

# Consumption smoothing in Russia

Geranda Notten

(Graduate School of Public and International Affairs, University of Ottawa)

and

Denis de Crombrughe

(Maastricht University School of Business and Economics)

## Abstract<sup>1</sup>

Using a panel from the Russian Longitudinal Monitoring Survey (1994-2004), this paper investigates to what extent Russian households have been able to maintain their living standards while suffering income shocks. Consumption smoothing is modelled by means of an equilibrium correction mechanism, which disentangles short-run dynamics and long-run equilibrium adjustments. GMM estimation is used to control for individual household effects in the presence of dynamics. Additionally, we differentiate between food and non-food consumption, positive and negative shocks, rural and urban areas, and several levels of poverty risk. We find that dynamics are important in the consumption equation, and that estimates are sensitive to imputation errors in home food production. No strong claims can be made regarding heterogeneity in smoothing behaviour.

## 1 Introduction

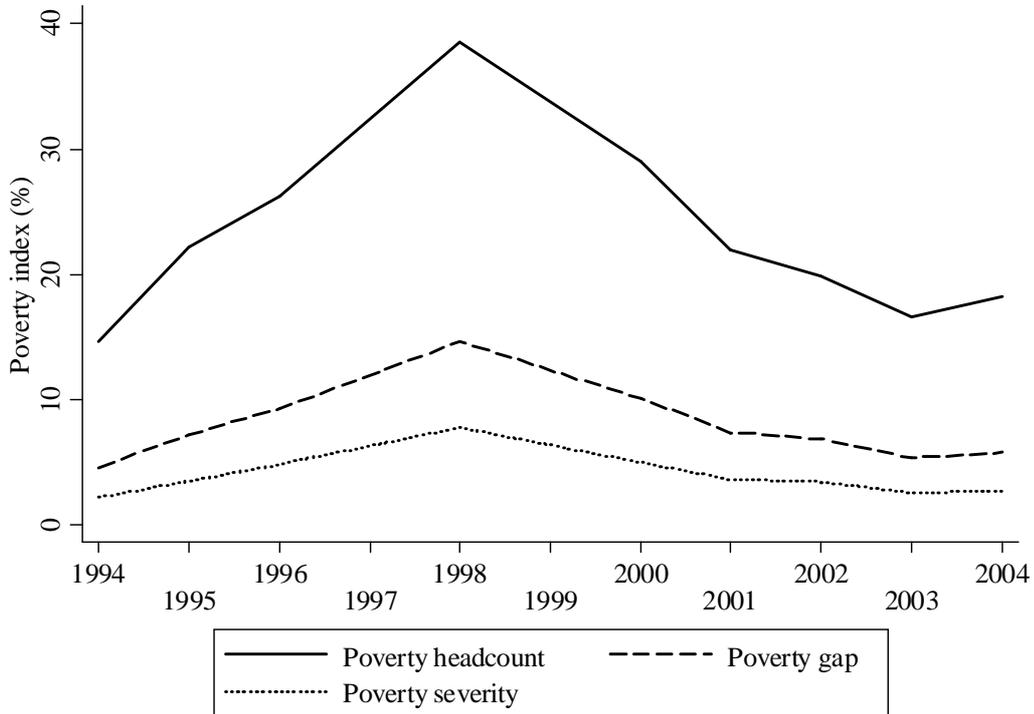
For Russian households after 1990, the transition from a centrally planned economy to a market based economy has been full of bumps, potholes and off the road experiences. Unemployment was a practically unknown phenomenon in pre-transition Russia. The closing down or privatization of the large Soviet-era industrial and agricultural companies resulted in mass unemployment and decreased job security. New skills and habits had to be acquired almost overnight and those still in a job faced wage arrears and forced leave arrangements. The cutting down of subsidies on food and energy caused surges in the cost of living. The fledgling market economy had barely been showing some hesitant signs of recovery when the 1998 financial and economic crisis stroke. It led to defaults on domestic and foreign debts followed by a wave of bankruptcies, a devaluation of the ruble and a collapse of the stock market (Brown, 1999; Buchs, 1999; Sapir, 1999; Slay, 1999). Many households lost (large parts of) their lifetime savings. In addition to the systemic uncertainty accompanying the upheaval, Russians were exposed to the usual range of idiosyncratic risks including illness, disability or death of a household member, crime, job loss or crop failure. Not before 1999 did a new period of sustained economic growth begin, with the restructuring process going on.

---

<sup>1</sup>An earlier version of this article circulated as a working paper (Notten & de Crombrughe, 2007). We are grateful to Chris Elbers, Jan Willem Gunning, Sybrand Schim van der Loeff, Tony Shorrocks, and two anonymous referees for their useful suggestions. Additionally we thank, as required, 'the Russia Longitudinal Monitoring Survey Phase 2, funded by the USAID and NIH (R01-HD38700), Higher School of Economics and Pension Fund of Russia, and provided by the Carolina Population Center and Russian Institute of Sociology for making these data available.' JEL classification: D12, D31, D91, I32.

**Figure 1: Poverty trends in Russia (1994 – 2004)**

Source: RLMS cross-sectional surveys



Perhaps unsurprisingly for such a volatile period, poverty statistics show that many Russian households were unable to maintain their previous consumption levels. We illustrate this with the Russian Longitudinal Monitoring Survey (RLMS, 1994 to 2004, to be described more extensively in the sequel). Figure 1 graphs the evolution of three poverty index series, a poverty headcount, a poverty gap, and a poverty severity measure<sup>2</sup>. The trends in Figure 1 are consistent with those reported by the World Bank (2004). From 1994 the poverty headcount in our panel of households rose to a peak of 38 percent in 1998, after which it steadily declined to about 18 percent in 2004. Clearly, privation was a common experience. About 58 percent of households in our sample experienced at least one episode of poverty between 1994 and 2004, while 10 percent had expenditures below the poverty line *every* time they participated in the survey. Even among households remaining above the poverty line, many had living standards barely higher than subsistence levels; a majority reported average expenditures less than twice the poverty threshold. For many of those, an episode of poverty would have been no more than one adverse shock away.

How do people cope with such a volatile economic environment? There is ample evidence that household consumption expenditures vary less over time than household income. Shocks affecting the earnings of household members do not (and certainly not instantaneously) affect consumption to the same extent. This observation was at the origin of Friedman's (1957) permanent income hypothesis. From a policy-making point of view, the ability of a household to

<sup>2</sup>The three measures belong to the Foster-Greer-Thorbecke (1984) class of decomposable indexes, with power parameter  $\alpha = 0, 1, 2$ , respectively. Consumption expenditures and poverty lines will be defined more precisely in Section 3. The reported poverty measures are calculated in terms of individuals rather than households, and are estimated using the RLMS nationally representative cross-sections. The longitudinal data used for estimation in later sections concern households.

deal with income uncertainty and other risks is an important aspect of its well-being. This is particularly relevant in places where large numbers of people live close to subsistence levels. Townsend (1994) and Ravallion and Chaudhuri (1997) were landmark contributions to the understanding of consumption smoothing processes in the developing world. Recently, thanks to the wide availability of the Russia Longitudinal Monitoring Survey (RLMS), a number of studies have focussed on the case of Russia; see Stillman (2001), Skoufias (2003), Mu (2006), and Gerry and Li (2008). These studies confirmed that only a fraction of all income shocks is transmitted to consumption expenditures.

What smoothing mechanisms do households use? A number of coping strategies is theoretically possible and all of them may be applicable to some extent. First, households may try and maintain their consumption levels by depleting savings, selling assets or durable goods, obtaining credit or borrowing money. Such strategies can be understood as forms of intertemporal ‘self-insurance’ (Skoufias, 2003). Secondly, people facing hardship may obtain help from public institutions and social networks like family, friends, communities and employers, through formal or informal transfer arrangements and insurance policies (Deaton, 1997; Fafchamps and Lund, 2003). The principle on which those arrangements and policies are based is that of ‘risk sharing’ among members of society. Thirdly, people will be stimulated to cultivate any available land, produce basic goods and services in autarchy, adjust their labour supply, or deliver productivity efforts. However, such reactions endeavour to maintain income rather than consumption per se.

Although the empirical literature on intertemporal consumption smoothing is extensive, there are more questions on which it provides no closure. One such question is how consumption smoothing in the short run relates to the necessity of balancing the accounts in the long run. There is a tendency for empirical analysis to focus on the instantaneous responses of consumption to shocks, apparently on the presumption that households adjust at once and that the adjustment does not depend on previous income and consumption levels or accumulated debts. This is not a realistic simplification. When faced with a decline in income, households may be able to maintain their consumption standards for a while in the hope that the decline is reversed, but if that hope proves vain the imbalance will become unsustainable. At some point in time savings and assets will be depleted, loans will have to be repaid, assistance from family and friends will dry up, and consumption expenditures will ultimately have to be adjusted in accordance with the permanent income stream.

A second incompletely answered question in the literature concerns heterogeneity. Exactly how much do households differ in their ability to smooth out income shocks, in particular as a consequence of unequal access to financial markets or social networks? Previous papers have shown how the individual resources and possessions of a household and the surrounding social structures and institutions interact to determine its scope for risk management (Banerjee and Duflo, 2007; Notten, 2008). For instance, the lack of financial or physical collateral will constrain many households in their access to official loans. Some of these may still be able to obtain advances from relatives on the basis of ‘social capital’ (kinship). Regardless of the availability of collateral, credit rationing may occur if recent institutions are incomplete or do not yet function properly, while old institutions or services are being abolished or drastically reformed. Since individual resources and local environments differ, everyone will be able to overcome shocks in different ways and to a different extent. It is plausible that such differences matter more than ever in a period of economic transition.

A third question which is not exhaustively treated is how to deal with the consequences of irregularities and measurement errors in income and consumption data. Infrequent purchases and the standard attenuation bias are not our main concerns here. In many household surveys, income and expenditure totals contain common imputed components representing the production and consumption of home-produced goods, known as ‘autoconsumption’. (This mainly consists of food, and is the dominant component in the case of subsistence agriculture.) Im-

putations are by definition approximations and, since they occur on both the income and the expenditure sides of the equation, any errors of appreciation they include will tend to exaggerate the estimated income elasticities (hence offsetting the classical attenuation bias). Moreover, imputation errors are likely to be serially correlated.

This paper proposes some new answers to those three questions. Using the economic transition in Russia as a case study, our aim is to contribute to the broader empirical literature on consumption smoothing behaviour. In a nutshell, the object of this literature is to find out what determines the ability of households to protect their ordinary consumption expenditures from systemic and idiosyncratic income shocks. Our main points are the following.

First, we investigate whether shock response behaviour is instantaneous, as implicit in previous contributions, or whether it has repercussions over a longer period. To introduce the possibility of delayed adjustments we propose an equilibrium-correction model (ECM) of consumption. Although it is linear in the variables, the ECM specification provides for flexible dynamics and introduces an explicit distinction between short-term reactions and long-term equilibrium adjustments.

Secondly, in regard to heterogeneity, we test for differences in consumption smoothing behaviour associated with the following factors: (i) the sign of income shocks (windfalls or setbacks); (ii) the type of environment (rural or urban); (iii) the household's rank in the distribution of living standards (as approximated by the level of consumption expenditures); and (iv) previous experience of poverty spells.

Thirdly, we implement a natural instrumental variable to deal with the issue of imputation errors exaggerating the correlation between income and expenditures and biasing the estimated income elasticities. This approach is general and easily applicable provided, as is frequently the case, the imputed amounts are known.

Beyond these three methodological points, the ultimate objective of the paper remains to assess the factual ability of a particular population of households to steady their consumption expenditures. Our findings confirm that during the transition, Russian households found ways to smooth their *food* expenditures to a considerable extent, especially in the short run. For a 10 percent income shock the adjustment of food expenditures was as low as 1.5 percent in the short run and 2.3 percent in the long run. Previous studies of Russia may have somewhat underestimated short-run smoothing abilities, and did not consider long-run implications. We can draw no comparable conclusions in the analysis of *non-food* consumption because specification tests give unfavourable outcomes.

Our investigation of possible heterogeneity produces only limited evidence that households reacted differently to positive and negative income shocks. The downward adjustment of food expenditures in case of a negative shock was only slightly stronger than the upward adjustment in case of a positive shock. Similarly, we obtain limited evidence that the response of food consumption to income shocks is stronger in urban environments than in rural areas, indicating that smoothing is somewhat easier in the latter. The differences between parameter estimates for the various subgroups in the sample are small, and the hypothesis of structural homogeneity of behaviour is not strongly rejected. Finally, we find that the correction for the endogeneity of income due to errors of approximation in the imputations for home production has a large impact on the estimates.

The remainder of this paper is structured as follows. Section 2 introduces the ECM specification for consumption expenditures and discusses the empirical implications for our econometric model. Section 3 outlines how we constructed the household panel from the available 1994-2004 RLMS data, and how we defined the main variables. Section 4 reports our estimates, including specification tests and sensitivity analyses. Section 5 presents tests for the presence of several possible types of heterogeneity in consumption smoothing abilities. Finally, in Section 6, we try to assess the significance of our findings for the comprehension of the dynamics of consumption.

