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Consumption and Income Inequality During the Transition to a Market Economy: Poland, 1985–1992

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This paper challenges the conventional wisdom that income and consumption inequality in Poland increased substantially following the economic transition in 1989–90. Using microdata from the 1985–92 Household Budget Surveys, we find that overall income inequality increased in 1989 but subsequently declined to pre-transition levels. The distribution of consumption reveals a similar pattern. Social transfers are shown to have played an important role in mitigating increases in overall income inequality during the transition. However, the relative well-being of different socioeconomic groups was altered and, despite the reasonably good targeting of transfers, there were clear winners and losers in the transition process. [JEL D31, J31, O15]

Poland experienced a sudden economic transformation in late 1989 and early 1990 that has become known as the "big bang." The noncommunist government that took power in 1989 ended food price controls in August 1989 and ended price controls on most other products in January 1990. This led to substantial inflation and changes in relative prices. Other aspects of the reforms, including reductions in state orders for manufactured goods and restraints on credit for state-owned enterprises,

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along with external shocks such as increased import competition and the collapse of the Council for Mutual Economic Assistance (CMEA) trade bloc, also contributed to large declines in real GDP (of 11.4 percent in 1990 and 7.0 percent in 1991 according to IMF estimates).[1]

The conventional wisdom is that the process of transition to a market economy has been accompanied by great increases in income inequality, both in Poland and in most of the other formerly centrally planned economies of Eastern Europe. For instance, in a cross-country study, Milanovic (1998) reports that, between 1987–88 and 1993–95, the Gini coefficient for household per capita income rose in 17 of 18 Eastern bloc countries. He notes that the average Gini increased from 0.24, a level similar to that in the Scandinavian and Benelux countries, to 0.33, a level similar to that in Canada and the United Kingdom. To put such an increase in historical perspective, it is roughly three times as great as the increase reported for the United States in the 1980s by Atkinson, Rainwater, and Smeeding (1995). For Poland, the Organization for Economic Cooperation and Development (OECD, 1997) reports that the Gini increased from 0.249 in 1989 to 0.290 in 1993, after which it stayed relatively flat through 1996.[2]

In this paper, we provide new evidence on changes in inequality in Poland during the transition. The main difference between our work and that of previous authors (reviewed in Section I) is that we have obtained for the first time direct access to the detailed microdata of the Polish Household Budget Survey (HBS) conducted by the Polish Central Statistical Office (CSO)[3] for the years 1985–92.[4] Prior work on inequality in transition economies has been based primarily on aggregate data about income distributions that are published by the statistical bureaus of the various countries. But, as we discuss in Section II, the published aggregate income data for Poland and other transition economies do not correspond to conventional economic measures of household income. However, at least for Poland, meaningful income measures can be constructed using the household level microdata.

Using the HBS microdata, we find no evidence that income inequality increased in Poland in the first three years following the big bang. For instance, we find that Gini coefficients actually declined from 1989 to 1992. Interestingly, while our Ginis for 1992 are quite similar to those reported by the CSO and OECD, we obtain much higher Ginis for the pre-1990 period. We conclude that

[1]There is some controversy over the relative importance of various factors in generating the output decline in Poland. Calvo and Coricelli (1992) and Commander and Coricelli (1992) stress the contraction of credit to state enterprises. Because most of these enterprises are loss making, a contraction in credit would force them to reduce their scale of operation. On the other hand, Berg and Blanchard (1994) argue that an aggregate demand contraction was the more important cause of the output decline. The basis for this claim is the finding that finished goods inventories increased after the big bang.

[2]Milanovic (1998) reports that the Gini for Poland increased from 0.256 in 1987 to 0.284 in 1993 (first half). This is somewhat smaller than the increase implied by the OECD figures, but nevertheless substantial by historical standards.

[3]Or, in Polish, the *Główny Urząd Statystyczny*, commonly referred to as GUS.

[4]At the time we began our study, the Polish CSO had never before released the HBS microdata. A long negotiation process by the first author during 1992–93 led to its release. Subsequently, the microdata for just the first six months of 1993 were released to the World Bank, and these data are used in World Bank (1995) and Milanovic (1998).

the published aggregate statistics seriously understate the degree of inequality that existed before the big bang. As a result, most of the post–big bang increase in inequality that is present in the aggregate statistics appears to be spurious.

In the HBS microdata we are able to distinguish between pre- and post-transfer income. We find that inequality in pre-transfer income did in fact increase substantially in the transition. Thus, it appears that transfer programs were quite successful in mitigating any increases in inequality. We find that these programs are well targeted in the sense that most transfers go to those at the low end of the income distribution. This is true even though transfer programs in Poland, as in other transition economies, tend to be based on class rather than income.

Another important difference between our work and that of previous authors is that we examine consumption inequality as well as income inequality. To the extent that households can smooth consumption over time, consumption inequality is certainly a more interesting measure. It is again our access to the detailed microdata that allows us to examine consumption inequality in a meaningful way. As we discuss in Section II, the aggregate consumption figures that were published by the Polish CSO, as well as by other former communist countries, did not correspond to conventional economic measures of consumption. After constructing reasonable consumption measures from the microdata of the HBS, we again find no evidence of increased inequality during the transition.

One reason for the interest in the changes in inequality that may be occurring in transition economies is that, to the extent that inequality has been increasing, it may create social unrest and political pressures that could stall the transition process. Our results suggest that, at least in Poland, such concerns may have been exaggerated. The existing social safety net appears to have done an adequate job of limiting the impact of transition on inequality.

Although we find no evidence of increases in overall inequality, our access to the HBS microdata enables us to examine whether certain socioeconomic groups have been relative winners or losers in the transition. We find that the transition did have significant distributional impacts across broadly defined socioeconomic groups. Some groups also experienced large increases in within-group inequality. For instance, among households for which labor income is the primary source of income, income differentials by education level of household head increased rapidly after the big bang. Gorecki (1994) previously noted such a pattern in the aggregate data released by the CSO. Before the transition, the wage structure in Poland was highly compacted, with wages of college-educated white collar workers little different from those of manual workers. Soon after the big bang, those with a college degree became much more concentrated in the upper quantiles of the income distribution, while those with only primary education became much more concentrated in the lower quantiles. Such a widening of across-group income differentials is to an extent desirable, as it implies an enhanced incentive for human capital investment. But it also raises concerns that dissatisfaction and social unrest may be a problem among those groups that have fared poorly.

In the next section, we describe the prior research on income inequality in Poland during the transition in more detail. Then, in Section II, we describe the HBS data. As we explain there, the Polish HBS is of higher quality and was collected according to a more consistent methodology over the transition period than the microdata for any of the other former communist countries. Thus, while the Polish case is interesting for its own sake, an analysis of the HBS data also provides the best hope for arriving at conclusions about the effects of transition on consumption and income distributions that may be generalizable.

In comparing the relative welfare of households with different levels of income or consumption, an important consideration is that an adjustment needs to be made for household size and, more generally, for the demographic composition of households. Most previous studies on inequality in transition economies have used per capita measures or equivalence scales constructed using industrial country data. An additional contribution of this paper is the construction of a full set of equivalence scales for Poland, which differ in some important respects from those based on industrial country data.

Section III describes our procedure for constructing equivalence scales. Section IV presents our main empirical results on the evolution of inequality. Section V analyzes income and consumption mobility. Section VI concludes.

I. Review of Prior Research

Several other studies have examined income inequality in Poland during the transition. But they report rather contradictory results, even though they all use income data from the HBS. For instance, OECD (1997, Figure 22, p. 86) reports that the Gini based on household per capita income for Poland is 0.25 for 1989, drops to 0.23 in 1990, and then rises substantially to 0.26, 0.27, and 0.29 over the period 1991–93. In contrast, Gorecki (1994) also finds a drop in inequality from 1989 to 1990, but finds no evidence of a subsequent increase in 1991. Similarly, Milanovic (1993) reports Gini values of 0.260, 0.255, and 0.247 for 1989–91. Thus, the OECD figures imply a very large increase in income inequality in 1991, while the Milanovic and Gorecki figures do not show this. The OECD (1997) and Milanovic (1998) figures are consistent, however, in implying that large increases in inequality had occurred by 1993.

The prior studies were based on aggregate statistics published by the CSO, with the exception of Milanovic (1998), who had access to the microdata for just the first six months of 1993.[5] The Gini values in the studies cited above were thus approximated using aggregate data on the income distribution published by the CSO in the annual publication *Budzety Gospodarstw Domowych*, which we henceforth refer to as the *Surveys*.[6] The accuracy of these approximations is certainly subject to question.

[5]A more recent paper by Torrey, Smeeding, and Bailey (1999) uses a sample that constitutes about 45 percent of the full HBS sample and that is available through the Luxembourg Income Survey (LIS), but only for selected years. Using the LIS data, these authors report income Gini coefficients of 0.217 for 1987, 0.248 for 1990, and 0.242 for 1992.

[6]The *Surveys* report the number of households in each of several per capita income ranges, along with the average per capita income within each range and the average number of persons per household within each range. The number of income ranges reported differs by year. This difference in reporting may itself account for some change in the Gini over time.

A more important point is that the aggregate income statistics reported by the CSO, as well as those reported in household budget surveys done in other former communist countries, differ in a number of important ways from measures of income that would be considered economically meaningful in the West. For example, for farmers, income includes gross farm revenues, rather than net revenues. This is an important problem, because approximately one-fourth of Polish households are either farm households or mixed farmer/worker households. In light of this, one must question any results on income inequality based on the aggregate data. Because we have access to the detailed microdata, we are able to make important adjustments to income in order to obtain a meaningful measure (in this example, by calculating net farm income).[7]

Furthermore, the aggregate consumption figures published by the Polish CSO, as well as by other former communist countries, do not correspond to Western-style measures of consumption. Rather, they correspond to something like total money outflows. For instance, for farm households, consumption includes farm investment and purchases of supplies. An indication of the strange nature of the aggregate consumption data is provided by Milanovic (1998, p. 41), who reports that for 1993 the Gini for consumption is 0.31, which substantially exceeds the Gini of 0.28 that he calculates for income. Also, on page 33 he reports that for 1993 the ratio of consumption to income is 1.30, an unreasonably high figure.

It is again our access to the detailed microdata that allows us to examine consumption inequality in a meaningful way. Once we make necessary adjustments to the categories that are included in consumption, we find the more plausible results that consumption Ginis are generally smaller than income Ginis and that the aggregate consumption to income ratio falls in the 0.894 to 0.955 range during 1985–92.

Note that previous research on inequality in Poland and other transition economies has relied almost exclusively on Gini coefficients to measure inequality. In this paper, we provide a more detailed characterization of changes in the income and consumption distributions. We examine alternative entropy measures besides the Gini, we examine quantile ratios, and we examine kernel density estimates of the income and consumption distributions. In addition, prior studies have generally used household per capita income rather than accommodating household economies of scale by using equivalence scales. We examine the sensitivity of our results to choice among a number of alternative equivalence scales.

