GROWTH AND INEQUALITY: EVIDENCE FROM TRANSITIONAL ECONOMIES

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Abstract

The transitional economies of Eastern Europe (EE) and the former Soviet Union (FSU) experienced a dramatic increase in income inequality in the 1990s. In this paper I investigate the causes of unprecedented changes in income distribution using a unique panel of inequality estimates for 24 transitional countries for the period 1989-1998. The fixed effects model is used to control for unobservable country-specific effects that result in a missing-variable bias in cross-sectional studies. The relationship between income inequality, measured by Gini coefficient, and per capita GDP is shown to be positive for EE, but negative for the FSU. Economic liberalization, privatization and deindustrialization are found to have contributed to the rise in income inequality in the transitional region. Hyperinflation also makes the distribution of income more unequal. I do not find strong support for unemployment and the size of government consumption affecting income distribution. While civil conflicts increase income inequality, the extent of political rights and civil liberties is not found to directly affect income distribution.

JEL Classification: D31, O15, O57.

Keywords: economic growth, income distribution, transitional economies, panel data.

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I. Introduction

The relationship between income inequality and economic growth has received much attention in the economic literature. The impact of economic development on income inequality, however, remains ambiguous. Even if found to be significant in univariate regressions of income inequality on per capita GDP, the parameter estimate on aggregate income loses its strength and can even reverse sign when other explanatory factors or country-specific dummies are introduced (Deininger and Squire, 1998).

However, a common trait of the previous studies linking income inequality and economic growth is that they concentrated primarily on what happens to income distribution during the process of *development*, that is of rising per capita income. In contrast, the countries of EE and the FSU witnessed a sharp contraction in output during the initial stage of the transition.[1] This decline has been accompanied by a marked increase in income inequality, though, not at a uniform rate across the region. In many transitional economies inequality has reached levels comparable to that observed in highly unequal countries of Asia and Latin America.

These developments in transitional countries pose many intriguing questions. What is the role of economic *decline* (and recovery) in changing

[1] In many countries of the FSU the outputs declined to 30-40 % of their pre-transition levels.

income distribution? What specific factors lie behind a noticeable increase in income inequality over the transition? How well do the same factors explain the changes in inequality across different counties?

I attempt to answer these questions in this paper using a unique panel of inequality estimates constructed for 24 transitional countries of EE and the FSU and embracing the period from 1989 to 1998. The fact that the combined population of these countries exceeds 400 million people makes the understanding of the factors driving the changes in income distribution go far beyond a purely research interest. Although it is often argued that policy makers should be more concerned about absolute poverty than income inequality, there are several reasons why one may (or should) care about the latter as well. At a given rate of economic growth, more unequal distribution of income would be associated with a lower rate of poverty reduction, assuming, of course, that the poor participate fully in sharing the gains from growth. Moreover, as suggested in many studies (e.g., Alesina and Rodrik, 1994; Birdsall et al., 1995; Deininger and Squire, 1998; Persson and Tabellini, 1994; Sylwester, 2000; Easterly, 2001), an unequal income distribution might itself be detrimental to long-run economic growth for a variety of reasons.[2] The most common arguments for this are that an unequal distribution of income creates pressure for re-distributional policies, and hence distorts

[2] A number of recent studies (e.g., Forbes, 2000) suggest, however, that in the short run the relationship between inequality and growth could be positive.

incentives for working and investing; that it leads to abuse of power by the elite and to sociopolitical instability and, thus, harms the investment environment; and finally that, in the presence of imperfect capital markets, it reduces opportunities for accumulating human capital (such as education and health) and physical assets. From a social welfare point of view, it has also been argued that both utilitarian and non-utilitarian views of welfare suggest that income inequality reduces aggregate well-being.[3] These considerations leave no doubt that inequality indeed matters, and in this paper I investigate which factors underlie the trends in inequality observed in transitional economies.

There is a growing amount of research which attempts to explain the rise in income inequality during the transition. Many existing studies try to figure out the possible factors behind the changes in the distribution of income using either theoretical models of transition (Aghion and Commander, 1999; Ferreira, 1999; Milanovic, 1999) or a Gini decomposition analysis (by income component or recipient) applied to a single country or a set of countries (Garner and Terrell, 1998; Milanovic, 1999; Yemtsov, 2001). Yet a third approach employs cross-country regressions to examine why income inequality is different *across* countries at a given point in time (the World Bank, 2000).

[3] Grun and Klasen (2001) applied a set of inequality-adjusted indicators of well-being to measure aggregate welfare in transition countries. They found that an adjustment for inequality significantly influences the ranking of transition countries in terms of their absolute levels of well-being and the achievements in well-being over time.

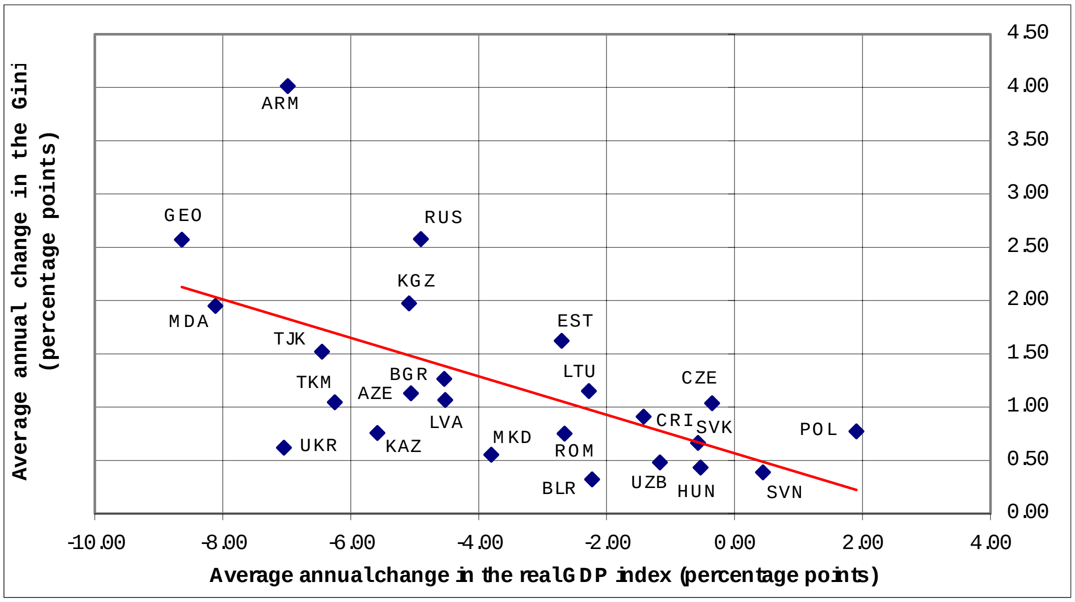
This paper represents the first attempt to identify factors underlying the changes in income inequality over time *within* countries rather than to explain differences in inequality levels across countries. Until now a lack of compatible time series data with sufficient geographical coverage ruled out the possibility of doing this, and I undertake the task using the assembled panel of inequality estimates comparable over time and across countries. I use panel data estimation methods to control for unobservable country-specific effects that result in a missing-variable bias in cross-sectional studies.

The remainder of the paper is structured as follows. In Section II I present some evidence on the evolution of income inequality and economic growth during the transition. Section III discusses potential determinants of rising inequality in transitional countries with a reference to existing literature. Section IV describes the data used in the empirical analysis. Section V is devoted to model specification and description of the estimation technique. Section VI describes regression results. In Section VII I examine the robustness of results. Section VIII offers a conclusion and presents some policy implications of the findings.

II. Growth and Inequality during the Transition

Figure 1 shows the relationship between the changes in income inequality and changes in real GDP during 1989-1998. It clearly suggests that better growth performers experienced much smaller increases in income inequality.

Figure 1. The dynamics of income inequality and GDP growth in transitional economies, 1989-1998



Source: Author's calculations using a constructed panel of inequality estimates and the Real GDP index from the TransMONEE2000 database, UNICEF, Florence.

There is substantial variation in the regional performance, with the transitional economies of EE performing much better, both in terms of economic growth and distributional outcomes, than the FSU countries. However, there are significant differences within these two groups of countries as well.

Although quite illustrative, inequality and growth dynamics presented in Figure 1 may be misleading as they do not fully reflect what happened at different stages *within* this period. For instance, given the evidence presented in Figure 1, one may mistakenly conclude that Poland (POL) and Slovenia (SVN) were growing consistently through the 1990s while other countries were declining, and that inequality was uniformly trending upwards during the period. Therefore, in Table 1 I present the evidence on the evolution of inequality and growth separately for economic decline and economic recovery episodes. This analysis gives us a better idea of the relationship between income distribution and economic growth.

