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Abstract

Elasticities of demand for meat imports in Russia are estimated using an AIDS model. The model differentiates among sources of imports as well as kinds of meat, but since the number of observations on Russian imports is limited, an improved block-substitutability restriction is introduced to conserve degress of freedom. The estimates of expenditure elasticities are positive for beef, pork, and chicken imported from western countries, and for beef and chicken, are larger than one. The expenditure elasticities are negative for beef and pork imported from former Soviet trade block countries. (Chicken is not imported from these countries.) Consistent with logic, the (compensated) cross-price elasticities indicate that products imported from different sources are substitutes. These estimates are perhaps the first available for the Russian economy, and not surprisingly, they indicate that declining real incomes in Russia mean decreasing meat imports from western countries.

ELASTICITES OF DEMAND FOR IMPORTED MEATS IN RUSSIA

This paper provides estimates of elasticities of demand for imported meat in Russia, which hitherto have not been available. Demand elasticities are a necessary input for analysis of trade and welfare policies, but applied econometric work with data from Russia's transitory economy faces two challenges: limited number of observations and potential complexity of models. Thus, it is necessary to find a flexible functional form for a demand system that allows for reasonable restrictions to reduce the number of estimated parameters.

The model should be flexible enough to take into account possible differences in trade for Russia between western exporters on the one hand and its former Soviet trade block partners on the other. The model should also accommodate possible substitution effects among different kinds of meat. The popular Armington (1969) trade model allows for source-differentiation, but it is restrictive otherwise. Specifications commonly used in studies of domestic demand are flexible, but are likely to have a degrees-of-freedom problem when products are differentiated by both kinds and sources of imports.

A model developed by Yang and Koo (1994) allows for direct-price effects among groups of products and among different imports within each group, but some restrictions in their model are difficult to justify by economic theory. In this paper we improve the restriction of block-substitutability (BLSUB) introduced by Yang and Koo. The improved block-substitutability (IBLSUB) (a) is consistent with economic theory, (b) further reduces the number of parameters to be estimated, and (c) is supported by the data in our sample. This model is used to estimate demand elasticities for imported meats using quarterly observations for 1994.1 through 1998.2.

The paper is organized as follows. The next section discusses the development of the model to be estimated. We show how the model relates

to the existing literature. As indicated, degrees of freedom is an important

problem for fitting a demand system to Russian data. Thus, we devote a

section to the question of degrees of freedom for alternative versions of

Almost Ideal Demand System (AIDS) models. Then, the data, estimation

procedures, and evaluation methods are outlined. The paper concludes

with a discussion of the empirical results.

Model Development

Overview

The early literature on estimation of import demand elasticities was mostly

concerned with individual countries and large aggregates of commodities. This was

justified by the interest of researchers in predicting gross trade flows and evaluating

the impact of exchange rate fluctuations on balance of payments (Sarris, 1981).

However, when research shifted to analyzing intervention policies and assessing the

degree of competitiveness of different exporters, the methodology shifted towards

microeconomic foundations.

One of the most popular models, emphasizing the importance of

disaggregation, is the Armington trade model. It allows for source-differentiation

by distinguishing goods not only by kind, but also by place of origin. Among the

factors contributing to the popularity of the Armington model is its ease of use.

The Armington model assumes separability with respect to kinds of

products, constant elasticity of substitution (CES), and homotheticity. An

implication of separability for econometric models is that the demand for each

group of products can be estimated independently from other groups. With CES,

elasticities of substitution are identical and constant for all products within a group,

thus implying weak separability with respect to sources. Homotheticity means that

all income elasticities are the same and unitary, which implies that market shares of

importing countries are not affected by the size of these countries' markets. This, in turn, implies 'homothetic separability' (Alston, et al., 1990) with respect to sources of import within each group of products.

In his original article, Armington argues that these assumptions are innocuous. Alston et al. suggest, however, that if inappropriate, these restrictions result in omitting relevant explanatory variables, and consequently, in introducing bias in the estimates of elasticities. They test for homotheticity and separability among import sources using data from the international cotton and wheat markets. Both parametric and non-parametric tests reject Armington's assumptions. Alston et al. do not, however, test for separability among groups of products (only one product is considered in their study).

While the Armington specification dominated the import demand literature, more flexible functional forms for estimating demand systems became available and extensively used in domestic demand studies. The AIDS model of Deaton and Muellbauer (1980a,b) is one of them. De Gorter and Meilke (1987) are among the first authors to use the AIDS specification in the context of estimating source-differentiated demand for imported products. Although all imports of wheat considered in their study are aggregated into a single commodity, the common assumption of weak separability between import and domestic demand is relaxed. Thus, their model distinguishes between two sources, and their results contrast with earlier findings and indicate the importance of source-differentiation.

Both de Gorter and Meilke and Alston et al. deal with a single kind of commodity (wheat or cotton). The former drew some criticism (von Cramon-Tabuadel, 1988) for not distinguishing among kinds of wheat. Theoretically, formulation of source-differentiated AIDS model for more than one good is straightforward. In practice, however, such a model will quickly grow in size. For four groups of products and five sources of imports in each group, an unrestricted AIDS model will have 20 equations and 20*(20+2)=440 parameters to estimate. As

we shall see, even the standard assumptions of adding-up, homogeneity and symmetry may not be sufficient to solve the degrees-of-freedom problem.

The first attempt to construct a model that allowed for both source-differentiation and direct cross-price effects among similar products was made by Yang and Koo. They start with an AIDS model and introduce an assumption of block-substitutability, which reduces the number of parameters to be estimated. In imposing block-substitutability, however, the standard assumption of adding-up is violated, and the procedure does not take full advantage of further increasing the degrees of freedom. An improved assumption of block-substitutability (IBLSUB), introduced in this paper, will make the source-differentiated AIDS model a better tool for international demand studies when working with relatively few observations.

Block-Substitutability and Weak Separability

Block-substitutability is best explained by considering the underlying budgeting process. In the Armington model, consumers are assumed to allocate their expenditures in two stages. In the first stage, they decide how much of each kind of good (beef, pork, chicken, and other meat products) should be purchased. During the second stage, expenditures on each good are allocated among the different sources of imports (German beef, Irish beef, etc.). Once a decision is made during the first stage of the allocation process, it is assumed to be irreversible. It cannot be altered during the second stage.

As is well known, weak separability is a necessary and sufficient condition for the second stage of the two-stage budgeting process. An important implication of weak separability in the Armington model is that the substitution effect between any two products in different groups (say, Chinese pork and US chicken) is limited to that of the income effect of the price change. No direct cross-price links are allowed (Alston et al.).

The standard AIDS model is summarized in Appendix A, and the two-stage budgeting ideology of the Armington model can be expressed in the LAIDS framework as follows (Edgerton, 1996):

(1a) $\quad w_i = \alpha_i + \sum_j \gamma_{ij} \ln(p_j) + \beta_i \ln\left(\frac{E}{P*}\right)$ (between-group allocation),

(1b) $\quad w_{ih} = \alpha_{ih} + \sum_k \gamma_{ihik} \ln(p_{ik}) + \beta_{ih} \ln\left(\frac{E_i}{P_i*}\right)$ (within-group allocation).

During the first-stage expenditures are allocated over kinds of meat i, j in (1a). E represents the amount of total expenditures and $P*$ represents the Stone price index for all meat products. The price and the budget share of the ith meat are p_i and w_i respectively. In this stage, products of the same kind, but imported from different sources, are assumed to be perfect substitutes.

