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IZA DP No. 4233

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June 2009

Forschungsinstitut
zur Zukunft der Arbeit
Institute for the Study
of Labor

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ABSTRACT

Inequality and Volatility Moderation in Russia: Evidence from Micro-Level Panel Data on Consumption and Income[*]

We construct key household and individual economic variables using a panel micro data set from the Russia Longitudinal Monitoring Survey (RLMS) for 1994-2005. We analyze cross-sectional income and consumption inequality and find that inequality decreased during the 2000-2005 economic recovery. The decrease appears to be driven by falling volatility of transitory income shocks. The response of consumption to permanent and transitory income shocks becomes weaker later in the sample, consistent with greater self-insurance against permanent shocks and greater smoothing of transitory shocks. Comparisons of RLMS data with official macroeconomic statistics reveal that national accounts may underestimate the extent of unofficial economic activity, and that the official consumer price index may overstate inflation and be prone to quality bias.

JEL Classification: E20, J30, I30, O15, P20

Keywords: inequality, income, consumption, transition, Russia

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[*] We thank Steven Stillman and conference and seminar participants at IZA, the New Economic School, and the PIER IGIER Fifth International Conference for useful comments.

1. Introduction

Modern macroeconomists are increasingly relying on the analysis of environments with heterogeneous agents. Many macroeconomic questions can only be asked (and answered) in the context of multi-agent environments. These richer macroeconomic models require a correspondingly rich set of empirical facts that come from micro data and incorporate information on distributions in addition to the usual aggregates. The goal of this paper is to provide a comprehensive set of cross-sectional and time series stylized facts for the Russian economy and a systematic study of multiple dimensions of inequality.

Since the late 1980s, Russian economy has been subject to substantial macroeconomic volatility, with a long phase of severe output contraction, periods of high and variable inflation, and a subsequent period of recovery. At the same time, Russia has tremendous regional diversity. The combination of these factors presents unique opportunities for studying both cross-sectional and time-varying dimensions of inequality. Fortunately, high quality data are available to explore these opportunities: a large, nationally representative panel study of Russian households that incorporates economic variables, the Russia Longitudinal Monitoring Survey (RLMS).

This paper includes multiple dimensions of inequality, with particular focus on consumption and income. We construct the key variables describing the economic behavior of Russian households and individuals and analyze their cross-sectional dispersion and time series patterns. Specifically, we create time-varying distributions of individual earnings and labor supply, as well as household-level income, expenditure, and consumption.

We would like to highlight two main results. First, almost all measures of cross-sectional inequality in income and consumption started falling during 2000-2005, after staying relatively high during 1994-1998. Second, the measured fall in inequality is mostly due to the moderation of the transitory shocks to household income and consumption.

The recent period of falling inequality was preceded by an initial rise in the early 1990s that accompanied Russia's transition from a centrally planned to market economy (e.g., Commander *et al* 1999, Galbraith *et al* 2004). However, the level of inequality at the end of our sample is still higher than it was during the socialist era. Interestingly, poor households do not appear to fall behind during the economic recovery – the lower tail of the expenditure distribution does not diverge from the middle as the economy expands. The latest level of inequality that we find is typical for a middle income country. For example, the Gini coefficient in 2005 was about 0.38-0.40, which is just slightly above the mean value of Gini coefficients for after-tax household income and consumption from upper middle income countries.[1,2]

Some features that set the Russian economy apart from more developed countries turn out to be important for the analysis of inequality. One such feature is home production of food. Our results indicate that home-grown food has a large equalizing effect on income and consumption. The effect is large, because poorer rural households are also the ones that grow a lot of food for own consumption. Another unique feature of the Russian economy is its geographic diversity. Accounting for regional differences in the cost of living (that vary by a factor of 2.7 in Russia) is shown to have a sizeable equalizing effect. Other important features of the Russian transition, such as underreporting of income, wage payment delays, irregularities in government transfer payments, and forced in-kind substitutes in lieu of wage payments also explain some of the inequality trends.

The comparison of income and expenditure inequality reveals further differences from developed economies. In developed economies, expenditures are usually distributed more equally than income, which is attributed to consumption smoothing possibilities. This turns out not to be the case for Russia, where expenditure inequality is almost as high as income inequality. We argue that the relatively high expenditure inequality reflected peculiar patterns of consumption smoothing during the downturn. Households facing irregular wage and transfer

[1] Our results on inequality levels have to be taken in the context of our sample. We think that the RLMS, like most household surveys, may under-represent the very rich individuals who own capital assets in Russia. This is evident from the negligible financial asset holdings of most RLMS respondents. The studies that attempt to adjust for super-rich typically document much higher levels of inequality. For example, Guriev and Rachinsky (2006) find that the income Gini coefficient for the city of Moscow is 0.625, and Aivazian and Kolenikov (2001) report a Gini coefficient of 0.55-0.57 based on parametric estimation of the uncensored expenditure distribution. We find some evidence that suggests divergence between the super-rich and the rest of the population in 2003-2005 (see Section 2 for further discussion).

[2] The comparisons are made using the Inequality Database of the World Institute for Development Economics Research.

payments, high inflation, and undeveloped financial markets used less conventional mechanisms such as food storage to smooth consumption. Food inventories were built up when income was received to insure against irregular wage payments.

We further look at the inequality dynamics between groups in our sample. We find the comparison of economic experience between urban and rural populations is particularly interesting. The rural population has a more restricted choice of jobs, which limits occupational mobility during transition. In addition, the workers with highest earnings potential might have migrated to cities. However, we do not find evidence that income and consumption of the rural population fell behind. The rural group did not seem to do relatively worse during the downturn, although during the recovery the rural population exhibited a slower growth rate in consumption of durables.

More broadly, we have found almost no evidence of convergence or divergence between groups based on observables, such as education, location, household composition, and age. The reduction in inequality during economic recovery resulted mostly from the moderation in the residual volatility of income and consumption growth.

We examine the reasons for the observed fall in residual income volatility by exploiting the panel dimensions of the data. In particular, we decompose the income process into permanent and transitory components and estimate their effect on consumption. We document that the fall in residual income volatility is mostly due to a fall in the variance of transitory income shocks.[3] Over time, consumption response to both permanent and transitory income components becomes weaker. This is consistent with better insurance against income shocks and hence better consumption smoothing later in the mid 2000s.

Apart from the analysis of inequality trends, we examine the trends in the levels of income and expenditure. The time trends show a 40 percent drop in real per-capita expenditure and a 50 percent drop in real hourly wages during 1994-1998. Recent literature has argued that the drop in Russian real output during the transition has been overstated due to exaggeration of the Soviet output and mismeasurement of the unofficial economy in the 1990s (Schleifer and

[3] Stillman (2001) finds that RLMS expenditures respond strongly to transitory shocks during 1994-1998.

Treisman 2005) or due to overstatement of inflation by the official CPI (Gibson *et al* 2004). To detect possible sources of CPI bias, we examine food prices and quantities from RLMS and find that the composition of food consumption has not changed much. Consequently, the CPI substitution bias within the food category is likely to be small. We do find, however, substantial disagreement in food CPI between RLMS and National Income and Product Accounts (NIPA), with a 25 percent discrepancy in the cumulative inflation during 1994-1998, but not much discrepancy afterwards. In addition, there is evidence of quality bias in the official CPI.

The paper's goal of documenting a comprehensive set of macroeconomic facts for Russia links it to many bodies of literature in macroeconomics, labor economics, development economics, and transition economics. In the interest of space, the literature survey below is necessarily incomplete, and it merely catalogues some of the related recent work by topic. Our analysis is closely related to the growing empirical literature that analyzes the joint evolution of income and consumption distributions (Cutler and Katz 1992, Attanasio and Davis 1996, Blundell and Preston 1998, Slesnick 2001, Krueger and Perri 2006, Heathcote *et al* 2007, Blundell *et al* 2008, etc.). There is also a growing body of research on inequality in developing countries. We find this literature particularly relevant for our study as it emphasizes the importance of measurement issues, urban-rural differences, home production, and income underreporting in understanding inequality in developing countries (e.g., Chen and Ravallion 1996, Deaton 1997).

Several papers document changes in income inequality in Russia in the 1990s. These studies establish a number of important facts for the early transition period: rising income inequality, significant income mobility, large regional variation, and insufficient government transfers to offset an increase in wage inequality (Commander *et al* 1999; Milanovic 1999; Flemming and Micklewright 2000). The rise in income inequality is mainly attributed to compositional shifts from the old state sector to the new private sector, liberalization of wage setting, liberalization of prices and trade, and macroeconomic volatility. Some studies argue in favor of inequality measures based on expenditures (Aivazyan and Kolennikov 1999, Jovanovic 2001). They find a significant share of the transitory component in shocks to expenditures, high instability, and a slight downward trend in expenditure-based inequality. Our findings are in general agreement with these studies. We extend previous analyses in a number of ways. We consider a longer time span covering recent years, provide a variety of measures and decompositions of inequality, investigate sources of inequality, and examine the co-movements between income and consumption using the panel aspects of the data.

The rest of the paper is organized as follows. In Section 2, we describe the data, provide basic information on the levels of consumption, income and labor market participation, and compare these statistics with official data. In Section 3, we document the trends in inequality in individual labor market outcomes over 1994-2005. In Section 4, we construct and report consistent time series for a variety of measures of consumption and income inequality at the household level. Section 5 decomposes the income process into transitory and permanent components and investigates the interaction of consumption and income inequality at the household level. In Section 6, we examine the role of regional disparities and dispersion of prices in generating inequality and discuss the possible sources of CPI bias. Our concluding remarks are in Section 7.

2. Data Overview

Sample and variables

The analysis in this paper uses the RLMS, which is a panel dataset that includes detailed information on measures of income, consumption, household demographics, and labor supply. The RLMS is organized by the Population Center at the University of North Carolina in cooperation with the Russian Academy of Sociology. The data are collected annually, and our panel includes 10 waves during the period 1994-2005, with the exception of 1997 and 1999, when the survey was not administered.[4] There were approximately 8,343-10,670 individuals who completed the adult (age 14 and over) questionnaire and 3,750-4,718 households who

[4] In all plots except Figure 2, the 1997 and 1999 values are 2 point linear interpolations of the data points in adjacent years.

completed the household questionnaire in each round. These individuals and households reside in 32 oblasts (regions) and 7 federal districts of the Russian Federation.[5]

The RLMS sample is a multi-stage probability sample of dwellings. The response rate is relatively high: it exceeds 80% for households and about 97 percent for individuals within the households. The sample attrition is generally low compared to similar panel surveys in other countries, partly owing to lower mobility and infrequent changes of residences.[6] To account for the panel attrition, all statistics reported in this study are weighted using the RLMS sample weights that adjust not only for sample design factors but also for deviations from the census characteristics. For comparability with other countries in this volume, we restrict our estimation sample to households in which at least one individual is 25-60 years old. Appendix 3 shows the size and composition of the estimation sample.

The variables employed in our study are carefully constructed and made not only internally comparable across different waves but also externally consistent with standard variable definitions in macroeconomic literature. We provide thorough treatment of missing values, influential observations, non-response, and other common problems of micro data. We also take into account important Russia-specific phenomena that influence our variable definition and data analysis such as wage payment delays in the 1990s, production of food at home, high regional diversity in cost of living, as well as peculiarities of the transition to a market economy. The detailed procedures of variable construction are documented in Appendix 1.

Economic conditions

Economic conditions in Russia affect our interpretation of income and consumption data in important ways. During the 1994-2005 period, Russia continued its transformation from a centrally planned system into a market economy. New integrated markets have emerged and new institutions of private ownership and property rights have been established.

[5] Russia had 89 regions and 7 federal districts as of December 1, 2005. The RLMS sample consists of 38 randomly selected primary sample units (municipalities) that are representative of the whole country.

[6] To deal with attrition, RLMS replenishes its sample on a regular basis by adding new dwellings, especially in the areas of high mobility such as Moscow and other large cities. To maintain the panel, RLMS partially attempts to collect information on those who moved out of the sample dwellings but live in the same location. More details on sample design, attrition, and replenishment are available at http://www.cpc.unc.edu/projects/rlms.

This transition to a market economy was accompanied by extreme macroeconomic disturbances, both real and nominal. Our sample period features two distinct phases: the downturn in 1994-1998 and the post-1998 period of rapid recovery. Panel A of Figure 1 shows that the early 1990s, following price liberalization in 1992, was a period of hyper-inflation. The end-year inflation rate in 1994 was 214 percent. The 1998 inflation spike (84 percent) corresponds to the government default on sovereign debt and the abrupt devaluation of the national currency, the ruble. In the downturn, real per-capita income and expenditures fell by about 40 percent (see panels B-D). Employee compensation and public transfers were paid irregularly, and were delayed by 3 to 5 months, on average. In the recovery phase, real per-capita income and expenditure growth was around 9 percent annually, and inflation stayed relatively low (10 to 20 percent).

Composition of income

The composition of household income during the sample period remained relatively stable, although there are important differences with Western industrialized economies. Panel B of Figure 1 compares four different measures of household after-tax monthly income during 1994-2005: (1) yL = labor income, (2) $yL+$ = net private transfers + yL, (3) y = capital income + $yL+$, and (4) yD = public transfers + y. Labor income, yL, is by far the largest income source; it accounts for 82 percent of household after-tax disposable income, yD, on average. In addition to labor income, yD includes income derived from financial assets, net private transfers, and public transfers. Income derived from financial assets is negligible; there is only a tiny difference between y and $yL+$. Net private transfers are contributions in money and in kind received from friends, relatives, and charitable organizations minus contributions given to individuals outside the household unit. Although net private transfers should not (and do not) affect average disposable income (the gap between $yL+$ and yL is negligible), gross private transfers are significant: private transfers received amount to 9 percent of disposable income, making them a potentially important channel of risk-sharing. Average public transfers are also large and amount to 13 percent of disposable income. The share of public transfers has increased after 2001, as evidenced by the growing gap between *yD* and *y* in panel B.

Composition of expenditures

Household consumption is constructed from numerous disaggregated categories of expenditures. Non-durable items, *c*, include 50 subcategories of food at home and away from home, alcoholic and non-alcoholic beverages, tobacco products, expenses on clothing and footwear, gasoline and other fuel expenses, rents and utilities, and 15-20 subcategories of services such as transportation, repair, health care services, education, entertainment, recreation, insurance, etc. Durable consumption is based on purchases of durable items within the last 3 months. All consumption measures are converted to a monthly base. To keep the coverage of consumption consistent across years, we exclude expenditure categories that became available only in recent years, such as washing supplies, personal hygiene items, books, sporting equipment, internet, and wireless phone services.[7]

Food is the biggest expenditure category for most households. The share of food purchases in aggregate non-durable expenditures starts from a high of nearly 70 percent in 1994 and gradually falls to 49 percent in 2005 (see also Figure 1C). One peculiar feature of Russian households is that many of them grow agricultural products on their subsidiary plots for own consumption. In 1994, about 10 percent of total food consumption (by market value) was home-grown, and by 2005 the share of food grown at home fell to 5 percent (see also Figure 1C). Despite declining in aggregate importance, home production of food significantly affects measures of inequality, because it is concentrated among the poorer rural households (see Section 4).

The share of durables was around 14 percent of aggregate expenditures, *cD*, during 1994-2002, but has increased significantly after 2003 (see also Figure 1D). Expenditures on durables

[7] The share of excluded expenditure categories is about 3% of total consumption expenditures in 2001-2004 and 5% in 2005. The 2 percentage point increase in 2005 is explained by adding expenditures on internet and cell phones in the 2005 RLMS questionnaire. The omitted expenditure categories do not affect the measures of consumption inequality.

tend to be concentrated at high income levels. 76 percent of households report no durable purchases within the last 3 months.

For 2000-2005, our dataset has a self-reported market value of owner-occupied housing. If we take the annual housing services flow to be 5 percent of its market value, the share of owner-occupied housing will equal roughly 11 percent of total consumption, *cD+*. The share of housing consumption is relatively stable over time because the aggregate market value of housing is growing at roughly the same rate as aggregate expenditures, *cD* (see Figure 1D).

Income underreporting

Two data facts lead us to believe that the aggregate income obtained from RLMS is likely to be underestimated. The first fact is the negligible share of capital income. This could be due to income underreporting but also due to the underrepresentation of the very rich individuals in the RLMS. To get a sense of the underestimated capital income, we can take the estimate of personal wealth of Russian billionaires and millionaires (1.4 times national GDP) from Guriev and Rachinsky (2006) and multiply it by a conservative rate of return on wealth (6 percent). If this is correct, the super-rich should earn about 8.4 percent of GDP, which we miss in our data.

The second fact is that for those who are in the sample, expenditures are consistently above reported income throughout the whole period (Figure 1D). This gap cannot be attributed to dissaving, as most households have negligible stocks of financial assets. We believe that income is under-reported because of tax evasion. For example, Gorodnichenko *et al* (2009) studied the gap between consumption and income in the RLMS data set and found the gap to be significantly larger in districts where respondents believed that other people do not pay their taxes. Over time, the gap between consumption and income seems to narrow, and the narrower gap may correspond to the effect of the 2001 tax reform, credit market development, and other factors (see Gorodnichenko *et al* 2009).

Since we do not have an independent estimate of the extent of income under-reporting (except the consumption-income gap itself), it seems to be more informative to compare expenditure, rather than income, levels between RLMS and other data sources.

Comparison with national accounts

We first compare income and expenditure levels between RLMS and official National Income and Products Accounts (NIPA). To make comparisons with national statistics, one must be careful about using compatible data definitions. The RLMS measure of household disposable income (*yD*) is after taxes and transfers *given*, and it excludes in-kind consumption, such as owner-occupied housing and home-grown food. The corresponding NIPA measure is disposable income for the "household account" after taxes and transfers minus in-kind consumption (Goskomstat 2007a). Similarly, the RLMS measure of consumption that we select for comparison purposes (*cD*) corresponds to the NIPA measure of household final consumption expenditures on durable and non-durable goods and services without imputed in-kind expenditures (Goskomstat 2007a). For comparability purposes, we use the full unrestricted sample.

