Land Inequality and Political Violence

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The American Political Science Review, Vol. 83, No. 2 (Jun., 1989), 577-596.

Stable URL:
http://links.jstor.org/sici?sici=0003-0554%28198906%2983%3A2%3C577%3ALIAPV%3E2.0.CO%3B2-E

The American Political Science Review is currently published by American Political Science Association.

LAND INEQUALITY AND POLITICAL VIOLENCE

Considerable research effort has been invested in establishing the appropriate relationship between patterns of land distribution and political violence. In an article in the June 1988 issue of the Review, *Manus I. Midlarsky proposed and tested a new measure of the distribution of land, which he called "patterned inequality." He presented supporting evidence with data from Latin American and Middle Eastern countries. In this controversy, Midlarsky's analysis is challenged by Edward N. Muller, Mitchell A. Seligson, and Hung-der Fu. They advocate an alternative measure of land inequality, test its effect on levels of political violence in Latin America, and find it wanting. In his rejoinder, Midlarsky offers new analytical support for his claims.*

It is frequently asserted that land inequality in agrarian societies is a major cause of rebellion. Cross-national tests of the land inequality hypothesis typically have used the Gini concentration ratio as the operational definition of land inequality. The most comprehensive test to date, conducted with data from 83 countries, found little support for it. A correlation (*r*) of only .24 was observed between the Gini ratio of land concentration circa 1970—weighted by the size of the agricultural labor force—and the death rate from political violence during 1973–77 (Muller and Seligson 1987, 435).

Three studies have recently attempted to uncover the theoretical explanation for the lack of an association between the Gini ratio of land concentration and political violence. Two of these studies insist that land distribution is in fact a major cause of political violence but that the failure to find strong and robust correlations is a function of inappropriate index construction. The third study, our own (Muller and Seligson 1987), views land distribution as important only when linked to the creation and maintenance of income inequality, the variable that has a direct bearing on political violence.

One proponent of the hypothesis that maldistribution of land is a major cause of rebellion argues that the Gini ratio is inappropriate because it is insensitive to the kind of inequality that is most likely to spawn peasant-based insurgency. Prosterman and Riedinger (1987) hypothesize that landlessness, rather than inequality in the distribution of land, is the most important cause of deeply felt grievances among peasants. In their view, the key to predicting the likelihood of large-scale internal violence in agrarian societies is the percentage of landless peasants in the entire population. The Gini land concentration ratio does not speak to the question of landlessness, of course, since it is based only on the landed population.

To support their thesis Prosterman and Reidinger (1987) cite a number of examples of twentieth-century revolutions in societies with substantial proportions of landless peasants (pre-1911 Mexico, pre-1917 Russia, pre-1936 Spain, pre-1941 China, pre-1952 Bolivia, pre-1959 Cuba, and pre-1961 South Vietnam). Although they claim that their index of landlessness is "remarkably well correlated with instances of major instability" (p. 25), Prosterman and Reidinger did not report correlation coefficients and made no effort to determine the predictive accuracy of the landlessness measure. Nor did they conduct tests to see if an association between landlessness and political violence held when other alternative causes of political violence were taken into account. When the 1973–77 rate of death from political violence is regressed on landlessness data circa 1970 (from Prosterman and Reidinger 1982), the adjusted r-squared value is only .04. And when income inequality is controlled for, the association between landlessness and political violence is not statistically significant (Muller and Seligson 1987, 436). Landlessness thus has trivial predictive power in cross-national perspective and is irrelevant also in the multivariate context. In short, it is a variable that does not seem to merit further consideration at this time.

The second proponent of the importance of land scarcity for political violence is Midlarsky (1982, 1988). He argues that a principal cause of political violence in agrarian societies is a process of bifurcation in the pattern of landholdings. Given a fixed supply of land, such factors as high birth rates and the absence of primogeniture among the peasant population will lead to increased subdivision of small farms, and the average size of the smallholders' properties will shrink. If large farms remain relatively intact, it is a classic case of immiserization, in which the rich remain rich while the poor become poorer. Midlarsky (1988) calls this a condition of "patterned inequality" and claims that apart from agreement at the extremes, the Gini concentration ratio is insensitive to differences in patterned inequality and therefore yields low associations with violence. He reports evidence purporting to show that patterned inequality in land distribution is a much more powerful determinant of political violence in Latin America and the Middle East than the Gini ratio.

There are serious flaws in both his measure of patterned inequality and his test of the hypothesis that patterned inequality is a major cause of political violence. Nevertheless, the concept of bifurcation is of real theoretical interest because it suggests an alternative to the equal-share concept of equality that underlies such frequently used measures as the Gini ratio.

The concept of bifurcated inequality, as we will define it here, focuses attention specifically on the difference between the average farm size in the upper and lower extremes of the distribution of landholdings—the gap between rich landlords and poor peasants. Bifurcated inequality is characteristic of the traditional lord-and-peasant type of agrarian society, in which most of the land is concentrated in a few great estates, while there are many very small farms. The condition of bifurcated inequality is reflected only imperfectly by the Gini ratio, which measures the extent to which the cumulative distribution of landholdings departs from a norm of perfect equality and therefore is sensitive to inequalities present in all parts of the distribution.

The hypothesis that bifurcated inequality will be associated with political violence more strongly than is the Gini concentration ratio is predicated on a rejection of the relevance of the equal-share concept of equality as a cause of political violence. Specifically, the assumptions are (1) that owners of middle-size farms will not be unhappy with their share of landholdings even though it may be considerably less than that of large landowners and (2) that owners of small farms will be unhappy with their share only if it is extremely small relative to that of large landowners. Instead of comparing the extent to which the entire distribution of landholdings departs from an equal-share norm of equality, the relevant comparison for measuring bifurcated inequality would be the ratio of the average size of a large farm to the average size of a small farm.

After reviewing the limitations of Midlarsky's patterned inequality measure of bifurcation, we will propose an alternative measure and compare it with Midlarsky's patterned inequality. Then we will conduct tests of Midlarsky's argument that bifurcated inequality is a relevant cause of cross-national variation in rates of political violence in Latin America.

Patterned Inequality

Midlarsky (1988) stipulates that a condition of patterned inequality occurs when the lower end of the distribution of landholdings fits an exponential model and the upper end fits a log-exponential model. The combination of an exponential lower distribution and a log-exponential upper distribution is assumed to be a result of a process of bifurcation in the distribution of landholdings such that the average size of large farms remains stable while that of small farms declines. The exponential-lower-and-log-exponential-upper combination is presumed to represent a high degree of inequality resulting from the process of bifurcation. Distributions that do not conform in the lower part to the exponential or in the upper part to the log-exponential are presumed to represent a low degree of bifurcated inequality.

Midlarsky's strategy of measuring the degree of bifurcation by comparing the fit between theoretical and observed distributions has several limitations. First, it is indirect, based on a set of assumptions about the correspondence between the subdivision of smallholdings and the form of the lower and upper parts of the distribution of landholdings. Since bifurcation, in the sense of the gap between the average size holdings of the rich and poor, is easy to measure directly, there is no need for a measure that merely infers bifurcation from patterns of land distribution.