Once w_i becomes known (as does E_i), the second-stage allocation of group expenditure (E_i) over sources of origin begins, equation (1b). The variable w_{ih} is the expenditure share for i^{th} kind of meat imported from country h. E_i denotes expenditures on i^{th} kind of meat and P_i* is the geometric average of prices of this meat imported from different sources (group i Stone index). The variable p_{ik} is the price of ith kind of meat imported from country k.

In contrast, Yang and Koo's assumption of block-substitutability does not require two-stage budgeting. Expenditures are allocated simultaneously over all products under consideration. This allows for direct cross-price effects among the products belonging to different groups. Their model assumes, however, that while allocating expenditures among different sources of the same good, consumers do not distinguish among sources of other goods. For example, when expenditures are allocated among different sources of beef imports, the pork produced in Germany is perceived to be identical in all respects to the pork produced in the United States.

Since Yang and Koo's source-differentiated LAIDS model does not assume two-stage budgeting, the model can be written as:

(2a) $w_{ih} = \alpha_{ih} + \sum_k \gamma_{ihik} \ln(p_{ik}) + \sum_{j \neq i} \gamma_{ihj} \ln(p_j) + \beta_{ih} \ln\left(\dfrac{E}{P*}\right)$, where

(2b) $\ln(p_j) = \sum_k \bar{s}_{jk} \ln(p_{jk})$.

Here, $\bar{s}_{jk}$ is the average *group* share of jth kind of meat imported from k^{th} country.[1] The variable p_j can be interpreted as the average price of jth kind of meat, and γ_{ihj} is then the cross-price effect of this price on the budget share of product i_h.

    The system of equations (2a) explicitly recognizes cross-price links among meat products from different groups (the third term on the right-hand side (RHS)). At the same time, these links are not source-specific as they are within the groups (the second term on the RHS).

    One shortcoming of the BLSUB model, as formulated in (2), is that the individual equations of the system have different variables on the RHS. This also occurs if each meat group does not have the same number of sources. Consequently, it is not clear how the adding-up and symmetry restrictions can be imposed on their model. Yang and Koo modify the standard theoretical restrictions in the following way:

(3a) $\sum_i \sum_h \alpha_{ih} = 1, \quad \sum_i \sum_h \beta_{ih} = 0, \quad \sum_h \gamma_{ihik} = 0,$
       $\sum_i \sum_h \gamma_{ihj} = 0, \forall j \neq i ; \quad$ (adding-up)

(3b) $\sum_k \gamma_{ihik} + \sum_{i \neq j} \gamma_{ihj} = 0, \forall i, h ; \quad$ (homogeneity)

(3c) $\gamma_{ihik} = \gamma_{ikih}, \quad \forall i ; \quad$ (within-group symmetry)

    Although the restrictions of homogeneity are the same as in the standard LAIDS model (see Appendix A), the restrictions of adding-up and symmetry are different. Specifically, symmetry is imposed only on the within-group cross-price parameters, while adding-up is imposed both on between-group cross-price

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[1] Yang and Koo use $\ln(p_j) = \sum_k w_{jk} \ln(p_{jk})$ in place of (2b). This is clearly a typo, since group shares s rather than market shares w should be used in calculating the average price of the jth kind of meat, which is what (2b) does. We use average shares $\bar{s}$ in order to reduce the problem of simultaneity (see Eales and Unnevehr (1988) for details). Average group shares $\bar{s}_{jk}$ are computed from T observations as follows: $\bar{s}_{jk} = 1/T \sum_{t=1}^{T} s_{jkt}$.

parameters (fourth term in (3a)) and within-group parameters (third term in (3a)). In the standard LAIDS model, which does not differentiate among sources, symmetry is imposed on all cross-price parameters while adding-up is usually imposed only on the parameters of one residual equation.

Thus, the Yang and Koo approach has two major shortcomings. First, because not all symmetry restrictions can be imposed, the gain from block-substitutability in terms of degrees of freedom becomes less obvious. Second, the within-group adding-up restriction is not justified by economic theory. In particular, theory does not require that within-group cross-price parameters add to zero (third term in (3a)).

To see why this is so, recall that the adding-up restriction follows from the budget constraint. For the import demand model which differentiates both among kinds of goods and sources of imports, the budget constraint is:

$\sum_i \sum_h p_{ih} g_{ih}(E,P) = E,$ where E stands for total expenditures, p is a vector of prices, and g (E,P) represents quantity of good i imported from country h as a function of expenditures and prices. For Yang and Koo's within-group adding-up restrictions to be satisfied, we must be able to observe a within-group budget constraint as well, i.e. $\sum_h p_{ih} g_{ih}(E_i,P) = E_i, \forall i$, where E_i stands for within-group expenditure. This constraint will hold under the assumption of two-stage budgeting, but Yang and Koo assume one-stage budgeting and have total expenditures in all equations of the model (see (2a)). Hence, their within-group adding-up restriction implies a within-group budget constraint of the form: $\sum_h p_{ih} g_{ih}(E,P) = E, \forall i$. This says that each group's expenditures must be equal to the total expenditures on all products, which is obviously incorrect.

In the following section, a modified version of (2) is suggested.

Imposing Block-Substitutability via Restrictions on LAIDS

We rewrite the AIDS model so that it explicitly distinguishes among kinds of goods and sources of origin:

$$(4) \qquad w_{ih} = \alpha_{ih} + \sum_j \sum_k \gamma_{ihjk} \ln(p_{jk}) + \beta_{ih} \ln\left(\frac{E}{P}\right).$$

Here subscripts i and j denote kinds of meat products, and subscripts h and k denote sources of origin; w_{jk} and p_{jk} are respectively the budget share and the price of jth kind of meat imported from k^{th} country; E represents total expenditures on meat, P is the linearized Stone price index for all meat products; α, β and γ are parameters.

Equations (2a,b) result from imposing on (4) the following restrictions, which is essentially Yang and Koo's block substitutability:

$$(5a) \qquad \gamma_{ihjk} = \bar{s}_{jk} \gamma_{ihj}, \quad \forall j \neq i.$$

Rewriting (5a) in terms of another product r from group j,

$$(5b) \qquad \gamma_{ihjr} = \bar{s}_{jr} \gamma_{ihj}, \quad \forall j \neq i.$$

Combining these two expressions, block-substitutability can be imposed on (4) in terms of the following set of restrictions:

$$(5c) \qquad \gamma_{ihjk} = \frac{\bar{s}_{jk}}{\bar{s}_{jr}} \gamma_{ihjr}, \quad j \neq i, k \neq r \text{ , where } \bar{s}_{jk}, \bar{s}_{jr} \text{ are average group shares, and}$$

$\gamma_{ihjk}, \gamma_{ihjr}$ are parameters of (4). Since the values of $\bar{s}_{jk}$ and $\bar{s}_{jr}$ can be computed

(see footnote 1), this restriction is easy to impose.

The BLSUB model, written as nested within the standard LAIDS specification, has the same variables on the right-hand side of all equations. Thus, the theoretical restriction of adding-up does not have to be imposed separately on the within-group and between-group cross-price parameters. However, it is still not possible to impose symmetry on the between-group cross-price coefficients. Since symmetry is the most powerful tool for reducing the number of estimated parameters, this problem is worth pursuing.

The parameter γ_{ihj} in (2a) relates the price of an aggregate good j with the budget share of a disaggregated product i_h, making the imposition of symmetry impossible. If we could adjust (2a) in such a way that γ_{ihj} is replaced with γ_{ij}, which

relates two aggregate goods, symmetry would be easy to impose. The work of

Hayes, et al. (1990) provides a clue to the solution of this problem.

