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Adjustments in Demand During Lithuania's Economic Transition

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ABSTRACT

Significant economic changes have occurred in Lithuania since 1992. Major market reforms were introduced through phased liberalization of producer and consumer prices. Panel data from a Lithuanian Household Budget Survey conducted during 1992-94 were used to estimate demand parameters for aggregated commodity groups. The evidence from the household survey shows the relative stability of demand parameters. Estimated price and expenditure elasticities are consistent with levels of low income countries. Households have been relatively responsive to changes in prices and total expenditure in their purchases during the period of transition.

ADJUSTMENTS IN DEMAND DURING LITHUANIA'S ECONOMIC TRANSITION

1. Introduction

The collapse of the Soviet system and the indpendence of the smaller republics in the early 1990s led to a series of economic reforms associated with the transition to more market-oriented economies in the former Soviet Union. Many of the countries adopted economic reforms to liberalize prices and withdraw government subsidies. As a result, these societies experienced rapid price and income changes, accompanied by other changes associated with the ongoing market reforms. Although the transition policies have had outcomes specific to each country, for many the results imposed severe hardship on a significant portion of the society (see, for example, Cornia).

Major economic reforms in Lithuania have been underway since 1990, when Lithuania regained independence. The early adoption of economic reforms makes evidence from this country useful to on-going evaluation of reforms both for Lithuania as well as for other emerging market economies. With Lithuanian independence came a period of economic contraction, rapidly rising prices, and other economic and social adjustments (OECD 1996). Initially, increases in wages and social benefits partly compensated for the increased prices. Following very high inflation rates in 1991 and 1992, inflation rates moderated to annual levels of 45 percent in 1994 and 36 percent in 1995. Real wages in the public sector fell sharply through 1993, and since then have improved (OECD).

Knowledge about the structure of consumer demand provides a sound basis both for policy formulation and evaluation. For example, basic demand parameters give information needed for an effective design and targeting of social assistance and other programs, as well as for evaluating the impact of economic reforms on households and general well-being. These are key issues for countries undergoing signficiant economic adjustments.

The main objective of this research was to evaluate consumer demand behavior in Lithuania during the economic transition period and to assess the effects of the economic changes on Lithuanian households. We estimate a preference consistent demand system and use monthly household budget survey data for the period July 1992 to December 1994 to estimate the demand parameters. The period of study is one in which households experienced large and fundamental changes in income, absolute and relative prices, level of state support, and stability of employment and income. Because differing distributional effects are likely to be present, it is important that we account for differences across households over time in the estimation process.

2. Data

The Lithuanian Household Budget Survey (HBS) provided information on the household level effects and behavior during the initial stages of the economic transition. The HBS, introduced in 1992, was designed to be nationally representative of Lithuanian households (Šniukštiné, Vanagaite and Binkauskiené 1996) and replaced the traditional Soviet Family Budget Survey (Atkinson and Micklewright 1992). The design included monthly surveys of households; households were included in the sample for 13 months. This allowed for a sample rotation with 1 of every 13 households replaced each month. The stratified survey design included strata for urban (Vilnius and other urban) and rural areas and income levels in urban areas. The income levels were set by ad hoc intervals in 1992 and 1993, and by deciles in 1994 (Cornelius 1995). Although the survey marked a significant improvement over the earlier survey, in practice the implementation suffered from problems associated with the sample not being fully random and from nonresponse problems associated with households not completing the full 13-month period of inclusion in the survey design. However review of the data with other, aggregated, consumption data did indicate the data to be a good measure of consumption trends, and representative of the national population. In total, about 1500 households were surveyed each month.[1]

Panel Data Construction

The 1992-94 HBS provided a representative sample of households each month. However, because of the fact that the procedure for household replacement (replacement of a household dropping out of the survey by another household of similar type) was not tightly controlled, the survey design did not allow for uniquely identifying households from month to month. Therefore, it was not possible to fully exploit the panel structure of the data at the household level. For this analysis, monthly household data for the period July 1992 through December 1994 were used to create a panel data for 40 *representative* household groups, defined by location (urban/rural), household size, and level of total (per capita) expenditures.

The construction of the panel was done in the following way. First, households were classified into five quintiles on the basis of per capita total household expenditures. Second, within each per capita expenditures quintile, households were classified into rural and urban households. The two levels of classification (quintiles and urban/rural) yielded ten groups of households (5×2). Each group of households was then further classified according to household size. The classification took account of the distribution of household sizes in the whole sample, and yielded a reasonably balanced distribution of observations in different cells in the three-level classification. Once the classes were selected, the means of different variables in each of the cells were used as representative values of the corresponding variables in the data set. The above procedure yielded 40 observations for each of the 30 months of data, and, therefore, 1,200 observations in total.

