An analysis of food demand and household food security in CEE: evidence from Slovakia

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Summary

We investigate the food security situation of Slovak households in terms of both access to food and quality of the diet consumed by estimating food demand system and diet diversity demand models using household budget survey data over the period 2004-2010. In most samples demand for meat and fish and fruits and vegetables is expenditure and own-price elastic. On average all five food groups investigated are found to be normal goods. Rural and low-income households appear more expenditure and price sensitive compared to the urban and high-income ones. Results from quantile regressions indicate that income has a positive while uncertainty has a negatively effect on the diversity of the diet as the effects are stronger in more vulnerable, low income and rural consumer subsamples. Overall the food security situation in Slovakia appears to have improved over time, since the country's EU accession.

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1. INTRODUCTION

Food security is an important dimension of household well-being. Therefore, food demand has been actively researched for over a century both in developed and developing countries as the focus has usually been on how income and prices influence household food expenditure and consumption patterns. Policy makers dealing with food security issues are often interested in studies that examine the response of households to price and income changes. While predominantly food demand analyses have been concerned with situations in developing countries, there are also several food demand studies employing household data from developed European countries (e.g., Molina, 1994 for Spain;Banks et al., 1996; 1997 for the UK; Moro and Sckokai, 2000 for Italy; Abdulai, 2002 for Switzerland). However, food demand responses in the middle-income former socialist countries, now new member states of the European Union (EU), have not been widely studied with micro data.[1] As under-nutrition and malnutrition exist to a considerable degree in both developed countries and developing and transition countries a study of the food security situation in the EU new member states (NMS) is timely.[2]

Food supply and demand in Europe have been importantly influenced by the Common Agricultural Policy (CAP), which is driven by the EU's commitment to support long-term food supply and meet the European and growing world food demand (European Commission, 2010). As a result of CAP and rising incomes the share of European household expenditure on food has been steadily declining over the years. However, international food prices have recently risen and are likely to remain high primarily because of the escalating cost of inputs and surging world demand. In 2005, a year after the accession of the first wave of NMS, food expenditure in the EU was between 10% and 35% of total household consumption budget, with the smallest shares in the EU-15 and the largest in the NMS (EEA, 2005). Consequently, the price index for food in the EU rose by almost 20% between 2005 and 2012 (Eurostat, 2012). Rising food prices create serious difficulties, especially for vulnerable, low-income households that spend a substantial proportion of their income on food.

Because a large number of vulnerable households are located in the NMSs, this paper aims at shedding light on the food security situation of households in Slovakia, a middle-income east European NMS with well performing economy, and the lowest income inequality in the EU (Eurostat, 2013); thus findings from Slovakia can be considered a upper bound of the indicators for the food security situation in the NMS. Documenting and understanding food security outcomes is useful for several reasons: to identify the food-insecure, characterize the nature of their insecurity (seasonal versus chronic), monitor changes in their

[1] Exceptions are studies by Janda et al. (2009) who estimate a complete demand system using Czech household budget survey data and Moon et al. (2002) who study the demand for food variety in Bulgaria; there are also a few partial demand analysis on selected food groups (e.g., Hupkova et al., 2009 and Zetkova and Hoskova, 2009 for Slovakia; Szigeti and Podruzsik, 2011 for Hungary).

[2] In Europe, about 5% of the overall population is at risk of malnutrition, and among vulnerable groups—the poor, the elderly, and the sick—this percentage is even higher (Reisch et al., 2013). In the NMS malnutrition and general poverty is the highest; for instance, in 2011, poverty rate ranged between 20% in Slovakia and 40% in Romania as poverty rates considerably differ between urban and rural areas and across income groups.

circumstances, and assess the impact of potential interventions. According to USAID (1992) there are two main dimensions to the definition of food security: access(conditional on availability) and utilization (whether a population will be able to derive sufficient and balanced nutrition during a given period).

As a first stage of our analytical framework we follow Banks et al. (1997) and employ the Quadratic Almost Ideal Demand System (QUAIDS) augmented with demographic and other controls to examine the household food demand patterns, and thus availability and access to food, across income groups and types of region. An important contribution of the paper is the combination of using extended QUAIDS methodology and household ``longitudinal data from Slovakia. Compared to other demand systems, QUAIDS is more appropriate since it allows for non-linearity in the Engel curves which is commonly the case when analysing aggregate commodity food demand system at household level. The fact that we use household (micro) data is important because managing food security requires not only understanding how policies influence the availability of food and income at national level but also how individual households can cope with income and price shocks. Furthermore, as a second stage of our framework we analyse household diet diversity demand functions, which provide information on food utilisation. We apply both OLS and quantile regressions, capturing the heterogeneity in behaviour across subsamples.

Our analysis of Slovak household demand patterns suggests that food security situation has improved since Slovakia's EU accession. However, food commodities important for healthy diet such as meat and fish and fruits and vegetables remain expenditure and own-price elastic. In terms of diet diversity, economic uncertainty importantly impacts, especially, low income households. There also is important heterogeneity in sensitivity to income and price shocks across subsamples of rural and urban and low- and high-income households that need to be taken into account by policy-makers. The rest of the paper consists of methodology, data, and results sections and a conclusion.

2. METHODOLOGY

Within the food security analysis framework, there is an association between food access and diet diversity at household level. The magnitude of the association increases with improving the food access; for example, Jackson (1984) shows that diet diversity measured as the number of food commodities consumed increase with income and expenditure and Hoddinott and Johannes (2002) demonstrate a link between the mean level of caloric availability and diet diversity. Therefore, our analysis of food security proceeds in two stages; first, we analyse access to food by the means of a demand system (QUAIDS) and second, we set up a framework for diet diversity analysis.[3] Taken together the two stages generate results capable of qualifying the food security situation of Slovak households in terms of both access to food and quality of diet.

2.1. Quadratic Almost Ideal Demand System

Several demand systems have been popular for modelling the allocation of total expenditures among commodities given certain budget. These include the Linear Expenditure System (LES) (Stone, 1954), the Rotterdam model (Barten 1964), the Indirect Translog System (ITS) (Christensen et al., 1975), and the Almost Ideal Demand System (AIDS) (Deaton and Muellbauer, 1980). LES is unable to describe demand behaviour consistent with the Engel's law where as income increases a good can change from normal to

[3] Nutrition science experts argue that as global food supply system is facing serious challenges from economic crises (and climate change), there are increasing constraints to the nutritional well-being of the populations, especially the poor. To cope, vulnerable populations prioritise consumption of calorie-rich but nutrient-poor food. Consequently, dietary quality and eventually quantity decline, increasing micronutrient malnutrition (or hidden hunger) and exacerbating pre-existing vulnerabilities that lead to poorer health, lower incomes, and reduced physical and intellectual capabilities (e.g., Bloem et al., 2010). In this context diet diversity is shown to be an important indicator of quality the diet (Drescher et al., 2007; Brinkman et al., 2010; Thorne-Lyman et al., 2010; Iannotti et al., 2012).

inferior one. The Rotterdam model is consistent with demand theory; however, since it is not derived from specific utility or expenditure function, the model is inconsistent with utility maximising behaviour. ITS has the advantage of a flexible functional form but poses a major estimation problem due to relatively large number of independent parameters. AIDS satisfies the restrictions of demand theory and its estimation is less complicated than other models.

