IMF Working Paper

WP/99/14

INTERNATIONAL MONETARY FUND

Research Department

Consumption and Income Inequality in Poland During the Economic Transition

Prepared by Michael Keane and Eswar Prasad*

Authorized for distribution by Peter Isard

January 1999

Abstract

This paper challenges the conventional wisdom that income inequality in Poland increased substantially following the economic transition in 1989–90. The results, based on micro data from the 1985–92 Household Budget Surveys, indicate that overall income inequality increased during the initial stages of the transition but then declined to pre-transition levels. Consumption distributions reveal a similar pattern. However, earnings inequality did increase markedly after the transition and the relative well-being of different socio-economic groups was altered. Absolute poverty levels increased during the transition, but this increase is attributable to declines in mean income and consumption rather than to changes in inequality.

JEL Classification Numbers: D31, J31, O15

Authors' E-Mail Addresses: mkeane@atlas.socsci-umn.edu; eprasad@imf.org

---

*Keane: University of Minnesota and New York University. We thank the staff at the Polish Central Statistical Office—especially Wiesław Łagodzinski and Jan Kordos—for assistance with our many inquires about the Polish Household Budget Survey data used in this study. We also thank Krzystof Przybylowski and Barbara Kamińska for excellent assistance with the translations of the survey instruments from Polish, and Vincent Koen for providing some of the CSO aggregate data. We received helpful comments from Benedict Clements, Mark de Broeck, Vincent Koen, Thomas Krueger, Branko Milanovic and seminar participants at Johns Hopkins and CEMFI.

Contents Page

Figures

I. INTRODUCTION

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, 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 the Gini coefficient for household per capita income rose in 17 out of 18 Eastern bloc countries when 1993–95 is compared to 1987–88. He notes that the average Gini increased from .24, a level similar to that in the Scandinavian and Benelux countries, to .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 U.S. in the 1980s by Atkinson, Rainwater, and Smeeding (1995). For Poland, OECD (1997) reports that the Gini increased from .249 in 1989 to .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. Our results challenge the conventional wisdom that inequality increased. The main difference between our work and that of previous authors is that we have obtained for the first time direct access to the detailed micro data of the Polish Household Budget Survey (HBS)

---

[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. Since 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 .256 in 1997 to .284 in 1993 (first half). This is somewhat smaller than the increase implied by the OECD figures, but nevertheless substantial by historical standards.

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 is published by the statistical bureaus of the various countries. But, as we discuss in section 2, the published aggregate income data for Poland and other transition economies do not correspond to Western style measures of household income. However, at least for Poland, meaningful income measures can be constructed using the household level micro data.

Using the HBS micro data, 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 the published aggregate statistics seriously understate the degree of inequality that existed *prior* to 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.

We do not have micro data after 1992, but this is not too great a limitation for two reasons. First, the published aggregate statistics imply that most of the increase in inequality in Poland occurred from 1990 to 1992. There was some additional increase in 1993, but after that the published inequality measures are rather flat. Second, even if we take the published aggregate statistics for 1993–96 as reliable, and allow that the inequality increase they imply for 1992 to 1993 is genuine, we can still estimate that roughly 90 percent of the total reported increase in inequality from 1989 to 1996 is spurious.

In the HBS micro data 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 despite the fact that transfer programs in Poland, as in other transition economies, tend to be class based rather than income based.

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 micro data that allows us to examine

---

[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 micro data. A long negotiation process by the first author during 1992–93 led to its release. Subsequently, the micro data for just the first six months of 1993 was released to the World Bank, and this data is used in World Bank (1995) and Milanovic (1998).

consumption inequality in a meaningful way. As we discuss in Section 2, the aggregate consumption figures that were published by the Polish CSO, as well as by other former communist countries, did not correspond to Western style measures. After constructing Western style consumption measures from the micro data of the HBS, we again find no evidence of increased inequality during the transition.

One reason for 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 be exaggerated. The existing social safety net appears to be doing 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 micro data enables us to examine whether certain socioeconomic groups have been relative winners or losers in the transition. We find that income differentials by education level increased rapidly after the big bang. Gorecki (1994) previously noted such a pattern in the aggregate data released by the CSO. Prior to 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.

Another important issue is the extent to which the absolute level of poverty increased during the transition. A number of prior studies have examined income-based poverty rates but, as is well-known, these measures will tend to exaggerate poverty to the extent that income is variable over time and households can smooth consumption. We use the micro data from the HBS to construct consumption-based poverty rates. We do find that the absolute levels of poverty, as measured by both income and consumption, increased substantially during the transition, but that most of this increase is attributable to the decline in mean income and consumption rather than to changes in inequality.

In the next section we describe the prior research on income inequality in Poland during the transition in more detail. Then, in Section 3 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 micro data 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. Section 4 presents our empirical results, and Section 5 concludes.

II. Review of Prior Research

There exist several other studies that have examined income inequality in Poland during the transition. But they report rather contradictory results. This despite the fact that they all use income data from the HBS. For instance, OECD (1997) reports (see Figure 22, page 86) that the Gini based on household per capita income for Poland is .25 in 1989, drops to .23 in 1990 and then rises substantially to .26, .27, .29 over the 1991 through 1993 period. 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 .260, .255 and .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 micro data for just the first six months of 1993. 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,[5] which we henceforth refer to as the *Surveys*. The accuracy of these approximations is certainly subject to question.

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-quarter 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 micro data, we are able to make important adjustments to income in order to obtain a meaningful measure (in this example, by calculating net farm income).[6]

---

[5]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.

[6]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.

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, page 41), who reports that in 1993 the Gini for consumption is .31, which substantially exceeds the Gini of .28 that he calculates for income. Also, on page 33 he reports that in 1993 the ratio of consumption to income is 1.30, an unreasonably high figure.

It is again our access to the detailed micro data that allows us 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 .894 to .955 range during the 1985–92 period.

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.

Besides the prior work on income inequality, there also exist a number of studies of the evolution of poverty in Poland. For instance, Szulc (1994, 1995), using data from the HBS for 1980–1992, finds that the fraction of people with consumption below the official poverty line—the "social minimum" defined by the Institute of Labor and Social Affairs (ILSA)—was in the 15 to 20 percent range over 1985-9, increased to 34.2 percent in 1990 and to 40.3 percent in 1992. Szulc (1995) also reports changes in the poverty rate broken down by type of family, source of income (i.e., occupation), education level and age, but only for the pre-1990 period. Milanovic (1992) provides a similar analysis for the 1978–88 period, and Milanovic (1993) calculates poverty rates of 17.3 percent for 1989 and 34.4 percent for 1991[7]. An important limitation of these studies is that they look only at income—and not consumption-based measures of inequality. Also, the social minimums defined by the ILSA are

[7]Milanovic (1993) describes how the poverty rates are approximated from the aggregate statistics in the *Surveys*. Recall that the *Surveys* report the number of households in various income intervals. According to Milanovic (1993) footnote 5, "All individuals belonging to a given income class are considered poor if the upper income limit of that class, adjusted for the number of equivalent consumption units, is less than the poverty line.... if the poverty line is between the two limits, a proportion of the individuals is considered poor." He does not describe exactly how this proportion is determined.

generally considered too generous to represent a reasonable absolute poverty line (see OECD (1997), page 91). Hence, we will examine alternative poverty definitions.

III. THE HOUSEHOLD BUDGET SURVEYS

The Polish Central Statistical Office has been collecting detailed micro data 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 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 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 their 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 obtain a 16 category classification of total household expenditure, the categories being: (1) Food; (2) Alcohol and tobacco; (3) Clothing and footwear; (4) House purchases; (5) House construction; (6) Household nondurables (incl. energy); (7) Household durables (incl. 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 utilize 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 data bases, 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 (FBS) collected in the Soviet Union suffer from a number of severe problems. First, the data are not a representative sample of the population (as families were mainly selected 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 FBS do not provide true household or individual level income data.

After the breakup of the Soviet Union, some of the Former Soviet States (FSU) maintained the same primitive data collection methods, while others (including all 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 is poor throughout, in the latter because the improved data from after the breakup is 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 up until the present time. For this reason, the HBS is the highest quality and most consistent micro data 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.

The HBS is in fact also superior to most if not all Western micro data sets in the sense of providing detailed data on income by source and consumption broken down by detailed category within a single data source (e.g., data sets available for the United States or Britain typically focus on income or consumption, but not both). Apparently, there is a time honored tradition of doing detailed micro data collection in Poland.

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 Szulc, Milanovic, and Gorecki) 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 on site at the CSO. This greatly limited the kind of analysis that was feasible, for obvious reasons.

IV. Results

A. 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 versus -11.4 percent). Thus, there was no aggregate smoothing of the 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 1985–92 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 .926, which seems reasonable. The sample is 50 percent urban, and 57 percent of the household heads are males in the 31–60 age range; 54 percent of the households include a married couple, 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 it contains 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 five fixtures. In Table 2, we list the percentage of households with each of the five fixtures. Overall means for the durable stocks are not very meaningful because many of them change drastically over time.