Table 1. The dynamics of income inequality and GDP growth in transitional economies of EE and the FSU

Region/ Country	Population (mid-1997, millions)	Real GDP index (1989=100)		Avg. annual change in the Real GDP index (percentage points)		Avg. annual change in the Gini index (percentage points)	
		at the bottom of decline (*year*)	in 1998	decline period	growth period	decline period	growth period
1	2	3	4	5	6	7	8
I. FSU							
a) Baltic states							
Estonia	1.5	60.76*(94)*	75.70	-7.85	3.74	3.31	-0.65
Latvia	2.5	51.04*(95)*	59.30	-8.16	2.75	1.00	1.20
Lithuania	3.7	64.83*(94)*	79.53	-7.03	3.67	2.51	-0.26
b) Western CIS							
Belarus	10.2	62.69*(95)*	77.75	-6.22	5.02	0.31	0.44
Moldova	4.0	-	32.00	-8.11	-	1.95	-
Russia	147.0	-	55.89	-4.90	-	2.58	-
Ukraine	50.4	-	36.61	-7.04	-	0.69	-
c) Caucasus							
Armenia	3.8	31.63*(93)*	41.68	-15.29	2.08	8.42	-0.40
Azerbaijan	7.8	41.86*(95)*	49.40	-9.69	2.51	2.05	-0.33
Georgia	5.4	24.60*(94)*	31.70	-8.64	-	2.57	-
d) Central Asia							
Kazakhstan	15.3	-	61.20	-5.49	-	0.84	-
Kyrgyz Rep.	4.6	50.39*(95)*	60.30	-5.09	-	1.97	-
Tajikistan	6.0	39.19*(96)*	41.90	-6.44	-	1.69	-

Turkmenistan	4.6	41.99*(97)*	43.75	-6.25	-	1.04	-
Uzbekistan	23.6	83.36*(95)*	89.50	-3.18	-	0.48	-
II. Central EE							
Czech Rep.	10.3	84.58*(92)*	94.90	-5.14	2.53	0.33	1.46
Hungary	10.2	81.89*(93)*	95.20	-4.53	2.66	0.32	0.52
Poland	38.7	82.21*(91)*	117.15	-5.22	5.47	-0.34	1.33
Slovak Rep.	5.4	74.97*(93)*	99.60	-6.26	4.09	0.41	0.74
III. South EE							
Bulgaria	8.3	63.69*(97)*	65.90	-4.54	-	1.27	-
Romania	22.6	74.99*(92)*	82.08	-8.34	3.30	0.63	1.37
IV. FY							
Croatia	4.7	59.54*(93)*	77.70	-10.11	3.65	0.35	1.36
Macedonia	2.0	67.99*(95)*	71.50	-5.34	0.79	0.60	0.41
Slovenia	2.0	82.04*(92)*	103.90	-5.99	3.66	0.37	0.40

Source: Author's calculations using a constructed panel of inequality estimates and the Real GDP index and population data from the TransMONEE2000 database, UNICEF, Florence.

Note: "-" in Column 3 means that by the end of 1998 a country under consideration continued to decline. South EE also includes Albania; the former Yugoslavia (FY) also includes Yugoslavia, FR and Bosnia-Herzegovina. These countries are not included in the table due to the lack of data.

Several major observations emerge out of the data in Table 1. First, no single country escaped economic decline and an increase in income inequality (except Poland) at the start of the transition (see Columns 5 and 7, Table 1).

Second, after the sharp economic decline in the initial period, most of the countries started to recover at some later stage. In general, Eastern Europe and the Baltic states began growing in 1992-1994, while the non-Baltic FSU countries started to grow later or continued to decline as of 1998 (see Column 3, Table 1).

Third, it appears that the economic recovery in the FSU countries was generally associated with declining income inequality. Conversely, recovery in EE countries was accompanied by rising income inequality, although at very modest rates (see Columns 6 and 8, Table 1). This is a very interesting observation since it indicates that the mechanisms behind the inequality trends in EE and the FSU are not necessarily the same.

In what follows I discuss potential determinants of the changes in income distribution in transitional countries. This discussion serves as a basis for the choice of variables used later in the empirical analysis. Most of the factors that I consider are those commonly found in the literature on the determinants of cross-country inequality, while others are specific to the transitional region circumstances that I expect to be influential in explaining the pattern of income inequality.

III. Potential Determinants of Rising Inequality in Transitional Countries

There is a vast amount of literature on the determinants of income inequality that considers both the individual (e.g., increasing returns to skills) and macro (e.g., inflation, political democracy) level factors affecting income distribution. In this paper I focus on the latter, although the former might be equally important.

The main factors that I anticipate to affect income inequality in transitional countries are: the level of economic development (measured by per capita GDP), macroeconomic conditions (inflation, unemployment), government involvement in the economy (government consumption, social transfers), structural changes (economic liberalization, privatization, deindustrialization), and forces outside economic domain (political freedom, civil conflicts).

Many attempts to identify a link between income inequality and the level of economic development have been undertaken since the seminal work of Kuznets (1955), who argued for an inverted U-shape relationship between income inequality and economic development. Although several studies (e.g., Paukert, 1973; Ahluwalia, 1976) have found a support for such a relationship, most of the recent research does not find economic development to affect income distribution (e.g., Anand and Kanbur, 1993; Deininger and Squire, 1998; Ravallion, 1995).

However, the striking economic decline in EE and the FSU countries in the initial years of the transition, and the subsequent economic recovery are expected to have had significant implications for income distribution. That is because economic decline and recovery were associated with dramatic and heterogeneous shocks to real incomes, the changes in the real value of social transfers, and other developments in social and economic conditions. Figure 1 provides strong support for anticipating a negative relationship between income inequality and economic development for transitional countries. Nevertheless, as the evidence from Table 1 indicates, this relationship is hardly universal across countries.

Inflation may have a strong redistributional impact through its effect on individuals whose nominal incomes are not adjusted proportionally to increases in prices; mostly state sector employees, pensioners and beneficiaries of various social benefits. That would be an argument for a positive relationship between income inequality and inflation. However, inflation may also have an equalizing impact on income distribution through a progressive tax system by pushing wage earners into higher tax brackets, thus implying less inequality in disposable income. These two effects may well counterbalance each other. In a study of the determinants of inequality for OECD countries (Gustafsson and Johansson, 1999), inflation was not found to be significant in explaining inequality. That may not be the case for transitional economies, however, as most of them experienced a sharp rise in inflation at the start of the transition.[4] Moreover, the progressivity of the inflation tax is unlikely to be a mechanism at work in most of the transition countries due to a high occurrence of tax evasion. Hence, I expect inflation to be positively associated with income inequality in the transition region.[5]

Destruction of the old economic system and significant structural changes during the transition caused a substantial rise in unemployment across the region. In many countries the unemployment rate grew from virtually zero to 10-15

[4] In our sample the mean annual inflation is 248 per cent, and the maximum is 9,750 per cent (Turkmenistan, 1993).

[5] The regressivity of the inflation tax and imperfect indexation have been found, using micro-level data, to increase income inequality in Brazil during the years of high inflation (Ferreira and Litchfield, 1998).

percent even when measured by the number of officially registered unemployed.[6] Unemployment is likely to largely affect those in the lower percentile of income distribution. Milanovic (1998) indicates that unemployment in transition countries increased the most amongst women, young people, and those with lower education. A negative impact of unemployment on income distribution has been confirmed in a number of studies of industrialized countries (Gustafsson and Palmer, 1997; Weil, 1984), and I anticipate unemployment to have an inequality-increasing effect in transitional economies as well.

In times of economic hardship and increasing unemployment, government-financed projects (e.g., construction) may provide a source of employment and income (with low-skilled labor probably benefiting the most), which serve as a buffer to widening income inequality.[7] The size of the public sector is found to reduce inequality in cross-country studies by Stack (1978) and Boyd (1988). Government involvement in the economy, measured as a share of government consumption in GDP, decreased between 1989 and 1998 in 8 out of 15 states of the FSU (including the Baltic states), increased in 5 states, and was practically unchanged in the rest. In the EE region government consumption has declined in

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[6] The official unemployment statistics are likely to significantly understate the real level of unemployment. Indeed, unofficial estimates indicate significantly higher levels of unemployment. Moreover, many people are registered as employed even when they are not receiving payment for their labor.

[7] Government consumption, however, can affect the distribution of income not only on the expenditure side, but also through the tax collection.

only two countries, while in the remainder it has either grown or has been relatively stable. In this paper I look at the effect of government consumption on income distribution in transitional economies.[8]

Centrally planned economies were dominated to a various extent by state enterprises with administratively set wages. The overwhelming predominance of the state sector in EE and the FSU economies is widely regarded as a main reason for low income inequality in the region before the transition.[9] The process of transition brought about a massive expansion of the private sector and the share of the private sector employment.[10] This process is likely to increase income inequality due to the wage differential between the state and private sectors. Moreover, the distribution of earnings within the private sector is usually more unequal than in the state sector. That privatization can lead to rising income inequality is argued in theoretical models of transition by Milanovic (1999) and Ferreira (1999). However, due to the poor data on the scope of privatization in transitional countries, the impact of rising private sector on income distribution has not been empirically tested; until now in this paper.

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[8] Another mechanism for government to influence income distribution is through social security transfers. These transfers are found to reduce inequality in a cross-country study of both advanced and developing countries by Milanovic (1994), but appear insignificant in a study of OECD countries by Gustafsson and Johansson (1999). I do not analyze the impact of social transfers in this paper due to the lack of data.

[9] The level of income inequality in EE and the FSU before the transition was widely considered to be lower than in the rest of the world (Atkinson and Micklewright, 1992).

[10] In many transitional countries the share of private sector in total economy has increased from barely existing to 50-60 %.

Economic liberalization also led to profound changes in the sectoral composition of the economy. There is a clear trend for the industrial sector to shrink, while the evidence for the agricultural sector is mixed -- in some countries its relative importance has declined, while in others it has increased.[11] The share of industry in total output in the region declined on average by 25 percent from 1989 to 1998, and in several countries the drop was even more profound. For instance, in the ten years after 1989 the share of the industrial sector declined from 52% to 32% in Poland, 58% to 33% in Slovak Republic, 59% to 25% in Bulgaria, and 50% to 35% in Russia. It is very likely that the declining industrial sector employment may have an inequality-increasing impact due to an outflow of labor to sectors with higher wage differentials, for instance, services.[12] A negative relationship between industrial sector employment and income inequality is confirmed in studies of industrialized countries by Gustafsson and Johansson (1999) and Levy and Murnane (1992), and in this paper I investigate the effect of deindustrialization on income distribution in the transitional region.

The process of economic transition in EE and the FSU was generally accompanied by the expansion of political democracy. Although a common

[11] Diminishing relative importance of the industrial and/or agricultural sector was offset by the growing relative importance of services sector, which in many countries has increased 1,5 - 2 times.