Panels A and B of Figure 2 compare *yD* and *cD* (in per capita terms) with their counterparts from NIPA. Consumer expenditures in RLMS and NIPA are close during most of the sample period[8], while reported disposable income in RLMS is up to 30 percent lower than the official figures. The big discrepancy in income levels across the two sources is expected, since NIPA expenditure and income data are internally consistent and adjusted for under-reporting,[9] and RLMS reported income is much lower than expenditures. This comparison supports income under-reporting as a possible explanation for the consumption-income gap in the RLMS and also points to expenditure data as potentially more informative about the level variables.

The close agreement between RLMS and NIPA expenditure numbers in panel B contrasts sharply with similar comparisons for the U.S. where household surveys tend to underestimate national aggregates by more than 30 percent. The analogous comparisons for the UK produce a

[8] The 1998 discrepancy can be explained by the fact that RLMS has been conducted just right after the August financial crisis while NIPA's numbers are averaged over the year.

[9] NIPA eliminates the discrepancy between reported income and consumption by construction. Disposable income is constructed as a sum of household aggregate expenditures and savings, and the difference between imputed disposable income and the officially reported income is included in the income accounts as unobserved labor compensation.

less significant discrepancy of 5 percent (Attanasio *et al* 2004). Our finding is somewhat surprising, since RLMS likely under-represents the very rich households that consume out of capital income. However, it is possible that that the official statistics make an insufficient adjustment for shadow economic activity, so the discrepancy between NIPA and RLMS expenditures is not as large as one may have expected.

Starting in 2003, RLMS consumption expenditures show slower growth than NIPA expenditures. As explained above, this difference in trends may indicate the growing gap between the RLMS sample and the super-rich individuals. Part of the gap may also be due to an upward trend in consumption of goods that RLMS data does not consistently track, such as internet and cell phone services. However, new consumption categories added to RLMS over the years account for at most 5 percent of aggregate expenditures, and their growth is not enough to account for the difference in trends after 2003. Finally, a small portion of the gap (up to 1.6 percent of aggregate expenditures per capita) can be explained by the replacement of one of the wealthiest oil-based regions in the North by the middle income region in Siberia in the 2003 RLMS sample (this was the only episode of regional sample replacement during the 1994-2005 period).

Comparison with the Household Budget Survey

We also compare RLMS with another official data source, the Household Budget Survey (HBS). HBS is the core Goskomstat source for published statistics on income differentiation and the composition of income and consumption. The HBS micro files are not publicly available. It is worthy of note that Goskomstat does not publish the actual income levels from HBS possibly for the reasons of massive under-reporting. Instead, it imputes money income as the sum of household expenditures and changes in financial assets (Goskomstat 1999).

Panels C and D of Figure 2 show the trends in consumption of food (including food grown at home) and non-food items, respectively. The statistics reported in Panels C and D are the average monthly consumption expenditures per household member. The RLMS expenditures are about 20 percent higher than its HBS counterpart, with the discrepancy being larger for non-food items. Some of the discrepancy is due to RLMS survey timing: the HBS reports average monthly consumption in a given year while RLMS reports last month consumption at the end of year. Then RLMS numbers should be larger when there is an upward trend in consumption and smaller when there is a downward trend. It is also plausible that consumer expenditures in HBS are underreported if people are reluctant to reveal their actual level of well-being in an official survey that asks, among other expenditures, the amount of taxes paid (RLMS does not ask about taxes, nor is it linked to any government agency).

Overall, RLMS appears to be a reliable data source for examining the inequality trends in labor market outcomes, reported income, consumption, with the common caveats of income underreporting and underrepresentation of the super-rich.

3. Inequality in Labor Market Outcomes

Since labor income is the most prevalent income source, the inequality in labor market outcomes is crucial for understanding the overall income inequality. This section takes a closer look at the dynamics of inequality in individual wages and labor supply, emphasizing the key differences between major population groups.

Aggregate labor market trends

We start with an overview of aggregate trends in wages and employment. Several studies observed that during the downturn period in Russia, the decline in employment and hours of work was small while the wage decline was large relative to the output decline, in contrast to Central and Eastern European transition economies (Boeri and Terrell 2002, World Bank 2002). We find that the post-1998 economic growth was also accompanied by significant wage adjustments and relatively small changes in employment and working hours.

Hourly real wage level experienced dramatic movements, down 48 percent, or 10 percent per year, during the downturn and up 87 percent, or 9 percent per year, during the recovery (Figure 3A). Panel A of Figure 3 shows actual hourly wage, defined as the ratio of actual labor earnings received last month from all regular jobs to actual hours worked, and compares it to contractual hourly wage (available 1998-2005), which is the ratio of average monthly labor earnings in the last 12 months to usual hours of work per month. The actual wage is higher than contractual wage, partly because actual hours are lower. Male wages appear to be more responsive to output fluctuations: male wages declined faster in downturn, but they also grew more rapidly in recovery.

In contrast to wages, hours of work do not vary considerably over time (Figure 3B). Even in the downturn, an average employed person worked more than 40 hours per week. The response of hours to the 1998 financial crisis was minimal. Usual hours of work are relatively high (48 hours in all jobs for males), and they are bigger than actual hours because of temporary absence from work due to illness, vacation, maternity leave, involuntary unpaid leave, and other reasons. Females typically work 5-6 hours less per week than males. The share of full-time workers does not change much in response to output fluctuations: it increases slightly over time for both genders, with a somewhat larger overall rise for females during 1994-2005 (Figure 3C).

Employment-to-population ratio in Russia is high by international standards. However, it declined significantly for males from 94-96 percent in 1985-1990 to 86 percent in 1994, and then down to 79 percent in 1998 (RLMS 2000, retrospective questions). In the growth period, the ratio did not revert to pre-crisis levels and stayed relatively constant at 83-84 percent for 25-59 age group (Figure 3D). On average, the employment rate for females is 8 percentage points lower than that for males, which is a smaller gender gap compared to 14 percentage points in the U.S. for the same age group (U.S. Bureau of Labor Statistics 2006). Figure 3D also shows that the official employment rate is lower than that in RLMS in the 1990s, but the difference between the two data sources vanishes in later years.

Earnings and wage inequality

Our sample starts in 1994, in the middle of an economic contraction in Russia that lasted almost a decade. Available evidence suggests that earnings inequality increased in the years preceding our sample period. This increase was associated with the transition to a market economy (Commander *et al* 1999). We estimate that the Gini coefficient for earnings increased from 0.28 in 1985 and 0.32 in 1990 to 0.48 in 1995 (RLMS 2000, retrospective questions).[10] The 90/50 ratio climbed from 2.2 in 1990 to 3 in 1995, while the 50/10 ratio rocketed from 2 to 4 in just five years.

During our sample period, however, wage inequality ceased to grow, as can be seen in Figure 4. This figure compares four different measures of inequality for individual earnings and hourly wages, both actual and contractual, in 1994-2005.[11] Actual earnings received last month are much more variable than contractual earnings. Part of the reason is delayed wage payments, which were a major phenomenon during 1994-1998.[12] Wage arrears tend to exaggerate earnings inequality. For example, some workers in the sample may have received several months of back pay and others received no pay in the reference month, generating income dispersion that is higher than dispersion in annual incomes. At the peak of wage arrears in late 1998, 62 percent of Russian workers reported overdue wages averaging 4.8 monthly salaries per affected worker (Earle and Sabirianova Peter, forthcoming). Consequently, the difference in actual and contractual earnings inequality was the largest in 1996-1998. Wage arrears subsided in later years, although they did not disappear entirely: about 12 percent of all employees reported delays in wage payments in 2005. Because of this, and also due to seasonal and irregular employment, actual earnings still show higher inequality than contractual earnings in later years. In Figure 4, measures of inequality for hourly wages and earnings are close because of low dispersion of working hours.

[10] This dynamics of the Gini coefficient is consistent with other studies. For example, Flemming and Micklewright (2000) report an increase in the Gini coefficient for per capita income from 0.27 in 1989 to 0.41 in 1994 based on the Household Budget Survey. They note, however, that inequality could have been larger in the Soviet period after accounting for significant in-kind subsidies (e.g., free housing).

[11] The observations on contractual earnings are available starting in 1998. For 1994-1996, we construct contractual earnings from the data on actual earnings and answers to questions about accumulated overdue wage amount and number of months of overdue pay, according to the method proposed by Earle and Sabirianova (2002). See Appendix 1 for details.

[12] Other reasons for excessive volatility of actual earnings in 1994-1998 include widespread temporary layoffs in the form of unpaid involuntary leaves and forced in-kind payments in lieu of wages owed. The use of involuntary leave peaked in 1996, when 15.8 percent of employees had average leave duration of about eight weeks. In-kind substitutes for money wages peaked in 1998, with 15.4 percent of workers affected (World Bank 2002). Adding these forced substitutes to actual earnings extends the bottom of the distribution of positive last-month income receipts and thus increases the overall dispersion.

According to most measures in Figure 4, inequality in wages and individual earnings has been declining over the sample period. The Gini coefficient for contractual earnings declined from 0.48 in 1995 to 0.41 in 2005 and the variance of logs decreased by 0.17. The decline in earnings inequality is more pronounced in the bottom half of earnings distribution: while the 90/50 ratio hardly changed over the sample period, the 50/10 ratio fell sizably from 4 to 2.5.

It may seem unusual that inequality at the bottom of the distribution was declining during an economic contraction. One explanation is that the timing of contraction (that started around at least as early as 1991) differed by income groups: for example, the dramatic rise in 50-10 ratio prior to our sample period suggests that low income workers suffered the most during the first years of market reforms. Several factors may have contributed to the decline in earnings inequality at the bottom of the distribution that continued after 1998: oil-driven growth that created labor demand in low-skill industries such as mining and construction, enhanced competition for workers (e.g., the number of employers increased dramatically), improved compensation in the public sector, etc.; each of these factors deserve a separate study.

Although the inequality indices remained higher than their pre-transition levels, the overall inequality decline is quite remarkable, and the reasons for it merit further research in the future. This trend is consistent with international macroeconomic data showing a negative contemporaneous correlation between income inequality and economic growth for less developed countries (Barro 2000).

Many Russians may be surprised to find that inequality has declined given the emergence of the conspicuous wealthy elite and a popular belief in the rising gap between rich and poor. We note, however, that adding the super-rich to the RLMS data will not affect the Kuznets ratios in Figure 4. There still might be a valid concern that upwardly mobile high earners may have left the addresses surveyed by the RLMS interviewers, and that those who stayed are self-selected low earners. Some of the issues with panel attrition are addressed within the survey itself by adding new dwellings to the sample and adjusting the sample weights.[13] The fact that

[13] To assess the importance of non-random exit from the survey on the measures of inequality, we re-weighted observations by giving a larger weight to observations with a higher probability of exit. The adjusted weight is calculated as *L.weight* × 1/(1- *Pexit*), where *L.weight* is the sample weight from the previous round and *Pexit* is the

our sample is unlikely to be skewed towards poor is also supported by the RLMS aggregate expenditure levels that are close to NIPA and exceed the HBS levels reported in Section 2.

Wage premia

The analysis of between-group wage inequality reveals several interesting results. They are reported in Figure 5 that shows aggregate trends in wage premium associated with education, gender, and experience. The male education (college/non-college) premium is substantial (about 50 percent on average), although it is smaller than the current education premium in the U.S. (e.g., Autor *et al* 2008, Eckstein and Nagypal 2004). The education premium has been rising since 1995 but dropped after 2002.

The gender premium in monthly earnings is large (up to 69 percent in 2000), even though it declined to 51 percent in recent years. The gender differences in hourly wages are smaller (35-47 percent) due to fewer hours of work by females. The level is comparable to the U.S. gender premium in the 1970s (e.g., Blau and Kahn 2000).

Remarkably, the male experience premium is negative, and it is below the female experience premium (Figure 5C). The age-earnings profile reaches its peak at age 33 for males (44 for females), whereas male earnings growth in the U.S. continues until much later ages (e.g., Heckman *et al* 2008). This unusual earnings profile may be partly attributed to the obsolescence of skills of Soviet-era workers.[14] However, if obsolete skills were the sole driving force of the negative experience premium, one would expect the experience premium to be low at first and to rise gradually over time as the old-era workers move out of the labor force. In fact, the male experience premium stays negative and roughly constant throughout the sample period. Another explanation for this result is that dramatic economic changes during both contraction and growth periods generated a wage premium for younger workers because they are more mobile and more

probability of exit from the survey estimated from a flexible probit regression that includes a wide range of controls for individual characteristics. We found that adjustment for non-random exit barely changes the magnitude and the trend slope of earnings inequality.

[14] Consistent with this, Guriev and Zhuravskaya (2008) find evidence of a big shift in life satisfaction by cohort: individuals who finished their education just before the transition report much lower life satisfaction than similar individuals who finished their education just after. This jump in life satisfaction could, perhaps, reflect brighter lifetime earnings prospects of workers educated under the new regime.

adaptive. We also think that deteriorating health, particularly for males, could be a contributory factor to the negative experience premium. The life expectancy of Russian males has dropped by 6.6 years, from 64.2 to 57.6, in just five years prior to 1994 (Brainerd and Cutler 2005). To the extent that this signals deteriorating health of males in their 50s, the "physical decay" of human capital could drag down the experience premium.

The residual inequality trends down over time, which is expected since the overall inequality is declining while the various wage premia for observable characteristics stay roughly constant (Figure 5D). By way of comparison, the residual wage inequality has an upward trend in the U.S. (e.g., Autor 2008, Lemieux 2006).

Gender differences in labor market outcomes

Figure 6 presents gender comparisons of inequality in hourly wages and hours worked. Wage inequality is higher among males than females, which is found in the U.S. data too (e.g., Eckstein and Nagypal 2004). Measures of wage inequality for both genders trend down over time, although the decline in inequality is more pronounced for males (this is again consistent with a higher responsiveness of male wages to output fluctuations). Consequently, the differences in wage inequality between genders become less noticeable by the end of the sample period (Figure 6A). Contractual wages show less dispersion than actual wages for both genders.

Hours worked are considerably less variable than wages (note that panels A and B have different scale). Females have slightly more variable hours, perhaps due to higher prevalence of part-time work. Dispersion of hours falls during 1994-1996 and stays stable afterwards.

The bottom two panels of Figure 6 show the correlations between hours and wages for males and females. These correlations are negative for both genders, which is probably due to a downward bias induced by a measurement error in hours, known as "division bias" (Borjas 1980). There is no clear time trend in the correlation between wages and hours for either gender.

Overall, the observed group differences in labor market outcomes behave in expected ways, with the exception of the negative male experience premium. We now turn to the analysis of inequality across households.

4. Inequality in Household Income and Consumption

This section analyzes the aggregate trends in income and consumption inequality at the household level. We first examine inequality in household labor earnings and then show the contributions to inequality from financial income, private transfers, government transfers, and home production. We also compare income inequality to consumption inequality and discuss possible reasons for the observed differences.

Inequality in household labor earnings

The RLMS data have several sources of information on household labor earnings. Our preferred measure of labor earnings, *yL*, is aggregated from individual responses on after-tax contractual labor earnings (see Appendix 1 for details). We note that Russian households are rather large and often include multiple generations of adults and extended family. The average number of adult members (14+) is 2.6, and it is not rare for a household to have more than two earners (see Appendix 3 for the sample composition of households). In this case one needs to be particularly careful when aggregating individual responses to the household level and should adjust for non-response.[15] However, since the RLMS response rate within the household is fairly high (about 97%), this adjustment does not affect the mean and the variance of labor earnings (e.g., compare *yLc* and *yL* in Figure 7A). Figure 7A shows that over time, the dispersion (var-log) in total contractual earnings across households is trending downward.

Another measure plotted in Figure 7A is the variance of the logarithm (var-log) of household actual labor earnings received last month (*yLa*). These earnings are reported on behalf of all household members by the reference person. While contractual earnings are monetary, actual earnings also contain non-monetary compensation (including forced in-kind substitutes for cash payments) that may introduce additional variability to household earnings. The dispersion in actual labor earnings has been declining after 1998, but its magnitude is considerably higher in comparison to contractual earnings, especially in the second half of the 1990s. As Figure 7B exhibits, the 1996-1998 period have the largest share of working

[15] We impute labor earnings for non-respondents using their demographic characteristics known from the roster files and location.

households affected by wage arrears (67-68 percent) and forced in-kind substitutions of payments (17-20 percent), which are the two most likely contributors to high earnings volatility (see also footnote 9). By 2005, the difference in dispersion between actual and contractual earnings reduces significantly, but it does not disappear entirely, possibly due to a measurement bias of one-person reporting, irregular employment, and residual wage arrears.

The variance of the log of labor earnings can be decomposed into parts accounted for by observable components based on the following regression:

$$\ln(yL_{ht}) = \beta_{0t} + \beta_{1t}D_{ht}^{H} + \beta_{2t}D_{ht}^{L} + \beta_{3t}D_{ht}^{E} + f_t(a_{ht}) + u_{ht}\,, \tag{1}$$

where yL_{ht} is contractual labor earnings of household h in year t, β_{0t} is year-specific intercept, D_{ht}^{H} is a set of dummies for household composition (e.g., categories for size, number of children, and number of seniors), D_{ht}^{L} is a vector of location characteristics such as an urban dummy, a dummy for Moscow and St. Petersburg, and 7 dummies for federal districts, D_{ht}^{E} denotes a set of dummies for educational attainment of the head of household, $f_t(a_{ht})$ is a quartic polynomial in age of household head, and u_{ht} is the error term (see Appendix 1 for details on how these components are constructed). The equation is estimated separately for each year. The observables explain a significant portion of inequality; however, the residual inequality remains large (46-62 percent, as shown in Figure 7C). The relative magnitude of residual inequality is similar to the one in developed countries. Figure 7D plots the contributions of observable components to the overall dispersion of household labor earnings. Location and household composition factors contribute the most to the observed inequality; education contributes some but age contributes close to zero. Because of its importance for inequality in Russia, we will consider the effect of location on inequality in more detail in Section 6.