B. Equivalence Scales

As noted above, most of the prior work on income distributions in Poland has simply looked at per capita household income, and not attempted to account for household economies of scale by employing equivalence scales. The exception is the work by Szulc (1994, 1995) who analyses 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 micro data 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. 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:

$$w_h = \alpha_1 + \beta \log (x_h / k_h P) + \Sigma \gamma_{1j} \log p_j \qquad \textbf{(1)}$$

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 2) alcohol and tobacco, 3) nonfood commodities, and 4) services. Note that the γ_{1j} must sum to zero to satisfy zero degree homogeneity in prices and total expenditure. The aggregate price index P is defined by $\log P = \alpha_0 + \Sigma \alpha_k \log p_k + (1/2)\Sigma\Sigma\gamma_{kj} \log p_k \log p_j$. But, as noted by Deaton and Muellbauer (1980), share weighted aggregate price indices will tend to be highly correlated with P. Thus, we estimate (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:

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

We then specify $-\beta \log k_h = \Sigma_j \phi_j D_{hj}$, where the D_{hj} are dummy variables indicating whether household h has characteristic j, where j indexes the set of demographic categories

---

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

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

listed in Table 1. 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 to 60 age range, forms the omitted category. We estimated (2) including quarter dummies. Given the estimated food share equation, we estimate the equivalence scale k_h as:

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

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

A potential problem with estimation of (2) is that denominator bias is present if nondurable consumption is measured with error. Thus, we have estimated (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 .64 to .84.

Table 3 reports OLS and 2SLS estimates of the food share equation. Note that the coefficient on log real nondurable consumption changes from -.195 to -.263 when we instrument. In the 2SLS regression the three relative price terms taken together imply a coefficient of .1897 on the log price of food. This implies that a one 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 .186 for the own food price coefficient using annual British data for 1954–74.[10] This agrees with our estimate to the third decimal place. Their estimate of the real nondurable consumption coefficient was -.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

[10]Deaton and Mullbauer (1980) examined allocation of expenditure across eight nondurable commodity categories. Thus, they treat durables as weakly separable, as we do here.

respective sample means. The average food share over the whole sample period is .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 .55 to .66, between 1989:4 and 1990:1. This is the immediate impact of the drop in real incomes following the big bang.

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, following the liberalization of food prices, 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 .55 in 1989:1 up to .70 in 1989:4, and that it would have then plummeted to .62 in 1990:1 and further to .49 in 1992:4. 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 .56, .56, .61 and .58. But changes in household demographics over the sample period had little effect on food shares.

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 2nd stage food share regression separately by year, and constructed equivalence scales separately for each year of the sample. We found 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.

C. Inequality Measures

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 five alternative equivalence scales. These are the food share based, CSO, OECD and McClements scales reported in Table 4, along with the simple per capita scale obtained by

---

[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. Since there was no properly functioning market for housing (either rental or owner occupied) we cannot impute the true level of housing consumption.

dividing household income by household size. Note that the four 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 1985–88, and that inequality actually fell from 1989 through 1992. The Gini based on the food share scale implies a somewhat sharper decline in inequality in 1989–92 (from .280 to .233) than do the Ginis based on the other three scales.

The Ginis based on simple on per capita household income are consistently about .015 to .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 .25 in 1989 to .27 in 1992. In contrast we obtain a decline from .293 to .267 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 8 per capita household income intervals. Based on that data, Milanovic (1998) calculates an approximate Gini of .252. Using the same data, we obtain a similar Gini value of .248.[12] This compares to the value of .292 that we calculate from the HBS micro data. Thus, prima facie, it appears that use of grouped data does lead to downward bias in the Gini. However, if we take the HBS micro data for 1987, group households into the same 8 per capita income intervals, and approximate the Gini based on that information, we obtain .281. 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 in 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 their household income measure. If we do the same, then for 1987 we obtain a Gini of .287. 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) reports (see Figure 22, p. 86) Gini values based on household per capita income of .25 and .23. Similarly,

---

[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* contains 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 Gini based on this data by assuming that all households in an interval are at the mean of per capita income for that interval.

Milanovic (1993) reports Gini values of .260 and .255. All these figures are based on various aggregate income decile data provided by the CSO, and we are uncertain of the sources of the (minor) discrepancies. Our Ginis based on per capita household income for those two years are much higher, at .293 and .281 respectively. If we group our data into deciles and then approximate the Ginis we get .286 and .278, again not a large change. And if we leave in gross farm income instead of net farm income we get .296 and .284. 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 .26-.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.

We would argue that the inequality measures that we have calculated directly from the HBS micro data 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. Since the choice of equivalence scale appears to make little difference to our results, we will henceforth report results using the food share based scale 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 .402 to .428 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 .280 to .233 during 1989–92 indicates. Nevertheless, it is possible that the growth in inequality of

---

[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.

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 in Section E.

Now we turn to examination of changes in consumption inequality. Again, the Ginis based on the four equivalence scales that allow for household economies of scale all show a similar pattern. Inequality grows from 1985–89 and then declines from 1989–92. The decline from .264 to .232 indicated by the food share based scale is again sharper than for the other scales. Similarly, the Gini for nondurable consumption declines from .220 in 1989 to .206 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. In fact, some important differences do emerge. The mean log deviation for both income and nondurable consumption does spike up in 1989, a turbulent year of hyperinflation that saw the removal of food price controls in August. But, aside from that spike, it is rather flat. In contrast, the coefficient of variation of income, which was rather flat from 1986–89, drops sharply in 1990–92. Thus, we find evidence of decreasing income inequality after the big bang only with the measures that are sensitive to changes near the median or at the high end of the distribution. Furthermore, the coefficient of variation for nondurable consumption is similar in 1991–92 to its levels in 1987–98. So at the high end of the consumption distribution there is little indication of change. Despite these differences, it is important to bear in mind the big picture: none of the inequality measures show an increase in inequality in either income or nondurable consumption after the big bang.

D. 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 for real household income for the years 1988, 1989, 1990, and 1990.[14] An Epanechnikov kernel with a bandwidth of 4000 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.

---

[14]No adjustment is made for household size.

The change in the shape of the densities between the years 1988–99 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 in section 4.3. 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 to 58 thousand 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 to 58 thousand zloty range, and found that these households receive over 80 percent of their income from pensions (80.5 percent in 1988, 82.2 percent in 1992). These percentages drop off quickly as household income rises above the 58 thousand zloty level. The percentage 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 to 58 thousand zlotys got 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 zloty in 1988 to 131,563 zloty in 1992, the mean real pension actually rose from 29,811 to 35,258. This resulted from legislation that took effect in 1991 that made pensions substantially more generous. Hence, it is clear from our results 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 importantly to the reductions in inequality measures that we have noted.[15]

E. Quantile Ratios and Shares

Another common way to summarize changes in inequality is to examine quantile ratios. Unlike the scalar inequality measures considered in Section 4.3, examination of a set of

---

[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.

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

Figure 5 reports values of the .10, .25, .50, .75 and .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 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.

F. 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 in the transition. In this section we turn to an examination of how different groups fared in terms of wages, 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 farmer/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 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

---

[16]The reason for the difference in the scales is that the mean numbers of persons in worker, farmer, worker/farmer and pensioner 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.

importantly to a reduction in inequality (see also the discussion in Section D).[17] The main impetus behind the improved relative position of pensioners was a substantial increase in pension levels that took place in 1991.

Next we consider how various education categories fared. In Table 7 we report standard human capital earnings functions (see Willis (1986)) separately for each year of the sample. These log earnings regressions are estimated using employed workers aged 25–60 who report labor income as their primary source of income. Of particular interest are the coefficients on the five educational category dummies (primary school only is the omitted category). Note that the college degree coefficient increases from .485 in 1989 to .599 in 1992. This implies an increase in the mean wage of college graduates relative to primary school graduates of approximately 12 percent. The relative mean wage of high school graduates also increases, but by a smaller amount. The coefficient on some high school drops from .135 in 1989 to essentially zero in 1992, indicating that mean earnings of high school drop outs falls to the same level as for those with only primary school.

Table 8 reports the fractions of households that fall in each quintile 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 quintile. 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 quintile, but by 1992 this had fallen to 9.5 percent.

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

G. Quantile Regressions

We follow Buchinsky (1994) in noting that a more complete picture of how various groups fared during the transition can be obtained by use of quantile regression. 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 labor income on demographic characteristics of workers. These characteristics were dummies for the six education categories (with primary school the omitted category), labor market experience (i.e., age-education-six), experience squared,

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

urban and sex. The sample included full time workers with no pension income in the 25–60 age range (just as in the linear earnings regressions reported earlier). Regressions for the .10, .25, .50, .75 and .90 quantiles were run.