[12] Note that if the labor force moves from the remaining state-owned industrial sector to the private sector, the potential effect of this on income distribution would be captured by the privatization variable.

argument in the literature is that the higher degree of political democracy should be accompanied by a more equal distribution of income (e.g., Gradstein et al., 2001; Rodrik, 1999), the existing evidence does not show any robust relationship between democracy and inequality in a cross-country regression analysis. Here I investigate whether political democracy affects income inequality in transitional countries.

A number of countries in EE (republics of the former Yugoslavia) and the FSU (Armenia, Azerbaijan, Georgia, Tajikistan, Moldova, and Russia) experienced persistent internal conflicts over the last decade. Since civil conflicts are likely to have strong distributional consequences I analyze their impact on income inequality in the transitional region.

The data used in the empirical analysis and their sources are described in detail in the next section.

IV. The Data

I construct a panel of inequality estimates using time-series data on income inequality across transitional countries. The majority of observations are drawn from the UNU/WIDER-UNDP World Income Inequality Database (WIID) (Version 1.0, September 2000), which to date represents the latest and most extensive data on inequality for both developed and developing countries.[13] In addition, I augment these data with a few observations from Milanovic (1998) (mainly for 1989) and the latest household surveys conducted by the World Bank (2000) (mainly for 1998-1999).

To minimize problems with data comparability across countries and over time I require inequality data that I select for the panel to be based on the same living standard indicator, have the same sample and enumeration unit, be drawn from nationally-representative surveys, and, whenever possible, come from one source.[14] Income inequality is measured by the Gini coefficient with individuals representing the unit of analysis. The coefficients are calculated based on household per capita income. The compiled panel of inequality estimates represents perhaps the most consistent and extensive coverage available for transitional countries to date. It consists of 149 observations covering 25 countries in transition from 1989 to 1999.[15] However, due to either missing observations on other variables, or the deletion of observations based on the influence diagnostics tests (as discussed below), only 129 out of 149 originally assembled Gini coefficients are used in the estimation. A detailed description of the data on

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[13] About half of the WIID database is formed by K. Deininger and L. Squire's 1996 database. However, for the transitional region most of the data in the database come from the UNICEF/IRC TransMonee2000 Database, Florence.

[14] In the original WIID database most of the countries are represented with multiple time series of inequality estimates that often are not compatible.

[15] As there are only 5 observations for 1999, in the estimation they are used as 1998 values.

income inequality used in the empirical analysis and their sources is presented in Table A1 in the Appendix.

The constructed panel of inequality estimates is far from being perfect, however, as not all of the above comparability requirements could always be met, and the resulting inequality measures are still subject to potential measurement error problems. The use of panel data and panel data estimation methods (to be discussed), however, help diminish some problems with data consistency. The country-specific intercepts in the fixed effects model setting can absorb, among all other unobservable characteristics, the differences in inequality definition across countries (Deininger and Squire, 1998).[16] Nevertheless, the use of panel data cannot remedy all data limitations, and thus the empirical results must be treated with some degree of caution.

I now turn to the definitions and sources of data on explanatory variables. The level of economic development is measured by PPP-adjusted per capita GDP in constant 1992 USD. The data on PPP-adjusted per capita GDP in current USD come from the World Bank World Development Indicators (WDI) 2000 database, and they are then deflated to 1992 prices using the U.S. GDP deflator.

Inflation is measured as the annual percentage change in the consumer price index (CPI) (end-year). As CPI-based inflation is not available for all countries in

[16] This is true, however, only if these differences are *systematic*. I am thankful to a referee for pointing this out to me.

our sample for 1989 and 1990, the GDP deflator inflation is taken for those years instead. Finally, as neither of the mentioned above indexes could be obtained for Croatia (1989), Macedonia FYR (1989, 1990), and Slovenia (1989) inflation there is measured by the food price index (a sub-index of the CPI). All inflation data are drawn from the World Bank WDI 2000 database.

Unemployment represents a share of the labor force that is without work but available for and seeking employment. However, as I have mentioned, unemployment data for transitional countries may substantially understate the actual scope of unemployment. Nonetheless, as no better alternative is available, official estimates are used in most cases. Unemployment data are taken from the EBRD Transition Report 2000, which provides further reference on the origination of the data for each country.

General government consumption, expressed as a fraction of GDP, refers to all current spending for purchases of goods and services (including wages and salaries). It also includes most expenditures on national defense and security, but excludes government military expenditures that are part of government capital formation. As such, government consumption represents a good measure of the government's involvement in the economy. The data on government consumption come from the World Bank WDI 2000 database.

Industrial employment represents a share of industry in total employment. I was able to obtain only 105 observations covering 24 countries (from the EBRD Transition Report 2000). As other explanatory variables contain more observations, the use of these employment data in the model estimation would substantially reduce a number of observations on other variables. Since the sample is relatively small, I consider that inappropriate. Therefore, I use a share of industry value added in GDP as a proxy for industrial sector employment.[17] This provides us with a substantially larger number of observations. The data come from the World Bank WDI 2000 database.

Private sector employment equals the number of people employed in the private sector as a percentage of total employment. Data availability, however, represents a severe constraint here, as practically no data prior to 1993-1994 exist. I have managed to collect 51 observations using IMF country reports, a number not sufficient for our purposes. Thus, in the regression analysis I use a share of the private sector in GDP as a proxy for the private sector employment. Since I have found high correlation between the size of the private sector and the private sector employment in our sample (for those observations that are available), and in view of the lack of an alternative, such a proxy is considered to be justifiable. These

[17] As a check of how good this proxy is I estimate the model with the industrial employment variable as well. The results (discussed below) indicate that industry value added is a very good proxy indeed.

data and their more extensive descriptions are available in the EBRD Transition Report 2000.

Economic liberalization (which is largely reflected in structural changes) is measured with the Cumulative Liberalization Index (De Melo et al., 1996), which reflects the progress with economic reforms on several fronts: internal (price) liberalization, external (foreign trade) liberalization, and the extent of privatization and banking sector reform.[18]

The progress in the introduction of political rights and civil liberties during the transition is measured using the Index of Political Freedom (IPF), which represents an arithmetic average of the political rights and civil liberties indexes (Freedom House, 2001). The political rights index reflects the extent to which people in a country can participate in the political process. The civil liberties index measures the freedoms to develop views, institutions, and personal autonomy apart from the state.

The effect of civil conflicts on income inequality is measured using a dummy variable set equal to one for each year since an internal conflict has taken place in a given country.[19] In our sample the countries affected by civil conflicts are Croatia, Macedonia FYR, Armenia, Azerbaijan, Georgia, Moldova, Tajikistan

[18] The data are updated up to 1998.

[19] The assumption is that the effect of civil conflict (if any) on income inequality is likely to persist for a certain period.

and Russia.[20] The historical information on civil conflicts in the region is obtained from the Reuters Foundation.[21]

Table A2 in the Appendix provides the descriptive statistics of the data used in the regression analysis. Table A3 shows the matrix of the Pearson correlation coefficients.

V. Model Specification and Estimation

The primary interest of this study is to explain the changes in income inequality in transitional economies, and I thus estimate income inequality as a function of various potential explanatory variables presented below. The base model specification is:

$$GINI_{(it)} = \alpha_i + \beta_0 * GDPPC_{(it)} + \beta_1 * GDPPC\_S_{(it)} + \beta_2 * INFL_{(it)} + \beta_3 * UNEMP_{(it)} + \beta_4 * CONSG_{(it)} + \beta_6 * INDVA_{(it)} + \beta_7 * PRIVS_{(it)} + \varepsilon_{(it)}; \quad (1)$$

$i = 1, \ldots, N; t = 1, \ldots, T;$

where i represents country index, t denotes time period, GINI is the Gini coefficient of income inequality, α_i is a country-specific intercept, GDPPC is PPP-adjusted GDP per capita (1992 constant USD), GDPPC_S is its squared value, INFL is annual inflation as measured by the year-to-year change in the consumer price index, UNEMP is a share of unemployed in total labor force, CONSG is

[20] In the case of Russia I refer to the military conflict in Chechnya.
[21] See http://www.alertnet.org.

general government consumption as a percentage of GDP, INDVA is industry value added as a percentage of GDP, PRIVS is the private sector share in GDP, and $\varepsilon_{(it)}$ is an error term. The assumption on $\varepsilon_{(it)}$ is that $\varepsilon_{(it)} \sim IID(0, \sigma_e^2)$. All variables (including the Gini coefficient) enter the regressions in the natural log form. The natural log of (1+INFL/100) is used for INFL variable in the estimations to deal with negative and very high values of INFL. The natural log of (1+UNEMP) is used for UNEMP variable in the estimations since the unemployment rate equals zero for many countries at the start of the transition. A squared value of GDPPC (with GDPPC expressed in the natural log form) is included into the regression to account for the potential quadratic relationship between income inequality and per capita GDP.

In view of the large body of literature exploring the effect of income distribution on economic growth, one may be quick to point out the possible problem with the given model specification arising from the potential existence of a reverse causality between inequality and growth. I argue that the transition economies of EE and the FSU represent a unique case when the possibility of causality from income distribution to economic growth can be ruled out, at least for the period under investigation, since the reasons for the economic collapse and subsequent recovery in the region had clearly nothing to do with the distribution of income. This is confirmed by the Granger causality tests (Granger, 1969). I have tested whether income inequality Granger-causes economic growth using from one up to five lags, which provides us with 87 to 22 observations respectively. In *none* of the cases did the test statistic indicate that I could reject the null hypothesis that the Gini coefficient *does not* Granger-cause per capita GDP.[22]

I estimate equation (1) using the assembled panel for 24 transitional countries covering a period from 1989 to 1998. The use of panel data produces several well-known advantages. The most important is that it allows one to control for unobservable time-invariant country-specific effects that result in a missing-variable bias, an often-encountered problem when cross-section data are used. This problem is recognized in Bourguignon and Morrison (1997), Bruno et al. (1995), Deininger and Squire (1998), Forbes (2000), Ravallion (1995), and other studies.