## 2 The adjustment process of household consumption

### 2.1 Background

The notion of consumption smoothing is not entirely clear-cut. The most straightforward interpretation refers to the empirical finding that consumption is less volatile over time than income. However, the term is used more often to indicate a degree of insulation of consumption from income shocks, which depends as much on the covariance between consumption and income as on their variances. The term can also describe deviations from an efficient process of intertemporal consumer optimisation in which rational adjustments are effectuated immediately and no predictable changes are delayed (Hall, 1978). In any case, to analyse smoothing we need to make a distinction between short-run and long-run adjustment. In the long run spending is constrained by a life-time budget constraint. In the short run (assuming convex preferences) consumers wish to level their living standards over time.

In this paper, consumption smoothing refers to the empirical ability of a household to smooth its consumption expenditures, that is, the extent to which it actually succeeds in isolating its consumption expenditures from all kinds of recorded income shocks.

The premise of many consumption smoothing models is that households display precautionary behaviour, motivated by the threat of being unable to satisfy the needs of its members in the future or the desire to maintain a minimal living standard under all foreseeable circumstances. In a world with complete markets and perfectly functioning financial institutions, households would have the possibility to formally insure against income risks. In the real world, and certainly in a country in transition, households can insure their living standards very partially at best.

Disruptions may affect household (real) income either directly or indirectly. Job loss, wage payment arrears and involuntary (unpaid) leave have a direct impact on household income. Events such as illness, disability or death of a household member also have a direct impact when the person involved was actively participating in income generating activities. Even when such events involve non-active family members there might be an indirect impact on the household income through adjustments in the internal household task division. Active members may have to reduce labour supply so as to free time for caring activities or other household tasks. Extremes in climate (like droughts or floods), epidemic diseases, crop failures, apart from the direct damage they may cause to households, often also reduce the amount of home produce that can be sold or consumed. Inflation or price adjustments affect the real value of consumption that can be attained, particularly if (as is usually the case) income does not or not fully compensate for price increases.

The empirical measurement of consumption smoothing abilities in developing countries has been inspired by two strands of intertemporal consumer choice models, often labelled risk-sharing and permanent-income or life-cycle models (Deaton, 1992, 1997). In the first strand of models opportunities for risk sharing between risk adverse consumers arise because their income distributions differ across 'states of the world'. Depending on the formalisation of the mechanism, the risk sharing may be achieved either through support networks within a community (kinship, social network, village economy) or by means of a market on which state contingent claims (so-called Arrow securities) are traded. In the case of full sharing, household consumption will grow at the same rate as that of the community. Shocks to a household's resources, be they positive or negative, will be absorbed in the common pool. This yields the testable 'perfect insurance' hypothesis that, when controlling for changes in community resources, household consumption should be unresponsive to income changes. While general prosperity will benefit all members, the insurance community will be powerless when faced with aggregate shocks (an economic crisis, a drought, the bankruptcy of a large firm). Such risk sharing models are the theoretical basis underlying many empirical studies of consumption smoothing (among others Altonji *et al.*,

1992; Cochrane, 1991; Mace, 1991; Townsend, 1994; Ravallion and Chaudhuri, 1997), including some concerning Russia in particular (see below).

The second strand of models derives from Milton Friedman’s permanent income theory, which postulates that consumption is determined by the present value of life-time resources. Forward-looking consumers use savings and credit markets to stabilise their consumption over future time periods (see, for instance, Deaton, 1992). The testable implication is that anticipated income changes should not affect consumption at all. Only unforeseen shocks (‘innovations’) may have an effect on consumption, depending on how persistent they are perceived to be. If a shock is transitory, consumption will hardly be affected; if a shock is permanent, consumption will change commensurately as consumers review their life-time consumption plans.

Friedman’s permanent income model has been extended in various ways in order to allow for the presence of market failures such as liquidity constraints, credit rationing, or information problems blocking access to credit for certain groups of consumers. Indeed, liquidity constrained or low wealth consumers may find themselves unable to buffer consumption expenditures, even with respect to anticipated income changes. Buffer stock models predict that some liquidity constrained consumers faced with a negative income shock could even reduce consumption by an amount exceeding the shock because their aim is to maintain a sufficient precautionary stock of cash. The closely related life-cycle models have been used to explain not only consumption smoothing over the life-time, but also smoothing patterns over shorter intervals such as agricultural production cycles (within a year), business cycles (year to year), and working life (Browning and Crossley, 2001). In the case of Russia in transition, a careful empirical implementation of the permanent-income, life-cycle model is that of Stillman (2001).

## 2.2 Static specification

As emphasised by Stillman (2001), a direct test of the validity of the permanent-income, life-cycle model would require either unrealistic simplifications or information that is not available. Therefore, attempts to measure smoothing behaviour typically rely not on full structural specifications but on relatively simple (linear) reduced form models. Deaton (1992, 1997) reviews earlier contributions in that area. We review here only the main empirical specifications that have been used in the recent studies of Russian consumption smoothing behaviour by Stillman, (2001), Skoufias (2003), Mu (2006), and Gerry and Li (2008).<sup>3</sup>

These studies share a similar basic behavioural equation which can be summarised as follows:

$$\ln c_{i,t} = \beta_1 \ln y_{i,t} + \sum_{j=1}^J \gamma_j x_{j,i,t} + \sum_{\ell=1}^L \delta_\ell D_{\ell,i,t} + \nu_i + \varepsilon_{i,t} \quad (1)$$

where  $\ln c_{i,t}$  and  $\ln y_{i,t}$  denote the logarithms of consumption and income per capita, respectively, for household  $i$  in period  $t$ ;  $x_{j,i,t}$ ,  $j = 1, \dots, J$ , denote so-called ‘taste shifters’, household characteristics affecting needs and hence the marginal utility of consumption;  $D_{\ell,i,t}$ ,  $\ell = 1, \dots, L$ , are binary indicator (dummy) variables identifying each local community  $\ell$  separately in each survey round  $t$ ;<sup>4</sup>  $\nu_i$  is an unobserved household-specific individual effect; and  $\varepsilon_{i,t}$  is an independent ‘idiosyncratic’ white noise (or, at least, serially uncorrelated) disturbance term. The income elasticity of consumption,  $\beta_1$ , is the key parameter measuring the extent (or rather the lack)

<sup>3</sup>These studies all use data from the Russia Longitudinal Monitoring Survey (RLMS) albeit with differences in the included waves, households, variables and definitions. For space reasons we limit our discussion to those aspects that are relevant to the purpose of this paper.

<sup>4</sup>It is not obvious exactly how to give an operational definition of the notion of ‘local community’. One convenient way is to identify the local community with one of the survey’s primary sampling units. In the RLMS, each of these comprises a geographical area that is roughly equivalent to an administrative ‘raion’ in the Russian Federation.

of consumption smoothing. In the case  $\beta_1 = 0$ , consumption is entirely insensitive to income changes and in this sense all income shocks are ‘smoothed out’. At the other extreme, if  $\beta_1 = 1$ , income changes affect consumption expenditures proportionally; in this case no smoothing takes place at all over time or across households, at least at the frequency of observation. The interpretation of the empirical relationship will depend on whether we view smoothing behaviour as resulting from risk-sharing or from life-cycle behaviour or a combination of both. Unless the structural behaviour underlying equation (1) can be identified, there is no way that unambiguous evidence in favour of a single smoothing mechanism can be obtained. The interest of equation (1) is that it provides a framework to estimate the behavioural (in)sensitivity of consumption to income changes, which can be interpreted as the extent to which households are able to isolate consumption from income shocks, whatever the means.

### 2.3 Risk-sharing interpretation

Under the *risk sharing* interpretation of smoothing behaviour, the theory would be that a household’s consumption grows or declines at the same rate as that of its risk-sharing community. The dummies  $D_{\ell,i,t}$  capture effects specific to local community  $\ell$  at time  $t$ , hence they control for community-level uninsurable shocks. The taste-shifters  $x_{j,i,t}$  measure household characteristics affecting consumption needs, such as demographic composition. An estimate of  $\beta_1$  can be used to test the full insurance hypothesis. If risk-sharing institutions function perfectly then idiosyncratic income shocks affecting a household should leave its consumption entirely unaffected ( $\beta_1 = 0$ ). A positive income elasticity ( $\beta_1 > 0$ ) indicates less than full insurance.

While the precise model specifications and estimation strategies vary, the three papers by Skoufias (2003), Mu (2006), and Gerry and Li (2008) test and reject the full insurance hypothesis under the risk-sharing interpretation. We will discuss them briefly in this subsection. For this purpose, consider a first-differenced version of equation (1):

$$\Delta \ln c_{i,t} = \beta_1 \Delta \ln y_{i,t} + \sum_{j=1}^J \gamma_j \Delta x_{j,i,t} + \sum_{\ell=1}^L \delta_\ell \Delta D_{\ell,i,t} + \Delta \varepsilon_{i,t}. \quad (2)$$

First-differencing is one straightforward way to eliminate the unobserved household effects  $\nu_i$  from equation (1) without making restrictive assumptions about their dependence on the observed household characteristics.

Skoufias (2003) differenced incomes and expenditures in equation (1), eliminating  $\nu_i$  and replacing  $\varepsilon_{i,t}$  by its first difference  $\Delta \varepsilon_{i,t}$ , but he did not difference the  $x_{j,i,t}$  and  $D_{\ell,i,t}$  terms and so estimated a variant of the following type:

$$\Delta \ln c_{i,t} = \beta_1 \Delta \ln y_{i,t} + \sum_{j=1}^J \gamma'_j x_{j,i,t} + \sum_{\ell=1}^L \delta'_\ell D_{\ell,i,t} + \Delta \varepsilon'_{i,t}. \quad (3)$$

Estimating this by Pooled OLS, Skoufias found an income elasticity of total consumption expenditures of 0.20. Instrumenting income changes using household-level shock indicators (wage payment arrears, forced leave and unemployment) he obtained a slightly higher estimate of 0.23 (with some indication that the higher estimate was due to urban households only). Replacing the (undifferenced) community dummies ( $D_{\ell,i,t}$ ) by the *change* in average community income resulted in a very similar elasticity of 0.24 with respect to average community income. Splitting consumption into food and non-food expenditures suggested that the community income effect was due exclusively to food expenditures. Skoufias concluded that he found ‘strong evidence of partial insurance and community risk sharing in food consumption’ (p.79).

Mu (2006) used a demeaning (‘within’) transformation to eliminate the  $\nu_i$ , and instrumented income with a refined set of household-level shock indicators. His fixed-effects instrumental-variables estimate of the income elasticity of total household expenditure came out practically

identical to that of Skoufias at 0.23. He further expanded equation (1) with additional terms interacting income with indicators of initial wealth and human capital (education) levels. These terms revealed some differences in smoothing ability between households with different stocks of physical assets and levels of education, at least after the panel was split in rural and urban samples. Wealth seemed to increase the smoothing ability of rural households significantly, whereas education did the same in urban households. For instance, Mu estimated income elasticities of 0.16 for urban households with primary or lower schooling, and 0.26 for rural households with low wealth levels; whereas in favourable cases elasticities as low as 0 were found.

Gerry and Li (2008) used the same differenced form of equation (1) as Skoufias, but then replaced income change in the specification by a combination of dummy variables indicating several types of idiosyncratic shocks (such as unemployment, wage arrears, forced leave) and coping mechanisms available to households (such as pensions and social benefits, sale of assets, help from relatives and home production of food). They focussed on the three RLMS rounds surrounding the Russian financial crisis of 1998, namely those of 1996, 1998 and 2000. They did not report an overall income elasticity of consumption but tested the full insurance hypothesis through the various estimated shock and response effects, using pooled OLS as well as quantile regression. The shock variables they included had large and statistically significant effects; for instance, the 1998 crisis depressed household consumption by over 60 percent, an unemployment spell by over 10 percent, and payment arrears by over 7 percent. Like those of the previous studies, these results seem to imply both a clear rejection of the full insurance hypothesis and the active use by households of available smoothing mechanisms.

## 2.4 Life-cycle interpretation

In contrast to those three contributions, Stillman (2001) based his tests on a permanent income, life-cycle hypothesis, and accordingly did not include the community dummies ( $D_{\ell,i,t}$  terms) in the model. Distinguishing between anticipated and unanticipated as well as between permanent and transitory income shocks, he focussed on unanticipated, transient, aggregate shocks by instrumenting income changes with a set of exogenous shock variables (oil prices, exchange rates, and community-level variations in wage and pension arrears). The model was estimated by 2SLS after a demeaning (within) transformation to eliminate the  $\nu_i$ . The estimates of the income elasticity of food and total non-durable expenditures were very high, ranging from 0.70 to 1.10, implying little smoothing and a strong rejection of the permanent income hypothesis. Remarkably, poorer households faced with an income shock not only reduced their consumption but simultaneously increased their savings. Since such precautionary behaviour is not predicted by the standard life-cycle model, Stillman interpreted it as the possible outcome of a buffer stock model in which credit-constrained consumers deal with the uncertainty of future incomes by achieving a target level of cash savings.