Table 9 reports the college and high school premia (relative to primary school) obtained from the quantile regressions. Note that the college premium jumps substantially from 1988 to 1991 at all quantile points. For example, median income was approximately 38 percent higher for college graduates than primary school graduates in 1988, and this premium increases to 54 percent in 1991. The changes in the college premia at the other quantile points are similar, except that the increases tend to be even greater at the higher quantiles.

In general, the high school premium also grows from 1988 to 1991. But here, the pattern that increases are greater at the higher quantiles is much more pronounced. For instance, a high school degree raised the .90 quantile of the income distribution by about 20 percent in 1988 and 32 percent in 1992, an increase of 12 percentage points. But the amount by which a high school degree raises the .10 quantile stays flat at roughly 22 percent throughout the whole sample period.

The main new pattern revealed by the quantile regressions is that both the high school and college premia increased more at the higher quantiles than at the lower quantiles during the sample period. For example, in 1985 a college degree raised the .10 quantile by about 38 percent but only raised the .90 quantile by about 32 percent. By 1992 the figures were 47 and 64 percent respectively, so the increase was more than three times as great at the .90 quantile.

Table 10 reports returns to experience, defined as the derivative of log real household income with respect to labor market experience of the household head. Since experience enters the regressions as a quadratic, we report the derivatives evaluated at 5, 15 and 25 years of experience. The striking finding here is that returns to experience are much smaller than found in Western data sets—both before and after the "big bang." In fact, a striking break is also apparent in 1990—after that, the returns to experience become even smaller.

The figures in Table 10 are in percentage terms. So, for example, in 1985 for people with five years of experience, an additional year raises median earnings by about 2.90 percent. For the United States in 1985, Buchinsky (1994, p. 419) reports a figure of 4.97 percent, which is about 70 percent greater. Also, for those with five years of experience the return to experience appears to be slightly greater at the higher quantiles. For the United States, Buchinsky found the exact opposite pattern—i.e., that for new labor market entrants the return to experience is (usually) greater at the lower quantiles.

For people with 15 years of experience in 1985 the return to experience at the median is quite small (1.64 percent), and by 25 years it is almost zero (0.37 percent). Results are similar at other quantile points. In contrast, Buchinsky (1994) p. 420 finds for the United

States in 1985 that the return to experience for people with 15 years of experience is about 2.9 to 3 percent at all quantile points, which is nearly double the rate in Poland.

These results suggest important differences between the wage structure in Poland and that in western capitalist countries in the period prior to the big bang. But the differences remain large in 1992. Although by that time education premia rose to a level more in line with those found in western countries, the divergence in returns to experience became even greater.

It is also interesting to examine how changes in the overall well being of households were influenced by the education level of the household head. To examine this issue, we also ran quantile regressions of log real quarterly household *per equivalent* income on characteristics of the household head. Table 11 reports conditional quantiles of log real household income based on the education level of the household head. Note that the log earnings for all education groups drop substantially at all quantile points from 1989 to 1990. The drops tend to be larger at the higher quantile points (e.g., at the .90 quantile they range from 30 to 33 percent, while at the .10 quantile they range from 25 to 28 percent). Thus we see a decline in within education group inequality 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 continues to fall from 1990 to 1992, with the drops much more pronounced at the higher quantiles. In contrast, for households headed by college graduates, the .10, .25 and .50 quantiles recover a bit, while the .75 and .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 .90 quantile falls. Thus, the only group of households that experiences actual recovery in earnings after 1990 is the group with a college educated head.

H. Within and Between Group Decompositions of Inequality

In this section we address the question of the extent to which inequality is within vs. between group, 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 12, 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 than between group, which is not surprising given the coarse nature of the grouping.

The interesting finding in Table 12 is that the changes in within-group inequality are very different for the different groups. For instance, consider the mean log deviation of income, which is sensitive to changes at the low end of the distribution. As we noted in Section C, this measure blips up in 1989 but is otherwise rather flat throughout the sample period. But this masks the clear trends that exist for different groups. For households headed by workers, the mean log deviation increases substantially after the big bang—e.g., from 5.8 in 1988 to 7.5 in 1992. In contrast, for households headed by farmers and farmer/workers, the mean log deviation declines substantially from 1988 to 1992. The same pattern emerges for nondurable consumption, and when we look at half the square of the coefficient variation as the measure of inequality. It is possible that the general perception that inequality increased in Poland after the big bang stems at least in part from the increase in inequality among workers observed in Table 12.

I. Poverty Rates

In this section we consider changes in poverty rates during the transition. Of course, the poverty line is a rather arbitrary concept. But there is widespread agreement that the poverty lines developed by the Institute of Labor and Social Affairs in Warsaw are now "overly generous" (see OECD (1997, p. 91), Milanovic (1998, p. 66). Using these poverty lines, Szulc (1994, 1995) calculates that the percentage of households in poverty rose from 16.7 percent in 1989 to 34.2 percent in 1990, and further to 40.3 percent in 1992. But the poverty line appears to lose its meaning in the local context when such a large fraction of the population is counted as poor.

We have calculated poverty lines by first constructing the median of per equivalent household income using pooled data for the whole 1985–92 period. Then, we alternately define a household as being in poverty if it has per equivalent income below either one-half or two-thirds of that median. In the first panel of Table 13, we report the fraction of the population living in households with per equivalent income below each of those thresholds in each year. For instance, the fraction of the population below the one-half median threshold jumps from 3 percent in 1989 to 9 percent in 1990, but then stays fairly stable. Thus, while poverty jumped in the immediate aftermath of the big bang, it did not grow over the subsequent three years.

It is interesting to examine the extent to which transfer payments alleviate poverty. We have conducted the experiment of removing transfers from household income, and redistributing the transfers equally to all households based on their number of equivalent units. Such an experiment of course assumes away any behavioral response of households to the change in transfer rule, but it does reveal the extent to which transfers alleviate poverty in a purely accounting sense. It is interesting that in 1992 the fraction of people below the two-thirds median threshold drops from 32 percent to 26 percent as a result of transfers, while the fraction below the one-half median threshold drops from 19 percent to 9 percent. Perfect targeting of transfers would imply that the percentage below the one-half median threshold should be reduced to zero before the percentage below two-thirds median is reduced at all. Thus, the fact that transfers appear to do less to reduce the fraction of people below the one-half median threshold suggests that targeting to the least well off households could be substantially improved.

Previous studies on poverty in Poland have not reported poverty rates based on consumption data. As has been noted by a number of authors (see, e.g., Triest (1998)), poverty rates for the U.S. and other Western countries are generally found to be lower when consumption data are examined rather than income data, presumably because much of the variance in income is transitory and households attempt to smooth consumption over time. In the last two columns of Table 13 we report poverty rates based on household (per equivalent) nondurable consumption, again using the one-half and two-thirds of the median as thresholds. These poverty rates are indeed lower than those based on income, but not substantially so.

However, for some types of households, the income and consumption based poverty rates do differ greatly. For instance, in 1992 the income based poverty rates (based on two-thirds of the median) are 20 percent, 39 percent, 32 percent and 36 percent for households headed by workers, farmers, farmer/workers and pensioners respectively. The corresponding consumption based rates are 17 percent, 30 percent, 32 percent and 28 percent. Thus, for households headed by farmers and pensioners, the consumption based poverty rates are indeed substantially lower than the income based rates.

V. Conclusions

We conclude by comparing our evidence on changes in inequality in Poland with the evidence from previously available aggregate statistics. Table 14 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 micro data 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 almost identical values for 1992 (.267 versus .270). But our calculations imply that a much higher level of inequality was present in Poland in 1989 (prior to the big bang) than do the CSO-OECD figures (.293 versus .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 prior to 1990.

Since we do not have the HBS micro data for the period after 1992, we cannot be sure what it implies for subsequent changes in inequality. Note that the GUS-OECD figures imply that inequality jumped substantially from 1992–93, and then stayed fairly flat through 1996. But we are uncertain whether the jump in 1992–93 is genuine, because there were substantial changes in the HBS sample design in that year. First, the sampling scheme was updated to obtain a more representative sample of the self-employed (in earlier years this group was under-represented). If that were the only change, we would suspect that the observed jump in inequality from 1992–93 stemmed from increased representation of this group. However, it is perhaps implausible that such a large jump in inequality resulted only from improved representation of the self-employed, because in 1993 they only made up 3.4 percent of the sample (rising to 4.0 percent in 1996). And, in fact, another important change in survey design also occurred in 1993—income reporting was changed from quarterly to monthly. We would expect that the distribution of monthly income is more unequal than that of quarterly income. Thus, much of the reported increase in inequality between 1992 and 1993 may be due to the shift to monthly income reporting.