To control for unobservable country-specific characteristics I introduce country-specific intercepts in the fixed effects model setting. The addition of fixed effects to the model also helps alleviate potential heteroscedasticity problems stemming from possible differences across countries (Greene, 1997).

There might be another reason for preferring the fixed effects model to the random effects model. A crucial assumption for the random effects model is that

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[22] The results of these tests are not shown, but are readily available upon request from the author.

country-specific terms (α_i) are uncorrelated with the other explanatory variables. Its violation makes random effects estimates biased and inconsistent (Greene, 1997). The use of the fixed effects model avoids this problem as individual effects are allowed to be correlated with other regressors. To test whether the country-specific effects are correlated with the exogenous variables I conduct a Hausman test.[23] While the test statistic suggests that I cannot reject the null hypothesis for the base model, the results of a Hausman test are found to be sensitive to the model specification and sample selection. Therefore, I prefer to use the fixed effects model.

More importantly, the fixed effects model is chosen since the main goal of this study is to investigate what factors have caused substantial changes in income inequality over time within countries rather than to explain variation in inequality across countries.[24] Thus, the use of the fixed effects estimator, which is also called the "*within*" estimator, is very appropriate since it allows one to focus on how changes in *within-country* characteristics are related to changes in *within-country* inequality. The fixed effects model is also more suitable when the focus is on a specific set of countries and the inference is restricted to these countries (Baltagi, 1996, p.10). Moreover, the country-specific effects have been found to be

[23] H_0 is that there is no correlation. The test statistic is distributed as X_k^2, where k denotes the dimension of the slope vector β (Baltagi, 1995, p.68).

[24] The random effects estimator combines the within and between estimators, thus giving some weight to the cross-country variation.

statistically significant.[25] This implies that if the model is estimated without taking them into account (i.e., if a single overall intercept is included instead of country dummies) the estimated coefficients will be biased.

The fixed effects estimation technique, however, is not perfect. First, random effects estimates are more efficient than fixed effects ones given that all necessary assumptions are satisfied. Second, the fixed effects model is very costly in terms of the lost degrees of freedom, which may represent a particular problem for the relatively small sample. To overcome this problem I estimate the model in deviations from the country means.[26] This *within-countries* estimator is identical to the least squares dummy variable (LSDV) estimator obtained if a dummy variable is included for each country (as in the original formulation of equation (1)), but the resulting R^2 is lower (Greene, 1997, p. 619).

Third, it has been argued (Barro, 1997, p. 37; Temple, 1999) that the fixed-effects technique eliminates the cross-sectional information and, hence, lowers precision of the estimates. This problem is, of course, especially acute if most of

[25] F-test of the null hypothesis that country-specific effects are not significant yields an F-value of 6.14, which is higher than a critical value of 2.07. This means that I can reject the hypothesis that there are no country-specific effects of omitted variables. The same test was conducted for time effects. In this case the $F_{9,\ 88}$ statistic was 2.10, indicating that it is not possible to reject the hypothesis of no time-specific effects. Hence, the use of the one-way model is appropriate.

[26] Specifying the original formulation of equation (1) as: $y_{it} = \alpha_i + \beta' X_{it} + \varepsilon_{it}$, the formulation in terms of deviations from the country means becomes: $(y_{it} - \overline{y_i}) = \beta'(X_{it} - \overline{X_i}) + (\varepsilon_{it} - \overline{\varepsilon_i})$, where $\overline{y_i} = \Sigma y_{it} / t$; $\overline{X_i} = \Sigma X_{it} / t$; $\overline{\varepsilon_i} = \Sigma \varepsilon_{it} / t$.

the variation in the data is due to cross-country differences. While in non-transition countries most (approximately 90%) of the variation in inequality is due to variation across countries (Deininger and Squire, 1996; Li et al., 1998; Quah, 2001), in transition countries a substantial source of the variation in inequality over the last decade is attributable to the profound changes in inequality over time (see Table A2, Appendix). Hence, the use of the fixed effects estimation for transitional countries seems to be appropriate. In addition, if one bears in mind that inequality comparisons across countries are likely to be much less reliable than inequality comparisons for a single country over time, despite all efforts to assemble the inequality estimates that are as consistent as possible both across space and time, the reliance on mostly over-time variation in inequality is even desirable.

Given all considerations outlined above I prefer to use the fixed-effects model since its advantages seem to outweigh its weaknesses given the data and research purposes. Nevertheless, to check the robustness of the results to the estimation technique the random effects model is also estimated. The empirical results are described later in the paper. The panel that I estimate is unbalanced as a number of time series observations differ across countries. However, assuming that observations are missing randomly, consistency of the fixed effects estimator is not affected. The fixed effects model is estimated with OLS, which, given the assumed properties of residuals, is the best linear unbiased estimator (Hsiao, 1986).

VI. The Regression Results

I first estimate a "full" version of equation (1) with both the log of per capita GDP (GDPPC) and its squared value (GDPPC_S) included among the explanatory variables. This allows capturing a potential threshold effect in the relationship between income inequality and per capita GDP.

It is worth noting that up to now most of the attempts to test for a U-shaped relationship between income inequality and economic development have used cross-sectional regressions. This approach may be conceptually incorrect when studying the intertemporal relationship between income inequality and per capita income. If one wants to see whether inequality changes with economic development longitudinal data are needed (Deininger and Squire, 1998). Here the use of the panel data provides an obvious advantage over purely time-series or cross-sectional data.

However, to test for any kind of a quadratic relationship between income inequality and per capita GDP is not the main purpose of this study. Here I undertake an attempt to identify specific factors behind the changes in inequality in the transitional region. While economic growth represents a good aggregate measure of the economy's health since it reflects the outcome of multiple complex processes taking place at all levels of the economy, it alone does not seem to be a satisfactory explanation of the inequality pattern. That is why in equation (1) I also introduce other potential explanatory forces into play. In addition, I estimate an alternative specification of equation (1) where I include only GDPPC (but not GDPPC_S) among the regressors to test for the linear relationship between income inequality and per capita GDP. Finally, since some of the explanatory variables in the model appear to be significantly correlated, I try several alternative specifications to investigate the robustness of the parameter estimates.[27] The regression results from estimating different modifications of equation (1) are reported in Table 2.[28]

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[27] I note here that the test of multicollinearity for the linear model indicates that the highest condition index equals 4.17, which is below a cutoff point of 10 suggested in the literature. Hence, I do not have a problem of severe collinearity in the estimation of the model.

[28] To detect outliers and influential cases I have conducted influence diagnostics such as the studentized residuals, the "hat" matrix, the COVRATIO statistic, DFFITS and DFBETAS (Belsley et al., 1980; Bollen and Jackman, 1985). I then deleted those observations that were detected influential by at least 3 tests. These observations turned out to be Moldova (1990), Russia (1998), Tajikistan (1998), and Uzbekistan (1989). However, when the model is estimated with these outliers included the results are very similar to those reported in Table 6.2.

Table 2. Fixed-effects Estimates from the Regression of the Gini Coefficient on Selected Explanatory Variables: All countries

Explanatory Variable	Full Model (quadratic relationship)	Full Model (linear relationship)	Reduced Model 1 (quadratic relationship)	Reduced Model 2 (quadratic relationship)
1	2	3	4	5
Intercept	-	-	-	-
GDPPC	- 4.190***	- 0.157**	-3.650***	- 4.912***
	(1.084)	*(0.063)*	*(1.131)*	*(1.208)*
GDPPC_S	0.239***	-	0.203***	0.272***
	(0.064)		*(0.067)*	*(0.072)*
INFL	0.033***	0.034***	0.024*	-
	(0.013)	*(0.013)*	*(0.013)*	
UNEMP	-0.001	-0.023	0.035*	0.086***
	(0.021)	*(0.021)*	*(0.020)*	*(0.019)*
CONSG	-0.014	-0.023	-0.061	-0.090*
	(0.047)	*(0.050)*	*(0.048)*	*(0.052)*
INDVA	-0.288***	-0.350***	-0.375***	-
	(0.072)	*(0.074)*	*(0.073)*	
PRIVS	0.105***	0.091***	-	-
	(0.027)	*(0.028)*		
Number of countries	24	24	24	24
Number of observations	129	129	129	129
R^2 adj.	0.65	0.61	0.61	0.52
F-value	32.66	31.97	33.33	12.14
Estimated threshold level (PPP-adjusted GDP per capita, 1992 USD)	6405	-	8015	8337

Note: All variables are in the natural log form. Standard errors are presented in parentheses. The model is estimated in deviations from the group means. The Yule-Walker (iterated) method was used to correct for serial correlation and heteroskedasticity.

*- significance at 10% level; **- significance at 5% level; ***- significance at 1% level (two-tailed tests).

The regression results indicate a statistically strong relationship between income inequality and per capita GDP. There is stronger support for a quadratic rather than a linear relationship.[29] The parameter estimates on GDPPC and GDPPC_S indicate that the relationship between income inequality and economic growth depends on where a country stands in terms of its per capita GDP. More specifically, they suggest that for a country below (above) some threshold level of development the process of economic growth would be associated with falling (rising) income inequality. It is worth noting that many transition countries (mostly those in the FSU) had either already been below the estimated threshold level of per capita GDP (see Table 2) at the start of the transition, or slipped below this level as a result of an economic decline, and thus are expected to have negative relationship between growth and inequality. The linear specification indicates that a 10 percent decline in per capita income would increase the Gini coefficient at the mean by 0.48 percentage points (see Column 3, Table 2). That income inequality might increase during recessions was confirmed in a number of studies of the United States (Meier, 1973; Metcalf, 1969; Thurow, 1970). So, given that economic decline in many transition countries reached an unprecedented scale, the adverse changes in the distribution of income are not surprising.

[29] I have also estimated the model with GDPPC and GDPPC_S being the only explanatory variables. The parameter estimates (not shown here) on both terms are statistically significant at 1% and 5% levels respectively (negative for GDPPC and positive for GDPPC_S). The F-test statistics for their joint significance is 45.71 (significant at a 1% level). Adjusted R^2 for the model is equal to 0.46.