Based on non-parametric analysis of consumer expenditure patterns Banks et al. (1996; 1997) show that the correct approximation of Engel curves requires a higher order logarithmic term of expenditure and propose QUAIDS which nests AIDS and also satisfies the restrictions of demand theory.[4] QUAIDS thus allows as income increases a good to change from normal to inferior one. Household preferences follow the indirect utility function:

$$\ln V = \left\{\left[\frac{\ln m - \ln a(\mathrm{p})}{b(\mathrm{p})}\right]^{-1} + \lambda(\mathrm{p})\right\}^{-1}, \tag{1}$$

where the term [ln$m$ - ln$a$(p)]/b(p) is the indirect utility function of the PIGLOG[5] demand system, m is household income, and a(p), b(p)and λ(p)are functions of the vector of prices p. To ensure the homogeneity property of the indirect utility function, it is required that a(p)is homogenous of degree one in p, and b(p)and λ(p)arehomogenous of degree zero in p. The price index ln$a$(p) has the usual translog form

$$\ln a(\mathrm{p}) = \alpha_0 + \sum_j \alpha_j \ln p_j + \frac{1}{2}\sum_i \sum_j \gamma_{ij} \ln p_i \ln p_j,$$

b(p) is a simple Cobb-Douglas price aggregator defined as

$$b(\mathrm{p}) = \prod_i \ p_i^{\beta_i},$$

and λ(p) is defined as

$$\lambda(\mathrm{p}) = \sum_i \ \lambda_i \ln p_i, \text{ where } \sum_i \ \lambda_\mathrm{i} = 0.$$

By applying Roy's identity to the indirect utility function,Equation (1), the budget shares in the QUAIDS are derived as

$$\omega_i = \alpha_i + \sum_j \ \gamma_{ij} \ln p_j + \beta_i \ln\left[\frac{m}{a(\mathrm{p})}\right] + \frac{\lambda_i}{b(\mathrm{p})}\left\{\ln\left[\frac{m}{a(\mathrm{p})}\right]\right\}^2. \tag{2}$$

For theoretical consistency and to reduce the number of parameters to be estimated adding-up, homogeneity and symmetry restrictions are commonly imposed. The fact that $\sum_i \ \omega_i = 1$, called the adding-up condition, requires that $\sum_i \ \alpha_i = 1, \sum_i \ \beta_i = 0$, $\sum_i \ \lambda_i = 0$ and $\sum_i \ \gamma_{ij} = 0 \ \forall j$. Moreover, since demand functions are homogeneous of degree zero in (p, m), $\sum_j \ \gamma_{ij} = 0 \ \forall j$. And Slutsky symmetry implies that $\gamma_{ij} = \gamma_{ji} \forall i \neq j$. These conditions are trivially satisfied for a model with n goods when the estimation is carried out on a subset of n - 1 independent equations. The parameters of the dropped equation are then computed from the restrictions and the estimated parameters of the n - 1 expenditure shares.

Majority of previous studies extend the system with demographic variables following Pollak and Wales (1981) where the demographic effects shift the intercept α_i in equation (2). However, we follow the scaling approach introduced by Ray (1983) which has been implemented by Poi (2012) into QUAIDS. This approach has the advantage of having strong theoretical foundations and generating expenditure share equations that closely mimic their counterparts without demographics. For each household the expenditure function $e(\mathrm{p}, \mathrm{z}, u)$, underlying the budget shares is written as the expenditure function of a reference household $e^R(\mathrm{p}, u)$, scaled by the function $m_0(\mathrm{p}, \mathrm{z}, u) = \overline{m}_0(\mathrm{z})\varphi(\mathrm{p}, \mathrm{z}, u)$ to account for the household characteristics where zrepresents a vector of s characteristics and u is direct utility. The first term of m_0,($\overline{m}_0$(z)) measures the increase in a household's expenditures as a function of z, not controlling for any

[4] Because usually data on food demand are presented as aggregates across commodities, the commodity group Engel curve will depend on the income levels at which commodities in the group enter the budget, and Jackson (1984) shows that the expenditure share on the group need not be monotonic. This suggests that flexible functional forms (Blaylock and Smallwood, 1982), such as QUAIDS can be an important tool for analysing aggregate commodity group Engel curves, and in demand analysis generally.

[5] Demand with expenditure shares that are linear in log total expenditure alone have been referred to as Price-Independent Generalised Logarithmic (PIGLOG) by Muellbauer(1976).

differences in consumption patterns. The second term ($\varphi(\mathrm{p},\mathrm{z},u)$)controls for differences in relative prices and the actual goods consumed.For example, a household with two adults and two infants will consume different goods than one comprising four adults.

Furthermore, we extend vector z with a food expenditure control the rationale for which is the following. In estimating a food demand system the implicit assumption is that the consumer's utility maximisation decision can be decomposed into two separate stages where in the first stage, the allocation of total expenditure between food and other commodity groups (housing, transport, entertainment, etc.) is decided. In the second stage, the food expenditure is allocated among different food groups.[6] The price and expenditure elasticities obtained from such atwo-stage budgeting process are conditional elasticities in the sense that a second-stage conditional demand system is estimated. To obtain unconditional elasticity estimates correction for the first stage budgeting decision is needed.[7]Given data limitation and the fact that structured two-stage budget allocation offers only an approximation under restrictive conditions, we opt for a reduced form single-stage specification, where besides standard demographic variables, the share of food expenditure in the total household budget is also added to vector z. The addition of the share of food expenditure in the total budget as a control in the budget share equation offers an alternative approximation of the budgeting process that is consistent with the weak separability assumption and its implications for the Slutsky substitution term (Okren and Alston, 2011, p.12).

The budget share equation (2) augmented with vector z becomes:

$$\omega_i = \alpha_i + \sum_j \gamma_{ij} \ln p_j + (\beta_i + \eta'_i \mathrm{z}) \ln \left[\frac{m}{\bar{m}_0(\mathrm{z}) a(\mathrm{p})}\right] + \frac{\lambda_i}{b(\mathrm{p}) c(\mathrm{p},\mathrm{z})} \left\{\ln \left[\frac{m}{\bar{m}_0(\mathrm{z}) a(\mathrm{p})}\right]\right\}^2, \quad (3)$$

where $c(\mathrm{p},\mathrm{z}) = \prod_j p_j^{\eta'_j z}$, η'_jrepresents the j^{th} column of parameter matrix η. The adding-up condition requires that $\sum_j \eta_{sj} = 0 \ \forall s$.

Following Banks et al. (1997) the expenditure and price elasticities are obtained by partially differentiating Equation (3) with respect to ln$m$ and lnp_j respectively:

$$\mu_i \equiv \frac{\partial \omega_i}{\partial \ln m} = \beta_i + \eta'_i \mathrm{z} + \frac{2\lambda_i}{b(\mathrm{p}) c(\mathrm{p},\mathrm{z})} \ln \left[\frac{m}{\bar{m}_0(\mathrm{z}) a(\mathrm{p})}\right] \text{and} \quad (4)$$

$$\mu_{ij} \equiv \frac{\partial \omega_i}{\partial \ln p_j} = \gamma_{ij} - \mu_i \left(\alpha_j + \sum_k \gamma_{jk} \ln p_k\right) - \frac{\lambda_i (\beta_j + \eta'_j \mathrm{z})}{b(\mathrm{p}) c(\mathrm{p},\mathrm{z})} \left\{\ln \left[\frac{m}{\bar{m}_0(\mathrm{z}) a(\mathrm{p})}\right]\right\}^2. \quad (5)$$

Then the expenditure and the uncompensated price elasticities are computed as $e_i = \mu_i / \omega_i + 1$ and $e_{ij}^u = \mu_{ij} / \omega_i - \delta_{ij}$ respectively; δ_{ij} represents Kronecker delta taking value 1 if i=j and 0 otherwise. Using the Slutsky equation, we can finally compute the compensated price elasticities: $e_{ij}^c = e_{ij}^u + e_i \omega_j$.

2.2. Diet diversity under uncertainty

It is established in the literature that as incomes increase consumers tend to increase not only the quantity but also the number of goods consumed (Theil and Finke, 1983; Jackson, 1984). Following Jackson (1984), we specify a (expected) utility function $u(q)$ defined for any vector of quantities q in some food commodity set N

$$u(q) = u(q_1, q_2, \dots, q_n). \quad (6)$$

[6] The assumption about (weak) separability of the food expenditure decision from other expenditure choices can be motivated by Maslow's (1943) hierarchy of needs theory where substitubaility between goods in different groups is limited. Furthermore, we note that assuming weak separability leads to modelling the demand for food commodities as a function of food expenditure, rather than total expenditure (von Haefen, 2000; Okrent and Alston, 2011).

[7]There are problems with the first stage allocation since it is not possible to replace the prices of the goods in a group with a single price index without imposing restrictive conditions (Gorman, 1959). Michalek and Keyzer (1992) and Edgerton (1997) show that under weak separability of preferences and price index for each group that is not too sensitive to changes in the utility function, the two-stage budgeting process leads to *approximately* correct budget allocation.