The common conclusion emerging from the studies reviewed here is that Russian households in the late 1990s found some ways to smooth their consumption but they still remained vulnerable to many shocks. A common limitation is that they focus on the instantaneous responses of consumption to income, implying that households understand the exact nature of shocks as they occur and adjust their behaviour independently of previous income and consumption levels. This is a convenient but bold simplification. Households will need time to find out the transient or definitive character of a shock. While they may be able to maintain their consumption standards temporarily in the hope that the decline is reversed, if that hope proves vain then at some point in time assets will be depleted, loans will have to be repaid, and assistance from family and friends will dry up. Ultimately, consumption expenditures will have to be adjusted and made sustainable for the longer term. We therefore extend the empirical model discussed above with an equilibrium-correction mechanism allowing for delayed adjustments and long-term dynamics.

## 2.5 Equilibrium Correction

Equilibrium-correction models, or error-correction models (ECM) as they were originally called, were introduced in the context of aggregate consumption functions by Davidson *et al.* (1978) and related formally to the concept of cointegration by Engle and Granger (1987). They are representations of autoregressive distributed lag models designed to be easily interpretable in terms of adjustments to long-run economic equilibria and short-run deviations from equilibrium. The specification is general enough to encompass the more restrictive distributed lag and partial adjustment mechanisms. In terms of consumption  $c_{i,t}$  and income  $y_{i,t}$  per capita as defined before, we start from a first-order autoregressive-distributed-lag process including individual effects and observed household characteristics:

$$\begin{aligned} \ln c_{i,t} &= \alpha \ln c_{i,t-1} + \beta_1 \ln y_{i,t} + \beta_2 \ln y_{i,t-1} \\ &+ \sum_{j=1}^J \gamma_j x_{j,i,t} + \sum_{\ell=1}^L \delta_\ell D_{\ell,i,t} + \nu_i + \varepsilon_{i,t}. \end{aligned} \quad (4)$$

An interesting reparameterisation emphasising the ECM interpretation of this process reads

$$\begin{aligned} \Delta \ln c_{i,t} &= \beta_1 \Delta \ln y_{i,t} - (1 - \alpha) (\ln c_{i,t-1} - \theta \ln y_{i,t-1}) \\ &+ \sum_{j=1}^J \gamma_j x_{j,i,t} + \sum_{\ell=1}^L \delta_\ell D_{\ell,i,t} + \nu_i + \varepsilon_{i,t}, \end{aligned} \quad (5)$$

where we have imposed the stability condition  $\alpha < 1$ , and defined

$$\theta \equiv \frac{\beta_1 + \beta_2}{1 - \alpha} \quad \text{for} \quad \alpha < 1. \quad (6)$$

The term  $(\ln c_{i,t-1} - \theta \ln y_{i,t-1})$  in equation (5) can be interpreted as the (relative) deviation of consumption from its equilibrium level in the previous period; as long as this term is nonzero, even in the absence of a current change in income, there is pressure for adjustment (since we imposed  $\alpha < 1$ ). The coefficient  $\theta$  parameterises the income elasticity at equilibrium; it is easily seen to be the cumulative ‘long-run’ response of consumption to income changes, whereas  $\beta_1$  measures only the immediate or ‘short-run’ response. Consistency with a life-time budget constraint implies that  $\theta > 0$ , hence also  $\beta_1 + \beta_2 > 0$ . The coefficient  $[-(1 - \alpha)]$  of the deviation from equilibrium is a negative ‘speed of correction’ parameter. If  $(1 - \alpha)$  is close to 0 adjustment is (very) slow; whereas if  $(1 - \alpha) = 1$  the speed is fast and the lagged adjustment is completed in a single period.<sup>5</sup>

Equations (2) and (3), unlike the ECM in (5), ignore errors of adjustment inherited from the past. Either the necessary readjustments are relegated to the idiosyncratic error term  $\varepsilon_{i,t}$ , disregarding the likelihood of induced serial correlation; or it must be the case that  $\alpha = 1$  (as proscribed here), in which case the data contain no information on any relationship binding income and consumption *levels* in the long run (whether or not such a relationship still makes sense).

In equation (5), we can give consumption smoothing a precise quantitative meaning. Since  $\beta_1$  measures the short-run response of consumption to income changes, the extent of ‘short-run smoothing’ can be defined as  $(1 - \beta_1)$ ; that is the proportion of income shocks that is not transmitted to consumption expenditures within the same year. By analogy, since  $\theta$  distinctly measures a long-run response, one could interpret  $(1 - \theta)$  as the extent of ‘long-run smoothing’; that is, more precisely, the proportion of income shocks that never gets transmitted to

<sup>5</sup> $(1 - \alpha) > 1$  or  $\alpha < 0$  would imply a form of overshooting, but we did not encounter such a situation.

consumption expenditures. However, smoothing is essentially a delaying tactic. A conceptually attractive measure of smoothing behaviour could therefore be based on the delay of adjustment, that is, the time it takes for income shocks to be transmitted to consumption. The *mean delay of adjustment* is easily calculated from an ECM with regular coefficient values, as follows:

$$\bar{\ell}_y \equiv \frac{\alpha\beta_1 + \beta_2}{(\beta_1 + \beta_2)(1 - \alpha)} = \frac{1}{1 - \alpha} \left(1 - \frac{\beta_1}{\theta}\right) \quad \text{for} \quad \beta_1 \geq 0, \beta_2 \geq 0, \beta_1 + \beta_2 > 0, \quad (7)$$

and  $0 \leq \alpha < 1$ .

This is a decreasing function of  $\beta_1$ , and an increasing function of  $\alpha$  and  $\beta_2$  (or  $\theta$ ); in other words, the mean delay parameter is higher the larger the extent of short-term smoothing ( $1 - \beta_1$ ), and the lower the correction speed ( $1 - \alpha$ ). Instantaneous adjustment is only possible if  $\alpha = \beta_2 = 0$ . The closer  $\alpha$  comes to 1, the longer households can delay the adjustment of their expenses to their ‘permanent’ income stream. Hence  $\bar{\ell}_y$  measures how long consumers faced with a crisis can put off tightening their belts, on average.

In ECM models the distinction between short-run and long-run effects is natural, quantifiable and testable. In general, long-run income elasticities are thought to be (substantially) larger than short-run ones, and we may therefore expect  $(1 - \beta_1)$  to be (substantially) larger than  $(1 - \theta)$ . Smoothing behaviour (delaying) is what causes the difference between the short and the long run. In the aggregate and in the cross-sectional dimension, household income and household consumption seem to move more or less proportionally, suggesting unit elasticities in the long run. If it is the case that  $\theta = 1$ , then the mean delay of adjustment  $\bar{\ell}_y$  is clearly the more relevant measure of the extent of household smoothing behaviour.

## 2.6 The GMM solution to the estimation problem

The estimation of the ECM in equations (4) or (5) is not without problems:  $\ln c_{i,t-1}$  is necessarily correlated with the unobserved household effect,  $\nu_i$ , and so is any further lag of  $\ln c_{i,t}$ . As shown by Nickell (1981), this correlation is not obliterated by the usual within transformation (demeaning). Arellano and Bond (1991), Arellano and Bover (1995), Blundell and Bond (1998), and others have developed a large family of consistent and more or less efficient Generalised Method of Moments (GMM) estimators for this type of dynamic panel equation. The attractiveness of the GMM approach depends (among other things) on the efficient use of available instruments, the optimal weighting of the orthogonality conditions between these instruments and the idiosyncratic disturbances, and above all a large sample size in the cross-sectional dimension of the panel.

The household unobserved effects,  $\nu_i$ , are first removed by an appropriate transformation; since the within (demeaning) transformation fails to eliminate dynamic panel bias, the obvious appropriate transformation is first differencing. However, as an alternative, Arellano and Bover (1995) proposed an adaptation of the within transformation which has some advantages over first differencing in panels with data gaps, namely the ‘forward orthogonal deviation’. This is the deviation obtained by subtracting the average of all available *future* observations from the current observation and applying a simple scale factor. We indicate either transformation, first differencing or forward orthogonal deviation, by three floating dots in superscript ( $\overset{\dots}{\cdot}$ ), as in the following transformed version of equation (4):

$$\begin{aligned} \overset{\dots}{\ln} c_{i,t} &= \alpha \overset{\dots}{\ln} c_{i,t-1} + \beta_1 \overset{\dots}{\ln} y_{i,t} + \beta_2 \overset{\dots}{\ln} y_{i,t-1} \\ &+ \sum_{j=1}^J \gamma_j \overset{\dots}{x}_{j,i,t} + \sum_{\ell=1}^L \delta_\ell \overset{\dots}{D}_{\ell,i,t} + \overset{\dots}{\varepsilon}_{i,t}. \end{aligned} \quad (8)$$

If  $\varepsilon_{i,t}$  is white noise and first differencing is used, then  $\overset{\dots}{\varepsilon}_{i,t} \equiv \Delta \varepsilon_{i,t}$  is a first-order moving average process or MA(1), implying first-order serial correlation (though none of a higher order)

in the transformed disturbances. In contrast, when forward orthogonal deviations are used, the transformed disturbances retain the serial uncorrelatedness (and even the white noise property) possessed by  $\varepsilon_{i,t}$ . The lagged dependent variable  $\ln c_{i,t-1}$  is still endogenous in equation (8), but the panel itself provides valid instruments in the form of  $\ln c_{i,t-2}$  and further lags (untransformed). As the longitudinal dimension of the panel increases, more and more lagged forms can be added as instruments, and actually care must be exercised not to let their number (which grows quadratically with  $T$ ) become disproportionate.

Instead of removing the unobserved effects,  $\nu_i$ , by transforming the data, there also remains the possibility of estimating equation (4) as it is in levels, using as instruments lagged differences that are uncorrelated with the  $\nu_i$ . Blundell and Bond (1998) explored the conditions under which such instruments would be valid, and designed a ‘system GMM’ estimator combining moment conditions from both the differenced and the level formulations. It is arguable that this approach would allow a more efficient use of our RLMS data, where the cross-section is large, the time dimension is very short, consumption levels vary up and down without excess persistence, and we expect the levels to be informative.

In principle, the GMM approach outlined here can resolve the ‘dynamic panel bias’ that complicates the estimation of a consumption equation including lagged levels. However, before we proceed, it is opportune to mention another potential source of bias. The income variable  $y_{i,t}$  is certain to be measured imperfectly and likely to contain endogenous adjustments. On both counts, it should not be treated as exogenous and must itself be instrumented. The classical attenuation bias is not our only (or even main) concern here. In the RLMS data, income and consumption both contain a substantial imputed component representing the production and consumption of home-produced goods (mostly food). Even if the reported quantities of home produce are precise, the imputed prices are unlikely to be accurate. Since the imputed component is present on both the left and the right-hand side of the equation, it may cause an upward bias in the current income coefficient (hence offsetting the ordinary attenuation bias). Furthermore it is plausible that such imputation errors are serially correlated. To resolve this problem we propose to instrument the income variable with its non-imputed part (‘income net of imputations’). Clearly, omitting imputed components from the instrument will not resolve any measurement error or endogeneity bias due to non-imputed income and more action will be needed. We will return to this issue in the following sections.

### 3 Data and measurement issues

#### 3.1 Data

Our data base is the Russia Longitudinal Monitoring Survey (RLMS), Phase II, covering the period 1994-2004.<sup>6</sup> The RLMS is known as the first nationally representative random sample of the Russian population (Heeringa, 1997; Zohoori *et al.*, 1998; Mroz *et al.*, 2005). It was constructed as a complex multi-stage probability sample of about 4000 ‘dwelling units’ (addresses) covering a huge territory. There were survey rounds in most years though unfortunately not in 1997 and 1999. Data collection took place almost systematically in the last three months of the year. The year by year poverty measures in Figure 1 were based on the separate (cross-sectional) surveys, but to investigate the dynamics of consumption we exploit the longitudinal dimension of the data base. After the necessary matchings are done our longitudinal sample consists of an unbalanced panel of 7753 households, 5575 of which are usable in the sense that the necessary consumption, income and household composition data have been recorded for at least two consecutive survey rounds.

---

<sup>6</sup>Detailed information is provided on the RLMS website, currently at the following url: <http://www.cpc.unc.edu/projects/rlms-hse>.

A point of concern is that of attrition in the panel. Analyses of the response rates show that these are satisfactory (typically above 80 percent), although less so in Moscow and St-Petersburg, and amongst small households with young members (see Heeringa, 1997). Some households are inevitably lost from the panel as a consequence of moving house, splitting up, or other more or less natural causes of attrition.<sup>7</sup>

The RLMS attempts in a number of ways to prevent nonresponse and other drop-outs from affecting the representativity of the survey sample. From 1996 on, households who moved away from the originally sampled address were tracked if possible to their new address and included in subsequent surveys. ‘Lost’ households were in principle replaced by ‘fresh’ households who moved into the originally sampled address after this was vacated. So-called ‘offspring’ households, newly created when a participating household split up, were also invited to participate. Fresh contingents of households from Moscow and St. Petersburg were added to the RLMS sample precisely to compensate for the relatively high metropolitan drop-out rates.<sup>8</sup> It is therefore reasonably likely that the representativeness of the survey sample was maintained if not improved over the years.