According to Hayes, et al. if groups of products i and j are quasi-separable, it

is possible to write:

(6a) $\quad \gamma_{ihjk} = s_{ih} s_{jk} \gamma_{ij}, \quad \forall j \neq i$.

Using subscripts r and q to denote some other sources of imports, we can also write:

(6b) $\quad \gamma_{irjq} = s_{ir} s_{jq} \gamma_{ij}, \quad \forall j \neq i$.

Combining (6a) and (6b) and replacing actual group shares with average ones, we

get a set of restrictions, which can be imposed on (4):

(6c) $\quad \gamma_{ihjk} = \dfrac{\bar{s}_{ih} \bar{s}_{jk}}{\bar{s}_{ir} \bar{s}_{jq}} \gamma_{irjq}, \quad \forall j \neq i$. (Appendix A provides an example of how block-

substitutability can be imposed in practice.)

To summarize, the three models discussed above are:

LAIDS. A source-differentiated import demand model formulated in terms of the

standard LAIDS model with adding-up, homogeneity and symmetry imposed. This

model can be written as follows:

(7) $\quad w_{ih} = \alpha_{ih} + \sum_{j} \sum_{k} \gamma_{ihjk} \ln(p_{jk}) + \beta_{ih} \ln\left(\dfrac{E}{P}\right), \quad \forall i,h$ (source-differentiated

LAIDS model);

(8a) $\quad \sum_{i} \sum_{h} \alpha_{ih} = 1, \quad \sum_{i} \sum_{h} \beta_{ih} = 0, \quad \sum_{i} \sum_{h} \gamma_{ihjk} = 0, \ \forall j,k$ (adding-up);

(8b) $\quad \sum_{j} \sum_{k} \gamma_{ihjk} = 0, \ \forall i,h$ (homogeneity);

(8c) $\quad \gamma_{ihjk} = \gamma_{jkih}, \forall i,h,j,k$; (symmetry).

BLSUB. A source-differentiated import demand model under the assumption of

block-substitutability as formulated by Yang and Koo: equations (2a), (2b), (3a),

(3b), and (3c), above.

IBLSUB. A source-differentiated import demand model under the improved assumption of block-substitutability formulated in terms of LAIDS model: equations (7), (8a), (8b), (8c) above, and

(8d) $\quad \gamma_{ihjk} = \dfrac{\bar{s}_{ih}\bar{s}_{jk}}{\bar{s}_{ir}\bar{s}_{jq}} \gamma_{irjq}, \quad \forall j \neq i; \forall h, r, k, q \quad$ (block-substitutability).

The formulae for elasticities are similar to those derived by Chalfant (1987) and assume constant group shares (see Appendix A).

Adequacy of Restrictions

The purpose of imposing block-substitutability is to reduce the number of parameters to be estimated. Thus, it is important to ensure that the number of restrictions is sufficient to estimate the models under consideration. A rough rule of thumb is that the average number of unrestricted parameters per estimated equation should be no more than 40% of the number of observations. This is a modification of a rule of thumb suggested in Belsley, et al. (1980, p.17).

In this section the rule of thumb is used to compare the "data-efficiency" of the LAIDS, BLSUB and IBLSUB specifications of the source-differentiated import demand model. We define a data-efficiency model as the most parsimonious model that is a statistically adequate representation of the data and that at a minimum satisfies the rule of thumb.

In this context, we address the question of the minimum number of observations required to fit alternate models. If N is the number of equations in a model including one residual equation and if K is the number of groups of products considered in the model not including the residual group, then since one equation will be dropped in estimating the system, $L=(N-1)/K$ is the average number of sources per group. Define T as a minimum number of observations required to satisfy the rule of thumb for a given model, then T can be approximately expressed as a function of N and K or as a function of K and L. The latter is more convenient. Thus, given number of groups of products in a model (K) and the average number of

sources of imports in a group (*L*), we can calculate (approximately) the number of observations necessary for the rule of thumb to be just satisfied. Table 1 contains a summary of results for the three models.

Table 1. Calculation of minimum number of observations required to satisfy the Rule of Thumb[a]

| Average number of unrestricted parameters per equation	
LAIDS	$(N+4)(N-1)\big/2N$
BLSUB	$(N-1)(N+2K^2+K-1)\big/2KN$
IBLSUB	$(5KN-5K+N^2-2N+1+K^3-K^2)\big/2KN$
Rule of thumb equation	
LAIDS	$100(N+4)(N-1)\big/2NT=40$
BLSUB	$100(N-1)(N+2K^2+K-1)\big/2KNT=40$
IBLSUB	$100(5KN-5K+N^2-2N+1+K^3-K^2)\big/2KNT=40$
Minimum number of observations *(obtained by* **solving 'rule of thumb equations' for *T* with respect to *K* and *N)***	
LAIDS	$5(N+4)(N-1)/4N$
BLSUB	$5(N-1)(N+2K^2+K-1)\big/4KN$
IBLSUB	$5(5KN-5K+N^2-2N+1+K^3-K^2)\big/4KN$
Minimum number of observations *(obtained by* **solving 'rule of thumb equations' for *T* with respect to *K* and *L)***	
LAIDS	$5(LK+5)LK/4(LK+1)$
BLSUB	$5LK(L+2K+1)\big/4(LK+1)$
IBLSUB	$5K(5L+L^2+K-1)\big/4(LK+1)$

[a] For the AIDS model with no block-substitutability imposed, all formulae are exact. All other formulae give approximate results and are exact only when all groups of products have the same number of sources of imports.

In the graphs 1 and 2, T is plotted as a function of K, given L=2 and 4 respectively. For example, if there are five groups in a model and each group has, on average, four sources of import, then for the rule of thumb to be satisfied, LAIDS, BLSUB and IBLSUB models would require at least 30, 18 and 12 observations respectively. As discussed below, we have 18 observations, four product groups (K=3), and two sources (L=2) for three of the four products. Thus, the minimum number of observations to meet the rule of thumb is LAIDS, 12; BLSUB, 10; IBLSUB, 9. It is, of course, preferable for T to exceed these minima.

Graph 1:   Minimum Observations to Satisfy the Rule of Thumb by Number of Products, for L=2 (average number of sources per group)
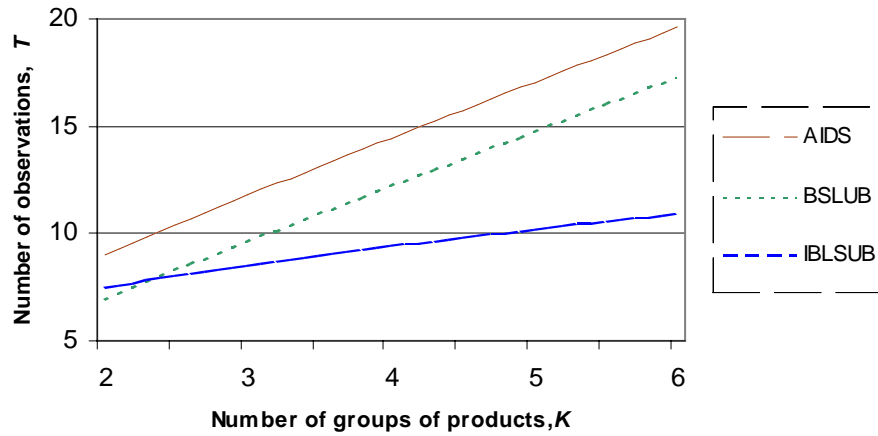
Graph 2:   Minimum observations to satisfy the Rule of Thumb by Number of Products, for L=4 (average number of sources per group)
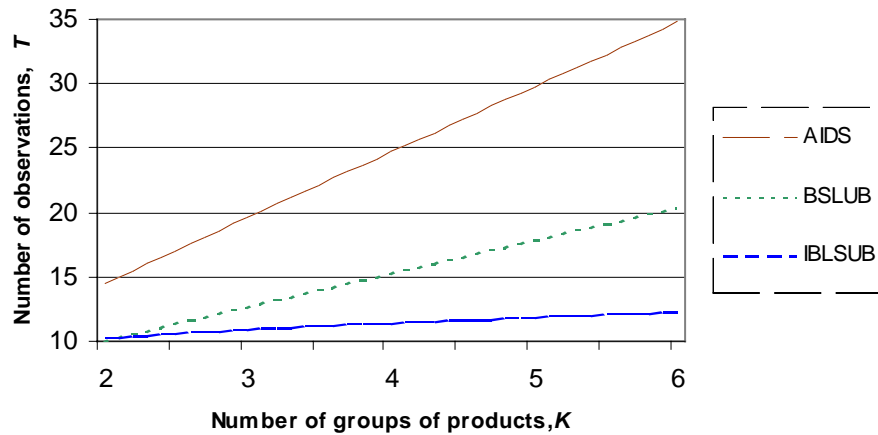
Data, Estimation, and Evaluation Methods