Data on Expenditures

The survey instrument was used to collect data on household demographics, expenditures, and income. Because the reference period for expenditures varied by type of expenditure, expenditures were aggregated on a monthly basis. Table 1 provides descriptive information on the survey data for the months July 1992 through December 1994.

For estimation purposes, expenditures on various goods and services were aggregated into the following six categories: food (including food away from home and the value of home-produced food and gifts); housing (including rent, building or purchase of houses, utilities such as heating, electricity, water, gas, etc.); household furnishings (including furniture, drapes, durable consumer goods, cleaning of households, and maintenance of household appliances, etc.); clothing (including dresses, fabrics used for making dresses, personal accessories, jewelry, cleaning, repair and maintenance of items included in this category); entertainment and education; and other nonfood consumer items (including all other consumption expenditures not included in other categories). This classification was maintained in accordance with the published consumer price indices of various subgroups of consumption expenditures as published by the Department of Statistics of the Government of Lithuania.

The category *other nonfood expenditures* is an aggregation of personal care, health care, transportation and communication, and other expenditures in the classification of the price indices published by the Lithuanian government. We aggregated these four categories into *other nonfood expenditures* by using the average expenditure share of each category of expenditure as the relevant weight.

Prices

The price indices for food, housing, household furnishings, clothing, entertainment and education, and transportation and communication (included in other) came from those published by the Department of Statistics. All these indices are fixed weight indices that use May 1992 as the base year. *Total household expenditures* is used to measure the purchasing power of the household. This is the sum of expenditures on food (including valued nonpurchased foods), housing, household furnishings, clothing, entertainment and education, and other nonfood expenditure. The shares of the different expenditure groups in total household expenditures are shown in Table 1.

Nonpurchased Foods

The survey provided information on the quantity and value (monetary expenditure) of the food items purchased by each household. It also provided data on the quantity of food consumed that was not purchased (such food came from home production, gifts, free food, etc.). To properly quantify the total food consumption, it was necessary to evaluate these quantities of food (nonpurchased food) at the market price. In the absence of information on prices of individual items, we divided the expenditure on the purchased quantity of each item by the corresponding quantity to obtain the unit-value. This unit value was then used as the price of the item in question in order to assign values to the nonpurchased food consumed by each household. Separate unit values were calculated for rural and urban households to account for the potential price differences encountered by the rural and urban households. Thus, the nonpurchased quantities were evaluated separately for the rural and urban households by using the appropriate unit value of each commodity. Nonpurchased food represented about 30 percent of total expenditures, with a relatively higher share for rural households, and lower for urban.

3. Model and Empirical Specification

The approach used in the specification and estimation of the demand model takes advantage of the panel structure of the data to account for variation across households and over time. Because of the limitations of the data on price indices, estimation of the full error components model was not possible;[2] however, we estimated the fixed effect model (otherwise known as Least Squares Dummy Variable Model, LSDVM). Within this framework, the effects of cross-sectional variation and time-specific effects are captured by allowing the intercept terms of the demand equations to vary across cross-section units and across time.

Formally, the model we estimated can be described as follows[3]. Let us assume that the system of equations is represented by

$$y_{it} = \bar{b}_0 + m_i + l_t + \sum_{k=1}^{K+1} b_k x_{kit} + e_{it},\ \text{i = 1, 2, …, N, t = 1, 2, …, T} \qquad [1]$$

where y_{it} is the dependent variable and x_{kit} are the K+1 explanatory variables (K goods and the total expenditure). In this formulation, $\bar{b}_0$ is the average intercept, μ_i represents the difference of this mean value $\bar{b}_0$ from the intercept term corresponding to the i[th] household, and λ_t represents the difference of the mean intercept $\bar{b}_0$ from that for the t[th] time period. At any given time period, the parameter μ_i represents the influence of the variables that vary across the cross-section of households, but remain constant over time while the parameter λ_t represents the influence of those factors that are common to all households and change over time. The cross-section effect, μ_i, and the time effect, λ_t, are assumed to be fixed. Here N denotes the number of cross-sectional units, and T denotes the number of time periods. We assumed that the vector of disturbances corresponding to the i[th] cross-sectional unit, $\mathbf{e_i}$, has the property $E[\mathbf{e_i}] = \mathbf{0}$, $E[e_i e_i'] = \sigma^2 \mathbf{I}_T$, and $E[e_i e_j'] = \mathbf{0}$ for $i \neq j$.