Second, there are many possible intermediate patterns between the two extreme types, and it is not clear how to rank-order them in regard to differing amounts of bifurcated inequality and differing amounts of expected violence. Midlarsky asserts that a fit of the exponential model to the upper part of the distribution and no fit of either model to the lower portion represents less of a bifurcated inequality than a fit of the log-exponential model to the upper part and no fit of either model to the lower part. But the reason for this difference is not explained. Another intermediate pattern is a fit of the exponential model to the lower part of the distribution and no fit of either model to the upper part. Midlarsky does not try to rank-order this pattern. Still another intermediate pattern involves a "near" fit to one of the theoretical distributions. If near fits are allowed, not only is the number of possible intermediate patterns increased substantially but even the criterion for determining distinct patterns becomes vague, since a near fit in one observer's opinion might not be a near fit to another.

A third problem with the patterned inequality measure is that the definition of the upper and lower ranges of land distribution is not constant or systematically derived across the countries Midlarsky studies. The upper and lower ranges are determined by inspection. This element of subjectivity in the measure introduces the possibility that selection of the data points to be included in the upper and lower ranges could be biased for or against a given hypothesis.

Although Midlarsky reported the upper and lower ranges only for the case of El Salvador, it is possible to infer those ranges from the degrees of freedom given in his Table 2, since the total number of data points used to compare the fit between the observed and theoretical distri-

Table 1. Midlarsky's Upper and Lower Ranges for the Log-Exponential and Exponential Analyses of Land Distribution in Latin America

Country	Year	Total Data Points[a]	Upper Part: Data Points	Upper Part: Range (hectares)	Lower Part: Data Points	Lower Part: Range (hectares)
Argentina	1960	9	5	200.1–10,000	—	—
Bolivia	1950	10	6	20.1–2,500	4	.1–20
Brazil	1960	4	5	—	—	—
Chile	1950	11	5	100–4,999.9	5	.1–49.9
Colombia	1960	16	5	50–2,500	6	.1–4.9
Costa Rica	1960	17	4	350–2,449.99	—	—
Cuba	1952	8	5	10–1,000	—	—
Dominican Republic	1960	10	5	4.7–1,572.5	—	—
Ecuador	1954	11	5	50–2,500	—	—
El Salvador	1960	13	5	50–2,499.99	5	.1–4.99
Guatemala	1950	11	4	447.2–8,943.99	—	—
Haiti	1970	18	5	5.2–25.8	—	—
Honduras	1960	10	5	35–1,749.99	—	—
Mexico	1960	9	7	10.1–5,000	2	.1–10
Nicaragua	1960	10	5	35–1,749.99	—	—
Panama	1960	9	4	50–1,000	—	—
Paraguay	1960	17	8	50–19,999.9	—	—
Peru	1961	15	5	50–2,500	4	.1–3
Uruguay	1960	11	5	200–9,999	—	—
Venezuela	1960	22	5	50–2,499.9	—	—

Source: Wilkie and Perkal 1985, 34–38.
[a]All categories –1, since the midpoint of the largest category cannot be estimated.

butions is two more than the degrees of freedom. The upper and lower ranges selected by Midlarsky for testing the exponential and log-exponential models in Latin American countries are listed in Table 1.[1]

There is extreme and inexplicable variation in Midlarsky's definition of what constitutes the upper range of large farms in each country. In the Dominican Republic large farms are those with 4.7 hectares of land or more. Large farms are defined as those with at least 20 hectares in Bolivia. At least 35 hectares are required for a landholding to be considered large in Honduras and Nicaragua. In Ecuador, El Salvador, Panama, Paraguay, Peru, and Venezuela the definition of a large farm is one with 50 hectares of land or more. In Argentina and Uruguay at least 200 hectares are required for a farm to be considered large. Large farms in Costa Rica are those with at least 350 hectares. And in Guatemala a large farm is defined as one with nearly 450 hectares or more. Most troubling about the apparently arbitrary selection of the upper ranges is that they contradict the basic land tenure patterns for some countries. For example, Argentina, with 457 million hectares of land in farms in 1960, is defined by Midlarsky as having a *lower* limit for his category of large farms than does Costa Rica, a country with only 2.7 million hectares of land in farms in 1963. Using the same source consulted by Midlarsky (Wilkie and Perkal 1985, 25), one finds 5.1 times as much arable land per capita for the agriculturally employed population in Argentina as in Costa Rica. More to the point, the average farm in the largest size cohort is 3.1 times larger in Argentina than Costa Rica. Yet, Midlarsky considers 200-hectares farms in Argentina to be large while Costa Rican farms need to be at least 350 hectares to be considered large.

Midlarsky determined by inspection that the lower range did not fit either the exponential or log-exponential distribution in most instances. Since the degrees of freedom were not reported for these cases, it is impossible—with one exception—to determine the boundaries of the lower range. The exception is Mexico, where seven of the nine data points were used to test the fit of the log-exponential model to the upper range. This means that the lower range must have comprised the two smallest categories, farms of up to 5 hectares and of 5.1–10 hectares. Otherwise, the lower range is up to 3 hectares in Peru, less than 5 hectares in Colombia and El Salvador, up to 20 hectares in Bolivia, and less than 50 hectares in Chile.

Since the operational definition of what constitutes the upper and lower ranges of land distribution is so variable across countries, it is impossible to compare them systematically. In some countries the upper range is restricted to extremely large farms of more than 200 hectares (Argentina, Costa Rica, Guatemala, and Uruguay), whereas in other countries it is expanded to include even farms as small as 5 hectares (the Dominican Republic). In some countries the total distribution is divided into only an upper and a lower part (Bolivia and Mexico), whereas there is apparently a middle part in the others. In one country (Chile) the lower range encompasses farms up to a size (50 hectares) that is included in the upper range in at least seven countries. Thus the fit of the observed to the theoretical distributions is essentially idiosyncratic, and there is no reliable way of grouping them into categories of patterned inequality.