The utility function is maximised subject to budget constraint, $\sum p_i\, q_{i,} = m$ and non-negativity constraints $q_i \geq 0$ where p_i is the price for the i^{th} food commodity and m is income. The following Kuhn-Tucker conditions should be satisfied

$$\frac{\partial u}{\partial q_i} - \lambda p_i = 0\ if\ i \epsilon S\ ,\ q_i > 0 \text{ and} \quad (7)$$

$$\frac{\partial u}{\partial q_i} - \lambda p_i < 0\ if\ i \epsilon \bar{S}\ ,\ q_i = 0, \quad (8)$$

where λ is the Lagrangian multiplier, S is the set of commodities purchased, and $\bar{S}$is the set of commodities not purchased; thus, in cardinality notation $|N| = |S| + |\bar{S}|$. The above conditions lead to the following (Marshallian) food demand function

$$q_i = q_i(p', m), \quad (9)$$

wherep' is a vector of food prices.

An important result of Jackson's (1984) analysis is that the number of food commodities in set S is also a function of food prices and income (food expenditures). Let $s_h = |S|$denotes the number of different food commodities consumed by household $h$which is a measure of diet diversity (D) at household level (e.g., Jackson, 1984; Stewart and Harris, 2005). Then $D_h = s_h$ is a function of food prices and food expenditures, i.e.

$$D_h = s_h = f_h(p', m_h), \quad (10)$$

where m_h is total household disposable income and $f_h$is household specific diet diversity function which accounts for the household characteristics and circumstances affecting diet choices.

The count of food items consumed is one measure of diet diversity but there are alternative ways of measuring diversity. A measure, which has become popular in the diet diversity economics literature (e.g., Thiele and Weiss, 2003; Drescher and Goddard, 2011; Hertzfeld et al., 2014) is the Berry index (Berry, 1971), $BI = 1 - \sum \omega_i^2$, where ω_i is the budget share of the i^{th} (disaggregate) food commodity specified in a manner similar toEquation (3).[8] It thus follows that this measure of diet diversity is also a function of food prices, income (expenditure), and household characteristics

$$D_h = BI_h = f_h(p', m_h). \quad (11)$$

Given the focus of the paper on food security, the analysis of diet diversity needs to be linked to decision making under uncertainty. Looking into implications of uncertainty for the dietary choices and quality of diet of risk-averse households is consistent with the demand analysis in the previous section. There the estimated expenditure and price elasticities measure the sensitivity of households to market shocks and thus provide insight into the access of households to food in uncertain market environment. Therefore, it is only logical to also ask what the impact of uncertainty on household diet diversitychoices would be.

Our starting point in answering the question is the neoclassical economics framework for decision making under uncertainty where concavity of the expected utility function is equivalent to consumer's (household's) risk aversion. The more concave the expected utility function the more risk averse the consumer - a property captured by the well-known Arrow-Pratt measure of absolute risk aversion:

$$r(q) = -\frac{u''(q)}{u'(q)}, \quad (12)$$

where$u'(q)$ and $u''(q)$ are the first and second derivative respectively of the utility function. The interpretation of $r(q)$ is that a consumeris more risk averse the larger the value of $r(q)$is and that she/he is less willing to accept a (small) gamble on the amount of her/his consumption. For example, if consumer has monetary income M (consumption is an increasing function of income, i.e., $q=q(M)$) and there is some probability π that she/he will lose an amount of income L in the future, a risk averse consumer will want to

[8] The Berry index formulation implies that diversity is higher when more foods are eaten in equal (quantity or expenditure) proportions such that a higher value of the index indicates a more balanced diet. The Berry index is also known as the Simpson index (Stewart and Harris, 2005) and is closely related to the well-known Hirschman-Herfindahlindex (Theil and Finke, 1983).

purchase insurance in order to avoid the (potential) loss, thus forgoing some consumption at present.[9] It is established in the literature that absolute risk aversion decreases with income (wealth), i.e., as consumers become wealthier they are willing to accept more (monetary) gambles (Pratt, 1964).

Furthermore, the Pratt's theorem formulates the conditions under which one consumer can be said to be more risk averse than another for all levels of wealth. Thus, if consumer A is more risk averse than consumer B then A would be willing to pay more to avoid a given risk than B would. Each consumer's risk premium is defined by the condition that the expected utility of a risky income with no insurance should be equal to the utility of the expected income minus the risk (insurance) premium. For small variation in income Pratt (1964) has shown that the risk premium (rp) is a function of the consumer's degree of absolute risk aversion, $r(q)$ and the variance of income. Then it can be said that consumer A is (globally) more risk averse than consumer B if $rp_A > rp_B$ for all levels of wealth and variation in income.

Therefore, taking uncertainty into account the household's diet diversity choice can be modelled as

$$D_h = f_h(p', m_h, rp_h). \tag{13}$$

We empirically implement the household diet diversity demand function by specifying an estimating equation where household diet diversity (D) is explained by household risk premium (rp), income, prices, and household demographic characteristics (household size and composition, education level of household head, etc.); as controls we also add year, season, and region dummy variable sets. The household risk premium is not directly observable in our data and therefore we rely on our (compensated) price elasticity estimates at household level which we obtain from the QUAIDS analysis.To use the available information in the diet diversity regression analysis, we aggregate the estimated price elasticities into a single measure by the means of a conventional factor analysis,which produces a variable with standard normal distribution (see Rizov et al., 2015, Appendix 1 for details). Considering that household level compensated price elasticities capture the sensitivity of individual households to price and income shocks they appear to be a good proxy for household risk premium capturing both the household risk aversion and the variance of expected income faced by each household.[10]

3. DATA

We apply our methodology to the Slovak Household Budget Survey (HBS) data. The HBS data is commonly used for social policy and the standard of living analysis, for defining consumer price index weights, and for estimating household consumption in the national accounts. Our dataset consists of seven annual rounds, from 2004 to 2010. The survey provides detailed information on household incomes and expenditures on food and non-food goods and services. The data also contain detailed information on quantities consumed by each household, its location and size as well as individual household member characteristics such as age, education, occupation, marital status. Each of our annual samples contains approximately between 4500 and 6000 households, however, the samples do not form a (real) panel as surveyed households are randomly selected from the population each round.

The information on food consumption is collected on a one-month recall basis in four waves, one for each of the four seasons in the year. We aggregate food commodities consumed into five food groups: cereals, meat and fish, dairy products and eggs, fruits and vegetables, and other food products. The other food products group comprises of food commodities such as fats, oils, condiments, and sugar. Rizov et al. (2015), Appendix 2 provide details on the aggregation of food commodities into groups. As economic theory does not provide any guidance on the number or composition of aggregated food groups, the construction of

[9] Under uncertainty a risk-averse consumer will reduce spending on food and thus *ceteris paribus* reduce consumption. Following Jackson (1984) the reduction would occur at both the intensive and extensive margins, thus resulting in reduction in diet diversity.

[10] In a recent paper Liu et al. (2014) study diet diversity in China and emphasise the importance of access cost for consuming more diverse diet. Considering their theoretical framework the risk premium in our analysis can be seen as confining effects of uncertainty with transaction cost effects even though our theoretical foundation is consumer optimal behaviour under uncertainty.

the food groups used in this analysis was influenced partially by past studies of the European food sector and by a classification reflecting the similarity (substitutability) of food items from a consumer's viewpoint. A major advantage to our food-grouping scheme is that it reduces the total number of parameters in the model and avoids the problem with zero consumption, thus making the demand system estimation simpler.

Since prices were not provided by HBS, implicit prices for individual food commodities were derived from the purchased quantity and expenditure data. Price indices for the aggregated food commodity groups were computed using the geometric mean with expenditure shares as weights (e.g., as in Abdulai, 2002). Each price obtained is effectively a value to quantity ratio, which is called 'unit value' by Deaton (1989). The price calculated this way is household specific, representing household purchase decisions. Thus, the variation in food-group prices is due to differences in the composition of items (goods) consumed in each commodity group and variation in prices of each good across households. The latter could be due to quality differences, seasonal effects, and regional market conditions.