In order to satisfy specific demands on data availability, Mu (2006) constructed a balanced, much more selective sample than ours; when he accounted for selectivity in his sample, he found that an attrition correction based on inverse probability weighting had very little effect on his estimates. Our panel is more inclusive, also incorporating newly observed households from later surveys, on the sole condition that they were observed in at least two consecutive rounds. Therefore we argue that our sample is if anything much less likely than Mu’s to be affected by attrition bias.

As controls for household demographics and other characteristics, our consumption equations will include counts of household members in each of six age-gender categories: children aged 0-6 (denoted  $x_{1,i,t}$ ), children aged 7-18 ( $x_{2,i,t}$ ), males aged 19-60 ( $x_{3,i,t}$ ), females aged 19-55 ( $x_{4,i,t}$ ), males aged 60 and above ( $x_{5,i,t}$ ), and females aged 55 and above ( $x_{6,i,t}$ ).<sup>9</sup> In addition, we use binary dummy variables (denoted generically by  $D_{\ell,i,t}$ ) to indicate the survey round ( $t$ ) and the regional location ( $\ell$ ) of the household’s dwelling.<sup>10</sup> Finally, poverty status was determined by comparing the household’s total expenditures (as defined in the next subsection) to a household-specific poverty line constructed by the RLMS project group. The RLMS poverty lines are based on regional age-and-gender-specific food baskets valued at regional prices, and adjusted for the demographic composition of the household; that is, they take into account the household’s composition and (regional) location.<sup>11</sup>

### 3.2 Composition of income and expenditures

Income data are retrieved from the RLMS-constructed income files; expenditures data from the RLMS-constructed expenditures files. We will now provide details on the construction of our

---

<sup>7</sup>A referee has pointed out that adverse income shocks may contribute to the decision of a household to move house and, hence, to drop out of the survey. Furthermore, Friebel and Guriev (2005) argue that ‘the widespread use of in-kind wages and wage arrears in Russia may be explained as an attachment strategy’ and restrict the mobility of workers; they also present some evidence corroborating this. Though beyond the scope of our paper, such hypotheses deserve attention.

<sup>8</sup>Indeed, in the early years of the RLMS, we find that panel households have somewhat lower income and expenditures than average and that rural areas are somewhat overrepresented, whereas from 1998 on such differences become much less clear-cut (Notten, 2008, pp. 135-137; Mu, 2006).

<sup>9</sup>Deaton (1997, p. 360) discusses the use of household composition variables as ‘taste shifters’.

<sup>10</sup>The RLMS partitions the Russian Federation into eight large regions. A finer indication of location is given by the 39 primary sampling units identified in the survey. In our tests, allowing for individual effects of the primary sampling units did not affect the coefficients of interest in a substantial way.

<sup>11</sup>The RLMS poverty lines are on average about 66 percent of the minimum subsistence level used by the Russian authorities to determine eligibility for means-tested benefits (Notten and Gassmann, 2008, p. 89).

household income and expenditure totals. We mention some data issues affecting both totals before discussing them separately.

First, it should be noted that different income and expenditure categories are based on different recall periods varying from a week for food purchases and a month for wage income, to three months for purchases of durables and one year for the amount of home produced crops. Subsequently all amounts are adjusted to a 30-day basis that broadly corresponds with the monthly income and consumption levels at the time of the interview.

Secondly, the value of food produced and consumed at home ('autoconsumption') is imputed by valuing the reported quantities of foodstuffs at local market prices. In Russia, home-produced food represents an important source of household income and consumption, especially but not only in the Russian countryside, in terms of both the number of households involved and the share of consumption or income (cash and in kind). Home-grown food is therefore included in the income and consumption definitions. This poses a specific measurement error issue since imputation errors will affect both sides of the equation equally.

Thirdly, there have been changes in the RLMS questionnaires over the years. Inconsistencies in questionnaire items may obviously induce spurious changes in consumption and income which will distort the evidence on consumption smoothing. After a thorough comparison of the questionnaires and examination of summary statistics we decided that it was safer to exclude expenditures on 'health' and 'other services' from our aggregates.<sup>12</sup>

On the income side, from a theoretical perspective, one would like to separate genuine shocks from induced income smoothing responses which attenuate the magnitude of the shock. In practice, it is hard to make a clear distinction between genuine shocks and induced household responses. To illustrate, the wage loss due to a household member losing her job could be compensated (partly) by another member working overtime. We did not attempt to identify such compensating behaviour and included all wage earnings in income. Similarly, although income from certain social protection arrangements like sick pay and unemployment benefits may be considered as post-shock income, they are unlikely to be endogenous to consumption and are included in our income definition.<sup>13</sup> Conversely, income from property and jewellery sales, transfers received from friends and relatives, and cash borrowings, were excluded because the use of these resources is more likely to be induced by the need to finance current consumption expenditures. Summing up, we calculate household income as the total household income from salary, rent, interest receipts, investments, pension benefits, unemployment benefits, child allowances, maternity benefits, apartment allowances, stipends, the value of home produced food (cash and in kind) and other income; excluding depletion of assets, property and jewellery sales, transfers received from friends and relatives, and cash borrowings.

On the expenditures side, the problem is to measure current household consumption. The standard ambiguities arise with real estate, durable goods and physical assets since they serve for future as well as current consumption, and may also be used as a means to store value (especially in times of high inflation and exchange rate instability; see Lokshin and Ravallion, 2004). A specific problem concerning housing expenditures is that no informed imputations are

---

<sup>12</sup>In 2000, the expenditure section of the household questionnaire was adjusted, including more detailed questions about expenditures on 'health' and 'other services'. Since that year, the 'health' category represents roughly 5 percent of the total budget, and 'other services' roughly 12 percent. The share of the original 'services' category remained unaffected, at about 14 percent of the budget before and after 2000. Hence it does not seem that introducing the new categories has displaced expenditures from the previously existing 'services' category.

<sup>13</sup>Less than 1 percent of the participants report receiving any unemployment benefits; we checked and found that excluding this source of income would not have a major impact on the estimates. We also note that in Russia during the 1990s, few benefits were means-tested. Since the year 2000 new income-tested benefits have been introduced (e.g., apartment benefits), and some previously existing benefits have become income-tested (e.g., child benefits). Empirical studies suggest that targeting errors are large (Gassmann, 2003; Notten and Gassmann, 2008).

provided by the RLMS for the value of housing to home owners. Finally, it is quite common for Russian households not to pay their rent or utility bills; tenants enjoy strong legal protection and the utility infrastructure makes it difficult to cut off single dwelling units (World Bank, 2003). Thus, unpaid bills seem to function as a sustainable way to make ends meet when cash resources fall short (that is, to relax the liquidity constraint).<sup>14</sup> For these reasons we chose to exclude from our consumption variable all expenditures on durables, luxury goods and housing. We calculated two subtotals: food expenditures, including both cash expenditures on foodstuffs and the estimated value of food produced and consumed at home; and non-food expenditures, which includes expenditures on tobacco, clothing, fuel, services, and meals taken outdoors (considering that these contain a large share of service rather than just food).

Table 1 presents summary statistics for the income and expenditure totals and subtotals (food and non-food) in per capita averages; all amounts are expressed in constant June 1992 ruble prices, and the total of excluded income and expenditure categories are also reported. Clearly, the 1998 crisis had a profound negative impact on household income and expenditures, in all categories. We will return to Table 1 in our discussion of measurement error.

**Table 1: Income and expenditures in the RLMS (1994-2004)**

Year	Number of households	Average Expenditures				Average Income			Ratio
		Total	Food	Non-food		Total	Incl.	Excl.	
				Incl.	Excl.				
1994	3586	3964	2344	943	677	3066	2740	326	1.29
1995	3441	3409	1978	970	461	2764	2161	603	1.23
1996	3320	3040	1630	963	447	2588	2066	522	1.17
1997	-	-	-	-	-	-	-	-	-
1998	3426	2240	1231	663	346	1859	1588	271	1.20
1999	-	-	-	-	-	-	-	-	-
2000	3634	2989	1271	838	880	2284	1865	419	1.31
2001	4039	3569	1380	1024	1165	2684	2302	382	1.33
2002	4213	3733	1352	1113	1268	3071	2559	512	1.22
2003	4295	4076	1337	1193	1546	3502	2766	736	1.16
2004	4246	4079	1327	1174	1578	3519	2986	533	1.16

Source: RLMS yearly survey rounds (cross-sections).

Notes: Entries are monthly per capita averages expressed in June 1992 rubles.

Excluded expenditures consist of the following RLMS categories:

durables, luxuries, rent, health, other services, miscellaneous,

bonds, savings, and other payments (insurance, loans).

Excluded income consists of the following RLMS categories:

property sales, insurance benefits, other benefits, help received

from family and friends, and sale of jewellery.

<sup>14</sup>See Freinkman (1998). A more detailed discussion can be found in Notten (2008, Chapter 6, ‘Managing Risks: What do Russian households do to smooth consumption?’).

### 3.3 Measurement errors

The possibility or rather the practical certainty of measurement errors in income and expenditures has already been mentioned. A first likely source of error is the inclusion in both income and consumption totals of imputed values of ‘autoconsumption’. This consists mostly, and in our case entirely, of food that is home-produced and home-consumed. Households report quantities of domestically produced items and these are subsequently valued at local market prices observed at the time of the survey. Since the same imputed values are included in both income and consumption, any imputation errors will cause a bias away from zero in the current income coefficient (hence reversing the ordinary attenuation bias). As suggested by Deaton (1997), we avoid the upward bias by instrumenting the income and lagged expenditures variables with their non-imputed parts. We point out that this treatment will not resolve any attenuation or endogeneity bias due to non-imputed income.

Previous studies used instrumental variables in order to prevent independent measurement errors (of the traditional variety) biasing tests towards accepting the full insurance hypothesis. Several authors cited earlier used information on wage arrears, pension arrears, forced leave, and unemployment to create instrumental variables for current income, on the premise that these events are associated with unexpected income shocks (Skoufias, 2003; Mu, 2006; Gerry and Li, 2008). We followed the standard GMM logic of exploiting information internal to the panel, and used lagged levels of the imputation-free part of income as instrumental variables for total income changes.<sup>15</sup>

A second measurement error issue concerns inconsistencies due to changes in the survey questionnaires. Some idea of their importance can be gleaned from Table 1. Recall that we excluded expenditures on ‘health’ and ‘other services’ from our consumption totals mainly because of such inconsistencies; we also excluded asset sales and borrowings from our income total because of their potential endogeneity. While the value of excluded income sources is rather modest and stable over time, the value of excluded expenditure rises from 17 percent in 1994 to 39 percent in 2004. Analysis of detailed expenditure categories suggests that this rise can be attributed in large part to the inclusion of new items in the 2001 questionnaire, leading to a higher recall of expenses on health and various services.<sup>16</sup> It is plausible that the extension of the questionnaire led to a substantial improvement in the accuracy of the data; nonetheless, inconsistencies over time are problematic for the estimation of a dynamic model, since variation in expenditure is liable to be interpreted as a response to variation in income. Given the magnitude of the increases, the two affected spending categories were excluded from the non-food expenditure variable.

A third type of measurement error is one that has plagued household budget surveys ever since that of Ducpétiaux, which Ernst Engel analysed in his famous 1857 article, namely the apparent ‘overconsumption’, or underreporting of income. Although not always discussed, let alone explained, this phenomenon seems to be quite common. The total income and expenditure categories in Table 1 give an indication of the scale of the problem in the RLMS. Average total expenditures are higher than average incomes in every survey round, and the implied negative savings represent up to one third of reported income. Such discrepancy suggests a systematic violation of the household budget constraint. In their studies, Atkinson *et al.* (1995), Deaton (1997), and Ravallion (1994) merely attribute the problem to the tendency of respondents to underreport income from informal or moonlighting activities.<sup>17</sup>

There is no detectable coherency between the implied (negative) savings rates and those

---

<sup>15</sup> Guariglia and Kim (2003) argued that the internal instruments should not be considered exogenous.

<sup>16</sup> Results not shown here are available from the authors on request.

<sup>17</sup> To obtain positive saving rates, Denizer *et al.* (2002) and Foley & Pyle (2005) purge the RLMS data of ‘acute dissavers’, i.e., households that report expenditures exceeding income by more than 50 or 100 percent, respectively.

reported in national income statistics for Russia. Official Goskomstat sources put the saving rate above 20 percent in the years 1994 to 1997 and fluctuating between 15 and 20 percent in the years 1998 to 2004. *Russian Economic Trends*, an independent monthly periodical discontinued in 2003, published figures corrected for double counting of hard currency sales that were between 6 and 10 percentage points lower than the Goskomstat rate.

