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Inequality, Transfers and Growth: New Evidence from the Economic Transition in Poland

Michael P. Keane and Eswar S. Prasad

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Abstract

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This paper challenges the conventional wisdom that inequality in Poland increased markedly during the economic transition. Income and consumption inequality actually declined in 1990-92 and rose only moderately above pre-transition levels by 1997. However, inequality in labor earnings increased markedly and consistently during 1990-97. Social transfer mechanisms, including pensions, helped mitigate increases in overall inequality and poverty. More importantly, these transfer mechanisms were well-designed to reduce political resistance to market-oriented reforms in the early years of transition, paving the way for rapid growth. Cross-country evidence from transition economies is consistent with this interpretation and with recent literature suggesting that inequality-reducing redistribution can enhance growth.

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I. Introduction

Among the most dramatic economic events of the early 1990s was the beginning of the process of transformation of countries in Eastern Europe from planned to market economies. These transition economies have had considerably different experiences in terms of the speed and success of transition and in terms of macroeconomic outcomes including output growth. But a widely held view is that, in all of these economies, the economic upheaval associated with the process of transition has led to substantial increases in inequality (see, e.g., Aghion and Commander, 1999).

In this paper, we challenge this conventional wisdom for one of the more successful transition countries—Poland. Using micro data from the Household Budget Surveys (HBS) conducted by the Polish Central Statistical Office (CSO), we examine the evolution of income and consumption distributions in Poland over the period 1985-1997. Our sample covers the first eight years of the economic transition that began with the so-called "big bang" reform of August 1989 to January 1990.[2] Thus, we are able to trace out the time path of income and consumption inequality for an extended period both leading up to and following the "big bang." Although we highlight changes in aggregate measures of inequality such as Gini coefficients to compare our results with those for other countries, the micro data enable us to provide a more detailed characterization of changes in Polish income and consumption distributions and over a longer period than any previous study of transition economies.

Contrary to conventional wisdom, we find no evidence that income and consumption inequality increased in the early years of the transition. In fact, our preferred estimate of the Gini coefficient for the overall individual income distribution actually declined from 0.256 in 1988 to 0.230 in 1992. It then began a gradual increase, reaching levels comparable to the pre-transition period in 1994-96 and then rising to 0.276 by 1997. To put an increase of 0.020 in the income Gini coefficient in perspective, it is only two-thirds as great as the increase reported for the U.S. in the 1980s by Atkinson, Rainwater, and Smeeding (1995). Viewed another way, it still leaves Poland with a Gini value closer to those of Scandinavian countries (around 0.25) than that of the U.S. (0.408) (see World Bank, 2000).

However, we find that inequality in labor earnings increased steadily and substantially during the transition period of 1989-1997. For instance, we estimate that the Gini measure of inequality for individuals in worker-headed households, based only on the labor earnings of those households, increased steadily from 0.252 in 1988, the last full year

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[2] The communist government ended food price controls as it left power in August 1989. The new Mazowiecki government implemented the Balcerowicz plan in January 1990. This ended price controls on most other products, leading 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 trade bloc, contributed to large declines in real GDP (of 11.6 percent in 1990 and 7.0 percent in 1991, according to IMF estimates).

prior to the transition, to 0.298 in 1997. This increase in the Gini coefficient for labor earnings (0.046) was more than twice that of the Gini for overall income (0.020). Analysis of individual earnings data, also from the HBS, indicates that earnings differentials across education levels increased rapidly during the transition, reflecting sharp increases in education premia. But the premium for labor market experience fell sharply after the transition and the position of older workers deteriorated relative to younger workers, consistent with the notion of rapid obsolescence of skills of older workers in a period of massive industrial restructuring

Furthermore, although we find no evidence of increases in overall inequality, an analysis of the relative positions of different socioeconomic groups indicates that there were indeed winners and losers in the process of transition. We find that social transfers played a key role in between-group income dynamics as well as in mitigating the increase in income inequality during the transition, particularly in the early phase. A marked increase in the generosity of public sector pensions in 1991 led to a substantial exit of older workers from the labor force onto the pension rolls in 1991-92 and improved the relative income position of pensioner-headed households. At the same time, other social transfers were increased from 3% of GDP in 1989 to about 5% by 1992. Together, these changes were sufficient to counteract the increase in earnings inequality. As Dewatripont and Roland (1996) point out, such increases in pensions and other social transfers can be rationalized as necessary to achieve initial political support for the "big bang" reform strategy. From 1993 onward, growth in transfers was halted and overall inequality began to rise gradually.

A substantial proportion of transfers was in fact directed not towards households at the bottom of the income distribution but towards the middle class and, via the increased generosity of pensions, to older workers who were potentially big losers in terms of employment and earnings prospects during the transition. Absolute levels of poverty did in fact increase during the transition and, while social transfers mitigated this increase, they did not entirely prevent it. Thus, although transfers may not have been well targeted from a welfare perspective, our results suggest that, from a political economy perspective, transfers may have been a critical component for ensuring social stability and setting the stage for rapid reforms, including enterprise restructuring, during the early years of the transition.

In the final part of the paper, we also provide cross-country evidence on inequality, social transfers and growth in the transition economies that is consistent with our interpretation of the Polish experience. Across 14 countries for which we can observe Gini values both prior to and several years after the start of the transition (i.e., in 1988-89 and 1995-97), the mean increase in the Gini is 0.095, which is several times larger than that observed in Poland. In fact, Poland had the least growth in inequality in this sample of countries but, at the same time, has experienced the fastest economic growth. Poland had cumulative GDP growth of 10.4% over the first 8 years of transition, compared to an average of –25.3% for our sample of 14 countries. We find that the correlation between growth and changes in inequality in transition economies has been strongly negative. This result holds up even when we control for a number of key factors that may help to explain growth, such as indicators of initial conditions and measures of policy reforms aimed at market liberalization—including establishment of property rights and other legal institutions, degree of price liberalization and privatization, etc.

The relationship between growth and inequality has been the subject of considerable debate in recent years (see the survey by Aghion, Caroli and Garcia-Penalosa, 1999). A traditional view is that higher inequality is associated with higher rates of growth. Kuznets (1955) presented evidence of a U-shaped relationship between inequality and per capita GNP, which he interpreted as evidence that inequality increases in the early stages of development and falls thereafter.[3] But more recent empirical work suggests a negative correlation between inequality and growth (see, e.g., Persson and Tabellini, 1994). Recent work in growth theory has rationalized this finding by showing that redistributive transfers can enhance growth in an environment characterized by significant liquidity constraints.[4] Also, in a political economy model, Alesina and Rodrik (1994) show that income redistribution can enhance growth by reducing political support for taxation of capital. And Perotti (1996) finds empirical support for the view that redistribution can enhance growth by fostering socio-political stability.

In our view, the evidence we provide on transfers and inequality in Poland is relevant to this literature on inequality, redistribution and growth. As noted above, we find that a high level of social (cash) transfers mitigated the increase in inequality in Poland during the transition. In fact, social transfers as a percent of GDP averaged 17.7% during 1990-1997, the highest level in any transition country. The mean level of transfers across the 18 countries for which we have data was 10.8%. The high level of transfers in Poland at least partially explains the fact that Poland had the smallest increase in inequality during the transition. In fact, Gomułka (1998) refers to a "Polish model" of transition "distinguished by an exceptionally large volume of social transfers, especially…pensions" that "…helped to reduce the social cost of reform, but is inhibiting Poland's ability to sustain rapid growth." This theme—that the level of transfers in Poland will hinder future growth—has been sounded by many authors, including OECD (1997). But such predictions have yet to be borne out. In 1998-99, Poland continued to experience more rapid growth than any of the other transition countries in our sample.

Given recent developments in growth theory, it is intriguing to speculate that a high level of transfers may actually have helped rather than hindered economic growth in Poland, especially in the early stages of transition. We conclude by presenting some cross-country

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[3] A standard argument is that inequality fosters growth in environments characterized by liquidity constraints, because only wealthy individuals can bear the sunk costs of starting industrial activities. Evans and Jovanovic (1989) provide some evidence that capital market constraints affect the decision to become an entrepreneur even in the U.S., a country with highly developed capital markets.

[4] For instance, Galor and Zeira (1993) turn on its head the argument that wealth concentration encourages growth when there are liquidity constraints. They present a model with borrowing constraints in which individual productivity is a concave function of human capital and show that redistribution of wealth from the rich to the poor enhances growth because the poor have a higher marginal productivity of investment. Related results have been obtained by Banerjee and Newman (1993), Aghion and Bolton (1996) and Benabou (1996).

evidence that suggests that the relationship between social transfers and growth in transition economies has in fact been strongly positive, which is similar to Perotti's (1996) finding for a different and larger sample of countries.

II. Review of Prior Research on Inequality in Poland

There exist a few other studies that have examined income inequality in Poland during the transition. But they report quite contradictory results. This despite the fact that they all use income data derived from the HBS and look at Gini coefficients for the individual income distribution, assigning to each individual the per capita income for the household in which he/she resides. For instance, based on statistics computed by the CSO, OECD (1997, p. 86) reports that the Gini for Poland was 0.25 in 1989, dropped to 0.23 in 1990 and then rose substantially to 0.26 in 1991 and to 0.29 by 1993. It then remained fairly stable in the 0.29 to 0.30 range through 1996. 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 (1999), using published data on income deciles for years prior to 1993 and the HBS micro data for 1993-5, reports that the Gini fell from 0.260 in 1989 to 0.247 in 1991. Like the OECD, he reports a very large jump in the Gini in 1993 to 0.298. But, in contrast to the OECD, his figures suggest that the Gini continued to rise very substantially after 1993, reaching 0.356 in 1995.[5]

To summarize, all three studies suggest that income inequality declined from 1989 to 1990. The CSO-OECD figures imply a very large increase in income inequality in 1991, while the Milanovic and Gorecki figures do not show this. The CSO-OECD (1997) and Milanovic (1999) figures are consistent, however, in implying that large increases in inequality occurred between 1992 and 1993. But the CSO-OECD figures indicate that inequality then stabilized, while the Milanovic figures imply that it grew substantially again in 1994-95.

What can account for this wide divergence in reported results? A problem with the studies cited above is that they do not all use the actual HBS micro data for the period prior to 1993. Rather, for the period prior to 1993, the Gini values in the studies cited above were approximated using aggregate data on quantiles of the income distribution published by the

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[5] Figures in EBRD (2000) are consistent with the OECD figures in that they imply that the Gini plateaued in the 0.29 to 0.30 range from 1995 onward. World Bank (1999, 2000) reports per capita income Ginis of 0.272 in 1992 and 0.329 in 1996. This stands between the OECD and Milanovic (1999) calculations in terms of the rise in inequality over this period. Torrey, Smeeding and Bailey (1999), using a sample that constitutes about 45% of the full HBS sample now available through the Luxembourg Income Survey (LIS) for selected years, report income Gini coefficients of 0.217 for 1987, 0.248 for 1990 and 0.243 for 1992. The LIS's attempt to use a standardized definition of income across country surveys could account for part of the difference between their results and those of other authors and the CSO.

CSO in the annual publication Budzety Gospodarstw Domowych, which we henceforth refer to as the *Surveys*.[6] The accuracy of these approximations is certainly an issue.

But a more serious problem is that in 1993 the CSO switched from quarterly to monthly data collection. Since income is typically more variable at the monthly than the quarterly frequency, this shift alone would have created a substantial increase in cross-sectional income inequality and in the Gini coefficient. Below we will argue that the switch to monthly income reporting accounts for most of the increase in inequality between 1992 and 1993 reported in both OECD (1997) and Milanovic (1999).

In the Appendix, we develop a technique for adjusting the 1993-1997 income and consumption data for the increased variability that may be attributable solely to the shift from quarterly to monthly reporting. The basic idea of our approach is to assume that income consists of a permanent or predictable component (determined by education, age and other observable characteristics of household members) plus a mean zero idiosyncratic component. We then assume that the variance of the idiosyncratic component would not have jumped abruptly between the fourth quarter of 1992 and the first month of 1993. Rather, we assume that the variance of the idiosyncratic component varies smoothly over time (measured in months) according to a polynomial time trend. We estimate this polynomial trend, along with a dummy for post-1992 that captures the discrete jump in variance that occurred with the change to monthly income reporting. Then, at the individual level, we scale down the idiosyncratic component of the post-1992 income data to eliminate this jump in variance.

Another potential problem with previous studies is that the aggregate income statistics reported by the CSO, as well as those reported by other former communist countries, differ in a number of important ways from economically meaningful measures of income. The official statistics appear to reflect total revenues or "inflows" since they include loans, dissaving, and cash holdings at the beginning of the survey period. For farmers, income includes gross, rather than net, farm revenues. This is an important issue as approximately one-fifth of Polish households are either farm households or mixed worker-farmer households. Access to the micro data enables us to make important adjustments in order to obtain a more meaningful measure of income (by excluding non-income revenue items and by calculating net farm income).[7, 8]

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[6] The *Surveys* report the number of households in each of several per capita income ranges, along with the average per capita income within each range, and the average number of persons per household within each range. The number of income ranges reported differs by year. This difference in reporting may itself account for some change in the Gini over time. Also, as described below, the income definition used by the CSO included some inappropriate items, and in some years the CSO made adjustments for family size, but this was not done consistently over time.

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

Both our procedure for adjusting for the spurious increase in inequality stemming for the switch to a monthly reporting interval, and our corrections for the definitions of income and consumption, rely on access to the HBS micro data. In particular, the variance correction requires access to the data for an extended period of time. Our study is unique in that it is based on the HBS micro data for a long sample period, extending from 5 years prior to the "big bang" to 8 years after. To our knowledge, no prior study of inequality in Poland has adjusted for the change in survey design in 1993, and most have not adjusted for the definitional problems noted above.[9]

Of the several improvements we make over previous studies (use of micro data for the pre-1993 period, correction of the income definition, and adjustment for the switch to a monthly sampling frame in 1993), it is our adjustment to the change in sampling frame in 1993 that has the greatest effect. We will argue that failure to account for this change caused prior studies to greatly overstate the increase in inequality in Poland. In fact, this adjustment is central to our finding that Poland has had the least increase of inequality of any transition country.