II. The Household Budget Surveys

The Polish Central Statistical Office has been collecting detailed microdata on household income and consumption at least since 1978, using fairly sophisticated sampling techniques. In the Polish HBS, the primary sampling unit is the household. A two-stage geographically stratified sampling scheme is used, where the first-stage

[7]It is possible to make some (but not all) of the necessary adjustments to income using information in the aggregate data on categories of income. Inconsistencies in the set of adjustments actually made may account for some of the discrepancies in Gini values reported in previous studies.

sampling units are the area survey units and the second-stage units are individual households. Households are surveyed every month for a full quarter in order to monitor their income and spending patterns, and supplementary information is collected from these households once every year. A certain fraction of the households interviewed in a quarter are interviewed in the same quarter of the following year, thereby adding a limited panel aspect to the data. The typical sample size is about 25,000 households per year (6,250 per quarter). The CSO uses the data obtained from these household surveys to create aggregate tabulations that are then presented in its monthly and annual Statistical Bulletins, or *Surveys*.

The HBS contains very detailed information on consumption. We have aggregated across many of the very detailed consumption categories provided in the surveys to classify total household expenditure into these 16 categories: (1) food, (2) alcohol and tobacco, (3) clothing and footwear, (4) house purchases, (5) house construction, (6) household nondurables (including energy), (7) household durables (including furnishings, appliances), (8) rent, (9) health, (10) hygiene, (11) education, (12) "cultural" durables (radio, TV, sporting goods, etc.), (13) recreation and tourism, (14) vehicles, (15) transportation, and (16) other expenditures. In this paper, we use a coarser breakdown in which the nondurable components of categories 4 through 16 are aggregated into two categories: nonfood commodities and services.

Information on sources and amounts of income is available for both households and individuals within each household. Total income is broken down into four main categories: (1) labor income (including wages, salaries, and nonwage compensation), (2) pensions, (3) social security and other transfers, and (4) other income. For farm households, farm income and expenditures, as well as consumption of the farm's produce, are also reported. Finally, the HBS also contains information on characteristics of the dwelling, stocks of durables, and demographic characteristics of all household members.

Using information obtained from other CSO publications and IMF databases, we have also extracted time series on prices corresponding to each of our 16 expenditure categories, as well as the nonfood commodities and services groupings mentioned above. Hence, we have been able to construct disaggregated measures of real consumption for each year.

To put the quality of the Polish HBS data in context, it is useful to discuss the limitations of the data sources available for other former communist countries. As discussed by Cornelius and Weder (1996), the Family Budget Surveys (FBSs) collected in the Soviet Union suffer from a number of severe problems. First, the data are not a representative sample of the population (because families were selected mainly on the basis of the industrial affiliation of their wage earners). Second, the income data are grouped, so only the fraction of the sample with income in various intervals is known. Thus, the FBSs do not provide true household- or individual-level income data.

After the breakup of the Soviet Union, some of the former Soviet states maintained the same primitive data collection methods, while others (including all of the Baltic states) adopted improved sampling methods in which individuals were chosen from the population register, with gender, age, and household size used as stratifying criteria. In either case, looking at changes in distributions over the transition period is problematic—in the former case because the data are poor throughout, in the latter because the improved data from after the breakup are not comparable to the Soviet-era data. Similarly, the Hungarian income data suffered from a substantial change in methodology in the early 1990s. And based on Flanagan (1995), it appears that data collection efforts in the Czech Republic have been sporadic over time.

In contrast, the Polish CSO remained well funded throughout the transition period, and collection of the HBS data using a fairly consistent methodology continued throughout the transition and continues today. For this reason, the HBS offers the highest quality and most consistent microdata available for any of the former communist countries. Thus, it provides the best hope for arriving at conclusions about the effects of economic transition on consumption and income distributions that may be generalizable.

This is the first study based on micro-level data from the HBS for years both before and after the big bang. Other researchers who previously used the data (such as Gorecki, Milanovic, and Szulc) had to either work with the aggregated information published by the CSO in the *Surveys*, submit requests for the CSO to calculate certain statistics for them, or work onsite at the CSO. This greatly limited the kind of analysis that was feasible, for obvious reasons.

Basic Statistics

We begin by presenting some basic statistics for Poland in the 1985–92 period. Table 1 reports changes in aggregate GDP, imports, exports, and consumption, as taken from the IMF *International Financial Statistics*, along with average household income and consumption, as taken from the HBS. A striking aspect of the aggregate data is that per capita consumption actually fell more than GDP in 1990 (–23.8 percent vs. –11.4 percent). Thus, there was no aggregate smoothing of the

Table 1. Selected Macroeconomic Indicators for Poland
(Annual percentage changes)

	1986	1987	1988	1989	1990	1991	1992
Aggregate Data							
Real GDP	4.2	2.1	4.0	0.3	–11.4	–7.0	2.6
Real consumption per capita	5.0	1.2	3.8	–0.3	–23.8	4.6	3.0
Import volumes	4.9	4.5	9.4	1.5	–17.9	37.7	13.9
Export volumes	4.9	4.8	9.1	0.2	13.7	–2.4	–2.6
Consumer price index	16.5	26.4	60.2	251.1	585.8	70.3	43.0
Employment (year end)	0.3	0.0	–1.0	–0.8	–6.2	–3.9	–3.1
Household Survey							
Per capita real income	1.7	–2.4	7.4	7.7	–25.9	3.3	–2.1
Per capita real total consumption	0.2	–0.2	4.6	3.6	–23.8	3.0	0.0

Notes: Aggregate data were obtained from various publications of GUS and the IMF. Aggregate consumption is deflated by the CPI. Net income and total consumption from household surveys are also deflated by the CPI.

adverse income shock, as is reflected by the large *decrease* in net imports. But, in 1991, consumption begins to bounce back (+4.6 percent) even as GDP continues to fall (–7.0 percent). This change is reflected in the very large increase in net imports. It is comforting that the HBS data show a similar pattern of consumption in 1990–91.

Table 2 reports a list of variables that we use extensively in our analysis, along with their overall means in the HBS. The total number of observations across all eight years from 1985 to 1992 is 203,620. Note that the mean of total real consumption is 149,610, and the mean of total real income is 161,574, where both variables are deflated by the aggregate CPI. The ratio is 0.926 (not shown in table), which seems reasonable. The sample is 50 percent urban, and 57 percent of the household heads are males in the 31–60 age range. Fifty-four percent of the households include a married couple (not shown in table), and the mean household size is 3.22. There are seven education categories reported, and the most common education levels for the household heads are primary school (35 percent), basic vocational training (31 percent), and high school or equivalent vocational training (19 percent).

An interesting feature of the HBS data is that they contain information on whether households own each of a list of 21 durable goods at the start of the interview period, and whether their house or apartment possesses each of 5 fixtures. In Table 2, we list the percentage of households with each of the 5 fixtures. Overall means for the durable stocks are not very meaningful because many of them change drastically over time.

III. Equivalence Scales

As noted above, most of the prior work on income distributions in Poland has simply looked at per capita household income, and has not attempted to account for household economies of scale by employing equivalence scales. The exception is the work by Szulc (1994, 1995), which analyzes poverty rates. He calculates equivalence scales based on estimating a demographically flexible Almost Ideal Demand System (Deaton and Muellbauer, 1980) for four categories of consumption using the microdata from several years of the HBS.[8]

We were concerned about estimating a complete demand system under conditions when rationing of certain commodities was probably an issue in some years, but where we do not observe the rationing regimes.[9] Thus, we choose to adopt the simpler Engel (1895) method, the basic idea of which is to assume that two households with the same food share are equally well off. Thus, implementation requires only the estimation of the food share equation, rather than a complete demand system. We examine food shares out of total nondurable consumption, because in Poland rationing was far more prevalent for durables than for other goods.

[8]The years are 1980–82, 1984–89, and 1990–92. The categories are (1) food, alcohol, tobacco, (2) clothing, footwear, hygiene, medical services, (3) house expenses and energy, and (4) transportation, education, entertainment, and other.

[9]Deaton (1981) discusses estimation of demand systems with known rationing regimes.

If durables are weakly separable from other goods in the utility function, then expenditure on durables only has an income effect on other demands, and this procedure is appropriate (see Pollak, 1971). Given the food share equation estimates, we obtain the equivalence scale as the relative expenditure necessary for a household of any given composition to achieve the same food share as a base household.

The almost ideal demand-type food share equation is

$$w_h = \alpha_1 + \beta \ln (x_h/k_h P) + \Sigma\, \gamma_{1j} \ln p_j, \tag{1}$$

where w_h is the food share of household h, x_h is nondurable consumption, k_h is the equivalence scale, and the p_j are the prices of food ($j=1$) and other goods. The other goods categories that we include are alcohol and tobacco, nonfood commodities, and services. Note that the γ_{1j} must sum to zero to satisfy zero degree homogeneity in prices and total expenditure. The P is an aggregate price index defined by $\ln P = \alpha_0 + \Sigma \alpha_k \ln p_k + (1/2)\Sigma\Sigma\, \gamma_{kj} \ln p_k \ln p_j$. But, as noted by Deaton and Muellbauer (1980), share-weighted aggregate price indices will tend to be highly correlated with P. Thus, we estimate equation (1) by replacing P with the aggregate price index for nondurable commodities (P^*) obtained from the *Surveys.* Imposing zero degree homogeneity and substituting the aggregate price index, we obtain

$$w_h = \alpha_1 + \beta (\ln x_h - \ln P^*) - \beta \ln k_h + \Sigma_j \gamma_{1j} (\ln p_j - log p_1). \tag{2}$$

We then specify $-\beta \ln k_h = \Sigma_j\, \phi_j D_{hj}$, where the D_{hj} are dummy variables indicating whether household h has characteristic j, where j indexes the set of demographic categories listed in Table 2. A base household consisting of a married couple with no children or other adults present, and where both the husband and wife are in the 31–60 age range, forms the omitted category. We estimated equation (2) including quarter dummies. Given the estimated food share equation, we estimate the equivalence scale k_h as

$$k_h = \exp[-\Sigma_j \phi_j D_{hj}/\beta]. \tag{3}$$

Note that the equivalence scale k_h equals one for the base-type household.

A potential problem with estimation of equation (2) is that denominator bias is present if nondurables consumption is measured with error. Thus, we have estimated equation (2) using both OLS and 2SLS. In 2SLS the instruments for $\log C_h$ are (1) the set of 18 household demographic dummies, (2) the education-level dummies for the household head, along with age and age squared of the household head, (3) an urban dummy, (4) the 21 durable holding dummies, (5) the five household fixture dummies, and (6) quarter dummies (to capture seasonals in tastes for food consumption). The first-stage regressions were run separately by year, and their R^2 values range from 0.64 to 0.84.