Cox and Wohlgenant (1986) argue that failure to adequately specify cross-sectional price effects could result in biased and misleading demand elasticities. This is because traditional Engel analysis may be inappropriate if prices are not constant in the cross section. In addition, prices in cross-sectional data are generally assumed to reflect quality effects which should be corrected for prior to estimation (Deaton, 1989). Specifically, price-income relationships are caused by differences in marketing services purchased; higher income households purchase more marketing services and, hence, pay higher average prices for commodities. Larger families generally pay lower average prices because of economies of size in purchasing and in household production-consumption activities. Cox and Wohlgenant (1986) propose a regression-based procedure for quality adjusting cross-sectional prices which is applied by several follow-up papers (notably, Park et al., 1996).

We follow the Cox and Wohlgenant's (1986) approach and quality adjust aggregate commodity prices in our data. However, instead of estimating regression residuals and then adding them up to regional price means we calculate median prices for narrowly defined sample segments whereby controlling for regional (supply), time (seasonality), and household characteristics variation. We define household segments by four quartiles of household net disposable income and size, as well as we control for presence of children in the household. The regional segments are formed by the eight main Slovak regions each divided into rural and urban component. Our approach has at least two advantages; it complies with the traditional Engel analysis wherequality adjusted prices are constant within narrowly defined segments and it avoids problems of estimated negative household prices.[11]

For our diet diversity analysis we compute two diet diversity measures as discussed in the methodology section using disaggregate food commodity consumption data. It is useful to consider more than one measure of variety, such as the count measure (CM)and the transformed Berry index (TBI).[12] An increase in CM would indicate that a household introduces new food commodities to its diet. However, TBI would provide information whether the new commodities and, possibly, other commodities are purchased in sufficiently larger amounts to affect the distribution of consumption shares. Van Trijp and Steenkamp (1990) provide an empirical comparison of methods for modelling diet diversity, and find only a weak correlation between measures similar to CMand TBI thus confirming our strategy to use the two measures. The evolution over time and correlation analyses of our two diet diversity measures, CM and TBI are presented in Rizov et al. (2015), Appendix 3.

In empirical studies, it is important to consider the time horizon over which diet diversity is measured. For ease of understanding, diet recommendations are often expressed in terms of a person's daily diet.

[11] Following Cox and Wohlgenant (1986) and Park et al. (1996) we estimated alternative quality adjusted prices; the QUAIDS results with these prices are similar to the results reported based on median prices at narrowly defined segments.

[12]Since the values of the Berry index (BI) lie in the interval between 0 and 1, the assumption of normality may not be fulfilled. To overcome this problem, a logistic transformation can be used (e.g., Greene, 1997) so that standard OLS regression can be estimated. The transformed Berry index (TBI) is $TBI = \ln\left[\frac{BI}{(1-BI)}\right]$.

However, references to daily intakes do not reflect the true goals of dietary recommendations which "apply to diets consumed over a reasonable period of time" (e.g., Shaw et al., 1996, p. 1). Moon et al. (2002) find that consumer preferences for diet diversity exhibit different patterns depending on the length of time allowed for consumption. Estimated correlation coefficients indicate that daily diet diversity deviates from that measured weekly and monthly as later two time dimensions appear to exhibit a similar pattern. Stewart and Harris (2005) adopt even one year time period in their analysis of fruit and vegetable diet diversity analysis. Therefore, our diet diversity measurescomputed on the monthly recall basis seem appropriate.

Table 1 reports summary statistics for the variables used in the QUAIDS estimations. It is evident that between 2004 and 2010 there was a significant change in real incomes and prices in Slovakia (in Rizov et al., 2015, Appendix 4 the evolution of the aggregate commodity group prices is also presented). Incomes almost doubled while the prices of cereals and diary increased more than twofold with prices of meat and fish, fruits and vegetables, and other food products increased more modestly which is reflected in the modest increase in total food expenditure. The household consumption patterns do not appear to have changed substantially over the period as evident from food expenditure shares which have remained quite stable as only the fruits and vegetables expenditure share shows a more significant increase. Detailed examination of the data suggests that the quantities consumed remained relatively stable too; the tendency for substitution of low-fat milk for whole milk is noteworthy though. This fact taken together with the noticeable increase in the fruits and vegetables expenditure share and the improvement in the diet diversity measures over time seems to indicate a shift of Slovak consumers towards a healthier diet which is an indicator of improved food security.

Table 1. Summary statistics of variables used in QUAIDS and diet diversity analyses

		2004		2010	
Variable	Definition	Mean	SD	Mean	SD
foodexp	Total monthly household food expenditure(€)	91.66	47.57	116.95	58.95
income	Net monthly household real income (€)	449.93	317.51	715.74	420.32
foodratio	Ratio of food expenditure and net income	0.24	0.13	0.19	0.12
$p_{cereals}$	Price of cereals (€)	0.81	0.15	2.22	0.22
p_{meat}	Price of meat and fish (€)	2.46	0.28	3.85	0.29
p_{dairy}	Price of dairy products (€)	1.30	0.28	2.78	0.35
p_{fruits}	Price of fruit and vegetables (€)	0.72	0.18	1.06	0.20
p_{other}	Price of other food (€)	2.01	0.50	3.05	0.71
$w_{cereals}$	Expenditure share on cereals	0.20	0.07	0.20	0.07
w_{meat}	Expenditure share on meat and fish	0.30	0.11	0.29	0.10
w_{dairy}	Expenditure share on dairy products	0.19	0.07	0.18	0.07
w_{fruits}	Expenditure share on fruits and vegetables	0.12	0.07	0.15	0.07
w_{other}	Expenditure share on other food	0.19	0.07	0.17	0.06
hh_size	Total household size	2.92	1.42	2.85	1.42
n_adults	Number of adults (above age 18)	2.22	0.97	2.44	0.82
n_children	Number of children (below age 16)	0.54	0.86	0.46	0.80
child	Dummy: 1 if a household has children	0.34	0.47	0.30	0.46
single	Dummy: 1 for a single member household	0.17	0.37	0.20	0.40
edu	Education of the household head; categorical scale from primary (0) to higher (3) education	1.99	0.52	2.03	0.49
gender	Gender of the household head; dummy: 1 if male	0.68	0.47	0.68	0.47
urban	Dummy: 1 if urban household and 0 otherwise	0.62	0.49	0.55	0.50
CM	Count measure of the food diversity	29.49	6.21	31.02	6.09

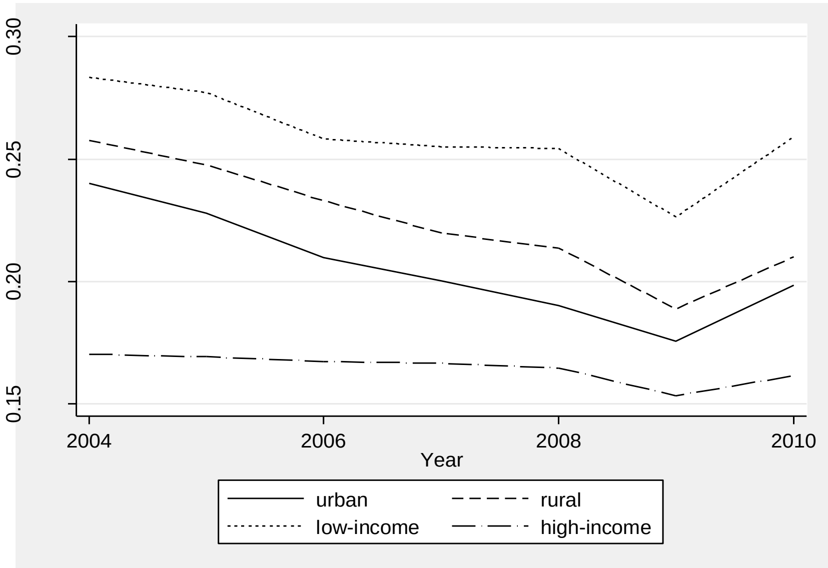
Table 1. Summary statistics of variables used in QUAIDS and diet diversity analyses, continued

		2004		2010	
Variable	Definition	Mean	SD	Mean	SD
TBI	Transformed Berry-index	2.47	0.34	2.55	0.31
rp	Risk premium (computed by factor analysis from the compensated own and cross price elasticities	1.47	0.48	0.41	0.17

