Social transfers and intrahousehold resource allocation: Evidence from Russia

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Abstract

I examine the effect of government transfers on intrahousehold resource allocation using panel data on household expenditures. I pay particular attention to child benefits, where the recipient of the funds differs from the target beneficiary. I find that the marginal propensity to spend on food is three times as high from child benefits money as it is from earned household income. Child benefits money improves both child and adult protein intakes in ways in which equivalent increases in other household income do not. I present evidence that the reason why government transfers are disproportionately spent on food in Russia is that women control this income source. Women appear to have systematically different resource allocation priorities from men.

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I Introduction

The purpose of this paper is to examine how intrahousehold resource allocation is affected by government transfers. I pay particular attention to child benefits, a type of transfer in which the recipient is not the target beneficiary, and to gender patterns in expenditure. Understanding how households use child benefits money is of substantial policy importance, given the high and growing rates of child poverty in many industrialised and developing countries. The broader question of the intrahousehold allocation of benefits has direct implications for economists attempting to understand household bargaining, as well as for policymakers aiming to reduce child poverty. If households view child benefit payments as 'earmarked' for children in some way that other income is not, providing child benefits will be more effective in improving the living standards of children than more generic forms of income support to low-income households, such as social assistance.

The issue of social transfer use is related to the recent theoretical and empirical literature on household resource allocation. This literature allows for distinct preferences over resource allocation amongst household members, and thus for changes in the distribution of resources within the household to have impacts on resource allocation. Browning, Bourgignon, Chiappori, and Lêchene (1994) estimate of a structural model of resource allocation in two-adult households with no children based on the theoretical work of Chiappori (1992). The Chiappori collective household model consider household members to have distinct utility functions, and thus to potentially allocate increments in income accruing to different members differently. Resource allocation is efficient. As well as the collective household model, a long-standing literature on cooperative and non-cooperative household bargaining models (see for example McElroy and Horney (1981), Pollak and Lundberg (1994a), Pollak and Lundberg (1994b)) has contributed to the prevailing view of the household as a consisting of multiple agents.

The theoretical literature on resource allocation in non-unitary households was partly inspired by empirical studies suggesting that income allocated to men has less effects on child health and wellbeing than income allocated to women.Thomas (1990) shows that income accruing to women has larger effects on child health than that accruing to men. Lundberg, Pollak, and Wales (1997) find that the 1977 reallocation of child benefits in the UK from fathers to mothers resulted in a significant increase in expenditure on children's clothing. To my knowledge, this is the first panel data study to examine how transfer income is spent within households. Duflo (2003)) examines the effectiveness of a generous pension program introduced in South Africa between 1991 and 1993 on the wellbeing of children living in households with pensioners. She finds that girls living in households where women (primarily their grandmothers) received a pension had significantly better anthropometric health scores than girls in households where no pension was received. No such effect on girls' health was found in households where a man received the pension. Thus one major positive result of the introduction of a generous old age pension in South Africa appears to be that girls in households with a pension-eligible woman (which are disproportionally poor) have better nutrition than they otherwise would.

The greater propensity to spend child benefits money than regular income on children has been termed a *labeling effect*. Economists who have looked at the so-called labeling effect of child benefits include Kooreman (2000) and Edmonds (2002). Kooreman finds that there is a large labeling effect of the Dutch social benefits system, whereas Edmonds finds no evidence of a labeling effect in the Slovenian child benefits system. Sahn and Gerstle (2003), use data from the 1994 Romanian Integrated Household Survey and estimate expenditure equations on child goods. The data used in all of these studies is cross-sectional, and so cannot account for inter-household hetergeneity of preferences. Aside from this work on child benefits labeling effects, a majority of the evidence on differences in the use of income from social transfers comes from analysis of US food stamp programs in the 1980s (see for example Devaney and Fraker (1986)). This research, which found that the marginal propensity to consume (MPC) on food from food stamps is much higher than from other income, spawned the so-called 'cash-out' hypothesis. This hypothesis centers on the form of benefits allocation (a stamp), and the idea that stigma resulting from the use of the stamp changes the marginal utility of a unit of food so received.

Evidence that child clothing expenditures are higher in households where child benefits make up a greater portion of income does not necessarily suggest that child benefits are better than more generic transfers in reducing child poverty, or the poverty of households containing children. While Sahn and Gerstle show that households receiving child benefits have slightly higher caloric intakes (per equivalent household) than otherwise, this does not necessarily imply that child nutrition improves with the benefit. Given the high levels of malnourishment amongst children in Russia and throughout the Former Soviet Union, and the fact that childhood nutrition deficiencies can have longterm consequences (in a way that lack of new clothes cannot), caloric diary information on protein consumption may be very helpful in understanding the true impact of child benefits on children. More calories does not imply better nutrition, as some high-calorie carbohydrates and fats may be inferior goods, and so it may be important to look at nutrient composition.

In this paper, I examine how child benefits and other social transfers are allocated, using panel data from the Russian Longitudinal Monitoring Survey (RLMS). I estimate an incomplete demand or expenditure system, using panel data techniques. I make use of several features of the Russian social security system in order to identify the effects of child benefits and other social transfers on the consumption of various household items. These features include: (*i*.) differences in real child benefits levels across regions within years, (*ii.*) differences in child benefits within regions across years, (*iii.*) differences in other social transfers (mainly pensions) both geographically and over time, (*iv.*) changes in real family income across years, and (*v.*) the fact that household composition is much more heterogeneous in Russia than in Western countries.

Amongst households in the 1994 RLMS, the mean number of female adults per household was 1.2, whereas for men it was 0.92. In 2000 disability-adjusted[1] male life expectancy in Russia was 56.2 years versus 66.4 years for females, one of the biggest sex gaps in the world, according to the WHO (2000). Similar large sex gaps in disability-adjusted life expectancy exist for the Ukraine (67.5 for females versus 58.5 years males), Belarus (67.2 years for females versus 56.2 years for males). According to the WHO, alcoholism and tobacco use are major factors in life expectancy, and in the sex gaps in life expectancy observed in the Former Soviet Union.

Government transfer income in Russia, as in many other countries, is strongly gendered. Child benefits are generally transfered to mothers. The largest social transfer payment in Russia is the old-age pension, which accrues disproportionately to women because of the large gender gap in life expectancy. According to the RLMS, one quarter of children under age 16 in 1994 lived with a pension-age female, while only 7% lived with a pension-age male. Changes in government transfer income in Russia imply exogeneous changes in the share of household income accruing to women. If there are systematic differences across the sexes in preferences, resource allocation patterns will generally be altered as the value of social transfers changes.

My main finding is that household revenue received as child benefits or social transfers is used differently than other sources of household revenue, particularly in the purchasing of food items. I find that the MPC on food from child benefits and other social transfers is three times as high as from other income. Households containing only female adults spend significantly more on food, and less on child activities and alcohol, than mixed-gender households with equivalent financial resources. Only in all-female households is the MPC from government transfers similar to that of earned household income. This suggests that when women in mixed-gender households have control of child and social transfer money, their spending this income on food in households is facilitated. I do not find that the MPC on child-specific commodities (childrens' supplies and activities, children's clothing) is greater from child benefits than from earned household income. Thus the results of this analysis support neither the Kooreman's labeling nor the cash-out hypothesis. Adult protein consumption is found to increase as much from child benefits as child protein consumption.

[1]Disability adjusted male life expectancy (DALE) summarises the expected number of years to be lived in full health. The WHO calculates DALE by subtracting expected years of ill-health from overall life expectancy.

The remainder of the paper consists of seven sections. In Section II I provide background information on changes in the socio-economic situation in Russia in the 1990s, paying special attention to the welfare of children. In Section III I discuss the organisation of child and pension benefits payments in Russia. Section IV introduces the Russian Longitudinal Monitoring Survey, the only household panel survey existing for Former Soviet Union countries. Section V discusses identification of the MPC from child benefits and other income, and examines consumption on assignable and household public goods from each source. In Section VI I examine how income and child benefits affect the diet of children and adults in Russia. Section VI is devoted to a discussion of how the results can be reconciled to those in the small existing literature on social transfers and intrahousehold resource allocation. Section VII concludes.

II Background

Reducing child poverty in the countries of the Former Soviet Union has become a high priority amongst international institutions in recent years. Since the transition to market economies in 1991, child poverty levels have risen dramatically. According to UNICEF (2003), 11 million children in Russia were living in poverty in 2002. While definitions of poverty are the subject of much methodological and political debate, there is broader agreement about the potential longterm consequences of childhood deprivation. In 1994, 12% of Russian children aged 0-7 were stunted (low weight for age), and 5% were wasted (low weight for height), according to RLMS data used in this study.

Malnutrition during early childhood is known to have strong negative effects on adult stature. Adult height is positively correlated with earnings, productivity, and cognitive skills in developing countries, and negatively correlated with premature mortality (see for example World Bank (1993)). Poverty within households is known to be a key factor in children leaving home for the streets, where they often drop out of school, become involve in substance abuse, and are vulnerable to exploitation. The rapid increase in the numbers of children living in the streets of all major cities of the Former Soviet Union since 1991 is a testament to the way in which economic crises are associated with family crises.

During the latter years of the Soviet Union, childhood malnutrition was relatively uncommon in Russia. The prevalence of employer-run cafeterias and school meals meant that families did not have to provide for all meals. The cultivation of vegetables at private plots (dachas) was prevalent, as it is today. However, real incomes have fallen substantially in transition. Free and subsidised meals programs have been eliminated, and the prevalence of single mothers has risen dramatically. In addition to child benefits paid to families, special benefits are also targeted to single mothers in Russia. According to Popkin, Mullan-

Harris, and Loshkin (2000), single parent households now represent about 25% of all Russian households.

III Variations in benefits levels and eligibility

Universal child benefits were introduced in Russia in 1991. The Russian/Soviet governments anticipated that economic transition would have strong negative effects on family budgets, and attempted to compensate. These benefits were to be paid out by the governments of the 32 regions of Russia. Prior to 1991, only very poor families received (means-tested) financial assistance for help with children. However, regional budget crises in Russia almost immediately compromised the universality of child benefits. Denisova, Kolenikov, and Yudaeva (2000) find that by 1996 only 33% of eligible families received child benefits. Misikhina (1999) provides evidence that relatively wealthy families received much larger fractions of benefits than poorer families. These benefits are sometimes important components of of household income. In 2000, the child benefit value was equal to 70% of the minimum wage.

Table I presents mean per capita child benefits levels by region, for 1994 and 2001. Regional and inter-temporal variation in these social transfers is large. In the Komi ASSR (Syktyvkar) in 1994, per capita received child benefits were 654 (June 1992) roubles, whereas they had dropped to 181 (June 1992) roubles in 2001 (Column (I) row (iv)). Syktyvkar is the capital of the Komi Republic in the Russian Far North, and is home to an important cellulose industry. Wages and salaries were traditionally much higher in the Far North than in other regions to the south because of 'compensation pay' for the harshness of the environment. As well, social transfer payments were comparatively generous. The cellulose industry experienced severe declines in the late 1990s and a loss in tax base. Even relatively prosperous areas, such as Moscow City (Column (I), row (ii)) experienced substantial declines in child benefits levels, from 220 roubles in 1994 to 85 roubles in 2000. Only in three regions, Chiliabinsk Oblast, Tomsk, and Tatarksaja ASSR, did real child benefits increase over the 1994.IV-2001.IV period.

From 1994, child benefits payments were made in the form of a single monthly payment for families. Child benefits levels, though universal, were structured according to the age of children. The age groupings used were 0-1.5, 1.5-6, and 6-16. Maternity benefits and benefits to non-working mothers with children under 1.5 years are financed by the Social Insurance Fund, not the regions or local authorities. In the RLMS it is not possible to distinguish child benefits from maternity benefits. However, I select households whose composition does not change from 1994 to 2001, so that maternity benefits for new children will not convolute the analysis of potential labeling effects.

In the second half of the 1990s regional governments began adopting new strategies to deal with the fiscal burden of paying child benefits. Some governments began to target benefits at poor families in 1995, with several others following suit in subsequent years. In 1998 the Russian State Duma passed legislation conforming to the prevailing practise. This legislation proclaimed that only families with incomes below the regional subsistence level could receive child benefits. Regions maintained the right to impose further restrictions. It was anticipated that targeting would help clear benefit arrears by limiting coverage to families most in need of support. In 1994, the mean monthly child benefit level per child (when received) was 473 June 1992 roubles, whereas in 2001 it was 217 rubles[2].