When parameters μ_i and λ_t are treated as fixed, as has been assumed in this study, one of the μ_i and one of the λ_t is redundant and we need to impose the restrictions $\Sigma_i \mu_i = 0$ and $\Sigma_t \lambda_t = 0$. The model can be estimated using the Least Squares method under the restrictions on μ_i and λ_t. Instead of estimating equation [1] directly with (N-1) cross-section dummies, and (T-1) time-specific dummy variables included, one can account for the μ_i and λ_t effects by transforming the model according to the following procedure. First, transform the dependent as well as each of the explanatory variables (y and the x's) as:

$$\tilde{z}_{kit} = z_{kit} - \bar{z}_{ki.} - \bar{z}_{k.t} + \bar{z}_{k..}\ , \qquad [2]$$

where $\bar{z}_{ki.} = \frac{\sum_{t=1}^{T} z_{kit}}{T}$, $\bar{z}_{k.t} = \frac{\sum_{i=1}^{N} z_{kit}}{N}$, and $\bar{z}_{k..} \frac{\sum_{i=1}^{N}\sum_{t=1}^{T} z_{kit}}{NT}$. Next, by using the transformed variables from [2], the new model can be written as

$$\tilde{y}_{it} = \sum_{k=1}^{K+1} b_k\ \tilde{X}_{kit} + \tilde{e}_{it},\ i = 1,2,\ldots,N,\ t = 1,2,\ldots,T \qquad [3]$$

Equations [1] and [3] share the same slope coefficients (β's) although [3] does not contain any dummy variable. Hence, one can apply the OLS to equation [3] in order to obtain estimates of the slope coefficients, β's. Once these coefficients are estimated, the $\bar{\beta}_0$, μ_i and λ_t coefficients in equation [1] can be obtained as follows

$$\bar{\beta}_0 = \bar{y}_{..} - \sum_{k=1}^{K+1} \hat{\beta}_k \bar{x}_{k..} \qquad [4a]$$

$$\hat{\mu}_i = \left(\bar{y}_{i.} - \bar{y}_{..}\right) - \sum_{k=1}^{K+1} \hat{\beta}_k \left(\bar{x}_{ki.} - \bar{x}_{k..}\right) \qquad [4b]$$

$$\hat{\lambda}_t = \left(\bar{y}_{.t} - \bar{y}_{..}\right) - \sum_{k=1}^{K+1} \hat{\beta}_k \left(\bar{x}_{k.t} - \bar{x}_{k..}\right) \qquad [4c]$$

To estimate the demand system parameters within the LSDV framework, we used the linear approximate version of the almost ideal demand system (AIDS) of Deaton and Muellbauer (1980). Although AIDS is intrinsically nonlinear in its parameters, the linear approximation version of the model (LA/AIDS) that uses Stone’s (expenditure) share weighted price index to simplify the estimation process has been a common practice in empirical studies. Apart from its aggregation properties that allow interpretation of demand parameters estimated from household data as equivalent to those estimated from aggregate data, the AIDS model is popular because of the availability of the linear approximate version that is linear in its parameters. Buse (1994) reviewed a number of studies that use the LA/AIDS model.

The AIDS model of Deaton and Muellbauer (1980) is derived from an expenditure function and can be expressed as

$$w_k = \alpha_k + \sum_{j=1}^{K} \gamma_{kj} \ln\left(p_j\right) + \beta_k \ln\left(\frac{x}{P}\right) + \eta_k \ , \qquad [5]$$

where w_k is the expenditure share of the k^{th} good, p_j is the price of the j^{th} good, x is the total (nominal) expenditure, η_k is the error term, and ln(P) is the general price index, defined by

$$\ln\left(P\right) = \alpha_0 + \sum_{j=1}^{K} \alpha_j \ln\left(p_j\right) + \frac{1}{2} \sum_{j=1}^{K} \sum_{k=1}^{K} \gamma_{jk} \ln\left(p_j\right) \ln\left(p_k\right) \ . \qquad [6]$$

To satisfy the properties of homogeneity, adding-up, and Slutzky symmetry, the parameters of equation [5] are constrained by $\Sigma_k \alpha_k = 1$, $\Sigma_k \gamma_{kj} = \Sigma_j \gamma_{kj} = 0$, $\Sigma_k \beta_k = 0$, and $\gamma_{kj} = \gamma_{jk}$.

