility. In fact, it is easy to verify that introducing a quadratic distance term or a square root does not alter the conclusion on the sign of dS/dw.

indicator variable I_B takes the following values:

$$\begin{aligned} I_B &= 1 \qquad \text{if } 2v^{-1}(\overline{u})B > [1/(n+1)](B+W) \\ &= 0 \qquad \text{otherwise.} \end{aligned}$$

Notice that in (A.9) we are making the working assumption that the majority type W always meets the minimum size requirement when all the α's participate. Remembering that $w = W/(B + W)$, the change in S when we decrease w by $\epsilon > 0$ around the point where I_B goes from 0 to 1 is given by

$$\Delta S = N_B(1 - w) - \epsilon(N_W - N_B).$$

For $\epsilon \to 0$, the second addendum (which captures the potential decrease in participation if the number of all-W groups exceeds that of all-B groups), is second order compared with the first one—hence $\Delta S > 0$. The rest of Proposition 3 is straightforward. □

Appendix 2: Variable Definition

The following is a list of the variables we use and their sources, followed by summary statistics. The data sources are abbreviated as follows: GSS stands for General Social Survey, cumulative file 1972–1994; CensusCD90 refers to the CDrom CensusCD+Maps by GeoLytics, Inc. (1996–1998) which contains data from the Summary Tape Files 3F of the 1990 Census. In all cases for variables constructed from the GSS, "no answer" and "not applicable" were coded as missing values. Unless otherwise stated, the source of a variable is authors' calculation on GSS data.

Member: dummy equal to 1 if respondent is a member of at least one group.

Member (excluding nationality): dummy equal to 1 if respondent is a member of at least one group other than a nationality group.

Member (excluding unions): dummy equal to 1 if respondent is a member of at least one group other than a union.

Cohort: year of birth of the respondent.

Age < 30: dummy equal to 1 if respondent is less than 30 years old.

Age 30–39: dummy equal to 1 if respondent is between 30 and 39 years old.

Age 50–59: dummy equal to 1 if respondent is between 50 and 59 years old.

Age $\geq$ 60: dummy equal to 1 if respondent is 60 years old or more.

Married: dummy equal to 1 if respondent is married.

Female: dummy equal to 1 if respondent is female.

Black: dummy equal to 1 if respondent is African-American.

Educ $<$ 12 yrs: dummy equal to 1 if respondent has less than twelve years of education.

Educ $>$ 16 yrs: dummy equal to 1 if respondent has more than sixteen years of education.

Children $\leq$ 5 yrs: dummy equal to 1 if respondent has children age five or less.

Children 6–12: dummy equal to 1 if respondent has children age six to twelve.

Children 13–17: dummy equal to 1 if respondent has children age thirteen to seventeen.

ln (real income): logarithm of respondent's family income (constant 1986 US$).

Full-time: dummy equal to 1 if respondent works full-time.

Part-time: dummy equal to 1 if respondent works part-time.

Size of place: logarithm of the size of place where respondent lives (thousands of people).

Med HH income: logarithm of median household income in MSA/PMSA where respondent lives [*Source:* authors' calculation on CensusCD90].

Med HH income^2: square of the logarithm of median household income in MSA/PMSA where respondent lives [*Source:* authors' calculation on CensusCD90].

Gini: Gini coefficient on family income in MSA/PMSA where respondent lives. Actual Gini coefficients were computed for the years 1970, 1980, 1990. The values for the remaining years in the sample were obtained by linear interpolation (and extrapolation for 1991–1994) [*Source:* authors' calculation on IPUMS 1%, Census 1970, 1980, 1990].

Race: racial fragmentation index in MSA/PMSA where respondent lives, defined in expression (16) in the text. The five categories used for the shares are the original Census categories: i) white; ii) black; iii) American Indian, Eskimo, Aleutian; iv) Asian, Pacific Islander; v) other [*Source:* authors' calculation on CensusCD90].