Comparisons of earnings inequality trends for individuals and households

It is informative to compare the dispersion of earnings at the individual level (*ec* on Figure 4A) and the household level (*yLc* on Figure 7A). In general, one would expect the distribution of household earnings to differ from the distribution of individual earnings due to the presence of multi-generational, multi-earner households. In RLMS, 56 percent of working households have more than one earner, and over 10 percent of working households have three earners or more. The resulting distribution of household earnings is strongly correlated with the number of earners in the household. For example, 85 percent of households in the lowest per capita earnings quintile are single-earner, while 27 percent of households in the highest per capita earnings quintile have three earners or more.

The dispersion of *yLc* is larger than the dispersion of *ec* throughout the whole period. Since secondary earners have virtually the same earnings dispersion as primary earners, the larger dispersion of household earnings implies the positive correlation in earnings levels among household members. The trends in household and individual inequality are also different: individual earnings inequality falls more rapidly after 2000 than household earnings inequality. The divergence in inequality trends between individual earnings and household earnings appears to be driven by the increased covariance between earnings of household members. At the same time, the share of multi-earner households in the sample does not change over time. The average income share of secondary earners in household labor income also stays stable.

Inequality in equivalized labor earnings

To account for the effect of household size on earnings inequality, we compute the equivalized household labor earnings, *yLe*, using the OECD equivalence scale.[16] The dispersion for log equivalized earnings is almost the same as raw dispersion because equivalized earnings are negatively correlated with household size (Figure 7C). Figure 8 presents several alternative measures of inequality in household labor earnings per adult equivalent. Similar to Figure 7C, the Gini coefficient and both Kuznets ratios for household equivalized earnings exhibit a downward trend in the recovery period and show rapid convergence in inequality between the actual and contractual measures of labor earnings after 1998. As explained above, the downward trend in household earnings inequality is less pronounced than the downward trend for individual earnings inequality.

[16] The OECD equivalence scale assigns a value of 1.0 to the head of the household, a value of 0.7 to each additional adult (17+), and a value of 0.5 to each child.

From wages to disposable income

Using the variance of the log, we analyze how income inequality changes as we add different components of household income. Figure 9A shows that the magnitude of dispersion and its trend hardly change as we move from hourly to monthly contractual earnings of household head. However, the earnings dispersion increases when we add earnings of other household members (*yL*). Again, we observe that inequality in household labor earnings falls more slowly after 2000 than inequality in individual labor earnings.

Taking household labor earnings as the base, we add income from other sources one at a time and report the corresponding inequality trend in panels B and C of Figure 9. For comparability purposes, all time series in Figure 9 are calculated on a consistent sample of working households with non-zero contractual earnings. Net private transfers further increase the dispersion of earnings throughout the whole period (the *yL+* line in panel B is above the *yL* line), conceivably, because they are made irregularly.[17] Income derived from financial assets is negligible for most households, which is why financial income has virtually no effect on inequality. Government transfers, on the other hand, play a significant role in reducing income inequality, especially after 1998 (see *yD* line in panel B). The spike in income inequality in 1996 could be explained by unusually high pension arrears and unemployment benefit arrears in that year. Having income from subsidiary farming at home (which includes both own consumption valued at market prices and sales of home grown food) also has a large equalizing effect on earnings distribution, as evidenced in panel C.[18]

The dispersion of disposable income of working families with one or more wage earners exhibits a downward trend since 1996. However, adding non-working families (about 11% of the sample) not only shifts the overall income inequality up but also alters the time trend (see Figure 9D). This is because non-working families whose income consists of small private or

[17] The correlation between net private transfers and household labor earnings is -0.14, suggesting that that more affluent households are likely to support other households, while not-so-affluent families are likely to receive support from others. However, unlike public transfers, sporadic lump-sum private contributions may cause sizeable movements in the resources available to households, which raise our measures of income inequality. In our view, private transfers would probably decrease inequality if they were measured on annual basis.

[18] A related study by Gottschalk and Mayer (2002) shows that income adjusted for the value of home production is more equally distributed than unadjusted income in the U.S.

public transfers are more likely to fall into the bottom end of income distribution. Over time, more non-working households report small positive income either because they started receiving public transfers or because payment of transfers became more regular. This may explain why income inequality for the pooled sample of working and non-working families does not decrease over time.

Inequality in consumption

Figure 10 presents the dispersion of our benchmark measure of consumption, non-durable expenditure for all households, working and non-working. We see that the dispersion of non-durable consumption increases significantly during the downturn and falls rapidly during the economic recovery. Other consumption variables follow this trend very closely, although their variance may have different magnitude. In particular, adding durable expenditures (*cD*) increases consumption variance while adjusting for services from owned housing reduces it (*cD+*). The equalizing effect of housing on consumption distribution is predictable since many households, especially older ones who are also poorer, inherited their housing from the Soviet era.

Figure 10 also presents decomposition of non-durable consumption inequality based on equation (1). Similarly to household earnings decomposition in Figure 7, the dispersion of equivalized consumption is slightly lower than the dispersion of raw consumption. The residual consumption inequality is large and follows the same time pattern as the raw measure of consumption inequality (Figure 10C). As was the case with income decomposition, the largest observable contributors to consumption inequality are household composition and location. Education of household head explains some of the consumption inequality, but age explains almost none (Figure 10D). By contrast, in the U.S. inequality across households typically grows with age. The lack of correlation between measures of inequality and age in Russia is also reflected in the flat life-cycle inequality profiles (see Appendix 4).

Comparison of income and consumption inequality

Figure 11 compares various measures of consumption and income inequality. While income inequality in the pooled sample of working and non-working households does not fall over time, consumption inequality rises during the downturn and falls during the recovery. One remarkable result is that consumption inequality actually exceeds income inequality in 1996-1998, which seems to be at odds with consumption smoothing. This fact may be driven by the tendency of Russian households to store food as a means of short-term consumption smoothing. Then expenditure would actually equal consumption plus "saving" in the form of food inventory change.

Why was food storage likely to spike in 1996-1998? We think that irregularly paid wages and transfers as well as volatile and unpredictable inflation made real household monthly income highly variable (e.g., note the difference between actual and contractual earnings inequality in Figure 7). In perfect financial markets, these income variations would be smoothed by changing the stock of household financial assets. However, most households in our sample do not hold significant financial assets, perhaps due to undeveloped financial markets or the low real rate of return associated with rampant inflation (recall Figure 1A). Instead, short-term consumption smoothing may have been done by adjusting food inventories: households that received several months of back pay purchased large quantities of storable food (i.e., flour, sugar, etc.) for future consumption. In this case we can have households that spend little and consume from their food inventories as well as households that spend a lot on food, but do not consume all of it. Thus, the presence of food storage can make expenditure inequality exaggerate consumption inequality. Consistent with this hypothesis, statistical decomposition of residual expenditure variance shows that its transitory volatility peaked in 1996-1998 (see Section 5).

In addition, income inequality may be subject to its own biases that would make it seem low relative to consumption inequality. As previously discussed in Section 2, income is likely to be underreported. To the extent that income underreporting varies by income level, underreporting can introduce a bias in measures of cross-sectional income inequality. For example, if higher income households report a smaller fraction of their income than the average household, cross-sectional measures of income inequality will be biased downwards. In particular, a downward bias in income inequality can explain why the 90/50 ratio in Figure 11B is higher for consumption than for income throughout the entire sample period.[19]

Expenditure inequality versus consumption inequality

Compared to the US, expenditure inequality in Russia is puzzlingly high relative to income inequality. For example, Heathcote *et al* (2008) report that in the US consumption inequality is three times lower than income inequality.[20] By contrast, expenditure and income inequality measures in Figure 11 are roughly comparable.

Part of the explanation for the apparently high expenditure inequality in Russia is that expenditure only partially captures the actual consumption. As noted in Section 2, many Russian households grow food on subsidiary plots and thus consume more food than their expenditure numbers suggest. Although the aggregate amount of food produced at home is fairly small (5-10 percent of non-durable expenditure), food production is concentrated among rural and poorer households. This can make expenditure inequality significantly overstate the true consumption inequality. It turns out that adjusting consumption for home-grown food produces a large equalizing effect on consumption distribution for all four measures of inequality, as can be seen in Figure 11, line *cH*. The impact of home-grown food on consumption inequality is particularly large at the lower end of the consumption distribution (compare panel B to panel C). Section 6 additionally shows that accounting for differences in the cost of living by location (i.e. using region-specific price deflators) reduces consumption inequality even further (see Figure 13A).

Thus, the distinction between expenditure and consumption in Russia is very important. Expenditure can be a noisy measure of consumption when households accumulate large inventories of goods (particularly, food) as a form of saving. Also, expenditure is not a complete

[19] There are also reasons to believe that income underreporting declined after 2001 (see Gorodnichenko *et al* 2009 for evidence). If this is the case, then the attenuation of income reporting bias towards the end of our sample period makes the true fall in income inequality even larger than that in Figures 7-9.

[20] Specifically, variance of the log of equivalized non-durable consumption is 0.24; variance of the log of equivalized household earnings is 0.75.

measure of consumption for households that heavily rely on home production. Consequently, expenditure inequality overstates consumption inequality. Although the aggregate amount of home produced food is small, its effect on inequality measures is quite substantial.

So far, our analysis of inequality measures relied on repeated cross-sections. In the next section, we will exploit the panel dimension of the data and investigate to what extent changes in income inequality translate into changes in consumption inequality.

5. Time Series Decomposition and Interaction of Income and Consumption Inequality

To understand the dynamics of inequality and the interactions between consumption and income, we need to identify the sources of uncertainty faced by households and to assess households' ability to smooth consumption. As a first pass, we exploit the panel aspect of RLMS and decompose the residual variability in consumption and income into permanent and transitory components. Specifically, we use a statistical model

$$\ln(s_{ht}) = X_{ht}\beta + u_{ht}^{(s)},$$

where s_{ht} is the variable of interest, such as income or consumption, and X_{ht} is the same set of controls as in equation (1). We decompose the residual term $u_{ht}^{(s)}$ into the sum of a transitory component and a permanent component that follows a random walk process:

$$u_{ht}^{(s)} = \alpha_{ht} + \varepsilon_{ht}, \qquad (3)$$

$$\alpha_{ht} = \alpha_{h,t-1} + \eta_{ht},$$

where $\varepsilon_{ht} \sim (0, \sigma_{\varepsilon,t}^2)$ is the transitory component and $\eta_{ht} \sim (0, \sigma_{\eta,t}^2)$ is the innovation in the permanent component. Note that the variances of the transitory and permanent components are allowed to be time-varying. Using the covariance matrix for the changes in $u_{ht}^{(s)}$ and an equally weighted minimum distance estimator, we estimate the time series for $\sigma_{\varepsilon,t}^2$ and $\sigma_{\eta,t}^2$. The estimation procedure is described in more detail in Appendix 5.

Variance of innovations to income

The estimates of $\sigma_{\varepsilon,t}^2$ and $\sigma_{\eta,t}^2$ are reported in Figure 12, Panels A-C. Each panel uses a separate income measure: individual labor earnings (*ec*) in panel A, household labor earnings

(yL) in panel B and household disposable income (yD) in panel C. The time pattern for variances is similar for all three income measures: the variance of innovations in the permanent component, $\sigma^2_{\eta,t}$, remained relatively stable while the variance of transitory component, $\sigma^2_{\varepsilon,t}$, declined considerably.[21] It appears that the fall in residual income inequality is primarily due to moderation of the transitory component.

The variances $\sigma^2_{\varepsilon,t}$ and $\sigma^2_{\eta,t}$ for permanent and transitory components of income are at least three times larger than comparable estimates for the U.S. (Heathcote *et al* 2008, Table 2). We do not think that this is due to differences in residual income inequality levels $u^{(y)}_{ht}$ between Russia and the US – the residual variance of log household earnings is 0.5 for the US and less than 0.6 for Russia (see Figure 7). Rather, the comparison seems to point to different sources of residual income inequality between Russia and the US. In Russia, residual inequality appears to be driven by high income mobility, making residuals $u^{(y)}_{ht}$ large and volatile. By contrast, in the US, the combination of high $u^{(y)}_{ht}$ and low $\sigma^2_{\eta,t}$ points to the dispersion of unobserved household fixed effects, α_{h0}, as playing a larger role.[22]

Variance of innovations to consumption

We use the same statistical procedure to perform the decomposition of household consumption. The results are reported in Figure 12D. Unlike transitory income variance, the variance of the transitory consumption component, $\sigma^2_{\varepsilon,t}$, does not start to fall until after 1998. The variance of transitory consumption is highest in 1996-1998. This seems to be consistent with our food storage story, as food inventory fluctuations would cause unexplained transitory consumption to be large.

It is remarkable that the permanent component of consumption is as volatile as the permanent component of income (Figure 12D). This sharply contrasts with recent trends in

[21] In addition to individual and household labor earnings and disposable income reported in Figure 12, we find that other income measures such as household earnings with income from home production have a similar trend.

[22] To take an extreme case, suppose that errors have no time series volatility, that is, the residual income variance that is due to household fixed effects stays constant over time. Then, the permanent-temporary decomposition will show zero variance for both ε and η. By contrast, if regression errors have a lot of time variation, this will lead to high variance in η, or ε, or both. In sum, the level of residual inequality is not necessarily related to the variances in the permanent-temporary decomposition

consumption and income inequality in the U.S., where dramatically increased income inequality did not translate into large increases in consumption inequality. Such divergence between the two inequality measures in the U.S. has been explained by developments in financial markets that allow more risk sharing and consumption smoothing (Krueger and Perri 2006) and by the changes in the persistence of income shocks (Blundell *et al* 2008). Russia witnessed significant advancements in financial markets (especially, consumer credit) towards the end of our sample period, yet we do not observe the divergence between consumption and income variance decompositions.[23] The high variance of permanent consumption innovation is even more puzzling given that Russian households had a variety of consumption smoothing tools such as saving, food storage, home production, variable labor supply, and extended family. On the other hand, the negative correlation between wages and hours and low savings are also consistent with the lack of insurance against income shocks (Heathcote *et al* 2007).

Response of consumption to income innovations

To look at possible changes in consumption smoothing patterns over time, we examine the response of consumption to innovations in the permanent and transitory components of income. We continue to assume that the income process is given by equation (3) and re-estimate the income equation jointly with a consumption equation that captures the impact of income innovations on residual consumption growth. We model the sensitivity of consumption to income components as in Blundell *et al* (2008):

$$\Delta u_{ht}^{(c)} = \phi_t \eta_{ht} + \psi_t \varepsilon_{ht} + \xi_{ht} - \xi_{h,t-1}, \tag{4}$$

The left hand side of (4) is the growth rate of residual household consumption. The first term in the right-hand side is the product of the permanent income innovation, η_{ht}, and the "loading" factor ϕ_t that measures the responsiveness of consumption to η_{ht}. Similarly, the second term, $\psi_t \varepsilon_t$, measures the response of consumption growth to a temporary income innovation, ε_{ht} given ψ_t capturing the sensitivity of consumption to ε_{ht}. The term $\xi_{ht} \sim (0, \sigma_{\xi,t}^2)$ absorbs

[23] Consumer credit more than doubled every year between 2002 and 2006 (Goskomstat 2006c).

measurement errors and unobserved household heterogeneity not attributed to income growth and other observables. See Appendix 5 for more details on the estimation procedure.

Blundell *et al* (2008) interpret loadings close to one[24] as indicating the lack of insurance against innovations in income. In contrast, if loadings are close to zero, then households have enough instruments (e.g., access to credit markets, self-insurance) to insulate consumption from income shocks. Loadings between zero and one can be interpreted as partial insurance.[25]

Table 1 presents the results from jointly estimating income equation (3) and consumption equation (4). The loading on the transitory component, ψ_t, is relatively small and falling over time, consistent with households being able to smooth temporary income shocks. The loading on the permanent income component, ϕ_t, is much larger, perhaps, indicating imperfect consumption insurance against permanent shocks. Nevertheless, ϕ_t is falling over time, consistent with an overall improvement in consumption insurance.

It is informative to compare the estimates in Table 1 to those reported in Blundell *et al* (2008) for the 1978-1992 U.S. data with similar estimation methodology. The 2005 estimates of ϕ_t and ψ_t from our Table 1 are close to those reported in Blundell *et al* (2008, Table 7): $\phi \approx 0.64, \psi \approx 0.053$. Thus the sensitivity of consumption to income innovations is about the same for Russian households in 2005 and US households in 1978-1992. By contrast, Russian households have a much higher variance of transitory and permanent income innovations: the Table 1 averages are $\sigma^2_{\varepsilon,t} = 0.204, \sigma^2_{\eta,t} = 0.088$ versus $\sigma^2_{\varepsilon,t} \approx 0.051, \sigma^2_{\eta,t} \approx 0.013$ for the US (Blundell *et al* 2008, Table 6).

The much more volatile income of Russian households makes potential welfare gains from consumption insurance much higher in Russia. Using our estimates, we can compute the variance of consumption growth that is due to innovations in income (Table 1, column 8),

$$\sigma^2_{cy,t} = \phi_t^2 \sigma^2_{\eta,t} + \psi_t^2 \sigma^2_{\varepsilon,t} .$$

[24] Loading coefficients cannot exceed 1 because this would violate household lifetime budget constraint.

[25] Blundell *et al* (2008) show that under certain restrictions the permanent income hypothesis implies ϕ=1 and ψ=0. That is, consumption should change by the same percentage as the change in the permanent income (ϕ=1), and it should not respond at all to transitory components (ψ=0).