The price index above, ln(P), is nonlinear in parameters. The linear approximation of the price index that is used here is Stone’s price index, defined by

$$P = \prod_{j=1}^{K} \left(p_j\right)^{w_j} \ . \qquad [7]$$

Expressed in logarithmic terms, the Stone's price index can be written as

$$\ln(P) = \sum_{j=1}^{K} w_j \ln(p_j) \quad . \qquad [8]$$

Substituting [8] in [6], and denoting *real expenditure*, $\left(\frac{x_t}{P_t}\right)$, by y_t, the LA/AIDS model becomes

$$w_k = \alpha_k + \sum_{j=1}^{K} \gamma_{kj} \ln(p_j) + \beta_k \ln y_t + u_k \quad . \qquad [9]$$

The demand system given by [9] is linear in parameters in α_k, β_k, and γ_{kj} and, therefore, simple to estimate.

The LA/AIDS formulation, such as that given by equation [9] is not derived from any well-defined system of preferences, and is only an approximation to the (integrable) nonlinear AIDS model. From that perspective, it is imperative that the LA/AIDS model (for that matter, any approximation of the nonlinear AIDS model) retain good approximation properties; however, as has been discussed by Moschini (1995), Stone's price index is not independent of the choice of any arbitrary unit of measurement for prices. Consequently, the estimated demand parameters from the LA/AIDS model that use this index may contain undesirable properties. To avoid the potential problems associated with the choice of units of measurement for prices, we follow Moschini's (1995) suggestion to define the price indices of each commodity group in units of the mean of the price series (i.e., $p_j^* = (p_j / \mu_{p_j})$, where p_j is the price index of commodity group j, and μ_{p_j} is the mean of p_j.

The LA/AIDS model, as specified in equation[9], does not account for cross-sectional individual effects or time-specific effects. In order to capture these effects, the demand system model specified in equation [9] is augmented to incorporate the individual and time-specific effects. It is assumed that the effects of cross-section or household-specific variables as well as those associated with time-specific variables are reflected in variations in the intercept terms of the demand equations. Under this assumption, the demand for k^{th} commodity group is specified as

$$w_{kit} = \bar{\alpha}_k + \mu_i + \lambda_t + \sum_{j=1}^{K} \bar{\gamma}_{kj} (\ln p_{jt}) + \bar{\beta}_k \ln(y_t) + \bar{u}_{kit} \qquad [10]$$

where w_{kit} is the expenditure share of k^{th} commodity group of household i specific to time period t, μ_i is the effect of household characteristic of ith household, λ_t is the effect of factors specific to time t, and $\bar{u}_{kit}$ is the error term which is assumed to be a random variable with zero mean and constant variance (σ^2). The parameters of the system satisfy the usual adding-up, homogeneity, and Slutzky symmetry restrictions.

The demand system equations as specified by equation [10] can be estimated directly with cross-equation restrictions imposed. An alternative but equivalent method that can be used to estimate the above model is to use the transformation defined by equation [2] and reformulate the model to eliminate the cross-sectional household and time effects. Each variable in the previous equation (i.e., $w_k,\ \ln(p_{jt}),\ and\ \ln(\frac{\mathrm{X_{it}}}{\mathrm{P_t}})$) can be transformed as defined in[2]:

$$\tilde{z}_{kit} = z_{kit} - \bar{z}_{ki.} - \bar{z}_{k.t} + \bar{z}_{k..}, \qquad [11]$$

where $\bar{z}_{ki.}$, $\bar{z}_{k.t}$, and $\bar{z}_{k..}$ are as defined earlier. Now, by using the transformed variables, the slope coefficients of the LA/AIDS model can be estimated from the following equation

$$\tilde{w}_{kit} = \sum_{j=1}^{K} \bar{g}_{kj} \ln(\tilde{p}_{jt}) + \bar{b}_k \ln(\tilde{y}_{it}) + \tilde{u}_{kit} \qquad [12]$$

where the tilde symbol on the variables reflects the fact that they have been transformed according to equation [11]. The slope coefficients of the demand model as specified in equation [10] are the same as those in equation [12].

Model Estimation

The demand system model [equation 12] is estimated with the data set described above. Once the slope coefficients are estimated, the intercept terms can be recovered using the method described earlier (i.e., using formulae given by equations 4a - 4c).

To comply with the homogeneity restriction of the AIDS model, one equation is deleted in the estimation process. The Slutzky symmetry restriction is imposed as a maintained hypothesis. Imposing these restrictions on equation [12], the demand model becomes as follows

$$\tilde{w}_{kit} = \sum_{j=1}^{5} \bar{g}_{kj} \ln(\tilde{p}_{jt}) + [1 - \sum_{j=1}^{5} \bar{g}_{kj} \ln(\tilde{p}_{jt})] \ln(\tilde{p}_{6t}) + \bar{b}_k \ln(\tilde{y}_t) + \tilde{e}_{kit}, \qquad [13]$$

where the notations are as defined earlier, $\tilde{\varepsilon}_{\mathrm{kit}}$ is the error term corresponding to the equation for the k^{th} commodity group, and $\bar{\gamma}_{\mathrm{kj}} = \bar{\gamma}_{\mathrm{jk}}$.