Inflation is found to increase income inequality, although the estimated coefficient in one of the specifications turns out to be only marginally significant.[30] The magnitude of the effect, however, does not appear to be large. A 10 per cent increase in inflation would raise inequality at the mean by at most 0.08 Gini points (see Table 2). I have also tested (using the LSDV specification) for a threshold effect of inflation as one would expect that it is only the inflation above a particular level that affects income distribution. Indeed, when I include among the explanatory variables in equation (1) dummies for inflation levels instead of INFL, a clear threshold effect is apparent. I find (see Column 2, Table 3) that hyperinflation (annual CPI exceeds five hundred percent) is associated with a 9.5 percent higher income inequality compared to the situation of the relative macroeconomic stability (annual inflation below 20 percent).

[30] The recent study by the World Bank (2000) has found that inflation *volatility* is also associated with distributional costs in the transitional region.

Table 3. Estimates from the Regression of the Gini Coefficient on Selected Explanatory Variables: All countries

Explanatory Variable	Model with inflation dummy (LSDV)	Model with CLI (LSDV)	Model with IPF (LSDV)	Model with war dummy (pooled regression)
1	2	3	4	5
Intercept	-	-	-	4.370***
				(0.365)
GDPPC	- 3.834***	- 2.993**	-4.005***	-
	(1.225)	*(1.217)*	*(1.230)*	
GDPPC_S	0.216***	0.162**	0.225***	-
	(0.073)	*(0.073)*	*(0.073)*	
Inflation > 500	0.091**	0.106***	0.089**	0.163***
dummy[1]	*(0.038)*	*(0.037*	*(0.038)*	*(0.036)*
UNEMP	0.007	-0.014	0.003	-0.002
	(0.025)	*(0.025)*	*(0.025)*	*(0.022)*
CONSG	-0.035	-0.044	-0.045	-0.049
	(0.054)	*(0.051)*	*(0.054)*	*(0.046)*
INDVA	-0.242***	-0.199**	-0.205**	-0.360***
	(0.081)	*(0.079)*	*(0.086)*	*(0.074)*
PRIVS	0.098***	0.020	0.097***	0.117***
	(0.032)	*(0.039)*	*(0.032)*	*(0.028)*
CLI	-	0.034***	-	
		(0.011)		
IPF	-	-	-0.016	
			(0.013)	
War dummy				0.125**
				(0.050)
Number of countries	24	24	24	24
Number of observations	129	129	129	129
R^2 adj.	0.999	0.999	0.999	0.61
F-value	4290.35	4573.36	4194.15	15.73
Estimated threshold level (PPP-adjusted GDP per capita, 1992 USD)	7255	10432	7293	-

Note: All variables are in the natural log form. Standard errors are presented in parentheses. Yule-Walker (iterated) method was used to correct for serial correlation and heteroskedasticity. [1] - excluded category is inflation < 20% annual. The coefficients on other categories (21-50, 51-100, 101-500) are not reported since they are not significant.

*- significance at 10% level; **- significance at 5% level; ***- significance at 1% level (two-tailed tests).

Unemployment rate does not seem to have an impact on inequality in the base model (see Column 2, Table 2). It is very likely, however, that the parameter estimate on unemployment is contaminated due to a high correlation of the unemployment rate with other indicators of structural changes, namely deindustrialization and privatization (see Table A3, Appendix). Indeed, when I eliminate PRIVS from the estimation, the coefficient on UNEMP becomes positive and marginally significant (see Column 4, Table 2). Finally, when INDVA and INFL are also omitted from the regression, the parameter estimate on UNEMP becomes statistically significant at a 1% level (see Column 5, Table 2). The magnitude of the coefficient suggests that a 1 percentage point increase in the unemployment rate at the mean would raise the Gini coefficient at the mean by 0.33 points.

The lack of robustness of the effect of unemployment on income inequality may also be due to the following factors. First, a likely inequality-increasing effect of growing unemployment can be counterbalanced by an increasing flow of unemployment benefits. Second, it is also possible that the increase in between-groups inequality stemming from larger unemployment could be offset by a more equal distribution of income amongst transfer recipients (Milanovic, 1999). Third, it has been argued that in developing countries the unemployment statistics can in fact reflect the extent of the informal sector employment and self-employment, which are generally associated with higher income inequality (Ferreira and Litchfield, 1998). Finally, the quality of the unemployment data can also be a simple and quite probable explanation of why the effect of unemployment is not very robust. These data, as was mentioned before, are mostly based on official unemployment records, which may severely underestimate not only the actual scope of unemployment, but also the changes in the rate of unemployment over time.[31] Thus, their use in the estimation may induce a substantial downward bias in the parameter estimate on unemployment.

With regard to government consumption, I do not find it to influence income distribution. Although the sign of the parameter estimate is as anticipated, the coefficient on CONSG is only marginally significant in one of the specifications (see Table 2). The *size* of government consumption (as a share of GDP) may have poor predictive power since the total effect of this factor on income distribution clearly depends on the composition of government expenditures and progressivity of taxes used to finance them. Also, the variation in the share of GDP devoted to public consumption is substantially higher across countries than over time (see Table A2, Appendix). The estimation method that relies on the intertemporal variation does not capture the potential effect of government consumption on the levels of income inequality *across* countries.

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[31] Unemployment data can also be a poor indicator of labor market conditions since in many transition countries adjustments in the labor market manifested themselves in underemployment

The parameter estimate on INDVA supports our hypothesis that deindustrialization increases inequality.[32] The coefficient from the full model (see Column 2, Table 2) suggests that a 10 percent decline in the share of the industrial sector would lead to a 0.88 percentage point increase in the Gini coefficient at the mean. The parameter estimate implies that in Ukraine, for example, where the share of industrial sector in total output dropped from 48% in 1989 to 34% in 1998, the Gini coefficient increased by 2.17 percentage points over the period due to this factor alone, thus explaining a third of the total increase in income inequality.

There is a statistically strong positive relationship between income inequality and the size of the private sector. The estimated coefficient suggests that a 10 percentage point increase in the share of the private sector in the economy would result in a 0.83 point increase in the Gini coefficient. The magnitude of the effect does not seem to be very large. However, if one bears in mind that transitional countries have witnessed a substantial growth of the private sector, it is clear that the rising private sector employment plays a crucial role in explaining the increase in income inequality.

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and non-payment of wages rather than in the shedding of labor.

[32] I note that when the model is estimated using the available observations on the industrial sector employment as a share of total employment (100 observations, 23 countries), instead of those on INDVA, the parameter estimate is nearly identical to the one on INDVA reported in Table 6.2. These estimation results are not reported here for brevity.

I next look at the effect of economic liberalization on income inequality using the Cumulative Liberalization Index as an explanatory variable in the regressions. The regression results (see Column 3, Table 3) indicate that the process of liberalization is associated with rising income inequality.[33] The CLI is by construction highly correlated with the structural indicators used in our analysis (see Table A3, Appendix). Nevertheless, the parameter estimate on CLI is significant at a 1% level even when all other variables are included. The coefficient on PRIVS, however, becomes insignificant in this case. The parameter estimate on CLI indicates that a 10 percent increase in the CLI at the mean would be associated with a 0.27 point increase in the Gini coefficient at the mean. The effect of economic liberalization on income distribution is highly robust to the model specification.

I next investigate the impact on the distribution of income of the factors outside economic domain. Transitional countries have made different progresses in the introduction of political rights and civil liberties during the transition, which makes it interesting to see whether the progress on the political freedom front has any implications for the distribution of income. To do this, I use the Index of Political Freedom (IPF) described above. It is worth noting that the IPF is highly correlated with some other variables (see Table A3, Appendix). For instance, it is

[33] This finding is in contrast to that reported in the World Bank (2000), where they use the same reform index as I do, but a smaller sample of countries (20) and a different estimation technique (a

positively associated with the progress in economic reforms and negatively with the occurrence of civil conflicts.[34] Noteworthy, the regression results do not indicate that the extent of political rights and civil liberties has an independent impact on income inequality (see Column 4, Table 3). That does not preclude, though, the possibility that the degree of political democracy may affect income inequality indirectly (Gradstein et al., 2001).

A number of countries in the transitional region experienced the periods of civil conflicts and wars over the last decade. Therefore, I also look at the impact of civil wars on income inequality by estimating the pooled regression with a dummy for civil conflicts constructed as discussed above.[35] Since this dummy is highly correlated with the level of per capita GDP (see Table A3, Appendix), I omit GDPPC and GDPPC_S from the estimation. The regression results (see Column 5, Table 3) indicate that civil conflicts are associated with a 13.3 percent rise in income inequality.[36]

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pooled regression rather than fixed effects).

[34] Note that the higher value of the IPF means *less* political rights and civil liberties.

[35] The fixed effects model cannot be estimated in this setting. Thus I perform a pooled regression. A pooled regression refers to an OLS regression with a single overall intercept.

[36] The civil war dummy remains significant at a 5% level even when GDPPC and GDPPC_S enter the regression. The estimated effect, however, is slightly lower in this case than the one reported.

VII. Sensitivity Analysis

In this section I investigate the robustness of my findings. I first check whether the results are sensitive to the definition of the dependent variable. Although most of the Gini coefficients in the data set are based on disposable income, there are several data points based on other welfare concepts (see Table A1 in the Appendix). As a first check of the robustness of the results described in previous section I perform the estimation using disposable income Gini coefficients only. Since several Gini indexes in our sample come from other sources than the main data series used (the WIDER database) (see Table A1 in the Appendix), I also examine the sensitivity of the findings to the omission of these observations. Finally, I estimate the model by including only those Gini coefficients that are based on the same welfare definition *within* countries.[37] The resulting parameter estimates are reported in Table 4.[38]

[37] Note that the estimation of the model in deviations from the country means requires that the Gini coefficients are as comparable as possible *within* countries over time but not necessarily across countries. For this reason the use of consumption-based Gini coefficients for Azerbaijan (see Table 6.A1) does not cause a problem since like is compared with like. The same applies to the disposable monetary income Gini coefficients for Romania and Macedonia.