The above variance can be used to estimate welfare gains from consumption insurance. Lucas (1987) shows a household with risk aversion parameter γ would be willing to sacrifice up to $\frac{1}{2}\gamma\sigma_{cy,t}^2$ share of their income to have a perfectly smooth consumption path. The Blundell *et al* (2008) parameters yield $\sigma_{cy,t}^2 \approx 0.0056$, while our Table 1, column 8 estimates of the same are 5-14 times larger. This means that Russian households should be willing to give up a 5-14 larger income share than the US households to achieve perfect consumption smoothing (assuming the same risk aversion parameter).[26] Intuitively, this result obtains because variance of consumption that is attributable to variance in income is much larger in Russia than it is in the US.

However, most of the variance in residual consumption growth is not attributable to income. Income innovations can only explain between 14 and 22 percent of variance in non-durable consumption growth (compare Table 1, columns 9 and 10). There may be several reasons for this. For example, consumption out of irregular unreported income and changes in food inventories would both be categorized as unexplained consumption growth. In addition, our chosen observables do not exhibit much time variation, but both consumption and income vary a lot over time, perhaps due to high occupational mobility. Finally, measurement errors and preference shocks could also contribute to unexplained consumption growth.

Overall, our findings suggest that income and consumption mobility was high in the early years of our sample, and that the ranking of households in the income distribution has been stabilizing in recent years. Despite recent improvements, households have had limited ability to smooth income shocks with financial assets, savings or other insurance instruments, and the benefit from providing access to such insurance probably remains substantial.

6. A Closer Look at Inequality Trends: the Role of Location and Prices

In the context of the Russian economy, two factors deserve special consideration as they can help with understanding of the observed inequality trends. This section takes a closer look at the role of geography and price dispersion.

[26] This result has an important caveat: if households smooth consumption with food storage, this would induce strong response of expenditure to transitory income shocks, but will not necessary imply non-smooth consumption.

Location effect

Russia is a large and diverse country, both geographically and economically. For example, monetary income per capita in the richest Russian region is 10.6 times larger than per-capita income in the poorest region in 2005 (Goskomstat 2007b). A similar maximum-to-minimum ratio across states in the U.S. is only 1.8 (U.S. Census Bureau 2007).

Location is the most important explanatory variable for the dispersion of earnings and consumption (see Figures 7 and 10). The substantial dispersion of the regional component of inequality may be associated with the large geographic variation in the cost of living. The 2005 ratio in the cost of fixed consumer goods between the most expensive region and the least expensive region was 2.7 (Goskomstat 2006a). With such inter-regional diversity, using a common national CPI may overstate the extent of inequality in both income and consumption. Indeed, using regional CPI and accounting for the regional differences in the cost of living move the magnitude of inequality down, but this adjustment does not affect the time trend (see Figure 13A).

The regional dispersion of expenditure may also be affected by uneven distribution of amounts of food grown at home between urban and rural households. While big city residents purchase more than 95 percent of their food at the store, residents of small towns and villages purchase about 80 percent at the store (less in early years) and grow the rest on their subsidiary plots. Consequently, rural households are likely to have a larger discrepancy between expenditure inequality and consumption inequality. Panels B and C of Figure 13 implicitly confirm this. The panels depict the variance of the log of non-durable expenditures for the two groups and the pooled sample using regional deflators, with and without food grown at home. Expenditure inequality is apparently much higher among the rural population (Panel B). By contrast, inequality in *consumption* that includes food purchased and grown at home is much more similar across urban and rural households (panel C).

While time trends in expenditure inequality for the two groups are similar, trends in consumption inequality diverge during economic recovery. In particular, consumption inequality among rural households shows no downward trend (panel C). This difference in trends is consistent with transition of rural households from subsidiary farming to professional farming, which could have made the amount of food grown for own consumption more unequally distributed.

Economic consequences of downturn and recovery may have differed between urban and rural populations. One would expect rural households to fall behind during the transition due to the lack of access to large and diverse labor markets that big cities offer and also because of possible migration of the ablest workers to cities.

Surprisingly, our data do not point to much divergence in the mean levels of income and consumption of the two groups until 2002. Figure 13D shows that the relative levels of disposable income (*yD*), expenditures (*c, cD*) and consumption (*cH*) stay fairly constant during 1994-2001. It is possible that rural households were already behind when our sample began – recall the discussion in Section 2. On the other hand, the relative consumption level of urban household was at its all-time high in 2002, 2004 and 2005, suggesting that rural households did lag behind as the economic recovery progressed. Particularly, the growth of durable consumption was stronger among urban households (*c* and *cD* lines in Figure 13D diverge after 2000).

The role of food grown at home in equalizing consumption is strikingly apparent in Figure 13D. Urban households, who spend 45 percent more than rural households, enjoy only 29 percent higher consumption, on average (compare *c* and *cH* lines).[27]

Comparisons of group income and consumption differences reveal important facts. On average, urban households report roughly 71 percent more disposable income than rural households, but their total expenditure is just 45 percent higher. Because saving rates of most households are fairly low, this leads us to suspect that income under-reporting is more severe among the rural households.

[27] Market value of home production probably overstates its net contribution to household welfare because of the cost of capital goods and materials and decreased leisure. Selling food is also likely to involve high transaction costs, making net income from home production lower than its value at market prices.

Price dispersion effect

Data on food prices available in RLMS allow us to investigate the effect of prices on the mean and the variance of real expenditures. It is important to examine price data for two reasons: (1) there is a concern that the growth rates of real output and expenditures may be mismeasured due to a bias in the official CPI (Gibson *et al* 2004), (2) expenditure inequality can arise as a result of either price dispersion or quantity dispersion, and only the latter captures inequality in actual consumption.

Using the data on food prices, we find that the official CPI substantially overstates inflation during 1994-1998. We also find that the effect of price dispersion on inequality is relatively small: inequality measures based on expenditures and those based on quantities purchased are close to each other.

Figure 1C shows that food expenditure fell faster than income during the 1994-1998 downturn (see Figure 1B and also the discussion in Mroz *et al* 2005), but never returned to its 1994 level during the economic recovery. Since food expenditures declined rapidly during the early, high inflation, period, one may suspect that the decline is not genuine and may be driven by a bias in the national CPI that we use to deflate food expenditures.

To check this, we construct the food CPI from RLMS prices and compare it to the NIPA deflator for food. Let p_{kt} denote the sample average unit price of food category k in year t, and q_{kt} denote the average physical quantity of food item k purchased in year t. Let $\overline{p}_k$ and $\overline{q}_k$ be the sample average price and the quantity purchased in the base year. Define a fixed-basket food CPI as

$$cpiF_{RLMS,t} = \sum_k p_{kt}\overline{q}_k \Big/ \sum_k \overline{p}_k \overline{q}_k$$

Figure 14A depicts the year-on-year growth rates of $cpiF_{RLMS,t}$ (with base year 2002) and the NIPA CPI deflator.[28] Unfortunately, the two deflators are not directly comparable, because we could not replicate the NIPA procedure without knowing the NIPA choice of base year and when it was changed. Predictably, two deflators in Figure 14A disagree most during the years

[28] To make it comparable, we calculate the NIPA inflation rate as the average monthly inflation rate weighted by the share of respondents interviewed in a given month.

with high inflation: the inflation rate derived from RLMS data on food prices is lower than the official CPI inflation in 1995 and 1998 years. As a result, between December 1994 and December 1998 food prices in RLMS grew by a factor of 4.5, compared to a factor of 5.6 according to NIPA. If the RLMS food deflator was used in place of the official CPI to compute real consumption, the 1998 value of aggregate consumption (*cD*) in Figure 1 would have been almost 25 percent higher.

The CPI deflator can also be subject to substitution bias if the composition of food consumption changes. To check whether this is the case, we compare two alternative measures of real food expenditures: nominal expenditures deflated by the fixed-basket food CPI, $cF_{RLMS,t}$

$$cF_{RLMS,t} = \sum_k p_{kt} q_{kt} \Big/ cpiF_{RLMS,t}$$

and the food quantity index, $qF_{RLMS,t}$

$$qF_{RLMS,t} = \sum_k \overline{p}_k q_{kt}$$

that weights current year quantities at base year prices. By construction, the ratio of the two expenditure measures equals the ratio of the CPI derived from the current year basket to the CPI derived from the base year basket. If $cF_{RLMS,t}$ and $qF_{RLMS,t}$ are substantially different, this is an indication of a time-varying consumption basket.

Figure 14B compares $cF_{RLMS,t}$ and $qF_{RLMS,t}$ and find that they are very similar to each other, suggesting that the food consumption basket was essentially fixed throughout the sample period. However, the real food consumption *cF* (computed with the official deflator) is substantially higher than both $cF_{RLMS,t}$ and $qF_{RLMS,t}$ during 1994-1998, indicating inflation overstatement by the official CPI.

Despite their apparent differences, all measures of real food expenditure show a decline over the sample period. The share of food in aggregate expenditures also steadily declined over the sample period (Figure 14C). From a viewpoint of static utility maximization, a change in demand may be driven either by a change in income or a change in relative price. We do not think that income change was driving the food demand, because real income was roughly the same in 1994 as in 2002, but food expenditures were much lower in 2002 than in 1994. The ratio of NIPA food deflator to NIPA non-food goods deflator (that excludes services), a proxy for the relative price of food, has also not changed (Figure 14D). Superficially, at least, there is no change in the relative price of food either, so the apparent fall in the share of food expenditures is puzzling.

We think that one explanation for this puzzle may be the lack of quality adjustment in the NIPA non-food CPI. The Soviet-era consumer goods were notorious for their low quality. If the relative quality of non-food goods rose over the sample period, this may have caused the quality-adjusted goods price to fall and the expenditure share to shift away from food towards non-food goods.

To summarize, our examination of food price and quantity data and its comparison with NIPA price indices point to evidence of quality bias. Gibson *et al* (2004) use RLMS food expenditure data to indirectly infer the total CPI bias from Engel curves and estimate that 2001 real GDP level may be understated by as much as 30 percent.

It is also important to check if the observed level of inequality in food expenditures arises from the difference in food prices that households face. Russia is a geographically diverse country with large variations in the price level by location. Other peculiar features of the Russian transition, including price liberalization, hyperinflation spikes, regional disintegration, imperfect markets, and elevated uncertainty, may contribute to relatively high price dispersion not only across locations but also within locations. As the economy stabilizes and markets develop, we may expect a decline in the level of price dispersion. On the other hand, the Soviet era products were fairly standardized with low quality variance. Over time, import penetration and domestic competition have brought new products of various qualities, thus increasing price dispersion. The resulting effect of transition on price dispersion is thus ambiguous. Figure 15A shows the trends in dispersion of food prices, overall and within location.[29] Price dispersion was high in 1994 but stayed constant afterwards, suggesting that counter-factors of dispersion cancel each other out.

[29] Price dispersion is calculated as $Var(P) = \sum_k Var(\ln p_{kt})\{[\sum_h q_{kh0} p_{kh0}]/[\sum_k \sum_h q_{kh0} p_{kh0}]\}$.

To control for cross-sectional price dispersion, we measure the inequality in food quantity index, $qF_{RLMS,t}$, that weights the quantity of food purchased at constant, base year prices. Figure 15B compares the variances of $ln(qF_{RLMS,t})$ and $ln(cF)$. Quantity dispersion does not have to be lower than expenditure dispersion, because the prices paid by individual households and the quantity of food that they purchased may be negatively correlated. In fact, Figure 15B shows that quantity dispersion is slightly higher than expenditure dispersion in 1995 and 1996 and slightly lower than expenditure dispersion in other years. Overall, the inequality trends for food expenditures and food quantity are similar.

Expenditure inequality may also be affected by regional differences in the cost of living. To control for this, we deflate food expenditures by *regional* CPI. The *cF-reg* line of Figure 15B shows the resulting measure of expenditure inequality. Predictably, using regional price adjustment has an equalizing effect on consumption distribution. Food expenditures with regional deflators show *less* dispersion than the food quantity index $qF_{RLMS,t}$, which indicates a negative correlation between the region-specific food prices and the quantity purchased.

7. Conclusions

We investigate the levels and the time trends of consumption and income inequality in Russia. The paper makes a number of contributions on issues of inequality measurement. We explain, for example, why consumption that includes home production, avoids underreporting of resources available to households, and is adjusted for regional variation in the cost of living should be a preferred inequality measure for Russian economy. We find that compared to its pre-transition level, inequality first rose and subsequently fell. The rise in inequality appeared to have happened during the price liberalization in the early 1990s while the fall started after 2000. The level of inequality in Russia is now very similar to that in the U.S. (e.g., Krueger and Perri 2006).

We uncover several important facts about inequality in Russia. First, poor households appear to gain from recent economic growth. Second, changes in key observable characteristics of households have a small contribution to the dynamics of consumption and income inequality.

The variance of permanent and transitory income components is much larger in Russia than in developed countries. Because of this, the fluctuations of consumption that are attributable to income shocks are larger in Russia as well. There are probably substantial gains from introducing insurance schemes to smooth consumption fluctuations. Third, recent moderation in consumption and income inequality and mobility appears to be driven by the decline in the volatility of transitory shocks. Fourth, unlike developed economies that presumably have rich consumption smoothing possibilities, expenditure and income inequality in Russia are not far apart.

Our results also point out some inconsistencies between RLMS and NIPA. In particular, comparisons of consumption levels across data sources suggest that there may be an insufficient adjustment for shadow economic activity in the official statistics. The growth rate of consumption in NIPA has recently become higher than that in RLMS, a phenomenon that was noted in other developing economies (e.g., Deaton 2005). The comparison of CPI levels reveals that NIPA may significantly overstate inflation, and that quality bias is potentially important.

Our analysis highlights several phenomena that merit further research. For example, the negative experience premium for males sharply contrasts with positive experience premium in other countries. Another puzzling finding is that income shocks explain a modest part of non-durable consumption variance. This could be due to consumer durables playing a bigger role in consumption smoothing, especially when financial markets are underdeveloped. There is theoretical work showing that income shocks can mostly be absorbed by durable consumption (e.g., Leahy and Zeira, 2005 and Stacchetti and Stolyarov, 2007). The panel structure of RLMS provides a natural data set for investigating the role of durable expenditure as a propagation mechanism for income shocks.

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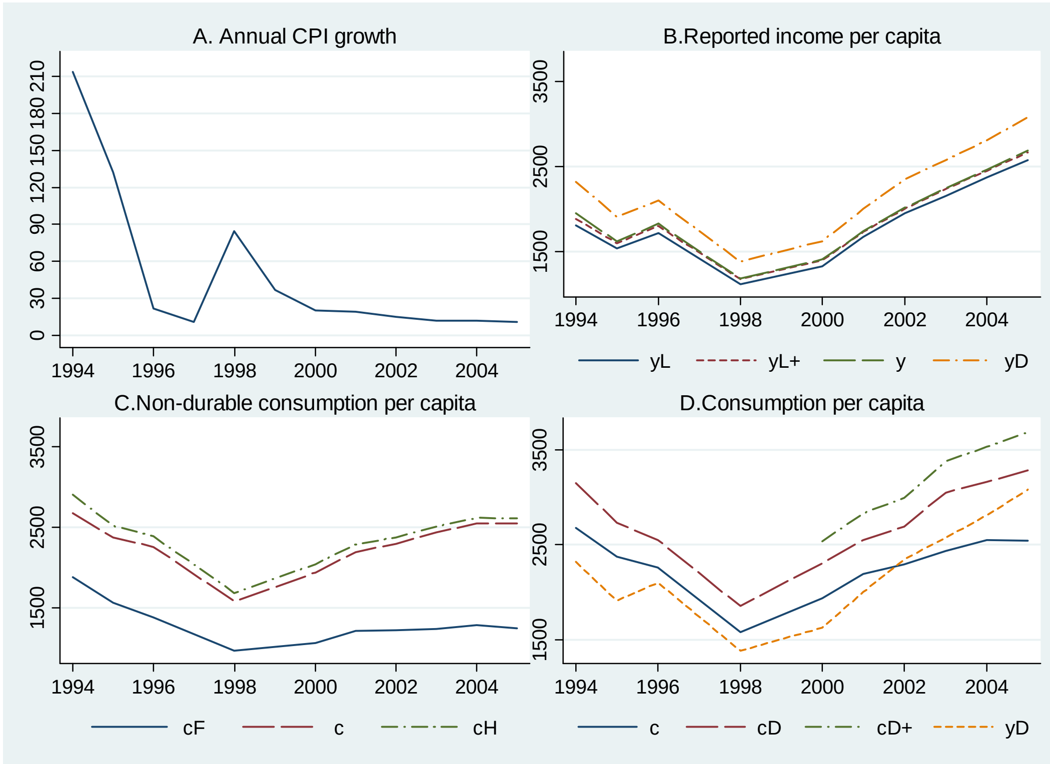
Table 1: Estimates of Consumption Response to Income Innovations.