The mid-1990s fiscal crises in Russian regions were not limited to child benefits. In 1996, during the run-up to Presidential elections, there was a sharp decline in the collection of the payroll taxes which funded pension payments (see Jensen and Richter (2000)). Approximately 14 million of Russia's 39 million pensioners experienced an extended period of non-payment of pensions. Using the 1994-1996 rounds of the Russian Longitudinal Monitoring Survey (RLMS), Jensen and Richter (2000) find that this pension crisis had a large negative impact on living standards of pensioners, effectively tripling poverty rates amongst this group. I will use information about pensions levels to aid in understanding whether there are common effects of social transfers intrahousehold resource allocation in Russia.

Child benefits were not exclusively paid in monetary form in the mid 90s. In many regions experiencing budget crises, portions of child benefits were 'paid' as credits on food in certain stores. Stores were able to reduce their tax arrears in this manner. Relatively little information is available on such benefits payments schemes in the RLMS data. Because the RLMS asks only about *money* received as child benefits, and *money* spent on different budget items, this feature of the benefits system cannot be included in the analysis. In this version of the paper, I exclude households who report receipt of government transfers in goods form in the month prior to an RLMS interview[3].

By 2001, means testing of child benefits had become pervasive in Russia. Different regions used different criteria, usually an income test combined with various means tests relating to the demography of the household. For a comprehensive discussion of these targeting schemes, and of the effectiveness of targeting criteria, see Denisova, Kolenikov, and Yudaeva (2000).

Figures I, II, and III present information child benefits receipt in Russia between 1994.IV and 2001.IV. (amongst households containing at least one under-16 child), for the sample period[4]. Figure I suggests that targeting schemes instituted in the latter half

[2]Author's calculations using RLMS 1994-2001.

[3]This amounts to excluding about 5% of households.

[4]Values are calculated using the estimation sample, to be discussed in the next section.

of the 1990s reduced eligibility. Figure II shows how benefits crises drove a wedge between eligibility and benefits receipt in the late 90s. Figure III shows that mean received benefits levels have varied widely over the sample period.

The conditions governing both eligibility and receipt of child benefits in Russia have been extremely complicated in the 1990s. However, all of the rules can be considered exogeneous to households and children. This feature of the benefits structure is the key to identifying the effects of child benefits on intrahousehold resource allocation. A multivariate analysis of the propensity to receive child benefits, presented in Appendix A, shows that there is no statistically significant relation between household income and propensities to receive benefits over the 1994-2001 period in Russia. Consistent with the situation described in this section, variation in propensities to receive benefits amongst households is principally due to regional and time variation. As shown in in Figure IV, eligibility rates are very high throughout the sample period. Eligibility is negatively related to household income, despite the fact that receipt is not.

IV Data

The RLMS consists of two panels, a four wave panel running between 1992 and 1994 and a (continuing) 6 wave panel running between 1994 and 2001. Both are nationally representative samples of Russian households and individuals. The first contains 6300 households and the second about 4000. These data were collected primarily to assess the health consequences of economic transition, although they contain detailed information on the working lives of individuals and on household income and expenditures. The 1994-2001 RLMS panel contains yearly information on prices of basic commodities and the availability of services in each of 157 communities in the RLMS. For examples of previous work using the RLMS, see Mroz and Popkin (1995) on poverty, (Popkin, Baturin, Kohlmeier, and Zohoori (1996)) on nutrition, and Sheidvasser and Benitez-Silva (1999) on returns to human capital. For more details on the RLMS panels see (*http://www.cpc.unc.edu/rlms/*).

The RLMS is the only panel survey for any of the Former Soviet Union countries, and provides comprehensive information on many aspects of people's lives. However, it does have some drawbacks. While community-level price data is collected on food items, it is not for clothing and other non-perishables. Separate information on child and adult clothing purchases was not collected until 1998. Thus it is not possible to estimate a full demand system for Russian households. However, the goal of the present analysis is more limited, and is that of assessing how social transfers, and particularly child benefits money, is spent within Russian households. In the following section I briefly outline the efficiency implications of the collective model of the household model and discuss the estimation procedure.

V Intrahousehold resource allocation

Many recent empirical tests of the unitary household model have rejected the key implication of the model that reallocation of income between spouses should have no effects on household resource allocations (see for example Lundberg, Pollak, and Wales (1997)). The advent of cooperative and non-cooperative bargaining models and collective models has generated a much richer set of predictions regarding intrahousehold allocations. A key implication of the collective household model is that household allocations be pareto efficient, despite the fact that changes in the allocation of income amongst adults in the household can have real effects on consumption.

Formally, consider a household with two adults and one child. Let e_f denote consumption of the female member (f), e_m denote consumption of the male member (m), and e_k denote consumption of the child (k). Let c refer to child benefits, n to total household earned income, s to social transfers (mainly pensions), and p to private transfers accruing to the household. Further, let $n = \sum_j n_j$, $p = \sum_j p_j$, $s = \sum_j s_j$, and $c = \sum_j c_j$ where $j \in \{m, f\}$.

Let the female household member have preferences such that

$$U(e_f, e_k)$$

and the male member have preferences such that:

$$V(e_m, e_k)$$

Denote the reservation utility of the male in the household as V_r such that

$$V(e_m, e_k) \geq V_r$$

The female's maximisation problem can be written as follows:

$$\max_{e_m, e_f, e_k} U(e_f, e_k)$$

subject to the constraint

$$n_m + s_m + p_m + c_m + n_f + s_f + p_f + c_f = e_m + e_f + e_k$$

and the attainment of the reservation utility of the male.

Further, let $z = n_f + s_f + p_f + c_f$ denote total resources available to the female and $h = n_m + s_m + p_m + c_m$ denote total household resources available to the male.

The optimisation problem of the female can be written as the Lagrangian:

$$L = U(e_f, e_k) + \lambda(z + h - e_m - e_f - e_k) + \theta(V_r - V(e_m, e_k))$$

The female's two constraints are the household resource constraint and the male's reservation utility constraint. The maximisation problem will now result in resource allocations which depend on the reservation of the male, V_r as well as the household's resource constraint $z + h$.

One may expect that V_r depends on the relative amount of financial power held by the male in the household, the ratio $(\frac{h}{z})$. An exogenous change in child benefits or other social transfers that accrue to one household member would then alter V_r, and thus generally alter resource allocations in the household after controlling for the pure income effect.

To illustrate this, let $V_r = f(\frac{h}{z})$ such that $\frac{\partial f(\frac{h}{z})}{\partial h} > 0$ and $\frac{\partial f(\frac{h}{z})}{\partial z} \leq 0$.

Consider a reallocation of benefits such that $c_f = a$ and $c_m = 0$, where previously $c_m = a$ and $c_f = 0$. This is the case studied by Lundberg, Pollak, and Wales (1997), in which a child benefit previously awarded to the father was replaced by an equivalent child benefit awarded to the mother in the UK in 1977. In this case we have that

$$\partial(z + h) = \partial z + \partial h = 0$$

Household income is unchanged. This implies that

$$\partial z = -\partial h$$

Then we can analyse the effects on V_r as:

$$\frac{\partial \frac{h}{z}}{\partial z} = \frac{1}{z}\frac{\partial h}{\partial z} - \frac{1}{z^2}$$

This expression simplifies to:

$$\frac{\partial \frac{h}{z}}{\partial z} = -\frac{1}{z} - \frac{1}{z^2}$$

The expression is negative. Thus, when z, the female's income, rises due to the change in the benefit structure, the reservation utility of the male will decrease $(\partial f(\frac{h}{z}/\partial z) <$

0). When male and female household members are allowed to have distinct preferences, resource allocation is affected by the change in benefits because

$$\frac{\partial e_m(z+h, V_r)}{\partial z} \mid_{\partial z = -\partial h} \neq 0.$$

The rise in the reservation utility of the female resulting from an increase in her share of household resources h, will change the efficient allocation of resources within the household. Note that such changes in household resource allocation are efficient.

These results differ from those that would be obtained under unitary household preferences. If the male and female have the same preferences, $U(e_f, e_k) = V(e_m, e_k)$, changes in the distribution of income between males and females should have no effect.

Lundberg, Pollak, and Wales (1997) explain their finding that the reallocation of child benefits from fathers to mothers in the UK increased expenditure on children's clothing in terms of systematic differences in preferences across genders within households. Because the policy change in allocation of child benefits was exogeneous, Lundberg et al. could be reasonably confident that higher child clothing expenditures associated with higher shares of female earnings in households were not due to systematic tendancies of mothers with preferences for child expenditures to work more.

In Russia child benefits are allocated to mothers. Pensions make up the other majority of other social transfers administered, and the vast majority of Russian pensioners are female. The gendered nature of government transfer income in Russia, coupled with the large degree of exogeneoous variation in the levels of these transfers in the 1990s, provides an opportunity to examine both how resource allocations respond to transfer income and the gendered nature of these responses. As well, the variation in child benefits levels both across regions and over time allows investigation of the Kooreman labeling hypothesis for child benefits.

The goal of this section is to investigate how and why changes in transfer income affect household resource allocation. I estimate random effects panel data models of monthly expenditure on several types of household items (assignable and household public goods). The budget lines for which I estimate expenditure equations are: child services, alcohol and tobacco, food, family services and recreation, child clothing and shoes, and adult clothing and shoes. Additional results regarding the influence of various household income sources on (*i*) private transfers (financial gifts) to non-household members and (*ii*) consumer durable consumption are presented in Appendix C.

The econometric specification adopted draws on that of Kooreman (2000) and Edmonds (2002), with the important exception that panel data estimators are employed. In both of these previous studies, reduced-form specifications of equations governing different types of household expenditures were adopted. Tests of the equality of coefficients of child benefits income and other income were performed after the estimation of the expenditure equations. These studies do not attempt to estimate full household demand systems, but rather focus on demands for child and adult-specific goods, namely child clothing, adult clothing, and alcohol and tobacco expenditures. I follow a similar approach to testing how social transfer income is spent, but estimate reduced-form models which permit account to be made for unobserved heterogeneity between households, household composition, and the aging of the household over the 1994-2001 panel.

A non-negligible fractions of 0s is observed amongst households in each of the budget lines under consideration. Censored regression models are preferable to those which assume normality in the distribution of expenditure in this case. The panel is an unbalanced one because it includes households which attrit before 2001, so long as household composition did not change prior to attrition[5].

Censored regression models with fixed effects have recently been significantly advanced by Honoré and Kyriazidou (2000). However, the properties of fixed effect panel tobit estimators are still less developed than random effects estimators. Fixed effects estimators for this model are still generally subject to incidental parameters problems (see Heckman (1981), Woolridge (2002)). For this reason, only the results using random effects estimators are here presented.

Let y_{it} refer to total expenditure by a household on a budget line. The random effects tobit (censored regression) panel data model estimated in this subsection takes is defined as follows.

Let

$$y_{it} = \omega c_{it} + \theta s_{it} + \xi p_{it} + \gamma_1 n_{it} + \gamma_2 n_{it}^2 + \beta' X_{it} + \kappa z_{rt} + u_i + \epsilon_{it} \tag{1}$$

Then define:

$$y_{it}^* = \begin{cases} y_{it} & \text{if } y_{it} > 0 \\ 0 & \text{otherwise} \end{cases} \tag{2}$$

for $t = 1994...2001.IV$, where i, j refer to households. Here c_{it} refers to child benefits, n_{it} to earned income, s_{it} to social transfers (mainly pensions), and p_{it} to private transfers in year t. In 1994, a mean of 77% of received social transfer income derived from pensions, and in 2001 the figure was 80%. In the estimation sample 23% of households contained at least one retirement-age individual in 1994, and in 85% of these cases the retired individual was female.

[5] For more on random effects models, and particularly estimation of variance-covariance matrices in the case of unbalanced panels, the reader is referred to Green (2000), p. 577.