Nevertheless, it is interesting to consider the result if we take the entire increase in inequality over 1992–96 implied by the CSO-OECD Ginis as genuine. That increase is from .270 to .300. If our Gini estimate increased by the same amount, it would be .297 in 1996. Thus, we would obtain an increase from .293 in 1989 to .297 in 1996, as compared to the increase from .249 to .300 indicated by the CSO-OECD numbers. By this reckoning, we have that approximately 90 percent of the increase in overall inequality reported for Poland since the big bang is spurious.

We also obtained a number of other interesting findings. We found that transfers 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 premia. In fact, we find that only households headed by college graduates have experienced substantial recovery in incomes since the big bang. Although education premia have risen since the big bang, returns to experience have fallen to levels well below those observed in Western economies. Finally, our analysis of poverty rates suggests that transfers could be better targeted toward the lowest income groups.

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 (end-year)	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 obtained from various publications of GUS and IMF. Aggregate consumption is deflated by the CPI. Net income and total consumption from household surveys also deflated by the CPI.

Table 2a. 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
Farm-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

Table 2b. Summary Statistics (concluded)

	Mean	Standard Deviation
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
Kid, 0-7	0.42	0.76
Kid, 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
Fixtures		
Running water	0.83	0.37
WC	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	

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)
kid, 0-7	0.021 *	(0.000)	0.026 *	(0.000)
kid, 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.

Table 4. Equivalence Scales as a Function of Household Composition

Household Type:	GUS	OECD	McClements	Food-Share Equations OLS	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 kid 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.

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

	1985	1986	1987	1988	1989	1990	1991	1992
	Gini Coefficients							
Total income								
Food-share based eqv. scale	0.256	0.273	0.270	0.278	0.280	0.261	0.244	0.233
Urban	0.202	0.208	0.201	0.205	0.227	0.220	0.215	0.214
Rural	0.328	0.330	0.318	0.327	0.316	0.291	0.263	0.254
CSO equivalence scale	0.257	0.274	0.275	0.279	0.280	0.265	0.254	0.249
McClements equivalence scale	0.253	0.270	0.268	0.275	0.276	0.259	0.245	0.236
OECD equivalence scale	0.257	0.274	0.273	0.278	0.280	0.263	0.249	0.241
Per capita	0.274	0.291	0.292	0.294	0.293	0.281	0.271	0.267
Income excluding transfers								
Food-share based eqv. scale	0.378	0.392	0.391	0.407	0.402	0.402	0.412	0.428
Total consumption								
Food-share based eqv. scale	0.232	0.238	0.247	0.250	0.264	0.244	0.235	0.232
Urban	0.224	0.224	0.232	0.237	0.253	0.238	0.235	0.226
Rural	0.241	0.249	0.255	0.258	0.270	0.248	0.230	0.229
CSO equivalence scale	0.239	0.246	0.256	0.257	0.267	0.251	0.250	0.252
McClements eqv. scale	0.231	0.238	0.248	0.249	0.262	0.243	0.238	0.238
OECD equivalence scale	0.234	0.241	0.251	0.252	0.264	0.246	0.242	0.242
Per capita	0.252	0.260	0.270	0.270	0.278	0.264	0.265	0.270
Nondurables consumption								
Food-share based eqv. scale	0.196	0.200	0.206	0.213	0.220	0.208	0.207	0.206
	Coefficient of Variation (variables adjusted by food-share based equivalence scales)							
Total income	0.596	0.682	0.652	0.664	0.674	0.579	0.576	0.497
Income excluding transfers	0.817	0.909	0.875	0.901	0.901	0.834	0.897	0.849
Nondurables consumption	0.380	0.387	0.403	0.406	0.423	0.391	0.402	0.400
	Half the Square of Coefficient of Variation (variables adjusted by food-share based equivalence scales)							
Total income	0.186	0.242	0.221	0.230	0.242	0.173	0.173	0.130
Income excluding transfers	0.334	0.416	0.385	0.406	0.409	0.348	0.409	0.362
Nondurables consumption	0.076	0.078	0.083	0.085	0.094	0.080	0.084	0.084
	Mean Log Deviation (variables adjusted by food-share based equivalence scales)							
Total income	0.080	0.082	0.080	0.087	0.104	0.082	0.085	0.081
Income excluding transfers	0.214	0.201	0.201	0.219	0.249	0.243	0.276	0.299
Nondurables consumption	0.061	0.064	0.067	0.070	0.077	0.065	0.069	0.068

Table 6. Quantile 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 quantile ranges for that variable.

Table 7. Earnings Regressions

(dependent variable--log monthly real labor income)

	1985	1986	1987	1988	1989	1990	1991	1992
Experience	0.032 *	0.030 *	0.027 *	0.027 *	0.031 *	0.024 *	0.026 *	0.024 *
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
Experience squared	-0.005 *	-0.001 *	0.001 *	-0.001 *	-0.001 *	0.000 *	0.000 *	0.000 *
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Male	0.251 *	0.269 *	0.278 *	0.246 *	0.246 *	0.229 *	0.222 *	0.209 *
	(0.005)	(0.005)	(0.005)	(0.005)	(0.006)	(0.006)	(0.007)	(0.011)
Private enterprise	0.103 *	0.112 *	0.120 *	0.081 *	0.111 *	0.158 *	0.199 *	0.146 *
	(0.021)	(0.016)	(0.015)	(0.015)	(0.013)	(0.012)	(0.010)	(0.017)
Urban	0.041 *	0.046 *	0.065 *	0.056 *	0.040 *	0.042 *	0.047 *	0.063 *
	(0.006)	(0.006)	(0.005)	(0.005)	(0.006)	(0.006)	(0.007)	(0.011)
College degree	0.432 *	0.452 *	0.457 *	0.471 *	0.485 *	0.508 *	0.571 *	0.599 *
	(0.010)	(0.010)	(0.009)	(0.011)	(0.012)	(0.012)	(0.012)	(0.020)
Some college	0.331 *	0.300 *	0.307 *	0.327 *	0.332 *	0.387 *	0.388 *	0.413 *
	(0.029)	(0.025)	(0.026)	(0.025)	(0.034)	(0.033)	(0.043)	(0.059)
High school	0.203 *	0.211 *	0.221 *	0.234 *	0.236 *	0.241 *	0.269 *	0.265 *
	(0.007)	(0.007)	(0.007)	(0.008)	(0.009)	(0.008)	(0.008)	(0.015)
Basic vocational training	0.121 *	0.127 *	0.119 *	0.120 *	0.106 *	0.095 *	0.100 *	0.100 *
	(0.007)	(0.007)	(0.006)	(0.007)	(0.008)	(0.009)	(0.008)	(0.014)
Some high school	0.080 *	0.152 *	0.122 *	0.144 *	0.135 *	0.092 *	0.121 *	-0.013
	(0.017)	(0.016)	(0.017)	(0.019)	(0.021)	(0.016)	(0.023)	(0.044)
Constant	10.637 *	10.668 *	10.618 *	10.679 *	10.769 *	10.488 *	10.420 *	10.312 *
	(0.017)	(0.018)	(0.016)	(0.017)	(0.020)	(0.021)	(0.022)	(0.036)
Adjusted Rsquared	0.393	0.389	0.416	0.357	0.286	0.256	0.306	0.308
Number of observations	16,812	18,328	20,116	19,735	19,536	19,410	17,863	6,857

Notes: The sample includes employed workers between the ages of 25 and 60; who report labor income as their primary source of income. The regressions also included 40 industry dummies. Robust standard errors are reported in parentheses An asterisk indicates statistical significance at the 5 percent level.

Table 8a. Fractions of Various Groups (based on education and age of household head) in Different Quantile Ranges

	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

Table 8b. Fractions of Various Groups (based on education and age of household head) in Different Quantile Ranges

	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.

Table 9. Education Premia

	Quantile = 0.10		Quantile = 0.25		Quantile = 0.50		Quantile = 0.75		Quantile = 0.90	
	COL	HS	COL	HS	COL	HS	COL	HS	COL	HS
1985	0.38	0.19	0.36	0.18	0.35	0.16	0.34	0.16	0.32	0.15
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)	(0.02)
1986	0.38	0.21	0.36	0.20	0.36	0.17	0.36	0.16	0.34	0.16
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)
1987	0.41	0.23	0.38	0.20	0.36	0.18	0.37	0.18	0.36	0.18
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)
1988	0.37	0.22	0.35	0.21	0.38	0.19	0.38	0.19	0.39	0.20
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.13)	(0.01)
1989	0.41	0.21	0.41	0.21	0.44	0.21	0.44	0.21	0.44	0.23
	(0.02)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)
1990	0.49	0.24	0.49	0.24	0.49	0.22	0.53	0.23	0.54	0.24
	(0.02)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.02)
1991	0.51	0.25	0.51	0.25	0.54	0.26	0.59	0.28	0.63	0.33
	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)
1992	0.47	0.22	0.49	0.23	0.51	0.24	0.57	0.28	0.64	0.32
	(0.04)	(0.02)	(0.01)	(0.01)	(0.02)	(0.02)	(0.03)	(0.02)	(0.04)	(0.02)

Notes: COL--college degree; HS--high school degree. The full set of regressors included an urban dummy, experience, experience squared, a dummy for sex, and six education dummies. The dependent variable is log real quarterly labor income.