III. THE HOUSEHOLD BUDGET SURVEYS

The CSO has been collecting detailed micro data on household income and consumption at least since 1978, using fairly sophisticated sampling techniques. In the 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-

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[8] In a similar vein, the aggregate consumption figures published by the Polish CSO, as well as by other former communist countries, often do not correspond to Western-style measures of consumption. Rather, they correspond to a measure of total outflows, including saving and repayment of loans. For farm households, consumption includes farm investment and purchases of supplies. An indication of the strange nature of the aggregate consumption data is provided by Milanovic (1998, p. 41), who reports that in 1993 the Gini for consumption is 0.31, which substantially exceeds the Gini of 0.28 for income. He also reports that in 1993 the ratio of consumption to income is 1.30, an unreasonably high figure. Our access to the detailed micro data enables us to make necessary adjustments to the categories that are included in consumption. We then find the more plausible results that consumption Ginis are smaller than income Ginis and that the aggregate consumption to income ratio falls in the 0.89 to 0.96 range during the 1985–97 period.

[9] 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 the first half of 1993 was released to the World Bank and this data is used in World Bank (1995) and Milanovic (1998). More recently, the data for 1993-96 have been obtained by researchers at the World Bank. A subsample of the HBS is also now available through the Luxembourg Income Survey (LIS) for 1987, 1990 and 1992. Thus, no prior researchers have had access to the micro data for the entirety of the extended period that we examine.

stage units are individual households.[10] Households were surveyed for a full quarter (until 1992) or for a full month (from 1993 onward) in order to monitor their income and spending patterns. Supplementary information on household demographics, durable good holdings, etc. is collected from the same households once every year. The typical sample size is about 25,000 households per year. 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 detailed information on sources and amounts of income for both households and individuals within each household. Total income is broken down into four main categories: labor income (including wages, salaries and nonwage compensation); pensions; social benefits and other transfers; and other income. Social benefits include income from unemployment benefits that were introduced in late 1989. A key point is that the data include measures of the value of in-kind payments from employers to workers, which have been an important part of workers' compensation in Poland and other transition economies. For farm households, farm income and expenditures, as well as consumption of the farm's produce, are also reported. There were no taxes on personal income until 1992. After that year, we use net incomes in the analysis.

In addition to the income data, the HBS also contains very detailed information on consumption. For this study, we aggregate the consumption information and only examine household total consumption and total nondurables consumption. Finally, the HBS also contains information on characteristics of the dwelling, stocks of durables, and demographic characteristics of all household members.

In the immediate aftermath of the big bang, Poland experienced rapid inflation and substantial relative price changes. Using information from various CSO publications and IMF data bases, we have extracted quarterly and, for 1993-97, monthly time series on prices that we use to deflate the income and consumption data. Our ability to match the frequency of the price data to the frequency of the survey data on income and consumption is important in the context of the large absolute and relative price changes that occurred during the transition.

Two important changes were made to the HBS survey design in 1993. We have already noted the change to monthly income and consumption reporting. The other major change was an attempt to obtain a more representative sample of the self-employed. This group's size is believed to have increased markedly since the transition began, resulting in its under-representation in the HBS data during the period 1990-92. In the next section, we examine the extent to which under-representation of the self-employed may have led to understatement of the extent of inequality in the early years of the transition.

[10] The sampling scheme was designed to obtain a survey sample that was representative of the underlying population. But the non-response rates differed across demographic groups, necessitating the use of sampling weights in order to achieve representativeness. We used sampling weights in our analysis where appropriate but these made little difference to any of the main results.

Table 1 reports sample means for some of the variables used extensively in our analysis of inequality.[11] Two interesting features are that the average share of income from transfers and the share of pensioner-headed households increase markedly after the transition. We discuss this in greater detail below. The demographic characteristics of households and household heads remain quite stable during and after the transition. The means of the education dummies indicate a small increase in average levels of educational attainment of household heads in the 1990s (a similar increase occurs in the general population as well).

IV. Inequality

In this section, we examine various aspects of inequality in Poland over the period 1985-1997. For the years 1993-1997, we use the income and consumption measures that are adjusted (using the procedure described in the Appendix) for the increase in idiosyncratic variance that occurred with the shift to a monthly reporting period.

The measures of inequality we examine are based on the distribution of individual income or consumption, unless explicitly noted otherwise. A key problem in inequality measurement is how to account for household composition and household economies of scale when measuring household well being, or when assigning individual income or consumption levels to household members. Most prior studies of income inequality in Poland and other transition economies have ignored these issues and simply assigned the per capita household income to each member of a household prior to measuring inequality in individual income.[12]

In an earlier paper (Keane and Prasad, 1999), we constructed food share (FS) based equivalence scales for Poland using the Engel (1895) method, which assumes that two households with different demographic composition are equally well off at income levels that enable them to have equal food shares (ratio of expenditure on food to total expenditure on nondurables). The equivalence scales we estimated exhibited somewhat greater household economies of scale than the scales typically used for western countries. Below we report our key results based on a number of alternative equivalence scales in order to ensure that our results are not sensitive to the choice of scale. Besides our own FS scale, we also use the OECD scale, the McClements (1977) scale (which is commonly used in Britain), and the simple per capita scale. Appendix Table B1 shows values of the alternative equivalence scales for a representative set of household types.

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[11] Note that the sample size falls in 1992. In that year, half of the total sample was used to test the new monthly survey; these data were considered unreliable and not made available to us.

[12] To the extent that there are household economies of scale, using per capita household income will exaggerate the well being of people in smaller households. And, to the extent that adults have greater expenses than children, use of per capita income will understate the well being of people in households with many children.

A. Measures of Overall Inequality

We first examine the evolution of summary measures of overall inequality. In all cases, we examine the distribution of individual income (or consumption), assigning to each individual the per equivalent (or per capita) income for the household in which the person resides. Table 2 reports Gini coefficients based on per capita incomes and incomes adjusted by the FS equivalence scale. The results in this table highlight the importance of adjusting for the change in survey frequency in 1993. Without this adjustment, for instance, the increase in the per capita income Gini from 1992 to 1993 is 0.045, which is far larger than the estimated increase of 0.021 we obtain using the adjusted data. Similarly, without the adjustment, the Ginis based on the FS equivalence scale would markedly overstate the increase in inequality that occurred between 1992 and 1993 (i.e., a Gini increase of 0.046 vs. 0.018). We use adjusted income and consumption measures for 1993-97 in all of the remaining analysis.

We also examine the Ginis with adjusted income but excluding the self-employed in 1993-96. The inclusion of the self-employed makes only a small difference to either set of Ginis and suggests that under-representation of the self-employed in 1990-92 is unlikely to have resulted in a significant downward bias in Gini coefficients for those years.[13]

The appropriate way of treating the self employed so as to maintain maximum comparability of the inequality measures over time is a difficult issue. For purposes of comparing two adjacent years like 1992 and 1993, it is probably best to exclude the self-employed, since their fraction of the population changed little in that short interval. But, for purposes of comparing the degree of inequality in 1997 with years prior to the big bang, it is best to include the self-employed, since the increase in the number of self-employed over that period is quite significant and could be an important source of increased inequality. Henceforth, we focus on results including the self-employed, recognizing that this generates a bit of a spurious jump in inequality in 1992-93 due to the slight change in sample composition.

Table 3 first reports Gini coefficients based on four alternative equivalence scales. Note that the three scales that account for household economies of scale (FS, McClements, OECD) produce very similar Ginis, typically differing only in the third decimal place. The Ginis based on all four scales indicate that inequality increased in 1989 compared to the level in 1985-88, but that inequality returned to pre-transition levels in 1990, and continued to decline in 1991-92. The Gini based on the FS scale shows the sharpest decline in inequality in 1989-92 (from 0.263 to 0.230) and the Gini based on per capita income shows the smallest decline (from 0.278 to 0.264), but Ginis based on all four scales exhibit the same basic pattern.

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[13] Since this group covers household heads engaged in a wide variety of businesses, households in this group do not systematically have higher income levels than the sample averages. In fact, the distribution of income among the self employed is just slightly more unequal than for the general population.

In short, inequality spiked up in the immediate aftermath of the big bang but, by 1992, was no higher than the levels seen before the transition. Starting in 1993, however, inequality begins to rise and, by 1997, is at a level higher than the peak attained in 1989. This pattern is robust to the choice of equivalence scale. It is important to note, however, that the increase in inequality even by 1997 is hardly dramatic. For example, using the FS equivalence scale, the Gini rises from 0.256 in 1988 (the year before the transition) to 0.276 in 1997. This increase of 0.020 is smaller than the increase of 0.03 reported for the U.S. in the 1980s by Atkinson, Rainwater and Smeeding (1995), or the increase from 0.326 to 0.361 reported for the United Kingdom from 1986 to 1991 in World Bank (1999, 2000).

Conventional wisdom suggests that inequality rose much more in Poland than our results suggest. All of our Gini coefficients, regardless of the equivalence scale on which they are based, imply a much smaller increase in inequality than is implied by official CSO-OECD (1997) figures for 1989-96 on which the conventional wisdom about the sharp increase in inequality after the transition appear to be based. Those figures imply that the Gini coefficient for per capita income rose from 0.249 in 1989 to 0.290 in 1993. In the same period, our per capita Ginis are rather flat, rising only from 0.278 to 0.282. For 1996, the OECD reports a Gini value of 0.300 while our value is 0.301. During 1989-1996 (the longest period for which we can compare results), the OECD figures imply an increase of 0.051 while our figures imply an increase of only 0.023. Thus, while the OECD figures imply an increase in inequality in Poland during the transition that is very large by historical standards, our figures imply an increase that is substantially smaller. Furthermore, our results using the FS scale, which we consider more reliable, imply essentially no increase in inequality over the 1989-1996 period (i.e., the Gini changes from 0.263 to 0.265).[14]

We also examined inequality based on income net of transfers (Table 3, row 5).[15] Interestingly, this reveals a very different picture. The Gini coefficient for income excluding transfers increased by 0.066 from 1988 to 1997, more than three times the increase in the Gini for overall income. Thus, it appears that transfers played a crucial role in inequality dynamics after the transition. We investigate this in greater detail below.

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[14] Note that, in the 1990s, our Ginis are closer to those computed by the CSO. In an earlier paper (Keane and Prasad, 1999), we described a detailed attempt we made to reconcile our Gini coefficients for earlier years with the CSO-OECD figures, which are also purportedly based on the HBS data. We did not succeed completely. The differences can, to a large extent, be attributed to (i) the CSO's use of "revenues" rather than incomes in earlier years; (ii) use of grouped data in calculating Ginis (in the 1980s, tabulated decile groups were used, with all individuals in a given decile group being ascribed the mean income level within that decile—in recent years, percentile groups have been used); and (iii) the apparent inconsistent use of equivalence scales over time (this is based on private correspondence with the CSO).

[15] Since transfers tend to be stable over time, the adjustment factors (used to adjust for the change in survey frequency in 1993-96) for income net of transfers were nearly identical to those we computed for income including transfers.

Rows 6-10 of Table 3 report results for consumption inequality (using, as noted earlier, adjusted consumption data for 1993-97). Consumption is a better measure of welfare than income, particularly as measures based on income could overstate inequality since they may reflect idiosyncratic income shocks that could be smoothed by households. As expected, the Gini coefficients for nondurables consumption are lower than those for income. Nevertheless, independent of the choice of equivalent scale, they show a pattern of changes in inequality almost identical to that based on income. Using total consumption reveals a similar picture.

We wished to examine whether our main results were sensitive to the choice of inequality measure. It is well known that the Gini coefficient is particularly sensitive to changes in a distribution near the median (see Atkinson, 1970). The coefficient of variation (and its monotonic transforms, one of which we use here) 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 inequality measures in the bottom 6 rows of Table 3, in order to determine if they tell a consistent story. In fact, they do. When we use either income or nondurables consumption, both these measures of inequality also show an upward spike in 1989, followed by a decline in 1990-92 to below the pre-transition level, and a subsequent steady increase in 1993-97 to a level modestly above that in the pre-transition period.

When we look at income net of transfers, both the coefficient of variation and mean logarithmic deviation show far greater increases in inequality over the transition period than for total income. This pattern is particularly interesting in the case of the CV measure, which is most sensitive to changes at the high end of the distribution. This result stems from the fact that transfers in Poland are focused not only at the low end of the income distribution but extend well into the high end. We give more details on the targeting of transfers below.

To summarize, we find no evidence to support the view, based on official statistics, of a sharp increase in total income inequality following the transition in Poland. Rather, we find that the increase in income inequality was modest compared, for instance, to increases observed in the U.S. and the U.K. in the 1980s and 1990s. Our results also differ markedly in terms of the timing of changes in inequality. The OECD-CSO figures imply that inequality grew tremendously from 1989 to 1993, and that it then stayed rather flat through 1996. Our results indicate that inequality actually fell from 1989-1992. But we find that inequality rose noticeably after 1993 and, especially, in 1996 and 1997. Thus, we find that most of the increase in inequality occurred several years after the "big bang," and long after the OECD-CSO figures imply the increase had already ceased.

This difference in timing has important implications for the interpretation of what occurred during the transition. The OECD-CSO figures for Poland, as well as the comparable figures for all other transition economies (e.g., Milanovic, 1999), are often interpreted as evidence that substantial increases in inequality are an inevitable concomitant of the process of transition to a market economy. Our results, however, indicate that the change in inequality during the first seven years of the transition in Poland was quite modest. Thus, our results suggest that changes in inequality during transition may not be inevitable but, rather, may result from particular policy choices. In later sections of the paper, we discuss in greater detail the role of social transfer policies in inequality dynamics.

Note that our results concerning the evolution of inequality over time were not at all sensitive to the choice of a particular equivalence scale. Hence, we use only the FS scale in all further analysis.[16] To this point, we have focused on the Gini coefficient and other summary measures in order to compare our results with those of other authors and the CSO. We now exploit our access to the micro data to provide a richer characterization of the evolution of inequality in Poland.