Here ω is the MPC from child benefits (c_{it}), θ the MPC from other social transfers (s_{it}), ξ the MPC from private transfers (p_{it}), and $\gamma_1 + 2\gamma_2$ gives the MPC from earned income (n_{it}) in the month prior to the RLMS interview. While γ_2 is generally very small, it is highly statistically significant and is thus included in the reported specifications. Expenditure on these budget lines is generally non-linear in income. However, perhaps because child benefits, private transfers, and social transfer income vary much less than does household income, I find that quadratic terms for these income sources are not necessary. Because of the benefits arrears crisis in Russia in the late 1990s, many households received no government transfers at times, and this makes it infeasible to estimate expenditure elasticities directly.[6]

Several recent tests of the collective model of the household have focused on labour supply responses of household members to changes in family non-labour income and to divorce laws (see Fortin and Lacroix (1997) and Chiappori, Fortin, and Lacroix (2004)). Standard micro theory predicts that the labour supply response to an increase in government or private transfers is negative due to wealth effects. It is important to note, however, that potential labour supply responses to transfer income, while resulting in $cov(n_{it}, c_{it}) < 0$, $cov(n_{it}, s_{it}) < 0$, and/or $cov(n_{it}, p_{it}) < 0$, would not cause endogeneity problems in the analysis of household demands to be undertaken here. Here c_{it}, n_{it}, p_{it}, and s_{it} are explanatory variables.

In X_{it} variables controlling for the age and composition of the household, and aging of members over time, are included. Only households whose composition remains constant during the panel are included in estimation. Changes in household composition would be expected to have effects on bargaining relations with households as well as effects on the structure of household demands, and would thus conflate income/social transfer effects. I only include households who have at least one child under age sixteen throughout the period of the panel. Under the universal system of child benefit all households with children under 16 received benefits. Note that no assumptions regarding household equivalence are necessary. Controls for the value of home production of each budget line consumed *within* the household are also included in all specifications. Home production which is sold is included in the value of earned household income.

The demand system estimated is incomplete. Here z_{rt} refers to local prices in interview site r at time t. Unfortunately, information on clothing prices, family services, and supplies, is not collected in the yearly RLMS community survey. Local price vectors are available only for alcohol and tobacco and for food. [7] For discussion of the construction

[6]Zero values occur both on the left and right hand sides. Estimation results using different functional forms show that the results presented are robust to these assumptions. These are available on request from the author.

[7]Thus I am unable to test cross-equation restrictions on the demand system, or to estimate the full

of the price vectors, the reader is referred to Appendix B.

In the random effects specification, u_j includes all time-invariant, household specific unobserved heterogeneity. I assume that u_j is Gaussian distributed ,and independent of observables. Here β and x_{it} are k X 1 column vectors, ϵ_{it} is distributed $N(0, \sigma_\epsilon^2)$, and u_i is distributed $N(0, \sigma_u^2)$. I assume that $E(u_i u_j) = 0$, $E(u_i \epsilon_{it}) = 0$, and $E(\epsilon_{it}\epsilon_{it+1}) = 0$ $\forall i \neq j$. Because a non-trivial fraction of households report zero expenditures on each of the household budget items, this is the chosen threshold. In this case it seems reasonable to assume that ϵ_{it} and ϵ_{it+1} are independently distributed. The available measures of income, expenditure, and childbenefits refer to the month prior to the RLMS interview, which is conducted yearly[8].

In the panel data censored regression model, the (Gaussian) distributed unobservable components are assumed to be time-invariant. Thus the likelihood function is somewhat more complicated than in standard censored regression frameworks. In the random effects model, the likelihood function for each household is the integral of a product. Define the indicator variable $w_{it} = I(y_{it}^* > 0)$.

For ease of exposition, I collect the observables in the term $\Psi' H = \omega c_{it} + \theta s_{it} + \xi p_{it} + \gamma_1 n_{it} + \gamma_2 n_{it}^2 + \beta' X_{it} + \kappa z_{rt}$.

The likelihood contribution for each household i is:

$$L_i = \int_{-\infty}^{\infty} \prod_{t=1}^{T_i} \left[f\left(y_{it}^* - \Psi' H - u_i, \sigma_\epsilon \right) \right]^{w_{it}} \left[\Phi\left(\frac{-\Psi' H - u_i}{\sigma_\epsilon} \right) \right]^{(1-w_{it})} f\left(u_i, \sigma_u\right) du_i \quad (3)$$

Quadrature techniques are used to calculate the integrals, so the full model can be estimated by regular maximum likelihood. In the reported estimations, Gauss-Hermite quadrature with 8 points is used. For more on the estimation of panel tobit models, the reader is referred to Honoré and Kyriazidou (2000).

Prior to discussing estimation results, I briefly outline characteristics of the estimation sample. Means of household income and social transfers for the full sample are presented in Table II. Comparing Columns (I) and (II) of row (i) shows that real household incomes have dropped on average nearly 30% through the sample period. As well, per capita child benefits have nearly halved, when received (row (ii)). While under the universal child benefits system operating in 1994, about 58% of all households with children under 16 actually received the child benefit in the month prior to the RLMS interview, this fraction had decreased to about 44% under the mix of targeting schemes operating in 2001 (row (iii), columns (I) and (II)). Table II also shows that private transfers, when received, were

demand system.

[8]Due to funding problems, the RLMS was not administered in 1997 or 1999. However, this is not expected to have any biasing impact on the results, and no corrections for this are made in estimation.

large relative to household income. While only about 1/4 of households reported receiving private transfers in either 1994 or 2001, the mean amount received was nearly half of mean household income in 1994. In 2001 mean private transfers, when received, were equal to mean household income.

Food expenditures (both consumed at home and eaten out) are the single largest budget item in Russian households. In 1994 mean food expenditure was 50% of mean total household income. In contrast, man expenditures on children's activities and supplies were relatively small fractions of mean incomes (2% in 1994), as were expenditures on alcohol and tobacco (4%) and services and recreation (8%).

Results are presented in Table III for the sample of households containing children 0-16. Column (I) presents the results for budget amounts spent on children's school supplies and recreational activities, column (II) presents results for the adult-assignable items alcohol and tobacco, Column (III) presents the results for food expenditure, and column IV presents the result for expenditure on family services and recreation. At the bottom of each column, Wald tests of the null hypothesis that the MPC from different sources of income (child benefits, other social transfers, private transfers, and earned income) is equal are displayed in Table IV.

The MPC on childrens's supplies[9] (Column (I)) and family services and recreation (Column (IV)) is not statistically different from that any of the other sources of household income considered. Column (II) suggests that the marginal propensity to spend money on adult-assignable items (alcohol and tobaccco) is slightly higher for child benefits than other income. The Wald test of the equality of the coefficients, $\omega = \gamma_1 + \gamma_2$ rejects at the 10% level. Households have MPCs on alcohol and tobacco which are about three times as high from child benefits money than either earned income or private transfer income. Interestingly, social transfer income (much of which accrues to female pensioners in 3-generation households) has a much lower MPC on alcohol than either child benefits money or earned income, suggesting that pensioners do not let their money be spent on 'sin' items.

However, the largest differences in consumption propensities from different income sources are found in Column (III) of Table III, which examines budget expenditures on food. Here ω is three times as large as $\gamma_1 + 2\gamma_2$. As well, social transfer income has a far higher MPC on food than either earned income or private transfer income. Child benefits and other social transfers are much more likely to be spent on food than either unearned income from private transfers or earned income.

Column (IV) presents results for expenditure on family services. Wald tests of the

[9]Respondents are asked "Did your family spend money in the last 30 days and if so how much?" on a list of items. This item is under the budget line "1. For child support and fees for children's attendance at preschools, schools, clubs, societies, payment for private lessons, rehearsals."

equality of expenditures from different income sources, presented in Column (IV) of Table IV, suggest that more of private transfer income is spent on these services than either type of social transfers or earned income.

Child benefits income generally are received by mothers on their children's behalf. Amongst households containing more than two adults (30% of the sample), much of social transfer income comes from pensions to elderly women. It is important to determine whether the differential propensities to consume food, alcohol and tobacco from child benefits, social transfer income, and earned income presented in Table III derive from systematic differences in expenditure priorities between the sexes. Because women exercise control over much of the government transfer income considered in the sample, the resource allocation effects observed may in fact be attributable to women having more access to household resources when these transfers are larger[10]. Such effects would be inconsistent with a unitary household model, but they are consistent with collective models of the household which allow adults in the household to maintain distinct preferences, and with models assuming cooperative or non-cooperative Nash bargaining.

I am interested in investigating whether systematic differences between the sexes in resource allocation priorities can explain high observed MPCs on food from government transfer income. In order to examine this hypothesis I estimate the same expenditure equations of the previous section with controls and interaction terms for all-female households. Just over ten percent of households in the sample are made up of all-female adults. If the same results hold in all-female households, systematic differences in preferences between men and women (the primary recipients of transfers) cannot be driving the results.

I Gender effects on intrahousehold allocations

Simple tabulations of the height for age scores of children in the 1994 RLMS suggest that the gender of adults plays a strong role in the nutritional status of children in Russia. Amongst children aged 0-5, height-for-age scores are 0.25 standard deviations less where fathers are present in the household, despite the fact that mean incomes are one quarter higher in these households and household size is essentially the same. In households containing only female adults, height-for-age scores are 0.56 standard deviations higher than in mixed-gender households, despite the fact that mean incomes are 40% lower. While unobserved features of these households may be driving some of these strong differences, I am able to control for unobserved heterogeneity between households in estimation of the effect of gender composition on resource allocations. In 1994, mean food expenditure

[10]Unfortunately, the RLMS survey does not contain information on who obtained each transfer. However, one knows the gender and age of all household members and whether or not pension income was received.

was 64% of mean income in all-female households, far higher than the 50% figure for the full sample.

Define ω_f as the specific effect of child benefits in all-female adult households. Similarly, let θ_f, ξ_f, and $\gamma_{1f}+2\gamma_{2f}$ represent these interaction terms for social transfers, private transfers, and household income, respectively. Table V presents the results of the estimation of censored panel regression models which take into account the gender composition of the household. As in Table III, I find that household income source is not important in decisions regarding expenditure on children's activities. However, for a given income and household composition, all-female households spend 308 (June 1992) roubles less on children's activities than those containing at least one adult male. Similarly, Column II shows that all-female households spend 728 (June 1992) roubles less per month on alcohol and tobacco for a given household income and composition. Tests of the equality of coefficients are presented in Table VI. Column (II) of Table VI shows that for all-female households income source does not affect alcohol and tobacco expenditures. However the results of the previous subsection are robust to the inclusion of these gender controls and suggest that households of mixed gender spend significantly more of child benefits money on alcohol and tobacco than other income.

Column (III) of Table V contains the results of expenditure equations for food. While all-female adult households spend significantly more on food for a given income level and household composition, differences in expenditure patterns across income type are not apparent for this subsample. Table VI shows that, while the general results of the estimation presented in Tables III and IV are robust to controls for adult gender composition, all-female households do not appear to alter their expenditures on food when income comes from different sources. These results suggest that (*i.*) women as a group prioritise food over expenditures such as children's activities and alcohol and tobacco, and (*ii.*) a major reason why the MPC on food from government transfer income is three times as high as from earned income in the general sample is due to women having control over this money. What is less clear is why alcohol and tobacco expenditures from government transfer income would be higher in mixed-gender households than in all female households. One explanation is that part of transfers controlled by women in mixed-gender households are used as gifts to men to 'keep the peace'. Note that, because I control for unobserved heterogeneity across households, systematic unobserved differences between all-female and mixed-gender households cannot be driving the general results. Thus the idea that alcohol and tobacco consumption is more enjoyable in mixed gender households can be eliminated as an explanation.

Column IV of Table V presents results for expenditure on family services. Family services include expenditures on transportation, the tailoring and repairing of shoes and clothes, the repair of household appliances, the construction and repair of housing, laundry, recreation services, medical treatment, vacation costs, and courses for adults. All-female households spend 9 of every ten roubles received as private transfers on such family services. Private transfer income is treated very differently than earned income in all-female households, and than in mixed-gender households. In all-female households private transfer income appears to be 'earmarked' for family services.

II Assignable clothing

The 1998 through 2001 rounds of the RLMS do contain information on child and adult-specific expenditures on clothing and shoes. Clothing may be a better child-assignable good that expenditure on children's activities, due to the fact that activities are generally continuous. Clothing requires no commitments to further consumption, and thus should be more responsive to transitory income changes. Given that the coefficient ω for adult assignable goods (alcohol and tobacco) for the full sample is in contrast to that predicted by the labeling hypothesis, it is important to look at these assignables.

Table VII presents the results of random effects tobit estimation of expenditures on assignable clothing, for the 1998-2001 period. Column (I) presents the results for expenditure on children's clothing and shoes, while Column (II) presents the specification for adult's clothing. Table VIII presents the results of statistical tests of the equality of coefficients. In neither case is the null hypothesis that child benefits money is treated the same way as other household income rejected using Wald tests. Thus little support for the labeling hypothesis is found from examination of these assignable expenditures. Still, there are some other results of interest. While social transfer income is a statistically significant determinant of expenditure on children's items, this is not true for adult's shoes and clothing. The MPC on adult clothing from private transfer income is significantly less than from earned income.