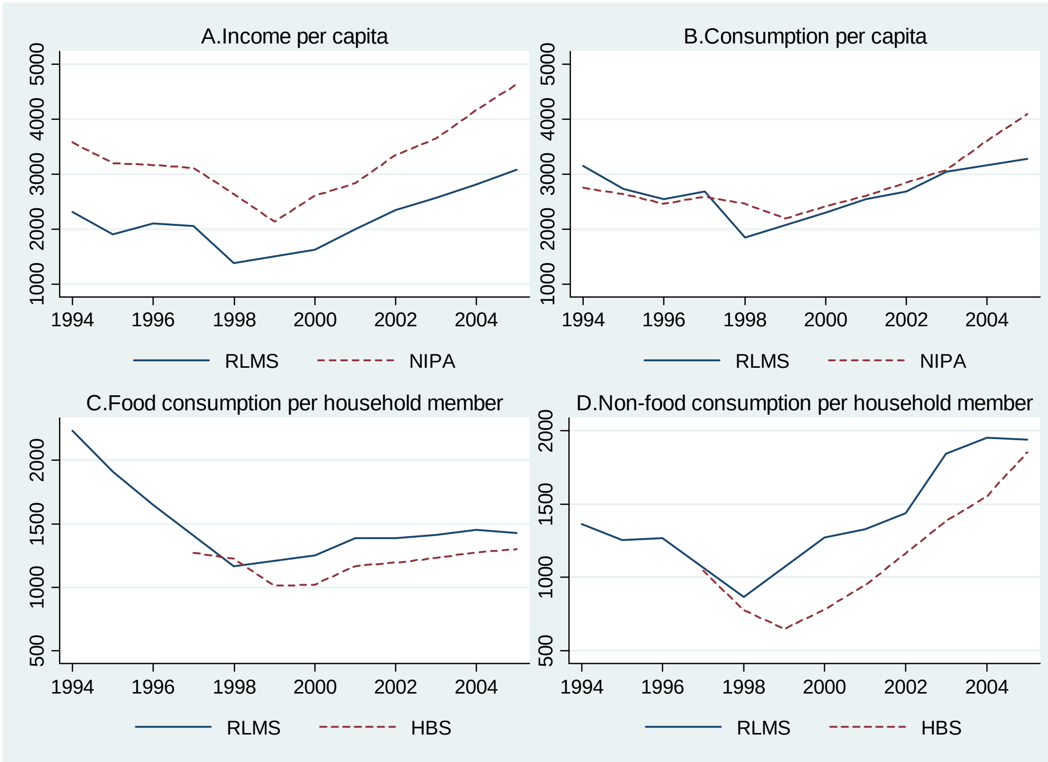
Year	Variance of innovation in a component of income		Loadings on components		Variance of residual consumption growth rate due to				
					Innovations in components of income			Measurement errors and unobserved household heterogeneity, $\sigma^2_{\xi,t}$	Total
	Transitory, $\sigma^2_{\varepsilon,t}$	Permanent, $\sigma^2_{\eta,t}$	Transitory ψ_t	Permanent ϕ_t	Transitory $\psi^2_t\sigma^2_{\varepsilon,t}$	Permanent $\phi^2_t\sigma^2_{\eta,t}$	Total $\sigma^2_{cy,t}$		
1995	0.257	0.069	0.125	0.908	0.0040	0.061	0.065	0.490	0.555
1996	0.305	0.069	0.125	0.908	0.0048	0.062	0.067	0.494	0.561
1998	0.238	0.092	0.174	0.801	0.0072	0.066	0.074	0.512	0.586
2000	0.213	0.115	0.151	0.561	0.0048	0.041	0.046	0.489	0.535
2001	0.184	0.093	0.040	0.568	0.0003	0.030	0.031	0.404	0.435
2002	0.166	0.107	0.028	0.559	0.0001	0.034	0.034	0.353	0.386
2003	0.169	0.071	0.036	0.629	0.0002	0.028	0.028	0.339	0.367
2004	0.153	0.085	0.032	0.718	0.0002	0.044	0.044	0.310	0.354
2005	0.153	0.085	0.032	0.718	0.0002	0.044	0.044	0.289	0.334
Average	0.204	0.088	0.083	0.708	0.0024	0.046	0.048	0.409	0.457

Figure 1: Trends in Household Income and Consumption



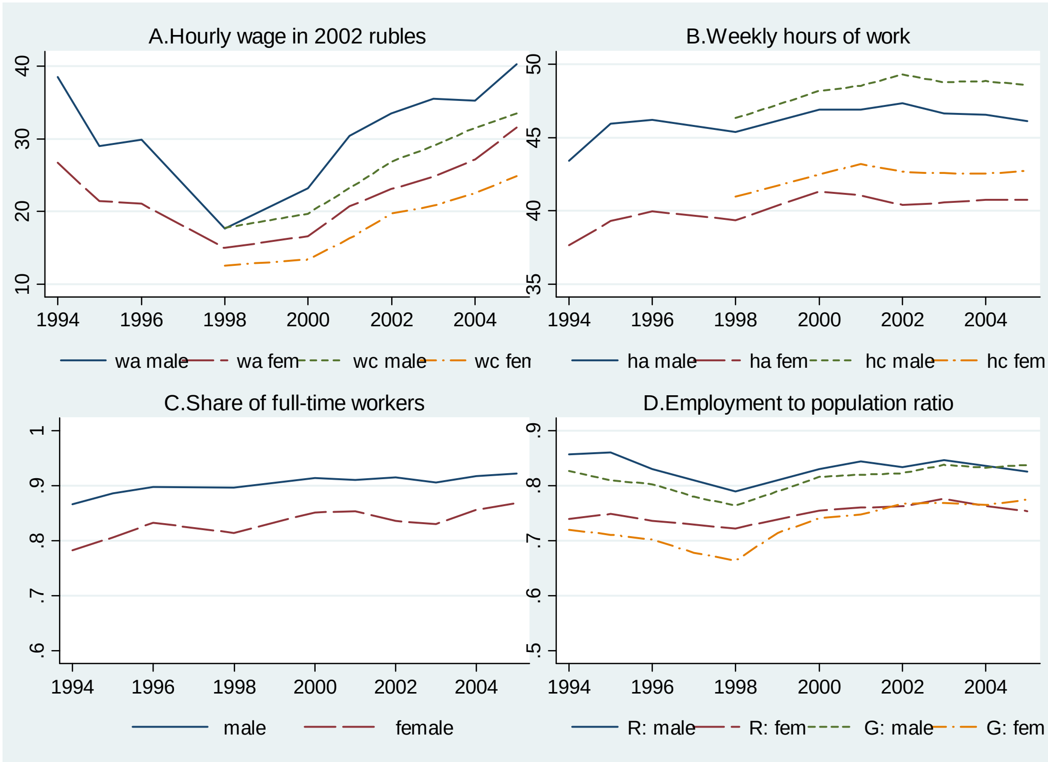
Notes: Panel A shows annual inflation rate using national end-year CPI from official sources. In remaining panels, all measures are in constant December 2002 prices (deflated using national monthly CPI and the date of interview). *yL* = household contractual labor earnings per month; *yL+* = *yL*+ net private transfers; *y* = (*yL*+) + financial income; *yD* = disposable household income = *y* + government transfers; *cF* = expenditures on food, beverages, and tobacco last week (multiplied by 30/7); *c* = household non-durable expenditures last month; *cH* = *c* + consumption of home-grown food; *cD* = *c* + expenditures on durables; *cD+* = *cD* + imputed services from housing.

Figure 2: Comparison of RLMS with Official Statistics



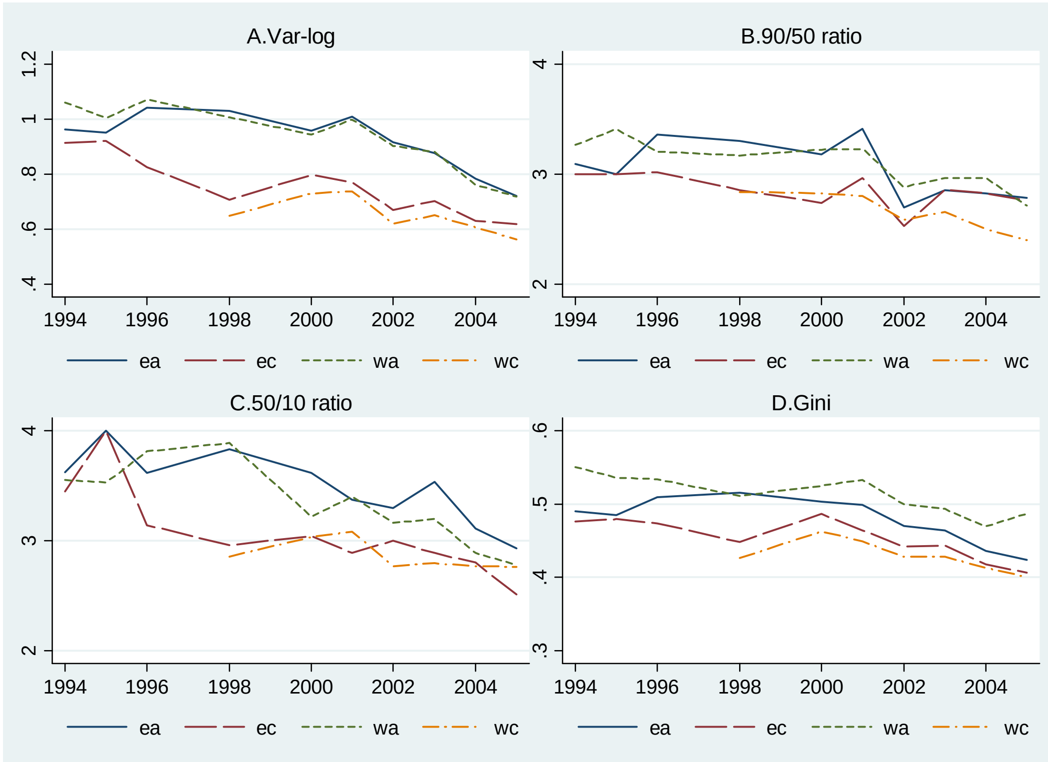
Notes: For comparability purposes, the following RLMS measures are selected: *yD* in panel A, *cD* in panel B, *cF* + consumption of home-grown food in panel C, *cD* - *cF* in panel D. The RLMS sample is unrestricted. All RLMS measures are deflated using monthly CPI and the date of interview. All NIPA and HBS measures are deflated using annual average CPI. RLMS income and consumption for 1997 are imputed using the lagged RLMS value multiplied by the 1997 growth rate from NIPA.

Figure 3: Trends in Labor Supply



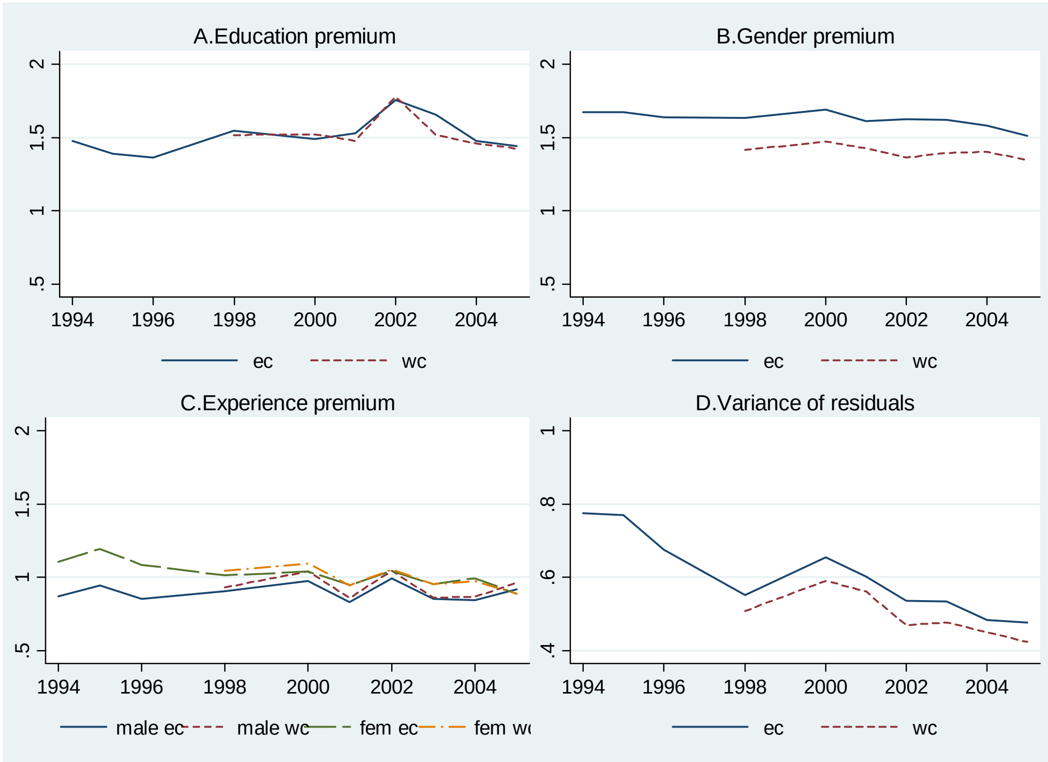
Notes: *wa* = hourly wage rate based on earnings received last month; *wc* = contractual hourly wage rate; *ha* = hours worked last month; *hc* = usual hours of work per month. All wages are deflated with national monthly CPI. Workers are considered full-time if actual hours at primary job were more than 120 hours in the reference month. Panel D compares employment-population ratios in the RLMS sample (R:) and official Goskomstat statistics (G:). Both ratios are calculated for age group 25-59.

Figure 4: Basic Inequality in Individual Wages



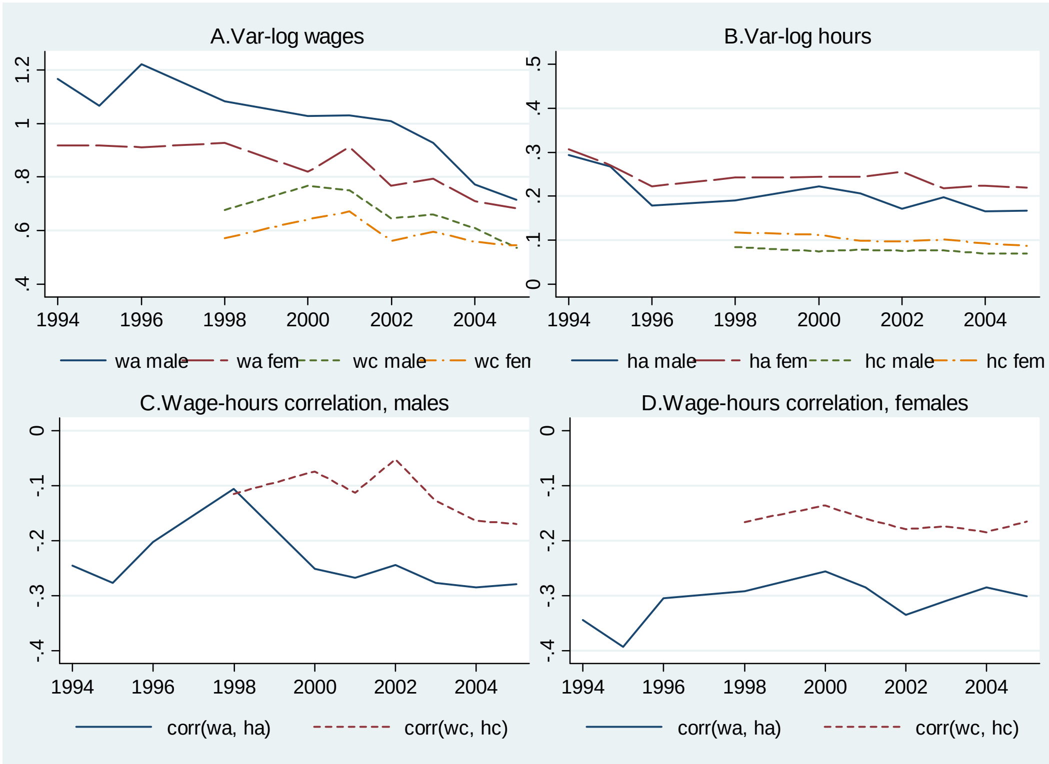
Notes: *ea* = actual individual labor earnings received last month; *ec* = contractual individual labor earnings per month; *wa* = hourly wage rate based on earnings received last month; *wc* = contractual hourly wage rate. Var-log is the variance of log earnings. All earnings are after-tax.

Figure 5: Wage Premia



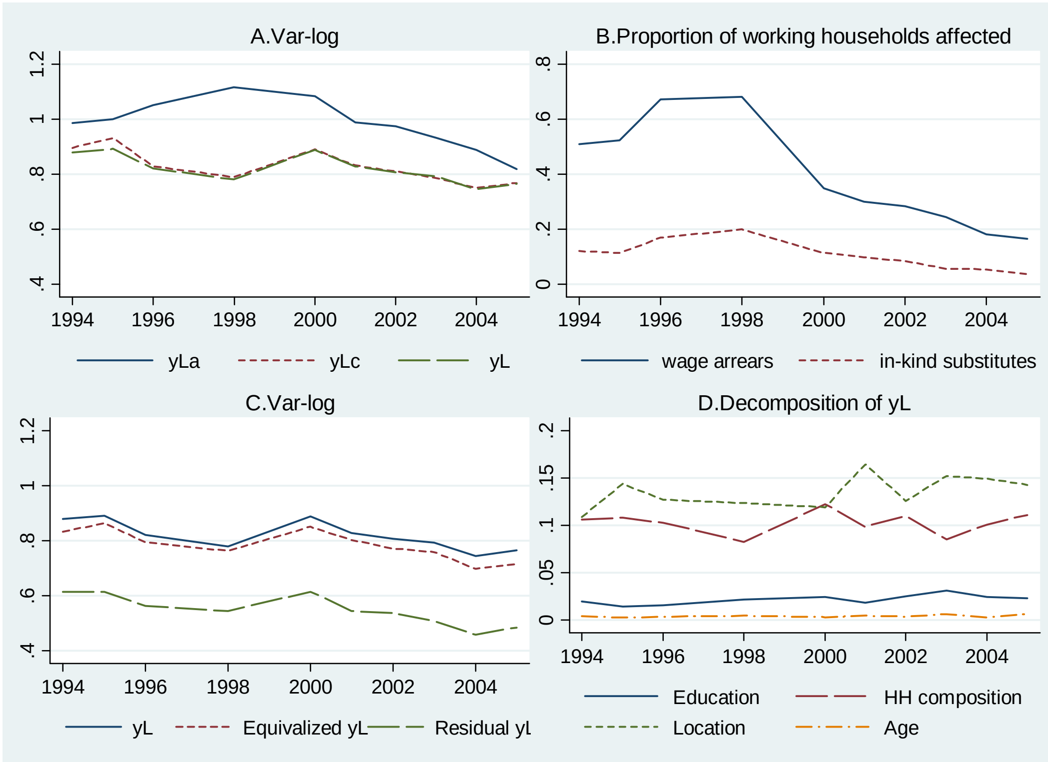
Notes: *ec* = contractual individual labor earnings per month; *wc* = contractual hourly wage rate. All earnings are after-tax. Education premium is the average wage of university educated males divided by the average wage of non university-educated males. Gender premium is the average wage of males divided by the average wage of females. Experience premium is the average wage of age group 45-55 divided by the average wage of age group 25-35. The variance of residuals is from equation (1).