Bifurcated Inequality

The data for our measure of bifurcated inequality are drawn principally from the 1950, 1960, and 1970 world censuses of agriculture conducted by the Food and Agriculture Organization (FAO) of the United Nations. The FAO (1981) reports information on landholdings for identical size categories of farms for each country: less than 1 hectare, 1 and under 2 hectares, 2 and under 5 hectares, 5 and under 10 hectares, 10 and under 20 hectares, 20 and under 50 hectares, 50 and under 100 hectares, 100 and under 200 hectares, 200 and under 500 hectares, 500 and under 1,000 hectares, and 1,000 hectares and over. We define a relatively small farm as one falling within the lower three levels of farm size—5 hectares or less. A middle-sized farm is defined as one falling within the middle four categories of farm size—5–99.9 hectares; and a relatively large farm is defined as one falling within the upper four levels of farm size—100 hectares or more. The size of the gap between large and small farms is measured by the ratio of the average number of hectares per large farm to the average number of hectares per small farm. This ratio of the average landholding of large farms to that of small farms—or land bifurcation ratio—is then weighted (multiplied) by the proportion of small farms in the country to yield a bifurcated inequality (BI) index (see the third, fourth, and fifth columns in Table 2). The weighting procedure is important theoretically because a bifurcated distribution of land will be more salient to the agricultural population, the greater the proportion of small farms.[2]

The index of bifurcated inequality is a quantitative variable that ranges from zero to a large number. If there are no large farms, the BI index will be zero. The BI index will be low if either the land bifurcation ratio is low or the proportion of

Table 2. Bifurcated Land Inequality in Latin American Countries

Country	Year	Hectares per Farm Small (< 5)	Hectares per Farm Large (≥ 100)	Land Bifurcation Ratio	Proportion of Small Farms	Index of Bifurcated Inequality	Gini
Argentina	1960	2.8	895.0	319.6	.16	51.1	.86
Bolivia	1950	1.4	2,775.9	1,914.4	.59	1,129.5	.94
Brazil	1950	2.6	642.0	246.9	.22	54.3	.83
	1960	2.5	565.8	226.3	.31	70.2	.83
	1970	2.2	498.9	226.8	.37	83.9	.84
Chile	1955	1.4	1,188.1	848.6	.35	296.8	.93
	1965	1.7	1,232.7	725.1	.49	355.3	.94
Colombia	1954	1.8	419.2	232.9	.60	139.7	.85
	1960	1.6	419.6	262.3	.63	165.2	.86
	1971	1.3	418.1	321.6	.60	193.0	.86
Costa Rica	1950	2.2	551.0	250.4	.39	97.7	.81
	1963	2.1	463.6	220.8	.39	86.1	.79
	1973	1.7	360.9	212.3	.43	91.3	.81
Dominican Republic	1950	1.5	509.0	339.3	.76	257.9	.80
	1960	1.2	513.3	427.8	.86	367.9	.80
	1971	1.5	381.5	254.3	.77	195.8	.82
Ecuador	1954	1.7	536.5	315.6	.73	230.4	.86
	1974	1.6	317.4	198.4	.65	129.0	.81
El Salvador	1950	1.4	382.0	272.9	.80	218.3	.83
	1961	1.2	360.5	300.4	.85	255.4	.84
	1971	1.2	330.1	275.1	.74	203.6	.81
Guatemala	1950	1.7	245.1	144.2	.88	126.9	.87
	1964	1.8	354.8	197.1	.87	171.5	.82
Haiti	1971	1.1	0	0	.96	0	.48
Honduras	1952	2.3	351.8	153.0	.56	85.7	.75
	1966	2.5	262.5	105.0	.68	71.4	.75
	1974	1.9	331.1	174.3	.64	111.6	.78
Mexico	1950	1.4	1,621.5	1,158.2	.73	845.5	.96
	1960	1.5	1,439.2	959.5	.66	633.3	.95
	1970	1.7	1,331.8	783.4	.51	399.5	.93
Nicaragua	1952	3.0	256.9	85.6	.35	30.0	.76
	1963	2.6	245.6	94.5	.51	48.2	.80
Panama	1950	2.1	342.5	163.1	.53	86.4	.71
	1960	2.2	262.4	119.3	.46	54.9	.74
	1971	1.4	293.9	209.9	.47	98.7	.78
Paraguay	1956	2.2	2,940.0	1,336.4	.46	614.8	.95
Peru	1961	1.5	1,484.4	989.6	.81	801.6	.94
	1972	1.5	1,184.4	789.6	.76	600.1	.91

TABLE 2 (*continued*)

Country	Year	Hectares per Farm Small (< 5)	Large (≥ 100)	Land Bifurcation Ratio	Proportion of Small Farms	Index of Bifurcated Inequality	Gini
Uruguay	1951	2.7	727.1	269.3	.14	37.7	—[a]
	1961	2.8	704.5	251.6	.15	37.7	.82
	1971	2.6	700.4	269.4	.13	35.0	.82
Venezuela	1950	2.1	1,531.1	729.1	.54	393.7	.94
	1961	2.3	1,160.4	504.5	.48	242.2	.92
	1971	2.1	941.7	448.4	.42	188.3	.91

[a]Insufficient data in the upper range for calculation of Gini.

small farms is low. The index of bifurcated inequality will be high if both the land bifurcation ratio and the proportion of small farms are high. Thus, where a process of bifurcation has created a polarized situation of many small farms and a few large farms, the index of bifurcated inequality will be high. At least one value for the index of bifurcated inequality can be calculated for every Latin American country except Cuba, and in 10 countries bifurcated inequality scores can be compared for the 1950s, the 1960s, and the 1970s. The relevant data are listed in Table 2,[3] along with the Gini concentration ratio.

In Midlarsky's (1988) view, El Salvador is the paradigmatic case of patterned inequality because it is supposed to have one of the most clearly bifurcated patterns of land distribution in Latin America. The average size of a landholding for small farms in El Salvador has been consistently one of the lowest in Latin America, initially 1.4 hectares in 1950, declining to 1.2 hectares during the 1960s and 1970s. However, the average size of a large farm in El Salvador has never been very great. Argentina, Bolivia, Brazil, Chile, Colombia, Mexico, Paraguay, Peru, Uruguay, and Venezuela all have a much greater average landholding among large farms than does El Salvador. Indeed, when the ratio of large to small farms is computed, the amount of land bifurcation turns out to be much greater in six other Latin American countries (Bolivia, Chile, Mexico, Paraguay, Peru, and Venezuela) than it is in El Salvador, which registers only intermediate values (see the Land Bifurcation Ratio column of Table 2).

Midlarsky contrasts El Salvador with Nicaragua, which is claimed to have a much lower level of bifurcation due to less land scarcity. This is an accurate observation. The average landholding of small farms in Nicaragua—initially 3 hectares in the 1950s, declining only to 2.6 hectares in the 1960s—was one of the three highest values in Latin America (small farms in Argentina and Uruguay being of similar size); while the average landholding of large farms in Nicaragua was among the three lowest values (Guatemala was similar and there were no farms of 100 hectares or more in Haiti). Consequently, Nicaragua had less land bifurcation than any other Latin American country except Haiti.

Costa Rica and Uruguay are presumed by Midlarsky to have the least amount of land bifurcation because in these countries the upper and lower parts of the distribution fit neither the exponential nor the log-exponential. However, as measured by the land bifurcation ratio, five Latin American countries (Guatemala, Haiti, Honduras, Nicaragua, and Panama) have had less land bifurcation than Costa Rica and Uruguay. Indeed, two of the three instances of very high patterned inequality (Colombia and El Salvador) have approximately the same land bifurcation ratio as the two instances of very low patterned inequality (Costa Rica and Uruguay). Thus, patterned inequality measures something other than land bifurcation.