III Assignable clothing with accounting for gender composition

Table IX I present the results of maximum likelihood censored panel regression estimation with controls for the gender composition of the household. Column (I) presents the results for children's clothing and shoes. I find that ω_f, θ_f, ξ_f and $\gamma_{1f} + 2\gamma_{2f}$ interaction terms, as well as the dummy for all-female households, are statistically insignificant. The source of household income is not important in spending on children's clothing in either all-female or mixed gender households.

Column (II) of Table IX presents results for the estimation for adults. As in Table VII, expenditure on adult clothing from private transfers is significantly less than from earned income. All-female households also spend 214.4 roubles less on adult clothing than mixed-gender households of the same income and composition. Given that the specification controls for time-invariant unobserved heterogeneity between households, these results cannot be attributable to substitutability of clothes amongst members in female households. Table X presents the results of tests of the equality of coefficients. The results for expenditures on assignable clothing strongly suggest that no labeling effects of child benefits exist in Russia. Private transfers accruing to households are generally not spent on adult-specific items.

In this section I found that the MPC on food from child benefits and other social transfer income is far higher than that from either earned income or private transfer income. I presented suggestive evidence that these higher expenditures on food from transfers are attributable to women having more control over these transfers than over other household income. I found that all-female households appear to have very different spending priorities from mixed-gender households. These results appear to be attributable to systematic differences across the sexes in preferences for food versus alcohol and tobacco. These results are robust to the specification of the time-invariant unobserved heterogeneity term, u_i, as a linear function of time-varying covariates, as in Chamberlain (1984). Results of these robustness checks are not presented here, but are available on request.

In the following section I examine whether or not the greater food expenditures generally found from transfer income translate to better nutrition amongst children and other household members.

VI More benefits, better nutrition?

In the 1994-96 RLMS surveys, caloric intake diaries were recorded for all household members. On the basis of these reports, the University of North Carolina-Chapel Hill and its Russian partners have constructed variables on the percentage of daily calories obtained from protein and fat, respectively. Given that more calories does not necessarily imply better nutrition, it is of interest to relate information on the fat and protein content of diets to household income and social transfers. I first examine child nutritional intake, and then that of the adult (above 16) members of households containing these children. Note that in the 94-96 period child benefits were still mainly universal.

In 2002 The US Food Nutrition Board released guidelines on Acceptable Macro-Nutrient Distribution Ranges (AMDR). The AMDRs for protein are 10-35% of the diet for adults, 5-20% for young children, and 10-30% for older children. In Table XII I present sample statistics on the diet composition of children and adults in Russia, for the 1994-96 period. Row (i) of Table XII shows that, for children, protein consumption in the diet declined significantly in between 1994 and 1996, from 12% to less than 9% of daily calorie intakes. These protein levels suggest that on average, children in Russia are marginal for obtaining adequate protein in the diet. In 1996 children in child benefits arrears regions had significantly lower protein consumption than non-arrears regions.

The AMDR recommendations for fat in the diet are 20-25% for adults, 30-40% for children under age 3, and 25-35% for children 4-18 years old. For both children and adults in Russia, fat intake appears to be in the mid to upper range of these recommendations. In 1994, children in benefits-receiving regions had significantly higher fat intakes than those in benefits-arrears regions. However, by 1996, this was no longer true.

For adults in households containing these children, a similar pattern emerges across years. Protein consumption was significantly less in child benefits-arrears regions than in benefits-receiving regions in 1996. Protein consumption in 1996 was below the margins of the AMDR recommended levels. However, protein consumption of adults is generally higher than that of children.

As mentioned previously, arrears in government transfers arose from regional budget crises associated with poor local economies. Thus low incomes levels, correlated with these budget crises, may be driving the finding that protein consumption was lower in 1996 in benefits arrears regions. The goal of the next subsection will be to separate the impact of child benefits from that of other household income in influencing diet composition.

I Child nutrition

Table XII presents the results of random and fixed effects estimation of the relation between child benefits, household income, and diet composition for children under 14 in 1994 in the RLMS panel. For details of the estimation of these models, the reader is referred to Green (2000). Columns (I) and (II) present results for protein intake for the random and fixed effects specifications, respectively. The Breusch-Pagan LM test rejects the null hypothesis that $Var(u_i) = 0$ (that there are no individual random effects) at the 1% level. However, the Hausman test, which tests the null hypothesis that these individual random effects are uncorrelated with other variables in the model, suggests rejection. Fixed effects estimators, which are consistent in this case (the random effects estimator is not) are preferable.

The F-test of the equality of the MPC from child benefits and other income rejects at the 10% level, under the fixed effects specification. For children, the consumption of protein increases substantially more from one rouble of child benefits income than it does from other household income. Note also that the positive effect of child benefits on calories from protein consumption is statistically significant at the 10% level. Statistical tests of the equality of coefficients are presented in Table XIII. An extra 1000 roubles of child benefits income translates to about a 0.2% increase in protein consumption in daily calories. Note, however, that the mean value of the child benefit per child (when received) was about 170 roubles in 2001. The effect, while statistically significant, is small. However, it does importantly corroborate the evidence on food expenditures from child benefits income presented in the previous section. Table III suggests that the MPC on protein from private transfer income, ξ, is significantly less than that from earned income according to the (preferable) fixed effects estimates.

A similar exercise is carried out in Columns (III) and (IV) of Table XII for the consumption of calories from fat. Here the Breusch-Pagan LM test again suggests random effects but the Hausman test is unambiguous, rejecting the null hypothesis that the differences in coefficients between the random and fixed effects specifications are systematic at the 1% level. Fixed effects estimates are consistent while the random effects estimator is not. Column (IV) of Table XIII shows that the F-test of the equality of the marginal propensities to consume from child benefits and other household income accepts the null hypothesis of equality at the 10% level. Child benefits appear to be associated with more protein for children, but not more fat.

II Adult nutrition in households containing children

A similar exercise was undertaken to examine adult nutrition amongst those in the households containing the children of the previous subsection. The goal is to understand the extent to which the higher MPC on food from child benefits translates to food that is only consumed by children.

The results of the analysis of fat and protein consumption of adults are presented in Table XIV. Columns (I) and (II) present the random effects and fixed effects results for protein consumption, respectively. The Hausman test (bottom of Column II) suggests that the fixed effects estimates are to be preferred (random effects estimators are not consistent in this case). In the fixed effect specification, the null hypothesis that child benefits income is related to protein consumption in the same way as income from other sources is rejected at the 1% level. The coefficients on both child benefits and earned income are all significant at the 1% level. A 1000 rouble increase in child benefits income is associated with a 0.2% increase in adult protein consumption as a fraction of dietary intake.

Columns (III) and (IV) of Table XIV present estimates for the relationship between fat in the diet, child benefits, and other income, for this group of adults. Again, the Hausman test suggests that the fixed effects specification is preferrable. In contrast with the results for children, the coefficient on child benefits is positive and significant under the fixed effects specification. As well, the F-test of the null hypothesis that the marginal propensities to consume fat from child benefits and other income are equal rejects at the 1% level. A 1000 rouble increase in child benefits is associated with a 0.9% increase in daily fat consumption of adults as a percentage in the diet. Changes in household income from social transfers also appear to have strong effects on fat consumption of adults. Table XV presents the results of statistical tests of the equality of the MPC from different income sources. The MPCs on adult protein and fat from private transfer money, ξ is significantly less than those from earned household income, according to the fixed effects estimates. The MPC on adult fat from social transfer income, θ, is significantly higher than that from earned household income and private transfer income.

In this subsection I have provided some evidence that child benefits money is not only associated with more food purchases, but also statistically-significant (although small) improvements in the protein intake of both adults and children in households. It is also of interest that, while social transfer income was found in the previous section to increase expenditure on food, it is not so strongly associated with increased protein intakes for either children or adults. Rather the higher MPC on food from social transfer income appears to translate into a significant increase in fatty foods consumed only by adults within the household. While both types of government transfer income appear to increase food expenditures, the type of food purchased appears to vary with the either nature of the transfer or the recipient. Given that it was found that fixed effects estimators are preferable to random effects estimators, it is unfortunately unfeasible to investigate the potential role of gender composition of the household in the diet using the panel.

VII Conclusions

In this paper I have used panel data from Russia to examine the effect of government transfers on household expenditure patterns, paying special attention to the gender composition of households. I have focused primarily on child benefits, which were universal in the early 1990s in Russia, but which became increasingly targeted towards the end of the decade. To my knowledge, this is the first paper to use household panel data to examine how household expenditure patterns respond to government transfer income and the effects of the gender composition of the household on allocations.

This paper concurs with a growing body of research which suggests systematic differences in preferences by gender. Several recent papers have shown distinct gender differences in attitudes to inequality, redistribution, and public goods provisions. Using subjective answers on redistribution from 1996 Russian survey data, Ravallion and Loshkin (2000) find that women favor redistributive policies more than men. Edlund and Pande (2002) show that the rise in divorce in the US since the 1960s is strongly associated with the rise in the fraction of women voting for left-wing political parties. Using 1990s data from the German Socio-Economic Panel, Edlund, Haider, and Pande (2004) find that transitions out of marriage make women, but not men, more sympathetic to left-wing political parties. Duflo and Chattopadhyay (2004) examine the effect of the gender of Pradhand members in India on decisions regarding public goods provisions. Using the natural experiment provided by a 1992 amendment to the Indian constitution which reserved seats for women, Duflo and Chattopadhyay (2004) finds that women who are elected under the reservation policy invest more in public goods closely related to women's concerns than do men. Interpreted in this light, the results of this paper suggest that women are more concerned with the nutrition of both children and adults in the household than are men. These findings are broadly consistent with Lundberg, Pollak, and Wales (1997), Duflo (2003), and Thomas (1990). In using panel data estimators which control for unobserved heterogeneity between households I show that these findings are not caused by systematic unobservable preference differences across household type.

One potential reason for greater preferences for alcohol and tobacco consumption in households containing men is simply that men are more likely to be addicted to alcohol and/or cigarettes than are women. According to the Russian Statistical Agency Goskomstat (*www.gsk.ru*), approximately 30% of men and 15% of women are addicted to alcohol. McKee, Bobak, Rose, Shkolnikov, Chenet, and Leon (1998) find that 65% of Russian men aged 18-24 and 73% of those aged 25-34 smoked in 1996. Amongst women, smoking is much more common among the young (27% amongst those aged 18-34) than among the middle-aged and elderly (5% amongst those aged 55 and older). Gender gaps in both alcohol and tobacco addiction are very large in Russia. Because substance addiction strongly affects consumption choices, the fact that a man is more likely to be addicted than a woman may be behind apparent gender differences in preference for consumption of these items.

I find that child benefits do not have statistically different impacts from earned income, private transfer income, or other social transfers, on the consumption of either child-specific supplies or child and adult clothing. None of these findings concurs with Kooreman's labeling hypothesis that child benefits are spent disproportionately on child-assignable goods. Thus it appears that the labeling effect of the child benefits system found for the Netherlands may be specific to either the culture or level of economic development in the Netherlands. Edmonds (2002) also finds no evidence of a labeling effect of child benefits using Slovenian data.

The findings of this paper raise some important policy concerns for Russia. There were several motivations given for the adoption of the universal child benefits system in Russia in 1991. One reason for the institution of child benefits was to help families with children, who were expected to experience income shocks during economic transition. Another reason was as a pro-natalist policy, to mitigate expected declines in the birth rate. This paper suggests that child benefits are useful on the first front. The finding that child benefits are largely spent on relatively high protein food, eaten by both adults and children, suggests that these benefits had positive effects in mitigating poverty within households. Targeting of child benefits in Russia has been instituted gradually across regions since 1996. The goal of targeting is to eliminate non-payment problems of cash-strapped regions, thus ensuring that those most in need obtain benefits. However, this paper suggests that, at the mean, both adults and children in households containing children had protein intakes at the margin of what the US Food Nutrition Board recommendations in 1996 in Russia, before targeting was widely instituted.

Finally, this paper is related to an ongoing debate in the development economics literature regarding the direction of causality between nutrition and incomes. Fogel (1994) advanced the thesis that income responds to nutrition in development processes. Bouis (1994), and Bouis and Haddad (1992) offered a re-examination of the relation between calories and income in this light. However, these findings have been controversial. Using data from the 1983 Maharashtran State portion of the Indian National Sample Survey, Shankar and Deaton (1996) suggest that nutrition is driven by income, and not *vice verse*. My finding that protein consumption of children improves both with household income and (more strongly) with social transfer income, supports the Shankar and Deaton (1996) view.