Table 3 reports OLS and 2SLS estimates of the food share equation. Note that the coefficient on log real nondurable consumption changes from –0.195 to –0.263 when we instrument. In the 2SLS regression, the three relative price terms taken together imply a coefficient of 0.190 on the log price of food. This implies that a

Table 2. Summary Statistics

	Mean	Standard Deviation
Real household income		
Total	161,574	127,026
Labor income	81,910	82,632
Transfers	39,728	36,906
Farm income	30,724	109,567
Other income	9,212	39,766
Real household consumption		
Total	149,610	102,273
Durables	19,035	59,106
Nondurables	130,575	69,207
Food	71,369	33,290
Household characteristics		
Urban	0.50	0.50
Number of persons in household	3.22	1.62
Primary income source of household		
Workers	0.49	0.50
Farmers	0.11	0.32
Farmers/workers	0.12	0.32
Pensioners, others	0.29	0.45
Household head characteristics		
Male, 18–30	0.11	0.31
Male, 31–60	0.57	0.49
Male, > 60	0.15	0.35
Female, 18–30	0.01	0.09
Female, 31–60	0.08	0.28
Female, > 60	0.08	0.28
Age	48.35	15.15
College degree	0.06	0.24
Some college	0.00	0.07
High school	0.19	0.39
Some high school	0.02	0.12
Basic vocational training	0.31	0.46
Primary school	0.35	0.48
Primary not completed	0.07	0.25
Demographic characteristics of other members of household		
Wife, 18–30	0.37	0.48
Wife, 31–60	0.09	0.29
Wife, > 60	0.09	0.28
Child, 0–7	0.42	0.76
Child, 8–12	0.30	0.61
Male, 13–17	0.14	0.40
Female, 13–17	0.14	0.39
Male, 18–30	0.12	0.37
Female, 18–30	0.16	0.40
Male, 31–60	0.01	0.11
Female, 31–60	0.22	0.44
Male, > 60	0.05	0.23
Female, > 60	0.12	0.33

Table 2. *(concluded)*

	Mean	Standard Deviation
Fixtures		
Running water	0.83	0.37
Water closet	0.72	0.45
Bathroom	0.70	0.46
Gas	0.63	0.48
Central heating	0.56	0.50
Number of observations (households)		
Total	203,620	
1985	21,560	
1986	25,475	
1987	29,510	
1988	29,287	
1989	29,366	
1990	29,148	
1991	28,632	
1992	10,642	

1 percent increase in the price of food, holding real expenditure (on nondurables) fixed, increases the food share by close to two-tenths of one percentage point. This is quite comparable to other estimates in the consumption literature. For instance, Deaton and Muellbauer (1980) obtained a value of 0.186 for the own food price coefficient using annual British data for 1954–74.[10] This agrees with our estimate to the second decimal place. Their estimate of the real nondurable consumption coefficient was –0.160, which is smaller than ours but still in the ballpark. As a sign of the quality of the HBS data, it is again comforting that we obtain estimates that look reasonably similar to ones in the established consumption literature.

In Figure 1, we examine how well our food share equation is able to mimic the actual changes in the average food share for Polish households over the 1985–92 period. The performance of the equation is strikingly good. We then break down the equation to examine the food share changes predicted by each of its four components (changes in real expenditure, relative prices, demographics, and seasonals), holding the other components fixed at their respective sample means. The average food share over the whole sample period is 0.58.[11] Now consider the effect of varying only real expenditure, holding other factors fixed. The model predicts an increase in the food share of 11 percentage points, from 0.55 to 0.66, between 1989Q4 and 1990Q1. This is the immediate impact of the drop in real incomes following the big bang.

[10]Deaton and Muellbauer (1980) examined allocation of expenditure across eight nondurable commodity categories. Thus, like us, they treat durables as weakly separable.

[11]The food share is so high largely because expenditures on housing are very small. During our sample period, the government provided heavily subsidized housing, and housing was rationed. Because there was no properly functioning market for housing (either rental or owner occupied), we cannot impute the true level of housing consumption.

Table 3. Food Share Equation
(Dependent variable: expenditure on food as a ratio to total expenditure on nondurables)

	OLS		2SLS	
log *csmn.*	–0.195*	(0.001)	–0.263*	(0.002)
log P_2 – log P_1	0.024*	(0.002)	0.029*	(0.003)
log P_3 – log P_1	0.005	(0.003)	–0.054*	(0.004)
log P_4 – log P_1	–0.117*	(0.001)	–0.165*	(0.002)
urban	–0.038*	(0.000)	–0.033*	(0.001)
hdmale, 18–30	–0.003*	(0.001)	–0.003*	(0.001)
hdmale, >60	0.002*	(0.001)	–0.010*	(0.001)
hdfem, 31–60	–0.020*	(0.001)	–0.024*	(0.001)
hdfem, 18–30	–0.029*	(0.003)	–0.027*	(0.003)
hdfem, >60	–0.041*	(0.001)	–0.075*	(0.002)
couple, 18–30	0.045*	(0.002)	0.072*	(0.002)
couple, 31–60	0.060*	(0.001)	0.089*	(0.001)
couple, >60	0.064*	(0.001)	0.081*	(0.002)
child, 0–7	0.021*	(0.000)	0.026*	(0.000)
child, 8–12	0.029*	(0.000)	0.036*	(0.001)
male, 13–17	0.034*	(0.001)	0.042*	(0.001)
fem, 13–17	0.025*	(0.001)	0.033*	(0.001)
male, 18–30	0.036*	(0.001)	0.049*	(0.001)
fem, 18–30	0.033*	(0.001)	0.048*	(0.001)
male, 31–60	0.046*	(0.002)	0.056*	(0.003)
fem, 31–60	0.054*	(0.001)	0.076*	(0.001)
male, >60	0.042*	(0.001)	0.054*	(0.001)
fem, >60	0.053*	(0.001)	0.068*	(0.001)
qrtrdum2	–0.001	(0.001)	0.000	(0.001)
qrtrdum3	0.043*	(0.001)	0.047*	(0.001)
qrtrdum4	0.014*	(0.001)	0.022*	(0.001)
constant	2.680*	(0.007)	3.538*	(0.018)

Note: Standard errors are reported in parentheses. An asterisk indicates statistical significance at the 5 percent level.

However, immediately following the big bang and proceeding through 1992, there was a substantial drop in the relative price of food. Figure 2 presents the price indices used in our analysis. Notice that the relative price of food rose substantially during 1989. Thus, holding other factors fixed, our model predicts that changes in relative prices would have sent the food share from 0.55 in 1989Q1 up to 0.70 in 1989Q4, and that it would have then plummeted to 0.62 in 1990Q1 and further to 0.49 in 1992Q4. In fact, by 1992 the relative price effect clearly dominates the real expenditure effect, and the food share is predicted to have dropped into the 50 percent range (as it in fact did).

The two other factors in the model are seasonals and demographics. The quarterly dummies are quite significant and generate a predicted seasonal pattern of 0.56, 0.56, 0.61, and 0.58. But changes in household demographics over the sample period had little effect on food shares.

Figure 1. Thailand: Interest Rate and Exchange Rate

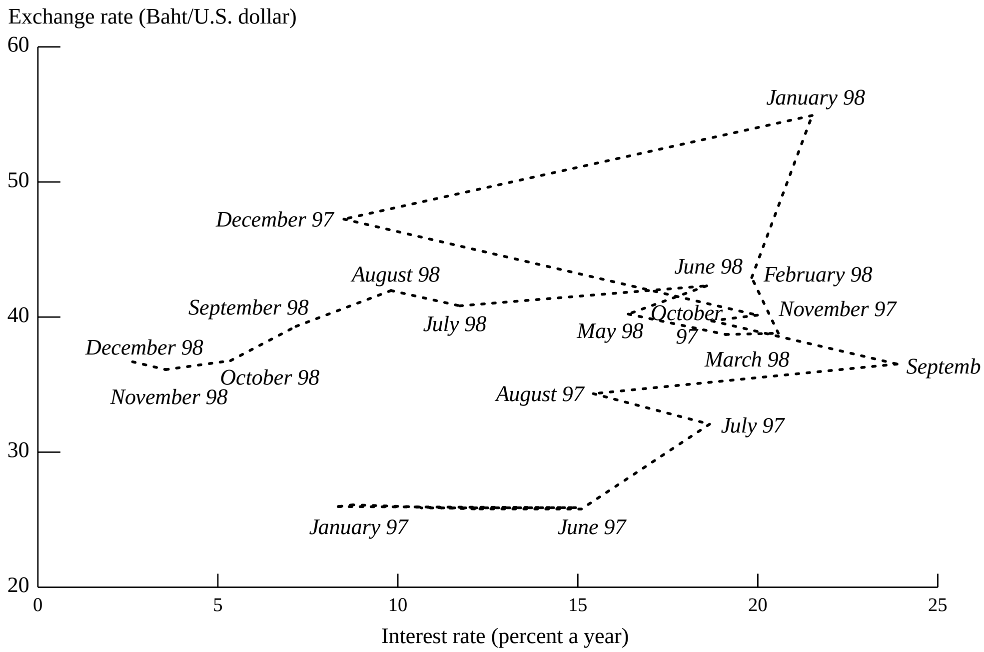
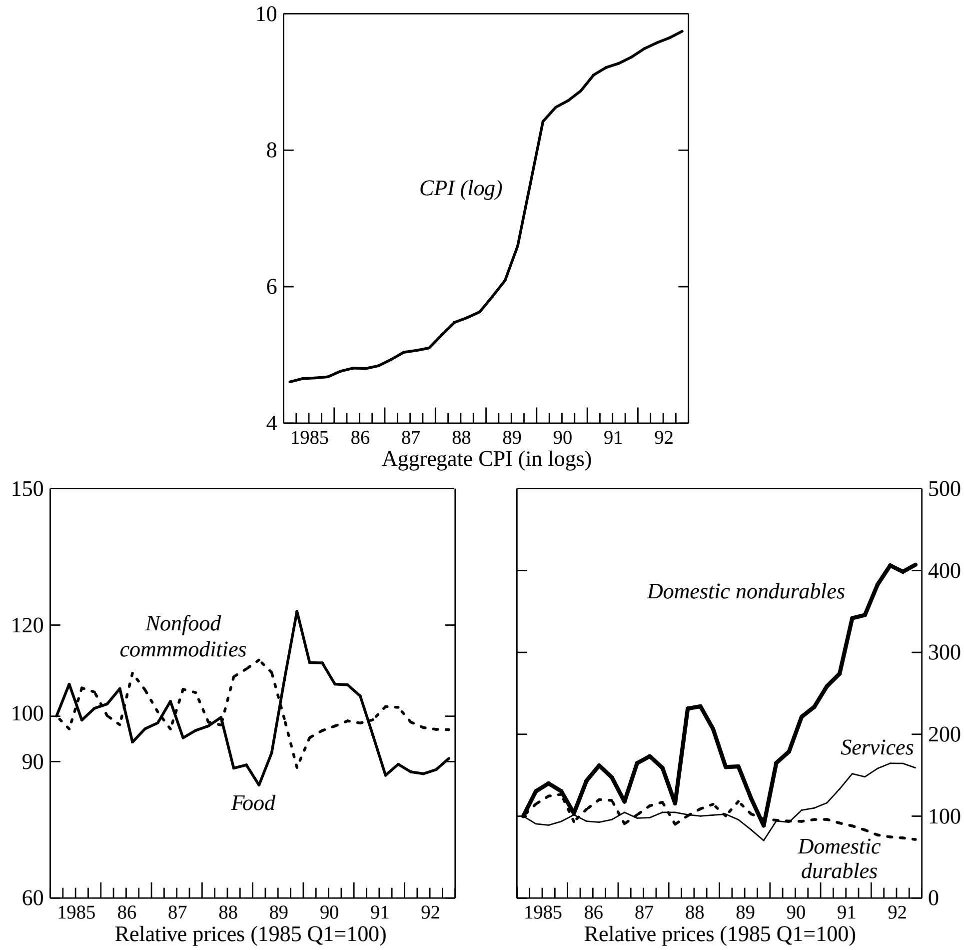


Figure 2. Aggregate CPI and Relative Prices, 1985–92



Note: Lower panels show price indexes relative to aggregate CPI.