Since the observations of the panel data used to estimate the demand system are group averages, estimation of equation [13] directly using the group averages would lead to problems of heteroscedasticity unless all group sizes are equal. In order to correct for heterscedasticity, each of the variables in the data set is transformed as

$$z_g^* = \sqrt{n_g} \, . \, z_g \qquad [14]$$

where the subscript g refers to the group g, and n_g is the number of households in group g. The transformed variables are then used to estimate the coefficients of the demand model.

4. Empirical Results

The demand system specified by equation [13] was estimated for six commodity groups (k=1, 2, …, 6); k = 1 is food expenditure, k = 2 is clothing, k = 3 is housing expenses, k = 4 is house furnishings, k = 5 is entertainment and education, and k = 6 stands for *other nonfood commodities and services*. This last group was the omitted group in the estimation of the system.

The model was estimated by the Iterated Seemingly Unrelated Regression (ITSUR**)** procedure by using SAS (1995). Visual inspection of the residuals as well as statistical tests did not suggest the presence of either heteroskedasticity or serial correlation[4].

The estimated coefficients of the demand system along with their t-ratios are presented in Table 2. Other coefficients of the system, including those of the deleted equation, can be recovered from the restrictions imposed in the estimation stage. As shown in Table 2, most of the coefficients estimated are statistically significant at conventional significance level.

The elasticities of demand (price and expenditure elasticities) were estimated by the methodology suggested by Green and Alston (1990, 1991). The Marshallian (uncompensated) elasticities are calculated as:

$$\text{Price elasticity} = \eta_{kj} = \frac{\partial \ln q_k}{\partial \ln p_j} = -\delta_{kj} + \frac{\partial \ln w_k}{\partial \ln p_j} = -\delta_{kj} + \frac{1}{w_k}\frac{\partial w_k}{\partial \ln p_j} \ ,$$

where δ_{kj} is the Kronecker Delta with value equal to one when k = j and equal to zero otherwise[5]. For the LA/AIDS model, the formulae for deriving the (uncompensated) price and expenditure elasticities are

$$\eta_{kj} = -\delta_{kj} + \frac{\gamma_{kj}}{w_k} - \frac{\beta_k}{w_k}\left(w_j + \sum_{l=1}^{K} w_l \ln p_l\right)\left(1 + \sum_{l=1}^{K} \beta_l \ln p_l\right)^{-1} \qquad [16]$$

and the expenditure elasticity, η_E, is

$$\eta_E = 1 + \frac{\beta_k}{w_k}\left(1 - \sum_{j=1}^{K} w_j \ln p_j\right) . \qquad [17]$$

We present the estimated uncompensated (Marshallian) elasticities in Table 3. Based on the model specification, these elasticities should be interpreted as long-run elasticities in which the estimated model has incorporated short-run dynamics in the adjustment of consumption expenditures to price and income changes.

The elasticities seem to be reasonable: own price elasticities are negative and many of the cross-price elasticities are positive. Some cross-price elasticities are negative and relatively large. Except for the category *other expenses*, the expenditure elasticities are near unity, with food being less than one (0.81). The own price elasticities range from -0.75 (food) and -0.57 (entertainment and education) to -2.91 (household furnishings). These values suggest the effects of the restrictive economic conditions on all households during this period.

Although these estimates appear reasonable based on other studies of demand, it is difficult to compare these estimated expenditure elasticities to other studies for two reasons. First, few estimates of these parameters for the former centrally planned countries are available in the literature. Second, the period under consideration is one in which there were large increases in prices and declines in real income along with socioeconomic dislocation of a significant share of the population. The relative effect of the price changes on the estimated model parameters is addressed below where tests of model specification and parameter stability are performed.

Tests of Model Specification and Parameter Stability

The empirical results presented in Tables 2 and 3 are based on the model specified by equation [13], for which it is assumed that cross-section household units and time variation have an important influence on household expenditure patterns. One can test for the significance of household effects (represented by μ_i) and time effects (reflected by λ_t) by testing the joint null hypothesis:

Null (H_0): $\mu_1 = \mu_2 = \ldots = \mu_{N-1} = 0$, and $\lambda_1 = \lambda_2 = \ldots = \lambda_{T-1} = 0$.

Alternative (H_A): Not H_0.

The test statistic for the this null and its alternative is the standard F-test.

The estimated value of the test statistic is 14.18, which exceeds the tabulated value of the F-distribution ($\alpha = 0.5$). Thus the null hypothesis of no cross-section and time effects is rejected at a conventional significance level.