B. Quantile Ratios and Shares

In this section, we examine income inequality by looking at quantile ratios and shares. Unlike the scalar inequality measures considered in section IV.A, examination of quantiles allows one to consider changes in inequality at various different points in the distribution. Figure 1 plots the 90-10 and 75-25 quantile ratios for each year over the sample period. The quantiles for individuals were calculated using real household income and nondurable consumption, both adjusted using the FS equivalence scale. The quantile ratios reveal some interesting patterns. After a brief spike in 1989, the 90-10 quantile ratio falls back to its pre-transition level before gradually increasing in the mid-1990s. However, note that the cumulative increase in the 90-10 ratio from the period 1985-88 through 1997 is only about 0.20, hardly a substantial increase. To put this in perspective, Gottschalk and Smeeding (1997) report a much greater increase of 1.04 (from 4.75 to 5.79) in the 90-10 ratio for the U.S. from 1980 to 1990. The 90-10 ratio for consumption follows a pattern very similar to that of the income ratio over the period 1988-97 (although, for reasons that are not clear, it exhibits an upward trend prior to the transition). The 75-25 quantile ratios for income and consumption are essentially unchanged over the sample period, indicating even greater stability in the middle part of these distributions. We also examined finer breakdowns of the 90-10 and 75-25 quantile ratios (e.g., the 90-50 and 50-10 quantiles ratios) and found that inequality was equally distributed above and below the median and that there were no significant changes in patterns of inequality that could be detected using these finer breakdowns of the data.

Table 4, which reports the shares of income and consumption going to each quintile of the respective distributions, provides an alternative perspective. The shares of income, total consumption, and nondurables consumption going to individuals in different quintile ranges have remained remarkably stable over time, except for a slight and transitory improvement in the relative position of the bottom quintiles right after the big bang. The evolution of income net of transfers is, however, dramatically different. The total share going to the bottom two quintiles fell from over 15 percent in 1985-7 to 13.3 percent by 1992 and further to 10.7 percent by 1997. This was mirrored by an increase in the share of the top quintile, from about 41 percent in 1985-87 to over 46 percent by 1997. These results confirm that transfers played an important role in the dampening of potential increases in inequality during the transition, especially at the lower end of the distribution.

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[16] We recomputed many of the later results in the paper using different equivalence scales. Although the levels of inequality were slightly affected by the choice of equivalence scale, as is the case in Table 3, patterns of the evolution of inequality over time were robust to this choice.

C. Kernel Density Estimates of Income and Consumption Distributions

To obtain a visual representation of changes in the shape and features of the entire distribution, we now examine kernel density estimates of household income and consumption distributions. Figure 2 (top panel) presents kernel density estimates for real household income for the years 1988, 1992, 1993 and 1995.[17] The density is calculated at the same 200 income points for all four years, and the first 125 are plotted in the figure. This covers at least 98% of the households in all four years. Figure 4 (lower panel) also 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% of the households.

The change in the shape of the densities between the year 1988 and selected years after the big bang 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 2 is not due simply to a contraction of the mean. To see this, consider taking the distribution for 1991 and multiplying all of the income figures by the ratio of mean income in 1988 to that in 1991. Such a transformation will preserve relative inequality measures, while equating mean income in 1991 with that in 1988. This enables us to directly compare the shapes of the distributions, abstracting from mean differences. The 1988 income density and the transformed densities for 1991 and also for 1995 are plotted together in Figure 3 (the vertical lines indicate the mean).

The most prominent features of Figure 3 are that, in moving from 1988 to 1991, the mass in the left tail is reduced, and the distribution becomes more peaked around the mode. This accounts for the declines in the Gini measures noted in Section IV.A. A key aspect of what happened during the transition becomes apparent if one compares the top panels of Figures 2 and 3. In Figure 2, we see that, as the overall income distribution shifted left, there was a support area at about 34 to 58 thousand zlotys (prices indexed to 100 in 1992Q4) below which household income tended not to fall. Because of the drop in mean real income from 1988 to 1991, the ratio of this support level to mean income increased. In Figure 3, 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% of their income from pensions (80.5% in 1988, 82.2% in 1991). 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% in 1988 and 26.8% in 1991. 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 178969 zloty in 1988 to 131563 zloty

[17] An Epanechnikov kernel with a bandwidth of 4000 was used for the kernel density estimation. No adjustment was made for household size.

in 1991, the mean real pension actually rose from 29811 to 35258. 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.[18] The lower panel of Figure 3, which compares the adjusted distributions for 1991 and 1995, shows that this effect was further accentuated through 1995.

D. Between-Group Changes in Inequality

We have found no evidence of an increase in overall inequality in Poland in the immediate aftermath of 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. We now turn to an analysis of how different groups fared in terms of relative income and consumption.

Figure 4 shows how median income and consumption evolved for four types of households differentiated by main income source of the household head: workers, farmers, mixed worker-farmers 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 looked at per capita income, "...pensions thus contributed strongly to increase inequality." But the per equivalent unit results in the lower panels tell a very different story.[19] They indicate that pensioner-headed households had much lower median income and consumption than other groups during the 1985-89 period, and that their relative position improved dramatically after the big bang so as to bring their income and consumption up to almost the same level as the next lowest group (farmers). As a result, we find that pensions contributed importantly to a reduction in inequality.[20] The main impetus behind the improved relative position of pensioners was a substantial increase in pension levels that took place in 1991. In fact, by 1997, the relative position of pensioner-headed households is inferior only to that of worker-headed households.

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[18] It is also worth noting that the fraction of households headed by pensioners (and other social benefit recipients) increased from about 28% in the 1985-89 period to 36% in 1992. Opting for the more generous pensions was apparently an attractive option for workers who did not fare well in the transition. We return to this issue later.

[19] 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.

[20] In a result that echoes ours, Garner and Terrell (1998) find that pensions substantially reduced inequality (as measured by income Gini coefficients) during the early transition years in the Czech and Slovak republics.

We also examined the fractions of households that fall in each quintile of the income distribution, conditional on education or age of the household head (results not shown here). One main finding was the substantial improvement in the relative positions of households whose heads have higher educational qualifications. For example, in 1989, 45.8% of households in which the head had a college degree were in the top quintile. This fraction rose to 58% by 1992 and further to 60.2% by 1997. By contrast, in 1989, among households in which the head had only a primary school education, 14.9% were in the top quintile, but this had fallen to 9.5% by 1992 and to 8% by 1997. Another striking result was the improvement of conditions for the old, which resulted from more generous pensions. Among households in which the head was over 60 years old, 39.2% were in the bottom quintile in 1989, but this dropped to only 24.3% by 1992. In contrast, the probabilities that a household with a young (18-30) or middle aged (31-60) head would fall in the bottom quintile of the income distribution increased over the same period.

E. Within-Group Changes in 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 the top panel of Table 5, we report decompositions of the former two inequality measures for income, grouping individuals by the 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. Both measures indicate that most of the increase in inequality during the transition was within group.

An interesting finding, which is apparent in the second panel of Table 5, is that the changes in within-group inequality were very different across different groups. For instance, Gini coefficients estimated separately for each group indicate a steady rise in inequality for individuals in worker-headed households, from 0.189 in 1988 to 0.248 in 1997. This increase of 0.059 in the Gini for individuals in worker-headed households is almost three times as great as the 0.020 increase in the Gini for the overall income distribution. The Gini coefficients in Table 5 indicate that within-group inequality actually fell among farmer and mixed worker-farmer households during the transition. There was also a modest increase in inequality within pensioner-headed households.

The most striking result here is the significant and steady increase in inequality among worker-headed households after 1988. The bottom two rows of Table 5 reveal that much of the increase in income inequality among worker-headed households can be attributed to increased inequality in labor income. When we look at labor income alone, the Gini increased from .252 in 1988 to .298 in 1997, an increase of .046. Thus, we see that inequality in labor earnings grew substantially more than inequality in the overall income distribution.

It is interesting to examine how overall income levels of worker-headed households were influenced by the human capital attributes of the household head. We ran quantile regressions of log real household income on characteristics of the household head (and an urban dummy). We do not report the results in detail here but only briefly summarize the main findings. Log incomes for households with heads in all education groups drop substantially at all quantile points from 1989 to 1990; these declines are greater at the upper quantile points, implying a slight reduction in within-education group inequality. However, by 1992, there is a clear divergence across groups. Households with a college-educated head experience a recovery in income; those with a high school-educated head have stable real incomes at most quantile points; and households headed by persons with lower educational qualifications experience a continuing decline in income. This divergence across groups is accentuated during 1994-96, confirming the earlier results that indicated rising inequality among worker-headed households after the transition.

F. Earnings Inequality

In order to gain more insight into the sources of changes in labor earnings inequality, we also examined the evolution of earnings for individual workers. These data are available in the HBS for all years except 1993 and 1997. We analyzed changes in the wage structure using OLS and quantile regression techniques. To conserve space, we do not present those results here but only briefly summarize the main findings that are relevant to this paper.

The most prominent result in the wage regressions was the sharp increase in education premia after the transition. Estimates of standard human capital earnings functions (see, e.g., Willis, 1986) indicated that the earnings premium for a college degree relative to a primary school degree increased from 47% in 1987 to 102% in 1996. The high school premium increased from 23% to 45% over the same period. These figures are also reflected in our earlier comments about the greater representation of households with better-educated heads in the upper quantiles of the income distribution as the transition progressed. Our finding of a sharp increase in education premia after the transition is consistent with that of Gorecki (1994), based on his examination of aggregate Polish wage data, and of authors who have examined the wage structure in other transition economies. For instance, Ham, Svejnar, and Terrell (1995) examine surveys conducted by the Federal Ministry of Labor in Czechoslovakia in 1988 and 1991. They find that the wage gap between university and elementary school graduates increased from 58% in 1988 to 63% in 1991. Based on her analysis of Russian data, Brainerd (1998) reports that, from 1991 to 1994, the marginal return to a year of education rose from 3.1% to 6.7% for men and from 5.4% to 9.6% for women.

The other main result in our wage regressions was that experience premia are estimated to have declined sharply in the early years of the transition. These declines were quite large at all quantile points of the distribution that we examined and were especially sharp for older workers. There was a slight recovery in experience premia in 1994-96; this recovery was greater for older workers while, for younger and middle-aged workers, experience premia remain below their pre-transition levels even by 1996.

Our results indicate that the returns to general human capital, reflected in education premia, rose markedly after the transition while the returns to experience, especially for older workers, declined sharply in the early years of the transition. This is consistent with the notion of rapid obsolescence of firm- or industry-specific skills during a period of rapid technological change and industrial restructuring (see Svejnar, 1996). Workers with higher levels of general human capital are better able to adapt to such changes, while older workers, who typically have higher levels of firm- or industry-specific human capital, face a sharp decline in their earnings potential. This, combined with the increased generosity of pensions, explains the surge in the number of pensioner-headed households in 1991-1992 that we noted in Table 1. Indeed, self-selection into retirement probably accounts for the recovery in experience premia for older workers that occurred after 1992, since a large number of older workers, particularly in the 55-65 age bracket, retired in 1991-92. The patterns of changes in earnings inequality that we have discussed here have important implications for understanding key aspects of the political economy of the transition process. This is the subject of the next section.

V. The Targeting of Transfers: A Political Economy Perspective

The analysis thus far has indicated that, while inequality in labor earnings did increase substantially among workers and worker-headed households, the overall rise in income inequality during the transition was quite effectively dampened by social transfer mechanisms. In this section, we provide a more detailed examination of the targeting of transfers.

First, we examine the extent to which transfers alleviated poverty. The poverty line is, of course, 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 "overly generous" (see OECD, 1997, p.91; Milanovic, 1998, p. 66).[21] Instead, we calculate poverty lines by first constructing the median of per equivalent household income using pooled data for the entire 1985-97 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. The first panel of Table 6 shows 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 jumped from about 2-3% in 1988-89 to 6% in 1990-92. This fraction rose further in 1993, peaked at 10% in 1994, and then declined moderately by 1997. A key finding is that, while poverty jumped in the immediate aftermath of the big bang, it did not increase much in the subsequent two years. Poverty rates based on the two-thirds median income threshold are higher but have the same time profile.

To analyze the targeting of transfers, we first conducted the simple 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

[21] Using these poverty lines, Szulc (1994, 1995) calculates that the percentage of households in poverty rose from 16.7% in 1989 to 34.2% in 1990, and further to 40.3% 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.

the extent to which transfers alleviate poverty in a purely accounting sense. In 1992, the fraction of people below the two-thirds median threshold drops from 27% to 20% (columns 4 and 2) as a result of transfers, while the fraction below the one-half median threshold drops from 16% to 6% (columns 3 and 1). 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 of the 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 have been substantially improved.

In short, while transfers did mitigate the increase in poverty during the transition, they could clearly have been better targeted if the goal was to prevent an increase in poverty.

We also examined poverty rates based on household (per equivalent) nondurable consumption, again using 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. Furthermore, the evolution of consumption-based poverty is much the same as for the income-based measures.[22]

A complementary approach to analyze the targeting of transfers is to regress transfers on income net of transfers. Results from nonparametric regressions (for households) for selected years are shown in Figure 5. The key observation from this figure is that there are substantial transfers even to households around and above the median of the distribution (the horizontal line shows median real household income based on the full sample). Clearly, from a welfare perspective, transfers could have been better targeted if the objective was to redistribute income to households near the bottom of the distribution of pre-transfer income. However, since individuals in the middle class tend to have a significantly higher propensity to vote than individuals at lower income levels, transfers targeted in this manner may have been more effective at "buying" the social stability that characterized the transition period, notwithstanding the disruptive effects of the economic transformation (see Roland, 1997, for a related analysis).

Another interesting aspect is the importance of pensions as a transfer mechanism. Pension expenditures and the size of the pension rolls increased enormously in the early years of the transition. As shown in Table 7, public expenditure figures indicate that total public pension expenditure as a percent of GDP rose from 8 percent in 1989-90 to almost 15 percent by 1992. The HBS data indicate a similar pattern, with the share of total income accounted for by pensions rising from 16 percent in 1989 to 25 percent in 1992. This is particularly interesting given the results from our wage regressions that showed a substantial decline in experience premia for older workers. Our view is that older workers who were

[22] For households headed by farmers, the consumption-based poverty rates are significantly lower than the income-based rates. The time profiles of poverty rates were broadly similar across different groups. Szulc (1995) reports changes in poverty rates broken down by type of family, income source etc., but only for the pre-transition period. Okrasa (1999) provides a more detailed analysis of poverty within specific groups during 1993-96.

adversely affected by the transition were cushioned by increasing the generosity of the pensions. Indeed, the replacement rate (average pension as a ratio of average wage) rose from about 52 percent in 1988-89 to 65 percent in 1991 and remained above 60 percent through 1997 (OECD, 1998).