Table 4 reports, for representative household types, the values of the household equivalence scales we obtain using the Engel method. For comparison, we also report equivalence scales used by the CSO, the OECD scale, and the scale constructed by McClements (1977), which is widely used in the United Kingdom. Note that our equivalence scales imply somewhat greater household economies of scale than do these other scales.

We also ran the second stage food share regression separately by year and constructed equivalence scales separately for each year of the sample. We do not report the results here but note that, for each type of household, the values of the scales changed little over time. This suggests that the changes in relative prices over the sample period had little effect on the relative cost of maintaining different types of households.

Table 4. Equivalence Scales as a Function of Household Composition

Household Type	GUS	OECD	McClements	Food-Share Equations OLS	Food-Share Equations IV
Single person households					
1 HD = Male, 31–60	0.54	0.59	0.55	0.74	0.71
2 HD = Male, 18–30	0.54	0.59	0.55	0.72	0.70
3 HD = Male, >60	0.54	0.59	0.55	0.74	0.68
4 HD = Female, 31–60	0.46	0.59	0.55	0.66	0.65
5 HD = Female, 18–30	0.46	0.59	0.55	0.63	0.64
6 HD = Female, >60	0.46	0.59	0.55	0.60	0.53
Married couples					
7 HD = Male, 31–60; Female, 31–60	1.00	1.00	1.00	1.00	1.00
8 HD = Male, 18–30; Female 18–30	1.00	1.00	1.00	0.91	0.92
9 HD = Male, >60; Female >60	1.00	1.00	1.00	1.03	0.92
Married couples with one child					
HD = Male, 31–60; Female, 31–60					
10 Male/Female, <7	1.23	1.29	1.17	1.12	1.10
11 Male/Female, 8–12	1.32	1.29	1.24	1.16	1.14
12 Male, 13–17	1.46	1.29	1.29	1.19	1.17
13 Female, 13–17	1.41	1.29	1.29	1.14	1.13
Married couples with older dependents					
HD = Male, 31–60; Female, 31–60					
14 Male, >60	1.54	1.41	1.40	1.24	1.23
15 Female, >60	1.46	1.41	1.40	1.32	1.29
16 Male, >60; Female, >60	2.00	2.00	1.80	1.63	1.59

Notes: HD indicates the head of household.

IV. Inequality

This section contains our main results on the evolution of inequality in Poland over the period 1985–92. It is worth emphasizing that our measures of inequality focus on the cross-sectional distributions of income and consumption and that, unless noted otherwise, our unit of analysis is the individual. After adjusting household income (or consumption) by an equivalence scale, we assign the same level of income (or consumption) to each individual in the household.

In Table 5 we report on the behavior of several alternative inequality measures over the 1985–92 period. The top panel reports Gini coefficients for household income based on four alternative equivalence scales. These are the food share–based, OECD, and McClements scales reported in Table 4, along with the simple per capita scale obtained by dividing household income by household size. Note that the three scales that allow for economies of scale all produce very similar Ginis, typically differing only in the third decimal place. The Ginis based on all four scales indicate that inequality grew from 1985 to 1988 and that inequality actually fell from 1989 through 1992. The Gini based on the food share

Table 5. Poland: Measures of Inequality, 1985–92

	1985	1986	1987	1988	1989	1990	1991	1992
Total income				**Gini Coefficients**				
Food share–based equivalence scale	0.252	0.254	0.246	0.256	0.263	0.250	0.235	0.230
McClements equivalence scale	0.249	0.253	0.246	0.254	0.261	0.249	0.238	0.234
OECD equivalence scale	0.253	0.257	0.250	0.256	0.264	0.253	0.242	0.238
Per capita	0.270	0.274	0.270	0.272	0.278	0.271	0.266	0.264
Urban	0.201	0.203	0.198	0.202	0.223	0.217	0.213	0.210
Rural	0.317	0.307	0.287	0.302	0.296	0.278	0.249	0.249
Income excluding transfers	0.373	0.375	0.368	0.385	0.384	0.389	0.404	0.416
Nondurables consumption								
Food share–based equivalence scale	0.196	0.200	0.205	0.211	0.219	0.209	0.208	0.205
McClements equivalence scale	0.197	0.202	0.208	0.214	0.220	0.210	0.213	0.212
OECD equivalence scale	0.200	0.207	0.212	0.217	0.224	0.214	0.218	0.217
Per capita	0.222	0.229	0.236	0.239	0.242	0.235	0.245	0.249
Total consumption	0.230	0.234	0.239	0.244	0.258	0.241	0.233	0.227
				Half the Square of Coefficient of Variation (variables adjusted by food share based equivalence scales)				
Total income	0.085	0.090	0.085	0.091	0.105	0.086	0.079	0.077
Nondurables consumption	0.066	0.068	0.070	0.074	0.081	0.068	0.072	0.068
Income excluding transfers	0.184	0.190	0.186	0.203	0.210	0.207	0.230	0.244
				Mean Log Deviation (variables adjusted by food share based equivalence scales)				
Total income	0.075	0.079	0.077	0.078	0.087	0.075	0.071	0.069
Nondurables consumption	0.060	0.062	0.064	0.067	0.074	0.062	0.064	0.064
Income excluding transfers	0.224	0.214	0.213	0.221	0.244	0.247	0.268	0.278

Notes: The inequality measures shown here are for the individual distributions of income and consumption. Household income and consumption are adjusted using the food share-based equivalence scale (unless indicated otherwise) and allocated equally to individuals in the household.

scale implies a somewhat sharper decline in inequality in 1989–92 (from 0.260 to 0.230) than do the Ginis based on the other three scales.

The Ginis based on simple per capita household income are consistently about 0.015 to 0.030 greater than those based on per equivalent income. Nevertheless, they show the same pattern of inequality growing from 1985 to 1988 and declining from 1989 to 1992. We noted earlier that OECD (1997) reports that the Gini based on per capita income grew from 0.25 in 1989 to 0.27 in 1992. In contrast, we obtain a decline from 0.278 to 0.264 when we use the per capita scale. Thus, the choice of equivalence scale is clearly not the cause of this difference in results. What does account for the difference between our Ginis and those reported by the OECD, or, for that matter, by Milanovic (1993, 1998) for this same period?

One potential source of difference is that prior studies approximated Ginis based on grouped income data. Consider the year 1987. In the *Survey* for that year, the CSO published data on the number of people in each of eight per capita household income intervals. Based on those data, Milanovic (1998) calculates an approximate Gini of 0.252. Using the same data, we obtain a similar Gini value of 0.248.[12] This is comparable to the value of 0.270 that we calculate from the HBS microdata. Thus, prima facie, it appears that use of grouped data does lead to downward bias in the Gini. However, if we take the HBS microdata for 1987, group households into the same eight per capita income intervals, and approximate the Gini based on that information, we obtain a value of 0.265. Hence, it appears that use of grouped data does bias down the Gini—but not by nearly enough to account for the substantially lower 1987 Gini value reported in earlier studies. The same pattern holds for other years.

Another potential source of difference between our Gini estimates and those in earlier studies is that prior studies used different definitions of income. As we noted earlier, the CSO includes gross rather than net farm income in its household income measure. If we do the same, then for 1987 we obtain a Gini of about 0.260. Thus, it appears that this difference in income definitions cannot account for much of the difference in Gini values.

Consider next the years 1989 and 1990. For those years, OECD (1997, Figure 22, p. 86) reports Gini values based on household per capita income of 0.25 and 0.23. Similarly, Milanovic (1993) reports Gini values of 0.260 and 0.255. All these figures are based on various aggregate income decile data provided by the CSO, and we are not certain of the sources of the (minor) discrepancies. Our Ginis based on per capita household income for those two years are much higher, at 0.278 and 0.271, respectively. If we group our data into deciles and then approximate the Ginis, we get slightly lower Ginis. And if we leave in gross farm income instead of net farm income, we get marginally higher Ginis. Thus, neither grouping nor the difference in income definition accounts for our much higher Gini values in 1989–90.

Strangely, in 1991–92 the discrepancies between our results and those in prior studies largely disappear. In those years our Gini values drop substantially, while those reported by the OECD rise substantially, and all the values fall in the 0.26–0.27 range.

At this point, we have been unable to determine why we obtain higher Gini values for years before 1991 than do prior studies based on aggregate income data from the CSO. But, at least mechanically, this difference explains why we find that inequality fell after the big bang while prior studies found that it increased: Essentially, our calculations suggest that income inequality was far higher before the big bang than the aggregate statistics from the CSO would indicate.

[12]We are unsure of the reason for the slight difference between our calculation and that of Milanovic (we tried to replicate his approximation method). The grouped data provided in the *Surveys* contain the number of households within certain ranges of household per capita income, the average of per capita income among households in each interval, and the average size of households within each interval. We approximate Ginis based on these data by assuming that all households in an interval are at the mean of per capita income for that interval.

We would argue that the inequality measures that we have calculated directly from the HBS microdata are more reliable than those calculated from the aggregate CSO statistics. Hence, we now leave off the comparison of our statistics with those from previous studies and go on to analyze our statistics in more detail. Because the choice of equivalence scale appears to make little difference to our results, we will henceforth report results using the scale based on food shares unless otherwise noted.

Consider now the rows of Table 5 that report separate Gini coefficients for the urban and rural populations. The Ginis for the rural population are consistently much greater than those for the urban population. Neither group shows any clear pattern of change in inequality during 1985–88. During 1989–92, there is a decline in inequality for both groups, but it is far greater among the rural population.

We next examine the role of transfer payments in reducing inequality. Strikingly, the Gini based on income excluding transfers increased from 0.384 to 0.416 during 1989–92. Thus, we find that actual income did grow more unequal after the big bang.[13] Yet, the transfer system more than compensated for this, as the decline in the Gini for total income from 0.263 to 0.230 during 1989–92 indicates. Nevertheless, it is possible that the growth in inequality of earned income has led at least in part to the general perception that inequality has risen. These results contradict the received wisdom that transfers in Poland have been regressive and have thus contributed to the increase in inequality (see, e.g., Milanovic, 1998, p. 49). We will explore this in more detail below.

Now we turn to examination of changes in consumption inequality. Again, the Ginis based on the three equivalence scales that allow for household economies of scale all show a similar pattern. Inequality grows from 1985 to 1989 and then declines from 1989 to 1992. The decline from 0.219 to 0.205 indicated by the food share based scale is again sharper than for the other scales. Similarly, the Gini for nondurable consumption declines from 0.258 in 1989 to 0.227 in 1992. It is also worth noting that, as expected, Ginis for nondurable consumption are well below those for income.

The Gini coefficient is sensitive to changes in a distribution near the median (see Atkinson, 1970). The coefficient of variation is more sensitive to changes at the high end of a distribution, while the mean logarithmic deviation is more sensitive to changes near the low end. We report these other inequality measures in the bottom panels of Table 5, in order to determine if they tell a consistent story. These measures echo the results based on the Gini coefficients. Both measures also reveal a sharp increase in inequality based on income net of transfers.