Because the data cover a period when the Lithuanian economy underwent fundamental and large changes we also evaluated the stability of the model parameters over the period of study, During such a period, it is possible that the behavioral relationships embodied in the demand parameters may have changed. Tests were performed from two different perspectives. First, we tested whether the coefficients showed significant change from one quarter to the next. Second, we examined whether the coefficients given by data for a particular quarter of a year were significantly different from those given by the data from the same quarter of another year.

For the first set of tests, we estimated the model with data from two successive quarters and tested whether the coefficients showed significant change between the two quarters. The testing procedure is follows: let $\mathbf{X}_i$ and $\mathbf{X}_j$ be the data vectors (including both dependent and explanatory variables) for two time periods, i and j. Earlier examination of residuals for heteroskedasticity and serial correlation in each of the residuals showed that disturbance terms individually satisfy the standard assumptions of zero mean and constant variance. However, we allowed variances corresponding to different equations to be different as well as contemporaneously correlated. Given these assumptions, first we estimated the demand system for each subperiod separately and obtained the variance-covariance matrices $\hat{\Sigma}_i$ and $\hat{\Sigma}_j$ of the estimated residuals and the upper triangular matrices $P_i = \hat{\Sigma}_i^{-\frac{1}{2}}$ and $P_j = \hat{\Sigma}_j^{-\frac{1}{2}}$. Then, we transformed the data vectors by multiplying $\mathbf{X}_i$ and $\mathbf{X}_j$ by P_i and P_j. Using the transformed variables, we re-estimated the model for each sub-period, i and j, as well as for the two period pooled together. Then we applied a Chow (1960) type test by using the sum of squares of the residuals from each of the estimated models that use the transformed data. Specifically, the test statistic is given by

$$\chi_c = \frac{\left(SSE_R - SSE_U\right)/K}{SSE_U / \left(N_1 + N_2\right) - 2K}, \qquad [19]$$

where SSE_R is the restricted sum of squares of the residuals obtained from the model estimated from the pooled data, SSE_U is the unrestricted sum of squares (obtained by adding the residual sum of squares of residuals from the two subsample estimates), N_i is the sample size in period i (120 in this instance for all i and j), and K is the number of coefficients estimated (20 in this instance). Under the null hypothesis that the coefficients are unchanged between periods i and j, the test statistic follows χ^2 with K degrees of freedom[6]. The estimated values of the test statistic and the 95% critical values are presented in Table 4.

Next, and in a similar manner, we tested whether the coefficients in one quarter were significantly different from those of the same quarter in another year. Specifically, starting with the third quarter of 1992 (the first three months data), we estimated the coefficients from the data for a particular quarter in different years and then tested whether the coefficients were statistically different between the same quarters in any two years. The results are shown also in Table 4.

It is evident from Table 4 that the estimated values of the test statistics are well below the 95% value for the test. This result is irrespective of whether one is conservative and uses an F-distribution or the χ^2-distribution for the test statistic under the null of parameter stability (see endnote 6). Hence, we have no evidence suggesting that the parameters of the demand system have undergone (statistically significant) change over the period of study.

Economic theory is based on consumers being free from money illusion. However, the large price changes give us a good opportunity to examine whether consumers behave in ways consistent with this assumption. Therefore, in addition to parameter stability, we also examined whether the expenditure data were consistent with the absence of money illusion. Zero-degree homogeneity restriction of the demand parameters would imply absence of money illusion on the part of the economic agents, so we formally tested the homogeneity restriction. The results, reported in Table 5, show that among the six expenditure categories, clothing satisfies the restriction at a 5% significance level, while housing, entertainment, and education satisfy the restriction at a 1% significance level. Test results for food, household furnishings, and other expenses clearly contradict the assumption of no money illusion. A likelihood ratio based test of zero-degree homogeneity for the entire system yields a test statistic of 597.00. Under the null of zero-degree homogeneity, the test statistic follows a Chi-square distribution with 5 degrees of freedom. Since the computed value of the test statistic (597.00) far exceeds the 95% value of Chi-square distribution with 5 df, the restriction is clearly rejected by the data. In sum, the expenditure data reveal presence of money illusion.

5. Discussion and Conclusions

Significant economic changes have occurred in Lithuania since the early 1990s. Major market reforms were introduced through phased liberalization of producer and consumer prices. Throughout the period, consumers devoted a relatively large share of total expenditures to food, and there is some evidence of an increasing share of budgets to housing.

During the early reform period examined here, households were relatively responsive to changes in prices and total expenditures in their purchases. Estimated price and expenditure elasticities are consistent with levels found in relatively low income countries. Evidence from the household survey shows the relative stability of the demand parameters. Since prices were rising rapidly during this period, the results suggest that households have made significant adjustments in their consumption patterns.