Note: All monetary values were transformed to Euros from Slovak crowns with the corresponding exchange rate and were deflated with CPI (base 2000=100). There are eight regions in Slovakia, Bratislava, Trnava, Trencin, Nitra, Zilina, BanskaBystrica, Presov, and Kosice which are approximately equally represented in the survey.
Source: Household Budget Survey of Slovakia; authors's caltulations

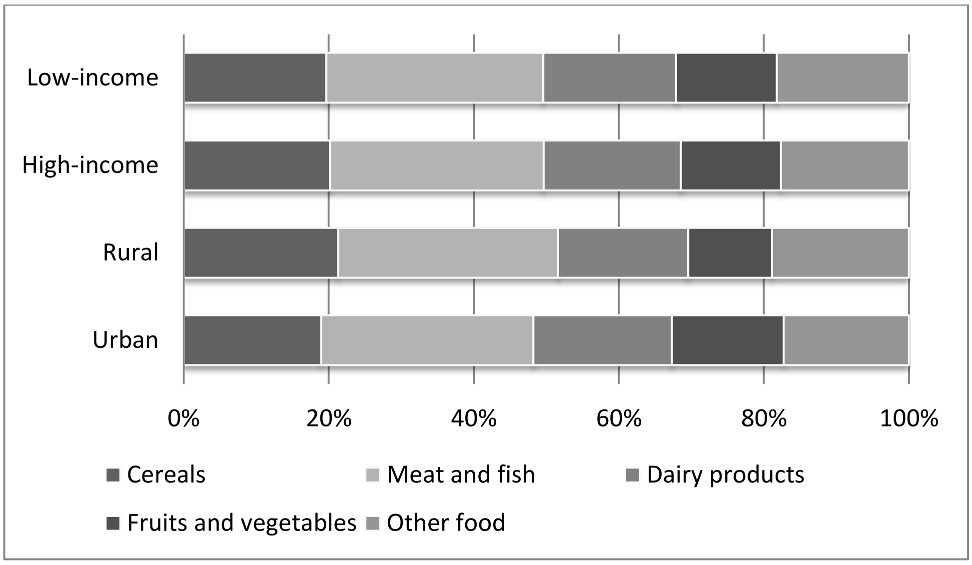
In terms of food security there is further evidence of improvement indicating the potentially important driving force – the rise of incomes. Figure 1 shows that the share of food expenditure in net income has been steadily declining since the Slovakia's accession to the EU in 2004. For the low-income subsample (households with income below the median) the ratio has dropped from 28% down to 23% in 2009 when the Euro was adopted, consequently followed by a modest hike in 2010. The trend for the high-income subsample is similar but the levels are quite different – the drop is from 17% to 15%, which is comparable with EU-15 levels. There are differences between rural (21% in 2010) and urban (20% in 2010) household food expenditure shares as these differences are less pronounced compared to the income-based subsamples while the declining trend is stronger confirming that the improvement in food security situation as indicated by the food expenditure share is a nationwide trend. There is also a relative homogeneity in terms of composition of the diet when comparing the rural and urban subsamples, and, interestingly, across income-based subsamples (see Figure 2).

Figure 1: Share of household food expenditure in net income, 2004-2010



Source: Household Budget Survey of Slovakia; authors's caltulations

Figure 2: Composition of the diet by household type, 2004-2010



Source: Household Budget Survey of Slovakia; authors's caltulations

4. Estimation and results

Our methodology underlines a two stage approach to the analysis of food security situation of Slovak households. To comprehensively analyse and understand the factors affecting both the access to food and the quality (diversity) of the diet we first estimate the price and income elasticities at household level which characterise the sensitivity of households to market shocks and thus the degree of households' constraints to access food. Second, we estimate diet diversity demand functions where key variables are household income and a measure of household's risk premium both describing the degree of households' constraints to consume diverse and healthy diet.

4.1. Food demand

We start our demand analysis by first estimating the Engel curves for the five food groups for the whole sample and by rural and urban subsamples using a non-parametric kernel regressionas in Banks et al. (1997); graphic presentation of the Engel curves can be found in Rizov et al. (2015), Appendix 5. The shapes of the Engel curves are consistent with the theory. An increase in income is associated with a monotonic decline in the share of expenditure on cereals while there is a positive relationship between income and the expenditure share of meat and fish suggesting that commodities from this food group are perceived as luxury. However, the patterns of the Engel curves for dairy products and for fruits and vegetables appear non-linear with inverted-U shape. The Engel curve for the other food products group is also highly non-linear. This preliminary analysis suggests that our choice of QUAIDS for estimating food demand behaviour in Slovakia is justified.

We estimate QUAIDS with Stata software using the code developed by Poi (2008; 2012). Parameter estimates are obtained for the full sample and for subsamples of rural and urban households and of low-income and high-income households by round. In the estimated samples large majority of own and cross-price parameters and linear expenditure parameters are statistically significant at conventional levels. The majority of the quadratic expenditure terms are also significant at 5% or better. Taken together the estimated expenditure parameters suggest that meat and fish, and for the rural households in early rounds also fruits and vegetables, are luxury. The demographic and regional control variables are generally significant and have the expected effects. For example, household size has a positive effect on the expenditure share of cereals and negative effect on the share of meat and fish. The effect of the expenditure ratio control is also highly significant in most equations and samples as it is, for example, positive in the cereals equations and negative in the meat and fish equations. The QUAIDS estimated parameters are reported in Rizov et al. (2015), Appendix 6.[13]

Table 2 reports compensated and uncompensated price elasticities and expenditure elasticities calculated from the QUAIDS parameters. These elasticieties are averages over the seven rounds (2004-2010) used. The expenditure elasticities of all food groups are positive as the largest in magnitude are the elasticities of fruits and vegetables (1.44) and meat and fish (1.22). Both compensated and uncompensated own-price elasticities are negative and thus consistent with demand theory. While all compensated own-price elasticities are smaller than unity in absolute value, the uncompensated own-price elasticities of meat and fish and fruits and vegetables are greater than unity revealing elastic demand. This finding is consistent with our results for expenditure elasticities and the effects of demographic variables and expenditure ratio. All compensated cross-price elasticities are positive albeit relatively small in magnitude suggesting that the respective food groups are weak substitutes, thus, confirming that our food group classification is appropriate.

Table 2. Average food demand elasticities, 2004-2010

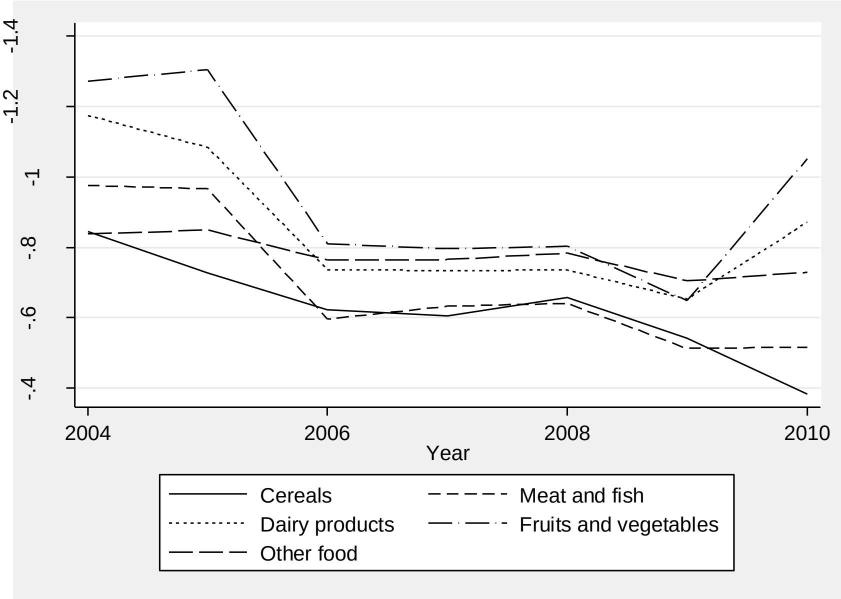
	C	MF	DP	FV	OF	
			Compensated price elasticities			Expenditure
C	**-0.61**	0.27	0.13	0.13	0.10	**0.92**
MF	0.21	**-0.69**	0.22	0.08	0.19	**1.22**
DP	0.08	0.40	**-0.86**	0.23	0.15	**0.68**
FV	0.30	0.04	0.35	**-0.96**	0.27	**1.44**
OF	0.05	0.39	0.12	0.22	**-0.78**	**0.73**
			Uncompensated price elasticities			
C	**-0.81**	-0.01	-0.04	0.00	-0.06	
MF	-0.04	**-1.06**	-0.01	-0.09	-0.04	
DP	-0.06	0.20	**-0.98**	0.13	0.03	
FV	0.01	-0.39	0.08	**-1.15**	0.01	
OF	-0.09	0.17	-0.02	0.12	**-0.91**	