Kernel Density Estimates for Income and Consumption

To obtain a visual representation of changes in the shape and features of the entire distribution, we now examine kernel density estimates of the income and consumption distributions. Figure 3 (top panel) contains kernel density estimates

[13]One cannot conclude from this that earnings potential grew more unequal. For instance, the labor earnings we observe are accepted rather than offered earnings. The accepted earnings distribution can grow more unequal even if the offered earnings distribution does not, simply because the nature of the selection into the pool of those who accept wage offers can change over time.

for real household income for the years 1988, 1989, 1990, and 1992.[14] An Epanechnikov kernel with a bandwidth of 4,000 is used. The density is calculated at the same 200 points for all four years, and the first 125 are plotted in the figure. This covers at least 96 percent of the households in all four years. Figure 3 (lower panel) contains kernel density estimates for real household nondurable consumption for the same four years. Reflecting the more compact distribution of consumption, the first 75 points cover more than 99 percent of the households.

The change in the shape of the densities between the years 1988–89 and the years 1990–92 is striking. Much of the change simply reflects the decline in mean income and consumption following the big bang. However, the change in shape observed in Figure 3 is not due simply to a contraction of the mean. To see this, consider taking the distribution for 1992 and multiplying all the income figures by the ratio of mean income in 1988 to that in 1992. Such a transformation will preserve relative inequality measures, while equating mean income in 1992 with that in 1988. The 1988 income density and the transformed density for 1992 are plotted together in Figure 4.

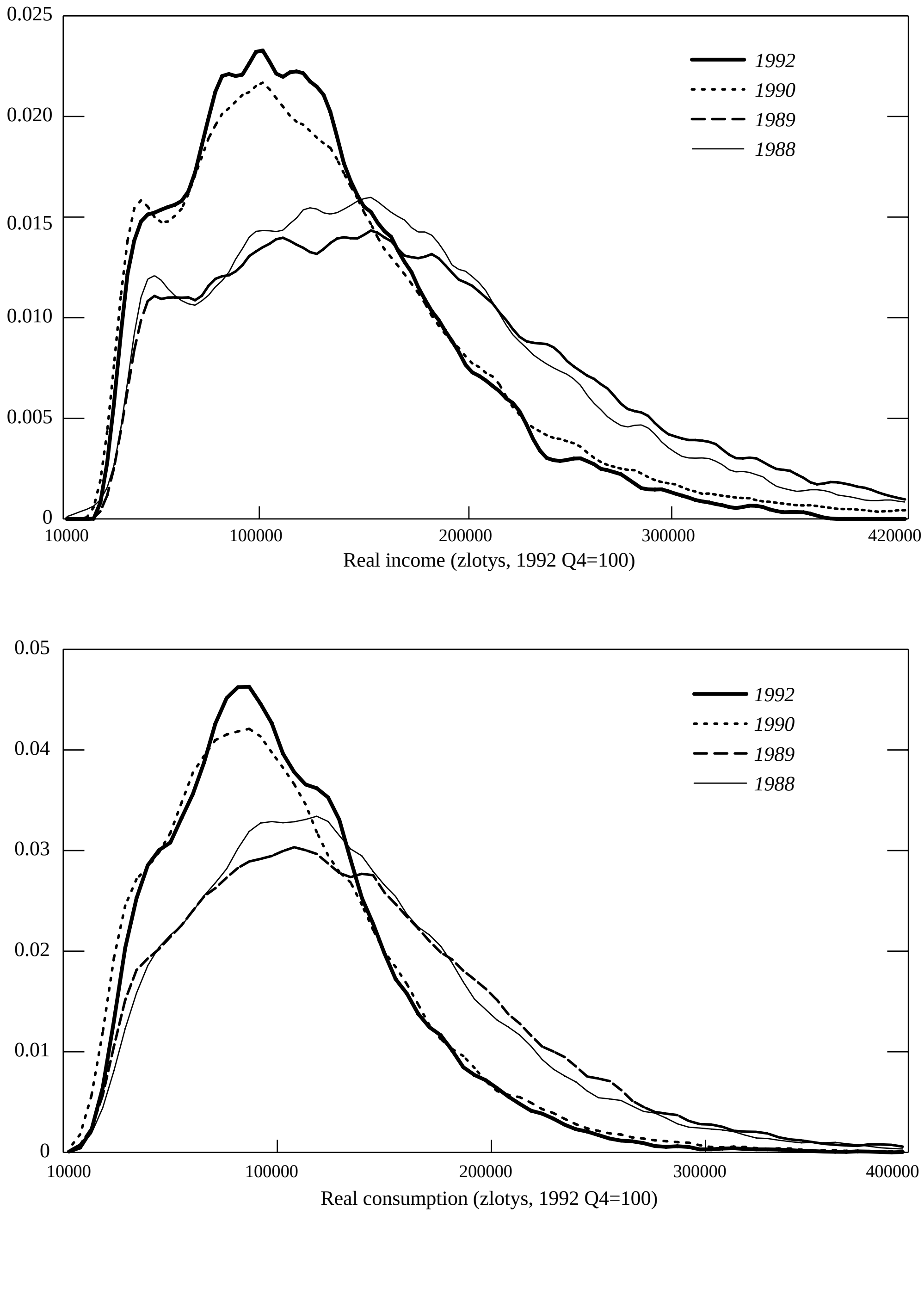
The most prominent features of Figure 4 are that, in moving from 1988 to 1992, the mass in the left tail is reduced, and the distribution becomes more peaked around the mode. This accounts for the declines in the various inequality measures noted above. A key aspect of what happened becomes apparent if one compares Figure 3 (top panel) with Figure 4. As the overall income distribution shifted left, there was a support area at about 34,000 to 58,000 zlotys (in 1992 fourth-quarter zlotys) below which household income tended not to fall. Because of the drop in mean real income from 1988 to 1992, the ratio of this support level to mean income increased. In Figure 4, this has the effect of shifting to the right the fat part of the left tail of the scale-adjusted income distribution.

We investigated the income sources of households with real income in the 34,000 to 58,000 zloty range, and found that these households receive more than 80 percent of their income from pensions (80.5 percent in 1988, 82.2 percent in 1992). The percentages drop off quickly as household income rises above the 58,000 zloty level. The proportion of total household income for all households coming from pensions was 16.8 percent in 1988 and 26.8 percent in 1992. Thus, the households with income in the support area of about 34,000 to 58,000 zlotys received a far higher share of income from pensions than the typical household. Furthermore, it is important to note that, while mean real household income fell from 178,969 zlotys in 1988 to 131,563 zlotys in 1992, the mean real pension actually rose from 29,811 to 35,258 zlotys. This resulted from legislation that took effect in 1991 that made pensions substantially more generous. Hence, our results suggest that the new pension law helped shift the fat part of the left tail of the income distribution to the right, and that this contributed significantly to the reductions in inequality measures noted above.[15]

[14]No adjustment is made for household size.

[15]It is also worth noting that the fraction of households headed by pensioners increased from approximately 15 percent in the 1985–90 period to roughly 25 percent in 1992. Opting for the more generous pensions was apparently an attractive option for workers who did not fare well in the transition.

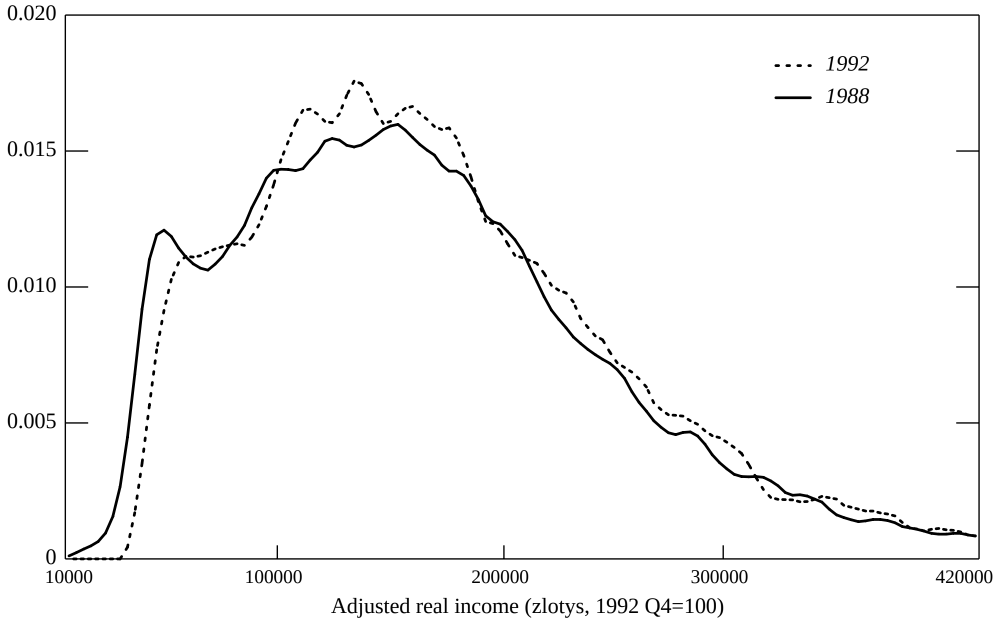
Figure 3. Kernel Density Estimates



Quantile Ratios and Shares

Another common way to summarize changes in inequality is to examine quantile ratios. Unlike the scalar inequality measures considered above, examination of a

Figure 4. Kernel Density Estimates



set of quantile ratios allows one to consider changes in inequality at various different points in the distribution.

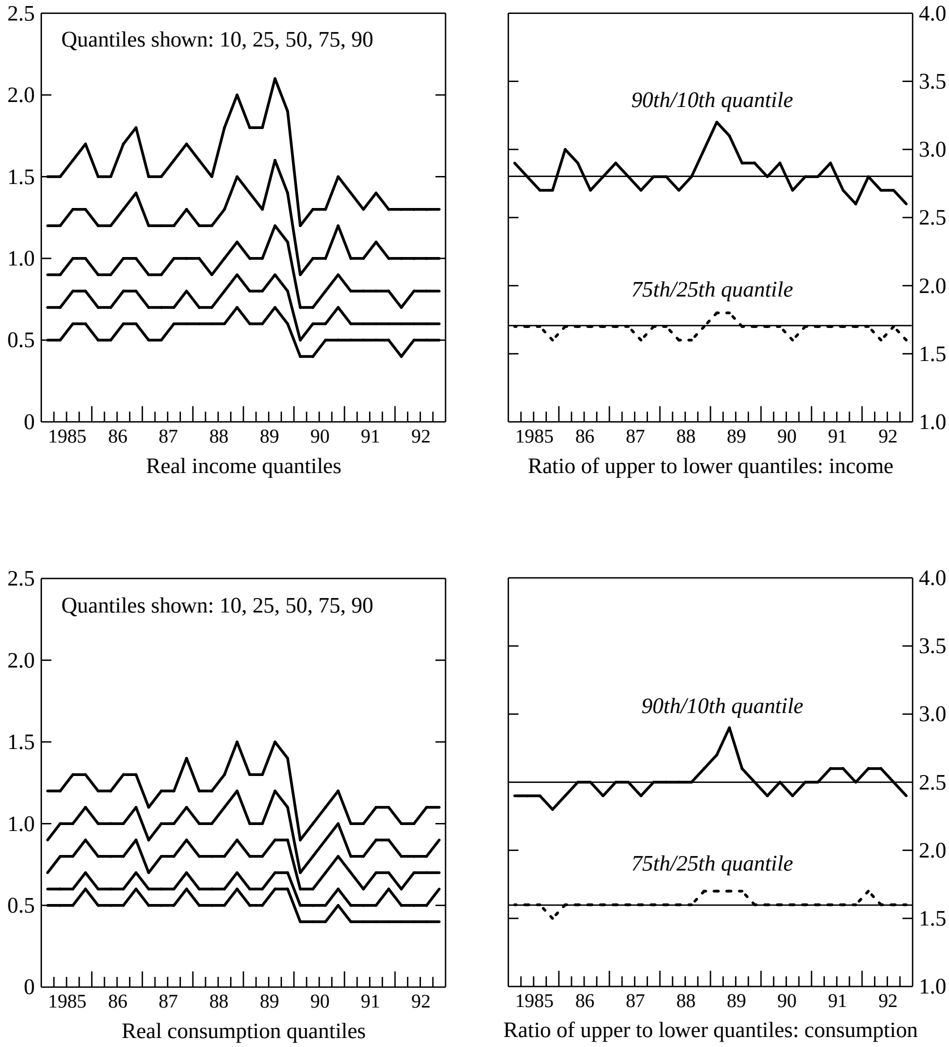
Figure 5 reports values of the 0.10, 0.25, 0.50, 0.75, and 0.90 income and consumption quantiles for each quarter over the sample period, as well as the 90/10 and 75/25 quantile ratios. The values are for real household income and nondurable consumption, adjusted using the food share–based equivalence scale. There are upward blips in both quantile ratios in late 1988 and early 1989, but there is little evidence of any trends over the sample period as a whole. If anything, the 90/10 ratio for income appears to drift slightly downward after 1989.