Apart from actual underreporting of incomes there are at least three other explanations for negative balances between income and consumption. One is that of negative savings by (continued) depletion of financial or non-financial reserves, including stocks of durable goods. In view of inflationary pressures, instability in the banking system, and uncertainty about monetary reform, it is possible that durable goods end up being used to store value and play the role of a wealth buffer. The second explanation is that of (increases in) wage arrears, with earnings due but not paid being seen as unspent income, hence a form of forced saving. The third is that inflation, combined with different recall periods for different categories of income and expenditures, may distort the balance. All three phenomena are likely to be common in transition Russia and can actually be traced in the RLMS data. However, all three seem to occur on far too small a scale to explain the observed extent of ‘overconsumption’.<sup>18</sup>

To illustrate the regularity of the problem in our data we graph the logarithmic income and expenditure series for the sequence of survey years in Figure 2. Panel 2A presents the evolution of overall averages, and Panel 2B the evolution of four percentiles (the 20th, 40th, 60th, and 80th) of the frequency distributions of household incomes and expenditures. We interpret the striking parallelism between the two totals as signifying that the extent of income underreporting is relatively constant as a proportion of the actual total. This encourages us to assume that we can treat the implied measurement error in a household’s income as proportionally constant over time. Its effect may accordingly be absorbed by the household individual effect  $\nu_i$  in (1) and (4).

A different check on the likelihood of measurement errors in incomes and expenditures is done by looking at changes in both quantities. There would be no issue of consumption smoothing if these amounts did not vary, but changes by a factor of 10, 100, or even more, raise suspicions. As a conservative measure, for the estimations reported below, we excluded observations implying income or total expenditure shocks beyond the range  $(e^{-5}, e^5)$ .<sup>19</sup>

The preceding discussion acknowledges that the RLMS income and expenditures data are far from perfect. Nonetheless, the RLMS is very seriously managed and constitutes by far the best available source of household information from the Russian confederation. We interpret the fact that the RLMS data have been used in numerous earlier studies as a confirmation of this view.

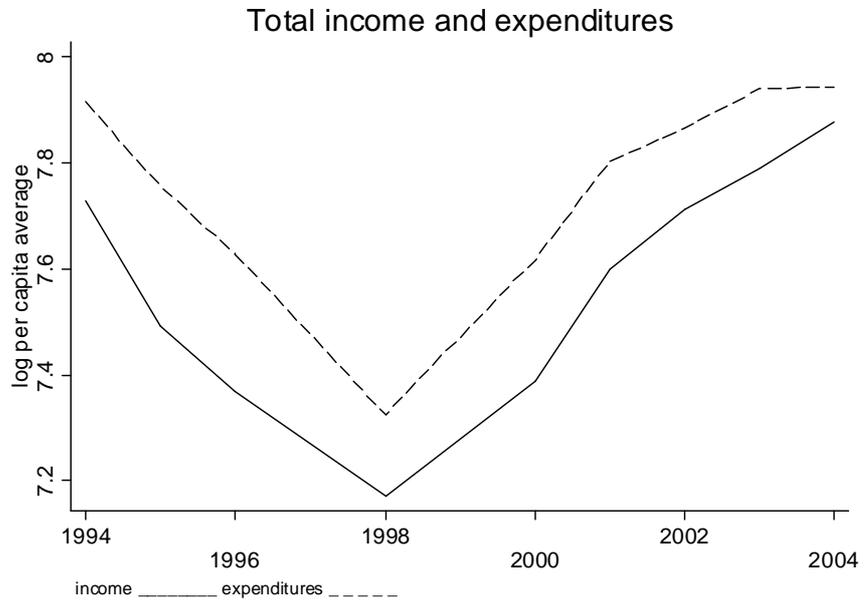
---

<sup>18</sup>Gregory et al. (1999) revise the official saving rate of 28.7 percent for 1994 downward and propose a smaller estimate of 12 percent; this is apparently calculated by imputing wage arrears and counting them as unspent current income in the RLMS.

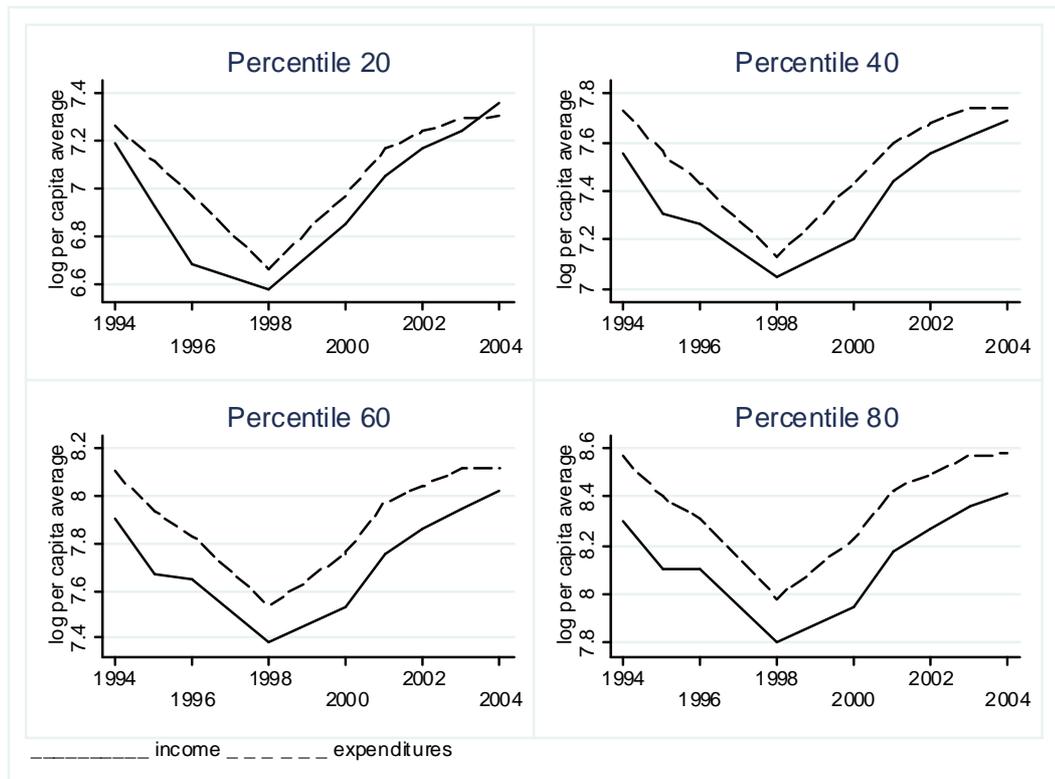
<sup>19</sup>A shock beyond the range  $(e^{-5}, e^5)$  means the change factor is smaller than 1/148 or exceeds 148. Depending on the sample considered, this concerns between 1 and 2 percent of the data.

Figure 2: Evolution of income and expenditures (log per capita)

Panel 2A: Averages



Panel 2B: Percentiles



There are reasons to believe that the quality of the expenditure data is better in the food than in the non-food categories. In comparison with non-food products, foodstuffs are by nature more perishable. In comparison with clothing and other durables and some services, food is usually purchased at regular intervals, much of it in non-lumpy quantities (except for staples), and it can be used (consumed) only once. An important point is that, in contrast to health and other services, the food items in the questionnaire have remained unchanged over the entire period under consideration. Finally, food is a large part of people’s daily needs and evidently the major component of household expenditures with respect to the issue of consumption smoothing. In the sequel, we will therefore focus the discussion primarily, although not exclusively, on food consumption.

## 4 Estimation results and sensitivity analysis

### 4.1 The irregular panel 1994-2004

In the preceding sections we prepared the ground for the estimation of the consumption equation. In this section we report the key estimates, some tests of the validity of the choices made, and some sensitivity analyses. There is a total of nine RLMS rounds available between 1994 and 2004; there are gaps in 1997 and 1999. Exploiting the irregularly spaced data fully is only feasible under stringent random-effects assumptions that are not justifiable here and that we therefore prefer to avoid.

We test our model on two different data sets. First, in this subsection, we use the most complete possible panel, comprising nine irregularly spaced waves (1994-1996, 1998, and 2000-2004, with gaps in 1997 and 1999;  $T = 9$ ). Since the year 1998 is isolated the 1998 data play no direct role in the estimates, although they will serve as instrumental variables when relevant. In the next subsection, we will see how the results are affected when we use the shorter but regular annual panel for 2000-2004 ( $T = 5$ ). Throughout, we will estimate separate models for food and non-food consumption.<sup>20</sup>

Beginning with the extensive data set we present OLS and IV estimates of equations (2) and (8) for food in Table 2; and GMM estimates of equation (8) in Table 3. For non-food, GMM estimates of equation (8) are reported in Table 4.<sup>21</sup>

Columns [i] to [iv] of Table 2 contain estimates of the static model of equation (2) applied to food consumption. This is similar to, although distinct from, the model of Skoufias (2003) and Gerry and Li (2008), who estimated equation (3). Column [i] reports pooled OLS estimates. The short-run income elasticity of food consumption is positive and statistically significant at any conventional level. A 10 percent income shock is predicted to affect the household’s food consumption by 2.3 percent. This compares closely with the pooled OLS estimate of Skoufias (2003), which was 2.0 percent.

---

<sup>20</sup>Let us point out that estimating the model for food and non-food expenditures separately, or for the aggregate of the two, imposes the same strong separability assumptions on the underlying utility functions.

<sup>21</sup>Detailed estimation output, including more tests, are available from the authors on request.

**Table 2: OLS and IV estimates for food expenditures 1994-2004**

	[i]	[ii]	[iii]	[iv]	[v]	[vi]
Method	OLS	IV	IV	IV	IV	IV
$\Delta \ln c_{i,t-1}$					0.064** (0.027)	0.063** (0.026)
$\Delta \ln y_{i,t}$	0.230*** (0.010)	0.240*** (0.017)	0.130*** (0.011)	0.074*** (0.025)	0.062 (0.037)	0.077 (0.053)
$\Delta \ln y_{i,t-1}$						0.011 (0.026)
$\Delta x_{1,i,t}$	-0.153*** (0.022)	-0.151*** (0.023)	-0.173*** (0.023)	-0.183*** (0.023)	-0.182*** (0.033)	-0.180*** (0.033)
$\Delta x_{2,i,t}$	-0.162*** (0.019)	-0.161*** (0.019)	-0.177*** (0.019)	-0.186*** (0.020)	-0.164*** (0.027)	-0.162*** (0.027)
$\Delta x_{3,i,t}$	-0.152*** (0.018)	-0.152*** (0.018)	-0.149*** (0.018)	-0.147*** (0.019)	-0.141*** (0.025)	-0.142*** (0.025)
$\Delta x_{4,i,t}$	-0.208*** (0.021)	-0.206*** (0.021)	-0.225*** (0.021)	-0.234*** (0.021)	-0.221*** (0.028)	-0.219*** (0.029)
$\Delta x_{5,i,t}$	-0.165*** (0.040)	-0.165*** (0.040)	-0.159*** (0.041)	-0.156*** (0.041)	-0.156*** (0.055)	-0.157*** (0.055)
$\Delta x_{6,i,t}$	-0.195*** (0.030)	-0.194*** (0.030)	-0.204*** (0.030)	-0.209*** (0.030)	-0.205*** (0.039)	-0.205*** (0.039)
IVs for:						
$\Delta \ln c_{i,t-1}$					$\ln \check{c}_{i,t-2}$	$\ln \check{c}_{i,t-2}$
$\Delta \ln y_{i,t}$		$\ln y_{i,t-1}$	$\ln \check{y}_{i,t}$	$\ln \check{y}_{i,t-1}$	$\ln \check{y}_{i,t-1}$	$\ln \check{y}_{i,t-1}$
$\Delta \ln y_{i,t-1}$						$\ln \check{y}_{i,t-2}$
IV count		12	12	12	11	12
OID deg.		0	0	0	0	0
Weak ID $F$		2427.8	3831.3	535.1	189.2	88.3
Observ.	18668	18668	18668	18668	10939	10939
Househ.	5575	5575	5575	5575	4336	4336
FD-AR(1)					0.000	0.000
FD-AR(2)					0.759	0.769

Notes: Robust standard errors in parentheses (with clustering by household).  
Significance tests: \*\*\* indicates  $p$ -value $<0.01$ , \*\*  $p$ -value $<0.05$ , \*  $p$ -value $<0.1$ .  
 $c$  is consumption,  $y$  is income,  $x_1 \dots x_6$  are counts of household members in six age-gender categories; survey round dummies are included (not reported).  
The IVs  $\check{y}_{i,t}$ ,  $\check{y}_{i,t-1}$ ,  $\check{y}_{i,t-2}$  denote income and its lags net of imputations.  
The IV  $\check{c}_{i,t-2}$  denotes consumption net of imputations and lagged twice.  
The IV count includes the demographic variables and relevant year dummies.  
OID deg. is degree of overidentification; Weak ID  $F$  is Kleibergen-Paap rank test for weak IV; FD-AR( $r$ ) is test ( $p$ -value) for serial correlation of order  $r$ .  
Source: RLMS 1994-2004 (entire period).

We have already pointed out the reasons for treating the income variable  $y_{i,t}$  as endogenous. Columns [ii], [iii], and [iv] of Table 2 report several versions of instrumental variable (IV or 2SLS) estimates with different instrumentation choices, based on lagged and/or imputation-free levels of income. These are essentially variants of the Anderson and Hsiao (1981) estimators. In column [ii] income changes are instrumented with the lagged total income level (including the imputed component). The lagged income level is an appropriate instrument if income is an accurately measured ‘predetermined’ variable. The instrument allows for endogenous income, in the sense that income changes may contain instantaneous responses to current shocks, but it does not resolve the common imputation error problem. It is not surprising then that, apart from increased standard errors, the instrumental variable estimates in column [ii] hardly differ from the OLS estimates in column [i].