Furthermore, since older workers had the most to lose from the privatization or closure of existing state-owned firms, giving them the option of moving on to the pension rolls may have been a key factor in removing a potential political obstacle to enterprise restructuring and privatization. This option, reflected in a relaxation of the pension eligibility requirements in 1990-91, was indeed exercised by a large number of workers, resulting in an increase in the number of newly granted pensions from about 0.6 million per year in 1988-89 to almost 1.4 million in 1991 (OECD, 1998, p. 65). Consistent with this result, we find that, in the HBS data, among households headed by a person in the 55-65 age range, the share of labor income in total income declined from 24 percent in 1989 to 12 percent by 1994, before recovering somewhat to 16 percent by 1997. In these years, the share of pension income in total income for these households was 64 percent, 74 percent, and 73 percent, respectively.[23]

Thus, transfers may have contributed not only to social stability but also to ensuring the conditions necessary for reforms such as privatization and enterprise restructuring that paved the way for high growth after the transition. As shown in the bottom panel of Table 7, this was accompanied by a substantial increase in the general government budget deficit in the early years of the transition. Although there was an attempt to hold the line on transfers in 1990, starting in 1991, the increased generosity of pensions and other social benefits led to a mushrooming of the deficit. This proved unsustainable and, by 1993, growth in transfer expenditures (as a percent of GDP) had been halted, although pensions and other social benefits were at a higher level than in the pre-transition years. The increase in aggregate inequality after 1993 is yet another indicator of how important the growth in transfers was in dampening the rise in overall inequality in the early years of the transition.

To summarize, the analysis in this paper highlights the role of policy choices, as embodied in transfer and other policies, on the dynamics of inequality during the transition to a market economy. In particular, we have argued that the increase in transfer expenditures (and, consequently, the budget deficit) during the critical early years of the transition may have played an important role in setting the stage for the successful economic transition in Poland.

VI. Inequality, Transfers and Growth: Some Cross-Country Evidence

Our detailed analysis of the Polish transition experience has suggested that, from a political economy perspective, the use of transfer mechanisms to mitigate the potential rise in

[23] Among households with heads in the 45-55 age range and in lower age ranges, there was a small drop from 1989 to 1992 in the share of income from labor income, but this was mostly offset by an increase in other social benefits rather than pensions. Among households with heads aged 65 and older, pensions constitute 85-90 percent of total income, with labor income accounting for barely 2 percent.

inequality during the transition to a market economy may have important implications for the success of the transition process. In this section, we expand our analysis to provide a cross-country perspective on the experiences of the transition economies of Eastern Europe in terms of inequality, social transfers and growth.

A prerequisite for the investigation is that we have available for each country two measures of income inequality: one for a year prior to the start of the transition and a second for a year several years after the start of the transition (so that the data do not simply capture the effects of the initial phase of transition on inequality). It is also important that the pre- and post-transition Gini values for each country be based on similar measures of income, similar sampling time frames, similar data sets, etc., so that the measures are reasonably comparable. Table 8 reports pre- and post-transition Gini values, obtained from 6 different sources that we believe reasonably satisfy these comparability criteria. The sources are Milanovic (1998, 1999), World Bank (1997, 1999, 2000) and OECD (1997).

The Gini coefficients in Table 8 are all for the respective individual income distributions, assigning to each individual the per capita income of the household. We have argued earlier that it would be more reasonable to use equivalence scales to accommodate household economies of scale and the different consumption needs of children versus adults. But only per capita income Ginis are available for most transition economies. Ginis based on labor earnings are available for more countries, but these would not account for the effect of transfers on the distribution of total income, which is our focus.

Some omissions from the table are noteworthy. We require that post-transition Gini values be in the 1995-7 period. As a result, we are unable to obtain post-transition values for the Slovak Republic, Uzbekistan, Turkmenistan and Moldova. Gini values for these countries are constructed by Milanovic (1998) for 1993, but we view this as too soon after the start of the transition for our purposes.

While Gini information for transition countries is scarce, there were 5 cases where we had Ginis for both 1988 and 1989 and two cases (besides Poland) where we had Ginis for both 1995 and 1996. In the former cases we took 1988 (the earlier year) and in the latter cases we took 1996 (the later year). Also note that the post-transition Gini values for Lithuania and Kazakhstan are for consumption rather than income. This probably understates the increase in income inequality in these countries. Since these countries also had poor growth performance, the effect is, if anything, to understate the negative correlation between GDP growth and changes in inequality that we find (see below).

Table 8 reports annualized cumulative GDP growth in the first 8 years of transition. This corresponds to the 1990-97 period for all eastern European countries except Romania, 1991-98 for Romania, and 1992-99 for Russia and the other Former Soviet Union countries. The table also reports the mean level of social (cash) transfers, as a percent of GDP, averaged over the period from the first year of the transition through 1997. Note that Poland and Slovenia are the only countries that surpassed pre-transition levels of GDP after 8 years. These countries also have among the highest average levels of social transfers (17.7% of GDP for Poland, 14.8% for Slovenia).

Finally, Table 8 also reports two variables that could be relevant for explaining the different growth experiences of the transition economies. The first is a summary measure of the EBRD transition indicators for each country, taken from the EBRD's 1995 Transition Report. This is a measure of government policies in terms of the degree of transition towards a market economy framework.[24, 25] The second variable is a measure of the difficulty of the initial conditions facing each country at the start of the transition. This variable, taken from de Melo, Denizer, Gelb and Tenev (1997, henceforth MDGT), is constructed using factor analysis and is based on the degree of industrialization, extent of initial macroeconomic imbalances, geographic orientation of trade and length of time under communism. We report the first common factor from their analysis. A higher score indicates more favorable initial conditions.[26]

Figure 6 plots cumulative GDP growth in the first 8 years of transition against the change in the Gini coefficient. A strong negative relationship is obvious, with those countries that have experienced the best growth performance also having the least increase in income inequality. The simple correlation is -0.86. The bottom panel of the figure also plots the relation between growth and government transfers as a percent of GDP, for all 18 countries

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[24] The EBRD report contains ten measures of the degree of transition to a market economy. Three of the measures relate to enterprises: the degree of large and small scale enterprise privatization, and the degree of enterprise restructuring (including elimination of soft-budget constraints). Three measures relate to markets and trade: the degree of price liberalization, the degree of trade liberalization and access to foreign exchange, and the extent of enforcement actions to prevent abuse of market power. Two measures relate to financial institutions: banking reform and interest rate liberalization, and the establishment of securities markets. And two measures capture the extent and effectiveness of the legal framework for securing property rights and regulating business activity. The measures are on a scale of 1 to 4+, with 1 indicating little progress and 4+ indicating a level comparable to that of western industrialized countries. We averaged the transition indicators within each of the 4 groupings (scoring a 4+ as a 5), and then averaged across those 4 scores, to obtain an overall measure of liberalization.

[25] We conducted an exploratory regression analysis to see what combination of the EBRD indicators would best explain growth. We tried regressions of cumulative GDP growth in the first 8 years of transition on many different summary measures of the ten EBRD indicators. We found that the equal weighting within each of the 4 broad categories, followed by equal weighting of the four broad categories, produced as good an Rsquared as any other procedure.

[26] It is worth noting that this measure is very highly correlated with distance from Western Europe. Thus, by using this variable to explain growth, one risks falling into the vacuous conclusion that the Central Asian countries did poorly because they are Central Asian countries. However, it is interesting that Uzbekistan did much better (relatively) than would be expected given its initial conditions, while Bulgaria and Latvia did much worse. We did not use the second factor from MDGT because it had a negligible partial correlation with GDP growth.

for which we were able to obtain transfer data. The relationship is strongly positive, with a simple correlation of 0.67 (0.61 in the subsample of 14 countries for which we have Gini coefficients). Note that finding a positive correlation between transfers and growth is particularly surprising given the blatant denominator bias driving the correlation in the opposite direction (higher output growth increases the denominator of the transfer to GDP ratio).[27] It is interesting that both of these results that we find here for transition economies have also been reported by authors such as Perotti (1996) for a different but much larger sample of industrial and developing countries.

These results are at least not inconsistent with recent developments in growth theory which imply that redistribution to enhance equality may actually enhance rather than dampen growth. But it is of course possible that some third factor explains both good growth performance and the maintenance of income equality in transition economies, and that the correlation we see between growth and the change in inequality has no causal interpretation. To investigate further, we tried regressions of growth on changes in the Gini coefficients, along with the EBRD measure of extent of transition to a market economy framework and the MDGT measure of initial conditions facing each country—the idea being that the extent of liberalization or the difficulty of initial conditions are plausible omitted factors that could explain both growth and changes in inequality.

These results, reported in Table 9, indicate that, individually, the EBRD transition indicator, the MDGT initial conditions measure, and the changes in the Gini coefficients are all highly significantly related to GDP growth (adjusted Rsquareds range from 0.57 to 0.72). In column 4, we include all three variables. Interestingly, only the EBRD transition indicator and the Gini difference are significant, while the initial condition indicator is not. This suggests that initial conditions did not matter for growth once subsequent policy choices (the transition indicators) are controlled for. The results in column 5 confirm this, since the adjusted Rsquared increases when the initial condition indicator is excluded from the regression. The coefficient estimates imply that market liberalization is positively associated with growth, while increasing inequality is negatively associated with growth.

Of course, a problem with these results is the potential endogeneity of the change in inequality. In an attempt to address this problem, in column 6 we instrument for the Gini difference using the initial condition indicator. This procedure relies on the (admittedly strong) identifying assumption that initial conditions do not directly affect growth once we control for subsequent policy choices (the transition indicator). Granted that, in column 6 we are identifying the effect of inequality on growth through variation in inequality outcomes that can be attributed to initial conditions differences (as opposed to changes in inequality that may have been caused by subsequent growth outcomes). Interestingly, the coefficient on the Gini difference does not change much and remains highly significant. In fact, this coefficient is quite stable whether we include initial conditions in the regression (column 4),

[27] Another problem is that higher transfers do not necessarily imply more redistribution. As noted by Commander and Lee (1998), transfers in Russia have actually become regressive in the transition.

use initial conditions to instrument for the difference (column 6), or exclude initial conditions entirely and run OLS (column 5).

A coherent interpretation of these results would be that greater progress towards a market economy framework enhances growth but that, conditional on the degree of liberalization, policy that maintains a greater degree of equality is more conducive to growth. Initial conditions do not affect growth directly once one controls for policy choices. But initial conditions do seem to matter through their effects on policy. In particular, countries with better initial conditions have not only made more rapid progress towards liberalization, but they have also tended to pursue policies that have resulted in smaller increases in inequality in the process. Only to the extent that better initial conditions have led to more liberalization or more equality have they enhanced growth. Of course, as noted above, this interpretation relies on a strong identifying assumption and other interpretations can not be ruled out.

We are naturally cautious about drawing strong conclusions from 14 data points. But, on the other hand, note that this is not really a "small sample" but rather the entire population of Eastern European countries experiencing the transition process (barring a few for which data are not available). One might feel we could obtain more conclusive results by exploiting the panel aspect of the data. But inspection of the data over time reveals that for each country there is tremendous persistence in growth performance, as well as in the degree of liberalization and the extent of increase in income inequality. That is, those countries that have relatively good growth performance, a relatively high degree of liberalization, and a relatively low increase in inequality tend to remain that way throughout the transition, and vice-versa. Thus, we do not feel that there is much to be gained by looking at these data as a panel. It would be an illusion to think there are many more than 14 independent observations to work with here.

Our overall interpretation of these results is that, consistent with the Polish experience that we have analyzed in detail in this paper, the use of social transfer mechanisms and other policies to buffer the potential increase in income inequality, especially in the critical early phase of transition, appears to be important in generating a successful transition to a market economy. We have argued that the Polish experience points to an interesting example of targeting of transfers that, while not necessarily ideal from the perspective of preventing an increase in poverty, may have been crucial for garnering political support for the drastic market-oriented reforms that facilitated Poland's strong growth performance in the 1990s.

VII. Concluding Remarks

This paper has argued, based on detailed analysis of household income and consumption data for the period 1985-97, that, in Poland, there is little evidence of a substantial increase in overall inequality during the transition that began in 1989-90. This contradicts the conventional wisdom that the process of transition to a market economy is inevitably accompanied by a surge in inequality. However, we did find that earnings inequality among workers increased substantially during the transition. We also documented that social transfer mechanisms played an important role in dampening the increase in overall inequality and in between-group income dynamics. We argued that, although the structure of transfers may not necessarily have been ideal from the perspective of preventing an increase in poverty, transfer mechanisms may have played a critical role in maintaining social stability and in reducing political resistance to the structural reforms that were undertaken in the early years of the transition and that facilitated Poland's subsequent strong growth performance. Finally, we presented cross-country evidence on inequality, transfers and growth in transition economies that, while not conclusive, is consistent with the notion that social transfers and other policies aimed at mitigating increases in inequality, especially in the critical early phase of transition, may be conducive to growth.

Accounting for the Change in Survey Frequency in 1993

A few important changes were introduced to the HBS in 1993. Starting in that year, households were surveyed for only one month rather than for a full quarter. In addition, the sampling scheme was modified to provide better coverage of the self-employed. Further, police, security and military personnel were included in the survey. Other aspects of the survey, such as the two stage sampling scheme and the structure of the survey instrument, were left essentially unchanged.

For the purposes of measuring cross-sectional inequality, the change in survey frequency is the most important change. In this appendix, we develop a technique for adjusting the 1993-1997 income and consumption data for the increased variability that may be attributable to the shift from quarterly to monthly reporting. We begin by assuming the following "statistical" or "forecasting" model for household income:

(1) $$Y_{ht} = \alpha_{1t} + \beta_t X_{ht} + \sigma_t \epsilon_{ht}$$

where Y_{ht} is income of household h in period t, X_{ht} is a vector of household characteristics used to predict household income, and ϵ_{ht} is the unpredictable or idiosyncratic component of household income scaled to have a standard deviation of unity. The time-specific standard deviation that scales the idiosyncratic component of household income is denoted by σ_t. Our objective is to estimate the increase in σ_t for the 1993-97 period that is due solely to the switch to a monthly reporting interval.