In Table 6 we report the shares of income and consumption going to each quintile of the respective distributions. Note that the share of total income going to the bottom quintile rose slightly over the 1989–92 period, while the share going to the top quintile declined. But for income net of transfers the pattern is reversed, again indicating that transfers served to reduce inequality after the big bang. For consumption the share of the bottom quintile also rose over 1989–92, while that of the top quintile fell.

Income and Consumption Patterns Categorized by Source of Income, Education, and Age

We have found no evidence of an increase in inequality in Poland in the first three years following the big bang, regardless of which of several inequality measures we consider. However, this does not mean that there were not winners and losers

Figure 5. Real Income and Consumption, 1985–92
(millions of zloty, 1992 Q4 prices)



Note: Household income and consumption are adjusted by equivalence scales.

in the transition. In this section, we turn to an examination of how different groups fared in terms of income and consumption.

In Figure 6 we report how median income and consumption moved for four types of households differentiated by main income source of the household head: workers, farmers, mixed farmers/workers, and pensioners. A notable feature of the results is that the use of equivalence scales is important. The per capita household income and consumption plots in the top panel suggest that pensioner-headed households moved from a middle position to being clearly better off than other households after the big bang. According to Milanovic (1998, p. 49), who looks at

Table 6. Quintile Shares of Income and Consumption

	1985	1986	1987	1988	1989	1990	1991	1992
				Total income				
Quantile range								
≤ 20	9.1	8.4	8.5	8.6	8.8	9.2	9.9	10.3
21–40	15.1	14.9	15.0	14.7	14.3	14.8	15.0	15.1
41–60	18.5	18.3	18.4	18.0	17.8	18.2	18.3	18.3
61–80	22.6	22.6	22.5	22.1	22.2	22.5	22.5	22.6
> 80	34.7	35.8	35.7	36.6	37.0	35.4	34.3	33.7
				Income net of transfers				
≤ 20	2.5	2.0	2.0	2.0	2.5	1.9	1.6	1.2
21–40	13.3	13.2	13.3	12.7	12.4	12.5	11.8	10.7
41–60	18.9	18.7	18.7	18.1	17.9	18.4	18.3	18.2
61–80	24.7	24.5	24.5	24.1	24.2	24.9	25.4	26.0
> 80	40.6	41.6	41.5	43.2	43.1	42.4	42.9	43.9
				Total consumption				
≤ 20	11.0	10.8	10.6	10.4	9.9	10.6	10.8	10.8
21–40	14.8	14.6	14.4	14.4	14.0	14.5	14.7	14.8
41–60	18.0	17.9	17.7	17.7	17.5	17.8	18.0	18.1
61–80	22.0	22.1	21.9	22.1	22.1	22.1	22.2	22.3
> 80	34.2	34.6	35.4	35.5	36.4	35.1	34.4	34.0
				Nondurables consumption				
≤ 20	11.9	11.7	11.5	11.2	10.9	11.4	11.5	11.4
21–40	15.6	15.5	15.4	15.2	15.1	15.4	15.4	15.5
41–60	18.7	18.7	18.6	18.5	18.5	18.5	18.5	18.6
61–80	22.4	22.4	22.4	22.5	22.7	22.5	22.4	22.5
> 80	31.5	31.7	32.1	32.6	32.8	32.2	32.2	32.0

Note: Each column indicates the share of aggregate income or consumption accounted for by persons in different quintile ranges for that variable.

per capita income, "pensions thus contributed strongly to increase inequality." But the per equivalent unit results in the middle panel tell a very different story.[16] They indicate that pensioner-headed households had very low income and consumption relative to other groups during the 1985–89 period, and that their relative position improved dramatically after the big bang so as to bring their income and consumption up to almost the same level as the next lowest group (farmers). As a result, we find that pensions contributed importantly to a reduction in inequality.[17] The main impetus behind the improved relative position of pensioners was a substantial increase in pension levels that took place in 1991.

[16]The reason for the difference in the scales is that the mean numbers of persons in worker, farmer, farmer/worker, and pensioners households are 3.59, 3.64, 4.55, and 1.88, respectively, while the mean numbers of equivalent units are 1.69, 1.77, 2.08, and 1.19, respectively.

[17]Milanovic (1998, p. 54) concludes that transfers in Poland were regressive overall and contributed to increased inequality. This also contradicts our findings above about the impact of transfers on the Gini coefficient.

Figure 6. Median Income and Consumption for Different Groups

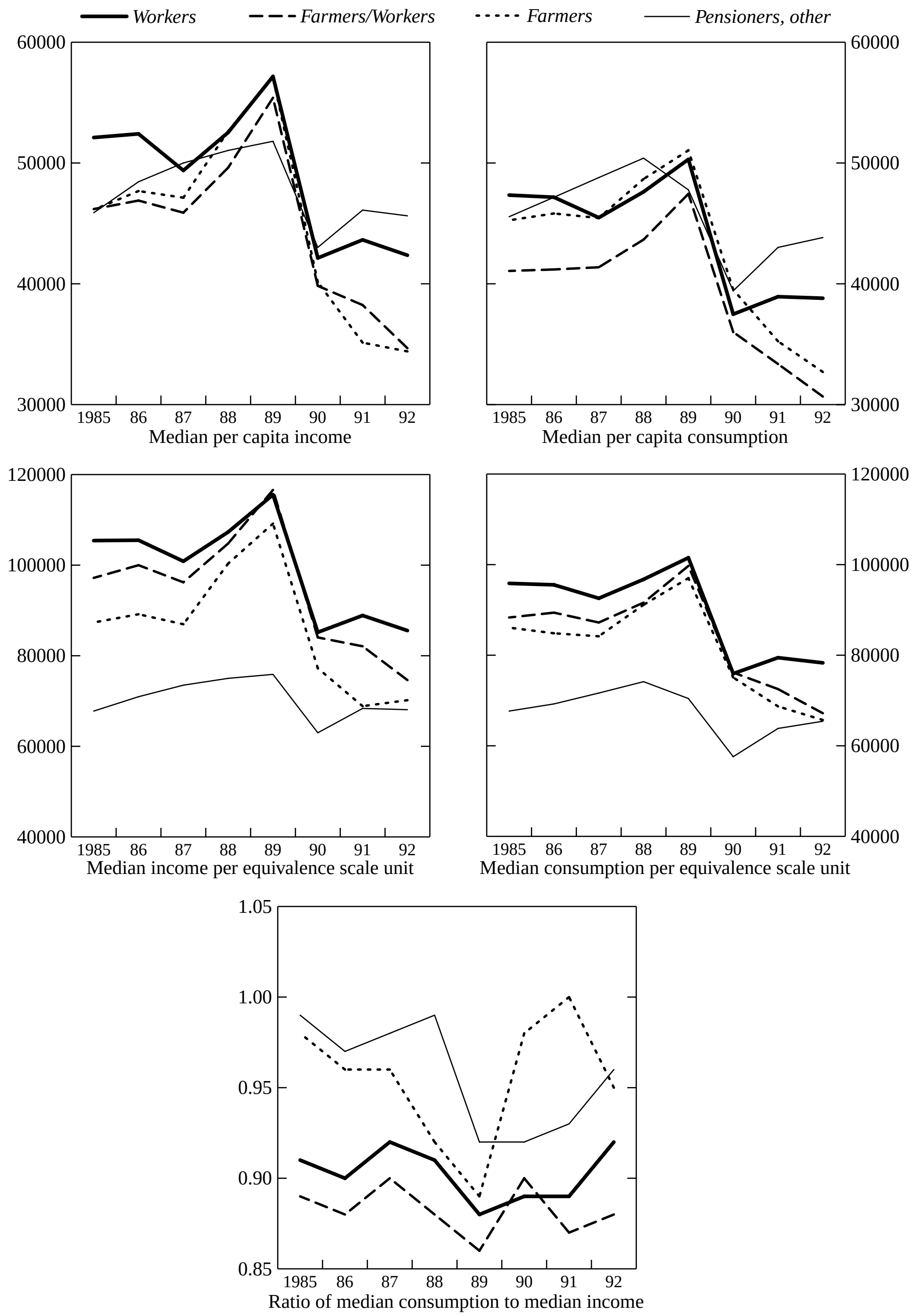


Table 7 reports the fractions of households that fall in each quantile of the income distribution, conditional on education or age of the household head. For example, in 1989, 45.8 percent of households in which the head had a college degree were in the top quantile. By 1992 the fraction rose to 58 percent. In contrast, in 1989, among households in which the head had only a primary school education, 14.9 percent were in the top quantile, but by 1992 the number had fallen to 9.5 percent.

Another striking feature is the improvement of conditions for the old, which resulted from more generous pensions. Among households in which the head was older than 60, 39.2 percent were in the bottom quantile in 1989, but the number dropped to 24.3 percent by 1992. In contrast, the probabilities that a household with a young (18–30) or middle aged (31–60) head would fall in the bottom quantile of the income distribution grew over the same period.

Quantile Regressions

We now examine how changes in the overall well-being of households were influenced by the education level of the household head, using quantile regression techniques.[18] This enables us to characterize in a parsimonious way the changes in the entire conditional distribution of income, as opposed to looking only at changes in the conditional mean. We ran quantile regressions of log real quarterly household per equivalent income on demographic characteristics of household heads. These characteristics were dummies for the six education categories (with primary school being the omitted category), labor market experience (i.e., age minus years of education minus 6), experience squared, location (urban/rural), and gender.