In column [iii] of Table 2, income changes are instrumented with current imputation-free income levels; the effect of a 10 percent income shock drops significantly to 1.3 percent. On the face of it, this is consistent with OLS containing an upward bias due to imputation errors in both income and consumption. In column [iv] income changes are instrumented with the lagged level of income net of imputed components. This instrument allows for endogeneity in contemporaneous income changes (fast responses to current shocks) in addition to the effect of common imputation errors.<sup>22</sup> The effect of a 10 percent income shock now drops further to 0.7 percent. Henceforth, we opt for the last type of instrumental variable.

The last two columns of Table 2 augment the consumption equation with dynamic terms. Column [v] introduces the lagged first-differenced consumption and Column [vi] adds the lagged first-differenced income. The resulting estimating equation is a case of equation (8) where the first-difference transformation is applied; first differences are instrumented with appropriate lagged levels. The short-term income elasticity loses statistical significance although it does not depart in magnitude from the value estimated in Column [iv]. The lagged consumption variable (which is instrumented with a further lag of the imputation-free consumption level) is significant at standard levels in both models; the lagged income variable fails to have a significant effect. Two serial correlation tests due to Arellano and Bond (1991) are calculated from the first-differenced data, labelled FD-AR(1) and FD-AR(2); in line with expectations, they display negative first-order autocorrelation in residual first differences, but no second-order autocorrelation, giving no indication of dynamic misspecification. However, the inclusion of the dynamics in the model comes at a cost. As more than 40 percent of the observations is lost, the IV estimator loses precision.

Also reported in the table is the number of instruments (including regressors treated as exogenous), the degree of overidentification, and the Kleibergen-Paap (2006) rank test statistic for weak identification; the latter is high in all columns, indicating that the instruments are far from weak.

---

<sup>22</sup>It would take another lag to resolve classical measurement error.

**Table 3: GMM estimates for food consumption expenditures**

Sample period	1994-2004		2000-2004	
Column	[i]	[ii]	[iii]	[iv]
$c_{i,t-1}$	0.093*** (0.027)	0.125*** (0.033)	0.074** (0.031)	0.110*** (0.038)
$y_{i,t}$	0.097*** (0.026)	0.151*** (0.041)	0.095** (0.041)	0.221** (0.088)
$y_{i,t-1}$	- -	0.048* (0.029)	- -	0.055 (0.034)
$x_{1,i,t}$	-0.198*** (0.022)	-0.176*** (0.025)	-0.161*** (0.030)	-0.129*** (0.036)
$x_{2,i,t}$	-0.145*** (0.015)	-0.130*** (0.017)	-0.148*** (0.025)	-0.128*** (0.028)
$x_{3,i,t}$	-0.130*** (0.015)	-0.134*** (0.015)	-0.135*** (0.023)	-0.135*** (0.022)
$x_{4,i,t}$	-0.175*** (0.017)	-0.166*** (0.018)	-0.202*** (0.026)	-0.182*** (0.029)
$x_{5,i,t}$	-0.099*** (0.028)	-0.105*** (0.027)	-0.158*** (0.044)	-0.163*** (0.044)
$x_{6,i,t}$	-0.174*** (0.025)	-0.170*** (0.024)	-0.219*** (0.037)	-0.221*** (0.036)
Observation count	13286	13286	8994	8994
Household count	4464	4464	3822	3822
IV count	35	35	19	19
Hansen $p$ -value	0.201	0.279	0.345	0.506
FD-AR(1) $p$ -value	0.000	0.000	0.000	0.000
FD-AR(2) $p$ -value	0.469	0.268	0.661	0.402

Notes: Robust standard errors in parentheses (clustering by household).  
Significance tests: \*\*\* indicates  $p$ -value $<$ 0.01, \*\*  $p$ -value $<$ 0.05, \*  $p$ -value $<$ 0.1.  
The data levels are transformed using forward orthogonal deviations.  
 $x_1 \cdots x_6$  are counts of household members in six age-gender categories.  
Year (survey round) dummies are included though not reported.  
The IV count includes the demographic variables and relevant year dummies.  
FD-AR(1),(2):  $p$ -values of Arellano-Bond tests for serial correlation of orders 1 and 2.  
Source: RLMS 1994-2004.

The ECM equation (8) is re-estimated in the first two columns of Table 3 using the optimal (relatively efficient) GMM method. Columns [i] and [ii] present the result without and with the lagged income term, respectively. A first efficiency gain should be achieved by using forward orthogonal deviations rather than first differences in equation (8); this reduces the loss of observations from over 40 percent to 29 percent. A second source of efficiency gain is the use of additional instruments; the number of instrumental variables (including those of the

GMM-type) increases from 12 to 35. A third improvement is expected from using an appropriate (theoretically optimal) weighting matrix for the different moment conditions involving the various instruments.<sup>23</sup> Both current income and lagged consumption are now statistically significant; even lagged income is marginally significant in Column [ii]. Again the autocorrelation tests of Arellano and Bond (1991), based on first-differenced data, remain in line with expectations. Since these estimations use more instruments than the minimum required for identification, the model is overidentified and we can test the validity of the overidentifying restrictions (moment conditions); the Hansen test indicates no contradiction and hence does not suggest that suspect instruments were included.<sup>24</sup>

Across estimations, the coefficients of the demographic ‘taste shifters’ ( $x_{1,i,t}, \dots, x_{6,i,t}$ ) vary relatively little. Since consumption expenditures and income are expressed in logarithmic per capita terms, the coefficients of  $x_{1,i,t}, \dots, x_{6,i,t}$  can be interpreted roughly as relative changes in per capita expenditures following a modification in household composition. These coefficients are all negative, confirming the existence of scale economies in a household’s food needs.<sup>25</sup> Most costly to feed (and soak) are adult men, including elderly ones. Somewhat less demanding are adolescent and especially young children. As the stereotype would have it, adult women seem to be the least voracious. Not all these differences are statistically significant but several are.

## 4.2 The annual panel 2000-2004

Next we investigate how sensitive the estimates are to the fact that there are gaps in the time period covered (1994-1996, 1998, 2000-2004). To this end we re-estimate equation (8) with and without the lagged income term using only the annual 2000-2004 observations. The results are reported in columns [iii] and [iv] of Table 3, following the comparable estimates for the less regular 1994-2004 panel. Although the results provide some indication that the lower number of observations reduces precision, they remain very similar, and no modification of our substantive conclusions is required. We can conclude that, as far as food consumption is concerned, we detect no obvious misspecification of the dynamic ECM model or violation of the conditions for consistency of the GMM estimator. The estimates seem robust in the sense that they are not very sensitive to the selection of the time interval and to moderately different definitions of the set of instruments used.<sup>26</sup>

In contrast to the findings on food consumption, our estimates using non-food consumption did not pass the specification tests. Table 4 reports the GMM estimates for the non-food total in both the 1994-2004 and the 2000-2004 sample periods. The complete specification produces *negative* income elasticities with relatively large standard errors. Marginal rejections by the Arellano-Bond second-order autocorrelation test (in first-differences) and strong rejections by the Hansen test point towards violations of the model assumptions. Excluding some of the instruments improved the outcome of these tests at the cost of losing precision in the estimates.

---

<sup>23</sup>The weighting matrix is obtained from a preliminary (first-step) GMM estimation round.

<sup>24</sup>In all cases, the Sargan test of the overidentifying restrictions was very close to the Hansen test and led to the same conclusion.

<sup>25</sup>As pointed out by a referee, since the denominator of per capita expenditures is household size, the finding of scale economies may to some extent be explained as a mechanical effect of the per capita calculation. More refined equivalence scales deserve attention but have drawbacks as well.

<sup>26</sup>To save space, not all results are shown here. See Roodman (2008) for an interesting discussion on ways of keeping the number of instruments limited without giving up too much information.

**Table 4: GMM estimates for non-food consumption expenditures**

Sample period	1994-2004		2000-2004	
Column	[i]	[ii]	[iii]	[iv]
$C_{i,t-1}$	0.149*** (0.019)	0.132*** (0.020)	0.123*** (0.026)	0.093*** (0.027)
$y_{i,t}$	0.113* (0.062)	-0.088 (0.098)	0.103 (0.069)	-0.598*** (0.197)
$y_{i,t-1}$	- -	-0.211*** (0.078)	- -	-0.346*** (0.089)
$x_{1,i,t}$	-0.050 (0.042)	-0.120** (0.051)	0.019 (0.060)	-0.150* (0.080)
$x_{2,i,t}$	0.133*** (0.032)	0.082** (0.038)	0.186*** (0.045)	0.078 (0.058)
$x_{3,i,t}$	0.128*** (0.035)	0.145*** (0.038)	0.069* (0.040)	0.085* (0.045)
$x_{4,i,t}$	0.064 (0.041)	0.037 (0.044)	-0.032 (0.052)	-0.122* (0.063)
$x_{5,i,t}$	0.044 (0.076)	0.063 (0.080)	-0.096 (0.102)	-0.076 (0.106)
$x_{6,i,t}$	-0.157** (0.063)	-0.169*** (0.065)	-0.106 (0.083)	-0.108 (0.088)
Observations count	12759	12759	8673	8673
Households count	4370	4370	3768	3768
Total IV count	35	35	19	19
Hansen $p$ -value	0.000	0.000	0.000	0.000
FD-AR(1) $p$ -value	0.000	0.000	0.000	0.000
FD-AR(2) $p$ -value	0.041	0.085	0.082	0.464

Notes: Robust standard errors in parentheses (clustering by household).  
Significance tests: \*\*\* indicates  $p$ -value $<$ 0.01, \*\*  $p$ -value $<$ 0.05, \*  $p$ -value $<$ 0.1.  
The data levels are transformed using forward orthogonal deviations.  
 $x_1 \cdots x_6$  are counts of household members in six age-gender categories.  
Year (survey round) dummies are included though not reported.  
The IV count includes the demographic variables and relevant year dummies.  
FD-AR(1),(2):  $p$ -values of Arellano-Bond tests for serial correlation of orders 1 and 2.  
Source: RLMS 1994-2004.

Although the proposed specification was rejected for non-food consumption, we also estimated the model using total consumption expenditures as the dependent variable (results not shown). Depending on the choice of instruments, the null hypotheses of the specification tests were marginally rejected at 5 or 10 percent significance levels. The point estimates, however, seemed relatively stable across specifications and remained close to those of the food consumption equation. In fact, they were very comparable to the results reported in the previous studies

of Russian consumption smoothing cited above.

It emerges that food and non-food consumption behaviour differ in transition Russia, and that the food component dominates in the analysis of (total) consumption smoothing. There is no shortage of potential explanations for differences in observed behaviour. Since many non-food items include an element of durability, one would actually not have expected the dynamics of adjustment to be identical for the two categories of goods. They may enter the household utility functions in asymmetric ways. Add to that differences in the recall period and in the method of recording expenditures and it will be clear that also the measurement error processes are likely to differ.

### 4.3 Further interpretations

As far as food consumption is concerned, the short-run effects in Table 2 are broadly consistent with the findings of Skoufias (2003), Mu (2006), and Gerry and Li (2008). The interpretation of the short-run income elasticity ( $\beta_1$ ) is the same in the static and in the dynamic models; it measures the instantaneous adjustment of consumption to income shocks. The full insurance hypothesis ( $\beta_1 = 0$ ) is rejected in each model, whatever the estimation method used. The estimates also confirm that current income shocks are not fully transmitted to current consumption ( $\beta_1 < 1$ ); households find ways to smooth their food consumption, or at least a large proportion of it.

At the same time, our findings suggest that the point estimates are sensitive both to the correction for imputation errors, and to the introduction of long-term dynamics. Controlling for imputation errors by the instrumental variable technique yields a substantially lower short-run income elasticity in the static model (viz., a reduction from 2.3 percent to 0.7 percent for a 10 percent income change). With the introduction of long-term dynamics the short-run elasticity comes back half-way (up to 1.5 percent in the complete specification).

The other parameters are also of interest in this relationship. We interpret the coefficient of the equilibrium correction term in equation (5),  $[-(1 - \alpha)]$ , as a negative correction speed. The coefficient is negative since households need to compensate for under- or overconsumption in the previous period, but it should not exceed 1 in absolute value if systematic overshooting is to be avoided. Hence, the meaningful values of the correction speed ( $1 - \alpha$ ) are in the range  $0 < (1 - \alpha) \leq 1$ . Estimated correction speeds, along with standard errors and 95 percent confidence intervals, are reported in Table 5. They are significantly larger than 0 and less than 1, confirming that delayed adjustments are important.

For regular values of the coefficients (as found here), the mean delay of adjustment ( $\bar{\ell}_y$ ) was defined by formula (7). As reported in Table 5, our estimated values (in years) of  $\bar{\ell}_y$  are 0.386 (a little over  $4\frac{1}{2}$  months) in the long 1994-2004 panel, and 0.321 (a little under 4 months) in the annual 2000-2004 panel. These are not long respite periods!