Ethnic: ethnic fragmentation index in MSA/PMSA where respondent lives, defined in expression (16) in the text. The ten categories used for the shares are obtained aggregating the original "first ancestries" from the Census as follows: (1) Arab; (2) Sub-Saharan African; (3) West Indian; (4) Race or Hispanic origin; (5) Canadian, United States, or American; (6) Austrian, Belgian, Dutch, English, French Canadian, German, Irish, Scotch-Irish, Scottish, Swiss, Welsh; (7) Czech, Hungarian, Lithuanian, Polish, Romanian, Russian, Slovak, Ukrainian, Yugoslavian; (8) French, Greek, Italian, Portuguese; (9) Danish, Finnish, Norwegian, Swedish; (10) other. Each share is computed as a share of people in that category over the total population in the MSA/PMSA (excluding people with "ancestry unclassified" and "ancestry not reported") [*Source:* authors' calculation on CensusCD90].

NGOV62: number of municipal and township governments in the MSA/PMSA in 1962 (*Source:* Cutler, Glaeser, and Vigdor [1999]).

MANSHR: share of the labor force employed in manufacturing in the MSA/PMSA in 1990 (*Source:* Cutler, Glaeser, and Vigdor [1999]).

NOBLKDINNER: dummy equal to 1 if respondent has not had a black person home for dinner in past few years. Original GSS survey question: "During the last few years, has anyone in your family brought a friend who was a black home for dinner?" Prompted answers coded in the GSS variable RACHOME: 1 = Yes; 2 = No; 8 = Don't know; 9 = No answer. Our variable takes value 1 if RACHOME = 2 and zero otherwise.

NORACCHNG: dummy equal to 1 if respondent says that he/she would not try to change racist rules in a club. Original GSS survey question: "If you and your friends belonged to a social club that would not let whites/blacks join, would you try to change the rules so that they could join?" Prompted answers coded in the GSS variable RACCHNG: 1 = Yes; 2 = No; 3 = Wouldn't belong to club; 8 = Don't know; 9 = No answer. Our variable takes value 1 if RACCHNG = 2 and zero otherwise.

RACSEGR: dummy equal to 1 if respondent strongly agrees that whites have a right to segregated neighborhoods. Original GSS survey question: "White people have a right to keep blacks out of their neighborhoods if they want to, and blacks should respect that right." Prompted answers coded in the GSS variable RACSEG: 1 = Agree strongly; 2 = Agree slightly; 3 = Disagree slightly; 4 = Disagree strongly; 8 = No opinion; 9 = No answer. Our variable takes value 1 if RACSEG = 1 and zero otherwise.

NOMOSTSCHOOL: dummy equal to 1 if respondent would not send children to school with most children of the opposite race. Original GSS survey question: "Would you yourself have any objection to sending your children to a school where most of the children are Whites/Blacks?" Prompted answers coded in the GSS variable RACMOST: 1 = Yes; 2 = No; 3 = Don't know. Our variable takes value 1 if RACMOST = 1 and zero otherwise.

RACTEACH: dummy equal to 1 if respondent thinks that racists should be allowed to teach. Original GSS survey question: "Consider a person who believes that blacks are genetically inferior. Should such a person be allowed to teach in a college or university, or not?" Prompted answers coded in the GSS variable COLRAC: 4 = Yes, allowed; 5 = Not allowed; 8 = Don't know; 9 = No answer. Our variable takes value 1 if COLRAC = 4 and zero otherwise.

NOBLKPRESID: dummy equal to 1 if respondent would not vote for black president. Original GSS survey question: "If your party nominated a black for President, would you vote for him if he were qualified for the job?" Prompted answers coded in the GSS variable 'RACPRES': 1 = Yes; 2 = No; 8 = Don't know; 9 = No answer. Our variable takes value 1 if RACPRES = 2 and zero otherwise.

NOMIXMARRIAGE: dummy equal to 1 if respondent is against mixed marriages. Original GSS survey question: "Do you think there should be laws against marriages between blacks and whites?" Prompted answers coded in the GSS variable 'RACMAR': 1 = Yes; 2 = No; 3 = Don't know. Our variable takes value 1 if RACMAR = 1 and zero otherwise.

NOBUSING: dummy equal to 1 if respondent opposes busing. Original GSS survey question: "In general, do you favor or oppose the busing of black and white school children from one school district to another?" Prompted answers coded in the GSS variable 'BUSING': 1 = Favor; 2 = Oppose; 8 = Don't know; 9 = No answer. Our variable takes value 1 if BUSING = 2 and zero otherwise.