Quarterly series of prices and volumes of meat imports for the period 1994.1 to 1998.2 are obtained from various issues of the Customs Statistics of International Trade. In order to avoid dealing with potential structural change, the sample does not contain any data generated during and after

the economic crisis of August 17, 1998[2]. All quantities are in thousands of kilograms and all volumes are in thousands of US dollars. Prices are obtained by dividing volumes by quantities and therefore, are in nominal US dollars per kilogram. A test for exogeneity of expenditures uses the Consumer Price Index (CPI) and real per-capita income. These series are found in various issues of Russia in Figures (1994-1998) and recalculated in terms of US dollars using the official exchange rate. Seasonality was not evident in the data series, and so the models do not include seasonal dummies.

Each commodity in the Customs Statistics bulletin is assigned a unique code according to the Commodity Nomenclature of Foreign Trade (CNFT). Four groups of meat products considered in this study are described in Table 2.

Although the Commodity Nomenclature has many more subgroups of meat products, the data were available mostly on these six categories. Occasional imports of other products were negligible. Beef, pork, and chicken form natural aggregates. While fish is an important dietary item, it is not included in the model, as Russia both imports and exports large volumes of fish and fish products.

The rationale for using a source-differentiated model in this study is our view that Russia's international trade patterns are strongly affected by non-market forces. In particular, the prevalence of barter and the existence of special customs regimes among the Newly Independent States (NIS) suggest that the model should explicitly distinguish between western and NIS exporters of meat products.

Hence, all groups of products used in this study are divided into two subgroups, depending on whether or not the exporting country is a NIS member.

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[2] On August 17, 1998 the Russian government announced a sovereign default on international and domestic loans. As a consequence of this move, US/R exchange rate, which had been stable for more than a year, quadrupled within just few months, causing major shift of consumer demand from imported to domestically produced commodities.

Table 2. Aggregation of Imported Meat Products by Kind

Aggregate	CNFT	Brief Description of the Commodity
Beef	0201	beef fresh and chilled
	0202	beef frozen
Pork	0203	pork fresh and chilled
	0204	pork frozen
Chicken	0207	meat and meat products of poultry
Residual	1602	other prepared and canned meat products

Table 3. Aggregation of Meat Imports by Kinds and Sources of Origin

	Group 1	Group 2
Beef	all foreign countries excluding NIS members (Germany, Denmark, Ireland, China, USA, France)	all NIS (Kazakhstan, Moldova, Ukraine)
Pork	all foreign countries excluding NIS members (Germany, Denmark, Canada, China, Netherlands, USA, France)	all NIS (Kazakhstan, Moldova, Ukraine)
Chicken	USA	all other exporters (Belgium, UK, Germany, Denmark, Netherlands, France)
Residual	Aggregated over all sources	Aggregated over all sources

This rule was not used for poultry due to the negligible volume of the intra-NIS trade. Rather, the poultry groupings are USA and all other. (The original research considered three levels of aggregation and 14 observations, see Soshnin (1998).)

The sample includes all exporting countries whose average share, in at least one group of products, exceeded three percent. A few exceptions to this heuristic rule were made so as to include countries with large import volumes in a few quarters. This results in the aggregation described in the Table 3. A summary of the sample statistics is provided in Table 4. The variable names defined in Table 4 are used subsequently to represent the prices and budget shares for the respective equations.

Note, despite high volumes of trade, Byelorussia is not among the NIS exporting countries. This is because after 1994, when it formed a union with Russia, no data on trade are available. This may be a problem, since the line between domestic and intra-NIS supplies is difficult to define. As far as trade is concerned, Byelorussia is very similar to the Ukraine, which is treated as an exporting country here. Econometrically, the absence of Byelorussia in the sample creates a potential errors-in-variables problem, which may lead to biased estimates.

In sum, the demand system consists of seven equations. There are two equations each for beef, pork, and chicken, representing different sources of imports for these meats. In addition, there is a residual meat import equation. The residual equation is deleted to achieve the adding-up restriction, and the model (with alternate restrictions imposed, as noted below) is estimated by an iterative Generalized Least Squares procedure in the TSP package.

The model is subjected to a number of misspecification tests, which are outlined in Appendix B. Table 5 provides a summary of the results of these tests: Wald statistics and t-statistics (for endogeneity test). In the table, the null hypotheses are abbreviated as:

ACOR: no autocorrelation is present;

HETR: no heteroscedasticity is present;

ARCH: no autoregressive conditional heteroscedasticity is present;

EXOG: expenditures are exogenous.

P-values indicate the minimum level of significance, at which the corresponding null hypothesis can be rejected. Thus, if we set the threshold of rejection at the traditional 5% level, all hypotheses with the associated *P*-value of 0.05 or less should be rejected.

The hypotheses of "no autocorrelation" and "no heteroscedasticity" are rejected in three out of six equations. The hypothesis of "no autoregressive heteroscedasticity" is rejected in two out of six equations. The hypothesis of exogeneity of expenditures is rejected in only one out of six equations.

Thus, autocorrelation, heteroscedasticity and autoregressive conditional heteroscedasticity may be a problem. To correct for autocorrelation, it is possible to try alternative dynamic specifications of the model. Eales and Unnevehr (1988) use first differences, while Assarsson (in Edgerton, et al.,1996, pp.196-203) discusses estimation of the AIDS model in error correction form. The correction for heteroscedasticity depends on its particular form. In principle, a variety of estimators are also available to correct for autocorrelation and heteroscedasticity.

The most problematic equations are the second (BF2) and fourth (PK2): the demand equations for beef and pork imported from NIS countries. This suggests a need for a deeper investigation of data and models for these particular sources, rather than using "mechanical" methods of correcting for autocorrelation and heteroscedasticity on the entire system of equations.