Table 8 reports conditional quantiles of log real household income based on the education level of the household head. Note that log income for all education groups drops substantially at all quantile points from 1989 to 1990. The drops tend to be larger at the higher quantile points (e.g., at the 0.90 quantile they range from 30 to 33 percent, while at the 0.10 quantile they range from 25 to 28 percent). Thus, we see a decline in inequality within education groups as measured by quantile ratios from 1989 to 1990. The interesting thing the table reveals is that for the vocational, primary, and some primary groups, income rises slightly in 1991 but then declines (below the 1990 levels) in 1992, with the drops much more pronounced at the higher quantiles. In contrast, for households headed by college graduates, the 0.10, 0.25, and 0.50 quantiles recover a bit, while the 0.75 and 0.90 quantiles hold steady. Hence, in 1991 and 1992, we see a further drop in inequality within each education group and an improvement in the relative position of the households headed by a college graduate. High school graduates hold steady at all quantiles in 1991–92, except that the 0.90 quantile falls. Thus, the only group of households that experiences a slight recovery in earnings from 1990 to 1992 is the group with a college-educated head.

[18] For an analysis of changes in the labor income of individual workers during the Polish transition, see Keane and Prasad (2001).

Table 7. Fractions of Various Groups in Different Income Quantile Ranges
(Based on education and age of household head)

	1985	1986	1987	1988	1989	1990	1991	1992
Quantile range				**College degree or some college**				
≤ 20	3.4	3.3	3.4	3.5	3.9	3.9	2.6	2.2
21–40	10.4	8.4	7.5	9.2	9.0	7.5	5.4	6.0
41–60	16.1	16.1	16.0	17.8	14.8	14.4	11.2	11.5
61–80	26.7	27.7	27.3	26.9	26.5	25.5	21.0	22.3
> 80	43.4	44.5	45.8	42.5	45.8	48.8	59.7	58.0
Fraction of annual sample	7.8	7.0	6.3	6.3	6.0	6.2	6.8	8.4
				High school				
≤ 20	9.7	9.9	9.2	10.6	11.0	11.0	9.5	9.9
21–40	17.8	16.7	15.4	16.8	16.8	16.2	12.9	15.1
41–60	22.6	22.2	22.3	21.7	21.4	20.9	20.7	19.2
61–80	25.4	26.9	27.0	26.3	25.2	25.0	25.9	27.6
> 80	24.5	24.3	26.0	24.6	25.6	26.9	31.0	28.2
Fraction of annual sample	21.1	18.7	17.7	18.2	17.6	19.0	20.0	22.8
				Some high school or vocational training				
≤ 20	12.9	13.8	14.1	13.0	14.2	17.5	17.5	19.1
21–40	19.8	19.5	20.0	19.6	19.2	19.9	20.8	21.2
41–60	23.8	23.1	23.2	23.1	23.0	21.5	22.2	22.9
61–80	23.5	22.9	22.4	23.5	23.3	21.7	22.6	20.4
> 80	20.0	20.8	20.3	20.7	20.4	19.5	16.9	16.3
Fraction of annual sample	29.1	30.3	30.9	31.9	33.0	35.2	34.1	34.0
				Primary school				
≤ 20	30.4	28.7	28.1	28.6	27.8	27.6	29.2	30.4
21–40	23.0	23.3	23.1	23.0	23.1	23.3	25.0	25.5
41–60	17.7	18.3	18.4	18.2	18.7	19.3	19.5	20.0
61–80	14.5	14.8	15.6	14.7	15.5	16.2	15.5	14.7
> 80	14.4	14.9	14.8	15.5	14.9	13.5	10.9	9.5
Fraction of annual sample	35.3	36.5	37.5	36.6	36.8	33.9	33.5	30.3

Within- and Between-Group Decompositions of Inequality

In this subsection, we address the question of the extent to which inequality is within versus between groups, and the extent to which each type of inequality changed over the transition. The single parameter generalized entropy measures of inequality can be additively decomposed into within- and between-group components (see Shorrocks, 1984). This family includes the mean log deviation and half the square of the coefficient of variation, but not the Gini coefficient. Hence, in Table 9, we report decompositions of the former two inequality measures for both income and consumption, grouping households by main income source of the household head. Notice that the vast majority of inequality is within group, rather

Table 7. *(concluded)*

	1985	1986	1987	1988	1989	1990	1991	1992
Quantile range				**Less than primary school**				
≤ 20	47.8	43.6	43.3	45.8	44.4	37.8	38.1	41.4
21–40	23.0	25.0	25.8	24.1	25.1	27.4	28.6	24.4
41–60	11.9	13.8	12.7	13.0	13.3	18.2	17.8	18.3
61–80	9.2	9.0	9.1	9.2	9.1	9.7	9.0	9.9
> 80	8.0	8.5	9.2	8.0	8.2	6.9	6.5	6.1
Fraction of annual sample	6.7	7.5	7.6	7.1	6.6	5.7	5.7	4.5
				Age 18–30				
≤ 20	10.1	12.6	13.0	11.1	11.7	15.6	14.1	16.0
21–40	17.7	17.3	17.4	17.4	18.0	17.0	17.3	18.9
41–60	22.7	22.8	22.2	22.1	22.2	19.7	21.7	21.7
61–80	25.1	22.6	22.6	24.2	22.9	22.0	23.4	20.4
> 80	24.3	24.7	24.9	25.2	25.2	25.7	23.5	23.0
Fraction of annual sample	12.8	13.0	13.0	11.4	10.4	10.9	10.7	10.2
				Age 31–60				
≤ 20	14.3	14.8	16.1	15.1	14.5	16.8	17.9	19.0
21–40	18.1	17.6	17.9	18.0	17.2	17.4	17.4	18.3
41–60	21.7	21.1	20.7	21.1	21.2	20.1	19.7	19.3
61–80	22.6	23.0	22.4	22.5	23.4	22.3	21.9	21.0
> 80	23.3	23.5	23.0	23.3	23.8	23.5	23.0	22.4
Fraction of annual sample	66.5	65.2	65.1	66.3	66.2	65.2	65.3	64.7
				Age > 60				
≤ 20	44.4	40.1	35.8	39.2	39.2	30.8	28.3	24.3
21–40	27.5	28.9	27.9	27.4	28.8	28.4	28.3	24.7
41–60	13.0	15.0	16.6	15.7	15.8	20.0	20.0	21.1
61–80	8.4	9.5	11.4	10.3	9.2	12.9	13.3	17.2
> 80	6.6	6.6	8.3	7.4	7.0	8.0	10.1	12.6
Fraction of annual sample	20.7	21.7	21.8	22.3	23.5	23.9	24.0	25.1

Note: Each column indicates the share of aggregate income or consumption accounted for by persons in different quantile ranges for that variable.

than between groups, which is not surprising given the coarse nature of the grouping. Both within- and between-group inequality rose in 1989. The total decline in inequality from 1989 to 1992 is attributable in about equal part to declines in both the within- and between-group components.

The lower panel of Table 9 shows the evolution of within-group inequality, as measured by the Gini coefficient, for different socioeconomic groups. The interesting finding is that the changes in within-group inequality are very different for the different groups. For instance, for households headed by workers, the Gini coefficient increased slightly after the big bang. In contrast, for households headed by farmers and farmers/workers, the Gini coefficient declined substantially from 1988 to 1992.

Table 8. Conditional Quantiles of Real Quarterly Household Income (in logs)
(Based on educational attainment of head of household)

	Quantile = 0.10				Quantile = 0.25				Quantile = 0.50			
	COL	HS	VOC	PS	COL	HS	VOC	PS	COL	HS	VOC	PS
85	11.20	11.05	10.96	10.82	11.44	11.28	11.19	11.08	11.66	11.51	11.41	11.33
86	11.23	11.07	10.98	10.85	11.47	11.31	11.20	11.10	11.70	11.54	11.44	11.36
87	11.24	11.07	10.98	10.85	11.47	11.31	11.21	11.09	11.70	11.53	11.42	11.34
88	11.28	11.12	11.06	10.92	11.50	11.34	11.26	11.16	11.73	11.57	11.48	11.40
89	11.36	11.16	11.06	10.95	11.61	11.40	11.29	11.20	11.84	11.64	11.53	11.46
90	11.09	10.89	10.78	10.68	11.32	11.12	10.99	10.93	11.57	11.36	11.23	11.17
91	11.21	10.95	10.82	10.72	11.42	11.20	11.03	10.95	11.68	11.43	11.26	11.17
92	11.13	10.89	10.76	10.66	11.36	11.13	10.98	10.90	11.61	11.37	11.20	11.11

	Quantile = 0.75				Quantile = 0.90			
	COL	HS	VOC	PS	COL	HS	VOC	PS
85	11.90	11.74	11.65	11.60	12.13	11.99	11.91	11.86
86	11.94	11.77	11.69	11.62	12.20	12.03	11.93	11.89
87	11.93	11.76	11.65	11.59	12.19	12.01	11.90	11.85
88	11.99	11.82	11.72	11.66	12.23	12.07	11.97	11.93
89	12.11	11.91	11.78	11.73	12.41	12.18	12.04	12.02
90	11.84	11.61	11.48	11.43	12.08	11.88	11.72	11.69
91	11.92	11.67	11.49	11.42	12.17	11.90	11.71	11.66
92	11.84	11.59	11.43	11.35	12.07	11.81	11.65	11.59

Notes: COL represents college degree; HS represents high school degree; VOC represents basic vocational training; PS represents primary school. The regressors in the quantile regressions included an urban dummy and the following variables based on household head attributes: experience, experience squared, a dummy for sex, and six education dummies. To generate the predicted quantiles, all independent variables except for the education dummies were set to their means over the full sample.

V. Income and Consumption Mobility

The Polish HBS also contains a limited panel aspect. A certain fraction of the households interviewed in each year are interviewed again the next year. We now exploit this panel aspect of the data to examine whether income and consumption mobility (e.g., the probabilities of moving across quintiles of the distribution) have changed. That is, in the transition to a free market economy, to what extent do household income and consumption become more variable over time?

In Table 10 we report quintile transition rates for total household income, adjusted using the food share–based equivalence scales. The transition matrices are presented for five pairs of years (1987–88, 1988–89, 1989–90, 1990–91, and 1991–92). Farmers and farmers/workers are excluded from the analysis because the income figures are reported quarterly and there are strong seasonals in farm income. There is a strong tendency for farm households who are interviewed during the harvest period to have high incomes in both years, while farm households interviewed in other periods tend to have low incomes in both years. As a

Table 9. Decomposition of Inequality Measures

	1985	1986	1987	1988	1989	1990	1991	1992
	Half the square of the coefficient of variation (× 100)							
Total	8.5	9.0	8.5	9.1	10.5	8.6	7.9	7.7
Between-group	1.4	1.2	0.8	1.2	1.8	0.9	0.6	0.6
Within-group	7.1	7.8	7.7	8.0	8.7	7.8	7.3	7.0
	Mean log deviation (× 100)							
Total	7.5	7.9	7.7	7.8	8.7	7.5	7.1	6.9
Between-group	0.8	0.6	0.4	0.6	1.0	0.5	0.3	0.3
Within-group	6.7	7.3	7.3	7.2	7.7	7.0	6.8	6.6
	Gini coefficients							
Workers	0.186	0.192	0.191	0.189	0.208	0.211	0.208	0.211
Farmers	0.475	0.483	0.478	0.496	0.440	0.420	0.366	0.321
Mixed, farmers/workers	0.272	0.279	0.276	0.285	0.271	0.253	0.229	0.220
Pensioners, other	0.211	0.212	0.203	0.205	0.214	0.206	0.210	0.203

Note: Socioeconomic groups are defined on the basis of the household's primary source of income.

result, inclusion of farm households will lead to an exaggeration of the degree of persistence in income.