Bifurcated Inequality and Political Violence in Latin America

Midlarsky (1988) argues that a bifurcated pattern of land inequality is a principal cause of political violence in Latin America. To test this hypothesis, he summed deaths from political violence (Taylor and Jodice 1983) over the years 1948–77, partitioned them into quintiles, and correlated them with the patterned inequality variable. Before computing the strength of association between these ordinal variables, Midlarsky adjusted the quintiles of deaths from political violence as follows: El Salvador was increased from the lowest quintile to the highest; Nicaragua was increased from the middle quintile to the highest; the Dominican Republic and Bolivia were reduced from the highest quintile to the second highest; Peru and Mexico were reduced from the second highest quintile to the middle; Haiti was reduced from the middle quintile to the second lowest; and Honduras was reduced from the second lowest quintile to the lowest. These changes were justified on the grounds that El Salvador and Nicaragua had experienced intense violence after 1977.

The "adjustments" in the level of political violence introduced by Midlarsky are subjective and arbitrary. They inflate the Tau_b correlation coefficient from .34 for the correct 1948–77 quintiles to .51 for his adjusted quintiles. Although the Tau_b value for the correct 1948–77 quintiles is significant at the .05 level, chi-squared (19.2 with 16 degrees of freedom) is not significant at even the .20 level. Since it is not obvious that patterned inequality is a legitimate ordinal variable, a conservative evaluation of the evidence would give greater credence to the lack of support for the hypothesis from the chi-squared test.

A second problem with Midlarsky's test of the inequality hypothesis entails his measurement of the dependent variable. Although deaths from domestic conflict events is a standard indicator of the magnitude or intensity of political violence in a country, gross death counts are hardly ever used because they are not comparable across countries differing in size of population. The usual procedure is to adjust directly for population by computing a death rate per some unit of population (typically one million population).[4] To construct a summary 1948–77 death rate variable we sum death counts over five-year intervals during that period, divide by midinterval population, and then compute the average. When the mean 1948–77 death rate from political violence is partitioned into quintiles, the Tau_b correlation between it and patterned inequality is .23, a value not significant at the .10 level (chi-squared is not significant at the .50 level).

The most appropriate way to test the hypothesis of a relationship between patterned inequality and political violence would be to dichotomize patterned inequality into the categories *absent* and *present* (or *nearly present*). This procedure at least avoids the level-of-measurement problem, although it does not solve the "apples and oranges" problem of inferring the presence or absence of patterned inequality from different definitions of the upper and lower parts of the land distribution.

The relationship between dichotomized patterned inequality and the 1948–77 political violence death rate is shown in Table 3. The point-biserial correlation (r for a quantitative and dichotomous variable) is quite small and not significant at the .10 level. The weakness of the relationship is illustrated by the fact that a majority of the instances of fits or near fits to patterned inequality—Peru, Chile, and El Salvador—had relatively low death rates from political violence.[5]

The patterned inequality variable is a seriously flawed indicator of bifurcation. Our index of bifurcated inequality is preferable because (1) it is a direct measure of the magnitude of bifurcation between large and small farms, (2) it is based on uniform definitions of the upper and lower ranges of land distribution, (3) it is a quantitative variable not subject to any ambiguity about level of measurement, and (4) it takes into account the proportion of small farms.

A plot of the relationship for Latin American countries between the mean index of bifurcated inequality during the 1948–77 period and the natural logarithm of the mean rate of death from political violence is shown in Figure 1. The death rate is logged in order to reduce the influence of extreme values. The correlation (r) between political violence and bifurcated inequality is .27, a value not significant at even the .10 level.[6] The reason for the weak association is that apart from Bolivia, all countries with very high levels of bifurcated inequality have had relatively low death rates from political violence during 1948–77 (Mexico, Paraguay, and Peru), while many countries with intermediate or relatively low levels of bifurcated inequality have had relatively high death rates from political violence (Argentina, Colombia, the Dominican Republic, Nicaragua, and Venezuela). Evidently there has been no long-term simultaneous association between bifurcated inequality and rates of political violence in Latin America.

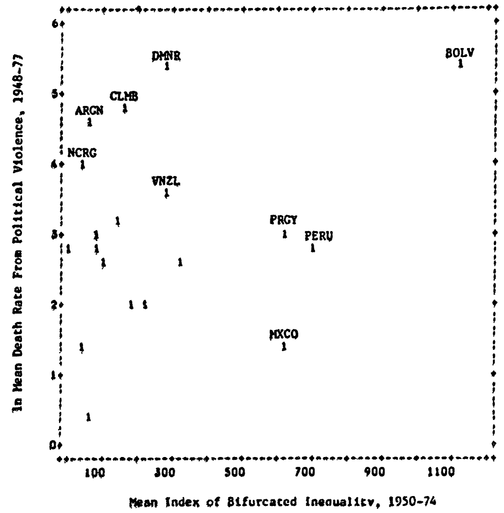
Another question is whether there is any lagged association between bifurcated inequality and political violence. To what extent would knowledge of a country's level of bifurcated inequality at one time predict its subsequent level of political

Table 3. Death Rate from Political Violence by Patterned Inequality: Latin America

Mean Death Rate from Political Violence, 1948–77	Patterned Inequality	
	Absent	Present
225.0		Bolivia
210.8	Dominican Republic	
116.5		Colombia
100.9	Cuba	
89.6	Argentina	
54.5	Nicaragua	
38.1	Venezuela	
21.4	Guatemala	
20.0	Paraguay	
17.2	Panama	
16.8	Honduras	Peru
15.3	Haiti	
12.8	Costa Rica	
12.6		Chile
6.7	Ecuador	
6.4		El Salvador
3.4	Uruguay	
3.2	Mexico	
.5	Brazil	

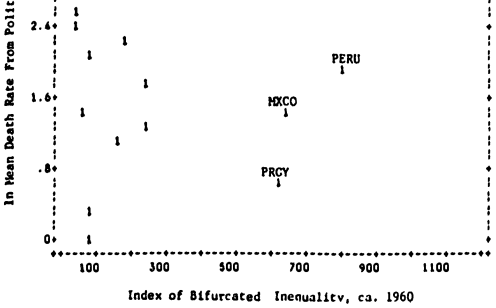
Note: $r = .23$.

Figure 1. Rate of Death from Political Violence by Bifurcated Inequality, 1948–77



violence? Although Midlarsky did not test the hypothesis of a lagged effect of bifurcated inequality on political violence, he stipulated in his theoretical discussion that "some time lag is required for the development of discontent by the peasantry" (1988, 496). Figure 2 shows a plot of the relationship between bifurcated inequality circa 1960 and the logged political violence death rate during 1968–77 for 18 Latin American countries.[7] The time-lagged correlation between political violence and bifurcated inequality is only .11. Argentina and Chile had relatively high rates of political violence during 1968–77 despite low-to-intermediate levels of bifurcated inequality circa 1960; Mexico, Paraguay, and Peru, again, had relatively low rates of political violence despite very high levels of bifurcated inequality; while Bolivia, with by far the highest level of bifurcated inequality, had only the third highest rate of political violence. Thus there is no predictive association between bifurcated inequality and political violence in Latin America.