Another question is whether the aggregations are appropriate. As mentioned above, the NIS aggregation does not include Byelorussia. Moreover, the "western" aggregation includes a number of large importers, and for pork includes China. We experimented with other models that involved greater disaggregation (some

Table 4. Summary of Sample Statistics, Russian Meat Imports, 1994.1-1998.2

Kind of meat	Source	Average vol. of import '000 USD	St. error	Average group share, %	Average market share, %
BEEF		**150752**	**54959**		**30,8**
BF1	Foreign	88586	43383	58,8	18,1
BF2	NIS	62166	20520	41,2	12,7
PORK		**108320**	**30363**		**22,1**
PK1	Foreign	100976	28532	93,2	20,6
PK2	NIS	7343	5744	6,8	1,5
CHICKEN		**156908**	**52204**		**32,1**
CH1	USA	118314	45719	75,4	24,2
CH2	Other	38594	12013	24,6	7,9
RESID		**73472**	**33329**		**15,0**
Total meats		*489451*	*106730*		

Table 5. Results of Misspecification Tests

Hypothesis	ACOR		HETR		ARCH		EXOG	
	Wald	P-value	Wald	P-value	Wald	P-value	t-stat	P-value
$\delta_{11}..\delta_{16}=0^a$	9.899	0.129	6.773	0.342	28.757	7E-05	1.996	0.046
$\delta_{21}..\delta_{26}=0$	20.689	0.002	13.709	0.033	5.780	0.448	-1.825	0.068
$\delta_{31}..\delta_{36}=0$	16.274	0.012	11.841	0.066	4.511	0.608	1.352	0.176
$\delta_{41}..\delta_{46}=0$	58.205	1E-10	23.529	0.001	28.108	9E-05	-0.583	0.560
$\delta_{51}..\delta_{56}=0$	9.375	0.153	10.135	0.119	7.460	0.280	-2.566	0.010
$\delta_{61}..\delta_{66}=0$	3.105	0.796	26.418	0.0002	8.296	0.217	0.882	0.378

[a]Parameter δ refers to equations (B2)-(B5). For EXOG the corresponding hypotheses are $\delta_1=0$, ..., $\delta_6=0$ and we use a t-test instead of Wald test.

of which are in Soshnin), but the lack of data is a limitation. Since we are not in a

position to check the quality of the data and since the small number of observations

limits our ability to explore larger models or use more complex estimators, we proceed

with the analysis of the models as proposed above.

Source-differentiated models provide a way for separating sources that is

important for policy-related research, and our modification of the Yang and Koo block-

substitutability restrictions allows for estimation of large models with relatively limited

data sets. Nonetheless, improper aggregations can seriously compromise the results.

The modified assumption for block-substitutability can be

formulated in terms of restrictions on the standard LAIDS model. Hence, it

is possible to use nested tests. The test amounts to estimating the LAIDS

model without and with block-substitutability; the latter is equation (8d). In

the test average group shares are treated as constants. Since we are dealing

with cross-equation restrictions, a system test should be used, but a system

Wald test is known to have a bias towards rejecting null hypotheses in small

samples. To deal with this problem, Bewley (1986) suggests the following

correction for a small sample linear model estimated as a system of

equations:

$$(15) \quad W^* = \frac{(T-k)}{T} W,$$

where T is the number of observations, k is the average number of parameters per

equation, and W is the value of the standard Wald test statistic. If q is the number of

restrictions being tested, W^* is distributed approximately as χ_q^2 under the null

hypothesis. Table 6 summarizes the results of tests for block-substitutability,

conditional on alternative models with varying restrictions imposed. Block-

substitutability is not rejected by the data.

We also test whether the standard assumptions of homogeneity and

symmetry are supported by data (Table 7). The results indicate that homogeneity

and symmetry are not supported by data. This conclusion does not change whether

Table 6. Tests for Block-Substitutability

	Given adding-up only	Given adding-up and homogeneity	Given adding-up, homogeneity and symmetry
Wald	13.391	13.369	6.582
P-value	0.768	0.769	0.681

Table 7. Tests for Homogeneity and Symmetry

	Homogeneity		Symmetry given homogeneity	
	Wald	*P*-value	Wald	*P*-value
Block-substitutability is not imposed (LAIDS)	41.352	2.5E-07	40.666	0.0004
Block-substitutability is imposed before homogeneity and symmetry (IBLSUB)	53.913	7.9E-10	14.679	0.023

Table 8. Goodness of Fit for LAIDS and IBLSUB Models

	LAIDS[a] R^2(adj)	**IBLSUB**[b] R^2(adj)
BF1	0.54	0.37
BF2	0.63	0.59
PK1	0.16	0.22
PK2	0.75	0.67
CH1	0.60	0.46
CH2	0.82	0.67
R^{2*}	0.51	0.42

[a]Adding-up, Homogeneity and Symmetry
[b]Adding-up, Homogeneity, Symmetry, Block-Substitutability

we use the standard LAIDS model or the same model with block-substitutability imposed (IBLSUB).

Empirical Results

Goodness-of-fit statistics are provided in Table 8. As a measure of goodness of fit for the system, we use McElroy's R^{2*}. Table 9 gives estimates of parameters of the source-differentiated LAIDS model with adding-up, homogeneity, symmetry and block-substitutability imposed. Table 10 gives estimates of the corresponding expenditure and price elasticities computed under the assumption of constant group shares. Comparable estimates without imposing block-substitutability are provided in Appendix C.

The estimates of both income and compensated own-price elasticities are reasonable for most equations of the model. The demand equation for beef and pork imported from NIS countries, however, have negative expenditure elasticities, and own-price elasticities are positive or not significantly different from zero. These are the same equations which have autocorrelated and heteroscedastic residuals.

While these results may be a consequence of model misspecification, the negative expenditure elasticities can be justified. Recall, depending on the sign and the magnitude of the expenditure elasticity, commodities can be classified as luxury, necessity and inferior goods. When expenditures on the entire group of products increase, the share of luxury goods (expenditure elasticity larger than one) in this group increases faster than the share of necessity goods. The share of inferior goods decreases.

With this background in mind, the estimates of within-group expenditure elasticities reveal an interesting pattern. While beef imports from western countries have a positive expenditure elasticity exceeding unity, the imports of beef from the NIS countries have a negative expenditure elasticity. The same pattern is found in the group of pork products. In particular, beef imported from western countries and chicken meat imported from both US and Europe appear to be luxury goods

Table 9.   Estimated Parameters of the Source-differentiated LAIDS Model With Adding-up, Homogeneity, Symmetry and Block-substitutability Imposed

	INT(α)	BF1	BF2	PK1	PK2	CH1	CH2	RES	EXPβ
BF1	-0.827	-0.051	0.037	0.094	0.007	-0.080	-0.029	0.021	0.072
	-1.014	*-1.177*	*1.627*	*3.391*	*3.391*	*-6.589*	*-6.589*	*0.598*	*1.139*
BF2	2.646	0.037	0.092	0.074	0.005	-0.063	-0.023	-0.123	-0.198
	5.484	*1.627*	*3.618*	*3.391*	*3.391*	*-6.589*	*-6.589*	*-4.584*	*-5.289*
PK1	0.837	0.094	0.074	-0.156	0.026	-0.018	-0.007	-0.013	-0.049
	1.436	*3.391*	*3.391*	*-2.110*	*2.721*	*-0.856*	*-0.856*	*-0.330*	*-1.080*
PK2	0.663	0.007	0.005	0.026	0.024	-0.001	0.000	-0.060	-0.050
	5.061	*3.391*	*3.391*	*2.721*	*4.338*	*-0.856*	*-0.856*	*-5.451*	*-4.954*
CH1	-3.294	-0.080	-0.063	-0.018	-0.001	0.009	0.006	0.148	0.272
	-4.830	*-6.589*	*-6.589*	*-0.856*	*-0.856*	*0.326*	*0.462*	*7.268*	*5.138*
CH2	-0.256	-0.029	-0.023	-0.007	0,000	0.006	0.038	0.014	0.028
	-1.222	*-6.589*	*-6.589*	*-0.856*	*-0.856*	*0.462*	*2.411*	*0.870*	*1.708*
RES	1.231	0.021	-0.123	-0.013	-0.060	0.148	0.014	0.012	-0.075
	1.460	*0.598*	*-4.584*	*-0.330*	*-5.451*	*7.268*	*0.870*	*0.187*	*-1.160*