Figure 6: Inequality in Labor Supply



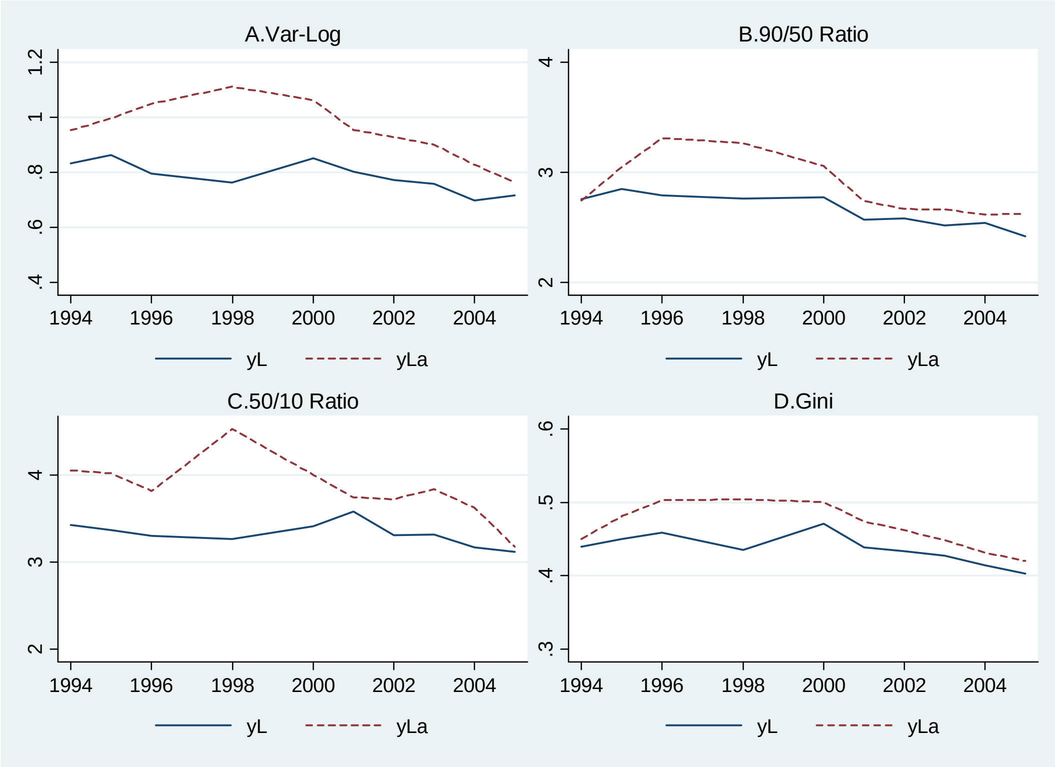
Notes: *wa* = hourly wage rate based on earnings received last month; *wc* = contractual hourly wage rate; *ha* = hours worked last month; *hc* = usual hours of work per month.

Figure 7: Household Earnings Inequality and Its Decomposition



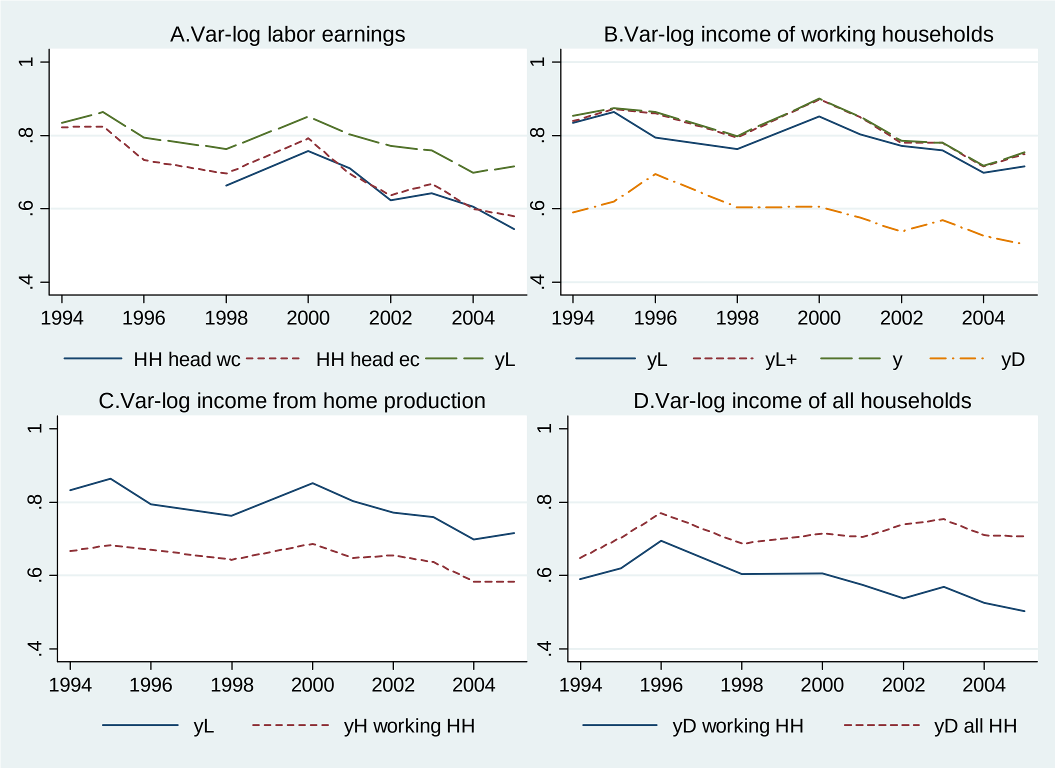
Notes: All earnings are after-tax and deflated using national monthly CPI. *yLa* = actual household labor earnings received last month; *yLc* = household contractual labor earnings per month; *yL* = household contractual labor earnings per month adjusted for non-response. Panel C reports the variance of log raw *yL*, the variance of log *yL* equivalized with an OECD equivalence scale, and the variance of residuals from equation (1). Panel D reports the variance of each observable component of equation (1).

Figure 8: Basic Inequality in Equivalized Household Earnings



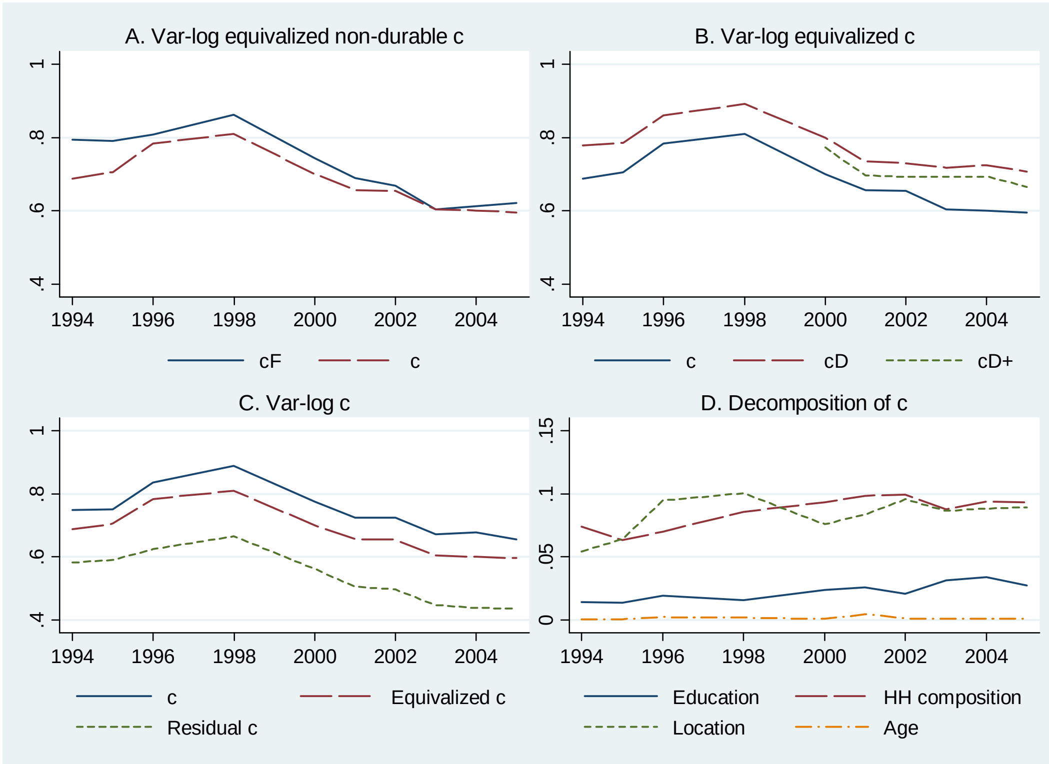
Notes: All earnings are after-tax, equivalized using an OECD equivalence scale, and deflated using national monthly CPI. *yLa* = actual household labor earnings received last month; *yL* = household contractual labor earnings per month adjusted for non-response.

Figure 9: From Wages to Disposable Income



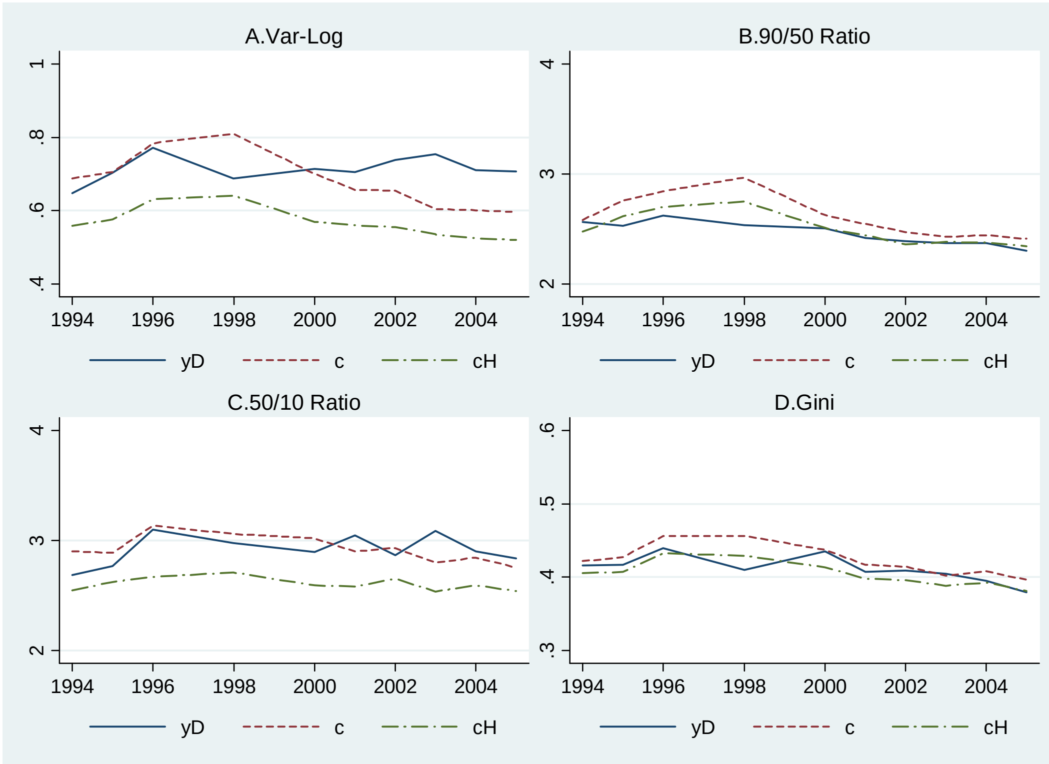
Notes: All income measures are after-tax and deflated using national monthly CPI. Measures at the household level are also equivalized using an OECD equivalence scale. *HH head wc* = contractual hourly wage rate of the head of household; *HH head ec* = contractual labor earnings per month of the head of household; *yL* = household contractual labor earnings per month adjusted for non-response; *yL+* = *yL*+ private transfers; *y* = (*yL+*) + financial income; *yD* = disposable household income = *y* + government transfers; *yH* = *yL* + income from home production. Working households include households with at least one wage earner. Var-log is the variance of the logarithm of income.

Figure 10: Consumption Inequality and Its Decomposition



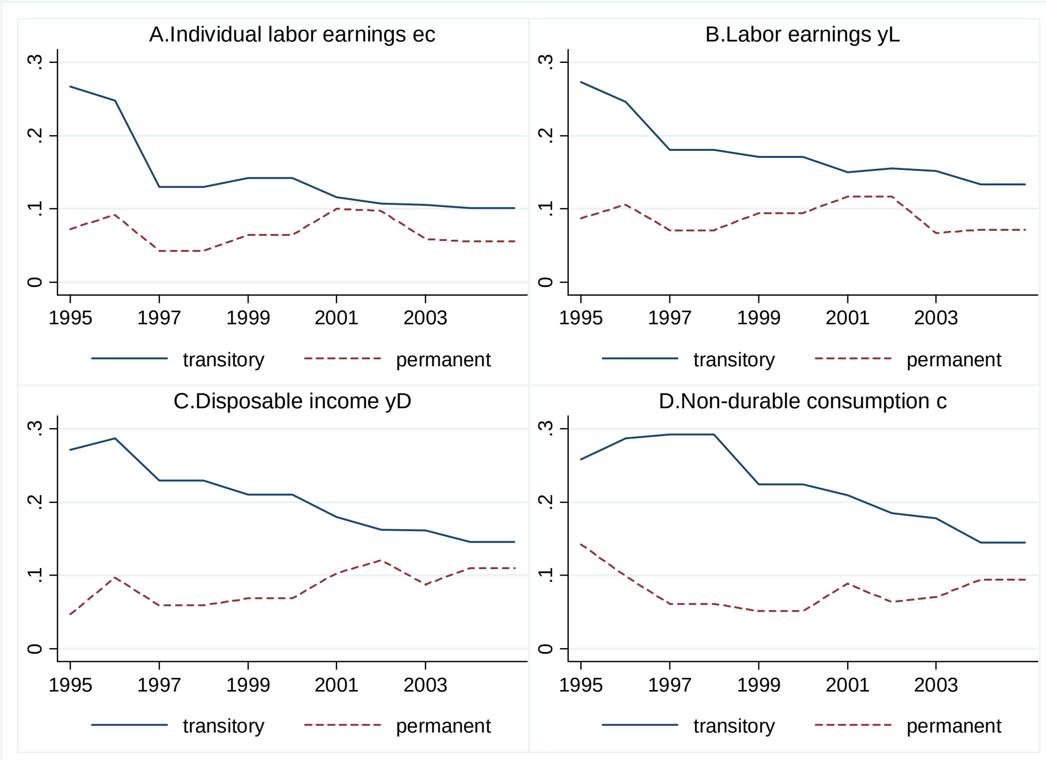
Notes: *cF* = expenditures on food, beverages, and tobacco last week (multiplied by 30/7); *c* = household non-durable expenditures last month; *cD* = *c* + expenditures on durables; *cD+* = *cD* + imputed services from housing. All consumption variables in Panels A and B are per adult equivalent. Panel C reports the variance of log raw *c*, the variance of log *c* equivalized with an OECD equivalence scale, and the variance of the residuals from equation (1). Panel D reports the variance of each observable component from equation (1).

Figure 11: From Disposable Income to Consumption



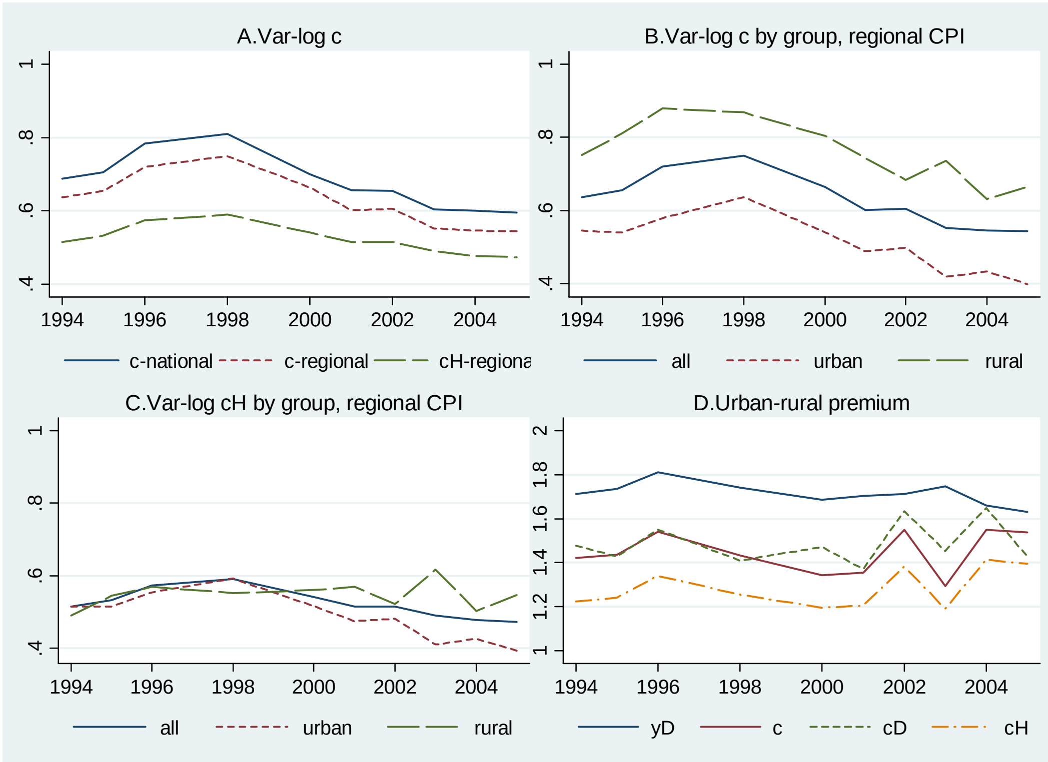
Notes: *yD* = disposable household income based on contractual labor earnings; *c* = household non-durable expenditures last month; *cH* = *c* + consumption of home-grown food. All measures are equivalized using an OECD equivalence scale and deflated with national monthly CPI.