Figure 2. Rate of Death from Political Violence by Bifurcated Inequality, 1968–77



Discussion

The search for a systematic relationship between patterns of land distribution and political violence has been long and frustrating. Supporters of the hypothesis of a linkage between the two will be disappointed by the results obtained in this analysis. While we have found no evidence to support Midlarsky's claim to have resurrected land distribution from the theoretical scrap heap, we do agree with his conclusion that "if one seeks to disclose relationships between inequality and political violence, the more sensitive procedure and one with stronger systematic variation with the dependent variable is to be preferred" (Midlarsky 1988, 504). This is why we prefer to rely on income inequality, which is not only associated significantly with political violence world-wide (Muller 1985; Muller and Seligson 1987) but also is correlated negatively and quite strongly with regime stability among democratic countries (Muller 1988).

Midlarsky has introduced a new measure of land distribution—patterned inequality—that in his opinion holds a major advantage over the Gini index.[8] We find his measure interesting conceptually but so flawed in construction as to be of little utility. We have proposed a more reliable and, in our opinion, more valid operational definition of bifurcation—a bifurcated inequality index, belonging to the family of measures of extremeness (see Alker and Russett 1966, 353–54)—which, however, was of no use in predicting violence.

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Edward Muller, Mitchell Seligson and Hung-der Fu have rendered a service to the debate over the origins of mass political violence. Their critique led me to discover two minor errors concerning citation of sources. Far more important, their basic misunderstanding led me to the discovery of a pictorial model of patterned inequality that I think goes far to explain both the origins of land distributions and the later appearance of mass political violence in regions with a relatively recent history of conquest and settlement, such as Latin America.

I will (1) directly address the issues raised by Muller and his colleagues; (2) briefly present the pictorial model of patterned inequality that has near-universal—or at least widespread—applicability, as shown not only by evidence presented in the original article but by its universal mathematical base; and (3) introduce arguments and systematic evidence to explain why the Gini index of land inequality has in the past demonstrated a marginal—although typically a significant—relationship with measures of political violence, why patterned inequality is superior, and why bifurcated inequality would not be expected to demonstrate such a relationship.

The Muller-Seligson-Fu Critique

Before proceeding to the heart of the Muller-Seligson-Fu critique and *their* emphasis (not my own) on bifurcated inequality, I should like to dispense with some of their procedural complaints. The first and most important of these is the variation in the upper range of large holdings across countries (Table 1). As was made plain in my article (p. 492), the theory of patterned inequality is based on a fair degree of context dependence. It is clear that given the different social and economic histories of countries and even of regions within a country, the pattern for the upper range will necessarily differ across countries. The differing geographic constraints of the various Latin American countries *must* eventuate in different cutoffs for the upper portion of the distribution and then, by extension, for the lower range as well. Mixing of the upper and lower patterns necessitated the omission of the middle portion in many instances.

Worse yet, Muller and his colleagues fundamentally misunderstand the relationship between the degrees of freedom and the cutoffs for the upper and lower ranges, despite clear indications in the article that such inferences cannot be made. They infer the cutoffs from the reported degrees of freedom for the chi-squared statistics despite the clear statement in footnote *a* to Table 2 (Midlarsky 1988, 498), "Where the expected value was below 1.5, adjacent categories were combined for the chi-square test until that figure was obtained." Such inferences cannot be made, because the degrees of freedom plus two, will actually correspond to the number of categories only where category combinations were unnecessary. Thus, whereas Muller, Seligson, and Fu infer that in the case of Costa Rica the upper range began at 350 hectares, in fact it was 199.5 hectares. In Ecuador and Panama, the upper range began at 20 hectares, not 50 as they inferred. And in Guatemala the land distribution pattern clearly dictated the log-exponential to begin at 1.4 hectares, not 450 as they inferred. For the sake of completeness, Table 4 includes the actual cutoffs for the upper and lower ranges. I excluded them from the original article on the grounds that six tables (plus appendix) were enough.

As can be seen in the fifth column, which gives the proportion of distribution area included in the upper range, it is not the cutoff per se that is important but the proportion of agricultural land area included in this range.[9] The values make in-

Table 4. Cutoffs; Data Sources; and Correlations between the Proportion of Distribution Area in the Upper Range, the Gini Index, and the Index of Bifurcated Inequality

Country	Year	Upper Range Cutoff (hectares)	Lower Range Cutoff (hectares)	Proportion of Distribution Area in the Upper Range ($r_{12} = .85$) (1)	Gini Index[a] ($r_{13} = .45$) (2)	Bifurcated Inequality[a] ($r_{23} = .61$) (3)
Argentina	1960	200.1[b]		89.58	.86	51.1
Bolivia	1950	20[c]	10[c]	99.37	.94	1,129.5
Brazil	1960	50[d]		86.99	.83	70.2
Chile	1950	100[b]	99.9[b]	92.60	.93	296.8
Colombia	1960	50[b]	4.9[b]	75.80	.86	165.2
Costa Rica	1960	199.5[b]		50.77	.79	86.1
Dominican Republic	1950	50.31[e]		53.33	.80	257.9
Ecuador	1954	20[b]		83.37	.86	230.4
El Salvador	1960	50[b]	4.99[b]	57.50	.84	255.4
Guatemala	1950	1.40[b]		96.69	.87	126.9
Haiti	1970	5.17[b]		19.91	.48	.0
Honduras	1960	35[b]		60.12	.75	71.4
Mexico	1960	10[b]		98.81	.95	633.3
Nicaragua	1960	35[b]		85.27	.80	48.2
Panama	1960	20[b]		77.52	.74	54.9
Paraguay	1960	5[b]		99.03[f]	.95	614.8
Peru	1961	50[g]	3[g]	86.21	.94	801.6
Uruguay	1960	200[b]		85.10	.82	37.7
Venezuela	1960	50[b]		92.01	.92	242.2

[a]After Muller, Seligson and Fu's critique, Table 2.
[b]Wilkie and Perkal 1985.
[c]Heath, Erasmus and Buechler 1969, 35.
[d]Roberts and Kohda 1967.
[e]Food and Agriculture Organization 1955.
[f]Area is for 1950, the only data for area given in the source.
[g]Food and Agriculture Organization 1966–70.

tuitive sense on examination. More important, they reveal strong systematic variation with the Gini index.

The criterion for choosing the cutoffs, it must be emphasized, was *not* a uniform numerical value across cases but the point in each instance at which the patterns of the log-exponential and exponential were observed to begin or end. This kind of uniformity is constant in pattern though different in number and is far more faithful to the uniqueness of the historical record than is Muller, Seligson, and Fu's choice of an arbitrary number to be applied across all cases. The issue, in other words, is not the size of a large or small farm (I never used that terminology) but the point of beginning or ending of a particular pattern. The use of the pattern allows for the assessment, by such interested observers as the peasantry, of change over time and direction of change.