The price elasticity of food, the major share of consumers expenditure, is relatively high (around-1.0). This level is observed despite allowances for home production of food. More detailed analysis of food expenditures shows shifts among foods, with increases in grains, vegetables and potatoes and decreases in dairy products in terms of shares of total food expenditures.

During this period housing prices were (artificially) low. As housing (including utilities) and food prices increased, households were reported to discontinue paying utility and housing bills in order to buy food. We observe price response not too far from unity, despite large, official increases to rent and utilities.

The stability of demand parameters, along with some evidence of money illusion, indicates the severity of price effects on Lithuanian households during this period. A system of social support and assistance programs designed to maintain “Minimum Living Levels” were especially important to lower income households, although it became increasingly difficult for the country to continue this level of social support. The results suggest the importance of putting in place longer run economic adjustment programs focused on income and labor policies in addition to shorter run targeted assistance to those most affected by price adjustments.

Table 1. Nominal Expenditure and Expenditure Shares: July 1992 - December 1994

Month	Nominal Expenditure		Expenditure Share					
	Total Expenditure	Per Capita Expenditure	Food	Clothing	Housing	Household Furnishings	Enter. & Education	Other Expenses
July 1992	127.18	50.90	66.59	9.45	3.81	3.58	1.65	14.92
August 1992	150.72	60.46	70.64	10.32	4.14	2.73	1.36	10.81
September 1992	135.00	54.36	71.27	10.02	3.42	2.82	1.50	10.97
October 1992	173.23	69.30	70.64	10.14	5.02	2.53	1.26	10.41
November 1992	179.62	72.62	72.73	6.44	7.13	1.62	1.81	10.27
December 1992	235.92	93.64	72.74	5.64	8.71	1.36	1.52	10.03
January 1993	204.60	82.05	71.51	5.44	10.61	2.06	1.33	9.05
February 1993	218.69	87.34	73.26	5.02	9.74	1.65	1.67	8.66
March 1993	243.17	97.53	71.51	6.18	8.85	1.71	1.57	10.18
April 1993	293.90	117.76	74.26	6.33	7.74	1.54	1.15	8.98
May 1993	299.40	120.60	74.05	6.46	6.85	1.61	1.47	9.56
June 1993	362.32	145.95	74.88	6.15	6.86	1.83	1.32	8.96
July 1993	356.94	143.49	75.76	5.75	6.83	1.66	1.04	8.96
August 1993	388.32	157.57	71.91	7.81	7.90	1.81	1.70	8.87
September 1993	377.85	152.68	72.62	7.08	7.66	2.02	1.54	9.08
October 1993	424.81	169.35	69.60	9.34	7.57	2.32	1.48	9.69
November 1993	453.13	182.04	67.76	9.08	10.24	1.71	2.15	9.06
December 1993	543.36	221.43	67.91	7.40	10.29	1.92	2.80	9.68
January 1994	483.50	195.23	69.10	6.69	11.39	2.21	1.87	8.74
February 1994	522.91	210.04	70.20	6.25	10.97	2.21	1.78	8.59
March 1994	548.13	221.77	69.00	7.07	11.09	2.14	1.90	8.80
April 1994	533.10	214.36	68.37	7.79	10.85	2.34	1.61	9.04
May 1994	523.25	210.35	68.22	7.39	10.23	2.07	1.77	10.32
June 1994	506.25	204.50	69.93	7.11	8.09	2.24	1.87	10.76
July 1994	546.82	219.63	71.24	6.69	8.40	2.00	1.51	10.16
August 1994	547.82	220.43	70.07	7.36	8.81	2.11	2.04	9.61
September 1994	538.01	217.74	69.49	6.81	9.87	1.99	1.85	9.99
October 1994	589.34	237.97	68.14	8.54	9.94	2.38	1.62	9.38
November 1994	602.23	242.10	67.06	7.77	11.74	1.97	2.15	9.31
December 1994	663.49	265.77	66.13	6.19	13.49	1.88	2.59	9.72

Source: Department of Statistics Publication, Government of Lithuania.

Note: The nominal expenditure figures are in Litas.