Figure 12: Permanent-Temporary Component Decompositions



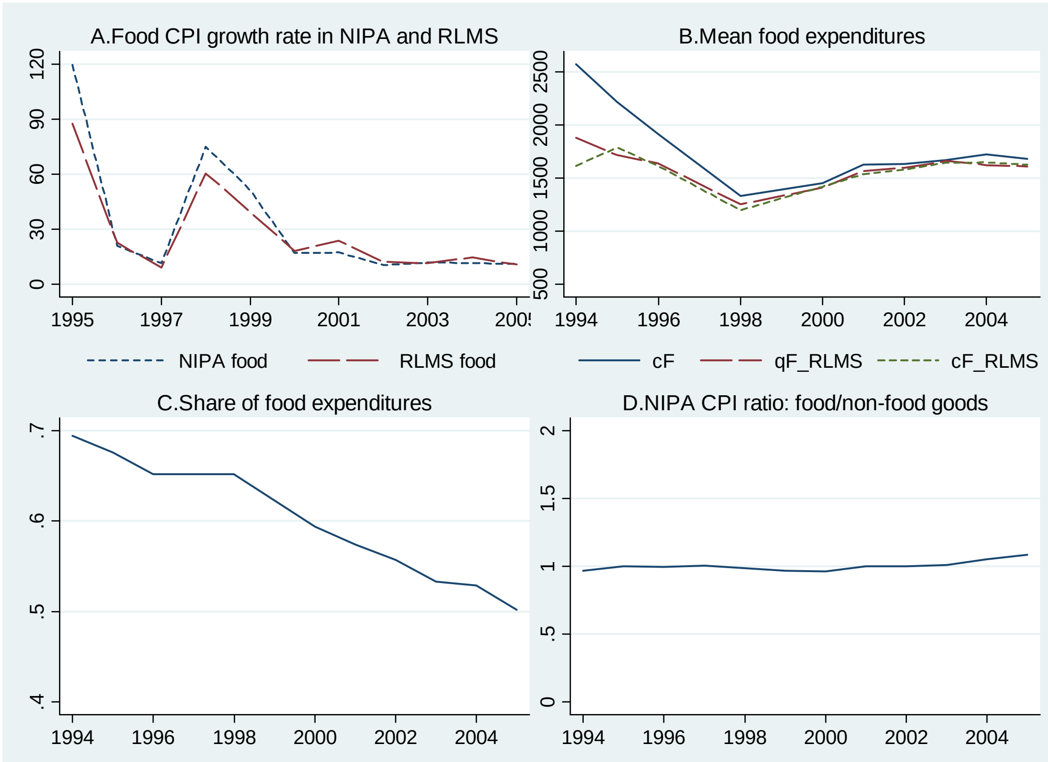
Notes: The figure reports the time series of estimated variance of permanent and transitory components. The estimated process is $u_{ht}=\alpha_{ht} + \varepsilon_{ht}$, $\alpha_{ht} =\alpha_{h,t-1} + \eta_{ht}$, where ε_{ht} is the transitory component and η_{ht} is the permanent component. In all specifications, u_{ht} is the residual from projecting the relevant measure of income or consumption on our baseline vector of observable characteristics of households; *ec* = contractual labor earnings of the household head; *yL* = household contractual labor earnings per month adjusted for non-response; *yD* = disposable household income based on contractual labor earnings; *c* = household non-durable expenditures last month. Values in 1998 and 2000 are adjusted for the fact that the permanent component is accumulated over two years. For both permanent and transitory components, 1997 and 1999 values are set equal to 1998 and 2000 values respectively.

Figure 13: Within-Group and Between-Group Inequality



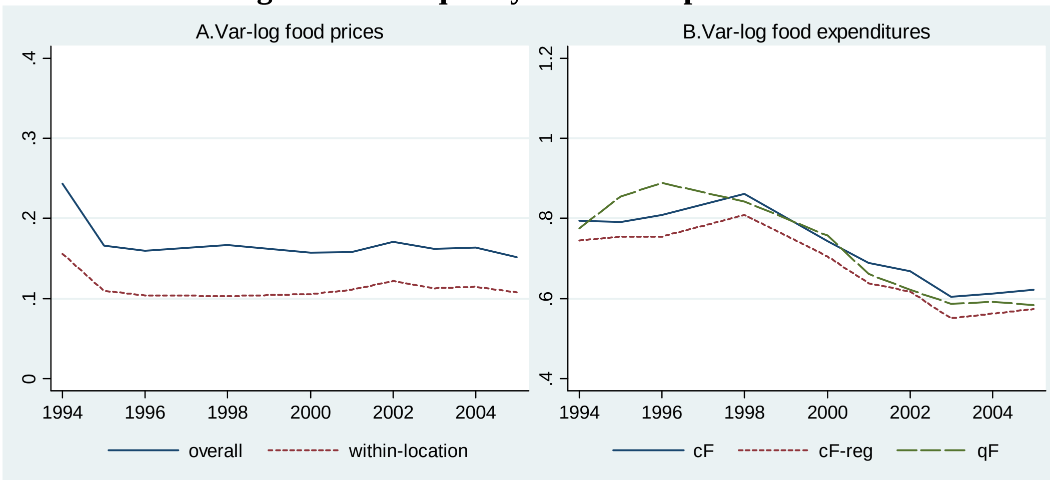
Notes: Rural location is defined as villages and small towns. *yD* = disposable household income based on contractual labor earnings; *c* = household non-durable expenditures last month; *cD* = *c* + expenditures on durables; *cH* = *c* + consumption of home-grown food. All measures are equivalized using an OECD equivalence scale and deflated with regional CPI unless indicated otherwise.

Figure 14: Trends in Food Expenditures



Notes: *cF_RLMS* = expenditures on food, beverages, and tobacco last week (multiplied by 30/7) deflated using national monthly CPI; *cF_RLMS* = expenditures on food, beverages, and tobacco last week deflated using RLMS food CPI; *qF_RLMS* = food quantity index in constant 2002 mean prices for each location. Panel C reports the share of food expenditures *cF* in aggregate consumption expenditures *cD*. All food expenditures are per adult equivalent.

Figure 15: Inequality in Food Expenditures



Notes: *cF* = expenditures on food, beverages, and tobacco last week (multiplied by 30/7) deflated using national monthly CPI; *cF-reg* = *cF* deflated using regional CPI and adjusted for regional differences in cost of living; *qF* = food quantity index in constant 2002 mean prices for each location. All food expenditures are per adult equivalent.

Appendix 1: Data Description

Description of RLMS sample

This study uses ten rounds of the Russian Longitudinal Monitoring Survey (RLMS) that was conducted in 1994-1996, 1998, and 2000-2005. RLMS was not conducted in 1997 and 1999. Time-series reported on the figures are linearly interpolated for missing annual data points.
The RLMS sample consists of the 38 randomly selected primary sample units (municipalities) that are representative of the whole country. They are located in 32 regions (or constituent subjects of the Russian Federation) and 7 federal districts. Russia had 89 constituent subjects and 7 federal districts as of December 1, 2005.

Sample restrictions

We restrict our sample to households in which at least one individual is 25-60 years old. The head of the household in the selected sample is the oldest working-age male or the oldest working-age female if no working-age males are present. If more than one person of the same age-gender is qualified for the head, then the reference person (or the first person surveyed in the roster files) is chosen.

General notes

1. All income variables are *after tax*.

2. All income and consumption variables are constructed on a *monthly basis*.

3. Summary statistics are weighted with individual and household sample weights provided in the RLMS.

4. When a household purchased the item but did not report the amount of the purchase, the missing amounts are imputed by regressing the log of expenditure on the complete interaction between year dummies and federal district dummies, controlling for the size of the household (5 categories), number of children 16 years old or younger (4 categories), number of elderly members 60+ (3 categories), and urban location. Because of the log dependent variable, the predicted values of expenditures are adjusted as $y = exp(\hat{\sigma}^2/2)exp\left(\widehat{\log y}\right)$. The subcategories with the largest number of missing values include utilities (2.12% of the sample), gasoline and motor oil (1.63%), transportation services (1.54%), and contributions to non-relatives (1.35%). Missing values for other subcategories are trivial.

5. Similar regression-based imputations are performed for missing subcategories of non-labor income and income from home production. Imputations of labor income are described in the table below. Although the share of missing values for each individual subcategory of non-labor income and expenditures is very small, altogether missing values affect about a third of surveyed households. Our imputation procedure is an improvement over the existing RLMS practice that treats missing values as zeros in computing aggregate income and expenditures.

Variable description and notes

	Variable Name	Definition	Notes
		Individual Earnings and Labor Supply	
ha	Actual hours of work last month	= hours worked last month at primary job + hours worked last month at secondary job + hours spent last month on regular individual economic activities (activities for which an individual is paid for regularly, such as sewing a dress, assisting with repairs, selling goods in a market or on the street, etc.)	Unusually high hours are top coded at 480 hours per month (16 hours per day*30 days)
hc	Usual hours of work per month	= 4 times usual hours in a typical week at primary job + 4 times usual hours in a typical week at secondary job + hours spent last month on regular individual economic activities.	*hc* is available in 1998-2005 only. Unusually high hours are top coded at 480 hours per month (16 hours per day*30 days).
status	Working status	= *full-time* if actual hours at primary job≥120, *part-time* if actual hours at primary job<120, *not working* if a respondent did not work last month at primary job, was not on a temporary leave, and was not engaged in regular individual economic activities	
ea	Actual labor earnings last month	= money received last month from primary job + money received last month from secondary job +	The variable is highly volatile during the period of wage arrears since a worker may

		money received last month from regular individual economic activities + payments in kind received last month from primary job + payments in kind received last month from secondary job	not receive any money last month or receive back payments for several months at once.
ec	Contractual labor earnings per month	*1998-2005, all employees*: = monthly average (over the last 12 months) after-tax labor earnings of an employee at primary job + money received last month from additional jobs for all employees in 1998-2005 *1994-1996, employees with wage arrears*: = total accumulated wage debt divided by the number of months of overdue wages + money received last month from additional jobs for employees with wage arrears at primary job in 1994-1996 *1994-1996, employees with no wage arrears*: = monetary portion of *wa* for employees with no wage arrears *All years, self-employed*: = monetary portion of *wa* for self-employed (or individuals reporting place of work other than an organization), including those involved in regular individual economic activities in all years.	1. *ec* does not include payments in kind. 2. Average monthly earnings are available for an employee at primary job in 1998-2005. 3. Implausibly low earnings below ½ of the official minimum monthly wage are recoded into missing (0.47% of positive earnings). 4. Implausibly high earnings are also recorded into missing if the residuals exceed five standard deviations from the mean after controlling for occupational categories, hours of work, age, age squared, years of schooling, and individual fixed effects (0.13% of positive earnings). 5. For household aggregation purposes, if a respondent worked last month at least one hour but has missing contractual earnings, missing values are imputed using occupational categories, hours of work, gender, age, age squared, years of schooling, urban location and federal district dummies (the share of imputed earnings is 7.8%).
wa	Hourly wage rate last	= *ea* / *ha*	

	month		
wc	Contractual hourly wage rate	= *ec* / *hc*	*hc* is available in 1998-2005 only; *wc* is calculated for non-imputed earnings
		Household Income	
yLa	Actual labor earnings received last month	After-tax payments received by all household members from all places of work in the form of money, goods, and services in the last 30 days as reported by the reference person of the household.	The variable is highly volatile during the period of wage arrears.
yLc	Contractual labor earnings per month	The sum of *ec* across all individual respondents within the household.	Such aggregation omits those adult household members who did not respond to an individual questionnaire; the response rate for working age individuals within the surveyed household is 96.5%.
yL	Contractual labor earnings per month adjusted for non-response	= *yLc* + imputed contractual labor earnings for working-age non-respondents within the household.	Labor earnings of working-age non-respondents are imputed as predicted earnings times the predicted probability of working using the full set of interactions between the four age groups (18-60) and two gender groups and controlling for urban and federal district dummies for each year separately.
yH	Labor earnings plus income from home production	= *yL* + 0.9*h*, where *h* is average monthly income from home-grown food in the last year defined as the sum of physical quantity of produced food items (minus items given away) multiplied by their mean price in a given region, 0.9 is the assumed labor share of home food production.	Mean prices are obtained in two steps. First, the household-specific market price of individual food item is calculated by dividing the cost of purchase by the amount purchased in the last 7 days. Then the mean price of individual food items is computed for each region (*oblast*) and year.

yL+	Labor earnings plus net private transfers	= *yL* + private transfers received last month – private transfers given to individuals outside the household unit last month.	"Private transfers received" include received alimonies and 11 subcategories of contributions from persons outside the household unit, including contributions from relatives, friends, charity, international organizations, etc. "Private transfers given" include alimonies paid and various contributions in money and in kind given to individuals outside the household unit (6 categories).
y	Household income before government transfers	= *yL* + net private transfers + financial income received last month.	Financial income includes dividends on stocks and interest on bank accounts.
yD	Disposable household income	= *y* + public transfers.	Public transfers include government pensions, state child benefits, stipends, unemployment benefits, and government welfare payments.
		Household Consumption	
cF	Market expenditures on food, alcohol and tobacco	Monthly expenditures on food, alcohol, and tobacco are computed as the sum of expenditures on individual items in the reference week multiplied by 30/7=4.286.	Items include 50 categories of food at home and away from home, alcoholic and non-alcoholic beverages, and tobacco products. See Appendix 2 for details of computation.
qF	Food quantity index	$qF_t = \sum_k \overline{p}_k q_{kt}$, where q_{kt} is the quantity of food item *k* purchased in year *t* and $\overline{p}_k$ is average price of item *k* for each location (psu) in the base year (2002).	
c	Non-durable expenditures	Sum of expenditures on non-durables in the last 30 days. Non-durable items include food, alcohol, tobacco, clothing and footwear, gasoline	

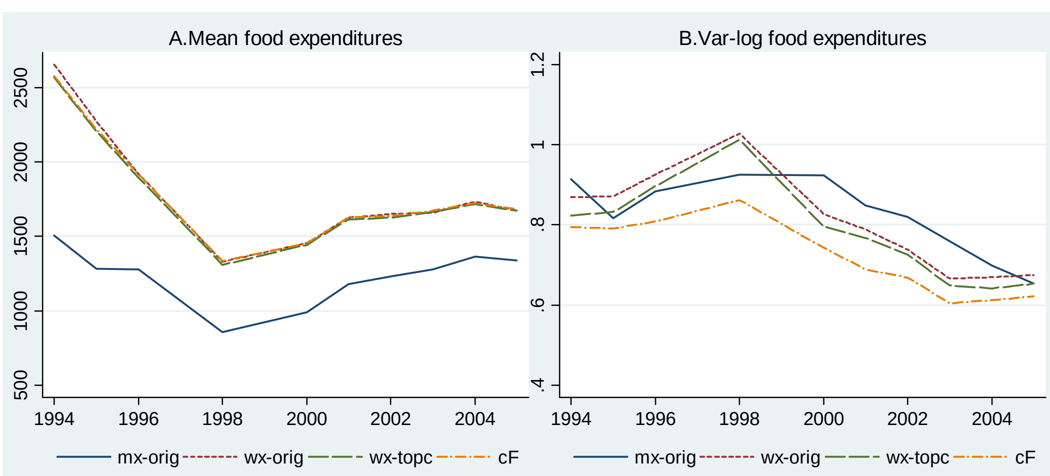
		and other fuel expenses, rents and utilities, and 15-20 subcategories of services (such as transportation, repair, health care services, education, entertainment, recreation, insurance, etc.).	
cD	Aggregate expenditures	= *c* + expenditures on durables in the last 3 months / 3. Durable items include 10 subcategories such as major appliances, vehicles, furniture, entertainment equipment, etc.	This is compared with purchases of goods and services from NIPA
cH	Non-durable expenditures plus consumption of home-grown food	= *c* + consumption of home-grown food, where the last term is calculated as average monthly quantities of consumed home-grown food items multiplied by their mean price in a given region.	Mean prices are determined in the same way as in *yH*.
cD+	Aggregate expenditures plus services from housing	= *cD* + imputed services from housing.	Imputed services from housing are calculated as 5% of the current housing market value divided by 12.
		Adjustments to Income and Consumption	
equiv	OECD equivalence scale	This equivalence scale assigns a value of 1.0 to the first adult household member, a value of 0.7 to each additional adult, and a value of 0.5 to each child 16 years old and younger.	
cpi_t	National monthly CPI	All income and consumption variables are deflated in prices of 2002 using monthly national CPI.	If the date of interview is in the first half of month, the previous month CPI is used. If the date of interview is in the second half of month, the current month CPI is used.
def_t	Regional deflator	Deflator that combines monthly national CPI, December to December regional CPIs, and the	To adjust for monthly inflation, the flow variables are expressed in December prices of

		regional value of fixed basket of goods and services.	each year by using a country average monthly CPI and the date of interview. Next, the annual (December to December) CPI for each 32 oblasts (regions) is applied to convert the flow variables into prices of December 2002. Finally, these real values are adjusted for regional differences in the cost-of-living by using the regional value of fixed basket of goods and services.
$cpiF_{RLMS,t}$	RLMS food CPI	$cpiF_{RLMS,t} = \sum_k p_{kt}\bar{q}_k \big/ \sum_k \bar{p}_k\bar{q}_k$, where p_{kt} denote the sample average unit price of food category *k* in year *t*; $\bar{p}_k$ and $\bar{q}_k$ are the sample average price and the quantity of food item *k* purchased in the base year.	
		Control Variables	
D^H	Household composition	Vector of household composition variables: 4 categories for the number of children 16 years old and younger (0, 1, 2, and 3+), 3 categories for the number of seniors 60 years old and older (0, 1, and 2+), and 5 categories for the number of household members (1, 2, 3, 4, and 5+).	
	Demographics	A female dummy and continuous age variable *a*.	
D^E	Schooling	A set of dummies for educational attainment of the head of household (incomplete secondary, secondary, vocational, technical, and university)	
D^L	Location variables	A set of dummies for 7 federal districts, a dummy for Moscow and St. Petersburg, and a dummy for urban location.	