[38] Unfortunately, I cannot test the robustness of our findings to the use of alternative measures of inequality, as only the Gini coefficients are available.

Table 4. Fixed-effects Estimates from the Regression of the Gini Coefficient on Selected Explanatory Variables: Robustness to the Definition of the Dependent Variable and the Choice of the Data Series

Explanatory Variable	Full Model *Disposable Income Gini coefficients only*	Full Model *The WIDER database Gini coefficients only*	Full Model *Gini coefficients based on the same definition within countries*
1	2	3	4
Intercept	-	-	-
GDPPC	- 6.623***	- 6.910***	- 5.727***
	(1.395)	*(1.243)*	*(1.293)*
GDPPC_S	0.388***	0.399***	0.334***
	(0.083)	*(0.073)*	*(0.077)*
INFL	0.032**	0.007	0.013
	(0.016)	*(0.013)*	*(0.015)*
UNEMP	0.023	0.025	0.019
	(0.025)	*(0.021)*	*(0.024)*
CONSG	-0.043	-0.079	-0.062
	(0.067)	*(0.051)*	*(0.060)*
INDVA	-0.179*	-0.287***	-0.205**
	(0.107)	*(0.069)*	*(0.101)*
PRIVS	0.133***	0.082***	0.113***
	(0.032)	*(0.029)*	*(0.029)*
Number of countries	18	21	23
Number of observations	100	110	113
R^2 adj.	0.58	0.66	0.55
F-value	18.44	29.48	18.35
Estimated threshold level (PPP-adjusted GDP per capita, 1992 USD)	5084	5758	5284

Note: All variables are in the natural log form. Standard errors are presented in parentheses. The model is estimated in deviations from the group means. The Yule-Walker (iterated) method was used to correct for serial correlation and heteroskedasticity.

*- significance at 10% level; **- significance at 5% level; ***- significance at 1% level (two-tailed tests).

Similarly to the results reported in Table 2 (quadratic specification), all estimations shown in Table 4 strongly support a U-shaped relationship between income inequality and per capita GDP. The estimated coefficients on GDPPC and

GDPPC_S are larger than their counterparts in Table 2.[39] Although the parameter estimates on INDVA and PRIVS differ somewhat from sample to sample (see Columns 2-4, Table 4), which is quite natural given the variation in the representation of countries and time periods across the samples, they are generally in line with those reported in Table 2. I note that the coefficient on inflation is significant in one case only (see Column 2, Table 4). In the light of our finding that there is a clear threshold effect in the relationship between inflation and inequality this result is not surprising, since the samples differ in the number of high-inflation observations.

I am aware of the literature that advises to make a regression-based additive adjustment of the Gini coefficients based on different concepts for the purpose of cross-country comparisons. This approach, however, hinges on a very strong assumption, namely that the differences in Gini coefficients based on different concepts are the same across countries and over time. For transition countries, this assumption clearly does not hold. For instance, the comparison of the consumption-based Gini coefficient with the disposable income-based one obtained from the *same* survey data indicate that the former is 2 percentage points *higher* than the latter in Poland, is of the same magnitude in Russia, and is 5 percentage points lower in Georgia (the World Bank, 2000). Thus, for these

[39] One may not conclude here on the quality of different data series, though, as the changes in the parameter estimates could be driven by exclusion of the observations for particular countries

countries such an adjustment can hardly be an improvement. In this situation the only solution is to use observations on inequality that are "as fully consistent as possible" (Atkinson and Brandolini, 2001).

That is what I attempt to do in this paper. It is important to note also that since I use the fixed effects estimation any adjustments to the Gini coefficients based on the *same* concept *within* countries would be cancelled out anyway. Moreover, when I estimate the pooled regression by including dummies for Gini coefficients based on different concepts I do not find those dummies to be significant.[40]

I also investigate the robustness of the results to the use of the random-effects rather than fixed-effects estimation technique. The regression results (see Table 5) indicate that the fixed-effects and random-effects estimates are generally very similar.

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and/or time periods.

[40] The comparison category is Gini coefficient based on the household per capita disposable income.

Table 5. Random-effects Estimates from the Regression of the Gini Coefficient on Selected Explanatory Variables: All countries

Explanatory Variable	Full Model (quadratic relationship)	Full Model (linear relationship)	Reduced Model 1 (quadratic relationship)	Reduced Model 2 (quadratic relationship)
1	2	3	4	5
Intercept	19.749***	6.638***	17.926***	20.524***
	(4.288)	*(0.464)*	*(4.218)*	*(4.682)*
GDPPC	- 3.427***	- 0.234***	-2.811***	- 3.673***
	(1.030)	*(0.046)*	*(1.011)*	*(1.115)*
GDPPC_S	0.190***	-	0.151***	0.197***
	(0.061)		*(0.060)*	*(0.066)*
INFL	0.027**	0.025*	0.021	-
	(0.013)	*(0.014)*	*(0.014)*	
UNEMP	-0.003	-0.017	0.027*	0.073***
	(0.020)	*(0.020)*	*(0.016)*	*(0.016)*
CONSG	-0.036	-0.057	-0.080*	-0.094*
	(0.047)	*(0.048)*	*(0.046)*	*(0.052)*
INDVA	-0.312***	-0.352***	-0.389***	-
	(0.073)	*(0.075)*	*(0.071)*	
PRIVS	0.080***	0.068**	-	-
	(0.027)	*(0.028)*		
Number of countries	24	24	24	24
Number of observations	129	129	129	129
R^2 adj.	0.64	0.62	0.63	0.53
Estimated threshold level (PPP-adjusted GDP per capita, 1992 USD)	8374	-	13422	11260

Note: All variables are in the natural log form. Standard errors are presented in parentheses.
*- significance at 10% level; **- significance at 5% level; ***- significance at 1% level (two-tailed tests).

I next verify to what extent the results could be driven by observations for a particular time period. The parameter estimates are found to be fairly robust to the removal of any single period from the estimation.[41] I note, however, that the coefficient on inflation shows to be insignificant when the model is estimated

[41] These results are not shown here, but available from the author upon request.

without data for 1993. That is not surprising since for most countries in the transitional region 1993 was the year of hyperinflation. Thus, the elimination of this year from the estimation is likely to substantially underestimate the impact of inflation on the distribution of income.

As the countries in the sample differ widely in their levels of development and growth experiences during the transition (despite being collectively referred to as transitional economies), it is necessary to investigate the robustness of the results to the regional coverage. I first test the robustness of the results with respect to countries by removing one country at a time. Although the values of the coefficients (not reported here) fluctuate slightly, their magnitudes and significance levels are largely in line with those reported.

In view of the countries' differences in institutional characteristics and macroeconomic performance during the transition, I then estimate the model separately for the FSU and EE regions. I do not argue that such a division of countries into sub-samples is perfect as countries within EE and the FSU regions are not homogenous, but it seems to be a natural choice in many respects. First, economic decline in EE was on average less profound and persistent than in the FSU. Second, income inequality in EE has increased much less than in the FSU. Third, in contrast to the FSU, most EE countries already had at least some rudimentary elements of the market economy (e.g., a private sector) before the transition, and were much more effective with reform implementation during the transition. Finally, social safety nets in EE during the transition are widely recognized to have been much stronger than in the FSU. The results of separate estimations for the FSU and EE are presented in Table 6.

Table 6. Fixed-effects Estimates from the Regression of the Gini Coefficient on Selected Explanatory Variables: EE versus the FSU

Explanatory Variable	Full Model *FSU (quadratic relationship)*	Full Model *EE (quadratic relationship)*	Full Model *FSU (linear relationship)*	Full Model *EE (linear relationship)*
1	2	3	4	5
Intercept	-	-	-	-
GDPPC	-2.766**	-3.424	-0.314***	0.252***
	(1.373)	*(2.679)*	*(0.095)*	*(0.089)*
GDPPC_S	0.146*	0.212	-	-
	(0.082)	*(0.154)*		
INFL	0.039**	-0.031	0.039**	-0.026
	(0.019)	*(0.020)*	*(0.019)*	*(0.020)*
UNEMP	-0.014	0.047*	-0.028	0.031
	(0.042)	*(0.024)*	*(0.042)*	*(0.021)*
CONSG	-0.006	-0.076	-0.021	-0.046
	(0.072)	*(0.061)*	*(0.073)*	*(0.055)*
INDVA	-0.301**	-0.233**	-0.339**	-0.203**
	(0.127)	*(0.094)*	*(0.129)*	*(0.091)*
PRIVS	0.060	0.100***	0.044	0.118***
	(0.051)	*(0.038)*	*(0.041)*	*(0.036)*
Number of countries	15	9	15	9
Number of observations	65	64	65	64
R^2 adj.	0.72	0.67	0.70	0.65
F-value	20.83	20.11	22.88	21.99
Estimated threshold level (PPP-adjusted GDP per capita, 1992 USD)	-	-	-	-

Note: All variables are in the natural log form. Standard errors are presented in parentheses. The model is estimated in deviations from the group means. The Yule-Walker (iterated) method was used to correct for serial correlation and heteroskedasticity.

*- significance at 10% level; **- significance at 5% level; ***- significance at 1% level (two-tailed tests).

A number of interesting observations can be made based on them. First, when the estimation is performed separately for the FSU and EE countries, a U-shaped relationship between income inequality and the level of economic development becomes less evident for the FSU and collapses completely for EE. The estimation of the re-specified model (excluding GDPPC_S to test for the linear relationship) indicates that inequality-development relationship is in fact linear within these regions. Moreover, the parameter estimate on GDPPC is negative for the FSU region, but positive for EE countries (see Columns 4-5, Table 6). These results are consistent with the found for the whole sample threshold effect in the relationship between income inequality and economic development. In fact, while the FSU countries mostly fell below the estimated threshold level of per capita GDP, EE countries were positioned mostly above the threshold. The estimated coefficients on GDPPC suggest that a 1000 USD increase in per capita GDP (constant 1992 prices) at the region-specific means would be associated with a 2.12 Gini point decrease in income inequality at the mean in the FSU and a 0.89 Gini point increase in income inequality at the mean in EE.