The long-term income elasticity of food consumption, defined in equation (6), is also reported in Table 5. It is estimated at 0.228 and 0.310 in the entire and the shorter period, respectively. These estimates are significantly larger than 0 and, more surprisingly, they are far smaller than unity (both substantively and statistically). They are, nonetheless, somewhat higher than their short-term counterparts, lending some support to the proposition that it becomes more difficult to ‘smooth’ consumption over the longer term.

**Table 5: Structural adjustment parameters for food consumption**

	Point estimates		95% Confidence Interval	
	Coefficient	Standard error	Lower bound	Upper bound
(a) Correction speed $(1 - \alpha)$				
Period 1994-2004	0.875	(0.033)	0.811	0.939
Period 2000-2004	0.890	(0.038)	0.816	0.965
(b) Mean delay of adjustment $\bar{\ell}_y$				
Period 1994-2004	0.386	(0.101)	0.188	0.583
Period 2000-2004	0.321	(0.076)	0.173	0.469
(c) Long-run income elasticity $\theta$				
Period 1994-2004	0.228	(0.082)	0.067	0.389
Period 2000-2004	0.310	(0.143)	0.029	0.591

Notes: Robust standard errors in parentheses (clustering by household).

Source: RLMS 1994-2004; estimates from Table 3 above.

The fact that we find long-term income elasticities far below unity means that we fail to reconcile the panel evidence with the cross-sectional and aggregate time-series evidence. If we construct a cross-section from the RLMS using ten-year differences, then the relationship between consumption expenditures and income is roughly proportional. Cross-sections based on short periods of one or two years exhibit elasticities somewhat lower than 1 but still strong. Similarly in the time-series dimension, the cross-sectional averages (or aggregates) exhibit a unitary long-term elasticity and lower short-term elasticities. But even short-term elasticities are a multiple of those found in the panel estimates.<sup>27</sup>

To the extent that the hypothesis  $H_0 : \beta_2 = 0$  holds, there would be no difference between the ECM model of equation (4) or (5) and a classical partial adjustment model. In any case, we have to conclude that the static specification of equation (2) is dynamically incomplete for food consumption, and this comment is bound to apply to the Skoufias specification in equation (3) as well.

In the case of non-food consumption, the situation is worse. Even the generality of the ECM is insufficient to capture the dynamics in the data. The problem may go unnoticed in a static equation for total consumption, especially when food expenditures dominate the budget. This result casts doubt on efforts to model the aggregate of food and non-food or consumable and durable expenditures, and the conclusions drawn from them.

It is important to keep in mind that the estimates reflect the outcome of a mix of consumption smoothing strategies, including all applicable combinations of self-insurance plans and risk sharing arrangements. Furthermore, as we pointed out in the introduction, the quality and functionality of the existing financial institutions constrain the available options. Given differences

<sup>27</sup>There is a similar puzzle concerning the time-series properties of income and consumption. In the aggregate, the series seem to have an autoregressive coefficient indistinguishable from unity. In our panel estimates, this is far from being the case.

in resources and environments, it is likely that all households will not be able to manage risks and absorb shocks exactly to the same extent. Hence in the next section we explore differences in smoothing responses across households.

## 5 Heterogeneity in smoothing behaviour

To investigate differences in smoothing responses across households we follow an exploratory approach. We estimate the food consumption ECM separately for various groups of households over the period 1994-2004, and compare the results across groups; in this way we also test whether the ECM specification is able to accommodate smoothing responses from households that find themselves in rather different circumstances.<sup>28</sup> The splits in the sample of households are guided by a set of indicators each of which is arguably associated with enhanced or reduced smoothing abilities: (i) the type of income shock (positive or negative), (ii) the environment (rural or urban area), (iii) average standard of living (low, middle, high), and (iv) previous experience of poverty spells (chronically poor, occasionally poor, never poor); the last two groupings will be crossed.

In a standard life cycle model with perfectly functioning financial markets, there is no reason for households to respond differently to positive or negative (unanticipated) income shocks. However, if access to credit or insurance is restricted, the urgency of adjusting to positive or negative shocks may no longer be the same. Households with no or limited access to risk management institutions may be forced to cut food expenditures faster in the face of an adverse shock than they will increase expenses after a windfall gain. Given the dramatic structural changes that characterise the economic transition in Russia, involving deep reforms of the financial markets, the government, the social protection system, and labour relations, one would expect that many Russian households did face restricted access to credit and insurance facilities. We test this proposition by separately estimating the ECM for households experiencing positive shocks and those experiencing negative shocks; the results are summarised in Table 6. We find the short-run income elasticity is very slightly higher for negative income shocks (1.4 percent for a 10 percent income shock) than it is for positive income shocks (1.3 percent for a 10 percent income shock); the small difference is not statistically significant. The long-run income elasticity, on the contrary, is just a little lower for negative shocks (2.1 percent) than for positive shocks (2.3 percent). The speed of adjustment ( $1 - \alpha$ ) in the equilibrium correction term is also somewhat higher for negative shocks. Thus the estimates suggest that households make somewhat faster and larger adjustments in their food expenses after a negative income shock; however, this result is far from conclusive as the standard errors imply overlapping confidence intervals for the parameters.

---

<sup>28</sup>An alternative approach is to introduce dummy variables and interaction terms; see, for an example, Mu (2006).

**Table 6: Overall and split-sample estimation by sign of shock**

	All shocks	Positive shocks	Negative shocks
$c_{i,t-1}$	0.125*** (0.033)	0.136*** (0.043)	0.129*** (0.038)
$y_{i,t}$	0.151*** (0.041)	0.128** (0.053)	0.141*** (0.033)
$y_{i,t-1}$	0.048* (0.029)	0.070* (0.040)	0.040 (0.032)
Observations count	13286	9082	6380
Households count	4464	3984	3444
IV count	35	35	35
Hansen $p$ -value	0.279	0.299	0.388
FD-AR(1) $p$ -value	0.000	0.000	0.000
FD-AR(2) $p$ -value	0.268	0.233	0.416

Notes: Robust standard errors in parentheses (clustering by household).  
Significance tests: \*\*\* indicates  $p$ -value<0.01, \*\*  $p$ -value<0.05, \*  $p$ -value<0.1.  
The data levels are transformed using forward orthogonal deviations.  
Demographic variables and year dummies are included in all estimations.  
The IV count includes the demographic variables and relevant year dummies.  
FD-AR(1),(2):  $p$ -values of Arellano-Bond tests for serial correlation of orders 1 and 2.  
Source: RLMS 1994-2004 (entire period).

The second proposition of heterogeneity we check is whether consumption smoothing behaviour differs between rural and urban households. Given the unequal levels of development in rural and urban areas, the access to and use of risk management institutions by Russian households may vary widely. Skoufias (2003) found that rural households have lower smoothing abilities than urban households, but he was concerned that these findings may have been driven by measurement error. Mu (2006) presented evidence suggesting that different factors determine smoothing abilities in rural and urban areas. While in urban areas higher consumption smoothing abilities are associated with higher levels of education (but not with wealth), in rural areas higher consumption smoothing abilities are associated with higher wealth (but not with education). However, neither Skoufias nor Mu controlled for measurement error in home food production. Although home food production is important for both rural and urban households in Russia (in 2003 half of the urban households reported using land for food production), the prevalence and importance of home food production is much larger for rural households (Notten, 2008 pp. 173-175). Table 7 reports our split-sample estimates for rural and urban households. For the urban sample, lagged income is not significant but the error correction coefficient is close to one (0.926). In the rural model the long-term income elasticity is significant and rather high with 3.7 percent, while the error correction coefficient is 0.813. The short-term income elasticity is considerably higher for rural households (2.0 percent) than for urban households (1.1 percent), suggesting that rural households reduce food expenses more strongly in response to current income shocks. The standard errors imply that 95 percent confidence intervals of the parameter estimates for the two subsamples would overlap slightly. The rural-urban split

is the most common one investigated in previous studies, even though the splitting line is not incontrovertible and remains a topic of research.<sup>29</sup>

**Table 7: Overall and split-sample estimation by type of area**

	All areas	Urban areas	Rural areas
$c_{i,t-1}$	0.125*** (0.033)	0.074** (0.038)	0.187*** (0.050)
$y_{i,t}$	0.151*** (0.041)	0.106** (0.053)	0.204*** (0.056)
$y_{i,t-1}$	0.048* (0.029)	0.001 (0.041)	0.094** (0.041)
Observations count	13286	8699	4587
Households count	4464	3046	1418
IV count	35	35	35
Hansen $p$ -value	0.279	0.820	0.429
FD-AR(1) $p$ -value	0.000	0.000	0.000
FD-AR(2) $p$ -value	0.268	0.324	0.571

Notes: Robust standard errors in parentheses (clustering by household).  
Significance tests: \*\*\* indicates  $p$ -value<0.01, \*\*  $p$ -value<0.05, \*  $p$ -value<0.1.  
The data levels are transformed using forward orthogonal deviations.  
Demographic variables and year dummies are included in all estimations.  
The IV count includes the demographic variables and relevant year dummies.  
FD-AR(1),(2):  $p$ -values of Arellano-Bond tests for serial correlation of orders 1 and 2.  
of the first and second order in residual first differences.  
Source: RLMS 1994-2004 (entire period).

Finally we will try and assess whether households with a lower living standard or having previously experienced one or more episodes of poverty have different ways of adjusting expenditures. Even though we focus on food, it is not entirely clear *a priori* what relationship to expect. On the one hand, households with a lower living standard would tend to have less resources to manage risks and more restricted access to risk management institutions; one would therefore expect them to respond more strongly to income shocks. On the other hand, it is sometimes argued that well-to-do households have more means to economise in planning, buying, conserving and preparing food, while unequipped households are at the mercy of a local and irregular supply.

We attempt a split of the sample of households according to their living standard (low, middle or high) and their previous experience of poverty spells (chronically poor, occasionally poor, never poor). Both criteria are derived from the history of actual household expenditures.<sup>30</sup> To construct a proxy for living standard, a series of ‘household-equivalent expenditures’ is generated as the ratio of actual household expenditures to the household-specific poverty line. The resulting

<sup>29</sup>We use the RLMS definition of *urban* and aggregate all other areas (including the ‘poselki gorodskogo tipa’ or ‘peri-urban’ areas) in a *rural* (or, strictly, ‘non-urban’) category.

<sup>30</sup>We disregard the implied endogenous selectivity on the grounds that its effects are likely to be minor. Only expenditures on non-durable goods are taken into account in these calculations.

ratio is then averaged over time. Thus, a time average below 1 indicates that a household lived in poverty during the transition period; and conversely for an average above 1. Then we group households into low (ratio below 1.5), middle (ratio between 1.5 and 2.5) or high (ratio of 2.5 or higher) living standard categories. The proportion of households in each category is roughly one third; from low to high, 30 percent, 37 percent and 33 percent, respectively.

Next we divide households into three categories: frequently poor, occasionally poor, and never poor. Frequently poor households are households observed to fall below the poverty line in at least three survey rounds (22 percent); occasionally poor households experienced one or two spells of poverty (36 percent); never poor households remained above the poverty line in all the survey rounds in which they participated (43 percent).<sup>31</sup>

The criteria of living standard and poverty frequency are then combined or ‘crossed’ into six subgroups: frequently poor (1); occasionally poor by low (2), medium (3), and high (4) living standard; and never poor by medium (5) and high (6) living standard.

\*\*\* INSERT TABLE 8 (in LANDSCAPE ORIENTATION) ABOUT HERE \*\*\*

The results are summarised in Table 8. Only in the three largest of the six subgroups, one or more of the estimated adjustment parameters are statistically significant: frequently poor (1); occasionally poor, medium standard (3); never poor, high standard (6). We attribute this result to the combination of a computationally demanding estimator (GMM) with relatively low numbers of observations in the subsamples. Focusing on the estimates for the larger subgroups, the results suggest that there is only a slight difference in food consumption smoothing behaviour between frequently poor and occasionally poor, middle standard households. High standard, never poor households tend to make larger short term and long term adjustments in response to income shocks; this result stood when we split the sample in only three living standard categories (results available on request). A higher income elasticity for high standard households is consistent with the explanation that well-to-do households have more means to economise on food expenditures than unequipped households. The large standard errors of the estimated income parameters, implying once again overlapping confidence intervals, indicate that these results should not be taken too literally.

## 6 Conclusion

In comparison to the models typically used in the empirical literature on consumption smoothing, this paper developed and tested a more flexible equilibrium correction model (ECM) of dynamic consumption smoothing behaviour. The ECM distinguishes between immediate and delayed adjustments of consumption expenditures to income shocks and thereby takes into account that income and consumption can temporarily deviate even though they should balance in the long run. We tested and rejected the full insurance hypothesis, namely, that current consumption is entirely detached from contemporary income shocks. The inevitability of a lifetime budget constraint suggests the necessity of a long-term connection between consumption and income. We therefore also investigated whether observed behaviour is consistent with a tendency of consumption and income to return to a long term equilibrium after income shocks. We failed however to detect the implied proportionality relationship between consumption and income at

---

<sup>31</sup>This classification is chosen to reflect a spells-based concept of poverty (Hulme & Shepherd, 2003). The ‘frequently poor’ category is comparable to the indicator for ‘at risk of persistent poverty’ proposed by Atkinson *et al.* (2002). Using the RLMS, Luttmer (2001) finds that one fifth of the poor in Russia are erroneously classified as a consequence of transitory shocks and noisy data.

the household level in the RLMS data, and thus to reconcile the panel evidence with stylised facts from cross-sectional or aggregate time-series analyses.