The excluded education dummy is for primary school degree. Hence, the coefficients reported above are interpretable as the income premia, in percent, for workers with a college or high school degree, respectively, relative to workers who have only a primary school degree. Robust standard errors are reported in parentheses below the coefficient estimates.

Table 10. Returns to Experience

	Quantile = 0.10			Quantile = 0.25			Quantile = 0.50			Quantile = 0.75			Quantile = 0.90		
Experience (in years):	5	15	25	5	15	25	5	15	25	5	15	25	5	15	25
1985	2.63	1.54	0.45	2.60	1.50	0.41	2.90	1.64	0.37	3.11	1.73	0.36	2.92	1.68	0.44
	(0.15)	(0.08)	(0.05)	(0.12)	(0.06)	(0.04)	(0.12)	(0.05)	(0.05)	(0.25)	(0.11)	(0.09)	(0.22)	(0.11)	(0.05)
1986	2.43	1.41	0.39	2.46	1.39	0.31	2.46	1.34	0.21	2.37	1.26	0.14	2.61	1.41	0.21
	(0.15)	(0.07)	(0.05)	(0.14)	(0.07)	(0.04)	(0.12)	(0.05)	(0.03)	(0.22)	(0.11)	(0.04)	(0.25)	(0.12)	(0.07)
1987	2.36	1.36	0.36	2.44	1.38	0.33	2.37	1.29	0.21	2.16	1.15	0.14	2.08	1.07	0.06
	(0.20)	(0.10)	(0.06)	(0.13)	(0.06)	(0.03)	(0.09)	(0.05)	(0.04)	(0.19)	(0.10)	(0.04)	(0.22)	(0.10)	(0.07)
1988	2.09	1.23	0.36	2.19	1.24	0.28	2.44	1.35	0.26	2.54	1.34	0.14	2.25	1.20	0.15
	(0.15)	(0.09)	(0.04)	(0.11)	(0.06)	(0.03)	(0.13)	(0.07)	(0.03)	(0.14)	(0.07)	(0.05)	(0.25)	(0.14)	(0.06)
1989	2.80	1.50	0.20	2.53	1.35	0.17	2.52	1.32	0.12	2.50	1.29	0.09	2.47	1.25	0.02
	(0.14)	(0.08)	(0.04)	(0.16)	(0.09)	(0.06)	(0.14)	(0.08)	(0.05)	(0.16)	(0.09)	(0.05)	(0.22)	(0.12)	(0.08)
1990	2.39	1.31	0.24	2.10	1.17	0.24	1.94	1.06	0.18	1.85	1.00	0.15	1.36	0.70	0.04
	(0.24)	(0.12)	(0.06)	(0.20)	(0.10)	(0.04)	(0.15)	(0.08)	(0.03)	(0.15)	(0.09)	(0.05)	(0.31)	(0.17)	(0.07)
1991	2.07	1.20	0.33	2.06	1.20	0.34	1.87	1.11	0.34	1.88	1.08	0.29	2.01	1.11	0.21
	(0.24)	(0.12)	(0.04)	(0.19)	(0.09)	(0.05)	(0.15)	(0.07)	(0.05)	(0.16)	(0.08)	(0.04)	(0.24)	(0.12)	(0.08)
1992	1.85	1.16	0.46	1.69	1.05	0.42	1.57	0.94	0.32	1.47	0.90	0.34	2.05	1.19	0.34
	(0.44)	(0.20)	(0.12)	(0.32)	(0.17)	(0.06)	(0.23)	(0.11)	(0.08)	(0.32)	(0.15)	(0.07)	(0.44)	(0.23)	(0.13)

Notes: The numbers reported above are derivatives of log real quarterly labor income with respect to the experience level of the worker, evaluated at the indicated experience levels. Standard errors for these derivatives, based on robust standard errors for the regression coefficients, are reported in parentheses. The full set of regressors included an urban dummy, experience, experience squared, a dummy for sex, and six education dummies.

Table 11. 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--college degree; HS--high school degree; VOC--basic vocational training; PS--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.

Table 12. Decomposition of Inequality Measures

	Income								Consumption							
	1985	1986	1987	1988	1989	1990	1991	1992	1985	1986	1987	1988	1989	1990	1991	1992
Mean log deviation (x100)																
Total	8.0	8.2	8.0	8.7	10.4	8.2	8.5	8.1	6.1	6.4	6.7	7.0	7.7	6.5	6.9	6.8
Between-group	0.9	0.8	0.6	0.9	1.4	0.5	0.4	0.5	0.4	0.3	0.2	0.3	0.5	0.3	0.3	0.3
Within-group	7.1	7.4	7.4	7.8	9.0	7.7	8.1	7.6	5.7	6.0	6.5	6.8	7.2	6.2	6.6	6.5
Workers	5.6	6.0	6.0	5.8	7.3	7.0	7.2	7.5	5.3	5.7	6.0	6.2	6.8	6.0	6.7	6.8
Farmers	13.5	15.2	12.2	13.7	18.4	11.9	14.7	9.8	7.0	7.2	8.1	8.8	9.0	7.7	6.9	6.5
Mixed, worker-farmers	8.7	6.7	8.1	10.1	8.9	8.5	7.4	8.1	5.6	6.1	6.5	7.0	7.2	6.0	6.1	6.1
Pensioners	7.0	6.9	6.8	6.5	7.0	6.2	7.1	6.9	6.5	6.2	6.5	6.7	7.3	5.8	6.4	6.1
Half the square of the coefficient of variation (x100)																
Total	18.6	24.2	22.1	23.0	24.2	17.3	17.3	13.0	7.6	7.8	8.3	8.5	9.4	8.0	8.4	8.4
Between-group	1.7	1.6	1.2	1.7	2.5	0.9	0.8	1.0	0.8	0.7	0.4	0.5	0.9	0.6	0.6	0.7
Within-group	16.9	22.7	20.9	21.3	21.8	16.4	16.5	12.0	6.9	7.2	7.9	8.0	8.5	7.3	7.8	7.7
Workers	7.0	7.1	9.4	9.0	11.0	9.2	9.6	10.3	6.1	6.6	7.3	7.2	7.8	6.9	7.8	8.2
Farmers	73.8	82.8	71.3	70.8	61.9	54.3	78.8	27.1	9.1	8.7	10.3	10.8	10.6	9.5	9.3	7.6
Mixed, worker-farmers	22.8	37.5	27.3	20.7	17.1	15.5	10.8	15.9	7.3	7.7	7.6	8.3	8.8	7.2	6.8	6.7
Pensioners	9.5	9.6	10.9	9.4	9.6	8.8	8.9	8.6	8.1	7.5	7.7	7.7	9.0	6.6	7.3	6.4

Notes: Socio-economic groups are defined on the basis of the household's primary source of income.

Table 13. Poverty Rates

Year	Income		Income excluding transfers; lump-sum redistribution of transfers		Nondurables consumption	
	≤ ½ median	≤ 2/3 median	≤ ½ median	≤ 2/3 median	≤ ½ median	≤ 2/3 median
1985	0.04	0.13	0.11	0.19	0.01	0.08
1986	0.05	0.13	0.11	0.19	0.02	0.09
1987	0.05	0.14	0.11	0.20	0.02	0.10
1988	0.04	0.10	0.10	0.18	0.02	0.09
1989	0.03	0.09	0.10	0.17	0.02	0.09
1990	0.09	0.25	0.16	0.30	0.06	0.22
1991	0.08	0.23	0.15	0.28	0.05	0.21
1992	0.09	0.26	0.19	0.32	0.06	0.22

Notes: Quarterly household income and consumption are adjusted by food-share based equivalence scales and deflated by the aggregate CPI. Each individual in a given household is then assigned the same level of income or consumption. The poverty lines based on median real income and real consumption are computed using data across all years. Each column indicates the fraction of the sample population below 1/2 or 2/3 of median real income or consumption, respectively. Median annual real income and real nondurables consumption at 1992:Q4 prices are, respectively, 3672 and 3024 in new zloty (10,000 old zloty = 1 new zloty). Using the OECD PPP exchange rate for 1992 (0.677 new zloty = US$1), this yields income poverty lines expressed in U.S. dollars of 2712 (1/2 median) and 3616 (2/3 median) per equivalent unit. The corresponding poverty lines based on consumption are 2233 and 2978. Poverty lines for different families can be constructed using the equivalence scales in the last column of Table 4. The poverty lines are the same for the first and second panels.