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Table I: Real Per-Capita Child Benefit Levels By Region, when received, 1994.IV and 2001.IV (in June 1992 Roubles)

region	1994	2001	region	1994	2001
St. Petersburg City	223.175	135.694	Tambov Oblast: Uvarovo CR	223.799	106.027
	(20.97)	(18.78)		(28.26)	(13.72)
Moscow City	237.087	86.926	Tatarskaja ASSR: Kazan	425.698	880.905
	(26.15)	(7.49)		(83.79)	(85.11)
Moscow Oblast	219.797	84.940	Saratov CR	209.043	64.929
	(23.36)	(1.85)		(25.45)	(18.55)
Komi ASSR: Syktyvkar	653.18	181.192	Saratov Oblast: Volskij Gorsovet Rajon	256.136	87.040
	(109.05)	(78.01)		(37.06)	(3.56)
Komi ASSR: Usinsk CR	452.203	194.884	Volgograd Oblast: Rudnjanskij Rajon	259.279	84.007
	(92.82)	(36.53)		(19.59)	(0.13)
Leningrad Oblast: Volosovskij Rajon	449.850	327.506	Kabardino-Balkarija, Zol'skij Rajon	192.419	92.922
	(99.58)	(243.57)		(14.21)	(6.64)
Smolensk CR	260.980	83.787	Rostov Oblast: Batajsk	273.653	105.042
	(58.77)	(0.31)		(75.75)	(21.26)
Kalinin Oblast: Rzhev CR	217.651	84.093	Krasnodar CR	334.547	120.575
	(29.482)	(0.19)		(54.85)	(0.00)
Tul'skaja Oblast: Tula	260.163	178.720	Stavropolskij Kraj: Georgievskij CR	200.783	166.379
	(44.37)	(62.04)		(29.46)	(87.39)
Kaluzhskaja Oblast: Kuibyshev Rajon	354.538	94.646	Krasnodarskij Kraj: Kushchevskij Rajon	316.481	246.899
	(111.59)	(10.47)		(62.70)	(73.62)
Gorkovskaja Oblast: Nizhnij Novgorod	334.684	840.008	Cheliabinsk	230.040	121.303
	(70.90)	(785.94)		(13.62)	(22.83)
Chuvashskaja ASSR: Shumerlja CR	293.305	91.653	Kurgan	327.819	84.478
	(41.73)	(7.89)		(66.35)	(11.98)
Pezenskaja Oblast: Zemetchinskij Rajon	196.739	85.745	Udmurt ASSR: Glasov CR	252.626	97.263
	(6.86)	(1.81)		(37.16)	(0.80)
Lipetskaja Oblast: Lipetsk CR	231.299	100.504	Orenburg Oblast: Orsk	217.315	100.018
	(26.23)	(18.49)		(8.93)	(10.28)
Perm Oblast: Solikamsk City Rajon	227.742	130.623	Altaiskij Krai: Kur'isnkij Rajon	221.150	102.859
	(8.80)	(26.45)		(18.27)	(7.70)
Cheliabinsk Oblast: Krasnoarmeiskij Rajon	219.143	238.768	Krasnojarskij Kraj: Krasnojarsk	281.399	188.098
	(17.95)	(122.03)		(51.71)	(72.27)
Tomsk City and Rajon	293.126	401.5217	Vladivostok	224.916	98.345
	(33.14)	(292.99)		(9.40)	(2.93)
Khanty-Manskiiskij AO: Surgut CR	463.828	360.411	Krasnojardkij Kraj: Nazarovo CR	413.147	126.605
	(70.64)	(181.52)		(108.75)	(75.97)
Altaiskij Kraj: Biisk CR	316.435	249.949	Amurskaja Oblast: Arkharinhskij Rajon	421.021	154.254
	(74.97)	(98.24)		(32.23)	(19.19)

Notes: Calculated from sample used in presented estimation. Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included.

Table II: Sample means of household income and social benefits, households with children under 16, RLMS 1994.IV-2001.IV

	1994	2001
real earned household income[a]	8483.92	5447.88
	(1012.82)	(267.19)
child benefits/child, when received	290.17	140.997
	(9.28)	(11.98)
social transfers, when received	1956.35	1537.22
	(87.72)	(69.14)
private transfers, when received	5792.99	5469.7
	(455.40)	(564.45)
fraction receiving child benefit	0.5753	0.4387
	(0.015)	(0.020)
fraction receiving social transfer	0.2603	0.352
	(0.013)	(0.019)
fraction receiving private transfers	0.2313	0.259
	(0.013)	(0.018)

Notes: Standard errors in parentheses. Real (June 1992 rouble) household income net of child benefits. Income and measures are for the 30 days prior to interview.

[a] Note that household income measure excludes child benefits, social, and private transfers.

Table III: Random Effects Panel Tobit Estimation: Expenditures by Household Income Source, all household types, RLMS 1994.IV-2001.IV

	children's activities	alcohol and tobacco	food	family services
child benefits, ω	0.0487	0.09068 **	0.3760 **	0.1232
	(0.031)	(0.029)	(0.117)	(0.108)
social transfer income, θ	0.0367	-0.01461	0.4565**	0.1321*
	(0.023)	(0.023)	(0.092)	(0.081)
private transfer income, ξ	0.0307**	0.03944**	0.1210**	0.1314**
	(0.004)	(0.004)	(0.017)	(0.015)
real earned hh. income, γ_1	0.0351 **	0.0361 **	0.1147 **	0.0744**
	(0.005)	(0.004)	(0.010)	(0.009)
real earned hh. income^2 [a], $\gamma_2 * 100000$	-0.0579 **	-0.0325 **	-0.0111**	-0.0075
	(0.013)	(0.008)	(0.001)	(0.001)
local price	no	yes	yes	no
σ_u	112.984	479.54	2183.76	967.01
	(0.000)	(25.26)	(105.02)	(95.73)
σ_e	907.256	841.49	3579.34	3452.37
	(18.10)	(13.715)	(46.83)	(46.35)
ρ	0.0152	0.2451	0.2713	0.0727
	(0.001)	(0.0210)	(0.021)	(0.014)
Wald $\chi^2(13)$	206.63	253.69	352.05	192.66
Prob $> \chi^2$	0.0000	0.0000	0.000	0.0000
uncensored	1502	2362	3800	3343
left censored	2384	1524	86	543
Number of observations	3886	3886	3886	3886
Number of households	804	804	804	804
Log likelihood	-13555.752	-20442.411	-36941.75	-32481.72

Notes: Full demographic controls constant, and controls for income in kind also included. Standard errors in parentheses.** significant at 5% level, * significant at 10% level. Real (June 1992 ruble) household income net of child benefits, social transfers, and private transfers.
[a]real earned hh. income^2 /100000

Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Controls for the number of children in the age groups 1-1.5, 1.5-6, and 6-16, the number of working-age adults, and the number of pension-age household members are included in all specifications. Controls for the receipt of any in-kind payments from workplaces are included, as well as controls for the value of home-consumed home production and a constant.
Income and expenditure measures are for the 30 days prior to the interview.

Table IV: Tests of equality of MPC coefficients, all household types, RLMS 1994.IV-2001.IV

	children's activities	alcohol and tobacco	food	family services
Test: $\omega = \gamma_1 + 2\gamma_2$				
Wald test $\chi^2(1)$=0	0.06	2.96	4.65	0.07
Prob> χ^2	0.8049	**0.0852**	**0.0311**	0.7858
Test: $\theta = \gamma_1 + 2\gamma_2$				
Wald test $\chi^2(1)$=0	0.18	4.80	13.69	0.41
Prob> χ^2	0.6754	**0.0285**	**0.0002**	0.5224
Test: $\xi = \gamma_1 + 2\gamma_2$				
Wald test $\chi^2(1)$=0	6.13	0.06	0.00	8.18
Prob> χ^2	0.0133	0.8046	0.9886	**0.0042**
Test: $\omega = \theta$				
Wald test $\chi^2(1)$=0	0.00	7.41	0.47	0.03
Prob> χ^2	0.9572	**0.0065**	0.5974	0.8732
Test: $\omega = \xi$				
Wald test $\chi^2(1)$=0	0.07	3.01	4.70	0.03
Prob> χ^2	0.7941	**0.0825**	**0.030**	0.8587
Test: $\xi = \theta$				
Wald test $\chi^2(1)$=0	0.06	5.15	12.93	0.00
Prob> χ^2	0.8008	**0.0233**	**0.0003**	0.9941

Notes: Tests refer to estimates presented in the previous table for the full sample of households.

Table V: Random Effects Panel Tobit Estimation: Expenditures by Household Income Source, interactions for all-female households, RLMS 1994.IV-2001.IV

	children's activities	alcohol and tobacco	food	family services
child benefits, ω	0.0625*	0.1018**	0.4099**	0.1644
	(0.033)	(0.031)	(0.129)	(0.118)
child benefits*all-female, ω_f	-0.1329	-0.0592	-0.1374	-0.2416
	(0.100)	(0.092)	(0.300)	(0.275)
social transfer income, θ	0.0275	0.0012	0.5081**	0.1729**
	(0.024)	(0.024)	(0.097)	(0.084)
social transfers*all-female, θ_f	0.0990*	0.0606	-0.1475	-0.2620
	(0.055)	(0.067)	(0.228)	(0.196)
private transfer income, ξ	0.0271**	0.0376**	0.1198**	0.1185**
	(0.004)	(0.004)	(0.017)	(0.015)
private transfers*all-female, ξ_f	0.2128**	0.0233	0.0623	0.8088**
	(0.028)	(0.032)	(0.123)	(0.1097)
real earned hh.income, γ_1	0.0315**	0.0336 **	0.1136**	0.0768**
	(0.005)	(0.004)	(0.010)	(0.008)
real earned hh.income*all-female, γ_{1f}	0.0461**	-0.0147	0.1067	-0.0113
	(0.021)	(0.024)	(0.087)	(0.076)
real earned hh. income ^{2a}, $\gamma_2 * 100000$	-0.0516**	-0.0300**	-0.0110 **	-0.0077**
	(0.012)	(0.007)	(0.001)	(0.000)
real earned hh. income ^{2a}*all-female, $\gamma_{2f} * 100000$	-0.0802	0.0285	-0.4216	-0.1339
	(0.266)	(0.075)	(0.294)	(0.265)
all-female household	-308.4219**	-728.1951**	636.2305**	176.6892
	(92.976)	(108.981)	(179.0919)	(317.0709)
local price	no	yes	yes	no
σ_u	8.8911	481.674	2166.691	917.0751
	(4.343)	(25.507)	(104.9105)	(96.75)
σ_e	908.9058	835.88989	3578.595	3434.237
	(17.774)	(13.637)	(46.864)	(46.070)
ρ	0.098	0.2492	0.2682	0.0665
	(0.010)	(0.021)	(0.020)	(0.013)
Wald $\chi^2(13)$	281.14	347.41	361.28	256.19
Prob $> \chi^2$	0.0000	0.0000	0.0000	0.0000
uncensored	1500	2358	3794	3337
left censored	2380	1522	86	543
Number of observations	3880	3880	3880	3880
Number of households	804	804	804	804
Log likelihood	-13520.04	-20359.538	-36879.471	-32397.469

Notes: Standard errors in parentheses.** significant at 5% level, * significant at 10% level. Real (June 1992 ruble) household income net of child benefits, social transfers, and private transfers. Controls for the number of children in the age groups 1-1.5, 1.5-6, and 6-16, the number of working-age adults, and the number of pension-age household members are included in all specifications. Controls for the receipt of any in-kind payments from workplaces are included, as well as controls for the value of home-consumed home production and a constant.
Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Income and expenditure measures are for the 30 days prior to the interview.