It is claimed that I distorted the findings by including Nicaragua and El Salvador in the highest quintile of the 1977 data. (In this connection, I was pleased that Muller and his colleagues confirmed the distinction between the revolutionary patterns for Nicaragua and El Salvador found by Kenneth Roberts and myself.) I regard their concern about the dependent variable to be of secondary importance, because even with their "corrected" data the relationship between patterned inequality and deaths from political violence was statistically significant. Their attempt to discredit their own finding by attacking the ordinal basis of patterned inequality—clearly explicated theoretically in my original article—is arbitrary, to say the least. In any event, the inclusion of Nicaragua and El Salvador in the highest quintile is justified by the arguments developed in my note 7 (pp. 506–7).

During the 1948–77 period, the Salvadoran, and especially Nicaraguan, revolutions were ongoing, with the former just beginning to enter its most violent phase (Leiken 1984) and the latter already well underway (LaFeber 1983). Very likely there was a time lag in the reporting of events—especially deaths due to political violence—that led to their omission from the data source. Indeed they were omitted from the later updated Inter-university Consortium for Political and Social Research listing (until 1982) that I rejected for this and the other reasons indicated in note 7. In any event LaFeber (1983, 239), for one, is clear about the tens of thousands of deaths occurring between 1972 and 1979 that simply did not appear in the 1977 data compilation. A similar argument exists for the Salvadoran revolution. It actually found its beginnings in the formation of the Popular Forces of Liberation (FPL) in 1970 and the gradual insurrection that followed, including the mass killing of demonstrators in San Salvador in 1975 (Leiken 1984, 115–17). Omitting this information would have been tantamount to ignoring the obvious.

The issue of using absolute numbers instead of deaths per unit population is a matter of analytic preference and varies especially according to the theoretical orientation of the analyst. Using death rate as the dependent variable entails the hidden but important assumption of a proportional or linear relationship between the likelihood of deaths due to political violence and population size. This assumption is frequently unwarranted. But if one assumes, as I do, that potentially long-term revolutionary processes—not isolated periodic disturbances—are the likely outcome of patterned inequality and that such processes are driven by a frequently nonlinear dynamic, then population size is much less relevant and may even mask important relationships (as in Muller, Seligson, and Fu's Table 3). Contrast the large numbers of deaths due to political violence in El Salvador or Nicaragua with the much smaller number of deaths in Cuba despite the larger population size of the latter country. The near-genocidal behavior of the revolutionary Khmer Rouge in Cambodia, even against its own people, is another case in point.

The core of the critique, though, is the creation of a measure of bifurcation based on theoretical notions that they inaccurately attribute to me. Perhaps the strength of the original argument may be gauged by my critics' having to generate such an artifact, whose lack of validity was then easy to demonstrate. I never put forward a theory of bifurcated inequality but rather of *patterned inequality*, in which bifurcations were *always* associated with the presence or absence of a particular pattern. This was crucial to the argument, as I made clear throughout the article but especially in the concluding paragraph (p. 505). There the *predictability* (italicized in the original) of the peasantry's impoverishment relative to another, largely unchanging distribution at the upper levels was a key factor in the willingness of the peasantry to engage in political violence.

It was made explicit in the article (pp. 494–98) that these distributions were the consequences of long-term processes that had begun centuries earlier. As a keen observer and interpreter of the environment, the peasant is in a unique position to assess changes or their absence therein. I even included the Irish illustration in my note 9 to demonstrate the subdivision over time in agricultural landholdings at the lower (peasant) levels, in contrast to the absence of such change at the upper levels of the land distribution. This option was not available for the Latin American holdings because of the relative recency of those data. A dynamic aspect found throughout the original article is clearly absent from Muller, Seligson, and Fu's emphasis on bifurcated inequality. The patterns that are the core of my theory are collapsed into two numbers—the means at the upper and lower ends—thus reducing the process of change captured by these patterns into static entities. Not only is the original dynamic theory transformed beyond recognition by Muller and his colleagues, but large quantities of information are lost through this procedure. Distributions with a total of 10 data points at the upper and lower ranges are reduced to 2. This loss of information *in itself* could yield the nonsignificant results they obtain.

In a real sense the essentially static cross-national method of analysis, which has likely contributed much to obfuscate the relationship between inequality and political violence, is here resurrected in order to attack the far more dynamic and process-oriented theory of patterned inequality. I find their critique puzzling because in a fine recent article on democracy and inequality, Muller (1988) explicitly opted for a longitudinal approach to correct the earlier (mistaken in his view) finding of the absence of any relationship between the two variables using a cross-sectional approach.

The Theory of Fractals

The adequacy of any theoretical framework is to be judged by a variety of considerations, not the least of which is the coherence and universality of its analytic core. And here the theory of patterned inequality rests on even stronger foundations than I had earlier suspected. The reason is that the overall pattern I had conjectured has a strong basis in the theory of fractals, which has only recently emerged as a branch of applied mathematics (Mandelbrot 1982). The Pareto-log-exponential distribution found to fit so many of the upper ranges of landholdings (Midlarsky 1988, 498) is a fractal distribution, as are many other natural and social phenomena. To my knowledge, this was the first identification of the upper range of land distributions as fractal, in tandem with the understanding of the upper ranges of income distribution in capitalist societies as Pareto-distributed (Lydall 1968) and therefore also fractal.

The essential characteristic of a fractal is self-similarity. As one reduces or increases the fractal pattern, it remains the same at all levels of observation. The pattern of a coastline and the cumulative winnings of a coin toss gambler with a fair coin are additional illustrations of this regularity (Feller 1968, 87; Mandelbrot 1982, 25–31, 240–41). Encroachments, such as the sea on the coastline or—more to the point here—conquests in Latin America after the Iberian invasions, illustrate this process. Indeed, there exists a model for conquest that predicts a fractal pattern in such instances of steady expansion. The model itself can be generated in an exceptionally versatile manner. All that is required is a specific rule set for steady expansion and a confined space over which the expansion will take place (Barnsley and Demko 1985). Even the random selection of outcomes as in the coin toss illustration, will after a time (say, 50 moves) yield a fractal pattern (Barnsley 1986). A variety of fractal patterns can be generated with very simple rule sets and in a random fashion. I will provide two illustrations, both of which are appropriate as models of conquest.

Consider a rule set for generating a particular configuration on R, a defined space (a plane in this instance, as on a computer screen) and for transitions from one state to the next on R. One example is given by a matrix consisting of a single one, and zeros that can be transformed to ones only under the certain conditions. The rule set G assigns a one if either there was a one in that cell at the previous move or there was a zero in that cell at the previous move but the neighborhood of that cell—consisting of all adjacent cells—now sums to one (mod 2).[10] In all other circumstances the cell entry remains a zero. At each move the cells are chosen randomly (say by a coin toss) as to direction of the move, then the cell entry is evaluated according to the preceding criteria. Eventually, after about 50 or so moves, a pattern begins to emerge not unlike that shown in Figure 3. Willson 1986 contains a more complete description of this procedure.