The own transition rates for the bottom three quintiles decline noticeably from 1987–88 to 1991–92. For instance, the sample probability of staying in the bottom quintile drops from 0.64 to 0.57, while the probability of moving from the bottom up to the second quintile increases from 0.21 to 0.25. For the fourth and fifth quintiles the transition rates are little changed. Overall, the results seem to show some increase in mobility, but it does not appear to be dramatic. Because Table 10 is rather hard to digest, we have attempted to summarize it using the simple regression in Table 11, in which the elements of the transition matrices are regressed on a set of dummy variables and time effects. The most interesting parameters are the coefficients on the interactions between calendar time and the dummies for whether the matrix element is one of the diagonals. These coefficients for income mobility (left column) show that the diagonals are trending significantly downward. The largest and only significant coefficient for an off-diagonal with time interaction is that for two off the diagonal. This coefficient indicates that the probability of a transition across two quintiles (either up or down) has trended upward.

Table 10 also reports transition rates for consumption. These transition matrices appear to be remarkably stable over time. This visual impression is confirmed in Table 11, which also reports regression results for the consumption transition matrix elements. Notice that there are no significant trends in the diagonal elements. Thus, we find that while income mobility has increased during the transition, consumption mobility has not.

A striking feature of the results is that the transition matrices for income and consumption look almost identical in 1992. Note, however, that this does not mean

Table 10. Quintile Transition Rates

		Income					Consumption				
		Quintile group in 1988									
		I	II	III	IV	V	I	II	III	IV	V
Quintile group in 1987	I	0.64	0.21	0.08	0.05	0.02	0.58	0.25	0.10	0.06	0.02
	II	0.22	0.40	0.21	0.12	0.05	0.25	0.33	0.24	0.12	0.06
	III	0.08	0.24	0.34	0.24	0.10	0.10	0.25	0.30	0.23	0.12
	IV	0.04	0.10	0.27	0.35	0.24	0.05	0.12	0.25	0.33	0.24
	V	0.02	0.05	0.10	0.25	0.59	0.02	0.05	0.11	0.26	0.56
		Quintile group in 1989									
		I	II	III	IV	V	I	II	III	IV	V
Quintile group in 1988	I	0.61	0.25	0.08	0.03	0.02	0.57	0.25	0.11	0.06	0.02
	II	0.24	0.35	0.24	0.12	0.04	0.25	0.32	0.24	0.14	0.06
	III	0.09	0.24	0.31	0.25	0.11	0.11	0.24	0.29	0.24	0.12
	IV	0.04	0.11	0.25	0.35	0.25	0.05	0.13	0.25	0.32	0.26
	V	0.02	0.05	0.10	0.25	0.58	0.03	0.06	0.12	0.25	0.54
		Quintile group in 1990									
		I	II	III	IV	V	I	II	III	IV	V
Quintile group in 1989	I	0.57	0.27	0.10	0.04	0.03	0.55	0.25	0.12	0.06	0.03
	II	0.25	0.34	0.21	0.14	0.06	0.25	0.32	0.22	0.14	0.07
	III	0.11	0.22	0.29	0.27	0.11	0.11	0.23	0.30	0.25	0.12
	IV	0.04	0.12	0.26	0.31	0.26	0.06	0.14	0.23	0.31	0.26
	V	0.03	0.05	0.14	0.24	0.54	0.02	0.06	0.14	0.25	0.53
		Quintile group in 1991									
		I	II	III	IV	V	I	II	III	IV	V
Quintile group in 1990	I	0.58	0.25	0.10	0.06	0.02	0.53	0.26	0.13	0.06	0.02
	II	0.24	0.33	0.23	0.15	0.05	0.27	0.30	0.22	0.15	0.06
	III	0.11	0.24	0.33	0.21	0.11	0.13	0.25	0.30	0.20	0.12
	IV	0.05	0.14	0.21	0.33	0.26	0.05	0.14	0.24	0.31	0.27
	V	0.02	0.05	0.13	0.25	0.56	0.02	0.06	0.11	0.29	0.53
		Quintile group in 1992									
		I	II	III	IV	V	I	II	III	IV	V
Quintile group in 1991	I	0.57	0.25	0.11	0.05	0.02	0.57	0.23	0.13	0.06	0.02
	II	0.25	0.34	0.25	0.11	0.05	0.25	0.33	0.24	0.13	0.05
	III	0.11	0.25	0.29	0.25	0.10	0.12	0.25	0.27	0.25	0.11
	IV	0.05	0.12	0.26	0.33	0.23	0.04	0.14	0.24	0.32	0.25
	V	0.02	0.04	0.09	0.25	0.59	0.02	0.05	0.13	0.24	0.57

Note: Only households surveyed in consecutive years are used for calculating income and consumption quintiles and quintile transition rates. Households with farm income or income from work on farms as the primary income source were excluded. Household income and nondurables consumption were deflated by the aggregate CPI and adjusted using food share based equivalence scales.

Table 11. Quintile Transition Rate Regressions

	Income	Nondurables Consumption
trend × *dum1155*	–0.008*	–0.003
	(0.003)	(0.003)
trend × *dum2244*	–0.008*	–0.003
	(0.003)	(0.003)
trend × *dum33*	–0.009	–0.005
	(0.005)	(0.004)
trend × *dum1off*	0.002	0.000
	(0.002)	(0.001)
trend × *dum2off*	0.004*	0.004
	(0.002)	(0.002)
trend × *dum3off*	0.001	–0.001
	(0.002)	(0.002)
trend × *dum4off*	0.001	–0.001
	(0.003)	(0.002)
dum1155	0.268*	0.257*
	(0.020)	(0.016)
dum2244	0.029	0.021
	(0.020)	(0.016)
dum1off	–0.101*	–0.060*
	(0.017)	(0.014)
dum2off	–0.241*	–0.193*
	(0.017)	(0.014)
dum3off	–0.295*	–0.248*
	(0.018)	(0.015)
dum4off	–0.319*	–0.282*
	(0.020)	(0.016)
constant	0.339*	0.306*
	(0.016)	(0.013)

Notes: The dependent variable consists of income (or consumption) quintile transition rates stacked into one vector with 125 observations (25 transition rates based on 5 quintiles, for 5 pairs of years). Let *qij* denote the transition rate from quintile *i* in year 1 to quintile *j* in year 2. The dummies in the regression are then defined as follows: *dum1155* (q11, q55); *dum2244* (q22, q44); *dum33* (q33); *dum1off* indicates transition rates one off the diagonal off the 5×5 transition matrix; *dum2off* indicates transition rates two off the diagonal, and so on. These dummies were also interacted with a time trend. The excluded dummy in the regression is *dum33*. Standard errors are reported in parentheses below the coefficients. An asterisk indicates statistical significance at the 5 percent level.

that consumption is as variable as income or that consumption closely tracks income. Table 12 reports variances of real household income and consumption over two-year periods, using the subsample of households that were interviewed in consecutive years. Note that the variance of consumption is only half that of income in 1992. Part of that difference is a scale difference, because mean consumption is 84 percent of mean income in that year. But the coefficient of variation for consumption is also less than for income. The explanation for how the quantile transition matrices for income and consumption can look so similar in 1992, despite the fact that consumption is less variable than income in absolute terms, is that the consumption quantiles are more compactly grouped together than are the income quantiles.

Finally, notice in Table 12 that the variance of income spikes up substantially in 1988–89 and 1989–90, the two most turbulent periods of the transition. The increase in variance for consumption is far less pronounced, suggesting some ability of households to smooth consumption over this period. We are exploring this issue further in ongoing work.

VI. Conclusions

We conclude by comparing our evidence on changes in inequality in Poland with the evidence from previously available aggregate statistics. Table 13 reports the statistics that are germane to this comparison. The first row of the table reports the Gini coefficients we have calculated from the HBS microdata for the 1985–92 period. These Ginis are for the distribution of individuals' incomes, using the per capita income of the household in which they reside. Although we calculated many alternative inequality measures, this is the one most comparable to published aggregate statistics. The second row of the table shows the Ginis calculated by the CSO for the OECD, and published in OECD (1997). Observe that we obtain similar values in 1991 and 1992 (0.267 vs. 0.270). But our calculations imply that a much higher level of inequality was present in Poland in 1989 (before the big bang) than do the CSO-OECD figures (0.278 vs. 0.249). Thus, we conclude that the increase in inequality in Poland for the first three years after the big bang that is implied by the CSO-OECD figures is spurious, resulting from serious understatement of the degree of inequality that existed before 1990.

We also obtained a number of other interesting findings. Social transfers appear to have played an important role in preventing increases in inequality during the transition. In fact, inequality in pre-transfer income did increase during the transition, but transfers more than counteracted this. An increase in the generosity of pensions was a particularly important factor in preventing increased inequality. There were important differences across socioeconomic groups in how inequality has changed. In particular, income and consumption inequality grew in households headed by workers, but declined in households headed by farmers and farmer/workers. A key factor in increasing inequality among workers was a substantial rise in education premiums. In fact, we find that only households headed by college graduates experienced a substantial recovery in incomes following the big bang.

Table 12. Variability of Household Consumption and Income

	1987–88	1988–89	1989–90	1990–91	1991–92
Var (*c*)	5.57	6.18	6.67	3.35	2.87
Var (*y*)	8.30	16.10	22.40	6.32	5.73
Var (*c*)/var (*y*)	0.67	0.38	0.30	0.53	0.50
Coeff. var (*c*)	0.270	0.277	0.323	0.255	0.231
Coeff. var (*y*)	0.278	0.364	0.478	0.288	0.275
Coeff. var (*c*)/ Coeff. var (*y*)	0.97	0.76	0.68	0.89	0.84
Observations	8,471	8,874	4,775	4,313	4,224

Notes: Variances of real nondurables consumption and real income were computed over two-year periods for each household surveyed in consecutive years. Sample means of these variances (divided by 10E+8) are reported here. Both variables are measured at 1992Q4 prices and were adjusted by equivalence scales. The coefficient of variation of nondurables consumption is computed as the square root of the variance of consumption shown in the table divided by the sample average of the two-year mean of consumption for each household surveyed in consecutive years; likewise for income.

Table 13. Comparisons with Per Capita Ginis Based on CSO Methodology

	1985	1986	1987	1988	1989	1990	1991	1992
HBS microdata— full distribution	0.270	0.274	0.270	0.272	0.278	0.271	0.266	0.264
CSO-OECD Ginis	—	—	—	—	0.249	0.230	0.260	0.270

Notes: The first row shows per capita Ginis calculated using the HBS microdata. The second row shows Gini coefficients calculated by the CSO for the OECD. Documentation from the CSO suggests that, for 1989–92, Ginis were computed using income decile groups based on per capita income. These Gini coefficients for 1989–92 were obtained from the OECD.

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