[a]real earned hh. income^2 /100000

Table VI: Tests of equality of MPC coefficients: all households with interactions for all-female households, RLMS 1994.IV-2001.IV

	children's activities	alcohol and tobacco	food	family services
		All households		
Test: $\omega = \gamma_1 + 2\gamma_2$				
Wald test $\chi^2(1)$=0	0.82	4.57	5.23	0.54
Prob> χ^2	0.3643	**0.0325**	**0.0222**	0.4625
Test: $\theta = \gamma_1 + 2\gamma_2$				
Wald test $\chi^2(1)$=0	0.03	1.70	16.22	1.27
Prob> χ^2	0.8737	0.1928	**0.0001**	0.2589
Test: $\xi = \gamma_1 + 2\gamma_2$				
Wald test $\chi^2(1)$=0	0.47	0.50	0.10	6.06
Prob> χ^2	0.4951	0.4773	0.7464	**0.0138**
Test: $\omega = \theta$				
Wald test $\chi^2(1)$=0	0.66	6.17	0.36	0.00
Prob> χ^2	0.4182	**0.0130**	0.5493	0.9548
Test: $\omega = \xi$				
Wald test $\chi^2(1)$=0	1.10	4.13	5.00	0.15
Prob> χ^2	0.2947	**0.0422**	**0.0254**	0.7007
Test: $\xi = \theta$				
Wald test $\chi^2(1)$=0	0.00	2.15	15.43	0.40
Prob> χ^2	0.9884	0.1425	**0.0001**	0.5270
		All-female households		
Test: $\omega + \omega_f = \gamma_1 + \gamma_{1f} + 2\gamma_2 + 2\gamma_{2f}$				
Wald test $\chi^2(1)$=0	2.28	0.07	0.03	0.29
Prob> χ^2	0.1313	0.7951	0.8562	0.5870
Test: $\theta + \theta_f = \gamma_1 + \gamma_{1f} + 2\gamma_2 + 2\gamma_{2f}$				
Wald test $\chi^2(1)$=0	0.68	0.38	0.36	0.56
Prob> χ^2	0.4108	0.5365	0.5500	0.4530
Test: $\xi + \xi_f = \gamma_1 + \gamma_{1f} + 2\gamma_2 + 2\gamma_{2f}$				
Wald test $\chi^2(1)$=0	20.81	1.09	0.07	42.76
Prob> χ^2	**0.000**	0.2965	0.7971	**0.0000**
Test: $\omega + \omega_f = \theta + \theta_f$				
Wald test $\chi^2(1)$=0	2.68	0.05	0.00	
Prob> χ^2	0.1013	0.8739	0.8165	0.9724
Test: $\omega + \omega_f = \xi + \xi_f$				
Wald test $\chi^2(1)$=0	9.59	0.04	0.09	13.44
Prob> χ^2	**0.0020**	0.8448	0.7638	**0.0002**
Test: $\xi + \xi_f = \theta + \theta_f$				
Wald test $\chi^2(1)$=0	3.40	0.00	0.53	21.53
Prob> χ^2	**0.0652**	0.9896	0.4672	**0.0000**

Notes: Tests refer to estimates presented in previous table.

Table VII: Random Effects Panel Tobit Estimation: Expenditures On Assignable Clothing by Household Income Source, RLMS 1998.IV-2001.IV

	children's clothes and shoes	adult clothes and shoes
child benefits ω	0.1492	0.4587
	(0.124)	(0.354)
social transfer income, θ	0.2718**	0.1991
	(0.122)	(0.350)
private transfer income, ξ	0.1139**	0.2027**
	(0.018)	(0.050)
real hh.income, γ_1	0.1492**	0.4092 **
	(0.020)	(0.058)
real hh. income $^{2\ a}$, $\gamma_2 * 100000$	-0.1326**	-0.4553 **
	(0.034)	(0.097)
local price	no	no
σ_u	1559.417	4411.187
	(114.25)	(380.028)
σ_e	2605.149	7231.78
	(65.04)	(197.636)
ρ	0.2627	0.2622
Wald χ^2 (9)	175.26	133.63
Prob > χ^2	0.0000	0.0000
uncensored	1234	971
left censored	503	766
Number of obs.	1737	1737
Number of groups	671	671
Log likelihood	-11982.937	-10548.292

Notes: Full demographic controls, constant, and controls for income in kind also included. Standard errors in parentheses.** significant at 5% level, * significant at 10% level. Real (June 1992 ruble) household income net of child benefits.

[a]real earned hh. income^2 /100000

Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Income and expenditure measures are for the 30 days prior to the interview.

Table VIII: Tests of equality of MPC coefficients, all households. Expenditures On Assignable Clothing, RLMS 1998.IV-2001.IV

	children's clothes and shoes	adult clothes and shoes
Test: $\omega = \gamma_1 + 2\gamma_2$		
Wald test $\chi^2(1)$=0	0.000	0.02
Prob> χ^2	0.9998	0.8906
Test: $\theta = \gamma_1 + 2\gamma_2$		
Wald test $\chi^2(1)$=0	0.99	0.35
Prob> χ^2	0.3208	0.5533
Test: $\xi = \gamma_1 + 2\gamma_2$		
Wald test $\chi^2(1)$=0	1.79	7.65
Prob> χ^2	0.1809	**0.0057**
Test: $\omega = \theta$		
Wald test $\chi^2(1)$=0	0.45	0.24
Prob> χ^2	0.5034	0.6212
Test: $\omega = \xi$		
Wald test $\chi^2(1)$=0	0.08	0.52
Prob> χ^2	0.7780	0.4726
Test: $\xi = \theta$		
Wald test $\chi^2(1)$=0	1.64	0.00
Prob> χ^2	0.2002	0.9921

Notes: Tests refer to estimation results of previous table.

Table IX: Random Effects Panel Tobit Estimation: Expenditures on assignable clothes and shoes, interactions for all-female households, RLMS 1998.IV-2001.IV

	children's clothes and shoes	adult clothes and shoes
child benefits, ω	0.2231	0.4227
	(0.146)	(0.420)
child benefits*all-female, ω_f	-0.3047	-0.2790
	(0.2798)	(0.793)
social transfer income, θ	0.2317*	0.0215
	(0.129)	(0.370)
social transfers*all-female, θ_f	0.2683	0.8741
	(0.292)	(0.347)
private transfer income, ξ	0.1153**	0.2028**
	(0.017)	(0.049)
private transfers*all-female, ξ_f	-0.0833	0.6030
	(0.192)	(0.544)
real earned hh.income, γ_1	0.1498 **	0.3881**
	(0.020)	(0.059)
real earned hh.income*all-female, γ_{1f}	0.0861	0.8062*
	(0.161)	(0.490)
real earned hh. income ^{2}/100000, γ_2	-0.1332**	-0.4275**
	(0.034)	(0.097)
real earned hh. income ^{2}/100000*all-female, γ_{2f}	-0.6896	-2.8731
	(1.063)	(3.2970)
all-female household dummy	-88.0650	-214.424*
	(458.795)	(90.908)
σ_u	1571.612	4254.389
	(115.413)	(379.636)
σ_e	2597.075	7227.305
	(65.275)	(197.539)
ρ	0.2680	0.2573
	(0.268)	(0.038)
Wald $\chi^2(13)$	177.90	141.01
Prob > χ^2	0.0000	0.0000
uncensored	1234	971
left censored	503	766
Number of observations	1737	1737
Number of households	671	671
Log likelihood	-11981.643	-10544.686

Notes: Standard errors in parentheses.$**$ significant at 5% level, $*$ significant at 10% level. Real (June 1992 ruble) household income net of child benefits, social transfers, and private transfers. Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Controls for numbers of children 0-1.5, 1.5-6, 6-16, number of working-age adults, and number of pension-age adults are also included, as well as controls for any in-kind transfers from workplaces and a constant. Income and expenditure measures are for the 30 days prior to the interview.

Table X: Tests of equality of MPC coefficients, all households. Expenditures on assignable clothes and shoes, interactions for all-female households, RLMS 1998.IV-2001.IV

	children's clothes and shoes	adult clothes and shoes
Test: $\omega = \gamma_1 + 2\gamma_2$		
Wald test $\chi^2(1)$=0	0.24	0.01
Prob> χ^2	0.6214	0.9351
Test: $\theta = \gamma_1 + 2\gamma_2$		
Wald test $\chi^2(1)$=0	0.39	0.95
Prob> χ^2	0.5331	0.3292
Test: $\xi = \gamma_1 + 2\gamma_2$		
Wald test $\chi^2(1)$=0	1.65	5.97
Prob> χ^2	0.1985	0.0145
Test: $\omega = \theta$		
Wald test $\chi^2(1)$=0	0.00	0.48
Prob> χ^2	0.9660	0.4877
Test: $\omega = \xi$		
Wald test $\chi^2(1)$=0	0.53	0.27
Prob> χ^2	0.4647	0.6025
Test: $\xi = \theta$		
Wald test $\chi^2(1)$=0	0.79	0.23
Prob> χ^2	0.3749	0.6287
	All-female households	
Test: $\omega + \omega_f = \gamma_1 + \gamma_{1f} + 2\gamma_2 + 2\gamma_{2f}$		
Wald test $\chi^2(1)$=0	1.12	1.46
Prob> χ^2	0.2891	0.2274
Test: $\theta + \theta_f = \gamma_1 + \gamma_{1f} + 2\gamma_2 + 2\gamma_{2f}$		
Wald test $\chi^2(1)$=0	0.70	0.01
Prob> χ^2	0.4044	0.9165
Test: $\xi + \xi_f = \gamma_1 + \gamma_{1f} + 2\gamma_2 + 2\gamma_{2f}$		
Wald test $\chi^2(1)$=0	0.77	0.34
Prob> χ^2	0.3796	0.5579
Test: $\omega + \omega_f = \theta + \theta_f$		
Wald test $\chi^2(1)$=0	2.01	0.36
Prob> χ^2	0.1564	0.5216
Test: $\omega + \omega_f = \xi + \xi_f$		
Wald test $\chi^2(1)$=0	0.13	0.55
Prob> χ^2	0.7174	0.4589
Test: $\xi + \xi_f = \theta + \theta_f$		
Wald test $\chi^2(1)$=0	2.23	0.29
Prob> χ^2	0.1350	0.5903

Notes: Tests refer to estimation results of previous table.

Table XI: Sample means of individual fat and protein intakes, households with children under 14 in 1994, RLMS 1994.IV-1996.IV

	children		adults	
	1994	1996	1994	1996
	% calories from protein			
All	11.927	8.771	12.745	9.025
	(0.069)	(0.118)	(0.058)	(0.101)
child benefits	11.989	11.749	12.782	11.712
	(0.092)	(0.124)	(0.075)	(0.142)
benefits arrears	11.835	7.799	12.695	8.146
	(0.109)	(0.144)	(0.990)	(0.120)
t-stat	1.0557	-15.0534	-0.7445	-15.6093
P>| t |	0.2912	0.0000	0.4566	0.0000
	% calories from fat			
All	32.342	29.850	33.851	31.880
	(0.212)	(0.231)	(0.168)	(0.188)
child benefits	31.957	30.182	33.870	32.485
	(0.270)	(0.391)	(0.217)	(0.341)
benefits arrears	32.862	29.691	33.825	31.595
	(0.339)	(0.286)	(0.267)	(0.225)
t-stat	2.1099	-0.9336	-0.1341	-2.0268
P>| t |	0.0350	0.3506	0.8933	0.0428

Notes: Standard errors in parentheses. Households must contain at least one child under 14 in 1994.IV-1996.IV.

Table XII: Random and Fixed Effects Estimates: Children's Marginal Propensity to Consume Nutrients RLMS 1994.IV-1996.IV

Children aged 0-14 in 1994.IV	**Protein (% calories)**		**Fat (% calories)**	
estimator	random	fixed	random	fixed
child benefits/10, ω	0.0016*	0.0021*	-0.0003	0.0046
	(0.001)	(0.001)	(0.897)	(0.003)
social transfers/10, θ	0.0007	-0.0005	0.0074*	0.0080
	(0.659)	(0.002)	(0.004)	(0.005)
private transfers/10, ξ	0.0001	-0.0003*	0.0021**	0.0009
	(0.000)	(0.000)	(0.000)	(0.001)
earned hh.income/10, γ_1	0.0002 **	0.0002 **	0.0008 **	0.0002
	(0.000)	(0.000)	(0.000)	(0.000)
earned hh. income ^{2}/100000, γ_2	-0.0002 **	-0.0002**	-0.0009 **	-0.0003
	(0.000)	(0.000)	(0.000)	(0.000)
age at start	yes	no	yes	no
sex	yes	no	yes	no
no. children	yes	no	yes	no
no. adults	yes	no	yes	no
food price	yes	yes	yes	yes
constant	yes	yes	yes	yes
σ_u	0.8703	2.3665	3.3120	6.6471
σ_e	3.1355	3.3533	8.9684	8.9899
ρ	0.0715	0.3325	0.1201	0.3534
Model Fit:				
Wald $\chi^2(6)$	34.24	–	107.43	–
Prob > χ^2	0.0003	–	0.0000	–
F test	–	3.11	–	4.31
Prob>F	–	0.0024	–	0.0001
no. observations	5118	5223	5069	5093
no. groups	1847	1853	1874	1851
Hausman Test:				
$\chi^2(3)$		23.50		44.22
Prob> $\chi^2(3)$		0.0014		0.0000
Breusch-Pagan LM Test:				
$\chi^2(1)$	7.41	–	49.23	–
Prob> $\chi^2(1)$	0.0065	–	0.0000	–

Notes: Standard errors in parentheses.$**$ significant at 5% level, $*$ significant at 10% level. Real (June 1992) roubles. Households must contain at least one child under 14 in 1994.IV-1996.IV. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 1996.IV are included. Income and expenditure measures are for the 30 days prior to the interview. Calorie consumption on day prior to interview.
1994 individual sample weights are used.