(t-statistics are italicized)

Table 10. Estimated Elasticities for the Source-differentiated LAIDS Model With Adding-up, Homogeneity, Symmetry and Block-substitutability Imposed

	Expenditure elasticities(η)	Compensated (Hicksian) own and cross-price elasticities						
		BF1	BF2	PK1	PK2	CH1	CH2	RES
BF1	1.415	-1.119	0.345	0.750	0.053	-0.228	-0.087	0.287
	3.880	*-4.503*	*2.605*	*4.674*	*4.728*	*-3.248*	*-3.455*	*1.383*
BF2	-0.527	0.462	-0.162	0.777	0.055	-0.251	-0.095	-0.786
	-1.826	*2.605*	*-0.828*	*4.613*	*4.664*	*-3.407*	*-3.604*	*-3.798*
PK1	0.761	0.632	0.489	-1.552	0.140	0.145	0.047	0.100
	3.447	*4.674*	*4.613*	*-4.319*	*3.049*	*1.388*	*1.249*	*0.533*
PK2	-2.324	0.613	0.474	1.916	0.637	0.149	0.048	-3.837
	-3.464	*4.728*	*4.664*	*3.049*	*1.703*	*1.482*	*1.338*	*-5.229*
CH1	2.160	-0.169	-0.139	0.127	0.010	-0.728	0.106	0.792
	9.565	*-3.248*	*-3.407*	*1.388*	*1.482*	*-6.412*	*1.813*	*9.145*
CH2	1.351	-0.191	-0.156	0.122	0.009	0.315	-0.438	0.339
	6.573	*-3.455*	*-3.604*	*1.249*	*1.338*	*1.813*	*-2.189*	*1.671*
RES	0.536	0.305	-0.625	0.127	-0.355	1.144	0.165	-0.762
	1.340	*1.383*	*-3.798*	*0.533*	*-5.229*	*9.145*	*1.671*	*-1.882*

(t-statistics are italicized based on the standard errors calculated under the assumption of constant budget shares)

(expenditure elasticity η=1.42, 2.16 and 1.35 respectively); pork imported from western countries and residual meats are necessity goods (η=0.76 and 0.54 respectively). All meats imported from NIS countries appear to be inferior goods (η= -2.32 and -0.53 for pork and beef respectively).

These estimates are obtained from a sample ending just before the Russian economic crisis in August 1998. We also estimated similar models using a smaller sample and obtained similar results (Soshnin). At that earlier time most economists believed that real incomes in Russia would be growing, and we thought that western exporters should be able to increase their market shares. It turned out, however, that after the crisis real earnings in Russia dropped significantly. According to our results, if this decline persists, the consequence will be a decline in market shares of western exporters and a growing market share for NIS exporters.

Turning to the cross-price effects, a positive value for compensated (Hicksian) cross-price elasticities suggest that the corresponding products are substitutes, while a negative elasticity suggests that they are complements (Deaton and Muellbauer, 1980b, p.45). In the context of trade, a positive compensated cross-price elasticity suggests that the products in question face competition from each other. Products are not competing, if their cross-price elasticity is not significantly different from zero. A negative cross-price elasticity is more difficult to explain. As far as meat products are concerned, all of them should be substitutes to a certain degree. The estimates of cross-price elasticities reported in Table 10 suggest that within each group, products imported from different sources are substitutes. Between-group cross-price elasticities suggest that beef and pork are substitutes as well, while chicken meat is a complement to beef. The latter result is difficult to justify.

Within-group cross-price elasticities are useful for evaluating the degree of competition among the exporters of the same kind of good. The estimates of within-group cross-price elasticities are positive and significant for all groups of

meat products. This suggests that importers are competing in their respective markets.

To summarize, with the exception of complementarity between beef and chicken meat, the results appear plausible. Imports of beef and pork from the NIS countries have negative or zero expenditure elasticities, which is a disadvantage at times of economic growth, but perhaps favorable when real incomes of consumers are declining. It is also apparent that the patterns of Russia's trade with the NIS countries are different from those with the western countries. Hence, when estimating demand models, aggregation over sources may introduce significant bias to the estimates of elasticities. This is particularly relevant for those meat groups with a considerable market share of NIS importers (beef for example). While this interpretation of findings is quite plausible, we remind readers that the estimated elasticities could be affected by the poor quality of data, the nature of aggregations by source, and by incorrect dynamic specification of the model. Certainly, room for additional research exists.

Appendix A

AIDS, LAIDS, and Block Substitutability

The standard AIDS model (Deaton and Muellbauer, 1980a,b) can be written in the form of share (w_i) equations:

(A1) $w_i = \alpha_i + \sum_j \gamma_{ij} \ln(p_j) + \beta_i \ln\left(\dfrac{E}{P*}\right)$, where

(A2) $\ln(P*) = \alpha_0 + \sum_l \alpha_l \ln(p_l) + \dfrac{1}{2} \sum_l \sum_m \gamma_{lm} \ln(p_l) \ln(p_m)$,

where subscripts i and j refer to the commodities under consideration. E represents total expenditures on these commodities; p_j is the price of the jth commodity; α, β and γ are parameters.

The non-linear price index $P*$ defined in (A2) can be approximated by the linear Stone price index (Deaton and Muellbauer, 1980b), which geometrically weights prices of the products under consideration. Since the weights used in calculating Stone index appear on the LHS of equations (A1), simultaneity may be a problem. To reduce its effect, we follow Haden (1990) in using average weights for calculating the Stone price index. Thus, (A2) becomes

(A3) $\ln(P*) \approx \sum_i \overline{w}_i \ln(p_i)$, where $\overline{w}_i$ is the average market share of ith commodity

defined as: $\overline{w}_i = \dfrac{1}{T} \sum_{t=1}^{T} w_{it}$, and T is the number of observations.

With (A3) replacing (A2), AIDS is referred to as a Linearized Almost Ideal Demand System (LAIDS). The computing formulae for LAIDS elasticities derived under the assumption of constant market shares are due to Chalfant:

(A4) $\eta_i = 1 + \dfrac{\beta_i}{w_i}$ for expenditure elasticity,

(A5) $\varepsilon_{ij}^u = -\delta_{ij} + \dfrac{\gamma_{ij}}{w_i} - \beta_i\left(\dfrac{w_j}{w_i}\right)$ for uncompensated (Marshallian) elasticities,

(A6) $\varepsilon_{ij}^c = -\delta_{ij} + \dfrac{\gamma_{ij}}{w_i} + w_j$ for compensated (Hicksian) elasticities.

Here $\delta_{ij} = \begin{cases} 0 \text{ if } i = j \\ 1 \text{ otherwise} \end{cases}$ is Kronecker delta.

The theoretical restrictions can be imposed on the demand system in the following way (Deaton and Muellbauer, 1980b):

(A7) $\sum_i \alpha_i = 1, \quad \sum_i \beta_i = 0, \quad \sum_i \gamma_{ij} = 0, \ \forall j;$ (adding-up)

(A8) $\sum_j \gamma_{ij} = 0, \ \forall i;$ (homogeneity)

(A9) $\gamma_{ij} = \gamma_{ji}, \forall i, j;$ (symmetry).