Figure 3. Fractal Pattern Emerging in the Pascal Triangle

```
1
11
1 1
1111
1   1
11  11
1 1 1 1
11111111
1       1
11      11
1 1     1 1
1111    1111
1   1   1   1
11  11  11  11
1 1 1 1 1 1 1 1
1111111111111111
1               1
11              11
1 1             1 1
1111            1111
1   1           1   1
11  11          11  11
1 1 1 1         1 1 1 1
11111111        11111111
1       1       1       1
11      11      11      11
1 1     1 1     1 1     1 1
1111    1111    1111    1111
1   1   1   1   1   1   1   1
11  11  11  11  11  11  11  11
1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1
11111111111111111111111111111111
1                               1
11                              11
1 1                             1 1
1111                            1111
1   1                           1   1
11  11                          11  11
1 1 1 1                         1 1 1 1
11111111                        11111111
1       1                       1       1
11      11                      11      11
1 1     1 1                     1 1     1 1
1111    1111                    1111    1111
1   1   1   1                   1   1   1   1
11  11  11  11                  11  11  11  11
1 1 1 1 1 1 1 1                 1 1 1 1 1 1 1 1
1111111111111111                1111111111111111
1               1               1               1
11              11              11              11
1 1             1 1             1 1             1 1
1111            1111            1111            1111
1   1           1   1           1   1           1   1
11  11          11  11          11  11          11  11
1 1 1 1         1 1 1 1         1 1 1 1         1 1 1 1
11111111        11111111        11111111        11111111
1       1       1       1       1       1       1       1
11      11      11      11      11      11      11      11
1 1     1 1     1 1     1 1     1 1     1 1     1 1     1 1
1111    1111    1111    1111    1111    1111    1111    1111
1   1   1   1   1   1   1   1   1   1   1   1   1   1   1   1
11  11  11  11  11  11  11  11  11  11  11  11  11  11  11  11
1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1 1
1111111111111111111111111111111111111111111111111111111111111111
```

Source: Chaitin 1987, 12.

The expanding ones can be viewed as *surrounding and occupying* choice territories (zeros, shown in the figure as blank spaces) for their own use as, say haciendas, while militarily *dominating* the remaining indigenous population centers (signified by the clusters of ones). Thus, there continue to exist in the clusters of ones militarily dominated local communities, and the conquerers make use of their tracts of land—large and frequently depopulated by disease—as they please.

The triangle below was generated by a different rule set (the Pascal triangle) that nevertheless obeys the general criteria, mentioned earlier, of steady expansion on a defined space. We begin with ones in the following positions:

1
1 1

Each entry will be the sum of the number directly above it (zero if none is explicitly there) and the number at its upper left. Thus, after several moves, one obtains

1
1 1
1 2 1
1 3 3 1
1 4 6 4 1

The pattern moves to the right and is not immediately obvious until all even numbers are treated as zeros and all odd numbers are ones, thus yielding Figure 3

(Chaitin 1987, 8-11). The zeros now occupy a set of triangles with a ratio of areas that is fractal, in this case Pareto-log-exponential. Zeros are removed from Figure 3, leaving blank spaces, to highlight the fractal pattern.

Note that the same or a similar pattern can be achieved either randomly, as in the first illustration, or by nonrandom methods, as in the second. In either case, a fractal pattern of triangles emerges with the precise ratio of areas—or, in this case, inequality relations—of the Pareto-log-exponential distribution. These are but two illustrations of the many (perhaps infinite) ways of yielding fractal patterns under very simple rule sets of expansion or conquest over a particular domain R.

Now the triangular pattern in Figure 3 is just as interesting for what it does *not* explicitly show as for what it does. In addition to the fractal pattern yielding a Pareto-log-exponential distribution, there exists the remainder of the area, indicated by ones, unoccupied although dominated militarily by the conquerer. This area is subdivided into triangles of equal area that would presumably be left to the indigenous peasantry after such a conquest. (It is immaterial whether one includes the isolated zeros in the "indigenous" triangles or opts for smaller triangles exclusive of the zeros; whichever option is chosen, all of the triangles have equal areas.) The peasantry would initially have equal land areas, which then, via the reproduction process and land subdivision, would yield the exponential distribution for the lower range of holdings. This process of subdividing initially equal holdings, thus yielding the exponential distribution, was conjectured earlier (Midlarsky 1982; Midlarsky 1988, 493-4; Midlarsky and Roberts 1985), but until the discovery of this large and rapidly growing body of applied mathematics it remained a conjecture with some empirical support but without direct analytic confirmation as shown here in Figure 3.

It must be emphasized that fractal patterns are the *typical* outcomes of random processes that obey rule sets along the lines of the two preceding illustrations (Barnsley 1986; Barnsley and Demko 1985; Willson 1986). Now it becomes clearer why so many of the distributions in the upper range of Latin American landholdings conformed to the log-exponential distribution (Midlarsky 1988, 498). These distributions resulted from conquest or expansion in one form or another, which, according to the fractal model, should yield the Pareto-log-exponential distribution. It is also clear why the cutoffs for the upper range *must* vary across the individual Latin American land distributions. Each of the fractal patterns emerges after expansion over a uniquely defined space, R, so the pattern that emerges is uniquely dependent on that space. Topography and any other environmental or social constraints affecting the direction of conquest also yield different fractal patterns that will inevitably result in different cutoffs for the patterns at the upper range. Cutoffs for the lower range also will vary now as the result of variations in the fractal pattern that determine the overall pattern of landholdings.

Understanding the Gini Index and Other Measures

It is interesting that despite the wide variation of upper-range cutoffs across countries—understood in the context of the theory of fractals—the proportions of land area resulting from these cutoffs show strong systematic variation with other measures, especially the Gini index. Table 4 gives Pearson correlation coefficients between the proportions of the total land area included in the upper range and the Gini values calculated by Muller and his colleagues in Table 2. This is an approximate analogue to the use of the upper 20% of the income range as a measure—really a surrogate measure—of inequality (e.g., Muller and Seligson 1987). The percentage of the variance held in common by the Gini index and proportion of land area included in the upper range is 72%, a rather substantial value. This finding, of course, empirically refutes one of the more serious claims of Muller and his colleagues, namely, that the variation in cutoffs for the upper range is without systematic foundation. Note that the Gini values are those given in Muller, Seligson, and Fu's critique. This condition stands in contrast to both the 37% of the variance held in common between their measure of bifurcation and the Gini index and the even-more-substantially-reduced 20% of the variance shared by the bifurcation index and the proportion of land area in the upper range (see Table 4).

These findings have several implications in addition to that of the systematic variation of the proportion of land area in the upper range. First, the Gini index is dominated by the upper range of holdings—as I suspected, given the vast difference in inequality between the log-exponential (upper) and exponential (lower) portions shown in the appendix (Midlarsky 1988, 505–6). Thus, the Gini index will tap the grosser sorts of inequalities inherent in the upper range, but largely ignore the pattern (mostly exponential) in the lower range under conditions of heavy land subdivision. But this is precisely the range of holdings that includes the peasant majority of the agricultural population that will be available for recruitment into revolutionary movements. The upper range of holdings, by definition, explicitly ignores this critical popular component. Given the correlations in Table 4, neither the Gini index nor the bifurcated inequality index would be sensitive to this popular sector, which is explicitly and separately taken into account by the patterned inequality measure. Apparently, the bifurcated inequality index is insensitive to all of the validation measures used both in this response and in the Muller, Seligson, and Fu critique. Patterned inequality, on the other hand, examines expected and observed patterns at both ends of the distribution range, thus accounting for inequalities across the broad spectrum of both landlord and peasant.