We begin by estimating equation (1) separately for each quarter from 1985-1992 and each month from 1993-1997.[28] The variables included in X_{ht} are controls for education level, age and sex of the household head, controls for presence of a spouse and age of the spouse, and controls for household size, urban residence, and primary income source of the household head. While the coefficients on most of the controls were stable over time in these regressions, one interesting aspect was that the education of the household head became more important as the transition progressed. A key feature of the results was that the R^2 values dropped sharply after the shift to monthly reporting in 1993. Presumably, the bulk of this drop is due to greater idiosyncratic variability of income, as well as greater relative importance of measurement error, when income is reported at monthly rather than quarterly frequencies.

Next, we assume that the standard deviations of the residuals from estimation of equation (1) follow the process:

(2) $$\ln \sigma_t = \pi_0 + \pi_1 t + \pi_2 t^2 + \pi_3 t^3 + \pi_4 YB_t + \pi_5 I[t>96] + \eta_t \quad t = 2, 4, \ldots, 95; \quad t = 97, 98, \ldots, 156.$$

[28] These results are not reported in the paper but are available from the authors.

Here *t* is a monthly time index. For the years 1985 through 1992, the data are quarterly, so t is assigned as the midpoint of the interval covered by each quarter (that is, t = 2, 4,…., 95). The variable I[t>96] is an indicator for the 1993-1997 sample period in which the data is monthly. Thus, π_5 captures the structural shift in the error standard deviation attributable to the shift to a monthly data frequency. The time polynomials capture the evolution of the error standard deviation over time due to changes in within-group income inequality, controlling for the group characteristics included in X_{ht}. The term YB_t controls for the effect of changes in mean income on the error standard deviation. Finally, η_t captures purely idiosyncratic period-specific changes in income variability.

Analysis of the residuals from equation (1) indicates that households with different primary income sources (i.e., worker-headed households, farm households, mixed farmer/worker households and pensioner households) have very different error variances, as well as different behavior of the error variances over time. Therefore, we estimate equation (2) separately for each household type, using the time series of residual variances from (1).

Estimates of equation (2) for each of the four household types are reported in Table A1, and plots of actual and predicted values of the dependent variable ln σ_t are presented in Figure A1. Observe that the estimates of π_5 are largest for farm and mixed farmer/worker households (0.369 and 0.201), smaller for workers (0.165) and smallest for pensioners (0.044). This is as we would expect, since the increase in income variability as a result of switching to a monthly survey period should be lowest for pensioners—whose primary income source is very stable from month to month—and greatest for farmers (with workers somewhere in between). The plots of the log residual standard deviations from the estimation of equation (1) clearly reveal the jump in the error standard deviation that occurs for each group in January 1993 (t = 97). It is this jump that the coefficients π_5 on the post-1992 dummy are picking up.

There are a number of other interesting aspects of the plots in Figure A1. For workers, note the cluster of large residual standard deviations in the five periods t = 32, 34, 36, 38, 40, which correspond to the fourth quarter of 1988 through the fourth quarter of 1989. There is then a precipitous drop in the residual variance in the period t = 42, which corresponds to the first quarter of 1990. This drop in the standard deviation is mainly a scale effect, resulting from the sharp fall in real incomes that occurred right after the big bang. The model tracks this drop in the standard deviation well due to the inclusion of the mean income level term.

Also note that the error standard deviation appears to rise substantially in 1996-97 for all groups. But this in itself does not necessarily mean that the within-group income equality began to grow rapidly in 1996-97. Part of the increase is purely a scale effect due to the fact that mean real income levels grew substantially in Poland in 1996-97. In Figure A2 (top panel), we plot the polynomial in time alone, holding YB_t fixed at the full sample mean, and setting I[t>96] to zero. These plots indicate that within-group income inequality did begin to grow substantially in 1996-97, even after controlling for changes in the scale of real income.

Including the time dummy (lower panel) shows the effect of the adjustment on within-group residual variances.

Finally, we adjust the income data for 1993-1997 to account for the increase in the idiosyncratic variance that we estimate occurred solely due to the shift to a monthly reporting period in January 1993. We define adjusted income for the 1993-97 period as:

$$(3) \qquad YA_{ht} = \alpha_{1t} + \beta_t X_{ht} + \{\sigma_t / \exp(\pi_5)\} \hat{\epsilon}_{ht} \qquad t = 97, 98, \ldots, 156.$$

Here $\hat{\epsilon}_{ht}$ is the estimated residual from equation (1) and π_5 is our estimate from equation (2) of the increase in the log of the residual standard deviation due to the switch to monthly income reporting. The scale factors $\exp(\pi_5)$ differ for households with each of the four primary income sources. They are 1.179, 1.446, 1.222 and 1.045 for worker-headed households, farm households, mixed worker-farmer and pensioner-headed households, respectively.

A problem we confront is that a representative sample of the self-employed was not obtained in the pre-1993 surveys. Some respondents in the pre-1993 surveys report that they are self employed. But representation of this group in the pre-1993 data is too small to obtain reliable estimates of the group- and time-specific residual variances. Besides, we could not be sure of the extent to which any change in variance for this group in 1993-97 is due to the shift to monthly reporting vs. increased representativeness of the self-employed sample. Thus, we simply assume the same scale factor for the self-employed as we do for workers (1.179). In any case, it turns out that our results on changes in inequality are not very sensitive to how we treat the self-employed, because they constitute only about 5-6 percent of the sample.

We also adjust the consumption data using the same procedure we use for income. We do not discuss those results in detail but simply note that the scale factors $\exp(\pi_5)$ for consumption are 1.108, 1.149, 1.118 and 1.086 for worker-headed households, farm households, mixed worker-farmer and pensioner-headed households, respectively. As expected, these are lower than the income adjustment factors and are quite similar across different groups. For pensioner households, the jump in consumption variability is greater than the jump in income variability with the switch to monthly reporting. While these households have more stable month-to-month income streams than other households, there is no obvious reason to expect their month-to-month variability in tastes for consumption to be lower. Finally, we also adjust the consumption data for the self-employed using the same scale factor we used for workers.

Table 1. Sample Means for Selected Years

	1988	1989	1990	1991	1992	1993	1995	1997
Real household income (shares)								
Labor income	0.52	0.53	0.51	0.49	0.49	0.50	0.52	0.56
Transfers	0.23	0.22	0.26	0.32	0.34	0.33	0.33	0.32
Farm income	0.18	0.19	0.16	0.12	0.12	0.11	0.11	0.08
Other income	0.06	0.05	0.06	0.06	0.05	0.07	0.05	0.05
Real household consumption (shares)								
Durables	0.13	0.14	0.11	0.10	0.08	0.08	0.08	0.10
Nondurables	0.87	0.86	0.89	0.90	0.92	0.92	0.92	0.90
Food	0.45	0.46	0.53	0.47	0.44	0.43	0.41	0.38
Household characteristics								
Urban	0.51	0.51	0.51	0.52	0.64	0.66	0.65	0.67
Number of persons in household	3.27	3.27	3.24	3.16	3.14	3.15	3.18	3.12
Primary income source of household								
Workers	0.55	0.55	0.53	0.50	0.49	0.44	0.42	0.42
Farmers	0.10	0.10	0.10	0.09	0.09	0.08	0.08	0.06
Mixed, worker-farmers	0.07	0.07	0.07	0.07	0.06	0.06	0.06	0.06
Pensioners, others	0.28	0.28	0.30	0.34	0.36	0.38	0.39	0.40
Self-employed	...	...	...	...	...	0.05	0.06	0.06
Household head characteristics								
Male, 18-30	0.11	0.10	0.10	0.10	0.10	0.10	0.10	0.11
Male, 31-60	0.58	0.59	0.57	0.57	0.57	0.59	0.59	0.58
Male, >60	0.13	0.14	0.14	0.14	0.14	0.13	0.13	0.13
Female, 18-30	0.01	0.01	0.01	0.01	0.01	0.01	0.01	0.01
Female, 31-60	0.09	0.09	0.09	0.09	0.09	0.09	0.09	0.09
Female, >60	0.08	0.08	0.08	0.09	0.09	0.08	0.08	0.08
Age	47.54	47.78	47.90	48.30	48.45	47.96	48.03	48.09
College degree	0.07	0.06	0.06	0.07	0.08	0.09	0.09	0.09
Some college	0.00	0.00	0.00	0.00	0.00	0.01	0.01	0.01
High school	0.20	0.19	0.20	0.21	0.23	0.24	0.24	0.26
Some high school	0.01	0.01	0.02	0.01	0.01	...	...	...
Basic vocational training	0.31	0.33	0.33	0.33	0.33	0.34	0.36	0.35
Primary school	0.34	0.34	0.32	0.32	0.30	0.28	0.26	0.25
Primary school not completed	0.07	0.06	0.05	0.05	0.04	0.04	0.03	0.02

Number of observations (households)

1985	21,560	1989	29,366	1992	10,642	1995	31,874
1986	25,475	1990	29,148	1993	31,966	1996	31,782
1987	29,510	1991	28,632	1994	31,942	1997	31,659
1988	29,287						

Notes: The components of income and consumption are shown as (mean) shares of total income and consumption, respectively.

Table 2. Effects of Changes in Survey in 1993 on Gini Coefficients for Income

	1985	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997
							Per Capita Income Ginis						
Baseline	0.270	0.274	0.270	0.272	0.278	0.271	0.266	0.264	0.285	0.298	0.294	0.301	0.319
Alternative Ginis for 93-97													
Without residual adjustment									0.309	0.323	0.318	0.327	0.339
Excluding self-employed									0.283	0.295	0.291	0.299	0.316
							Food-share Based Equivalence Scale						
Baseline	0.252	0.254	0.246	0.256	0.263	0.250	0.235	0.230	0.248	0.262	0.255	0.265	0.276
Alternative Ginis for 93-97													
Without residual adjustment									0.276	0.292	0.284	0.296	0.304
Excluding self-employed									0.243	0.257	0.250	0.261	0.274

Notes: The baseline Ginis include the self-employed (whose representation in the sample was increased in 1993) and incorporate adjustments for the change in survey frequency (from quarterly to monthly) in 1993. The procedure for adjusting the income data for 1993-97 is described in the Appendix.

Table 3. Poland: Measures of Inequality, 1985-97

	1985	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997
						Gini Coefficients							
Total income													
Food-share based eqv. scale	0.252	0.254	0.246	0.256	0.263	0.250	0.235	0.230	0.248	0.262	0.255	0.265	0.276
McClements equivalence scale	0.249	0.253	0.246	0.254	0.261	0.249	0.238	0.234	0.253	0.266	0.259	0.270	0.282
OECD equivalence scale	0.253	0.257	0.250	0.256	0.264	0.253	0.242	0.238	0.257	0.271	0.264	0.275	0.286
Per capita	0.270	0.274	0.270	0.272	0.278	0.271	0.266	0.264	0.285	0.298	0.294	0.301	0.319
Income excluding transfers	0.373	0.375	0.368	0.385	0.384	0.389	0.404	0.416	0.416	0.437	0.432	0.448	0.451
Nondurables consumption													
Food-share based eqv. scale	0.196	0.200	0.205	0.211	0.219	0.209	0.208	0.205	0.222	0.228	0.222	0.227	0.235
McClements equivalence scale	0.197	0.202	0.208	0.214	0.220	0.210	0.213	0.212	0.229	0.234	0.229	0.233	0.242
OECD equivalence scale	0.200	0.207	0.212	0.217	0.224	0.214	0.218	0.217	0.234	0.239	0.234	0.239	0.247
Per capita	0.222	0.229	0.236	0.239	0.242	0.235	0.245	0.249	0.262	0.268	0.264	0.268	0.277
Total consumption	0.230	0.234	0.239	0.244	0.258	0.241	0.233	0.227	0.247	0.254	0.247	0.262	0.271
						Half the Square of the Coefficient of Variation							
Total income	0.085	0.090	0.085	0.091	0.105	0.086	0.079	0.077	0.097	0.103	0.096	0.105	0.112
Nondurables consumption	0.066	0.068	0.070	0.074	0.081	0.068	0.072	0.068	0.088	0.093	0.085	0.091	0.099
Income excluding transfers	0.184	0.190	0.186	0.203	0.210	0.207	0.230	0.244	0.265	0.281	0.278	0.294	0.306
						Mean Log Deviation							
Total income	0.075	0.079	0.077	0.078	0.087	0.075	0.071	0.069	0.079	0.086	0.081	0.086	0.093
Nondurables consumption	0.060	0.062	0.064	0.067	0.074	0.062	0.064	0.064	0.067	0.064	0.056	0.055	0.082
Income excluding transfers	0.224	0.214	0.213	0.221	0.244	0.247	0.268	0.278	0.404	0.357	0.333	0.317	0.444

Notes: The inequality measures shown here are for the individual distributions of income and consumption. Household income and consumption are adjusted using the food-share based equivalence scale (unless indicated otherwise) and allocated equally to individuals in the household. Income and consumption data for 1993-97 are adjusted for the change in survey frequency.