Source: Household Budget Survey of Slovakia; authors's caltulations

The fact that the signs of several (thirteen out of twenty) compensated price elasticities are different from the signs of the uncompensated elasticities suggests that income effects are important in consumer demand decisions. The overall effect of price changes on demand responses is most relevant for capturing food security and aggregate welfare effects. Therefore, in Figure 3 we present the evolution of the compensated own-price elasticities for the five food groups over time. The general impression from Figure 3is that since 2004 the own-price elasticities have declined for all food commodity groups. This observation suggests that Slovak households have become less prone to food price shocks over the period of analysis. However, there is a pronounced hike in household price sensitivity around 2009-2010 – the period when Slovakia adopted the Euro currency and experienced effects from the global economic crisis.

[13] To formally test the validity of QUAIDS, we performed specification tests comparing restricted models with linear Engel curves for all food groups and the alternative models with quadratic Engel curves. The Chi-square tests rejected the restricted models in all samples. Similar tests confirm the validity of the demographic controls used. The test results are reported in Rizov et al. (2015), Appendix 7.

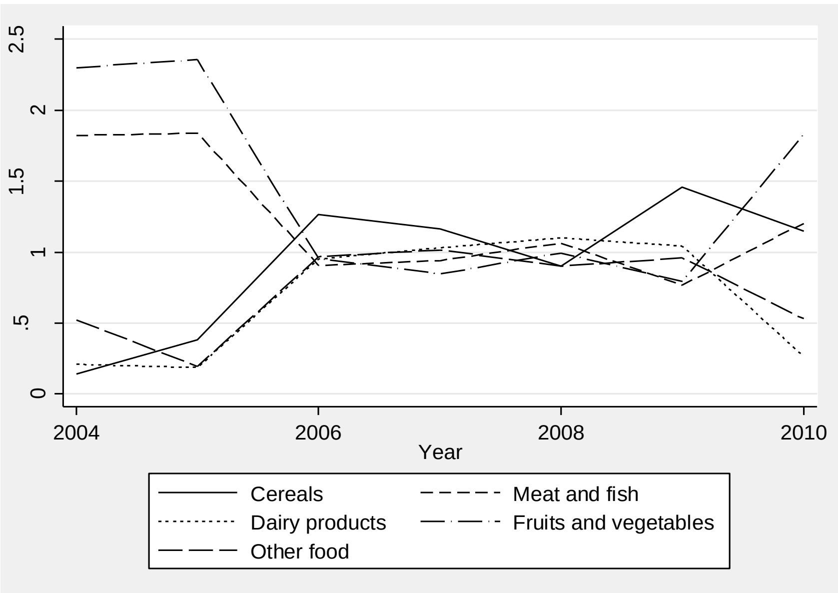
Figure 3: Compensated own-price elasticities, 2004-2010



Source: Household Budget Survey of Slovakia; authors's caltulations

Our results from the analysis by subsamples of households further demonstrate the substantial heterogeneity of demand responses. The compensated and uncompensated price elasticities and expenditure elasticities computed from the QUAIDS parameters for rural and urban and low-income and high-income households are reported in Rizov et al. (2015), Appendix 8. Generally, we can observe higher sensitivity and volatility of responses in the rural and low-income household subsamples throughout the period, since the Slovak EU accession. There is a substantial hike in the price sensitivity of meat and fish demand of low-income households since 2008, the beginning of the economic crisis. High-income households have experienced increased price sensitivity of their fruits and vegetables and meat and fish demand in the post-Euro period while urban household experienced similar effects on their demand for dairy products, fruits and vegetables and other food products.

Figure 4: Expenditure elasticities of aggregated food groups, 2004-2010



Source: Household Budget Survey of Slovakia; authors's caltulations

To sum up, an important result of our demand analysis is the observed reduction in price and expenditure elasticities over the period of analysis. Noteworthy is also the observed convergence of the five food group expenditure elasticities at relatively lower level as depicted in Figure 4. This suggests reduction in the relative income constraints on food consumption and diet composition choices. Following this logic one could argue that the quality of the diet has been improving over time with the convergence in the income elasticity magnitudes. We analyse the quality of the diet measured by diet diversity next.

4.2. Diet diversity

We estimate empirical specifications of the diet diversity function, Equation (13) for each of the two diversity measures - food count (CM) and transformed Berry index (TBI) - by the means of both OLS and quantile regressions (Koenker and Bassett, 1978; Koenker and Hallok, 2000).[14] The advantage of the quantile regression (QR) is that unlike the standard OLS which estimates the average relationship between the outcome variable and a set of explanatory variables based on the conditional mean function, QR describes the relationship at different points of the outcome variabledistribution. As it is likely that the effects of independent variables are different at different points of the diet diversity distribution the QR analysis is appropriate.[15] Furthermore, QR is more robust to non-normalerrors and outliers than OLS.

The estimation results from OLS and QR (for three quantiles – 0.1, 0.5, and 0.9) for both CM and TBI specifications are reported in Table 3 and show relationships consistent with theory. We find a significant positive effect of income on diet diversity as the OLS and QR median estimates are similar. Income has a stronger effect on diet diversity at lower quantiles. Following our theoretically motivated specification food prices are also included (as controls) and they indeed have significant effect on diet diversity in several cases however the directions of the effects is difficult to interpret as discussed by Thiele and Weiss (2003); therefore we do not discuss the price coefficients further.

[14] Only a few studies have applied quantile regressions for diet diversity analysis; for example Variyam et al. (2002) estimate demand for macronutriens in the USA and Drescher and Goddar (2011) analyse food diversity in Canada.

[15]Quantile regression can be specified as $Q_\theta(D|X) = X'\beta_\theta$, where D denotes the food diversity measure as a function of a set of independent variables, X within the θth quantile of the outcome variable *D*. The special feature of the quantile regression approach is that the set of coefficients of the independent variables, β_θ can differ across quantiles. The estimatorβ_θ of the quantile regression is obtained by minimizing the objective function $Q(\beta_\theta) = \sum_{i:y_i \geq x_i'\beta}^{N} \theta|y_i - x_i'\beta_\theta| + \sum_{i:y_i < x_i'\beta}^{N} (1-\theta)|y_i - x_i'\beta_\theta|$via Simplex method. We estimate our quantile regressions using Stata's qreg and sqreg commands and report bootsraped standard errors.