Appendix 2: Constructing Food Expenditures

This appendix describes the steps in constructing our measure of food expenditures.

1. RLMS food data contain information on the physical quantity and monetary value of last week purchases for 50 categories of food at home and away from home, alcoholic and non-alcoholic beverages, and tobacco products. We first create *wx-orig* as the sum of expenditures on these individual items multiplied by 30/7=4.286. Missing values for this measure are treated as zero.
2. The RLMS questionnaire also asks about the total sum of food purchases in the last 30 days (*mx-orig*). We discard this measure because of a potentially large measurement error, higher probability of underreporting, and ambiguity in the question (e.g., it is likely to exclude beverages and tobacco). We note, however, that the two measures of food expenditures have similar variance (compare *wx-orig* and *mx-orig* in figure below).
3. When a household purchased the item but did not report the quantity of the purchase, the missing quantities are imputed by regressing the log of expenditure on the complete interaction between year dummies and federal district dummies, controlling for the size of the household (5 categories), number of children 16 years old or younger (4 categories), number of elderly members 60+ (3 categories), and urban location. Because of the log dependent variable, the predicted values of expenditures are adjusted as $y = exp(\hat{\sigma}^2/2)exp\left(\widehat{\log y}\right)$. Missing values for food items are generally trivial.
4. We use top coding of unreasonably high prices in excess of 3 interquantile ranges above the mean prices in a given location as well as unreasonably high amounts (quantities) of food purchases (the top 99$^{\text{th}}$ percentile), conditional on the household structure and location. Top coding and imputations does not change the mean value and only slightly reduce the variance (see *wx-topc* in figure below)
5. It is very well known that inequality measures, especially those based on logarithms, are very sensitive to very low values. For that reason, we eliminate the bottom 1% of total food consumption (from purchases and home production) in constant 2002 prices (about 12 percent of the cost of the reference basket of 25 major food items reported by Goskomstat in 2002). While this procedure does not change the mean value of food expenditures, it predictably reduces the variance (see line *cF*).



Notes: All reported measures are per adult equivalent and deflated with national monthly CPI.

Appendix 3: Sample Composition

		Full sample	Restricted sample	Estimation sample
Year:	1994	9.34	9.61	9.66
	1995	8.89	9.09	9.07
	1996	8.82	8.94	8.75
	1998	9.00	9.01	8.91
	2000	9.42	9.24	9.23
	2001	10.64	10.35	10.42
	2002	10.97	10.74	10.81
	2003	11.09	10.92	10.96
	2004	11.07	11.17	11.21
	2005	10.75	10.92	10.99
Region:	Moscow and St. Petersburg	11.28	11.17	11.31
	North West	6.89	7.33	7.37
	Central	19.09	18.17	18.26
	Volga	17.72	17.42	17.39
	South	11.73	12.13	11.93
	Urals	14.17	14.60	14.59
	Siberia	9.41	9.45	9.41
	Far East	9.71	9.73	9.73
Number of household members: 1		18.39	7.58	7.18
	2	27.74	24.28	24.16
	3	25.34	30.83	31.07
	4	18.06	23.49	23.72
	5+	10.47	13.82	13.87
Number of children <16: None		56.99	45.63	45.25
	1	28.26	35.02	35.32
	2	12.23	15.99	16.09
	3+	2.53	3.36	3.34
Urban (excluding small towns)		68.91	69.55	70.01
		42,541	31,969	31,409

Notes: Restricted sample includes households in which at least one individual is 25-60 years old. Estimation sample includes households with non-missing values on disposable income. The sample composition is unweighted.

Appendix 4: Inequality over the Life Cycle

The peculiar age-earnings profile for Russian males with its negative experience premium (Figure 5C) underscores the importance of investigating the life-cycle pattern of inequality. One would like to separate out the age effect on inequality from time effects and cohort effects but these effects are collinear unless one imposes additional restrictions (see Heathcote *et al* 2008). Since none of the restrictions is entirely satisfactory, we present decompositions of age, cohort, and time effects under alternative identifying assumptions.

Suppose that the cross-sectional inequality moment $M(a,t)$ depends on age, a, time, t, and cohort effects, $t-a$, through a linear function. An inequality-age regression can separately identify one of these three effects, and the combined effect of the other two. We first perform inequality-age regressions controlling for time effects and assuming that there are no cohort effects. This specification confounds age effects with cohort effects, and the two cannot be separately identified. We regress the inequality moments on the set of age and time dummies:

$$M(a,t) = \sum_a \beta_a D(a) + \sum_t \beta_t D(t) + \varepsilon_{a,t},$$

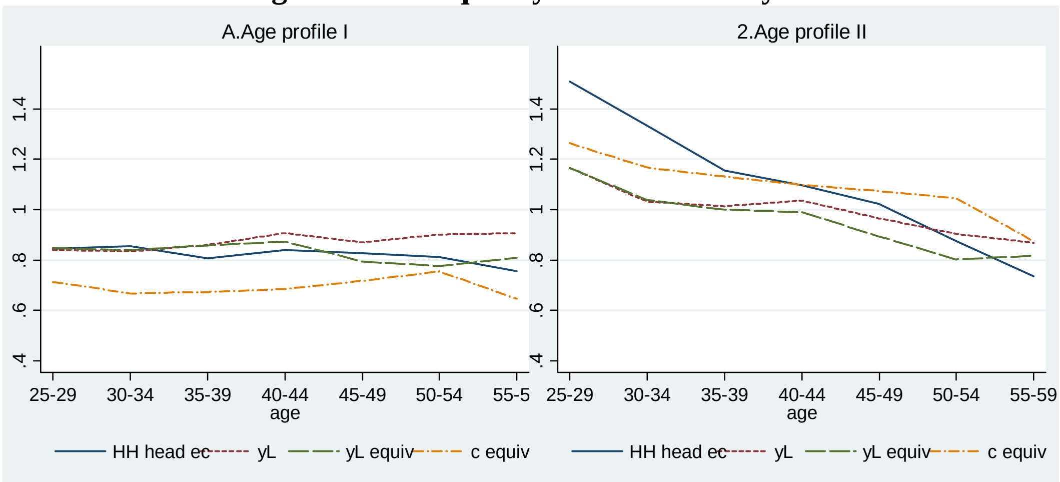
where $M(a,t)$ denote the variance of log income (or consumption) within age group a at time t. Panel A below shows the pattern of age dummies β_a. In almost all cases, the age-inequality profiles are essentially flat, with the exception of a slight decline in inequality among the oldest workers. The flat life cycle inequality profile can be interpreted as age effects and cohort effects roughly canceling each other out. The flat profile of age dummies is consistent with income and consumption decompositions in Figures 7 and 10, where age was found to have almost no explanatory power.

Panel B below reports the age coefficients β'_a from a different specification that assumes away time effects and regresses the cross-sectional inequality moments on age and cohort dummies:

$$M(a,t) = \sum_a \beta'_a D(a) + \sum_{t-a} \beta'_{t-a} D(t-a) + \varepsilon'_{a,t-a}.$$

Now the age-inequality profiles are downward-sloping, because time effects are confounded with age effects. In other words, if income and consumption inequality falls over time for a fixed cohort, the regression model categorizes this as an age effect. Our results potentially point to large time effects on inequality.

Figure A1: Inequality over the Life Cycle



Notes: Panel A depicts age profiles for the var-log controlling for year effects. Panel B depicts age profiles for the var-log controlling for cohort effects. All measures are deflated with national monthly CPI. *HH head ec* = contractual labor earnings per month of the head of household; *yL* = household contractual labor earnings per month adjusted for non-response; *yL equiv* = *yL* equivalized with an OECD equivalence scale; *c equiv* = household non-durable expenditures equivalized with an OECD equivalence scale.

Appendix 5: Time Series Decomposition of Income and Consumption Growth

Permanent-temporary decomposition

The procedure decomposes *residual* variation of income and consumption $u_{ht}^{(s)}$ into temporary and permanent components, where *s* denotes a measure of income or consumption. Using the notation in the body of the paper, the residual $u_{ht}^{(s)}$ from regression (1) can be decomposed into the sum of a transitory component and a random-walk permanent component:

$$u_{ht}^{(s)} = \alpha_{ht} + \varepsilon_{ht},$$
$$\alpha_{ht} = \alpha_{h,t-1} + \eta_{ht},$$

where $\varepsilon_{ht} \sim (0, \sigma_{\varepsilon,t}^2)$ is the transitory component and $\eta_{ht} \sim (0, \sigma_{\eta,t}^2)$ is the innovation in the permanent component.

Given $u_{ht}^{(s)}$, we form a vector of changes in the residual $\Delta u_{ht}^{(s)} = \eta_{ht} + \varepsilon_{ht} - \varepsilon_{h,t-1}$ (that equals the annual growth rate of s_{ht}). The full vector of growth rates for household *h* and measure s_{ht} is $g_h = [\Delta u_{h,1}^{(s)}\ \Delta u_{h,2}^{(s)}\ ...\Delta u_{h,T}^{(s)}]'$, where t = 0 is the first year in the panel and *T* is the last. The covariance matrix of vector g_h, which has $T(T-1)/2$ unique empirical moments, is

$$V \equiv \tfrac{1}{H}\sum_{h=1}^{H}(g_h - \overline{g})(g_h - \overline{g})'$$

where $\overline{g} = \tfrac{1}{H}\sum_{h=1}^{H} g_h$ is the average value of the change in the residual and *H* is the number of households in the sample.

Let Λ be the vector of parameters we to be estimated (i.e., the year-specific variances of innovations in permanent and transitory components of s_{ht}) and let $V(\Lambda)$ be the corresponding covariance matrix. Under the assumptions of our statistical model,

$$V(\Lambda) = \begin{bmatrix} \sigma_{\eta,1}^2 + \sigma_{\varepsilon,1}^2 + \sigma_{\varepsilon,0}^2 & -\sigma_{\varepsilon,1}^2 & 0 & \cdots & 0 & 0 \\ -\sigma_{\varepsilon,1}^2 & \sigma_{\eta,2}^2 + \sigma_{\varepsilon,2}^2 + \sigma_{\varepsilon,1}^2 & -\sigma_{\varepsilon,2}^2 & \ddots & 0 & 0 \\ 0 & -\sigma_{\varepsilon,2}^2 & \sigma_{\eta,3}^2 + \sigma_{\varepsilon,3}^2 + \sigma_{\varepsilon,2}^2 & \ddots & 0 & 0 \\ \vdots & \ddots & \ddots & \ddots & \ddots & \vdots \\ 0 & 0 & 0 & \ddots & \sigma_{\eta,T-1}^2 + \sigma_{\varepsilon,T-1}^2 + \sigma_{\varepsilon,T-2}^2 & -\sigma_{\varepsilon,T-1}^2 \\ 0 & 0 & 0 & \ddots & -\sigma_{\varepsilon,T-1}^2 & \sigma_{\eta,T}^2 + \sigma_{\varepsilon,T}^2 + \sigma_{\varepsilon,T-1}^2 \end{bmatrix}.$$

Two identification issues are apparent from the above expression for $V(\Lambda)$. First, $\sigma_{\varepsilon,0}^2$ is not identified separately from $\sigma_{\eta,1}^2$. Second, $\sigma_{\varepsilon,T}^2$ is not identified separately from $\sigma_{\eta,T}^2$. We follow the common practice of addressing these identification issues by imposing $\sigma_{\varepsilon,T}^2 = \sigma_{\varepsilon,T-1}^2$ and $\sigma_{\varepsilon,1}^2 = \sigma_{\varepsilon,0}^2$. After imposing these constraints, the vector of parameters to be estimated becomes $\Lambda = \{\sigma_{\varepsilon,1}^2, \sigma_{\varepsilon,2}^2, ..., \sigma_{\varepsilon,T-1}^2, \sigma_{\eta,1}^2, \sigma_{\eta,2}^2, ..., \sigma_{\eta,T}^2\}$.

Vector Λ is estimated by minimizing the distance between theoretical and empirical moments

$$\hat{\Lambda} = \arg\max_{\Lambda}(vech\{V - V(\Lambda)\})'(vech\{V - V(\Lambda)\}),$$

where the weight matrix is set to be the identity matrix.

Estimating consumption response to income innovations

The approach to estimating the response of consumption to income components is similar to that in Blundell *et al* (2008). The procedure uses (auto)covariances of income and consumption growth rates. As before, the residual in the income equation is assumed to follow the process

$$u_{ht}^{(y)} = \alpha_{ht} + \varepsilon_{ht},$$
$$\alpha_{ht} = \alpha_{h,t-1} + \eta_{ht},$$

The residual consumption growth

$$\Delta u_{ht}^{(c)} = \phi_t \eta_{ht} + \psi_t \varepsilon_{ht} + \xi_{ht} - \xi_{h,t-1},$$

is decomposed into the parts: the influence of permanent income innovation, the influence of temporary income innovation, and unobserved household heterogeneity. Let $g_h = \left[\Delta u_{h,1}^{(c)}, \Delta u_{h,1}^{(y)}, ..., \Delta u_{h,T}^{(c)}, \Delta u_{h,T}^{(y)}\right]$ denote the vector of income and consumption growth rates for household *h*. As before, define the empirical covariance matrix

$$V \equiv \tfrac{1}{H}\sum\nolimits_{h=1}^{H}(g_h - \bar{g})(g_h - \bar{g})'.$$

Let Λ be the vector of parameters we to be estimated (i.e., the year-specific variances of innovations in permanent and transitory components of income and transitory components in consumption as well as loadings ϕ_t and ψ_t) and let $V(\Lambda)$ be the vector of theoretical moments (i.e., the model equivalent of V). Under our statistical model, with *T*=3 (for example) we have

$$V(\Lambda) = \begin{bmatrix} \sigma_{\eta,1}^2 + \sigma_{\varepsilon,1}^2 + \sigma_{\varepsilon,0}^2 & \phi_1^2\sigma_{\eta,1}^2 + \psi_1^2\sigma_{\varepsilon,1}^2 & -\sigma_{\varepsilon,1}^2 & 0 & 0 & 0 \\ \phi_1^2\sigma_{\eta,1}^2 + \psi_1^2\sigma_{\varepsilon,1}^2 & \phi_1^2\sigma_{\eta,1}^2 + \psi_1^2\sigma_{\varepsilon,1}^2 + \sigma_{\xi,1}^2 + \sigma_{\xi,0}^2 & -\psi_1\sigma_{\varepsilon,1}^2 & -\sigma_{\xi,1}^2 & 0 & 0 \\ -\sigma_{\varepsilon,1}^2 & -\psi_1\sigma_{\varepsilon,1}^2 & \sigma_{\eta,2}^2 + \sigma_{\varepsilon,2}^2 + \sigma_{\varepsilon,1}^2 & \phi_2^2\sigma_{\eta,2}^2 + \psi_2^2\sigma_{\varepsilon,2}^2 & -\sigma_{\varepsilon,2}^2 & 0 \\ 0 & -\sigma_{\xi,1}^2 & \phi_2^2\sigma_{\eta,2}^2 + \psi_2^2\sigma_{\varepsilon,2}^2 & \phi_2^2\sigma_{\eta,2}^2 + \psi_2^2\sigma_{\varepsilon,2}^2 + \sigma_{\xi,2}^2 + \sigma_{\xi,1}^2 & -\psi_2\sigma_{\varepsilon,2}^2 & -\sigma_{\xi,2}^2 \\ 0 & 0 & -\sigma_{\varepsilon,2}^2 & -\psi_2\sigma_{\varepsilon,2}^2 & \sigma_{\eta,3}^2 + \sigma_{\varepsilon,3}^2 + \sigma_{\varepsilon,2}^2 & \phi_3^2\sigma_{\eta,3}^2 + \psi_3^2\sigma_{\varepsilon,3}^2 \\ 0 & 0 & 0 & -\sigma_{\xi,2}^2 & \phi_3^2\sigma_{\eta,3}^2 + \psi_3^2\sigma_{\varepsilon,3}^2 & \phi_3^2\sigma_{\eta,3}^2 + \psi_3^2\sigma_{\varepsilon,3}^2 + \sigma_{\xi,3}^2 + \sigma_{\xi,2}^2 \end{bmatrix}.$$

Again, there are two identification issues. First, $\sigma_{\varepsilon,0}^2, \sigma_{\xi,0}^2$ are not identified separately from $\sigma_{\eta,1}^2$. Second, $\sigma_{\varepsilon,T}^2, \sigma_{\xi,T}^2$ are not identified separately from $\sigma_{\eta,T}^2$. We impose $\sigma_{\varepsilon,T}^2 = \sigma_{\varepsilon,T-1}^2, \sigma_{\varepsilon,1}^2 = \sigma_{\varepsilon,0}^2$, $\sigma_{\xi,T}^2 = \sigma_{\xi,T-1}^2, \sigma_{\xi,1}^2 = \sigma_{\xi,0}^2$. Thus, our vector of parameters becomes

$$\Lambda = \{\sigma_{\varepsilon,1}^2, \sigma_{\varepsilon,2}^2, ..., \sigma_{\varepsilon,T-1}^2, \sigma_{\eta,1}^2, \sigma_{\eta,2}^2, ..., \sigma_{\eta,T}^2, \sigma_{\xi,1}^2, \sigma_{\xi,2}^2, ..., \sigma_{\xi,T-1}^2, \phi_1, \phi_2, ..., \phi_T, \psi_1, \psi_2, ..., \psi_T\}.$$

We estimate Λ by minimizing the distance between theoretical and empirical moments

$$\hat{\Lambda} = \arg\max\nolimits_{\Lambda}(vech\{V - V(\Lambda)\})'(vech\{V - V(\Lambda)\}),$$

where the weight matrix is set to be the identity matrix.