Table 4. Quantile Shares of Income and Consumption

Quantile range	1985	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997
							Total Income						
≤ 20	9.0	9.1	9.5	9.4	9.3	9.7	10.4	10.5	10.2	9.6	9.9	9.4	9.1
21-40	14.8	14.9	15.0	14.7	14.4	14.9	15.3	15.2	14.6	14.5	14.6	14.3	14.3
41-60	18.5	18.1	17.9	17.5	17.6	18.0	18.4	18.6	17.9	17.9	17.9	17.7	17.8
61-80	23.0	22.4	22.0	21.6	22.0	22.3	22.4	22.8	22.0	22.0	22.2	22.0	22.1
>80	34.6	35.5	35.7	36.8	36.7	35.1	33.4	32.9	35.3	35.9	35.5	36.6	36.8
							Income net of transfers						
≤ 20	2.4	2.2	2.2	2.2	2.7	2.2	1.9	2.0	1.7	1.2	1.3	0.9	0.5
21-40	12.7	12.9	13.3	12.7	12.5	12.5	12.1	11.3	11.2	10.8	10.7	10.0	10.2
41-60	19.0	18.8	18.7	18.0	18.1	18.6	18.9	18.3	17.9	17.8	17.8	17.3	17.7
61-80	25.3	24.4	24.0	23.7	24.0	24.8	25.5	26.0	24.9	25.1	25.2	25.1	25.1
>80	40.6	41.6	41.8	43.4	42.8	41.8	41.6	42.4	44.3	45.2	45.1	46.7	46.5
							Total consumption						
≤ 20	10.7	10.8	10.9	10.7	10.1	10.6	11.0	10.7	11.1	11.0	11.1	10.8	10.6
21-40	14.6	14.7	14.8	14.6	14.2	14.7	15.0	14.9	15.0	14.9	14.9	14.7	14.8
41-60	18.0	18.0	17.8	17.6	17.6	18.0	18.2	18.3	18.3	18.2	18.3	18.1	18.2
61-80	22.4	22.0	21.8	21.8	22.1	22.1	22.3	22.5	22.3	22.3	22.5	22.5	22.5
>80	34.3	34.4	34.7	35.3	35.9	34.7	33.5	33.6	33.3	33.5	33.2	33.9	33.9
							Nondurables consumption						
≤ 20	11.6	11.7	11.9	11.6	11.2	11.5	11.8	11.2	11.1	11.0	11.1	10.8	10.6
21-40	15.4	15.7	15.8	15.5	15.4	15.6	15.8	15.5	15.0	14.9	14.9	14.7	14.8
41-60	18.7	18.8	18.7	18.6	18.6	18.8	18.7	18.7	18.3	18.2	18.3	18.1	18.2
61-80	22.7	22.4	22.1	22.3	22.7	22.5	22.4	22.6	22.3	22.3	22.5	22.5	22.5
>80	31.7	31.3	31.5	32.0	32.2	31.7	31.3	32.0	33.3	33.5	33.2	33.9	33.9

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

Table 5. Decomposition of Inequality Measures for Income

	1985	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997
Half the square of the coefficient of variation (x100)													
Total	8.5	9.0	8.5	9.1	10.5	8.6	7.9	7.7	9.7	10.3	9.6	10.5	11.2
Between-group	1.4	1.2	0.8	1.2	1.8	0.9	0.6	0.6	1.5	1.4	1.1	1.5	1.4
Within-group	7.1	7.8	7.7	8.0	8.7	7.8	7.3	7.0	8.2	8.9	8.5	9.0	9.8
Mean log deviation (x100)													
Total	7.5	7.9	7.7	7.8	8.7	7.5	7.1	6.9	7.9	8.6	8.1	8.6	9.3
Between-group	0.8	0.6	0.4	0.6	1.0	0.5	0.3	0.3	0.8	0.7	0.6	0.8	0.8
Within-group	6.7	7.3	7.3	7.2	7.7	7.0	6.8	6.6	7.2	7.9	7.5	7.8	8.5
Gini coefficients													
All	0.252	0.254	0.246	0.256	0.263	0.250	0.235	0.230	0.248	0.262	0.255	0.265	0.276
Workers	0.186	0.192	0.191	0.189	0.208	0.211	0.208	0.211	0.222	0.234	0.228	0.240	0.248
Farmers	0.475	0.483	0.478	0.496	0.440	0.420	0.366	0.321	0.313	0.362	0.341	0.366	0.414
Mixed, worker-farmers	0.272	0.279	0.276	0.285	0.271	0.253	0.229	0.220	0.223	0.234	0.244	0.252	0.267
Pensioners, other	0.211	0.212	0.203	0.205	0.214	0.206	0.210	0.203	0.225	0.231	0.226	0.228	0.240
Urban	0.201	0.203	0.198	0.202	0.223	0.217	0.213	0.210	0.239	0.247	0.241	0.250	0.257
Rural	0.317	0.307	0.287	0.302	0.296	0.278	0.249	0.249	0.247	0.270	0.261	0.273	0.286
Ginis for worker-headed households													
Labor income	0.237	0.243	0.240	0.252	0.262	0.268	0.278	0.289	0.285	0.292	0.288	0.295	0.298
Income excluding transfers	0.230	0.232	0.230	0.243	0.255	0.257	0.264	0.270	0.271	0.280	0.274	0.287	0.291

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

Table 6. Poverty Rates

	Income		Income excluding transfers; lump-sum redistribution of transfers		Nondurables consumption	
Year	≤ ½ median	≤ 2/3 median	≤ ½ median	≤ 2/3 median	≤ ½ median	≤ 2/3 median
1985	0.03	0.10	0.10	0.17	0.01	0.05
1986	0.03	0.10	0.09	0.16	0.01	0.06
1987	0.04	0.11	0.10	0.17	0.01	0.08
1988	0.03	0.08	0.09	0.15	0.01	0.07
1989	0.02	0.07	0.08	0.14	0.01	0.07
1990	0.06	0.20	0.13	0.25	0.04	0.17
1991	0.05	0.17	0.13	0.24	0.04	0.16
1992	0.06	0.20	0.16	0.27	0.04	0.18
1993	0.08	0.23	0.16	0.29	0.04	0.17
1994	0.10	0.26	0.18	0.31	0.06	0.23
1995	0.09	0.24	0.17	0.30	0.06	0.22
1996	0.08	0.21	0.17	0.28	0.05	0.20
1997	0.07	0.18	0.15	0.25	0.05	0.19

Notes: 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, 3,374 and 2,829 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 2,492 (1/2 median) and 3,323 (2/3 median) per equivalent unit. The corresponding poverty lines based on consumption are 2,089 and 2,786. Poverty lines for different families can be constructed using the equivalence scales in the last column of Table B1. The poverty lines are the same for the first and second panels.

Table 7. Social Transfers

	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997
General Government Expenditures										
(in percent of GDP)										
Cash transfers to individuals	9.4	11.2	10.6	17.3	19.9	20.4	20.2	19.7	18.7	19.4
Pensions	7.1	8.2	8.1	12.2	14.8	15.0	14.9	14.5	14.3	14.4
Unemployment benefits	0.0	0.0	0.2	1.2	1.7	1.2	1.2	1.2	1.1	1.0
Other benefits	2.3	3.0	2.3	3.9	3.4	4.2	4.1	4.0	3.3	4.0
Mean Cash Transfers (HBS data)										
Total transfers	41154	41792	36254	44948	44694	43486	44171	44860	46786	48197
(avg. ratio to total income)	(23.4)	(21.8)	(26.3)	(32.2)	(33.6)	(31.6)	(32.8)	(32.7)	(32.4)	(31.3)
Pensions	29857	30497	27307	33520	33346	33172	34672	36240	38008	40715
(avg. ratio to total income)	(17.0)	(15.9)	(19.8)	(24.0)	(25.1)	(24.1)	(25.8)	(26.4)	(26.3)	(26.4)
Other cash benefits (incl. UI)	11280	11279	8927	11404	11323	10315	9498	8620	8777	7482
(avg. ratio to total income)	(6.4)	(5.9)	(6.5)	(8.2)	(8.5)	(7.5)	(7.1)	(6.3)	(6.1)	(4.9)
General Government Balance										
(in percent of GDP)	0.0	-7.4	3.1	-6.5	-6.7	-2.9	-3.0	-3.1	-3.4	-3.1
Real GDP (annual % change)	4.0	0.3	-11.6	-7.0	2.6	3.8	5.2	7.0	6.1	6.9

Notes: The data on real GDP and government expenditures are taken from various IMF sources. The figures in the middle panel (mean transfers in HBS data) are expressed in terms of 1992Q4 prices.

Table 8. Cross-Country Data

Country	Annualized cumulative GDP growth	Gini coefficients			Average cash transfers (% of GDP)	Transition indicators	Initial condition indicator
		Pre-transition	Post-transition	Difference			
Poland	1.25	0.272	0.301	0.029	17.7	14.3	1.18
Slovenia	0.47	0.174	0.223	0.049	14.8	13.0	1.24
Czech Republic	-0.29	0.194	0.254	0.060	12.1	14.7	1.43
Hungary	-1.15	0.248	0.308	0.060	16.5	14.7	1.47
Romania	-2.18	0.233	0.280	0.047	8.9	9.8	0.94
Estonia	-3.05	0.230	0.354	0.124	10.0	12.5	-0.33
Belarus	-4.23	0.228	0.288	0.060	8.9	8.3	-1.19
Kazakhstan	-5.03	0.257	0.354	0.097	6.9	8.3	-1.07
Bulgaria	-5.03	0.228	0.317	0.089	11.8	10.3	0.55
Lithuania	-5.65	0.225	0.324	0.099	9.6	10.5	-0.52
Russia	-6.14	0.238	0.380	0.142	7.5	9.7	-0.34
Kyrgyzstan	-6.67	0.260	0.405	0.145	12.4	10.3	-1.03
Latvia	-6.89	0.225	0.320	0.095	11.8	10.2	-0.46
Ukraine	-10.64	0.233	0.473	0.240	9.4	8.7	-0.91

Notes: Annualized cumulative GDP growth is measured over the first eight years of transition. The first year of transition is 1990 for Bugaria, the Czech Republic, Hungary, Poland, Romania, and the Slovak Republic; 1991 for Slovenia; and 1992 for the Baltics, Russia and other countries of the former Soviet Union. Data on Gini coefficients were taken from Milanovic (1998, 1999), World Bank (1997, 1999, 2000), OECD (1997) and, for Poland, from this paper. Data on average cash transfers from the transition years through 1997 are from Milanovic (1998). The transition indicator is a weighted average of the transition indicators in the EBRD's 1995 Transition Report (Table 2.1). The index of the difficulty of initial conditions (a higher score indicates more favorable conditions) is taken from de Melo, Denizer, Gelb, and Tenev (1997).

Table 9. Cross-Country Regressions

Dependent variable: Annualized cumulative real GDP growth in first 8 years of transition						
	(1)	(2)	(3)	(4)	(5)	(6)
Change in Gini *100	-0.520*	...	...	-0.365*	-0.381*	-0.442*
	(0.089)			(0.075)	(0.065)	(0.149)
Transition indicator	...	1.121*	...	0.587*	0.671*	0.599*
		(0.261)		(0.243)	(0.154)	(0.225)
Initial conditions index	...	...	2.617*	0.283	...	...
			(0.562)	(0.616)		
Constant	1.020	-16.388*	-4.124*	-6.990*	-7.757*	-6.373
	(0.967)	(2.952)	(0.551)	(2.751)	(2.107)	(3.731)
Adjusted Rsquared	0.72	0.57	0.61	0.83	0.89	0.88
Number of observations	14	14	14	14	14	14

Notes: A positive change in the Gini coefficient reflects an increase in inequality after the transition. The transition indicator captures progress in various dimensions of transition to a market economy. A higher value indicates better progress. The initial conditions index is defined such that a higher value indicates more favorable initial conditions at the start of transition. See Table 8 for a detailed description of the data used in these regressions. Regressions (1) through (5) are OLS specifications, while regression (6) is an IV specification with the initial conditions index used as an instrument for the change in the Gini coefficient. Standard errors are reported in parentheses. An asterisk indicates statistical significance at the 5 percent level.

Table A1. Regressions Using Income Residuals

	Workers	Farmers	Workers/farmers	Pensioners
Time	0.001	0.006	0.006	0.000
	(0.002)	(0.006)	(0.006)	(0.001)
Time ^2/10^3	0.000	0.188	-0.171	-0.002
	(0.000)	(0.092)	(0.092)	(0.023)
Time ^3/10^6	0.001	0.919	0.859	-0.009
	(0.010)	(0.365)	(0.364)	(0.095)
Mean real income	0.207	0.126	0.116	0.367
	(0.014)	(0.013)	(0.021)	(0.027)
Dummy for 93-97	0.165	0.369	0.201	0.044
	(0.028)	(0.102)	(0.098)	(0.026)
Adjusted Rsquared	0.88	0.65	0.51	0.85
Number of observations	92	92	92	92

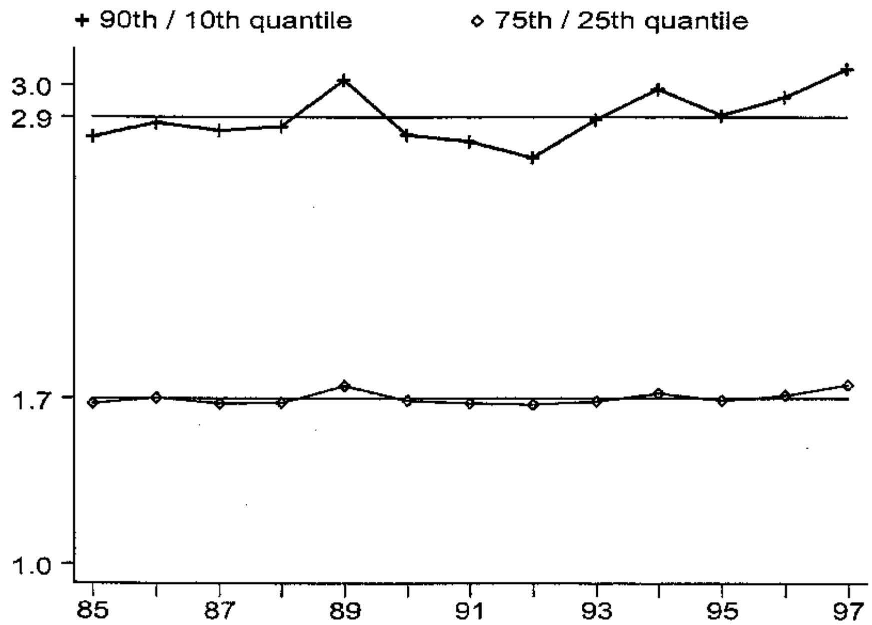
Notes: The dependent variable is the log standard deviation of the residuals from equation (1). Standard errors are reported in parentheses.