BLKNOPUSH: dummy equal to 1 if respondent thinks that blacks should not push. Original GSS survey question: "Here are some opinions other people have expressed in connection with black-white relations. Which statement on the card comes closest to how you, yourself, feel? The first one is: Blacks shouldn't push themselves where they're not wanted." Prompted answers coded in the GSS variable 'RACPUSH': 1 = Agree strongly; 2 = Agree slightly; 3 = Disagree slightly; 4 = Disagree strongly; 8 = No opinion; 9 = No answer. Our variable takes value 1 if RACPUSH = 1 and zero otherwise.

Appendix 3: Summary Statistics

	Mean	Std. dev.	No. obs.
Member	.72	.45	10031
Member (excl. nationality)	.71	.45	9922
Member (excl. unions)			9922
Cohort	1939.85	17.59	10031
Age <30	.25	.43	10031
Age 30–39	.24	.43	10031
Age 50–59	.13	.34	10031
Age >60	.20	.40	10031
Married	.56	.50	10031
Female	.56	.50	10031
Black	.16	.37	10031
Educ <12 yrs	.24	.43	10031
Educ >16 yrs	.22	.41	10031
Children ≤ 5 yrs	.19	.39	10031
Children 6–12	.22	.41	10031
Children 13–17	.18	.38	10031
ln (real income)	10.03	.93	10031
Full-time	.52	.50	10031
Part-time	.10	.30	10031
Size of place	4.28	2.18	10031
Med HH income	10.38	.14	10031
Med HH inc $\wedge$2	107.74	3.0	10031
Gini	.41	.03	10031
Race	.36	.14	10031
Ethnic	.67	.07	10031
NGOV62	89.84	90.78	10031
MANSHR	.17	.05	10031
NOMIXMARRIAGE	.19	.40	5901
NOBLKDINNER	.68	.47	4493
BLKNOPUSH	.32	.47	2885
NOHALFSCHOOL	.19	.39	5717
RACTEACH	.44	.50	4661
NOBLKPRESID	.12	.33	5872
NOBUSING	.73	.44	6303

APPENDIX 4: NONLINEARITIES IN INCOME

	(1)	(2)	(3)	(4)	(5)	(6)
ln (real income)	−1.789**	−1.818**	−1.806**	−1.835**	−1.821**	−1.853**
	(.550)	(.582)	(.553)	(.586)	(.546)	(.577)
ln (real income) ∧2	.196**	.204**	.198**	.205**	.199**	.207**
	(.060)	(.063)	(.060)	(.064)	(.059)	(.063)
ln (real income) ∧3	−.007**	−.007**	−.007**	−.007**	−.007**	−.007**
	(.002)	(.002)	(.002)	(.002)	(.002)	(.002)
Below poverty		.039*		.039*		.039*
		(.020)		(.020)		(.020)
Gini	−.901**	−1.029**				
	(.242)	(.244)				
Race			−.201**	−.227**		
			(.058)	(.056)		
Ethnic					−.256**	−.282**
					(.106)	(.108)
INDIV CONTROLS[a]	Yes	Yes	Yes	Yes	Yes	Yes
STATES	Yes	Yes	Yes	Yes	Yes	Yes
YEARS	Yes	Yes	Yes	Yes	Yes	Yes
No. obs.	10534	10181	10534	10181	10534	10181
Pseudo R^2	.10	.09	.10	.09	.09	.09
Observed *P*	.72	.72	.72	.72	.72	.72
Predicted *P*	.74	.74	.74	.74	.74	.74

*Denotes significance at the 10 percent level; **at the 5 percent level.

Marginal probit coefficients are calculated at the means. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.

a. Individual controls: all those listed in Table II.

APPENDIX 5: FIRST-STAGE REGRESSIONS (DEPENDENT VARIABLE: GINI)

	Column of reference in Table VI		
	[2]	[3]	[5]
NGOV62	.0001**	.0001**	.0001**
	(.0000)	(.0000)	(.0000)
REVIG62		.117*	
		(.063)	
MANSHR			−.092
			(.080)
CONTROLS[a]	Yes	Yes	Yes
STATES	Yes	Yes	Yes
YEARS	Yes	Yes	Yes
No. obs.	725	725	725
R^2	.67	.67	.67

*Denotes significance at the 10 percent level; **at the 5 percent level.

Marginal probit coefficients are calculated at the means. Standard errors are corrected for heteroskedasticity and clustering of the residuals at the MSA level.

a. Controls: means at the MSA/PMSA level of all individual controls listed in Table II, plus size, Hincmd, and Hinmd².

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