Table 3. Determinants of food diversity (pooled sample 2004-2010)

	ln(CM)				TBI			
Variable	OLS	Q(0.1)	Q(0.5)	Q(0.9)	OLS	Q(0.1)	Q(0.5)	Q(0.9)
ln(income)	0.03***	0.05***	0.04***	0.03***	0.06***	0.09***	0.06***	0.04***
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
ln($p_{cereals}$)	-0.12***	-0.18***	-0.10***	-0.05***	-0.01	-0.02	0.00	0.03**
	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
ln(p_{meat})	0.02***	0.03***	0.01	0.00	-0.06***	-0.05***	-0.06***	-0.06***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
ln(p_{dairy})	-0.04***	-0.02*	-0.05***	-0.04***	0.06***	0.07***	0.05***	0.04***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
ln(p_{fruits})	0.10***	0.14***	0.09***	0.05***	0.04***	0.06**	0.03**	0.03**
	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.03)	(0.01)	(0.01)
ln(p_{other})	0.08***	0.12***	0.07***	0.04***	0.02***	0.03**	0.01*	0.01
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
rp	-0.12***	-0.12***	-0.11***	-0.10***	-0.13***	-0.11***	-0.15***	-0.16***
	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
2005	0.04***	0.04***	0.05***	0.03***	0.04***	0.04***	0.03***	0.04***
	(0.00)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
2006	0.33***	0.34***	0.32***	0.28***	0.29***	0.28***	0.32***	0.36***
	(0.01)	(0.03)	(0.01)	(0.01)	(0.02)	(0.03)	(0.02)	(0.02)
2007	0.34***	0.35***	0.33***	0.28***	0.29***	0.28***	0.32***	0.35***
	(0.01)	(0.03)	(0.01)	(0.01)	(0.02)	(0.03)	(0.02)	(0.02)
2008	0.35***	0.37***	0.33***	0.28***	0.28***	0.27***	0.30***	0.33***
	(0.01)	(0.03)	(0.01)	(0.01)	(0.02)	(0.03)	(0.02)	(0.02)
2009	0.33***	0.32***	0.33***	0.30***	0.29***	0.27***	0.32***	0.35***
	(0.01)	(0.03)	(0.01)	(0.01)	(0.02)	(0.04)	(0.02)	(0.02)
2010	0.28***	0.29***	0.28***	0.25***	0.25***	0.23***	0.28***	0.32***
	(0.01)	(0.02)	(0.01)	(0.01)	(0.02)	(0.03)	(0.02)	(0.02)
qy2	0.01***	0.02***	0.01***	0.01**	0.03***	0.03***	0.03***	0.02***
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
qy3	0.03***	0.04***	0.02***	0.02***	0.02***	0.02*	0.01**	0.01*
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
qy4	0.01**	0.01	0.01**	0.01	0.01**	0.02**	0.02***	0.01*
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
TT	0.01	0.05***	-0.01	-0.02***	-0.02***	0.01	-0.02***	-0.05***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
TN	0.03***	0.09***	0.01**	-0.02***	0.00	0.04***	0.00	-0.02***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
NR	0.01**	0.06***	0.00	-0.03***	-0.05***	-0.01	-0.05***	-0.07***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
BB	0.00	0.04***	-0.01**	-0.03***	-0.02***	0.02	-0.02**	-0.05***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)

Table 3. Determinants of food diversity (pooled sample 2004-2010), continued

	ln(CM)				TBI			
Variable	OLS	Q(0.1)	Q(0.5)	Q(0.9)	OLS	Q(0.1)	Q(0.5)	Q(0.9)
PO	0.00	0.05***	-0.01**	-0.04***	-0.03***	0.02	-0.03***	-0.05***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
ZA	0.02***	0.07***	0.01**	-0.01***	0.01	0.05***	0.01	-0.02***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
KE	0.02***	0.06***	0.01**	-0.02***	-0.03***	0.01	-0.03***	-0.04***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
urban	0.03***	0.05***	0.04***	0.02***	0.09***	0.12***	0.08***	0.06***
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
edu	0.00	0.01	0.00	0.01**	0.02***	0.02*	0.02***	0.02***
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
hh_size	0.01***	0.01**	0.01***	0.00**	-0.02***	-0.02***	-0.02***	-0.03***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
single	-0.04***	-0.07***	-0.02***	-0.01***	-0.02***	-0.05***	-0.01	0.00
	(0.00)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)
child	0.01**	0.01	0.01	0.00	0.03***	0.04***	0.02***	0.02***
	(0.00)	(0.01)	(0.000	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)
cons	2.87***	2.40***	2.92***	3.25***	1.90***	1.28***	1.94***	2.38***
	(0.02)	(0.04)	(0.02)	(0.02)	(0.03)	(0.06)	(0.03)	(0.03)
N	33243	33243	33243	33243	33243	33243	33243	33243
(Pseudo)R^2	0.08	0.06	0.04	0.03	0.07	0.04	0.04	0.04

Robust standard errors are presented in parentheses; * p<0.05, ** p<0.01, *** p<0.001.
Source: Household Budget Survey of Slovakia; authors's caltulations

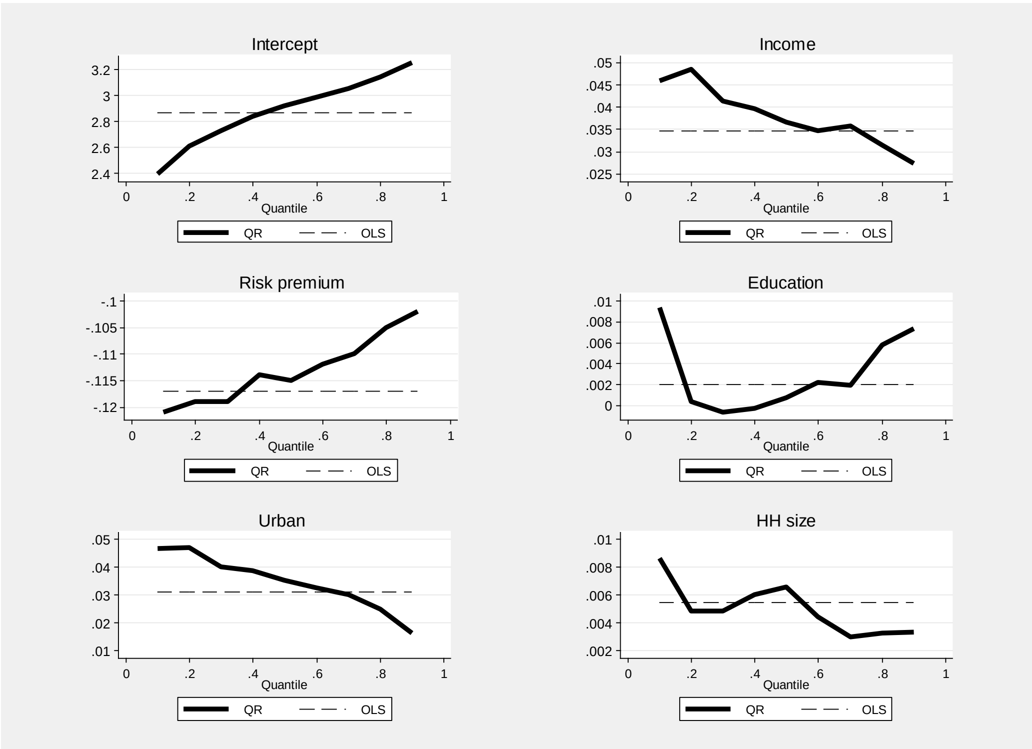
The variable of key interest according to our theoretical framework is the measure of the impact of uncertainty – the household's risk premium, *rp*. We find that the effects of *rp*are always significant negative as predicted by theory. The effects do not differ substantially between OLS and QR median estimates, while across quantiles the effects monotonically decline along the diet diversity distribution for the CM specification and increase for the TBI specification. The later finding is interesting and suggests that at the lower end of the diet diversity distribution households are more likely to adjust consumption at the extensive margin (reduce number of commodities consumed) rather than at the intensive margin (rebalance quantities consumed); the opposite behaviour is exhibited by households at the higher end of the diet diversity distribution.

Among demographic characteristics likely to influence household diet diversity, we consider the education level of the household head. The estimated effect is generally significant positive for the TBI specification as the OLS and QR median estimates are similar and the effect is stronger at higher quantiles of the diet diversity distribution. In the CM specification education does not show significant effect. Blisard et al. (2003) argue that better-educated consumers may be more aware of the importance of healthy eating and therefore spend money on more diverse (balanced) diet. Moon et al. (2002) find empirical evidence to support the argument. However, other studies have not found a strong relationship between education and diet diversity (e.g., Thiele and Weiss, 2003).