Table B1. Equivalence Scales as a Function of Household Composition

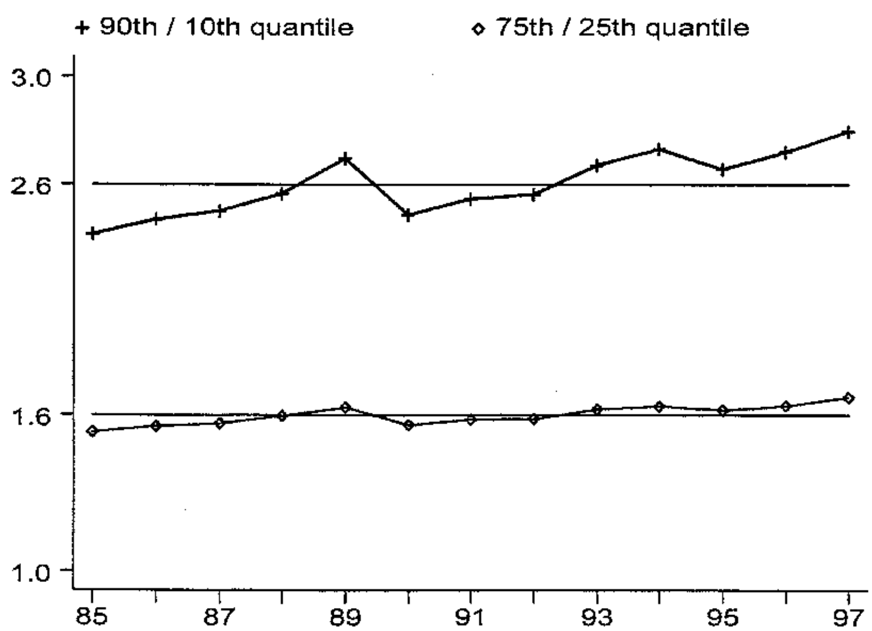
Household Type:	GUS	OECD	McClements	Food-Share Equations OLS	IV
Single person households					
1 HD = Male, 31-60	0.54	0.59	0.55	0.74	0.71
2 HD = Male, 18-30	0.54	0.59	0.55	0.72	0.70
3 HD = Male, >60	0.54	0.59	0.55	0.74	0.68
4 HD = Female, 31-60	0.46	0.59	0.55	0.66	0.65
5 HD = Female, 18-30	0.46	0.59	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. The equivalence scales shown in the last two columns are based on OLS and IV estimates, respectively, of food share equations using the HBS data. Equivalence scales based on IV estimates (column 5) are used in the paper.