Table 14. Comparisons with per Capita Ginis Based on CSO Methodology

	1985	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996
HBS micro data-- full distribution	0.274	0.291	0.292	0.294	0.293	0.281	0.271	0.267	--	--	--	--
CSO-OECD Ginis	--	--	--	--	0.249	0.230	0.260	0.270	0.290	0.300	0.290	0.300

Notes: The first row shows per capita ginis calculated using the HBS micro data. The second row shows gini coefficients calculated by the CSO for the OECD. Documentation from the CSO suggests that for 1989-93 ginis were computed using income decile groups based on per capita income. For 1994-96, it appears that centile income distributions were used to calculate ginis after adjusting household income by OECD equivalence scales. These gini coefficients for 1989-95 were obtained from the OECD and, for 1996, directly from the CSO.

Figure 1. Actual and Predicted Food Shares

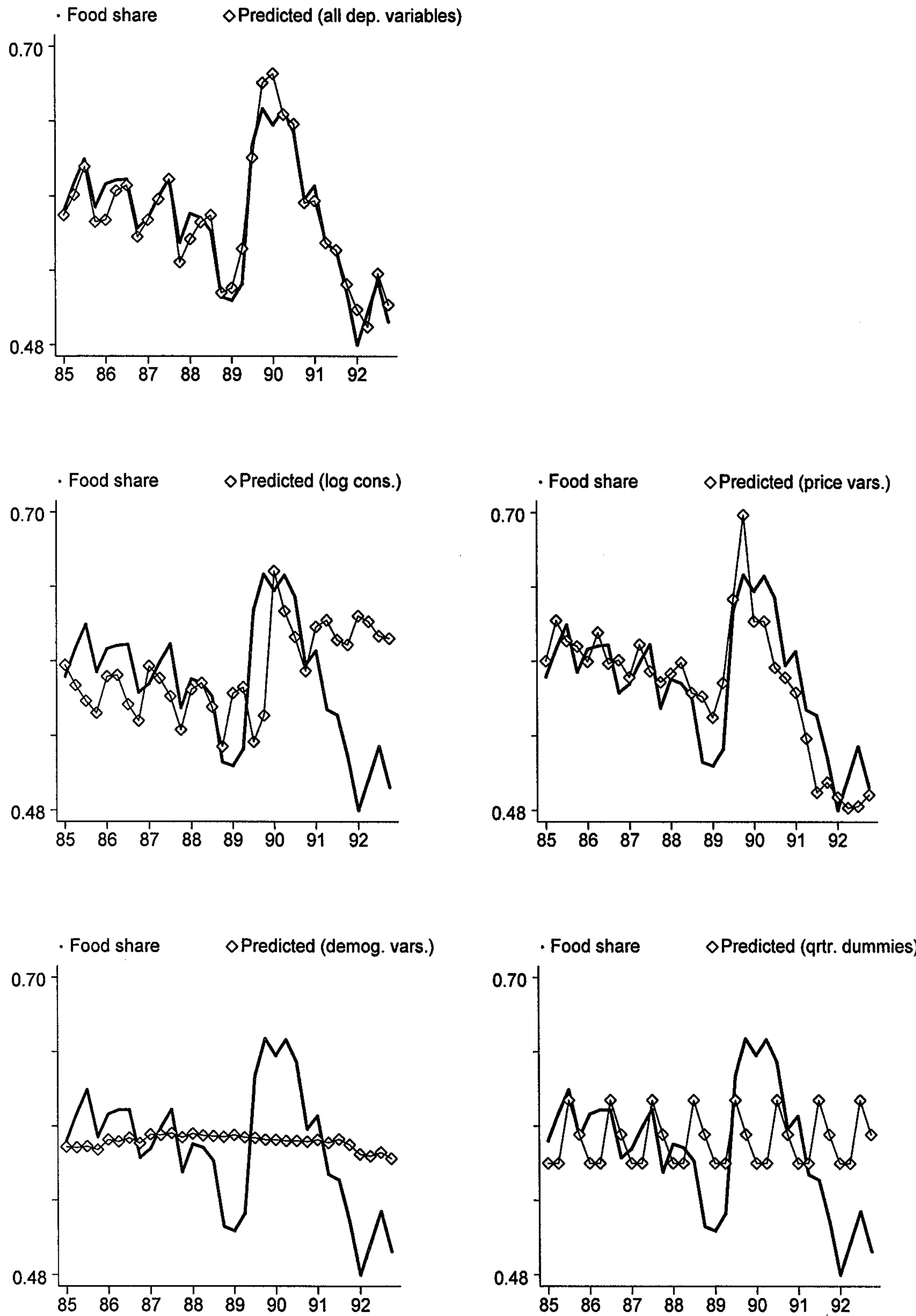
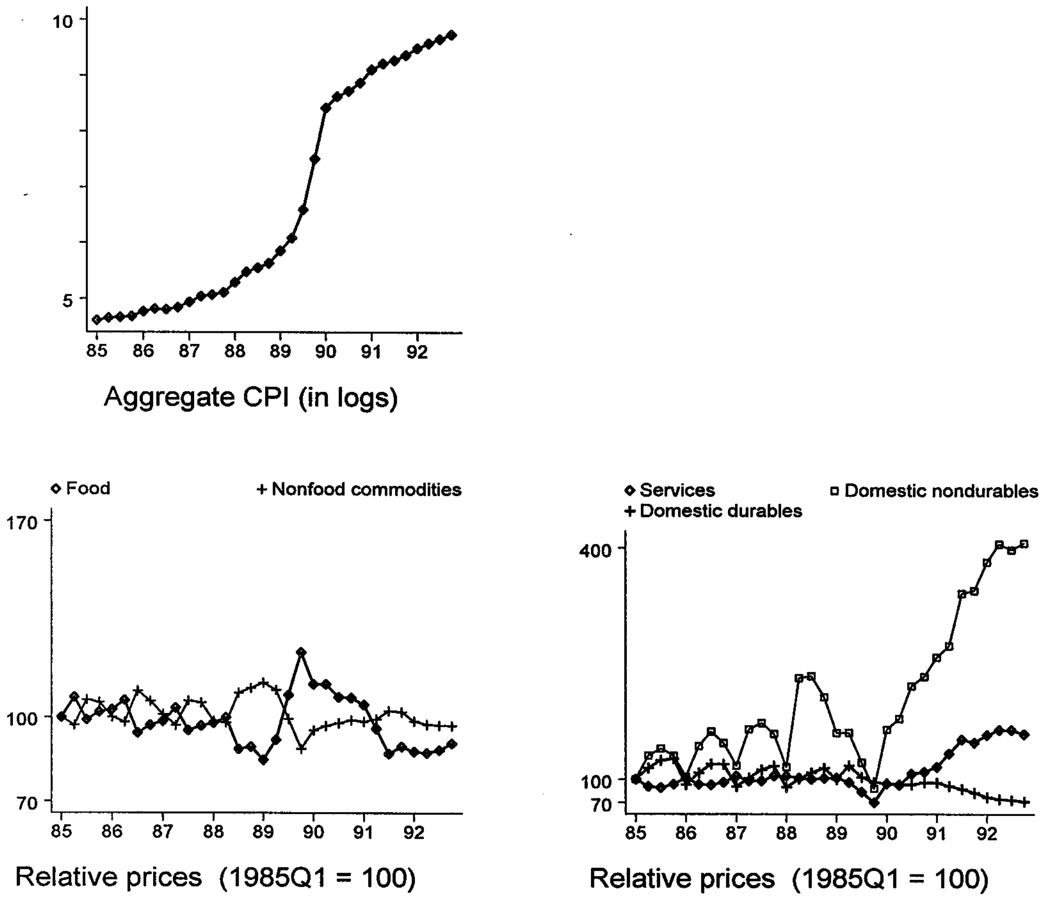
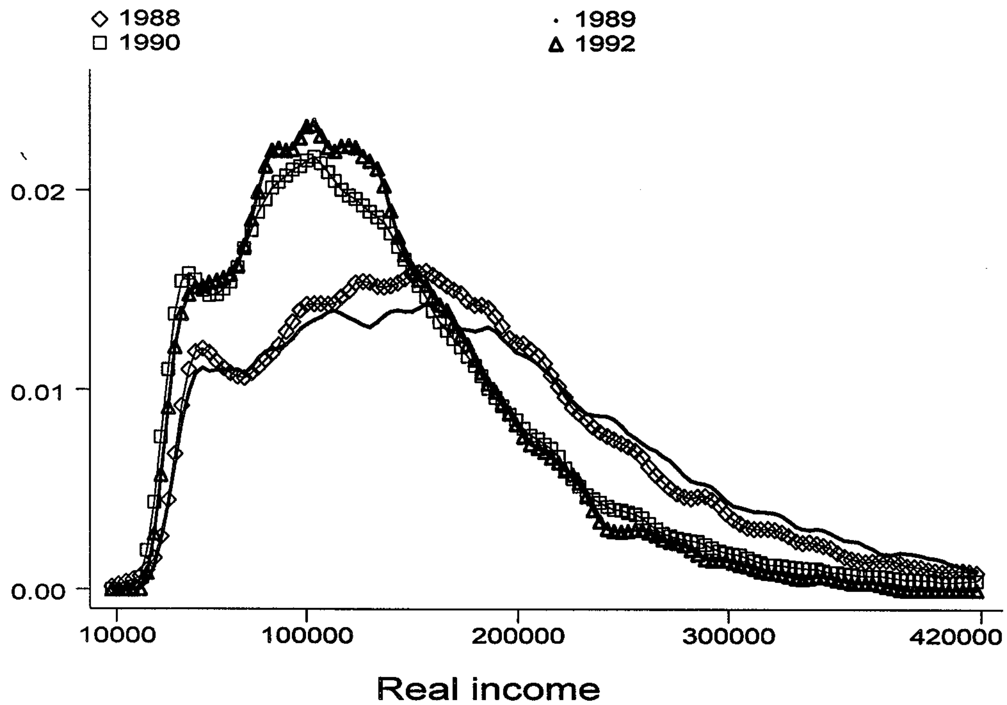