Table XIII: Tests of the equality of coefficients: Children's Marginal Propensity to Consume Nutrients RLMS 1994.IV-1996.IV

Children aged 0-14 in 1994.IV	**Protein (% calories)**		**Fat (% calories)**	
estimator	random	fixed	random	fixed
Test: $\omega/10 = \gamma_1/10 + \gamma_2/100000$				
Wald test $\chi^2(1)$=0	2.71	–	0.20	–
Prob> χ^2	**0.098**	–	0.6554	–
F(1,3366) test $\chi^2(1)$=0	–	2.73	–	1.77
Prob> χ^2	–	**0.090**	–	0.1833
Test: $\theta/10 = \gamma_1/10 + \gamma_2/100000$	0.08			
Wald test $\chi^2(1)$=0	0.08	–	2.35	–
Prob> χ^2	0.7708	–	0.1253	–
F(1,3366) test $\chi^2(1)$=0	–	0.16	–	2.28
Prob> χ^2	–	0.6864	–	0.1315
Test: $\xi/10 = \gamma_1/10 + \gamma_2/100000$				
Wald test $\chi^2(1)$=0	0.37	–	8.33	–
Prob> χ^2	0.5405	–	**0.0039**	–
F(1,3366) test $\chi^2(1)$=0	–	6.40	–	1.11
Prob> χ^2	–	**0.0115**	–	0.2929
Test: $\xi = \theta$				
Wald test $\chi^2(1)$=0	0.12	–	1.52	–
Prob> χ^2	0.7261	–	0.2175	–
F(1,3366) test $\chi^2(1)$=0	–	0.01	–	1.88
Prob> χ^2	–	0.9109	–	0.1704
Test: $\omega = \theta$				
Wald test $\chi^2(1)$=0	0.34	–	2.39	–
Prob> χ^2	0.5609	–	0.1219	–
F(1,3366) test $\chi^2(1)$=0	–	1.38	–	0.32
Prob> χ^2	–	0.2397	–	0.5728

Notes: Tests refer to coefficient values of previous table.

Table XIV: Random and Fixed Effects Estimates: Adult's Marginal Propensity to Consume Nutrients RLMS 1994.IV-1996.IV

Adults in households with children aged 0-14 in 1994.IV	**Protein (% calories)**		**Fat (% calories)**	
estimator	random	fixed	random	fixed
child benefits/10, ω	-0.0001	0.0023*	0.0033	0.0086 **
	(0.001)	(0.001)	(0.002)	(0.003)
social transfers/10, θ	0.0010	0.0004	0.0085**	0.0112**
	(0.001)	(0.002)	(0.003)	(0.004)
private transfers/10, ξ	0.0003*	-0.0001	0.0021**	0.0000
	(0.000)	(0.000)	(0.000)	(0.000)
earned hh.income/10, γ_1	0.0004 **	0.0005 **	0.0013 **	0.0004*
	(0.000)	(0.000)	(0.000)	(0.000)
earned hh. income ^{2}/100000, γ_2	-0.0004 **	-0.0005**	-0.0012 **	-0.0003
	(0.0000)	(0.000)	(0.000)	(0.000)
age at start	yes	no	yes	no
sex	yes	no	yes	no
no. children	yes	no	yes	no
no. adults	yes	no	yes	no
food price	yes	yes	yes	yes
constant	yes	yes	yes	yes
σ_u	1.7893	2.9944	4.1425	7.4154
σ_e	3.8980	3.9821	9.1702	9.1808
ρ	0.1740	0.3612	0.1697	0.3948
Model Fit:				
Wald $\chi^2(7)$	152.89	–	160.69	–
Prob > χ^2	0.0000	–	0.0000	–
F test	–	9.82	–	5.84
Prob>F	–	0.0000	–	0.0000
no. observations	9250	9250	8730	8816
no. groups	3298	3298	3282	3288
Hausman Test:				
$\chi^2(3)$		60.55		105.46
Prob> $\chi^2(3)$		0.0000		0.0000
Breusch-Pagan LM Test:				
$\chi^2(1)$	226.19	–	225.28	–
Prob> $\chi^2(1)$	0.0000	–	0.0000	–

Notes: Robust standard errors in parentheses.$**$ significant at 5% level, $*$ significant at 10% level. Real (June 1992) roubles. Households must contain at least one child under 14 in 1994.IV-1996.IV. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 1996.IV are included. Income and expenditure measures are for the 30 days prior to the interview. Calorie consumption on day prior to interview. Household level clustering implemented for standard errors.
1994 individual sample weights are used.

Table XV: Tests of the equality of coefficients: Adult's Marginal Propensity to Consume Nutrients RLMS 1994.IV-1996.IV

Adults in households with children aged 0-14 in 1994.IV	**Protein (% calories)**		**Fat (% calories)**	
estimator	random	fixed	random	fixed
Test: $\omega/10 = \gamma_1/10 + \gamma_2/100000$				
Wald test $\chi^2(1)$=0	2.69	–	0.77	–
Prob> χ^2	**0.1000**	–	0.3815	–
F(1,5946) test $\chi^2(1)$=0	–	8.25	–	7.81
Prob> χ^2	–	**0.004**	–	**0.0052**
Test: $\theta/10 = \gamma_1/10 + \gamma_2/100000$				
Wald test $\chi^2(1)$=0	0.16	–	4.56	–
Prob> χ^2	0.6872	–	**0.0327**	–
F(1,3366) test $\chi^2(1)$=0	–	0.00	–	7.26
Prob> χ^2	–	0.9908	–	**0.0071**
Test: $\xi/10 = \gamma_1/10 + \gamma_2/100000$				
Wald test $\chi^2(1)$=0	1.38	–	5.98	–
Prob> χ^2	0.2401	–	**0.0145**	–
F(1,3366) test $\chi^2(1)$=0	–	10.43	–	0.43
Prob> χ^2	–	**0.0012**	–	0.5125
Test: $\xi = \theta$				
Wald test $\chi^2(1)$=0	0.27	–	3.53	–
Prob> χ^2	0.6041	–	**0.0603**	–
F(1,3366) test $\chi^2(1)$=0	–	0.13	–	7.61
Prob> χ^2	–	0.7145	–	**0.0058**
Test: $\omega = \theta$				
Wald test $\chi^2(1)$=0	0.39	–	1.54	–
Prob> χ^2	0.5310	–	0.2147	–
F(1,3366) test $\chi^2(1)$=0	–	0.68	–	0.27
Prob> χ^2	–	0.4316	–	0.6056

Notes: Tests refer to estimation results of previous table.

Table XVI: Random Effects Panel Logit Estimation: Probability of obtaining child benefits, RLMS 1994.IV-2001.IV

Specification	I	II	III
real hh.income/10	0.0003**	0.0000	0.0000
	(0.0000)	(0.0000)	(0.0000)
real hh. income ^{2}/1000000	0.0005**	0.0000	0.0000
	(0.0000)	(0.0000)	(0.0000)
year dummies (5)	no	yes	yes
oblast dummies (7)	no	no	yes
σ_u	0.5926	0.7171	0.5381
	(0.0657)	(0.067)	(0.0743)
ρ	0.0965	0.135	0.081
Wald χ^2	68.41	346.00	409.19
Prob > χ^2	0.0000	0.0000	0.0000
Number of obs	3925	3925	3925
Number of groups	811	811	811
Log likelihood	-2573.272	-2398.209	-2346.213
LR Tests:		II nests I	III nests II
LR χ^2	–	343.72	105.88
prob> χ^2	–	0.000	0.0000

Notes: Standard errors in parentheses.$**$ significant at 5% level, $*$ significant at 10% level. Real (June 1992 ruble) household income net of child benefits. Controls for the number of children in the age groups 1-1.5, 1.5-6, and 6-16, the number of working-age adults, and the number of pension-age household members are included in all specifications. Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Income and expenditure measures are for the 30 days prior to the interview.

Appendix A: Benefits receipt and eligibility

In this appendix, I analyse the propensity to receive child benefits over the sample period. The results of a random effects logit model estimation are presented in Columns (I) through (III) of Table XVI. Column (I) presents the results which control for only household income[11] and demographics. Without controls for changes over time in the propensity to receive child benefits, it appears that there is a statistically significant positive relationship between income and benefits receipt propensities. However, this relationship is no longer statistically (nor economically) significant when time dummies (Column (II)) and regional dummies (Column III) are included. Likelihood ratio tests show that specification III is preferable to either I or II. Variation in benefits receipt appears to be mainly a function of regional budgets and of the overall macroeconomic situation in a year. The reasons why there appears to be a positive relation between benefits receipt propensities and household income in Column (I) are: (*i.*) incomes fell over the 1994-2001 period and (*ii.*) regional budget crises are associated with poor regional economic conditions.

[11]Note that here 'household income refers to *all* other household income net of child benefits.

Table XVII: Random Effects Panel Logit Estimation: Probability of household being eligible for child benefits, RLMS 1994.IV-2001.IV

Specification	I	II	III
real hh.income/10	-0.0002**	-0.0002**	-0.0002**
	(0.0000)	(0.0000)	(0.0000)
real hh. income ^{2}/1000000	0.0001	0.0001	0.0001
	(0.0001)	(0.0003)	(0.0003)
year dummies (5)	no	yes	yes
oblast dummies (7)	no	no	yes
constant	yes	yes	yes
σ_u	1.7596	1.8942	1.761
	(0.1264)	(0.136)	(0.0.138)
ρ	0.4848	0.522	0.485
Wald χ^2	74.10	161.32	188.07
Prob $> \chi^2$	0.0000	0.0000	0.0000
Number of obs	3925	3925	3925
Number of groups	811	811	811
Log likelihood	-1192.406	-1139.271	-1117.0386
LR Tests:		II nests I	III nests II
LR χ^2	–	102.79	43.49
prob$> \chi^2$	–	0.000	0.000

Notes: Self-reporting of child benefits eligibility. Standard errors in parentheses.$**$ significant at 5% level, $*$ significant at 10% level. Real (June 1992 ruble) household income net of child benefits. Controls for the number of children in the age groups 1-1.5, 1.5-6, and 6-16, the number of working-age adults, and the number of pension-age household members are included in all specifications. Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Income measures are for the 30 days prior to the interview.

Next, I turn to the multivariate analysis of benefits eligibility in the sample. The results of random effects logit estimation of the probability of being eligible for child benefits in a year are presented in Table XVII. In Column (I) I report specifications which control only for income and household demographics. In Column (II) I introduce year dummies, and in Column (III) I include year and time dummies. In all three specifications the probability of a household being eligible[12] is higher at lower income levels. While likelihood ratio tests suggest that specifications which control for time and regional variation are preferable, these factors do not tell the whole story. As would be expected under targeting, poorer households are more likely to be eligible for benefits.

A caveat to the results of this paper is that the RLMS does not contain information on whether non-receipt of benefits is due to lack of application, or to arrears. A question on *perceived* eligibility of respondents is included, but in an era of quickly-changing rules, individuals may not accurately know this information. Because I use panel estimators, I am able to control for unobserved preferences of households, which can be considered to include the latent variable governing the propensity of households to claim.

In Table XVIII below I examine the propensity to receive social transfers across households. Households who receive more social transfers have significantly lower earned incomes, after controlling for household composition, regional and time effects. Households with lower labour earnings have higher transfers. However, because some social transfers (pensions) are large, it is possible that labour supply effects of transfers are driving the observed relationship.

[12] Derived from self-reports.