In Table 4, which includes cutoffs for the upper and lower ranges and data sources for each country, two corrections are incorporated. First, the data for the Dominican Republic are from FAO 1950, not FAO 1960 as reported in the original article. The earlier date was chosen as the earliest most complete data source (to allow time for the peasantry to become mobilized), and the source was recorded as 1960 by an oversight. Second, the words *or earlier* were mistakenly omitted from the reference to the UCLA Latin American series cited in Table 2 of the original article. This omission led to the erroneous impression by Muller and his colleagues in note 1 that the upper range for Brazil was somehow misspecified. The earlier source for Brazil in the UCLA series is cited in Table 4 here.

Finally, a comment is in order on the Muller, Seligson, and Fu's preference for analyzing income inequality instead of land inequality. I have never juxtaposed land and income inequality and opted for one over the other as a source of political violence. Income inequality, indeed, may be related to political violence in ways not yet explored. At this time this is a subject of some controversy (e.g., Weede 1986, 1987).

My preference for land inequality as an initial focus of investigation emerges from several considerations. First, land is generally a fixed resource in mostly agrarian societies that can lead to zero-sum games between landlords and peasants. Income, on the other hand, can be generated in expanding economies—typically industrial—to avoid the zero-sum aspect that likely is critical to the onset of political violence (e.g., Midlarsky and Roberts 1985). Second, as a consequence of the relative fixedness of land area, one can make the scarcity assumption that is necessary for the deriviation of both the exponential and log-exponential distributions. In contrast, the assumption of proportional growth and a consequent log-normality have been found to fit the lower and middle ranges of income distributions, thus necessitating an entirely separate analysis of income inequality in relation to political violence. Clearly, this additional research should be carried out. It will very likely clarify the important, but as yet poorly understood, role of income inequality as a progenitor of societal discontent.

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Notes

1. We are unable to determine the upper range for Brazil because Midlarsky stated the degrees of freedom as three—implying five data points—whereas in Wilkie and Perkal 1985 only four data points are available for estimation.

2. A comparison of Costa Rica and El Salvador illustrates the point. Both countries have similar land bifurcation ratios. But in Costa Rica, where many farms are middle-sized (between 5 and 99.9 ha.), small farms are a relatively small proportion of the total; whereas in El Salvador the distribution of land is such that most farms are either relatively large estates or small subsistence plots, and small farms make up approximately four-fifths of the total (twice that in Costa Rica). Therefore, to construct a valid operational definition of the concept of bifurcated inequality, i.e., a condition in which there are a few great landlords, many poor peasants, and few middle-sized "yeoman" farmers, it is necessary to adjust the land bifurcation ratio for differences in the extent to which agrarian society is polarized between lord and peasant.

3. All data are from Tables 15.1, 15.2, and 15.3 of FAO 1981 except for Haiti, where the data are from Tables 2.2, 2.3, and 3.2; and Argentina, Bolivia, Chile, Guatemala, Honduras (1966 only), Nicaragua, and Paraguay, where the data are from Tables 302 and 303 of Wilkie and Perkal 1985. Some of the distributions reported in Wilkie and Perkal cannot be compared exactly with the standard size categories of the FAO, so we have used the nearest approximation. In the cases of Guatemala, Honduras, and Nicaragua the upper limit of small farms is 6.9 ha.; in the case of Guatemala, the lower limit of large farms is 44.8 ha.; in the cases of Honduras and Nicaragua, the lower limit of large farms is 70 ha.

4. An alternative but less-desirable procedure is to include population as an independent variable in a regression equation. (For critiques of this approach see Linehan 1980, 190–93; Muller 1985, 51–52).

5. Note that El Salvador ranks in the lowest quintile of political violence even when the death rate is used in place of simple death counts.

6. If the rate of death from political violence is not logged, the correlation is .42, a value marginally significant (.10 level). The moderate strength of the association is entirely due to the extreme case of Bolivia, however. If Bolivia is excluded, the correlation drops to −.06.

7. The years range from 1954 to 1966 except for Bolivia, where the 1950 score is used as an estimate of bifurcated inequality circa 1960. Information on bifurcated inequality circa 1960 is not available for Haiti.

8. Curiously, Midlarsky (1988, 504) claims that the Gini concentration ratio "is simply a mathematical measure (the area between one curve and another), without theoretical content." By contrast, in Alker and Russett's (1966, 358–62) comprehensive review of measures of inequality, Gini is considered to be one of the few measures with "explicit theoretical norms."

9. In Table 4, the cutoff for the upper range signifies all of the area above the tabled value, and the cutoff for the lower range signifies all of the area below the tabled value. In the actual testing, as in Tables 1 and 4 of the original article, the highest land area category listed in the source was excluded because it was open-ended and its midpoint therefore could not be estimated. However, this area is included in the upper range *pattern* by implication, and therefore is included in the values of the fifth column for proportions of the distribution area in the upper range.

10. The term *mod 2* refers to a cardinal modulus that gives the remainder after division by 2. Thus 5 mod 2 is equal to 1. In the illustration given in the text, all summations are divided by 2 and the cell entries therefore are either 1 or 0, depending respectively on whether the summations are odd or even. See Messer and Marshall 1986 (pp. 55–56) for more details.

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Forthcoming in September

The following articles, research note, and controversies, have been tentatively scheduled for publication in the September 1989 issue:

Brian J. Cook and B. Dan Wood. "Principal-Agent Models of Political Control of Bureaucracy." A Controversy.

Robert S. Erikson, Gerald C. Wright, and John P. McIver. "Partisan Elections, Public Opinion, and State Policy."

Steven E. Finkle, Edward N. Muller, and Karl-Dieter Opp. "Personal Influence, Collective Rationality, and Mass Political Action."

Charles H. Franklin and Leane C. Kosaki. "The U.S. Supreme Court, Public Opinion, and Abortion."

Micheal W. Giles and Thomas D. Lancaster. "Political Transition, Social Development, and Legal Mobilization in Spain."

Richard Gunther. "Electoral Laws, Party Systems, and Elites in Spain."

Gary C. Jacobson. "Strategic Politicians and the Dynamics of House Elections, 1946–86."

Charles O. Jones. "Mistrust but Verify: Memoirs of the Reagan Era." A Review Essay.

Kenneth Koford. "Dimensions in Congressional Voting." A Research Note.

Paul E. Peterson and Mark Rom. "Federalism, Welfare Policy, and Residential Choices."

Paul J. Quirk, "Cooperative Resolution of Policy Conflict."

Philip G. Roeder. "Modernization and Participation in the Leninist Developmental Strategy."

Barry R. Weingast. "Floor Behavior in Congress: Committee Power under the Open Rule."

Donald Wittman. "Arms Control Verification and Other Games Involving Imperfect Detection."