Undoubtedly more research will be required to find out exactly how complex the dynamics of adjustment need to be; and why the implied long-run income elasticity of (food and non-food) consumption expenditures is so much smaller than unity.

In our empirical estimation strategy we compared a range of short-term and long-term models, using several estimation techniques and making various adjustments to reduce the impact of measurement error on our estimates. To take these factors simultaneously into account we used a state of the art instrumental variable estimation technique, the Generalised Methods of Moments. The ECM was subjected to a number of specification and robustness tests. Our results broadly confirm the findings of previous studies on consumption smoothing behaviour in Russia, mainly the rejection of the full insurance hypothesis. However, they also show that dynamics are necessary in the consumption equation, and that smoothing estimates are sensitive to the measurement error caused by imputations in home food production. The heterogeneity tests in the last section of the paper provide some limited indication that consumption smoothing abilities may vary according to the type of shock (positive or negative), the level of urbanisation, standards of living and previous experience of poverty spells. However, small and statistically insignificant differences in the point estimates prevent strong claims with respect to heterogeneity in smoothing behaviour.

Concluding, the results presented in this paper very well illustrate that the task of gaining a better understanding of the empirical realities of consumption behaviour is highly complex and that, despite the many recent advances in the availability of data and estimation techniques, our power to answer deep questions is still limited.

## References

- Altonji, J.G., Hayashi, F. and Kotlikoff, L.J. (1992). ‘Is the extended family altruistically linked? Direct tests using micro data’, *American Economic Review*, 82, pp. 1177-1198.
- Anderson, T.W. and Hsiao, C. (1981). ‘Estimation of Dynamic Models with Error Components’, *Journal of the American Statistical Association*, 76, pp. 598-606.
- Arellano, M. and Bond, S. (1991). ‘Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations’, *Review of Economic Studies*, 58, pp. 277-297.
- Arellano, M. and Bover, O. (1995). ‘Another Look at the Instrumental Variable Estimation of Error-Component Models’, *Journal of Econometrics*, 68, pp. 29-51.
- Atkinson, A. B., Cantillon, B., Marlier, E. and Nolan, B. (2002). *Social indicators: the EU and social inclusion*, Oxford and New York: Oxford University Press.
- Atkinson, A. B., Rainwater, L., and Smeeding, T.M. (1995). *Income distribution in OECD countries: the evidence from the Luxembourg Income Study (LIS)*, Paris: Organisation for Economic Co-operation and Development.
- Banerjee, A.V. and Duflo, E. (2007). The economic lives of the poor, *Journal of Economic Perspectives*, 21, pp. 141-167
- Blundell, R. and Bond, S. (1998). ‘Initial Conditions and Moment Restrictions in Dynamic Panel Data Models’, *Journal of Econometrics*, 87, pp. 115-143.
- Brown, A. (1999). ‘The Russian crisis: Beginning of the end or end of the beginning?’ *Post-Soviet Affairs*, 15, pp. 56.
- Browning, M. and Crossley, T.F. (2001). ‘The life-cycle model of consumption and saving’, *Journal of Economic Perspectives*, 15, pp. 3-22.

- Buchs, T. D. (1999). 'Financial crisis in the Russian Federation', *Economics of Transition*, 7, pp. 687.
- Cochrane, J.H. (1991). 'A simple test of consumption insurance', *Journal of Political Economy*, 99, pp. 957-976.
- Davidson, J.E.H., Hendry, D.F., Srba, F. and Yeo, S. (1978). 'Econometric modelling of the aggregate time-series relationship between consumers' expenditure and income in the United Kingdom', *Economic Journal*, 88, pp. 661-692.
- Deaton, A. (1992). *Understanding Consumption*, Clarendon Lectures in Economics, Oxford: Oxford University Press and Clarendon Press.
- Deaton, A. (1997). *The analysis of household surveys: A microeconomic approach to development policy*, Washington, D.C.: Johns Hopkins University Press for the World Bank.
- Denizer, C., Wolf, H.C. and Ying, Y. (2002). 'Household saving in the Transition', *Journal of Comparative Economics*, 30, pp. 463-475.
- Engel, E. (1857). 'Die Productions- und Consumptionsverhältnisse des Königreichs Sachsen', *Zeitschrift des Statistischen Büreaus des Königlich Sächsischen Ministerium des Innern*, No. 8 u 9, 1-54. Reprinted as an appendix to 'Die Lebenskosten Belgischer Arbeiter-Familien früher und jetzt' in *Bulletin de l'Institut International de la Statistique IX* (1895), 1e livraison - Anlage 1, pp. 1-54.
- Engle, R.F. and Granger, C.W.J. (1987). 'Cointegration and Error Correction: Representation, Estimation and Testing', *Econometrica*, 55, pp. 251-276.
- Fafchamps, M. and Lund, S. (2003). 'Risk-Sharing Networks in Rural Philippines', *Journal of Development Economics*, 71, pp. 261-287.
- Foley, M.C. and Pile, W. (2005). 'Household Savings in Russia during the Transition', Middlebury College Economics Discussion Paper No 05-22, <http://sandcat.middlebury.edu/econ/repec/mdl/ancoec/0522.pdf>
- Foster, J., Greer, J. and Thorbecke, E. (1984). 'A Class of Decomposable Poverty Measures', *Econometrica*, 52, pp. 761-766.
- Freinkman, L. (1998). 'Russian Federation Housing and Utility Services: Policy Priorities for the Next Stage of Reforms', Report No. 17483-RU, Washington, D.C.: World Bank.
- Friebel, G. and Guriev, S.M. (2005). 'Attaching Workers Through In-kind Payments: Theory and Evidence from Russia', *World Bank Economic Review*, 19, pp. 175-202.
- Friedman, M. (1957). *A theory of the consumption function*, Princeton: Princeton University Press.
- Gassmann, F. (2003). *Review of Targeted Social Assistance in Russia; Description, Analysis and Recommendations*, Eindhoven: Gassmann Consulting.
- Gerry, C.J. and Li, C.A. (2008). Consumption smoothing and vulnerability in Russia. *Applied Economics*, iFirst 05 Dec 2008, doi: 10.1080/00036840701765403.
- Gregory, P., Mokhtari, M. and Schrettl, W. (1999). 'Do Russians Really Save That Much? Alternate Estimates from the Russian Longitudinal Monitoring Survey', *Review of Economics and Statistics*, 81, pp. 694-703.
- Guariglia, A. and Kim, B.Y. (2003). 'Wage arrears uncertainty and precautionary saving in Russia', *Economics of Transition*, 11, pp. 493-512.
- Hall, R. (1978). 'Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence', *Journal of Political Economy*, 86, pp. 971-987.
- Heeringa, S.G. (1997). *Russia Longitudinal Monitoring Survey (RLMS), Sample Attrition, Replenishment, and Weighting in Rounds V-VII*, Survey Design and Analysis Unit, Institute for Social Research, University of Michigan, Ann Arbor, MI, 48106-1248.

- Hulme, D. and Shepherd, A. (2003). 'Conceptualizing chronic poverty', *World Development*, 31, pp. 403-423.
- Kleibergen, F. and Paap, R. (2006). 'Generalized reduced rank tests using the singular-value decomposition', *Journal of Econometrics*, 127, pp. 97-126.
- Lokshin, M. and Ravallion, M. (2004). 'Household Income Dynamics in Two Transition Economies', *Studies in Nonlinear Dynamics and Econometrics*, 8, pp. 1-31.
- Luttmer, E. (2001). 'Measuring Transition Dynamics and Inequality in Transition Economies', World Bank Policy Research Working Paper No. 2549, Washington, D.C.
- Mace, B.J. (1991). 'Full insurance in the presence of aggregate uncertainty', *Journal of Political Economy*, 99, pp. 928-957.
- Mroz, T.A., Henderson, L. and Popkin, B.M. (2005). 'Monitoring Economic Conditions in the Russian Federation: The Russia Longitudinal Monitoring Survey 1992-2004', Report submitted to the U.S. Agency for International Development, Carolina Population Center, University of North Carolina at Chapel Hill, North Carolina.
- Mu, R. (2006). 'Income Shocks, Consumption, Wealth, and Human Capital: Evidence from Russia', *Economic Development and Cultural Change*, 55, pp. 857-892.
- Notten, G. (2008). *Measuring and Managing Poverty Risks*, Maastricht: Boekenplan, <http://arno.unimaas.nl/show.cgi?fid=11319>
- Notten, G. and de Crombrugghe, D. (2007). 'Poverty risk and consumption smoothing abilities in Russia', MPRA Paper No. 5314.
- Notten, G. and Gassmann, F. (2008). 'Size matters: poverty reduction effects of means-tested and universal child benefits in Russia', *Journal of European Social Policy*, 18, pp. 260-274.
- Ravallion, M. (1994). *Poverty Comparisons, Vol. 56*. Harwood Academic Press.
- Ravallion, M. and Chaudhuri, S. (1997). 'Risk and Insurance in Village India: Comment', *Econometrica*, 65, pp. 171-184.
- Roodman, D. (2008). 'A Note on the Theme of Too Many Instruments', *Oxford Bulletin of Economics and Statistics*, 71, pp. 135-158.
- Russian Economic Trends, 1998-2003. (Various issues; discontinued in 2003)
- Sapir, J. (1999). 'Russia's Crash of August 1998: Diagnosis and Prescription', *Post-Soviet Affairs*, 15, pp. 1-36.
- Skoufias, E. (2003). 'Consumption smoothing in Russia: Evidence from the RLMS', *Economics of Transition*, 11, pp. 67-91.
- Slay, B. (1999). An Interpretation of the Russian Financial Crisis. *Post-Soviet Geography and Economics*, 40, pp. 206-214.
- Stillman, S. (2001). 'The response of consumption in Russian Households to economic shocks', IZA Discussion Paper No. 411, Institute for the Study of Labor, Bonn.
- Townsend, R.M. (1994). 'Risk and Insurance in Village India', *Econometrica*, 62, pp. 539-591.
- World Bank (2003). '*Housing and Communal Services in Russia: Completing the Transition to a Market Economy, Final Report*', The World Bank Infrastructure and Energy Department, Europe and Central Asia Region, [http://siteresources.worldbank.org/INTRUSSIANFEDERATION/Resources/305499-1094736798511/518266-1095320717197/housing\\_eng.pdf](http://siteresources.worldbank.org/INTRUSSIANFEDERATION/Resources/305499-1094736798511/518266-1095320717197/housing_eng.pdf)
- World Bank (2004). *Russian Federation: Poverty Assessment*. 28923-RU, Poverty Reduction and Economic Management Unit, Europe and Central Asia Region, Washington D.C.: The World Bank.
- Zohoori, N., Mroz, T.A., Popkin, B., Glinskaya, E., Lokshin, M., Mancini, D., Kozyreva, P., Kosolapov, M. and Swafford, M. (1998). 'Monitoring the Economic Transition in

the Russian Federation and its Implications for the Demographic Crisis - the Russian Longitudinal Monitoring Survey', *World Development*, 26, pp. 1977-1993.

**Table 8: Overall and split-sample estimation by poverty experience and living standard**

	All	Frequently poor		Occasionally poor & Usual living standard			Never poor & Usual living standard			
		low	medium	high	low	medium	high	low	medium	high
$c_{i,t-1}$	0.125*** (0.033)	0.078 (0.049)	-0.010 (0.083)	0.006 (0.061)	0.004 (0.086)	0.045 (0.109)	0.303*** (0.082)			
$y_{i,t}$	0.151*** (0.041)	0.146** (0.074)	-0.001 (0.138)	0.152** (0.077)	0.141 (0.135)	0.075 (0.091)	0.263*** (0.083)			
$y_{i,t-1}$	0.048* (0.029)	0.059 (0.055)	-0.048 (0.088)	0.009 (0.055)	0.060 (0.103)	-0.013 (0.070)	0.132*** (0.050)			
Observations count	13286	2703	1086	2592	1043	1818	3124			
Households count	4464	784	470	767	306	663	1142			
Instrument count	35	35	35	35	35	35	35			
Hansen $p$ -value	0.279	0.502	0.154	0.681	0.183	0.280	0.251			
FD-AR(1) $p$ -value	0.000	0.000	0.000	0.000	0.000	0.001	0.000			
FD-AR(2) $p$ -value	0.268	0.414	0.442	0.840	0.471	0.321	0.169			

Notes: Robust standard errors in parentheses (clustering by household).

Significance tests: \*\*\* indicates  $p$ -value $<$ 0.01, \*\*  $p$ -value $<$ 0.05, \*  $p$ -value $<$ 0.1.

The data levels are transformed using forward orthogonal deviations.

Demographic variables and year dummies are included in all estimations.

The IV count includes the demographic variables and relevant year dummies.

FD-AR(1),(2):  $p$ -values of Arellano-Bond tests for serial correlation of orders 1 and 2.

Source: RLMS 1994-2004 (entire period).