A question of great interest is what makes economic growth push income inequality in different directions in the FSU and EE? It could be that explanation lies in the institutional environment in which the growth takes place. For instance, the levels of rent seeking and corruption in the FSU have been much higher than in EE. That may be one explanation of why income inequality in the former region has risen despite a dramatic economic decline.[42] It is also important to note that while the FSU countries were experiencing economic decline over most of the transition decade, the EE countries were growing.[43] The impact of economic recessions and recoveries on income distribution is not necessarily symmetric.

Inflation is found to have a significant impact on income inequality in the FSU countries, but not in EE. This result clearly comes from much higher inflation in the former region. In contrast to the FSU countries, many of which experienced hyperinflation (with annual inflation measured in hundreds percent), most EE countries witnessed inflation rates relatively modest by transitional standards.

Deindustrialization appears to be associated with rising income inequality in both EE and the FSU regions. This finding is robust to the alternative model specifications. The estimated coefficients from the linear specifications (see Columns 4-5, Table 6) suggest that a 10 percent decline in the share of industrial sector in the economy would be associated with a 1.17 and 0.54 percentage point

[42] I do not have sufficient longitudinal data on the levels of corruption in transitional countries to run the fixed effects regression. However, a simple regression of the average Gini coefficient during the transition period on the corruption perception index (Freedom House, 2001) for 16 transition countries of EE and the FSU indicates that higher corruption is associated with larger income inequality (the parameter estimate is significant at a 7% level). The regression results are not shown here, but available from the author upon request.

increase in the Gini index at the region-specific means in the FSU and EE, respectively.

The parameter estimates on PRIVS suggest that the growing private sector had an inequality-increasing effect exclusively in EE countries. This result is robust to the model specification. For instance, when INDVA is omitted from the regression, the coefficient on PRIVS becomes even more significant for EE, but remains insignificant for the FSU. The differential impact of privatization in two regions is striking given that the private sector expanded markedly in all economies. It is clearly a look at what makes privatization processes in EE and the FSU different that may provide the explanation. For instance, one consequence of the growing private sector in EE was a significant increase in returns to education and a rise in wage disparities (the World Bank, 2000). Conversely, privatization in the FSU countries did not substantially raise educational premiums, probably because of the excess supply of highly skilled labor.

Unemployment is not found to affect income distribution. The parameter estimate is only marginally significant for EE in one specification (see Column 3, Table 6). I have also tried several other specifications of the model and the results (not shown here) generally suggest that the exclusion of at least one variable reflecting the structural change in the economy, such as PRIVS or INDVA, from

[43] This is reflected in the composition of our sample, where the majority of observations for the FSU and EE countries cover the periods of economic decline and economic recovery, respectively.

the linear model makes the coefficient on UNEMP significant (at a 10% level) for EE, but not for the FSU.

The share of government consumption in GDP does not appear to explain the distributional outcomes neither in EE nor in the FSU.

With regard to the impact of CLI and IPF on income distribution in EE and the FSU, the regression results suggest that economic liberalization was associated with rising income inequality in both regions (see Columns 2-3, Table 7).

Table 7. Estimates from the Regression of the Gini Coefficient on Selected Explanatory Variables: EE versus the FSU

Explanatory Variable	Model with CLI (LSDV) FSU	Model with CLI (LSDV) EE	Model with war dummy (pooled regression) FSU	Model with war dummy (pooled regression) EE
1	2	3	4	5
Intercept	-	-	4.440***	4.510***
			(0.488)	*(0.626)*
GDPPC	- 0.386***	0.243***	-	-
	(0.105)	*(0.096)*		
Inflation > 500 dummy[1]	0.122**	0.006	0.208***	-0.026
	(0.059)	*(0.051)*	*(0.047)*	*(0.059)*
UNEMP	-0.050	0.001	-0.005	0.047
	(0.041)	*(0.025)*	*(0.036)*	*(0.031)*
CONSG	-0.082	-0.046	-0.024	-0.209***
	(0.084)	*(0.055)*	*(0.078)*	*(0.058)*
INDVA	-0.222*	-0.201*	-0.376***	-0.286**
	(0.133)	*(0.103)*	*(0.102)*	*(0.136)*
PRIVS	-	-	0.132***	0.068*
			(0.050)	*(0.040)*
CLI	0.039***	0.041***	-	-
	(0.019)	*(0.008)*		
IPF	-	-	-	-
War dummy	-	-	0.112*	0.193***
			(0.067)	*(0.061)*
Number of countries	15	9	15	9
Number of	65	64	65	64

observations				
R^2 adj.	0.999	0.999	0.69	0.64
F-value	2705.14	8450.25	15.71	10.57
Estimated threshold level (PPP-adjusted GDP per capita, 1992 USD)	-	-	-	-

Note: All variables are in the natural log form. Standard errors are presented in parentheses. Yule-Walker (iterated) method was used to correct for serial correlation and heteroskedasticity. [1] - excluded category is inflation < 20% annual. The coefficients on other categories (21-50, 51-100, 101-500) are not reported since they are not significant.

*- significance at 10% level; **- significance at 5% level; ***- significance at 1% level (two-tailed tests).

The extent of political freedom, however, does not affect income inequality in either region.[44]

I also look at the impact of civil conflicts on income distribution separately for EE and the FSU regions using the pooled regressions. The regression results (see Columns 4-5, Table 7) indicate that the periods of civil wars were associated with rising income inequality in both regions.

Finally, it is worth noting that the results obtained from the separate estimations for EE and the FSU must be treated with caution due to the relatively small regional samples.

[44] The results are not reported here, but available from the author upon request. I have also estimated several alternative specifications with IPF, but the parameter estimate on IPF is found to be insignificant.

VIII. Conclusions

The main goal of this paper is to identify the factors that caused dramatic changes in income inequality in the transitional countries of EE and the FSU throughout the 1990s. The empirical analysis is performed using a unique panel of inequality estimates that cover 24 transitional countries over the period 1989-1998. The econometric approach employs panel data estimation methods.

I find support for a normal U-shaped relationship between income inequality and per capita GDP for the transitional region as a whole. It suggests that for a country below (above) some threshold level of development economic growth is associated with falling (rising) income inequality. Specifically, the relationship between income inequality and economic growth is shown to be negative for countries of EE, but positive for those of the FSU. The results suggest that economic recovery-promoting policies may certainly have an equalizing effect on income distribution in some transition countries. However, at least in the short-run, there can be a trade-off between economic growth and income inequality in other countries.

Although undoubtedly important, the relationship between income inequality and economic growth does not represent the main focus of this study. I have searched for specific economic factors and non-economic forces that determine the changes in income distribution during the transition.

The empirical results indicate that economic liberalization and structural adjustments are associated with rising income inequality. More specifically, I find that a 10 percent increase in the Cumulative Liberalization Index at the mean is associated with a 0.27 and 0.34 percentage point increase in the Gini coefficient at the region-specific means in the FSU and EE countries, respectively. Deindustrialization has a strong impact on income distribution in both regions. A 10 percent decline in the share of industrial sector in the economy is related to a 1.17 and 0.54 percentage point increase in the Gini coefficient at the mean in the FSU and EE countries, respectively. Although the economies of both EE and the FSU regions have been substantially privatized during the transition, the evidence suggests that a rapidly growing private sector has contributed to rising income inequality in EE countries only. A 10 percent growth in the share of private sector in the economy is associated with a 0.31 percentage point increase in the Gini coefficient at the mean in these countries.

It is important to note that some increase in income inequality due to structural reforms associated with the transition from centrally-planned to market economy is largely inevitable and should not be considered in the negative light. These reforms may have sizable longer-term rewards by strengthening incentives, creating new jobs and fostering economic growth. Ultimately, it is better to be unequally rich than equally poor. Nevertheless, the policies aimed at facilitating the transition of workers from the public to the private sector, and from the manufacturing sector to services may be of paramount importance for the distributional outcomes of the reforms.

Although there is some evidence to suggest that unemployment may be positively associated with income inequality, the effect is not robust to model specification.

The degree of government involvement in the economy through government consumption generally does not seem to have an impact on income distribution.

I find that hyperinflation makes the distribution of income more unequal. This finding may certainly contribute to the explanation of why income inequality in the FSU countries (most of which experienced hyperinflation at the start of the transition) increased much more than in EE countries, where inflation levels have been relatively moderate. The important policy implication of this finding is that macroeconomic stabilization not only fosters economic recovery, but is also beneficial in terms of distributional outcomes.

Finally, I have also investigated the role of some forces outside economic domain in determining income inequality in transition economies. The empirical evidence indicates that civil conflicts are associated with rising income inequality. On average, they lead to a 13.3 percent higher income inequality. The extent of political rights and civil liberties, measured by the Index of Political Freedom, is not found to affect income distribution. Nevertheless, this index is strongly correlated with the indicators of structural changes in the economy, suggesting that political rights and civil liberties are likely to affect income distribution indirectly.

To conclude, the avenue of research undertaken in this paper appears promising, for it reveals forces influencing income distribution in transitional countries. However, I certainly have not exhausted all factors explaining the dynamics of income inequality in the transitional region, and thus further research here may be beneficial.