The assumption of block-substitutability, derived in the text, can be written as: $\gamma_{ihjk} = \dfrac{\overline{s}_{ih}\overline{s}_{jk}}{\overline{s}_{ir}\overline{s}_{jq}} \gamma_{irjq}, \quad \forall j \neq i$. To illustrate the process of imposing the block-substitutability restrictions, we assume that there are three groups of meat products in the model: beef, pork and chicken. All other meats are aggregated and form a residual product called "other meats". Further, there are exactly two sources of imports in each group except "other meats". For the purpose of illustration we take a close look at three equations: two beef sources and one pork source.

Letting subscripts B, P, C and R denote beef, pork, chicken and the residual group respectively, and letting subscripts 1 and 2 denote sources of imports within each group (sources do not have to be the same for different groups), the three unrestricted equations of interest are:

(A10) $w_{B1} = \alpha_{B1} + \gamma_{B1B1} \ln(p_{B1}) + \gamma_{B1B2} \ln(p_{B2}) + \gamma_{B1P1} \ln(p_{P1}) + \gamma_{B1P2} \ln(p_{P2})$
$$+ \gamma_{B1C1} \ln(p_{C1}) + \gamma_{B1C2} \ln(p_{C2}) + \gamma_{B1R} \ln(p_R) + \beta_{B1} \ln\left(\frac{E}{P}\right),$$

(A11) $w_{B2} = \alpha_{B2} + \gamma_{B2B1} \ln(p_{B1}) + \gamma_{B2B2} \ln(p_{B2}) + \gamma_{B2P1} \ln(p_{P1}) + \gamma_{B2P2} \ln(p_{P2})$
$$+ \gamma_{B2C1} \ln(p_{C1}) + \gamma_{B2C2} \ln(p_{C2}) + \gamma_{B2R} \ln(p_R) + \beta_{B2} \ln\left(\frac{E}{P}\right),$$

(A12) $w_{P1} = \alpha_{P1} + \gamma_{P1B1} \ln(p_{B1}) + \gamma_{P1B2} \ln(p_{B2}) + \gamma_{P1P1} \ln(p_{P1}) + \gamma_{P1P2} \ln(p_{P2})$
$$+ \gamma_{P1C1} \ln(p_{C1}) + \gamma_{P1C2} \ln(p_{C2}) + \gamma_{P1R} \ln(p_R) + \beta_{P1} \ln\left(\frac{E}{P}\right).$$

The same equations after imposing block substitutability, (8d), are given below. Parameters in the square brackets are the ones, which do not have to be estimated due to block-substitutability. Since these restrictions affect only cross-price parameters and the cross-price parameter of the residual group is already restricted by homogeneity, the irrelevant parameters are omitted in the following to preserve transparency:

$$(A13) \quad w_{B1} = \ldots + \gamma_{B1B1}\ln(p_{B1}) + \gamma_{B1B2}\ln(p_{B2}) + \gamma_{B1P1}\ln(p_{P1})$$
$$+ \left[\frac{\bar{s}_{P2}}{\bar{s}_{P1}}\gamma_{B1P1}\right]\ln(p_{P2}) + \gamma_{B1C1}\ln(p_{C1}) + \left[\frac{\bar{s}_{C2}}{\bar{s}_{C1}}\gamma_{B1C1}\right]\ln(p_{C2}) + \ldots,$$

$$(A14) \quad w_{B2} = \ldots + \gamma_{B2B1}\ln(p_{B1}) + \gamma_{B2B2}\ln(p_{B2}) + \left[\frac{\bar{s}_{B2}}{\bar{s}_{B1}}\gamma_{B1P1}\right]\ln(p_{P1})$$

$$+ \left[\frac{\bar{s}_{B2}\bar{s}_{P2}}{\bar{s}_{B1}\bar{s}_{P1}}\gamma_{B1P1}\right]\ln(p_{P2}) + \left[\frac{\bar{s}_{B2}}{\bar{s}_{B1}}\gamma_{B1C1}\right]\ln(p_{C1}) + \left[\frac{\bar{s}_{C2}}{\bar{s}_{C1}}\gamma_{B1C1}\right]\ln(p_{C2}) + \ldots,$$

$$(A15) \quad w_{P1} = \ldots + \gamma_{P1B1}\ln(p_{B1}) + \left[\frac{\bar{s}_{B2}}{\bar{s}_{B1}}\gamma_{P1B1}\right]\ln(p_{B2}) + \gamma_{P1P1}\ln(p_{P1})$$

$$+ \gamma_{P1P2}\ln(p_{P2}) + \gamma_{P1C1}\ln(p_{C1}) + \left[\frac{\bar{s}_{C2}}{\bar{s}_{C1}}\gamma_{P1C1}\right]\ln(p_{C2}) + \ldots.$$

Now to impose within-group symmetry on the beef products, we set $\gamma_{B2B1} = \gamma_{B1B2}$. To impose symmetry between the groups of pork and beef products, we set $\gamma_{P1B1} = \gamma_{B1P1}$ everywhere in (A15)[3]. Other symmetry restrictions can be imposed in a similar manner. Note, symmetry also could be imposed before block-substitutability, which is probably more convenient.

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[3] $\gamma_{ihjk} = \dfrac{\bar{s}_{ih}\bar{s}_{jk}}{\bar{s}_{jk}\bar{s}_{ih}}\gamma_{jkih} = \gamma_{jkih}$

Appendix B

Specification tests

Equations (7) from the text can be written in stochastic form as follows[4]:

(B1) $w_{it} = f_i(Z_t; \Theta) + \varepsilon_{it}, \quad i=1, 2, ..., m$ and $t=1,2, ..., T,$

where $m = N-1$ is the number of equations to be fitted in the model and T is the number of available observations; Z is a matrix of independent variables and Θ is a matrix of parameters to be estimated; f_i is a function of Z and Θ, and ε_{it} is the vector of contemporaneous errors. In this context, we test for autocorrelation, heteroscedasticity, autoregressive conditional heteroscedasticity and exogeneity of independent variables (in particular, exogeneity of group expenditures, as it often appears to be a problem).

When estimating a system of equations, a common practice is to run the tests separately for each equation of the system. Here, however, we use procedures similar to those described in Edgerton, et al. They argue, that in Godfrey's (1988) framework, system testing is quite simple, while single equation techniques may be misleading.

Test for autocorrelation. The modified version of the Breusch-Godfrey (BG) test for autocorrelation can be written as follows:

(B2) $\hat{\varepsilon}_{it} = f_i(x_t, \Theta) + \sum_{j=1}^{m} \delta_{ij} \hat{\varepsilon}_{j,t-1} + v_{it}, \quad$ for $i=1,...,m$

where m is the number of equations in the system (excluding the residual equation), and the test is for first-order autocorrelation The null hypothesis of no autocorrelation is $H_0 : \delta_{ij} = 0,$ for all j in each equation i. The residuals for the test are those computed from the LAIDS model with adding-up, homogeneity, symmetry, and block-substitutability imposed.

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[4] In order to keep notation simple, single subscripts are used whenever possible.

Test for heteroscedasticity and autoregressive conditional heteroscedasticity.

Due to lack of degrees of freedom, we do not use the standard White's heteroscedasticity test, which requires regressing squared residuals on all explanatory variables, their squares and cross-products. Instead, following Edgerton et al., who take Bickel's (1978) approach, possible heteroscedasticity is modeled by regressing the squared residuals on the predictions of budget shares from the same model:

$$(B3) \quad \hat{\varepsilon}_{it}^2 = \alpha_i + \sum_{j=1}^{m} \delta_{ij} \hat{w}_{jt} + \upsilon_{it}, \quad \text{for } i=1,...,m.$$

The null hypothesis of no heteroscedasticity is $H_0 : \delta_{ij} = 0,$ for all j in each equation i.