Table XVIII: Random Effects Panel Logit Estimation: Probability of obtaining social transfers, RLMS 1994.IV-2001.IV

Specification	I	II	III
real hh.income/100	-0.0038**	-0.0053**	-0.0054**
	(0.001)	(0.001)	(0.001)
real hh. income ^{2}/10000	0.0355	0.0521	0.0541*
	(0.035)	(0.034)	(0.033)
year dummies (5)	no	yes	yes
oblast dummies(7)	no	no	yes
σ_u	1.7286	1.8676	1.8220
	(0.113)	(0.121)	(0.120)
ρ	0.4759	0.5146	0.5022
Wald χ^2	456.77	459.34	465.23
Prob $> \chi^2$	0.0000	0.0000	0.0000
Number of obs	3886	3886	3886
Number of groups	804	804	804
Log likelihood	-1466.712	-1410.749	-1398.221
LR Tests:		II nests I	III nests II
LR χ^2		111.92	25.06
prob$>\chi^2$	–	0.000	0.001

Notes: Standard errors in parentheses.** significant at 5% level, * significant at 10% level. Real (June 1992 ruble) household income net of all benefits, private and public. Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Income and expenditure measures are for the 30 days prior to the interview. Controls for numbers of children 0-1.5, 1.5-6, 6-16, number of working-age adults, and number of pension-age adults are also included, as well as controls for any in-kind transfers from workplaces and a constant.

Appendix B: Construction of Price Variables

Throughout the 1990s large differences in the prices of goods across regions persisted. I construct a vector of food and alcohol/tobacco prices for each year in order to control for these differences in the estimation of the incomplete demand system.

The information I use in the construction of the food price vector is the following:

(*i.*) Community-level information on the prices of different foodstuffs for each year.
(*ii.*) Information on the expenditure on each food budget line across RLMS households in each year.

I calculate what the median-income RLMS household would have spent, given the same allocation of the food budget, in each of the RLMS sites within a year. I normalise the index so that the mean value across sites is 100 (although the mean value across households will not be.)

For alcohol/tobacco prices I undertake a similar procedure, using the alcohol/tobacco basket of the median income RLMS household in the year to derive the alcohol price vector.

This demand system estimated with such a price vector is incomplete. I do not have information on the prices of clothing, family services, or supplies, and am therefore unable to estimate the complete demand system and test cross-equation restrictions.

Appendix C: Private transfers and luxury goods expenditure from social transfers

In this appendix I attempt to shed light on how child benefits and other social transfers are related to luxury consumption spending and to private transfers (gifts) to other households.

In Table XIX I present results of random effects tobit estimation of the MPC from social transfers for four budget lines which might be considered 'longer term' or at least reflecting that basic needs of households are satisfied. In Column (I) I present the MPC from child benefits and other income on financial transfers sent to friends and relatives. As expected, these transfers are rising in household income. However, there is no statistically significant relation between child benefits money and these private transfers to other households. However, other social transfers (mainly pension income) have MPCs about four times as high as earned household income. Thus, the use of child benefits money diverges from other social transfer income when considering assistance (gifts) to non-resident friends and family. It is also worth noting that private transfers are significant in determining assistance to friends and family but essentially in the same way as earned household income.

Column (II) presents results for real expenditures on luxury items. Again the Wald test does not reject equality of coefficients for any of the household income types. However, it is of interest that only the coefficients on private transfer income and earned income are statistically significant, and that they are of essentially the same magnitude. The lack of statistical significance of the coefficient ω, reflecting child benefits money used on these budget lines, suggests that this child benefits income is not used in expenditures on either luxury spending or gift transfers to other households. Unfortunately, small sample sizes for all-female households households prevent clarification of the extent to which these effects are due to systematic gender differences in preferences.

Table XIX: Pooled Tobit Estimation: Private transfers and luxury expenditure, all household types, RLMS 1994.IV-2001.IV

	private assistance	luxury spending[a]
child benefits, ω	-0.1165	0.1573
	(0.234)	(0.493)
social transfers, θ	0.4412**	0.0475
	(0.143)	(0.384)
private transfers, ξ	0.1120**	0.3667**
	(0.026)	(0.052)
earned hh.income, γ_1	0.1240**	0.3435**
	(0.015)	(0.0643)
earned hh. income ^{2a}, $\gamma_2 * 100000$	-0.0118**	-0.2020**
	(0.002)	(0.102)
log likelihood	-8798.7345	-4703.9762
LR $\chi^2(13)$	109.20	107.53
Prob $> \chi^2$	0.0000	0.0000
	Tests of equality of coefficients	
Test: $\omega = \gamma_1 + 2\gamma_2$		
F- test(1,3871)	1.04	0.14
Prob$>F$	0.3081	0.7092
Test: $\theta = \gamma_1 + 2\gamma_2$		
F- test(1,3871)	4.62	0.58
Prob$>F$	**0.0317**	0.4472
Test: $\xi = \gamma_1 + 2\gamma_2$		
F- test(1,3871)	0.15	0.09
Prob$>F$	0.6949	0.7669
Test: $\omega = \theta$		
F- test(1,3871)	3.72	0.03
Prob$>F$	**0.0539**	0.8654
Test: $\omega = \xi$		
F- test(1,3871)	0.94	0.18
Prob$>F$	0.3318	0.6723
Test: $\theta = \xi$		
F- test(1,3871)	4.87	0.68
Prob$>F$	**0.0274**	0.4093

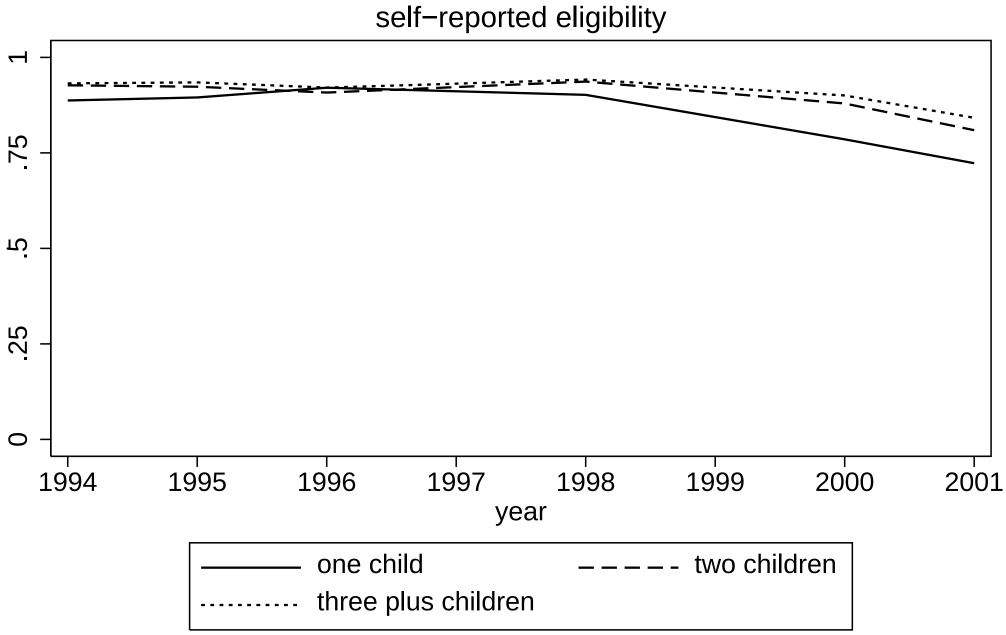
Notes: Demographic and geographical controls are included as in previous tables. Pooled tobit estimation is reported here because the ρ and σ_u values were found to be statistically insignificant in random effects panel estimation. Standard errors in parentheses.$**$ significant at 5% level, $*$ significant at 10% level. Real (June 1992 ruble) household income net of child benefits, other social transfers, and private transfers.

[a]real earned hh. income^2 /100000

Households must contain at least one child under 16 in all periods. Only households whose composition did not change between 1994.IV and either (*i.*) attrition, or (*ii.*) 2001.IV are included. Income and expenditure measures (with the exception of Column (II)) are for the 30 days prior to the interview.

[a] Data collected refers to spending on luxury goods (ie. cars, consumer durables) in 3 months prior to interview. The measure reported is divided by 3, giving a monthly measure.

Figure I: Child benefits eligibility by year and number of children



Source:RLMS 1994.IV−2001.IV

Figure II: Fraction of households receiving child benefits by year and number of children

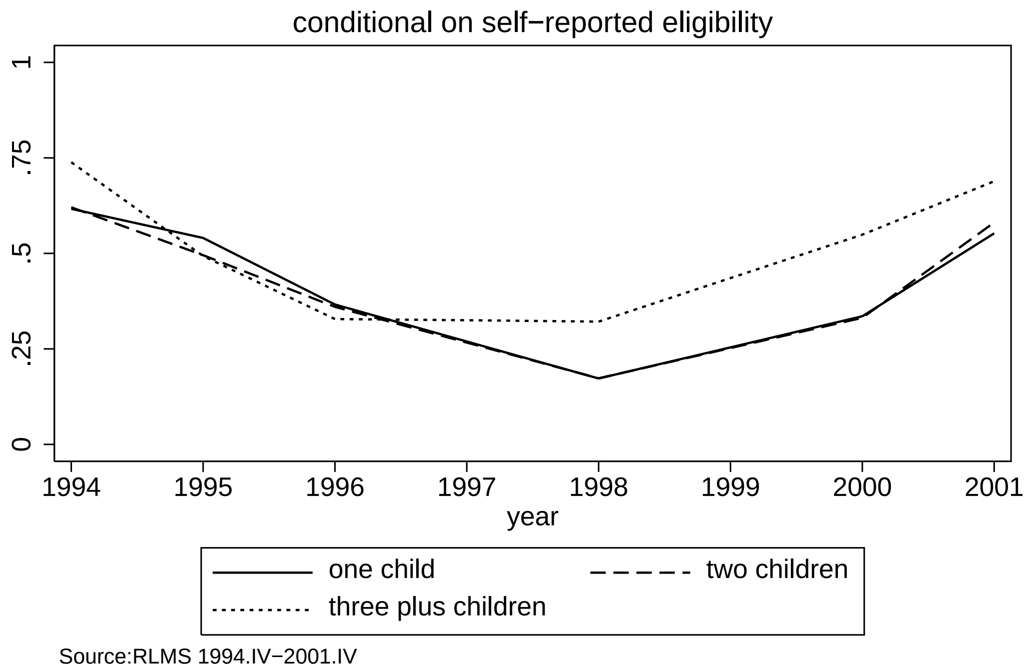
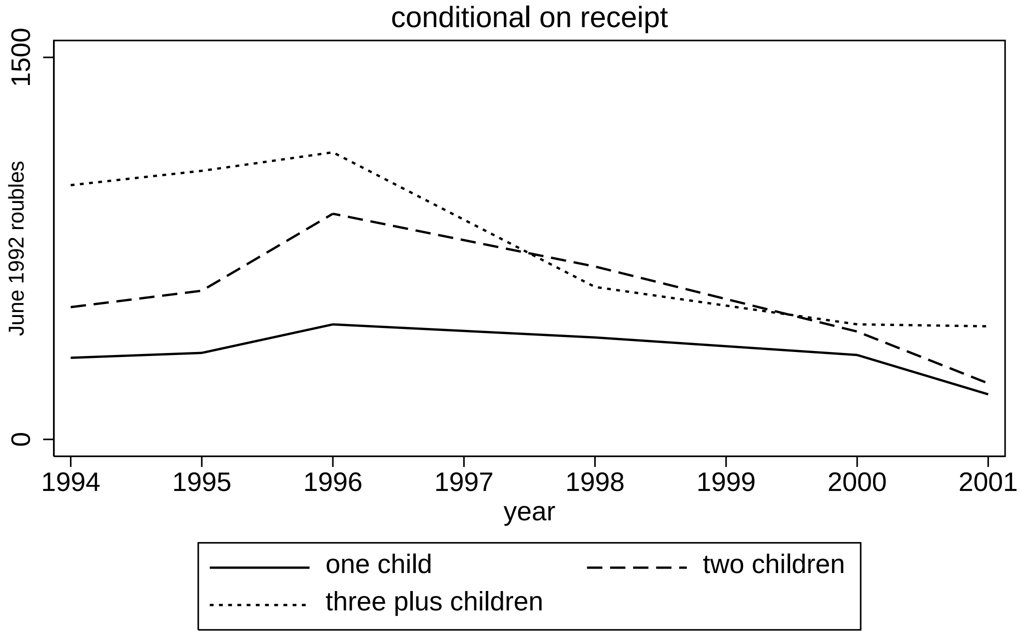


Figure III: Mean levels of received child benefits by year and number of children



Source:RLMS 1994.IV−2001.IV