Figure 2. Aggregate CPI and Relative Prices, 1985-92



Note. Lower panels show price indexes relative to aggregate CPI

Figure 3. Kernel Density Estimates



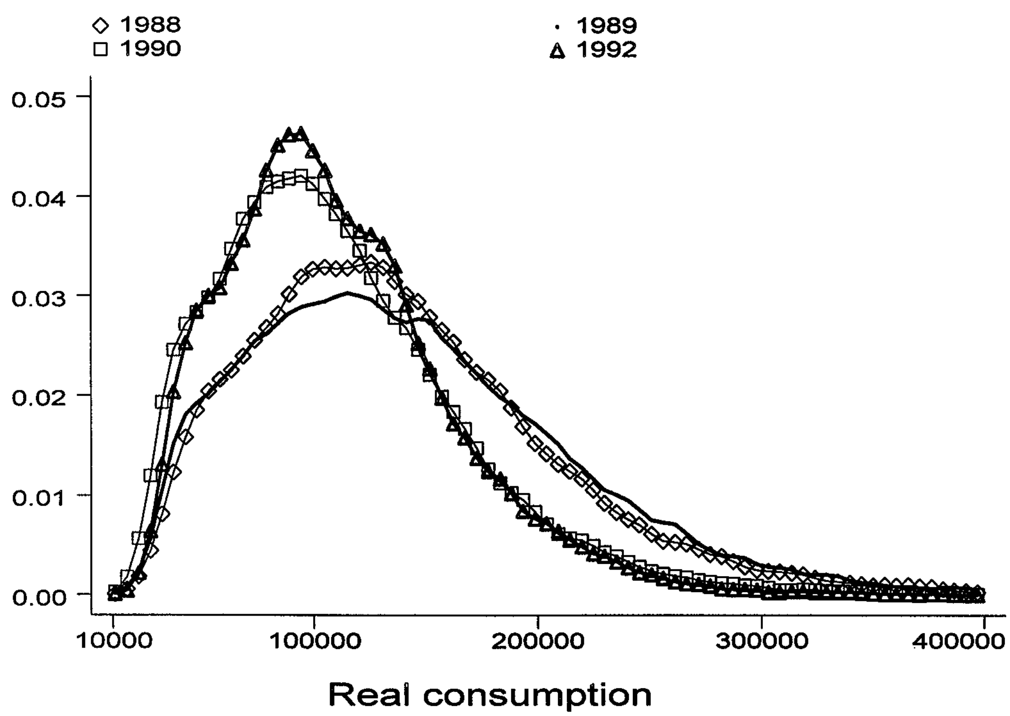


Figure 4. Kernel Density Estimates

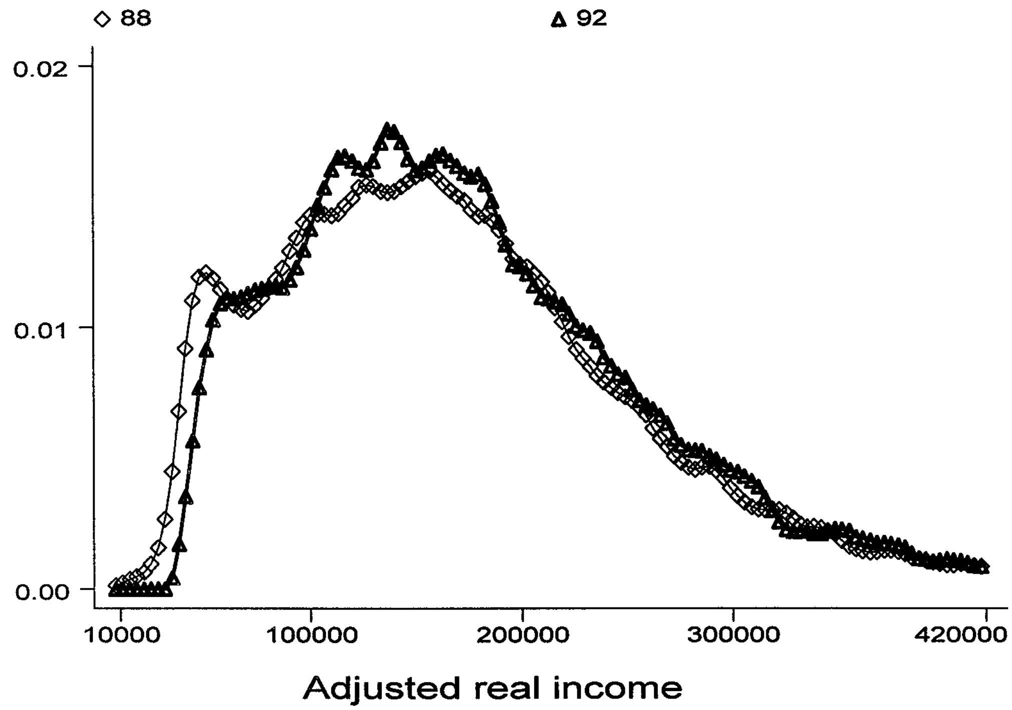
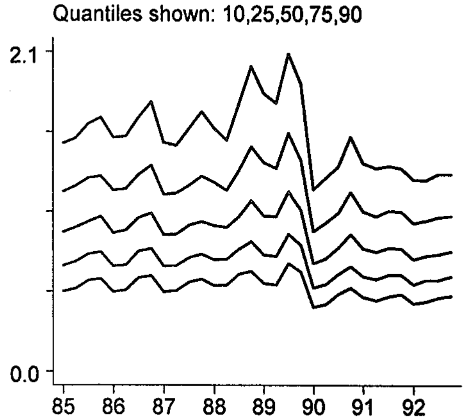
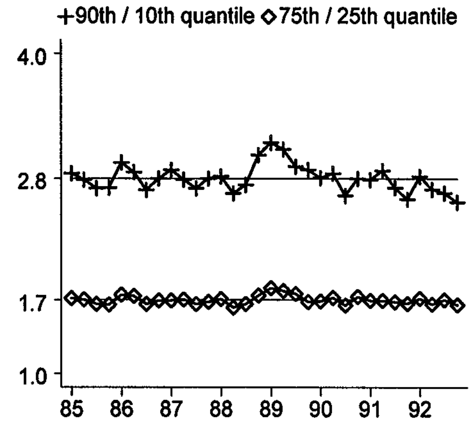


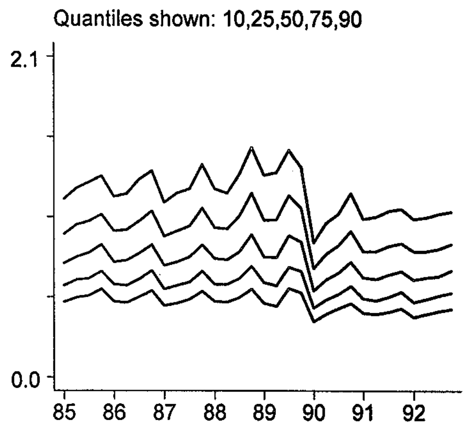
Figure 5. Real Income and Consumption, 1985-92
(millions of zloty, 1992Q4 prices)