Appendix

Table A1. Description of the Inequality Data Used in the Empirical Analysis

Region/ Country	Start	End	No. of obs.	Gini index (start)	Gini index (end)	Max. value (year)	Welfare measure	Area/ Population Coverage	Sample/ Reference unit	Source (Year)
1	2	3	4	5	6	7	8	9	10	12
I. FSU										
a) Baltic states										
Estonia	1989	1998	9	23.00	36.97	36.97 (98)	DI, GI (90)	All/All	HH/HH pc	WIID, BM (89)
Latvia	1989	1998	5	22.50	32.10	32.60 (97)	DI, GI (89)	All/All	HH/HH pc	WIID, BM (89)
Lithuania	1989	1999	6	22.50	34.00	35.04 (94)	DI, GI (89)	All/All	HH/HH pc	WIID, BM (89), WB (99)
b) Western CIS										
Belarus	1989	1999	5	22.80	26.00	26.00 (99)	DI, GI (89)	All/All	HH/HH pc	WIID, BM (89), WB (99)
Moldova	1993	1997	2	36.50	42.00	42.00 (97)	DI, GI (93)	All/All	HH/HH pc	BM (93), WB (97)
Russia	1989	1996	5	23.80	37.83	37.83 (96)	GI	All/All	HH/HH pc	WIID, BM (89)
Ukraine	1989	1999	6	25.80	32.00	32.00 (99)	DI, GI (97)	All/All	HH/HH pc	WIID, WB (99)
c) Caucasus										
Armenia	1990	1998	6	26.90	59.00	62.14 (95)	DI, GI (90)	All/All	HH/HH pc	WIID, WB (98)
Azerbaijan	1995	1999	2	44.00	43.00	44.00 (95)	CS	All/All	HH/HH pc	WB
Georgia	1989	1997	4	31.30	51.86	58.71 (96)	DI, GI (88,90)	All/All	HH/HH pc	WIID
d) Central Asia										
Kazakhstan	1990	1996	4	29.70	35.00	35.00 (96)	DI, GI (90,93)	All/All	HH/HH pc	WIID, WB (96)
Kyrgyz Republic	1990	1997	3	30.80	47.00	55.30 (93)	DI, GI (90)	All/All	HH/HH pc	WIID, BM (93), WB (97)
Tajikistan	1989	1990	2	31.80	33.40	33.40 (90)	GI	All/All	HH/HH pc	WIID
Turkmenistan	1989	1993	3	31.60	35.80	35.80 (93)	DI, GI (89,90)	All/All	HH/HH pc	WIID
Uzbekistan	1990	1994	3	31.50	33.00	33.30 (93)	DI, GI (90,94)	All/All	HH/HH pc	WIID
II. Central EE										
Czech Republic	1989	1997	9	19.36	27.64	28.14 (96)	DI	All/All	HH/HH pc	WIID
Hungary	1989	1997	7	21.41	24.58	24.58 (97)	DI	All/All	HH/HH pc	WIID
Poland	1989	1998	9	25.05	32.00	34.20 (97)	DI	All/All	HH/HH pc	WIID, WB (98)
Slovak Rep.	1989	1997	9	18.06	23.36	24.83 (96)	DI	All/All	HH/HH pc	WIID

1	2	3	4	5	6	7	8	9	10	12
III. South EE										
Bulgaria	1989	1997	8	24.47	34.59	34.78 (96)	DI, GI (89,91)	All/All	HH/HH pc, HH (89,98)	WIID
Romania	1989	1997	9	23.24	30.27	31.18 (95)	DI	All/All	HH/HH pc	WIID
IV. FY										
Croatia	1989	1998	4	25.10	33.30	33.30 (98)	DI	All/All	HH/HH pc	WIID
Macedonia	1990	1997	5	34.90	36.65	36.94 (96)	DI	All/All	HH/HH pc	WIID
Slovenia	1991	1998	4	22.71	25.00	25.05 (93)	DI	All/All	HH/HH pc	WIID, WB (98)

Note: UNU/WIDER-UNDP World Income Inequality Database, Version 1.0, September 12, 2000 provides further reference on the source and estimation methodology for each data point drawn from this database. The data by Branko Milanovic are from the Appendix 4 "The Original Income Distribution Statistics" of his book *Income, Inequality and Poverty during the Transition from Planned to Market Economy* (1998). The data by the World Bank are taken from the Appendix D "Poverty and Inequality Tables" of the book *Making Transition Work for Everyone: Poverty and Inequality in Europe and Central Asia* (the World Bank, 2000).

Disposable income (DI) is equal to *gross income* (GI) minus payroll and direct personal income taxes (PIT). Gross income consists of earnings from labor, cash social transfers, self-employment income, other income (gifts, income from property) and in-kind consumption (for instance, agricultural products grown on a household's plot of land). It is argued (Milanovic, 1998) that the difference between gross and disposable incomes is negligible for transition countries (especially for the pre-transition period) as gross income already excludes payroll taxes withdrawn at the source, and PIT is minimal (less than one percent of gross income). That allows one to use the Gini coefficients based on gross incomes as the benchmarks for the levels of income inequality observed before the transition (mostly for the FSU countries, for which the pre-transition disposable income Gini indexes are not available). It is very important to have these pre-transition observations in the sample since the evidence suggests that most of the variation in income inequality over time has taken place over the initial period of transition and economic collapse. As the first transition-period surveys were often conducted a few years into the transition process, by taking the estimates of inequality derived solely from these surveys one would significantly underestimate the changes in inequality over time (which is what I want to explain). The Gini coefficients for Romania and Macedonia are based on disposable monetary income, which does not include in-kind consumption. These Gini coefficients are likely to overestimate the *levels* of inequality, but not the *changes* in inequality. Note that two data points (Azerbaijan, 1995, 1999) are Gini coefficients based on *consumption* (CS). These observations are used due to the lack of alternatives. They are not found to influence the overall results. The Gini index for 1988 is used in the absence of 1989 data.

The data coming from the Family Budget Surveys (FBS) (mostly 1989 data in our sample) are not completely representative and may underestimate inequality as FBS excluded pensioner-headed households and households headed by the unemployed. However, the estimates of inequality obtained from transition-year surveys can also be downward-biased due to decreased response rates among the rich, inadequate coverage of informal sector incomes, etc. (for a detailed discussion of these and other data issues see Milanovic, 1998). It is not clear, though, how all these biases would, on the net, affect the *changes* in inequality. In any case, there is not much that one can do about these sorts of problems except trying to use observations that are "as fully consistent as possible" (Atkinson and Brandolini, 2001). That was a guiding principle in the compilation of the data.

Table A2. Descriptive Statistics and Variance Decomposition

Variable	Mean	Min	Max	Std. Dev.		
				Overall	Between	Within
GINI	30.5	17.8	62.1	9.21	7.87	5.14
GDPPC	6322.4	1649.9	13764.9	2789.40	2680.60	1191.45
INFL	247.7	-7.6	9750.0	945.22	669.02	793.09
UNEMP	7.9	0.0	38.8	7.39	6.84	3.89
CONSG	18.0	6.0	27.4	4.96	4.27	2.90
INDVA	38.2	15.3	67.4	8.71	5.30	7.02
PRIVS	38.4	5.0	75.0	21.44	13.64	17.55
CLI	2.6	0.0	7.7	2.11	1.21	1.81
IPF	3.4	1.5	7.0	1.60	1.44	0.89

Source: Author’s calculations.

Note: The number of observations for all variables is 129. The overall and within (over time) standard deviations are calculated over all 129 observations. The between (across countries) standard deviation is calculated over the means for 24 countries.

Key:

GDPPC = GDP per capita, PPP-adjusted USD in 1992 prices;
INFL = Inflation rate, measured by the year-on-year change in CPI (percent);
UNEMP = Share of unemployed in total labor force (percent);
CONSG = Government consumption as share of GDP (percent);
INDVA = Share of industry value added in GDP (percent);
PRIVS = Private sector share in GDP (percent);
CLI = Cumulative Liberalization Index (score);
IPF = Index of Political Freedom (score).

Table A3. The Matrix of the Pearson Correlation Coefficients

	GINI	GDPPC	INFL	UNEMP	CONSG	INDVA	PRIVS	WAR_D	CLI	IPF
GINI	1.00	-	-	-	-	-	-	-	-	-
GDPPC	-0.75 (0.00)	1.00	-	-	-	-	-	-	-	-
INFL	0.15 (0.10)	-0.25 (0.01)	1.00	-	-	-	-	-	-	-
UNEMP	0.26 (0.00)	-0.12 (0.18)	-0.05 (0.55)	1.00	-	-	-	-	-	-
CONSG	-0.24 (0.01)	0.14 (0.11)	-0.04 (0.68)	-0.01 (0.99)	1.00	-	-	-	-	-
INDVA	-0.62 (0.00)	0.41 (0.00)	0.14 (0.10)	-0.42 (0.00)	0.05 (0.58)	1.00	-	-	-	-
PRIVS	0.30 (0.00)	-0.11 (0.22)	-0.22 (0.01)	0.52 (0.00)	-0.12 (0.17)	-0.51 (0.00)	1.00	-	-	-
WAR_D	0.56 (0.00)	-0.43 (0.00)	0.35 (0.00)	0.09 (0.30)	-0.27 (0.00)	-0.29 (0.00)	-0.02 (0.86)	1.00	-	-
CLI	0.25 (0.00)	0.04 (0.68)	-0.27 (0.00)	0.68 (0.00)	0.04 (0.63)	-0.51 (0.00)	0.81 (0.00)	-0.05 (0.56)	1.00	-
IPF	0.11 (0.20)	-0.37 (0.00)	0.27 (0.00)	-0.49 (0.00)	-0.11 (0.22)	0.21 (0.02)	-0.57 (0.00)	0.31 (0.00)	-0.59 (0.00)	1.00

Note: All variables except WAR_D, CLI and IPF are in the natural log form. ln(1+INFL/100) and ln(1+UNEMP) are used for respectively INFL and UNEMP. WAR_D is a dummy variable equal to 1 for each year since an internal conflict has taken place in a given country. The values in parentheses indicate Prob. > | r | under H_0: Rho = 0; (0.00) means that the p-value is less than 0.005.

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