Table 2. Estimated Coefficients and t-ratios

	P_{food}	P_{cloth}	$P_{housing}$	P_{hfurn}	P_{entedu}	Real Exp.
Food	-0.0179	0.0068	-.0013	-0.0315	-0.0016	-0.0141
	(-7.60)	(3.23)	(-1.09)	(-13.99)	(-1.85)	(-6.15)
Clothing		-0.0214	-0.0274	0.0187	-0.0135	0.0035
		(-1.24)	(0.0187)	(1.90)	(-1.80)	(1.53)
Housing			-0.0043	0.0245	-0.0015	-0.0028
			(-1.54)	(7.20)	(-0.62)	(-2.14)
Household				0.0079	0.0027	-0.0295
Furnishings				(0.99)	(0.62)	(-11.38)
Entertainment					0.0046	-0.0006
& Education					(2.92)	(-0.66)

Note: Figures in parentheses are t-ratios.

Table 3. Estimated Price and Expenditure Elasticities

	Price Elasticity						
	Food	Clothing	Housing	Household Furnishings	Entertainment & Education	Other Expense	Expenditure Elasticity
Food	-1.0092	0.0103	-0.0012	-0.0392	-0.0017	0.0606	0.9816
Clothing	0.0533	-1.2759	-0.3505	0.2341	-0.1730	0.4646	1.0442
Housing	0.0355	-1.1172	-1.1757	1.0161	-0.0611	0.4235	0.8869
Household Furnishings	-0.0933	0.2230	0.2674	-0.8862	0.0352	-0.2104	0.6866
Entertainment & Education	-0.0582	-0.6962	-0.0790	0.1451	-0.7636	0.4866	0.9677
Other Expenses	0.1396	0.3894	0.1053	-0.3090	0.0994	-1.9701	1.5092

Note: Price and expenditure elasticities are computed at mean values of prices and expenditure

Table 4. Test of Parameter Stability

One Quarter vs. the Next Quarter		One Quarter vs. the Corresponding Quarter in Another Year	
Quarters Compared	Test Statistic	Quarters Compared	Test Statistic
1992 Q3 vs. 1992 Q4	0.558	92Q3 vs. 93Q3	0.868
1992 Q4 vs. 1993 Q1	1.048	92Q3 vs. 94Q3	0.658
1993 Q1 vs. 1993 Q2	0.690	93Q3 vs. 94Q3	0.505
1993 Q2 vs. 1993 Q3	0.548	92Q4 vs. 93Q4	0.663
1993 Q3 vs. 1993 Q4	0.527	92Q4 vs. 94Q4	0.796
1993 Q4 vs. 1994 Q1	0.566	93Q4 vs. 94Q4	0.421
1994 Q1 vs. 1994 Q2	0.784	93Q1 vs. 94Q1	0.884
1994 Q2 vs. 1994 Q3	0.534	93Q2 vs. 94Q2	1.263
1994 Q3 vs. 1994 Q4	0.669		

Note: The 95% tabulated value of χ^2 distribution with 20 degrees freedom is 31.41, and the 95% tabulated value of F^{20}_{200} is 1.57.

Table 5. Test of Zero-degree Homogeneity of the Demand System

	Test Statistic (F-Statistic)	Distribution of F-Statistic | H_0	95% Critical Value of F^1_∞	99% Critical Value of F^1_∞
Food	135.06	F^1_{NT-K}	3.84	6.63
Clothing	0.51	F^1_{NT-K}	3.84	6.63
Housing	5.35	F^1_{NT-K}	3.84	6.63
Household Furnishings	247.37	F^1_{NT-K}	3.84	6.63
Entertainment & Education	6.30	F^1_{NT-K}	3.84	6.63
Other Expenses	262.89	F^1_{NT-K}	3.84	6.63

ENDNOTES

[1]In January 1996, Lithuania's Household Budget Survey was replaced with a new survey which follows a methodology common to that developed for the European Community Household Panel (Verma and Clémenceau). The 1992-94 Household Budget Survey provides current and complete consumption and expenditure data for the period of interest.

[2]Estimation of the error components model or random effects model requires variation in prices across cross-section units.

[3] Judge, et al. (1985), chapter 13 provides detailed discussion on the estimation procedure.

[4] For heteroskedasticity, we implemented Bartlett's test [see Judge et. al., p. 447] on the residuals from each equation, and did not find any evidence of the problem. Also, we tested for serial correlation in the computed residuals. The test did not show strong evidence of serial correlation.

[5] There is some disagreement in the literature regarding the appropriate expression for price and expenditure elasticities. This disagreement stems from the linear approximation used to estimate the demand system rather than the full AIDS model being estimated (see Hahn 1994 on the issue). Because of this approximation, the estimated elasticities using any of the formulae available in the literature are also approximations to the true elasticities, and in finite samples these estimates may be subject to bias.

[6] This uses large sample distribution for the test statistic. If one is conservative in performing the test and chooses to use $F^{k}_{N_1+N_2-2K}$ as the distribution of the test statistic under the null, the conclusion would remain unchanged.

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