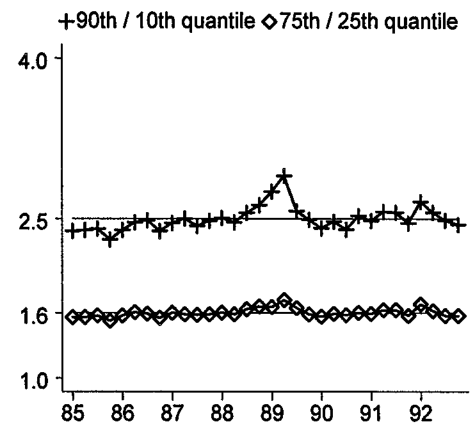
Real Income Quantiles



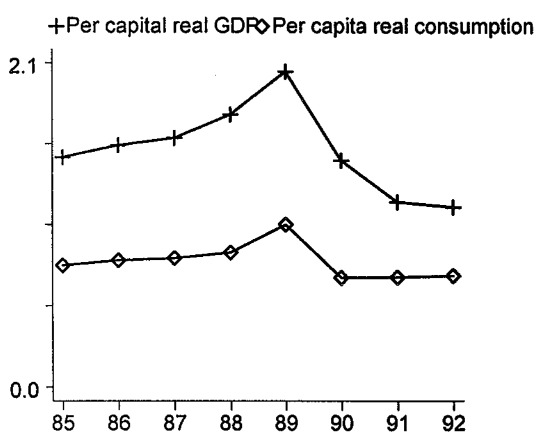
Ratios of Upper to Lower Quantiles: Income



Real Consumption Quantiles



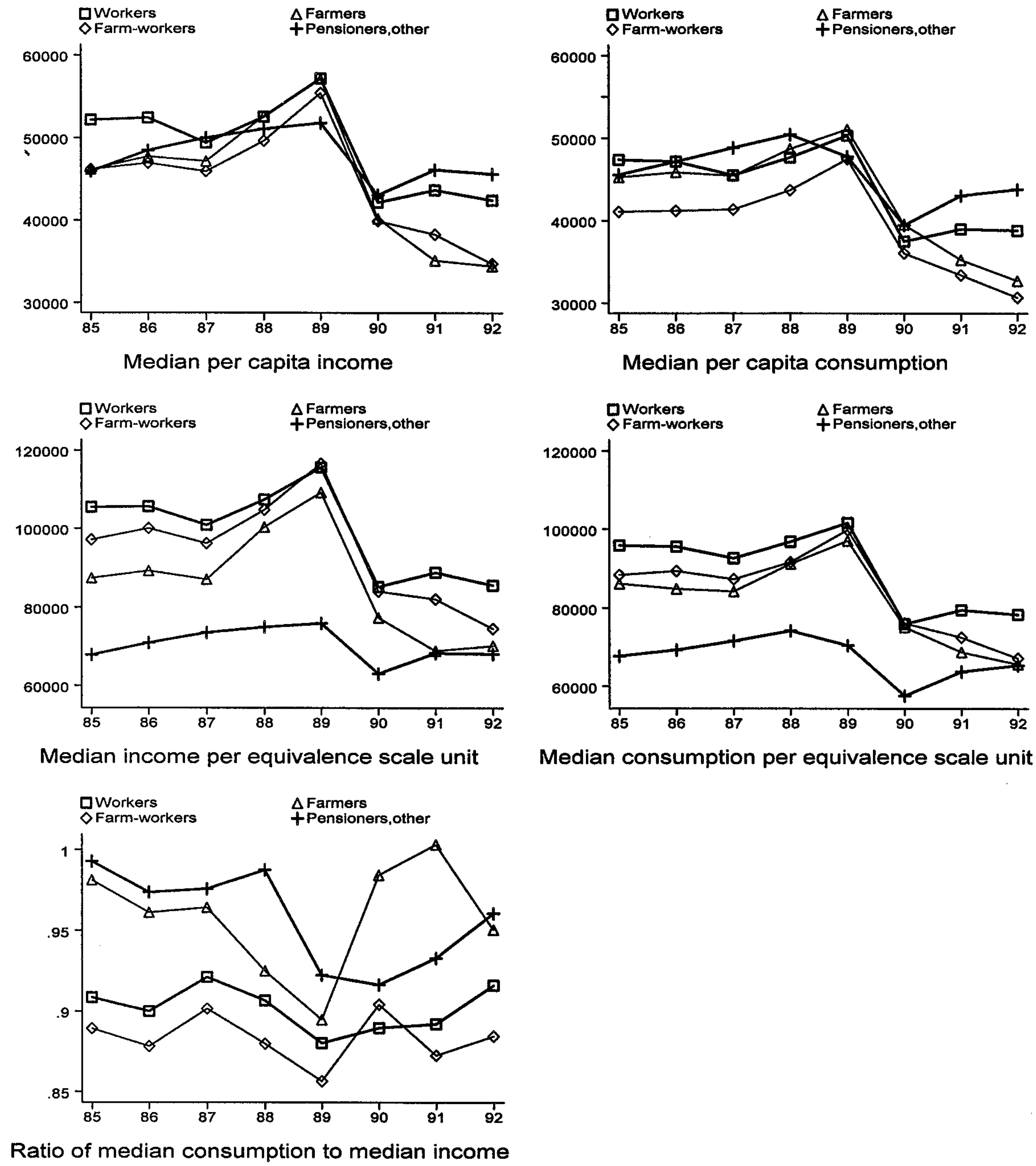
Ratios of Upper to Lower Quantiles: Consumption



Aggregate Output and Consumption

Note. Hhold income, csmn. Adjusted by eqv scales

Figure 6. Median Income, Consumption for Different Groups



References

Atkinson, Anthony B., 1970, "On the Measurement of Inequality," *Journal of Economic Theory*, Vol. 2, pp. 244–63.

Berg, Andrew, and Olivier Blanchard, 1994, "Stabilization and Transition: Poland 1990–91," in *The Transition in Eastern Europe Volume 1: Country Studies*, Olivier Blanchard, Kenneth Froot and Jeffrey Sachs (eds.) (Chicago: The University of Chicago Press).

Buchinsky, Moshe, 1994, "Changes in the U.S. Wage Structure 1963–1987: Application of Quantile Regression," *Econometrica*, Vol. 62:2, pp. 405–58.

Calvo, Guillermo, and Fabrizio Coricelli, 1992, "Stabilizing a Previously Centrally Planned Economy: Poland 1990," *Economic Policy*, Vol. 14.

Commander, Simon, and Fabrizio Coricelli, 1992, "Output Decline in Hungary and Poland in 1990–91," Working Paper #WPS 1036, Economic Development Institute and the Country Economics Department, The World Bank.

Cornelius, Peter and Beatrice Weder, 1996, Economic Transformation and Income Distribution: Some Evidence from the Baltic Countries," *IMF Staff Papers*, Vol. 43:3, pp. 587–604.

Deaton, Angus, 1981, "Theoretical and Empirical Approaches to Consumer Demand under Rationing," in Angus Deaton (ed.) *Essays in the Theory and Measurement of Consumer Behavior*, Cambridge University Press.

Deaton, Angus, and John Muellbauer, 1980, "An Almost Ideal Demand System," *American Economic Review*, Vol. 70:3, pp. 312–26.

Engel, E., 1895, "Die Lebenskosten belgischer Arbeiter-Familien früher und jetzt," *Bulletin de l'Institut International de Statistique*, Vol. 9, pp. 1–129.

Flanagan, Robert, 1995, "Wage Structures in the Transition of the Czech Economy," *IMF Staff Papers*, Vol. 42:4, pp. 836–54.

Gorecki, Brunon, 1994, "Evidence of a New Shape of Income Distribution in Poland," *Eastern European Economics*, pp. 32–51.

McClements, Leslie, 1977, "Equivalence Scales for Children," *Journal of Public Economics*, Vol. 8, pp. 191–210.

Milanovic, Branko, 1992, "Poverty in Poland, 1978–88," *Review of Income and Wealth*, Vol. 38:3, pp. 329–40.

Milanovic, Branko, 1993, "Social Costs of the Transition to Capitalism: Poland 1990–91," Working Paper #WPS 1165, Policy Research Department, The World Bank.

Milanovic, Branko, 1998, "Income, Inequality and Poverty during the Transition from Planned to Market Economy," manuscript, The World Bank.

OECD, 1997, "OECD Economic Surveys: Poland 1997."

Pollak, Robert, 1971, "Conditional Demand Functions and the Implications of Separable Utility," *Southern Economics Journal*, Vol. 37, pp. 423-33.

Shorrocks, Anthony F., 1984, "Inequality Decomposition by Population Subgroups," *Econometrica*, Vol. 52, pp. 1369–85.

Szulc, Adam, 1994, "Poverty in Poland During the Transition Period: 1990–1992 Evidence," *Statistics in Transition*, Vol. 5, pp. 669–82.

Szulc, Adam, 1995, "Measurement of Poverty: Poland in the 1980s," *Review of Income and Wealth*, Vol. 41:2, pp. 191-205.

Triest, Robert K., 1998, "Has Poverty Gotten Worse?," *Journal of Economic Perspectives*, Vol. 12:1, pp. 97–114.

Willis, Robert J., 1986, "Wage Determinants: A Survey and Reinterpretation of Human Capital Earnings Functions," in *Handbook of Labor Economics*, Vol. 1 (eds.) Orley Ashenfelter and Richard Layard (Amsterdam: North Holland).

World Bank, 1995, "Understanding Poverty in Poland."