Another demographic variable – the size of the household – we find to have opposite effects on diet diversity depending on the measure. In the TBI specification the effect is significant negative, while it is significant positive in the CM specification. Thiele and Weiss (2003) argue that reconciling the effects of household size on diet diversity, measured by count and share-based measures, is complicated as the two measures reveal different aspects of a consumption pattern. Kinsey (1990) notes that larger households, with three or more children, are considered a prime market for the basic food ingredients, traditionally provided by (cheap) grocery stores. For (sufficiently) small households, Lee and Brown (1989) find that an increase in household size will expand the variety of a household's purchases. However, Lee and Brown (1989) also find that the effect of introducing an additional member to a household will be smaller as the total size of the household increases, and would even become negative in larger households. A possible explanation is that introducing an additional person to a large family would increase the difficulty of coordination in preparing foods acceptable to all family members, and thus lead to more simplified, uniform diet that fits to a variety of heterogeneous tastes in a large family. The estimation results for two additional variables characterising household demographics – dummies for single households and households with children are consistent with our previous results, on household size, as generally single households consume less diverse diet while the presence of children leads to increase in diet diversity measured by TBI; the later effect is the strongest at lower quantiles of the distribution.

In all estimated specifications we have included controls for time (year dummy variable set) and location (regional dummy variable set) which show significant effects. The main results are that diet diversity throughout the period of analysis is higher compared to the reference 2004 year and that relative to the capital city Bratislava diet diversity (measured by TBI) is lower in other regions, except Trencin and Zilina. The control for seasonality (a set of four dummies) also shows significant effects as diet diversity generally appears lower during the winter compared to other three seasons, thus suggesting that consumers may be constraint in accessing some food commodities during the winter months. In terms of differences between rural and urban locations, there is a pronounced divide in diet diversity as urban households appear to consume a more diverse diet. This result is robust to diversity measure used and estimation technique.[16]

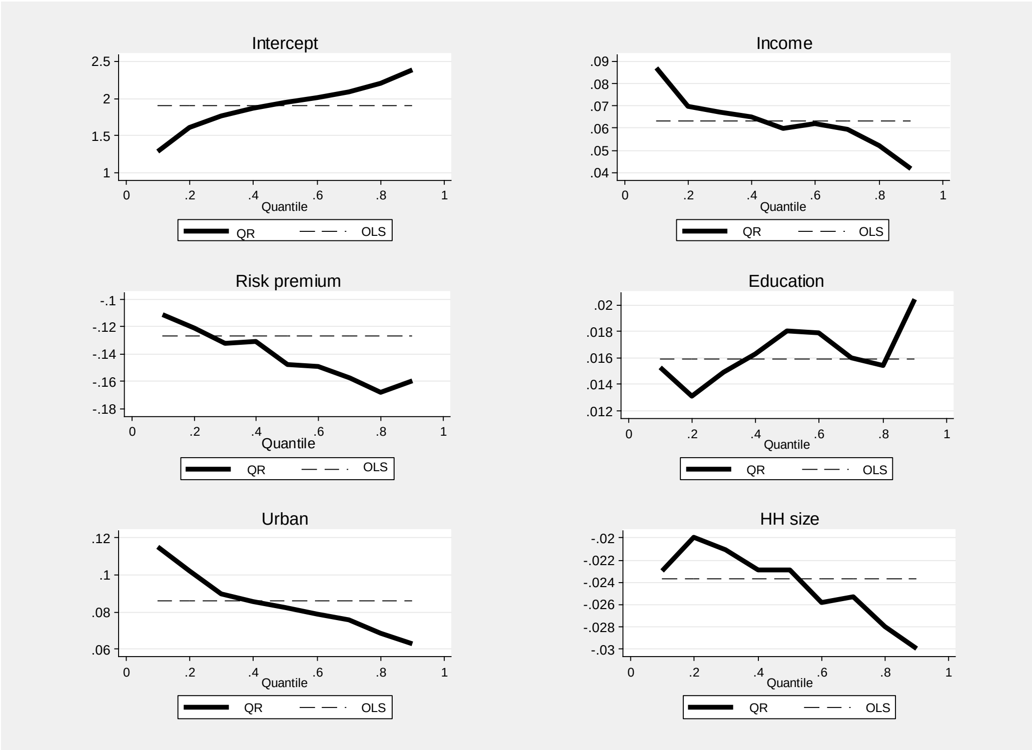
Figure 5.a: Effects of selected variables across quantiles, CM specification



Source: Household Budget Survey of Slovakia; authors's caltulations

[16] In Rizov et al. (2015), Appendix 9 we report Wald test results for the coefficient differences across quantiles, for both CM and TBI specifications. One can observe that large majority of key variables have differential effects across quantiles.

Figure 5.b: Effects of selected variables across quantiles, TBI specification



Source: Household Budget Survey of Slovakia; authors's caltulations

Figures 5a and 5b illustrate how the effects of income, risk premium, education, household size and location (rural vs. urban) on diet diversity vary over quantiles, and how themagnitude of the effects at various quantiles differ considerably fromthe OLS estimate. An exception is the effect of education which is relatively uniform along the diet diversity distribution. It is noteworthy that the intercept is quite large and increasing along the distribution showing that households in the higher quantiles *ceteris paribus* have stronger preferences for diverse diet.

5. CONCLUSION

We analyse the food demand patterns of Slovak households since the accession of Slovakia to the EU in 2004. Our study is one of the few food demand and diet diversity analyses for the new EU member states. Weapply a two-stage analytical framework where, in the first stage, we estimate QUAIDS and diet diversity quantile regressions, in the second stage, respectively. The Slovak longitudinal BHS data employed covering seven year period allow us to reveal changes in demand behaviour over time as well as cast light on the food security situation at micro level. In terms of food security a noteworthy nationwide trend is the continuous reduction in the food expenditure and income ratio. By 2010 the food expenditure ratio has dropped to about 16% for high-income households – a level comparable with demand patterns in the richer EU-15. The ratio is still quite high though, at about 26% for the low-income households.

Our first stage results show that Slovak households are price and income responsive as food expenditure patterns vary across types of household. All five food groups analysed have positive expenditure elasticities as their magnitudes suggest that cereals, dairy products and other food products are necessities while fruit and vegetables and meat and fish are luxuries for some groups of households. In line with demand theory, all own price elasticities are negative while a significant number of the cross-price elasticities are positive albeit smaller in magnitude suggesting that even though the commodities from the five food groups are substitutes the substitution possibilities might be quite limited. Furthermore, the results from subsamples by household type reveal that the demand sensitivity of low-income and rural households is higher compared with high-income, urban households.

In the second stage of our analysis we find that the diet diversity, measured by both food item count and Berry index, has been increasing since 2004 indicating again improving food security situation of Slovak households. Besides tastes which seem to be very important, income has a strong impact on diet diversity while the risk premium, proxied by an aggregate factor of the compensated price elasticities has a strong negative effect on diet diversity. The household demographic characteristics generally have trivial effects on diet diversity as expected. It is noteworthy that there is a pronounced seasonal pattern with lowest diversity of the household diet during winter months – finding suggesting that there are possible binding supply side constraints during that period. We also find dietary differences between rural and urban locations as well as between the capital city and the rest of the country, with notable exceptions. These later findings are consistent with the supply side constraint hypothesis.

Our findings are generally consistent with studies from other developed countries, where food security does not present a significant challenge. For example, Michalekand Keyzer (1992), Abdulai (2002), and Chern et al. (2003) find that for majority of the population food demand is price and income inelastic and food is perceived as necessity rather than luxury while diet diversity is positively affected by income and certain demographic characteristics (Thiele and Weiss, 2003; Drescher and Goddard, 2011).Considering the fact that in Slovakia average expenditure elasticities for all food groups surpass in magnitude the own-price elasticities, policy tools for enhancing income generating activities might be more effective compared to policies that are targeted at price reductions. Income-generation oriented policies would also be consistent with our second stage results where income has strong positive effect on diet diversity while the risk premium which is decreasing in income has a negative effect. Hence, in order to improve the household access to food and achieve diverse (and healthy) diet income-generation oriented policies would be appropriate which should also be complemented with policies for rural development and improvement of the food supply chains.

A final point on the generalizability of our findings and policy recommendations, considering the fact that Slovakia has been one of the most economically successful NMS during the period of analysis, the food security situation in other east European NMS could be relatively less optimistic. In support of the later conjecture is also the fact that in recent years Slovakia has had one of the lowest levels of inequality in the EU as measured by the Gini coefficient indicating relatively more favourable general welfare conditions while several other NMS such as the Baltic states, Bulgaria, and Romania rank quite high in terms of inequality.

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