Another possibility is to test for Autoregressive Conditional Heteroscedasticity (ARCH). This test involves regressing squared residuals on their lags and an intercept. According to Edgerton et al., this procedure can also be generalized for system testing:

$$(B4) \quad \hat{\varepsilon}_{it}^2 = \alpha_i + \sum_{j=1}^{m} \delta_{ij} \hat{\varepsilon}_{jt-1}^2 + \upsilon_{it}, \quad \text{for } i=1,...,m.$$

The null hypothesis of no ARCH is: $H_0 : \delta_{ij} = 0,$ for all j in each equation i.

Test for endogeneity. We use a modification of the Hausman-Wu (HW) test for endogeneity. In general, Hausman tests check for contemporaneous correlation between regressors and the error. This test involves finding an appropriate instrument for each regressor that is thought to be contemporaneously correlated with the error term, and checking whether this instrument has any explanatory power. The appropriate candidate for such an instrument must be a variable, which is contemporaneously uncorrelated with the error and highly correlated with the variable, for which it serves as an instrument. To ensure that these requirements are satisfied, an instrument can be constructed in the following way. Let Z_1 be a set of RHS variables, which we want to test for endogeneity, Z_2 be a set of RHS variables

which we *a priori* assume to be exogenous, and Q_1 be a set of appropriate instruments. The orthogonal projection of Z_1 on Z_2 and Q_1 then, is the best instrument for Z_1.

As a candidate for Z_1 we consider here only expenditures[5]. As a set of instruments (Q_1) for the expenditures on meat, we use the consumer price index (*CPI*) and real per capita income (*Y*) calculated in dollar terms. Then, the two stages of the HW test can be formalized as follows:

$$(B5) \qquad \ln\!\left(\frac{E_t}{P_t}\right) = \alpha + \sum_{i=1}^{m+1} \gamma_i \ln p_{it} + \beta_1 \ln\!\left(\frac{CPI_t}{P_t}\right) + \beta_2 \ln\!\left(\frac{Y_t}{P_t}\right) + v_{it}, \text{ (first stage)}$$

where $m+1$ is the number of products considered in the model (including residual product), P is the linearized Stone index and v_{it} is the classical error term.

$$(B5) \qquad w_{it} = f_i(x_t, \Theta) + \delta_i \hat{x} + v_{it}, \quad \text{for } i=1,...,m \text{ (second stage)}$$

where $\hat{x}$ is a 1-by-*T* vector of predicted values of the dependent variable in (14a). The null hypothesis that expenditures are exogenous is: $H_0 : \delta_i = 0,$ for each equation *i*.

Since all hypotheses discussed above can be formulated as functions of estimated parameters such as $g(\beta)=0$, it is convenient to use Wald (*W*) test statistic, readily available in the TSP package. As noted in the text, the problem with this statistic is that, when used for joint system tests, it is known to have a considerable bias towards rejection of null hypothesis in small samples. Thus, tests are conducted separately for each equation of a system. When testing for endogeneity a standard *t*-statistic is used instead of a Wald statistics.

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[5] Edgerton et al. points out that prices may be endogenous as well, though this problem is usually dismissed by the majority of authors, who either try to provide an *a priori* justification of the opposite, or ignore the problem altogether. One common justification, which is used in domestic food demand studies, is that food prices are likely to be regulated or otherwise determined in the world market. This explanation extends to the case of import demand.

Appendix C

Table C1. Estimates of parameters of the source-differentiated LAIDS model without block-substitutability (adding-up, homogeneity and symmetry imposed)

	INT(α)	BF1	BF2	PK1	PK2	CH1	CH2	RES	EXP(β)
BF1	-1.561	-0.161	0.010	0.128	-0.017	0.001	-0.081	0.120	0.129
	-2.073	*-3.047*	*0.345*	*2.353*	*-1.466*	*0.034*	*-4.824*	*2.438*	*2.205*
BF2	2.653	0.010	0.089	0.068	0.016	-0.053	-0.030	-0.100	-0.198
	5.444	*0.345*	*2.920*	*1.712*	*1.359*	*-2.270*	*-2.236*	*-3.080*	*-5.246*
PK1	0.541	0.128	0.068	-0.265	0.053	0.019	0.028	-0.032	-0.023
	0.823	*2.353*	*1.712*	*-2.738*	*2.718*	*0.438*	*0.987*	*-0.509*	*-0.444*
PK2	0.761	-0.017	0.016	0.053	0.015	0.000	-0.002	-0.066	-0.058
	5.582	*-1.466*	*1.359*	*2.718*	*1.453*	*-0.008*	*-0.228*	*-4.018*	*-5.533*
CH1	-2.679	0.001	-0.053	0.019	0.000	-0.047	0.011	0.069	0.223
	-4.139	*0.034*	*-2.270*	*0.438*	*-0.008*	*-1.289*	*0.801*	*1.773*	*4.429*
CH2	-0.298	-0.081	-0.030	0.028	-0.002	0.011	0.036	0.037	0.031
	-1.518	*-4.824*	*-2.236*	*0.987*	*-0.228*	*0.801*	*2.221*	*1.525*	*2.043*
RES	1.584	0.120	-0.100	-0.032	-0.066	0.069	0.037	-0.027	-0.104
	2.236	*2.438*	*-3.080*	*-0.509*	*-4.018*	*1.773*	*1.525*	*-0.376*	*-1.918*

 (t-statistics are italicized)

Table C2. Estimates of elasticities for the source-differentiated LAIDS model without block-substitutability (adding-up, homogeneity and symmetry imposed)

	Expenditure elasticities(η)	Compensated (Hicksian) own and cross-price elasticities						
		BF1	BF2	PK1	PK2	CH1	CH2	RES
BF1	1.744	-1.754	0.187	0.946	-0.083	0.242	-0.390	0.852
	5.168	*-5.763*	*1.121*	*3.008*	*-1.241*	*1.161*	*-4.010*	*3.013*
BF2	-0.533	0.250	-0.186	0.734	0.142	-0.173	-0.153	-0.614
	-1.825	*1.121*	*-0.795*	*2.379*	*1.520*	*-0.964*	*-1.473*	*-2.435*
PK1	0.891	0.797	0.461	-2.082	0.274	0.325	0.217	0.008
	3.609	*3.008*	*2.379*	*-4.427*	*2.876*	*1.573*	*1.554*	*0.026*
PK2	-2.826	-0.957	1.222	3.752	0.000	0.230	-0.031	-4.215
	-4.087	*-1.241*	*1.520*	*2.876*	*0.000*	*0.367*	*-0.065*	*-3.869*
CH1	1.951	0.178	-0.095	0.285	0.015	-0.966	0.127	0.456
	9.086	*1.161*	*-0.964*	*1.573*	*0.367*	*-6.211*	*2.126*	*2.756*
CH2	1.388	-0.854	-0.250	0.564	-0.006	0.376	-0.463	0.632
	7.313	*-4.010*	*-1.473*	*1.554*	*-0.065*	*2.126*	*-2.243*	*2.053*
RES	0.357	0.909	-0.488	0.010	-0.390	0.658	0.307	-1.006
	1.066	*3.013*	*-2.435*	*0.026*	*-3.869*	*2.756*	*2.053*	*-2.247*

(t-statistics are italicized based on the standard errors calculated under the assumption of constant budget shares)

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