Figure 1. Income and Consumption Quantile Ratios



Ratios of Upper to Lower Quantiles: Income



Ratios of Upper to Lower Quantiles: Consumption

Figure 2. Kernel Density Estimates

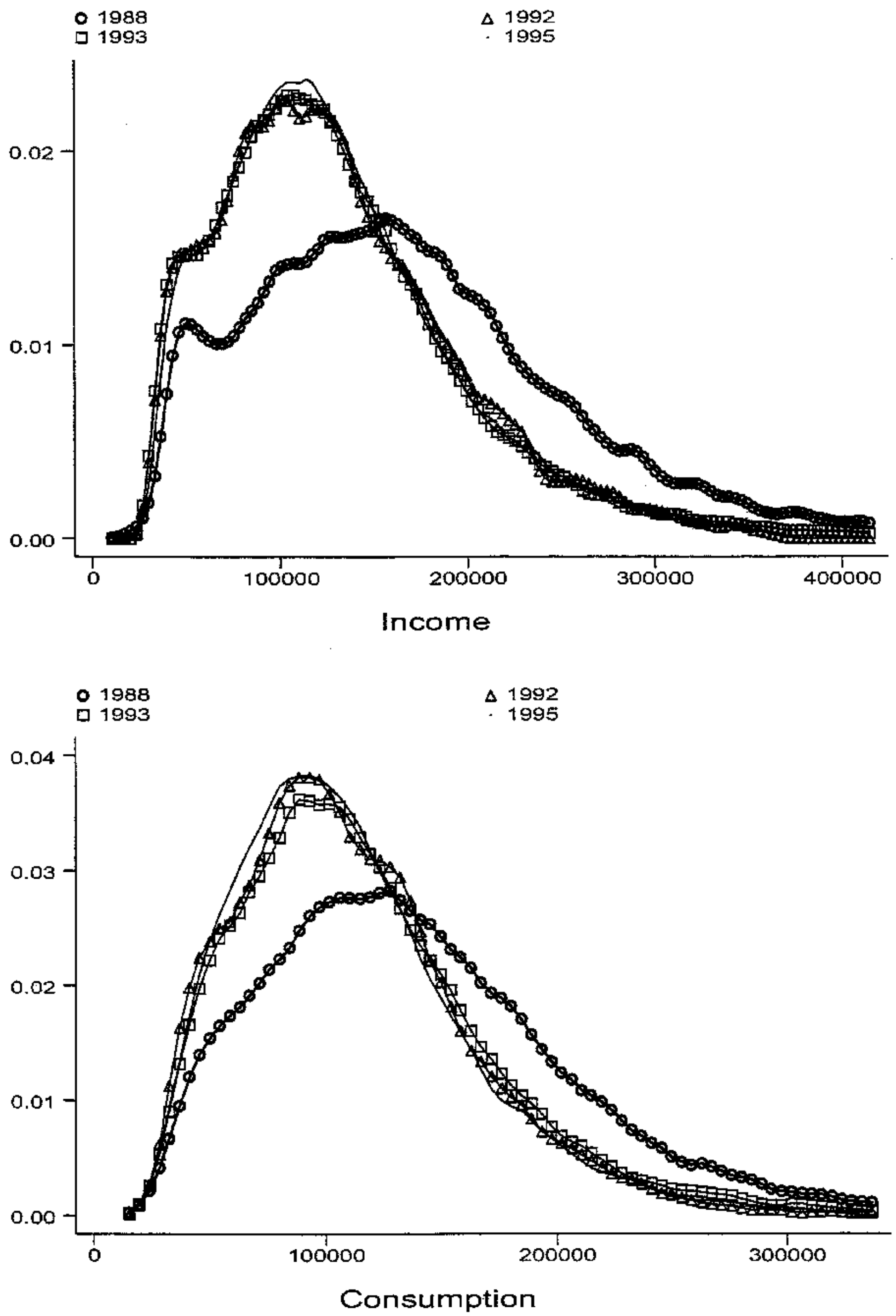


Figure 3. Kernel Density Estimates: Adj. Income

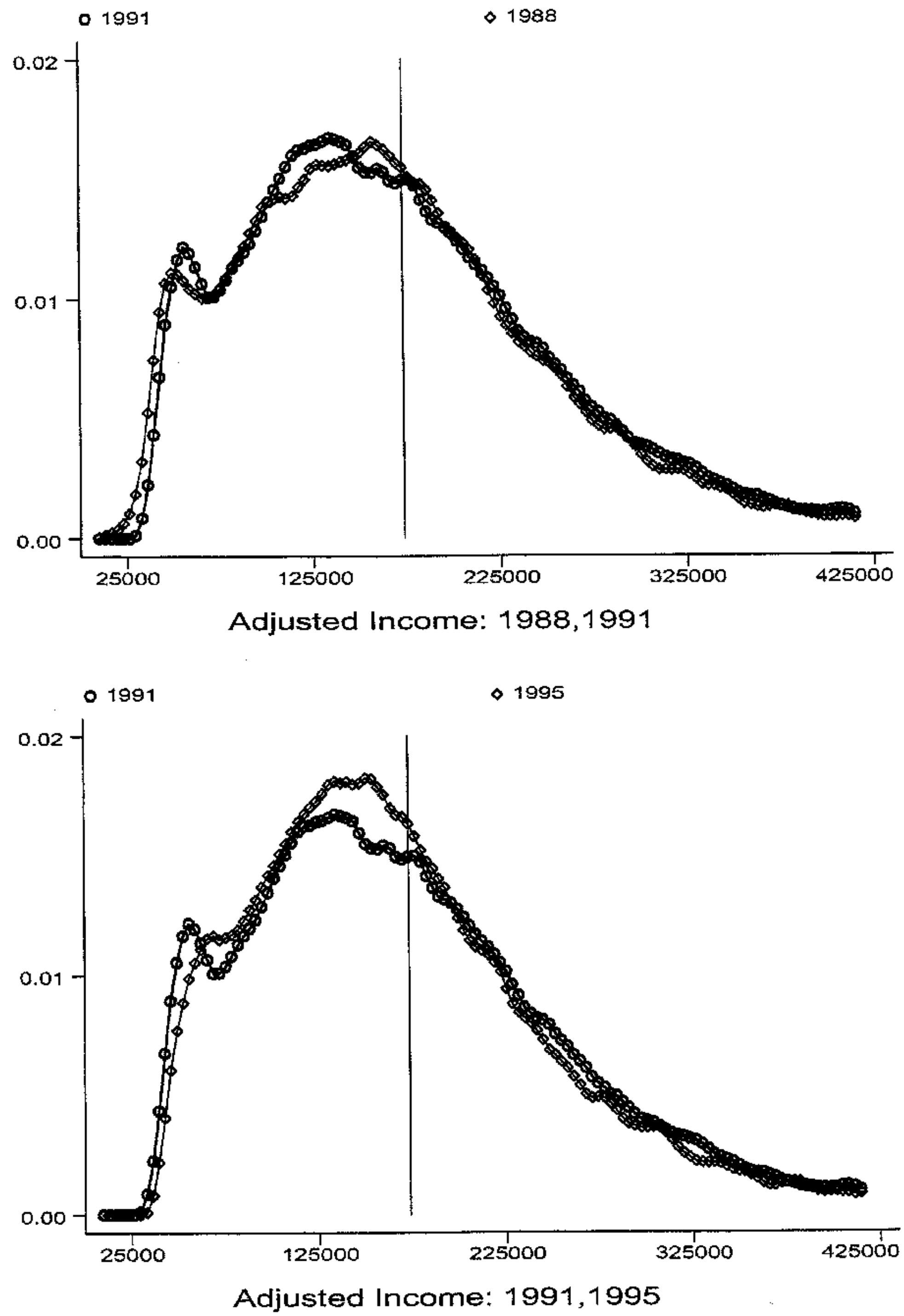


Figure 4. Median income, consumption for diff. socioeconomic groups

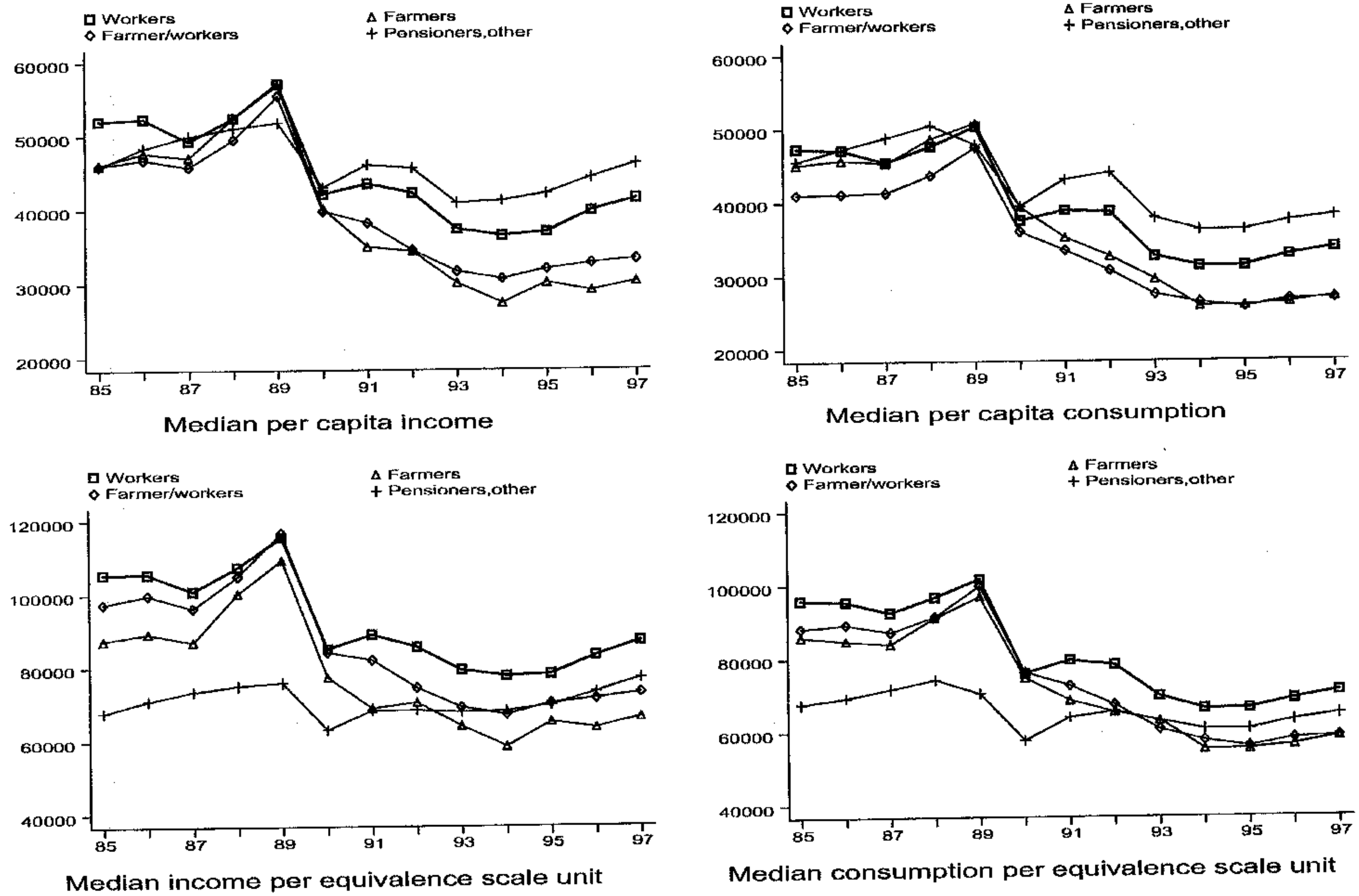


Figure 5. Transfers, Income Net of Transfers: Nonparametric Estimates

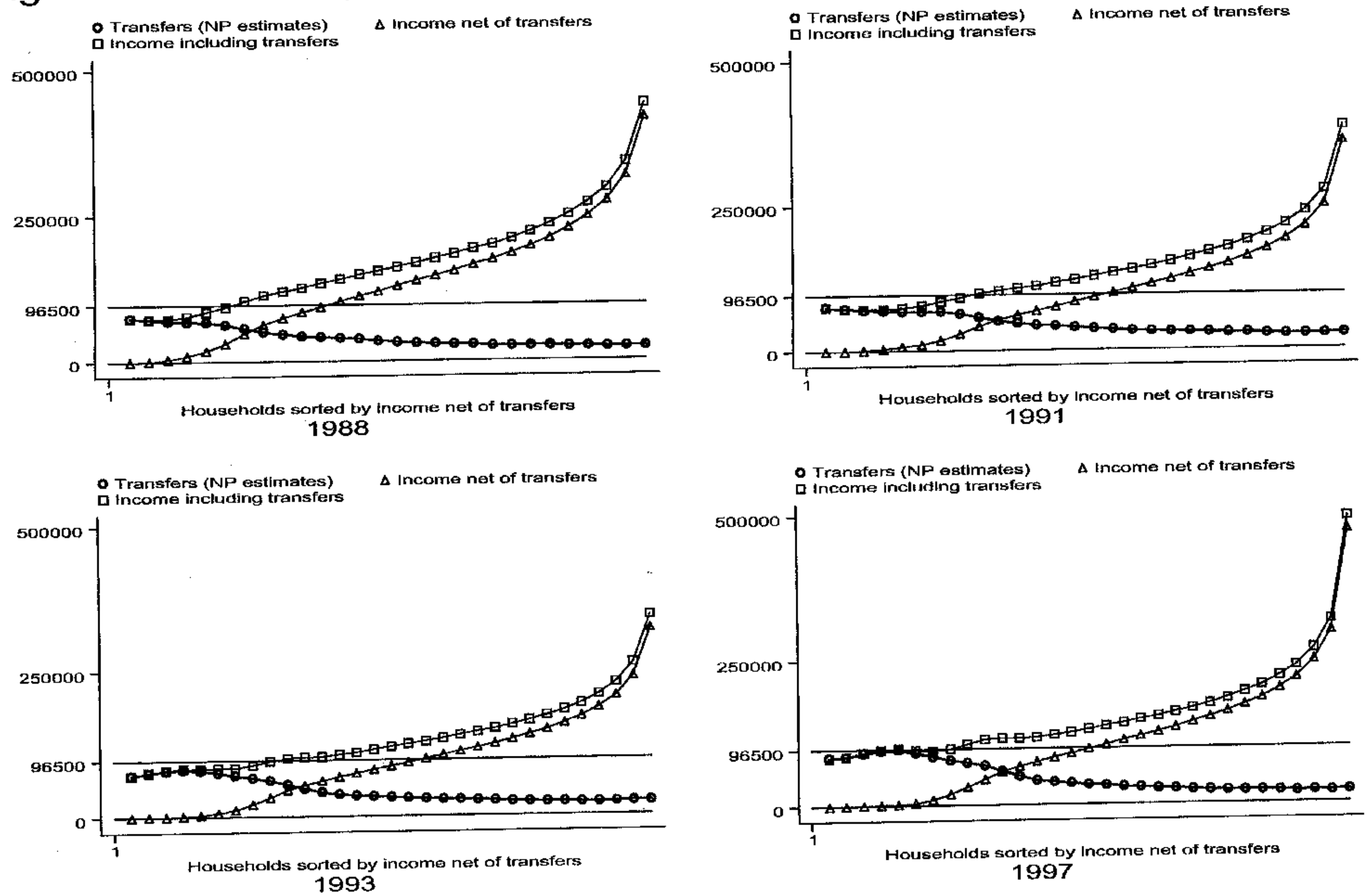


Figure 6. Transfers, Inequality and Growth During Transition

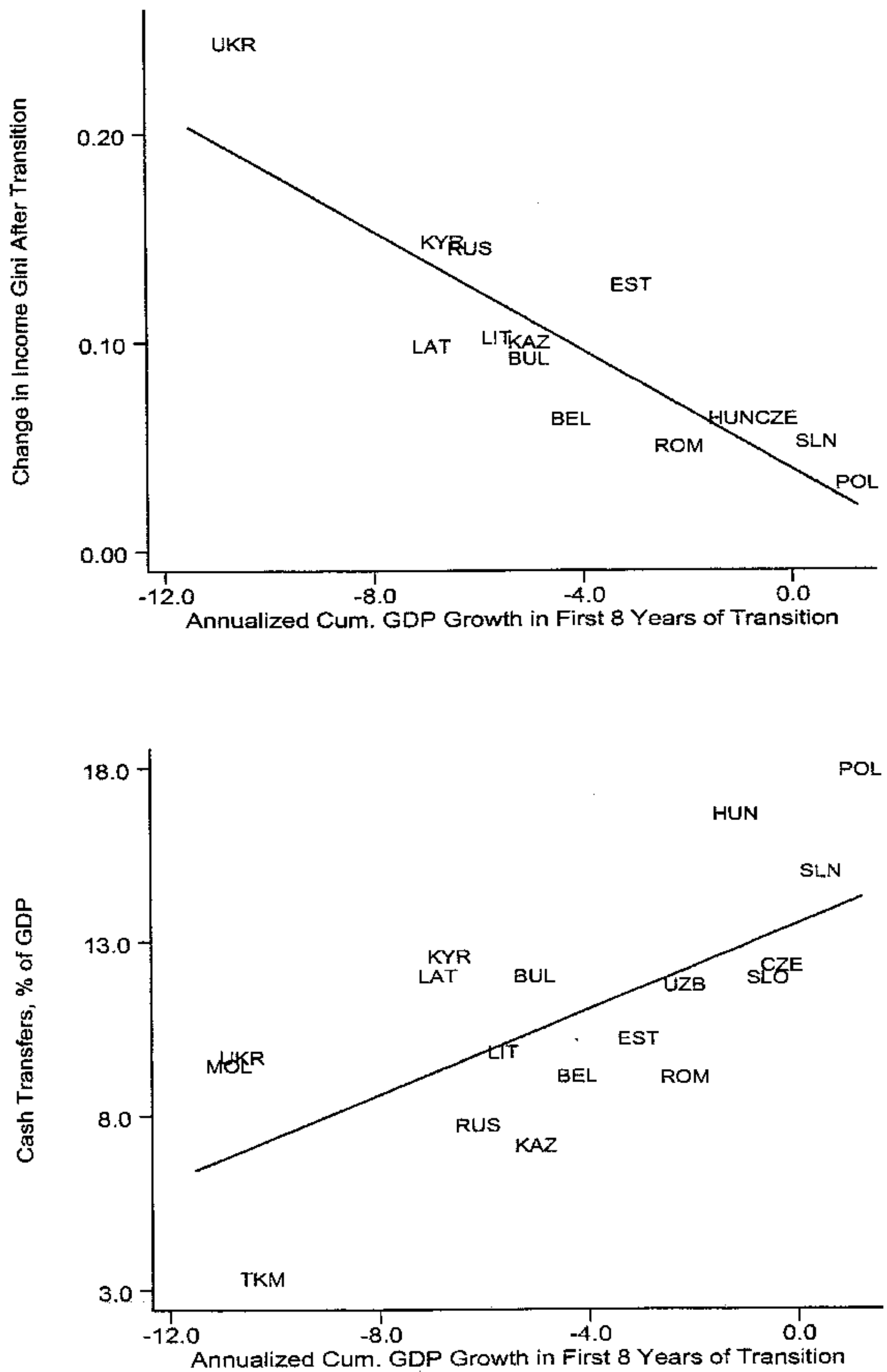


Figure A1. Income Residuals, Fitted Values from Stage II Regressions

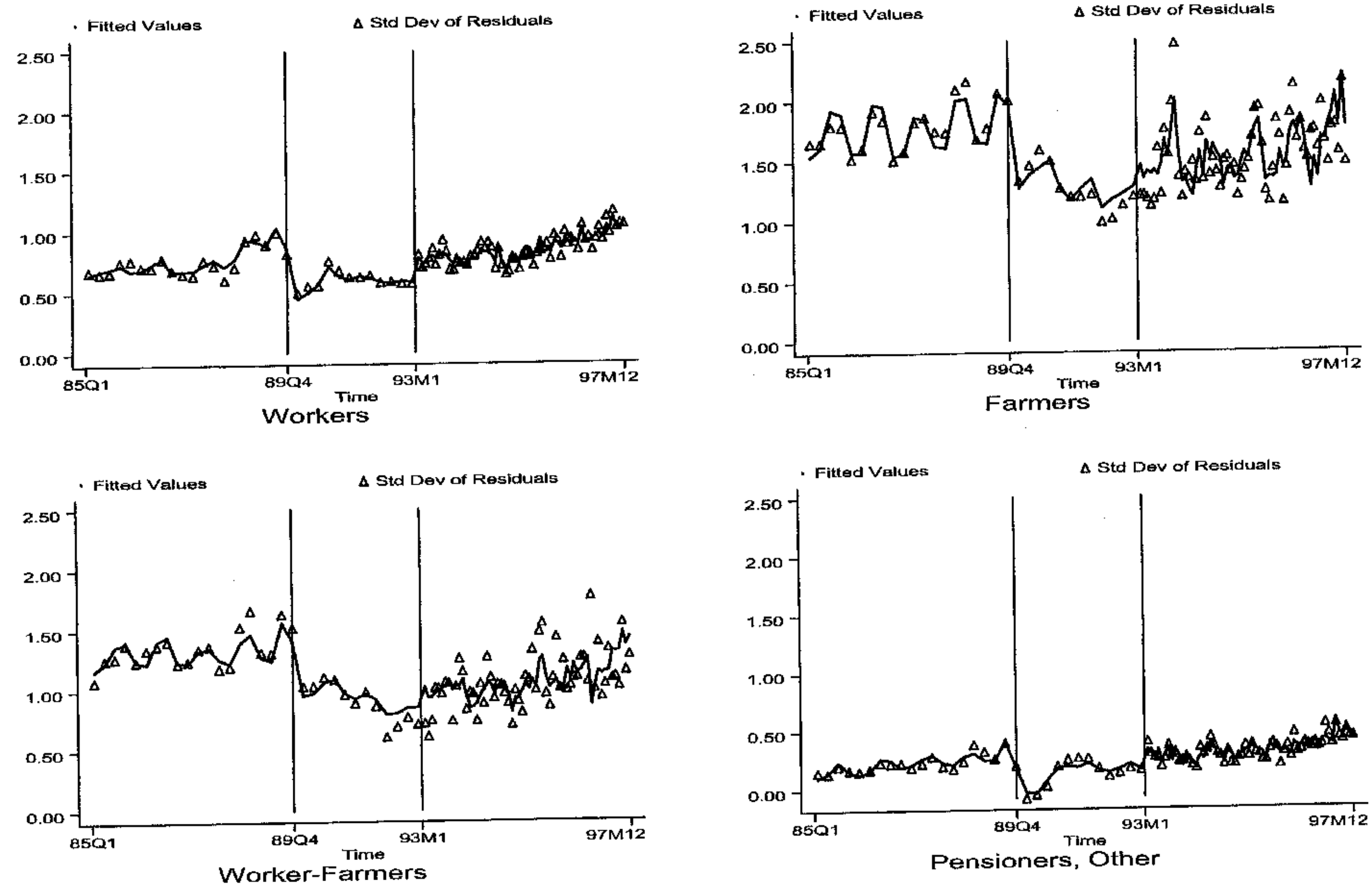
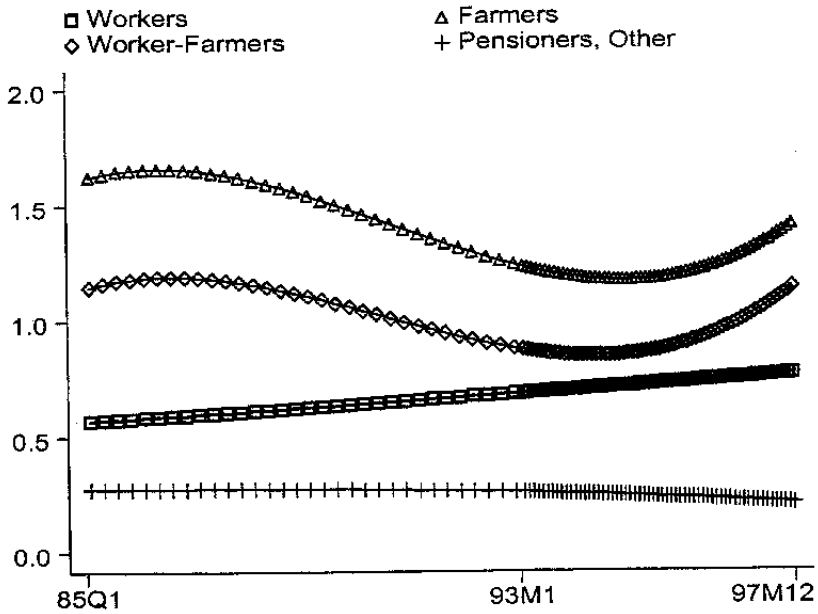
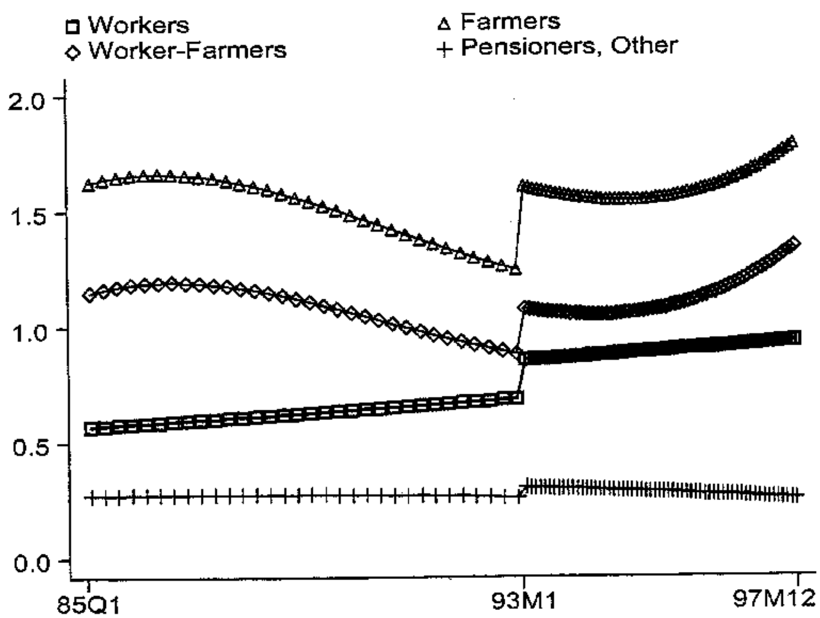


Figure A2. Time Polynomials, Stage II Regressions



Fitted Time Polynomials



Fitted Time Polynomials + Estimated Dummy for 93-97

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